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Competition in Large Markets

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Abstract

This paper evaluates the simplifying assumption that producers compete in a large market without substantial strategic interactions using nonparametric regressions of producers' choices on market size. With such *atomistic* competition, 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. I examine observations of restaurants' sales, seating capacities, exit decisions, and prices from 222 U.S. cities. Given factor prices and demographic variables, increasing a city's size increases restaurants' average sales and decreases their exit rate and prices. These results suggest that strategic considerations lie at the heart of restaurant pricing and turnover.

See Tables/table1.m

JEL Classification: L11, L81

Keywords: Atomistic Competition, Market Size, Free Entry, Exit, Nonparametric Test.

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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. This paper quantifies this informal suspicion 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 result is familiar from highly stylized models of Chamberlinian monopolistic and perfect competition, and this paper proves it in a very general framework without substantial restrictions on the market demand system, producers' cost functions, or the variables over which they compete. It requires nonparametric regressions to display no influence of market size on producers' choices given a sufficiently rich set of control variables. In fact, restaurants in larger cities have lower prices, exit less frequently, and have greater sales revenues. Even if one finds atomistic competition implausible ex ante, the theory and regressions together quantify how the strategic environment influences producers' observable choices. Such quantification is essential for extending the domain of strategically-oriented empirical industrial organization to large markets.

See Tables/table1.m

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 market twice as large with double the number of producers and the same distribution of their observable actions. Of course, the distribution of producers' actions could differ across large and small markets even without substantial strategic interaction if the production technology or consumers' tastes systematically change with market size. I eliminate dependence of demand and costs on market size in the model by assumption; and the regressions control for differences in production possibilities and consumers' tastes across U.S. cities with factor prices and demographic measures.

Standard Chamberlinian monopolistic competition cannot capture the idea that increasing market size makes competition "fiercer" by increasing the number of producers, but true oligopoly models with strategic interaction raise issues of dynamic game theory that are not necessarily central to a particular author's problem. This difficulty has led some authors to use a model with a continuum of producers and goods due to Ottaviano et al. (2002), in which the elasticity of any given good's demand decreases with the number of goods offered even though no two producers compete head-to-head. Asplund and Nocke (2006) and Nocke (2006) find that an otherwise-standard model of industry dynamics with this specification predicts that producer entry and exit rates *increase* with market size. I find the opposite to be true for U.S. restaurant markets. Apparently, the dynamic aspects of oligopolistic interaction that the model of Asplund and Nocke omits are unimportant for the industry they examined but nevertheless substantially lower restaurants' exit rates in larger markets.

Since models of atomistic competition predict that producer choices are invariant to market size, their testing can proceed even if either the available measure of market size is poorly measured or the market definitions are based on geography rather than cross-elasticities of substitution in demand. These practical difficulties reduce the test's power but do not change its size. Since its purpose is to uncover information about the competitive environment, the test also requires no information about how competitors strategically interact. For these reasons, it can serve as a first step towards adding oligopolistic interaction to the already rich literature that examines producer dynamics and productivity growth with census data. Nevertheless, it 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 (2009) 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* from the same sample of *MSA*s examined by Campbell and Hopenhayn (2005). The observations cover Restaurants, defined by the Census as those Eating Places which provide table service. Thus, the analysis excludes familiar counter-service chains like McDonalds, which are included instead in Refreshment Places.[1] Henceforth, the paper uses the improper noun "restaurants" to refer to the industry under study.

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*.[2]

[1] Results from applying this paper's analysis to data from Refreshement Places are summarized below.

[2] It would be desirable to examine more recent observations. Unfortunately, the Census has not

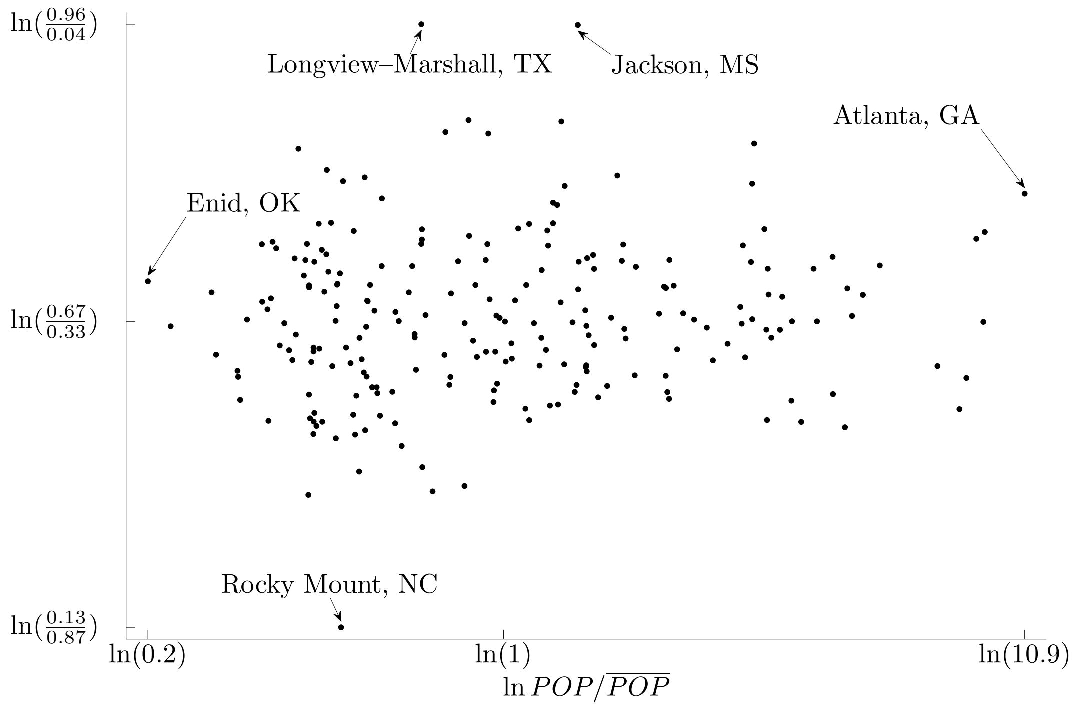

Figure 1: Logistic Transformation of the Share of Establishments with High Meal Prices(i)

Notes: (i) The figure plots the logistic transformation of the share of establishments with typical meal prices exceeding \$5.00 against the demeaned logarithm of *MSA* population.

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From these observations, I 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

published *MSA* level observations of these variables from the two most recent published Economic Censuses (1997 and 2002) as of this paper's completion (August 2009). The Miscellaneous Subjects volume for the 2007 Census of Retail Trade is scheduled to be published in December, 2010.

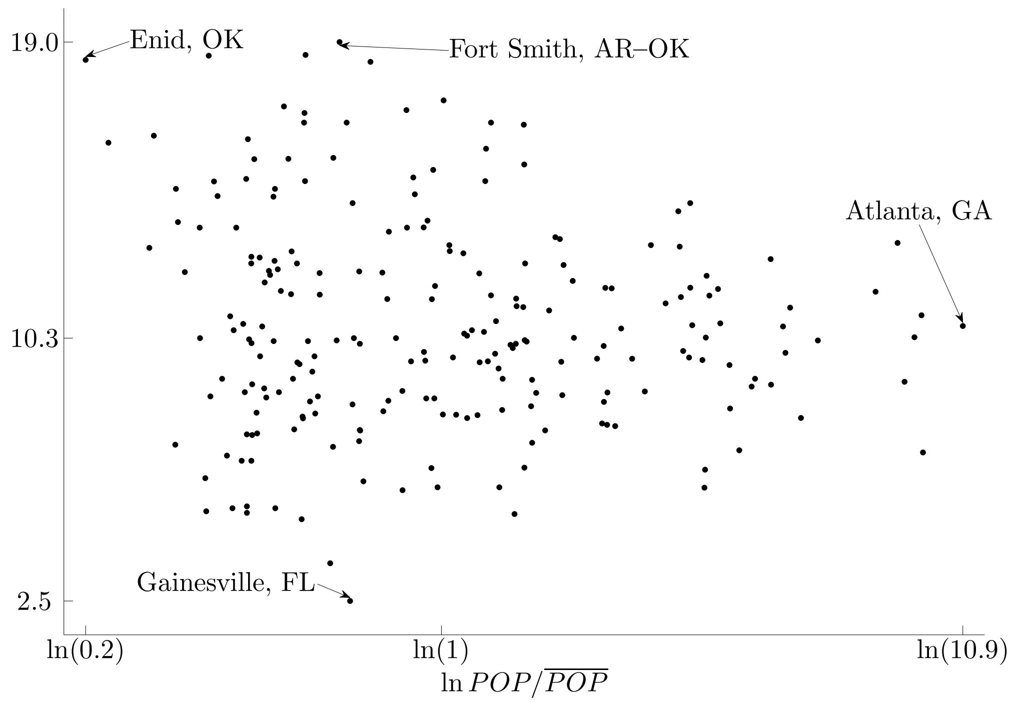

Figure 2: Restaurants' Annual Exit Rate in Percentage Points

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observations with extreme values 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. The correlation between the logistic transformation of $\mathbb{S}(\$5.00)$ the logarithm of *MSA* population equals 0.09.

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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 percent. The correlation between these variables equals −0.12.

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The exact exit rate for Enid, OK equals 18.52 percent. The exact maximum equals 19.05 percent.

The exact exit rate for Atlanta, GA equals 10.65 percent. The exact median equals 10.29 percent.

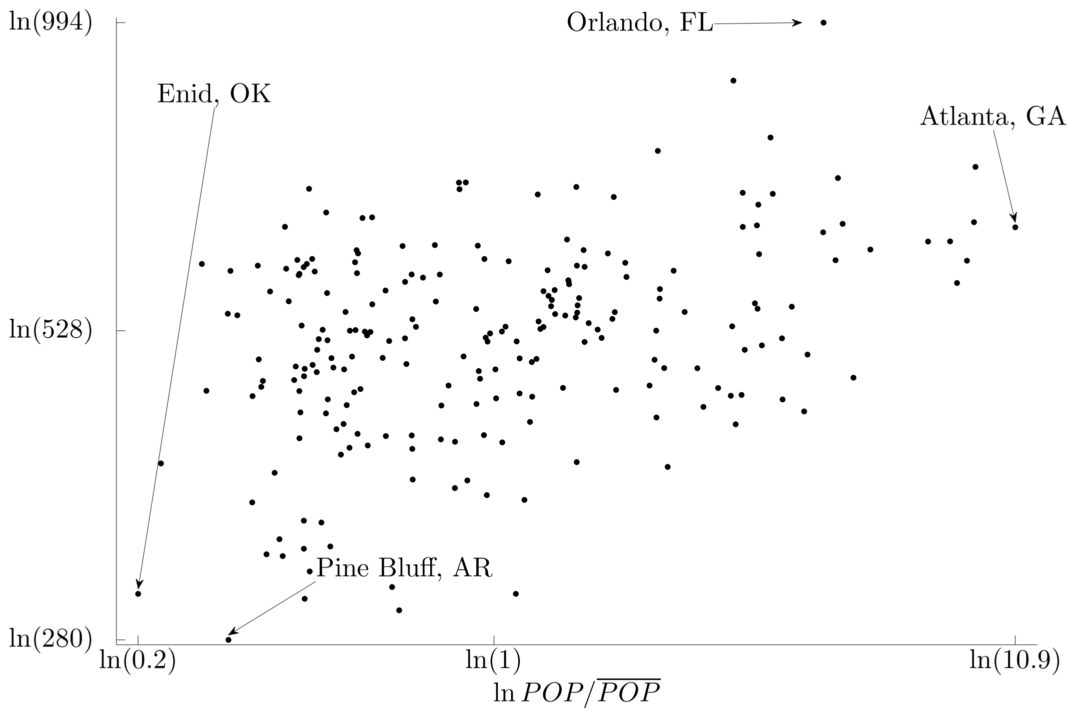

Figure 3: Logarithm of Restaurants' Average Revenue

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The other two variables of interest both measure average restaurant size, restaurants' average revenue and average seating capacity. 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 is evident in Figure 3. Figure 4 reveals little correlation between *MSA* population and average seating capacity.

Please see Table A1 for these four variables' conventional summary statistics.

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 demand associated with income, female labor-force participation, race, and education that could shift the the nature of producers' products and thereby

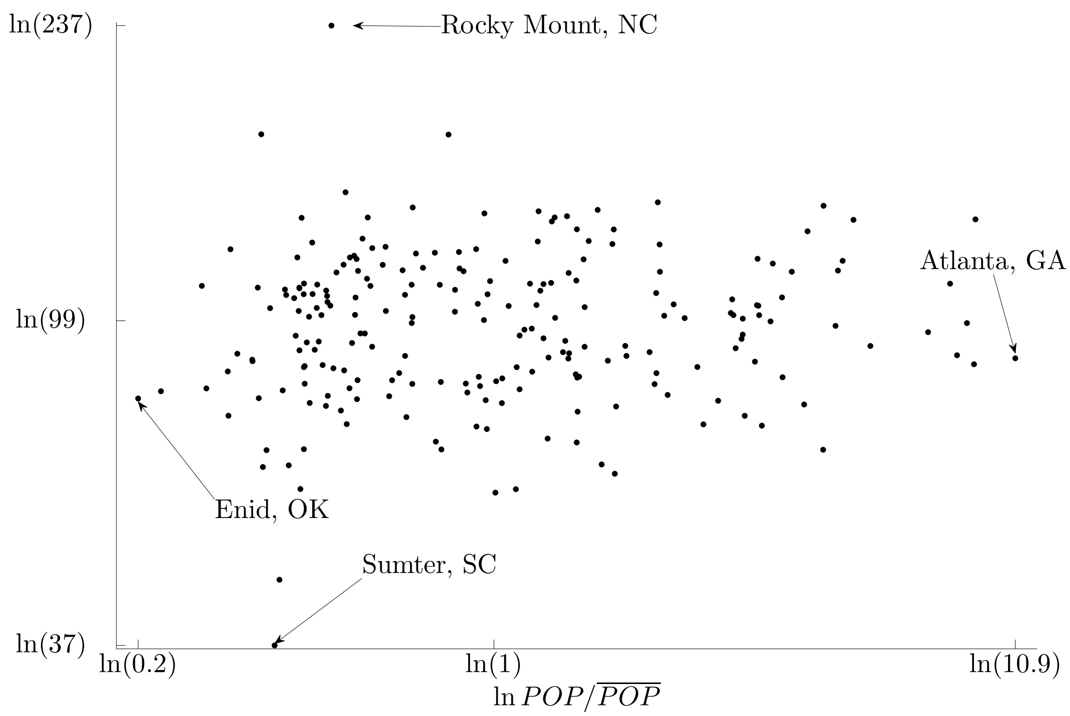

Figure 4: Logarithm of Average Seats per Restaurant

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indirectly influence their observable decisions. 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.[3]

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

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

where $f\left(\ln S, W\right)$ is the joint density function of $\ln S$ and W across markets and

[3] 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 equivalent to assuming that $\mathbb{S}_i(\$5.00) = e^{m(\ln S_i, W_i)+U_i}/(1+e^{m(\ln S_i, W_i)+U_i})$.

expectations are taken with respect to the same joint density function. Powell et al. (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. Powell, Stock, and Stoker's (1989) 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. Because this bandwidth applies to *all* of the regression's explanitory variables, the elements of W were all scaled by the standard deviation of $\ln S$, which equals 0.86. To increase the precision of the estimates' reports, *all entries in the table and in the text have been multiplied by* 100.

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The estimate of δ_S for the regression of $\mathbb{L}(\$5.00)$ equals -13.64 and is statistically significant at the 5 percent level. Thus, restaurants in larger markets charge *lower* prices given factor costs. To gauge the magnitude of this coefficient, consider doubling S given that $\mathbb{S}(\$5.00)$ is initially at its median value of 0.67. If we assume that $\partial m(\ln S, W)/\partial \ln S$ is constant, then this decreases $\mathbb{S}(\$5.00)$ to 0.65. The hypothesis that increasing market size lowers prices permeates empirical industrial organization, but to date only Syverson (2007) has verified that this is so for a particular industry. This finding that typical meal prices fall with market size complements his results.

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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. All of the demographic variables but mean personal income influence prices.

The estimate of δ_S for the exit rate is also negative, -0.70, and statistically significant at the 5 percent level. This implies that doubling S decreases restaurants' exit rate by 0.49 percentage points. As Campbell and Hopenhayn (2005) document, an increase in $\ln S$ strongly raises restaurants' average revenue. The estimated coefficient is 4.84, and it is statistically significant at the one-percent level.[4] The final dependent variable is the logarithm of average seats per restaurant. The estimated coefficient is positive, 1.82, but its standard error equals 1.97. Hence, these observations are

[4]This estimate differs greatly from that reported by Campbell and Hopenhayn (2005) for a similar regression. The discrepancy between the two largely reflects an error in Campbell and Hopenhayn's calculations. An erratum to that paper available at `http://dx.doi.org/10.1111/j.1467-6451.2007.00315.x` corrects that error.

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

	$\mathbb{L}(\$5.00)^{(iv)}$	Exit Rate	Revenue	Log Average Seats per Restaurant
Population	-13.64**	-0.70**	4.84***	1.82
	(5.93)	(0.31)	(1.77)	(1.97)
Commercial Rent	10.56*	-0.29	2.00	-2.29
	(5.97)	(0.31)	(1.53)	(1.85)
Retail Wage	14.63**	0.46	-0.84	-1.94
	(6.98)	(0.34)	(1.59)	(2.16)
Advertising Cost	-8.68	-0.28	-1.57	-0.56
	(5.40)	(0.26)	(1.66)	(1.95)
Income	6.72	-0.01	6.60***	3.53
	(6.78)	(0.36)	(1.98)	(2.54)
Percent Females in Labor Force	-16.27**	-1.04***	-1.33	2.49
	(6.36)	(0.34)	(1.88)	(2.50)
Percent Black	21.74***	0.75***	1.51	-3.91*
	(5.54)	(0.25)	(1.23)	(2.03)
Percent College	26.56***	0.22	8.80***	1.69
	(6.92)	(0.39)	(1.38)	(2.05)

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 $\star$, $\star\star$, and $\star\star\star$ indicate statistical significance at the 10, 5, and 1 percent levels (ii) The regressors are Population, the logarithm of MSA population in 1992; Commercial Rent, Median Rent per Square Foot for Strip Malls; Retail Wage, First-quarter Retail Payroll/March Retail Employment; Advertising Cost, the cost per 1,000 exposures of a column-inch in a Sunday newspaper; Percent Females in Labor Force, the percentage of women 16 and over in the labor force; Percent Black, the percentage of people who report that they are Black; and Percent College, the percentage of those 25 and over with a College Degree. The Commercial Rent and the Advertising Cost data come from Campbell and Hopenhayn (2005), and the Retail Wage is calculated from data in the 1992 Census of Retail Trade. Observations of the remaining variables come from calculations using the 1994 County and City Data book. (iii) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iv) Here, $\mathbb{L}(\$5.00)$ refers to the logistic transformation of the fraction of restaurants in an *MSA* with typical meal prices greater than or equal to \$5.00.

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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.

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The estimates in Table 1 depend on the particular measure of market size (population) and the bandwidth. 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 -10.70 with population density, and this is statistically significant at the ten percent level. 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 estimate differs little from its original value (-11.97 versus -13.64) but the increased standard error makes the estimate statistically insignificant.

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Relevant variables omitted from the regression could substantially influence the estimates, so Table 3 presents the estimates of interest for specifications with additional regressors. The first column reports estimates from a specification which adds the number of vehicles per household to the regression's right-hand side. For $\mathbb{L}(\$5.00)$, the coefficient falls to -12.02 and the standard error rises to 6.24. In consequence, the estimate's statistical significance falls to the ten percent level. Since the (unreported) coefficient on vehicles per capita itself is statistically insignificant, the standard error's increase appears somewhat symptomatic of an included irrelevant variable. Nevertheless, adding this variable leaves the inferences from the other regressions unchanged. It is reasonable to suppose that the division of demand between visitors and residents can influence the organization of a market, so the second column reports estimates from a model which includes a proxy for visitor density, the dollar value of hotel sales per capita. This variable itself has a positive and statistically significant effect only on Restaurants' Average Revenue. Just as when adding vehicles per capita, the coefficient for $\mathbb{L}(\$5.00)$ falls and its statistical significance drops to the ten percent level.

Table A2 reports the corresponding coefficient estimates for the additional regressors.

Table 2: Estimates with Alternative Market Size and Bandwidth Choices$^{(i,ii)}$

	Market Size Measures			Bandwidth Choices	
	Population	Population Density	Housing Units	$h = 1$	$h = 3$
$\mathbb{L}(\$5.00)^{(iii)}$	-13.64**	-10.70*	-13.16**	-11.97	-12.41**
	(5.93)	(5.53)	(6.02)	(7.83)	(5.73)
Exit Rate	-0.70**	-0.79***	-0.69**	-0.94**	-0.73**
	(0.31)	(0.28)	(0.31)	(0.39)	(0.30)
Log of Restaurants' Average Revenue	4.84***	4.42***	5.16***	5.80***	4.56***
	(1.77)	(1.68)	(1.79)	(1.97)	(1.70)
Log of Average Seats per Restaurant	1.82	2.33	2.16	2.82	1.97
	(1.97)	(1.89)	(2.02)	(2.54)	(1.89)

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. The underlying regressions include the right-hand side variables listed in Table 1. 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) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iii) 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.

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Table 3: Estimates with Additional Control Variables$^{(i,ii)}$

	Vehicles per Household	Hotel Sales per Capita	Percentage Students	Percentage Elderly	Percentage Low-Income	Percentage High-Income	Population Growth Rate
$\mathbb{L}(\$5.00)^{(iii)}$	-12.02*	-10.68*	-14.36**	-14.07**	-14.08**	-19.25***	-14.44**
	(6.24)	(6.26)	(5.91)	(6.08)	(5.98)	(6.48)	(5.94)
Exit Rate	-0.73**	-0.73**	-0.72**	-0.71**	-0.71**	-0.46	-0.70**
	(0.32)	(0.32)	(0.31)	(0.31)	(0.31)	(0.31)	(0.31)
Log of Restaurants' Average Revenue	4.94***	4.14**	4.90***	5.02***	4.68***	5.38***	4.20***
	(1.78)	(1.65)	(1.78)	(1.82)	(1.77)	(1.89)	(1.62)
Log of Average Seats per Restaurant	2.37	0.67	1.86	2.16	1.96	1.47	1.69
	(2.00)	(2.02)	(1.97)	(2.00)	(2.01)	(2.21)	(1.97)

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. The underlying regressions include the right-hand side variables listed in Table 1. 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) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iii) 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.

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The next four columns report results from regressions which include the percentages of College Students, the Elderly (65 years or older), Low-Income households (with incomes below the poverty line) and High-Income households (with income at or exceeding \$75,000). Only one of these variables (the percentage of High-Income households) has any impact on a reported inference (in the regression of the Exit Rate). The regression behind the final column's estimates includes the rate of population growth from 1980 to 1992. The theoretical capacity for population growth to influence the regressions is clear, but its addition does not greatly change the coefficients of interest. Overall, Table 3 indicates that the basic findings do not reflect a very obvious omission of a relevant variable.

I have undertaken four 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 -9.67, which is statistically significant at the ten percent level. 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 economically or statistically significant 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. The resulting coefficient (expressed in percentage points) equals -2.55 and is statistically significant at the five percent level. Finally, I have also constructed analogues of Tables 1, 2, and 3 for a sister industry, Refreshment Places. These are Eating Places *without table service*. Those estimates reveal no measurable effect of market size on typical meal prices, and the effect of market size on the exit rate changes to -0.51, which is statistically significant at the ten percent level. The remaining inferences from the data with restaurants also apply to refreshment places.

Table A3 reports these estimates and the results of estimating the linear model with instruments to account for possible measurement error in shopping-center rent.

The estimated coefficient equals 3.93 with a standard error of 5.47. Table A4 reports the complete regression estimates.

See Table A4.

See Tables A6, A7, and A8.

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3 A General Model of Atomistic Competition

Understanding results such as these from competition in large markets requires a model with many competitors. Some, such as the Salop (1979) circle model, highlight strategic interaction. I call these models of "Hotelling-style" monopolistic competition. Others, such as Dixit and Stiglitz (1977) and Spence (1976) place firms with price-setting decisions into anonymous competition without substantial stragetic considerations. I call these models of "Chamberlinian" monopolistic competition. In this section, I define and explore *atomistic competition*, which unites price-taking perfect competition with Chamberlinian monopolistic competition.

The results of the previous section clearly conflict with very basic models of atomistic competition. For example, in Dixit and Stiglitz (1977) the free-entry condition requires each firm's sales to equal the product of the exogenous fixed cost with consumers' constant elasticity of demand. Doubling the number of consumers leaves producers' average sales unchanged. In this section, I show that the abstraction from strategic interaction is the sole source of the conflict between such models and the data. Their 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.

3.1 Two Specific Examples

To make following the general model easier, this section begins with two specific examples of monopolistic competition. The first features Chamberlinian monpolistic competition, and the second has Hotelling-style monopolistic competition. The second example highlights how the absence of strategic interactions in the first generates scale invariance. The remainder of the section presents the general model of atomistic competition and its analysis, referring back to these examples to explain its moving parts.

3.1.1 Chamberlinian Monopolistic Competition

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 mass 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. 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 mass 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 and decreasing in N, so there exists a unique value of N that equates the ex-ante value of a new entrant with the entry cost.

As specified, the free-entry equilibrium displays scale invariance. Since each period's profit only depends on the individual firm's quality and on the ratio of the number of consumers to the number of active producers, the payoff functions $V_L(S, N)$ and $V_H(S, N)$ are homogeneous of degree zero. From this, it follows immediately that the free-entry equilibrium displays scale invariance: Doubling the number of consumers doubles the number of initial entrants and leaves the distributions of producers' decisions unchanged.

3.1.2 Hotelling-style monopolistic competition

The example of Hotelling-style monopolistic competition departs only slightly from the previous example; by replacing random matching of consumers with restaurants with the random matching of consumers with submarkets. The mass of submarkets is K. Each restauranteur chooses one submarket to serve upon entry. Retauranteurs post prices simultaneously, and the submarket's consumers make their purchases from the restaurant with the lowest quality-adjusted price. For simplicity, assume further that $i' = \infty$ so that firms' continuation decisions are trivial. Under these conditions, a restauranteur serving a submarket without competitors earns $(S/K) \times (w \times \pi_H + (1 - w) \times \pi_L)$. A restauranteur facing one or more competitors in its submarket earns a positive profit if and only if it is the only firm with $q = q_H$. In this case, the fortunate restauranteur charges slightly less than $p^O = \min\{q_H m/q_L, p_H^\star\}$, where $p_H^\star$ is a high-quality restaurant's unconstrained monopoly price. Since competition might constrain this firm's price, its profit per customer π^O cannot exceed π^H.

Because the submarkets are identical, the number of firms in the free entry equilibrium rises in steps as S increases. Define $\overline{S}_1 = i \times K/(w \times \pi_H + (1 - w) \times \pi_L)$. When $S = \overline{S}_1$, the return to operating as a monopolist equals the sunk cost of entry. Continuing for $M > 1$, define $\overline{S}_M = i \times K/(w \times (1 - w)^{M-1}\pi^O)$. If $S = \overline{S}_M$, then entry into a market with $M - 1$ competitors yields an expected profit equal to the sunk cost. These threshold values of S completely describe the free-entry equilibrium. If $S < \overline{S}_1$, then no firms serve any market, and if $\overline{S}_M \leq S < \overline{S}_{M+1}$, then exactly M firms serve each submarket.

The influence of market size on observed producer behavior can be appreciated most quickly by considering average prices. When S lies between $\overline{S}_1$ and $\overline{S}_2$, restaurants charge monopoly prices. Increasing S to a point between $\overline{S}_2$ and $\overline{S}_3$ induces a second entrant into each submarket. With probability $1 - 2 \times w(1 - w)$, Betrand competition produces a zero markup over the firms' common marginal cost, and with the complementary probability the markup is that associated with p^O. As long as q_H/q_L is small relative to the low-quality restaurant's monopoly markup, the resulting average price is below the average monopoly price. To summarize, increasing market size makes the submarkets more competitive and thereby lowers prices. Concomitantly, firms' average sales must rise with market size so that they can recover their sunk costs.

3.1.3 Empirical implications

Both of these examples are far too stylized for empirical work, but suppose for the moment that the Chamberlinian example 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 Chamberlinian 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 Hotelling-style example is at least qualitatively consistent with producers in larger markets charging lower prices, so it is at least possible to match the data better with such a model. The analysis of the general model demonstrates that the first example's inconsistency with the data arises from assumption of Chamberlinian competition rather than one of its other simplifying assumptions. Therefore, Hotelling-style competition marks out the only path towards better understanding these observations.

3.2 The General Model

The general model of atomistic competition 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. That 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 Chamberlinian example did not make restaurateurs' product choices explicit, but this can be remedied by assuming 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 Chamberlinian example, a represents a restaurant's early and late prices and its continuation probability and Z determines restaurants' qualities.[5]

For simplicity, we assume that if two or more entrants chose x, they both choose the same post entry action.[6] 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 $c(a', x'; A, Z, F, S)$.

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 Chamberlinian 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 are

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

[5]The Chamberlinian 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 low quality if the clockwise distance between x/N (interpreted as a location on this circle) and Z is greater than w or if $x > N$. 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 Chamberlinian example.

[6]As in the specific Chamberlinian 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.

Here, the restaurant's quality ($q(x,Z)$) is the function described in Footnote 5.

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.[7]

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*,*

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

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

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

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

Condition *(a)* of this definition ensures that the action profile $A^\star(x;Z,F)$ 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 $(A^\star, F^\star)$ to correspond to a subgame perfect Nash equilibrium with pure strategies in the entry stage.

3.2.1 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.

[7] 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 $\mathbb{R}^k$, 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 subgames, working with the strategy profile as defined here simplifies the model's exposition.

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 $\mu(M)$ to denote the Borel measure of $M \in \mathcal{M}$.

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

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

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

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

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 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 restauranteur'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 a real 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; and competition is "global" in the sense of Anderson and de Palma (2000). The Chamberlinian example satisfies this assumption. Any model in which firms' producing exactly the same product act as Bertrand competitors will satisfy Assumption A2.[8]

[8] A model with price-taking producers of a homogeneous good, such as Hopenhayn's (1992), could costlessly accommodate this assumption by defining a trivial product placement choice x on the real

Hart (1985) and Wolinsky (1986) propose a 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).(Wolinsky, 1986, page 493)

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) are also models of atomistic competition. However, the definition of atomistic competition is broad enough to encompass models without market power.

3.2.2 Intrinsic Scale Effects

Thus far, the model's specification does not rule out direct effects of the scale of the market on the model's primitives. Furthermore, there is no guarantee that a free-entry equilibrium even exists. For these reasons, this section proceeds now to present conditions on the model which address these and other economically uninteresting obstacles. I classify them as *regularity* conditions, because most useful models of competition between many producers satisfy them.

The first three regularity conditions eliminate direct effects of market size on the model's primitives.

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 Chamberlinian example, it follows from the uniform random matching of consumers with firms. In Salop's (1979) model of competition on the circle doubling the number

line and assuming that the cost of entry at a given "location" increases steeply with the number of entrants there.

of active producers while maintaining their uniform spacing *and holding their prices fixed* cuts each producer's revenue in half. Accordingly, that and similar models of Hotelling-style competition also satisfy a version of Assumption S1 appropriately restated to remove the presumption that each firm has measure zero when calculating aggregates.

Assumption S1 is closely related to the independence of irrelevant alternatives: Adding a producer to a market while holding all incumbents' prices and product characteristics fixed does not change the *relative* market shares of any two incumbents. The quadratic demand system of Ottaviano et al. (2002) generally violates Assumption S1, because adding a good increases the cross-good elasticity of substitution and thereby moves sales from high-priced to low-priced incumbents. In this sense, this quadratic demand system hard wires markups' decline with market size.

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*

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

then

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

Assumption S2 asserts that it 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 actions (such as a quantity setting game with no productive spillovers) 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 Chamberlinian 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 $F^T(x) \equiv F\left(g^{-1}(x)\right)$, 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*

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

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 (1977) 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. 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 scarce.

3.2.3 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 and is a Borel measurable function of x given any realization of Z and any Borel measurable entry profile F.*

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) $F_0(x)$*is a Borel measurable functions of* x *for any* Z*,*

(b) $A_0(x; Z, F)$ *is a Borel measurable function of* x *given any realization of* Z *and any Borel measurable entry profile* F*, and*

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

A model that violates E1 can make no equilibrium prediction for some market size, and a model that violates E2 makes no useful prediction for any market size.

In this specific sense, a model that violates either of these assumptions is not well defined. Parts (a) and (b) of E2 ensure that required integrals can be calculated, and part (c) 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 E1 and E2 are clearly true for this section's examples. Accordingly, I view both assumptions as regularity conditions that a model must satisfy before it could be taken seriously as an explanation for the data.

3.2.4 Market Size and Producers' Actions

The Chamberlinian example and many other models of monopolistic competition satisfy all of the above assumptions. Together, they place sufficient structure on the model to imply the following observational implication.

Proposition *If $S = 2^j \times S_0$, where j is a non-negative integer and S_0, A_0, and F_0 are 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 $G(A, Z, F)$ is as defined in Assumption A1.

The proposition says that for every measurable equilibrium with a given market size there exists a corresponding equilibrium for a market with twice as many consumers with identical distributions of producers' actions. Unless there exists more than one free-entry equilibrium *and* the equilibrium selection rule systematically depends on S, there can be no observable relationship between market size and the distribution of any producer choice.

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]

The proposition illustrates that the invariance of producers' decisions to the market's size in the Chamberlinian example extends well beyond its particular assumptions. All of the assumptions excepting A1 and A2 are regularity conditions. These keep the primitives of tastes and technology constant across markets of different size, and so their empirical validity depends on the richness of the control variables included in the regression. I interpret the fact that market size *does* influence restauranteurs' observable decisions even after accounting for cross-market differences in factor prices and a variety of demographic variables as a rejection of atomistic competition.[10]

3.2.5 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 extending the general model to allow for discrete dimensions of firms' product-placement decisions is straightforward.

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

[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 $t > 0$, 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.

[10]Since the proposition places no direct requirements on producers' cost functions above those in Assumptions S2 and S3, the invariance of firms' decisions to market size holds good even if the available production technology features returns to scale or scope.

entry or incumbents raise entrants' fixed costs. Bresnahan and Reiss (1990) apply this approach to concentrated retail automobile markets in isolated towns. Berry and Waldfogel (2006) 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 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.[13] This is because classical measurement error biases an OLS estimate of the coefficient multiplying the infected regressor *towards zero*. Under the null hypothesis of atomistic competition, the true coefficient equals zero, so the usual bias from measurement error vanishes. Put differently, 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.

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. In this sense, industries that satisfy his model's assumptions are natural oligopolies. As noted above in Subsection 3.2.5, it is notationally burdensome but straightforward to add pre-entry investments in quality to this paper's model. This extension

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

[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 equals 0.93, and its standard error equals 0.02.

[13] Bresnahan and Reiss (1990) 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.

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 necessity of the highest quality firm's guaranteed minimum market share for his results. That is, endogenous sunk costs are necessary but not sufficient for an industry to be a natural oligopoly. The observation that an industry's producers incur endogenous sunk costs does not imply that its firms are oligopolists, but 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 firms. These papers have focused on differences in firm growth and survival across the life cycle (as in Dunne et al. (1988)) and on the interaction of resource reallocation with the business cycle (as in Davis et al. (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 in more sparsely populated areas, and he interprets this as the result of more intense selection in highly competitive markets. As noted in the introduction, Asplund and Nocke (2006) create a model of such selection by incorporating markups which depend on the number of producers into an otherwise standard model of industry dynamics with Chamberlinian monopolistic competition. They confirm Syverson's finding by showing that Swedish hairdressing establishments are younger in larger markets. This paper finds the opposite relationship between market size and exit for U.S. restaurants. Together, this paper's theoretical results and those of Asplund and Nocke imply that dynamic aspects of strategic interaction substantially influence the rate of restaurant turnover.

5 Conclusion

Researchers' prior beliefs about the usefulness of different modelling approaches influence their investigations of industries. The relative simplicity of atomistic competition models makes them a tempting first choice for the empirical study of competition in large markets. However, those who believe that strategic interaction permeates all producers' choices have chosen to focus instead on industries with few competitors and relatively well-defined strategic environments. This paper's results are of use to both sorts of researchers. For those who regularly abstract from strategic interaction, the nonparametric test of atomistic competition can be used to subject this assumption to empirical scrutiny before proceeding to a more involved investigation.[14] For those unwilling to part from a strategic focus, the regressions indicate the dimensions of the data along which strategic interaction manifests itself quantitatively. This is an essential first step to extending the strategic analysis of competition to large markets.

The application of this paper's theoretical analysis 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. A particularly important aspect of this is that exit rates fall with market size. In related work, Yeap (2009) documents that the 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 (2006) apply this papers' results to test the assumption of atomistic competition in their structural model of new Texas bars' growth and exit decisions.

Appendix 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

(i) $F(x') \geq 2$ for some $x' \in \mathcal{X}$ or

(ii) $\{x | F(x) \neq F_1(x)\}$ 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 satisfies either (i) or (ii) above. Suppose that

$F(x) \in \{0,1\}$ and $F(x) = F_1(x)$ almost everywhere. Then

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

The first equality holds because F_1 and $F_0(C_p) + F_0^T$ 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

$$\begin{aligned}
&S_1 \times \rho(a', x'; G(A_1, Z, F), Z, N_F) - c(a', x'; A_1, Z, F, S_1) \\
&= S_0 \times \rho\left(a';, x';, G(A_0, Z, F_0(C_p)), Z, N_{F_0(C_p)}\right) - c(a', x'; A_0, Z, F_0(C_p), S_0)
\end{aligned}$$

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

Alternatively, if $F_0^T(x) = 1$, then we can construct a parallel argument to show that $A_1(x', Z, F)$ maximizes the firm's profit. Thus $A_1(x, Z, F)$ 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 $\pi(x'; A_1, F_1(\{x'\}), S) \leq 0$. Next, note that if $F_1(x') = 0$, then we know from above that $G(A_1, Z, F_1(\{x'\}))(a) = G(A_0, Z, F_0(\{x'\}))(a)$, 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 $\pi(x'; A_1, F_1(\{x'\}), S) \leq 0$ 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 $N_{F_1} = 2 \times N_{F_0}$, so 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 $F_1(x) + I\left\{x = x''\right\} - I\left\{x = x'\right\} = F_1(x) + I\left\{x = x''\right\}$ almost everywhere. From this and the fact that we have already verified condition (b) of an equilibrium's definition, we conclude that

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

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

Appendix Supplementary Tables

Table A1: Summary Statistics for Dependent Variables

	Mean	Standard Deviation	Correlation with Log Population
$\mathbb{S}(\$5.00)^{(i)}$	0.66	0.14	0.10
Exit Rate (in percentage points)	10.70	3.16	-0.12
Restaurants' Average Revenue (in $\$1,000$)	528.89	105.36	0.39
Average Seating Capacity	98.97	22.83	0.05

Notes: (i) Here, $\mathbb{S}(\$5.00)$ refers to the fraction of establishments charging \$5.00 or more for a typical meal.

Audit trail: Tables/table1.m → Tables/table1.tex

Table A2: Coefficient Estimates for Additional Control Variables[i,ii]

	Vehicles per Household	Hotel Sales per Capita	Percentage Students	Percentage Elderly	Percentage Low-Income	Percentage High-Income	Population Growth Rate
$\mathbb{L}(\$5.00)^{(iii)}$	4.07	2.38	-15.47	-8.36	10.00	$31.81^{\star\star}$	12.71
	(6.02)	(7.90)	(12.70)	(8.89)	(10.99)	(12.69)	(7.90)
Exit Rate	-0.25	0.05	-0.13	-0.58	0.64	$-1.48^{\star\star}$	0.44
	(0.31)	(0.35)	(0.58)	(0.44)	(0.54)	(0.63)	(0.36)
Log of Restaurants' Average Revenue	-0.71	$7.07^{\star\star\star}$	3.01	-0.79	$5.67^{\star}$	-2.46	$4.99^{\star\star\star}$
	(1.76)	(1.92)	(2.82)	(2.58)	(2.99)	(3.05)	(1.73)
Log of Average Seats per Restaurant	0.92	-0.40	1.44	-2.13	5.06	2.13	3.32
	(2.23)	(2.39)	(3.61)	(3.54)	(4.36)	(4.67)	(2.11)

Notes: (i) The table's entries are estimated density-weighted average derivatives of the indicated variable with respect to the given control variable. In addition to the given control variable, each underlying regressions includes the right-hand side variables listed in Table 1. 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) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iii) Here, $\mathbb{L}(\$5.00)$ refers to the logistic transformation of the fraction of restaurants in an *MSA* with typical meal prices greater than or equal to \$5.00.

Audit trail: Tables/table4.m → Tables/table4b.tex

Table A3: OLS and IV estimates(i,ii)

	OLS	IV
$\mathbb{L}(\$5.00)^{(iii)}$	-9.67*	-18.80**
	(5.34)	(8.48)
Exit Rate	-0.81***	-0.60*
	(0.28)	(0.32)
Log of Restaurants' Average Revenue	3.67**	3.96**
	(1.67)	(1.99)
Log of Average Seats per Restaurant	1.31	2.56
	(1.89)	(2.54)

Notes: (i) This table's first column reports ordinary least squares estimates of linear regression model analogous to those reported in the first column of Table 2. The second column reports estimates of the same linear model using the median value of owner-occupied housing and the median rent of renter-occupied housing to instrument for possible measurement error in the average shopping-center rent. (ii) 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) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iii) Here, $\mathbb{L}(\$5.00)$ refers to the logistic transformation of the fraction of restaurants in an *MSA* with typical meal prices greater than or equal to $5.00.

Audit trail: Tables/table3.m → Tables/table3b.tex

Table A4: Estimates with Alternative Price Statistics and the Number of Firms[i,ii]

	$\mathbb{L}(\$7.00)^{(\text{iii})}$	$\mathbb{S}(\$5.00)^{(\text{iv})}$	$\ln N^{(\text{v})}$
Population	-2.55**	-2.55**	0.93***
	(1.24)	(1.24)	(0.02)
Commercial Rent	2.63**	2.63**	0.06**
	(1.18)	(1.18)	(0.02)
Retail Wage	2.53*	2.53*	-0.02
	(1.35)	(1.35)	(0.03)
Advertising Cost	-1.87*	-1.87*	-0.02
	(1.03)	(1.03)	(0.02)
Income	1.56	1.56	0.10***
	(1.42)	(1.42)	(0.03)
Percent Females in Labor Force	-3.15**	-3.15**	0.03
	(1.33)	(1.33)	(0.02)
Percent Black	4.13***	4.13***	-0.09***
	(1.02)	(1.02)	(0.02)
Percent College	4.75***	4.75***	0.01
	(1.32)	(1.32)	(0.02)

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. See the notes for Table 1 for a description of the regressors. 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) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iii) Here, $\mathbb{L}(\$7.00)$ denotes the logistic transformation of the fraction of restaurants charging less than \$7.00 for a typical meal. (iv) Here, $\mathbb{S}(\$5.00)$ denotes the fraction of restaurants charging less than \$5.00 for a typical meal. (v) Here, $\ln N$ denotes the logarithm of the number of establishments serving the market.

Audit trail: Tables/table2.m → Tables/table2b.tex

Table A5: Summary Statistics for Dependent Variables: Refreshment Places

	Mean	Standard Deviation	Correlation with Log Population
$\mathbb{S}(\$5.00)^{(i)}$	0.26	0.09	0.11
Exit Rate (in percentage points)	9.65	3.19	0.08
Restaurants' Average Revenue (in $\$1,000$)	566.94	87.09	0.10
Average Seating Capacity	75.26	15.42	-0.03

Notes: (i) Here, $\mathbb{S}(\$5.00)$ refers to the fraction of establishments charging \$5.00 or more for a typical meal.

Audit trail: Tables/table5.m → Tables/table5.tex

Table A6: Nonparametric Regression Estimates: Refreshment Places[i,ii]

	$\mathbb{L}(\$5.00)^{(iii)}$	Exit Rate	Revenue	Log Average Seats per Restaurant
Population	4.01	-0.51*	2.88**	2.40
	(4.36)	(0.29)	(1.45)	(1.69)
Commercial Rent	-0.50	1.07***	-2.47**	-2.70
	(4.95)	(0.33)	(1.22)	(2.20)
Retail Wage	21.43***	0.29	0.31	0.19
	(5.05)	(0.40)	(1.37)	(1.88)
Advertising Cost	5.07	-0.43	2.19*	0.96
	(4.41)	(0.31)	(1.26)	(1.61)
Income	-6.86	-0.08	-0.92	-3.06
	(5.67)	(0.35)	(1.86)	(1.93)
Percent Females in Labor Force	3.17	-0.09	1.84	3.68
	(5.57)	(0.35)	(1.35)	(2.51)
Percent Black	2.68	0.49*	7.44***	1.60
	(4.05)	(0.29)	(1.03)	(1.67)
Percent College	6.11	-0.01	0.87	0.40
	(5.05)	(0.30)	(1.38)	(2.12)

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. See the notes for Table 1 for a description of the regressors. 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) To increase the precision of the estimates' reports, *all entries have been multiplied by 100*. (iii) Here, $\mathbb{L}(\$5.00)$ denotes the logistic transformation of the fraction of restaurants charging less than \$5.00 for a typical meal.

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Table A7: Estimates with Alternative Market Size and Bandwidth Choices: Refreshment Places$^{(i,ii)}$

	Market Size Measures			Bandwidth Choices	
	Population	Population Density	Housing Units	$h = 1$	$h = 3$
$\mathbb{L}(\$5.00)^{(iii)}$	4.01	2.83	3.68	4.78	4.10
	(4.36)	(3.81)	(4.44)	(5.40)	(4.37)
Exit Rate	-0.51*	-0.74***	-0.48	-0.70*	-0.50*
	(0.29)	(0.28)	(0.29)	(0.36)	(0.28)
Log of Refreshment Places' Average Revenue	2.88**	3.08**	2.40	5.71***	3.21**
	(1.45)	(1.30)	(1.49)	(1.78)	(1.32)
Log of Average Seats per Refreshment Place	2.40	0.11	2.36	3.02	1.68
	(1.69)	(1.77)	(1.77)	(2.13)	(1.76)

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. The underlying regressions include the right-hand side variables listed in Table 1. 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) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iii) 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.

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Table A8: Estimates with Additional Control Variables: Refreshment Places(i,ii)

	Vehicles per	Hotel Sales			Percentage		Population
	Household	per Capita	Students	Elderly	Low-Income	High-Income	Growth Rate
$\mathbb{L}(\$5.00)^{(iii)}$	4.22	4.16	3.87	3.77	4.70	3.13	3.78
	(4.49)	(4.33)	(4.40)	(4.48)	(4.39)	(4.82)	(4.45)
Exit Rate	-0.56*	-0.59**	-0.51*	-0.47	-0.48	-0.59*	-0.48
	(0.29)	(0.29)	(0.30)	(0.29)	(0.29)	(0.34)	(0.30)
Log of Restaurants' Average Revenue	4.23***	2.91**	3.02**	2.84*	2.39	2.57	2.55*
	(1.41)	(1.48)	(1.46)	(1.49)	(1.46)	(1.59)	(1.47)
Log of Average Seats per Restaurant	2.49	2.54	2.44	2.71	2.18	2.90	2.65
	(1.75)	(1.76)	(1.74)	(1.75)	(1.68)	(2.02)	(1.73)

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. The underlying regressions include the right-hand side variables listed in Table 1. 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) To increase the precision of the estimates' reports, *all entries have been multiplied by* 100. (iii) 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.

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