Efficiency Gain from Ownership Deregulation: Estimates for the Radio Industry*

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Abstract

Reducing fixed cost duplication—a common justification for concentrated market structure—motivated the US government to relax the number of radio stations a firm could operate in any local market. After deregulation the number of firms per market decreased. The implied cost saving depends on the per market fixed costs incurred by each firm. Using data from 140 markets we estimate upper and lower bounds to fixed costs using (i) an empirical model of gross profit and (ii) the assumption that the observed post-deregulation market structure is a Nash equilibrium. The estimates suggest that the efficiency savings were significant.

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1 Introduction

EFFICIENCY GAINS are a standard justification for tolerating a more concentrated market structure. The US government cited them as a major reason for relaxing ownership restrictions that placed limits on the number of radio stations a firm could operate both nationally and in any local market. In the subsequent change to market structure there was a fall in the number of firms per market, while the number of stations remained approximately constant, resulting in a reduced duplication of per-market fixed costs. This paper estimates bounds to the size of these fixed costs using a model of equilibrium market structure and computes the implied cost-side efficiency gains from the deregulation.

The regulations in place up to 1992 prevented any firm from operating more than 20 stations nationally and more than two in any local market. These regulations were criticized by the Federal Communications Commission (FCC) who argued that they prevented firms from enjoying potential local scale efficiencies. The government accepted this argument and relaxed the regulations in two stages: initially in 1992 and again in the 1996 Telecommunications Act. This Act abolished the national limit and increased the local limit to between five and eight stations depending on market size. There followed a period of restructuring which by 2000 had resulted in more concentrated local markets.

The firms that expanded in this period opted to cluster their new stations locally rather than spread them evenly across markets. Given that there are several reasons why a radio firm might wish to *avoid* local clustering—e.g. to avoid cannibalizing its own stations' audiences—these factors must be outweighed by other motives that favour local clustering. The most active firm in the 1996-2000 phase was Clear Channel Communications Inc. It justified the practice of local clustering as follows:

> We believe in clustering our stations in markets to increase our individual market share thereby allowing us to offer our advertisers more advertising options that can reach many audiences. We believe owning multiple radio stations in a market allows us to provide our listeners with a more diverse programming selection and a more efficient means for our advertisers to reach those listeners. By clustering our stations, we are also able to operate our stations with more highly skilled local management teams and eliminate duplicative operating and overhead expenses. [Source: Form 10K report for the Securities and Exchange Commission (2000), p7.]

This statement suggests that by operating several local stations a firm can both expand local market share and avoid duplication of per-market fixed costs. These fixed costs might include buildings, local advertising sales teams, expertise in serving local listeners with news and weather updates, etc. FCC (1992) detailed these efficiency gains as follows:

> consolidation of facilities, managerial and clerical staffs, sales, book-keeping, promotion, production, news[,] and other aspects of station operation. [FCC (1992) para 37].

To measure these fixed costs we specify a model of market structure for the radio industry. In the model a firm's output is its number of listeners—which are supplied to advertisers at a price per unit—and there are no marginal costs (i.e. no costs per-listener) so that a firm's gross profit is given by its revenue. Net profit follows after deduction of fixed costs, which are of two types: a fixed per-station cost of adding successive stations and a fixed per-market cost reflecting the overheads noted above.

Using survey data on audience sizes in 140 local markets we estimate (i) a listener choice system in which the number of listeners for any firm is a function of the number of stations it owns and the number owned by its rivals, and (ii) an advertising demand function relating the price of advertising to the number of listeners in the market. Together these give in each market a revenue model for each firm in terms of own and rival station numbers (allowing an endogenously determined equilibrium advertising price).

The revenue model is used to estimate bounds on fixed costs as follows. We assume that in each market the number of stations operated by each firm is a Nash equilibrium in a static simultaneous-move game with complete information; i.e. we assume a positive profit difference between the observed action of each firm and any alternative action it could have chosen (holding rival actions fixed). The profit differences are constructed using the (known) revenue model and the model of fixed costs (known up to parameters). The alternative actions include (i) not entering the market and (ii) a *replication* option of splitting into two firms (holding constant the total number of stations); these respectively imply natural upper and lower bounds to per-market fixed costs. Using the moment inequalities method in Pakes, Porter Ho and Ishii, hereafter PPHI (2006), we identify the set of fixed cost parameters that ensure a positive sign for the moments of the profit differences.[1] This set implies upper and lower bounds to fixed costs.

We find that per-market fixed costs are important and imply that imposing the pre 1992 ownership limits would result in an increase to fixed costs with a lower bound of around l0% of industry revenue. This finding is of interest in the debate about radio ownership regulation and more generally given the use of efficiency gains as a standard justification for industry consolidation.

The paper builds on a large literature on market structure. As emphasized in Sutton (1998), there are typically many possible equilibrium market structures, particularly when multi-product (multi-station) firms are present. Most of the existing papers use likelihood function estimation and therefore require a unique market structure prediction. Early work exploited the availability of mappings from the nonunique set of equilibria to some unique characteristic of the equilibria, such as the number of entrants or products (e.g. Bresnahan and Reiss (1991), Berry (1992), Berry and Waldfogel (1999) and Davis (2005)) but a drawback of these methods is the strong symmetry restrictions they impose on gross profits. As Seim (2006) shows, it is possible to relax these restrictions if an incomplete-information game is specified but a drawback of this approach is that Bayesian-Nash equilibrium results in ex-post regret, a feature which is difficult to justify in a model of equilibrium market structure. In our paper the method of moments technique allows for asymmetry in gross profits and multiple Nash equilibria, and does not impose any parametric form for the distribution of unobservable fixed costs.

We generalize the model used in Berry and Waldfogel (1999) which estimates fixed costs

[1] A related paper using inequality restrictions from Nash equilibrium in the context of an entry model is Andrews et al (2004).

for the radio industry in the *pre*-deregulation period. Our paper uses the same data sources but for the post-deregulation period. The questions that we study in this paper could not be analyzed using their model as it does not allow for multi-product firms or per-market—as well as per-product—fixed costs. Our model allows both of these features. Other relaxations include asymmetric market shares and a nonparametric distribution of unobservable fixed costs.

We also contribute to the literature on the radio industry in the post-deregulation period. Berry & Waldfogel (2001) and Sweeting (2006) analyze the programming choices of firms with multiple stations, Tyler Mooney (2006) looks at the impact of these effects on listener welfare, and Sweeting (2007) estimates the cost of format switching.

Finally, the paper is related to the literature on cost efficiencies from local clustering. Recently, Holmes (2006) and Jia (2006) have examined Wal-Mart's expansion decision and find significant cost advantages to Wal-Mart from concentrating stores in the same local region. More generally the paper is related to the literature on geographic location by multi-outlet (or multi-plant) firms (see Tovianen and Waterson (2005) and Ellison and Glaeser (1997)), and to the literature on the welfare consequences of entry regulation (see Schaumans and Verboven (2006)).

The paper is organized as follows. In Section 2 we give background information on the radio market and present some summary statistics on local market structure after deregulation. In Sections 3 and 4 we present a model of multi-station entry. In Section 5 we discuss the econometric assumptions. Section 6 presents the data and empirical results. Section 7 concludes.

2 Deregulation in the Commercial Radio Industry

The changes to entry regulations for the commercial radio market evolved as summarized in Table 1.[2] Up to 1992 the number of stations per owner was limited to two per market and 20 nationally. This was relaxed in 1992-1996 to 40 nationally and three or four per market. Finally, in February 1996, the Telecommunications Act abolished all national restrictions and increased market limits to 5 to 8 stations.[3]

After 1996 there was a rapid change to market structure.[4] The national picture is summarized in Table 2 which shows that the number of stations did not change much but there was a substantial reduction in the number of firms. The effect on local markets is summarized in Table 3 (using data for the 140 markets we later use to estimate the entry model). In line with most research on the radio industry, we use the local markets defined by Arbitron, a market research firm.[5] These markets correspond closely to metropolitan

[2]This paper studies commercial radio, which sells radio advertising time. Non-commercial radio is funded from public subscriptions and has a share of the total listeners of only a few percent. We do allow listeners to choose noncommercial radio stations in the listener choice model but do not analyze fixed costs or entry choices for firms running these stations.

[3]For more detail on policy background see Ekelund et al (2000).

[4]See FCC (2001) for a detailed discussion.

[5]Since 2003 FCC have defined markets using these the Arbitron market definitions. Up to 2003, however, they used a more complicated signal-contour method which corresponded approximately to Arbitron metro

Table 1: Government Limits on Station Ownership

	Maximum number of stations		
	Pre-1992	1992-1996	Post-1996
Nationally:	20	40	No limit
Locally:			
Markets with 1-14 stations	2	3	5
Markets with 15-29 stations	2	4	6
Markets with 30-44 stations	2	4	7
Markets with 45+ stations	2	4	8

Table 2: National Ownership: 1996 and 2001 (Source FCC(2001))

Nation-wide Data			
March 1996		March 2001	
Total #stations	10257	Total #stations	10983
Total #owners	5133	Total #owners	3836

Table 3: Changes in Market Structure 1996-2000

Data from 140 markets used in structural model		
	1996	2000
Total Radio Market Share (Source: Arbitron)	0.157	0.154
Total Number of Stations (Source: Arbitron)	2819	2961
Mean Market Concentration by station numbers (Source: Duncans):		
C_1	0.20	0.25
C_2	0.33	0.44
C_3	0.41	0.57
Median size of largest firm (#stations)	4	6

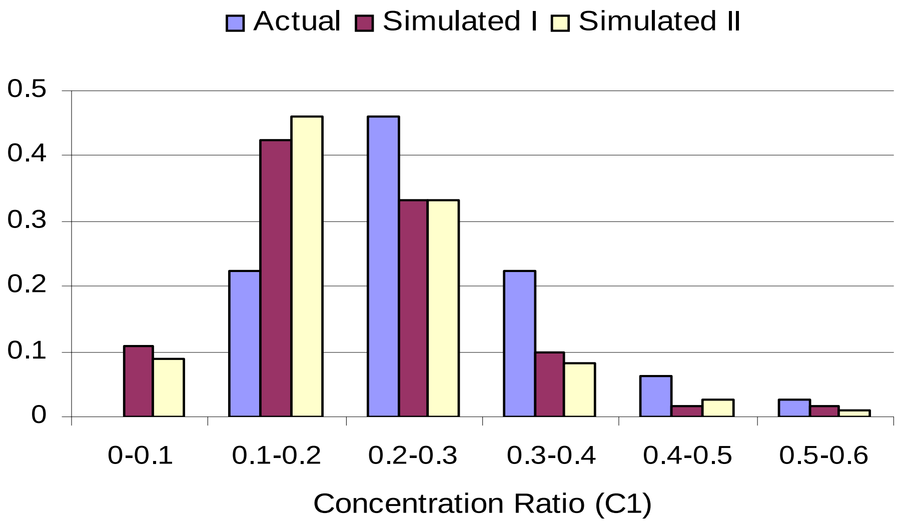

Figure 1: Histogram of actual and simulated C1 for year 2000 markets

Table 4: Greenstein Rysman Test for Local Clustering

Data from the 140 markets used in structural model		
	All stations in 2000	Stations Acquired 1996-2000
MT [H_o: $MT = 0$]	-19.88	-15.45
Standard deviation	0.47	0.46

areas and hereafter we refer to them simply as radio markets. Despite the change to market structure there was little change in either the share of the population listening to radio or the total number of stations broadcasting (see first two rows). The average concentration in these markets increased between 1996 and 2000 and maximum firm size increased. Concentration C_x is computed using station numbers, i.e. the number of stations run by the largest x firms as a share of the total number of stations broadcasting from within the market.

The concentration ratios observed in local markets are higher than would be expected if the stations in each market were allocated randomly to firms in proportion to firm size. To show this Figure 1 compares the distribution of C_1s (by station number) with those that are obtained simulating a random allocation. Specifically, suppose each of the n_m stations in market m is allocated to a firm by an independent random draw from a multinomial distribution where the probability p_j of a station being allocated to firm j is assumed equal to firm j's observed share of stations across all markets. We carry out two alternative simulations. Simulation I allocates stations operating in year 2000 using independent multinomial trials as described above while Simulation II conditions on stations that did not change ownership and only allocates stations that experienced a change of ownership between 1996 and 2000.

markets. We comment further on the differences between the signal-contour and the Arbitron market definitions in section 5.

As Figure 1 shows, the C_1 from either simulation is lower than the observed C_1, suggesting firms adopt a policy of local clustering.

A test statistic MT has been derived by Greenstein and Rysman (2004) which is normally distributed with mean zero under the null hypothesis that the stations are allocated to the firms independently as in the simulation model we described. A value $MT < 0$ indicates local clustering so that our results in Table 4 reject the null in favour of local clustering in two alternative specifications: one for all stations operating in the year 2000 (column 1) and the other for stations whose ownership changed since 1996 (column 2). We describe the construction of MT in Appendix 1.

Finally, we note that the evidence presented in FCC (2001, 2007) suggests that local market structure had largely stabilized by around 2000.[6]

3 Model Overview

The game is played separately in M independent markets ($m = 1, ..., M$); to avoid clutter in this section we drop the m subscript and consider a single market. A firm j derives revenue by supplying q_j listeners to advertisers at a market price p per listener. The firm can not set q_j directly but can influence it in a "lumpy" way through a choice of number of stations n_j. The number of people q_j listening to firm j's stations is given by the listener demand function:

$$q_j = q_j(n_j, \mathbf{n}_{-j})$$

where $\mathbf{n}_{-j}$ is the vector of station numbers for the other firms. The j subscript on $q_j()$ allows asymmetries between firms firms in their ability to attract listeners given $(n_j, \mathbf{n}_{-j})$. If j does not enter then $q_j(0, \mathbf{n}_{-j}) = 0$. Given any market structure $\mathbf{n} = (n_j, \mathbf{n}_{-j})$ the total number of listeners Q is

$$Q = \sum_{j \in \mathcal{J}^P} q_j(n_j, \mathbf{n}_{-j})$$

where $\mathcal{J}^{\ P}$ is the set of potential entrants. Advertisers pay a uniform price p per listener which declines along the advertisers' inverse demand curve for listeners $p(Q)$. Firm j has no marginal costs, i.e. costs that depend on the number of listeners q_j, but incurs the per-market fixed cost F_j to enter the market and a per-station fixed cost f_j, so profit for firm j is

$$\begin{aligned} \pi_j(n_j, \mathbf{n}_{-j}) &= p(Q) q_j(n_j, \mathbf{n}_{-j}) - f n_j - F_j \qquad (1) \\ &\equiv r_j(n_j, \mathbf{n}_{-j}) - f n_j - F_j \qquad (2) \end{aligned}$$

where $r_j()$ in the second line is a revenue function reduced to strategic variables.

[6] The charts in FCC (2001) and (2007) show rapid change up to 1998 and stability from about 200 onwards. Summarising this evidence FCC(2001, p6) states that: "This trend of fewer owners generally earning a larger percentage of market revenue [in local markets] ... has substantially tapered off over time. The large increase in concentration that occurred from March 1996 to November 1998 can be largely attributed to the relaxation of the local ownership rules required by the 1996 Telecommunications Act, as can the smaller increase that occurred from November 1998 to March 2000. The subsequent change from March 2000 to March 2001 is less pronounced.". More recently, FCC(2007) adds that "the [local] four-firm concentration ratio shows no substantial change between March 2002 and March 2007".

The firms act simultaneously with complete information. We consider only pure strategy Nash equilibria and allow nonuniqueness (a feature of simultaneous games with lumpy output). The market structure $\mathbf{n} = (n_j, \mathbf{n}_{-j})$ is a Nash equilibrium if and only if for each firm j the observed action weakly dominates any alternative action $a \in \mathcal{A}$, holding constant the actions of the other firms. We let $\Delta_a \pi_j(n_j, \mathbf{n}_{-j})$ denote the profit difference between the observed action and alternative a. The Nash equilibrium condition is then as follows:

$$\Delta_a \pi_j(n_j, \mathbf{n}_{-j}) \geq 0 \qquad \forall j \in \mathcal{J}^P \qquad \forall a \in \mathcal{A} \tag{3}$$

where $\mathcal{A}$ is the set of alternative actions. For the *actual* entrants $\mathcal{J} \subset \mathcal{J}^P$ the following four alternatives $\{R, L, U, D\}$ are included in $\mathcal{A}$:

1. *Ordered choice to right (a=R).* A unit increase to n_j is weakly dominated:

$$\Delta_R \pi_j(n_j, \mathbf{n}_{-j}) = r_j(n_j, \mathbf{n}_{-j}) - r_j(n_j + 1, \mathbf{n}_{-j}) + f_j \geq 0$$

2. *Ordered choice to left (a=L).* A unit decrease to n_j is weakly dominated:

$$\Delta_L \pi_j(n_j, \mathbf{n}_{-j}) = r_j(n_j, \mathbf{n}_{-j}) - r_j(n_j - 1, \mathbf{n}_{-j}) - f_j \geq 0$$

3. *Replicate upwards (a=U).* A replication of j into two independent firms (j and l) with identical cost and demand characteristics is weakly dominated:

$$\Delta_R \pi_j(n_j, \mathbf{n}_{-j}) = r_j(n_j, \mathbf{n}_{-j}) - [r_j(n_j - n_l, \mathbf{n}_{-j}) + r_l(n_l, \mathbf{n}_{-l})] + F_j \geq 0$$

   where n_l is divided between the two firms i.e. $n_l \in \{n_l : 0 < n_l < n_j\}$.

4. *Do not enter (a=D).* Not entering is weakly dominated:

$$\Delta_D \pi_j(n_j, \mathbf{n}_{-j}) = r_j(n_j, \mathbf{n}_{-j}) - f_j n_j - F_j \geq 0.$$

The ordered choice conditions respectively provide lower and upper bounds for per-station costs while the remaining two conditions respectively provide lower and upper bounds for per-market costs. Note that the above conditions relate only to actual entrants. We could have added for potential entrants the condition that no firm would have made a profit by entering; like the replication condition this gives a lower bound to per-market costs. However, for the purposes of estimating bounds to fixed costs we find it convenient to limit the conditions to the actual entrants (for reasons given in section 5); this does imply that the fixed costs we estimate are the costs of actual firms rather than all potential firms.

A simplifying assumption is that we do not explicitly model the firm's choice of individual station characteristics. Among stations the main form of product differentiation is format, e.g. Country, Classical, Talk, etc., for which discrete format codes exist. An extension to allow the firm a choice of format for each station would greatly complicate the model given the large number of distinct formats. A simpler alternative would be to allow the firm to choose a *number* of formats (as well as a number of stations) but in practice this is unnecessary because for any firm and market the number of formats is almost perfectly correlated with the number of stations (see Table 6 in Section 5) so that the inclusion of n_j in the listener demand function captures the effect of the number of formats. We find n_j to be a very good predictor of firm j's market share.

Finally in this section we note that in the larger markets in practice there is little room left on the spectrum for new stations.[7] In such cases the per-station costs f_j estimated using the ordered choice conditions include licence rents. Importantly, however, our measures of efficiency gain from deregulation will be based entirely on per-market fixed costs F_j which are not inflated by these rents.

4 Model Specification

In this section we fully specify the components of the profit function (1). We begin with listener demand $q_{jm}(n_{jm}, \mathbf{n}_{-jm})$ for firm j and market m. This is built up from a model of choice specified at the level of the individual consumer (i.e. potential listener) and individual radio station. In market m consumer $i = 1, ..., I_m$ has a choice set comprising all stations g ($g = 1, ..., \sum_{j \in \mathcal{J}_m} n_{jm}$) and the outside option of not listening to a radio station (denoted $g = 0$). The utility u_{ig} of consumer i in market m from listening to station g of firm j is:

$$u_{ig} = \alpha x_m + \xi_m + \beta x_{jm} + \xi_{jm} + \varepsilon_{ig} \equiv \delta_{jm} + \varepsilon_{ig} \tag{4}$$

where δ_{jm} is the mean utility for consumers in market m of listening to a station belonging to firm j. The remaining term ε_{ig} is an effect for consumer i and station g. The distribution of ε_{ig} determines the extent to which a new station is valued by consumers. We assume that this term is randomly distributed as for a two-level nested logit model, so the distribution of ε_{ig} depends on two parameters, namely (μ, λ); as we now show these parameters respectively determine the extent of business stealing and market expansion from a new station.[8]

At the lower level in the nested logit model the radio stations are grouped by *firm* $j = 1, .., J$. At this level parameter μ determines the extent to which ε_{ig} is positively correlated among the n_{jm} stations of a given firm. A high level of correlation (as $\mu \to 0$) implies a low expected benefit to any consumer from additional firm j stations.[9] This is clear from the expression for firm j's share s_{jm} of all radio listening:

$$s_{jm}\left(n_{jm}, \mathbf{n}_{-jm}\right) = \frac{\exp(\beta x_{jm} + \xi_{jm} + \mu \ln n_{jm})}{\sum_{k \in \mathcal{J}_m} \exp(\beta x_{km} + \xi_{km} + \mu \ln n_{km})} \qquad \text{for } \forall\ j \in \mathcal{J}_m \tag{5}$$

where the effect of n_{jm} on s_{jm} diminishes as $\mu \to 0$.[10]

[7] In a recent email correspondence James Duncan (the industry expert whose data is used in both studies) stated that station scarcity is likely to be a feature of the top 100 markets ranked by Arbitron (markets above New Haven in Appendix 3).

[8] For a discussion of the nested logit model see Berry (1994).

[9] We allow for correlation by owner to allow for the possibility that the content of stations is influenced by the owner; we do *not* expect that the consumer directly cares about (or knows about) the owner.

[10] Expression (5) is the standard form for the probability of choosing group j in a nested logit model when the n_{jm} options *within* any group j are symmetric up to ε_{gm} (i.e. have the same mean utility, δ_{jm}, for all g in j). Note that the model allows for asymmetries *between* firms. Also note that (5) can be specified more generally in n_{jm} if we drop explicit microfoundation at the individual station level; in section 6 we compare the fit of this expression with a specification where u_{jm} is quadratic in n_{jm}.

At the upper level in the nested logit model, choices are divided into two groups: the group comprising all radio stations and the group comprising (solely of) the outside option of not listening (which has utility $u_{i0} = 0 + \varepsilon_{i0}$). This grouping allows variation across consumers in taste for listening to radio—via positive correlation in the utility from radio options—determined by parameter λ. A high level of correlation (as $\lambda \to 0$) implies a new station would attract listeners mostly from existing radio stations. This can be seen from the following expression for the share S_m of m's population listening to radio:

$$S_m\left(\mathbf{n}_m\right) = \frac{\exp\left(\alpha x_m + \xi_m + \lambda V_m\right)}{1 + \exp\left(\alpha x_m + \xi_m + \lambda V_m\right)}, \tag{6}$$

which is the standard upper-level nested logit choice probability, where V_m is the log of the term from the denominator of (5):

$$V_m = \log\left\{\sum_{k\in\mathcal{J}_m} \exp(\beta x_{km} + \xi_{km} + \mu \ln n_{km})\right\}. \tag{7}$$

The number of listeners to firm j's stations is constructed as follows:

$$q_{jm}(n_{jm}, \mathbf{n}_{-jm}) = I_m \left[\frac{\exp\left(\alpha x_m + \xi_m + \lambda V_m\right)}{1 + \exp\left(\alpha x_m + \xi_m + \lambda V_m\right)}\right]\left[\frac{\exp(\beta x_{jm} + \xi_{jm} + \mu \ln n_{jm})}{\sum_{k\in\mathcal{J}_m} \exp(\beta x_{km} + \xi_{km} + \mu \ln n_{km})}\right]$$

for all $j \in \mathcal{J}_m$ where I_m is market population. An increase in n_{jm} increases both s_{jm} and (through V_m) S_m and the size of the two effects depends on parameters μ and λ respectively.

We now turn to the next component of the profit function (1): the advertisers' inverse demand function for listeners. We use the Berry and Waldfogel (1999) specification in which a uniform per-listener price of advertising p_m in market m is given by:

$$p_m(Q_m/I_m) \equiv p_m(S_m) = \exp\left(\gamma x_m + \omega_m\right)\left(S_m\right)^\eta \tag{8}$$

where x_m is observable market variables that affect the willingness to pay for advertising and ω_m is an unobservable market-specific effect. If $\eta < 0$ then the advertisers have a diminishing valuation of an advertising message reaching an individual listener as S_m increases. For a discussion of this specification see Berry and Waldfogel (1999).

Finally per-market and per-station fixed costs (F_{jm} and f_{jm} respectively) in (1) are the sum of a market effect and an unobserved effect, i.e.:

$$F_{jm} = \phi x_m + \tau^F_{jm} \tag{9}$$

$$f_{jm} = \sigma x_m + \tau^f_{jm} \tag{10}$$

where (ϕ, σ) are unknown parameters to be estimated, x_m is a vector of cost shifters (including a constant), and $(\tau^F_{jm}, \tau^f_{jm})$ are effects which are unobserved to the econometrician but known to the firms. We assume the unconditional means are zero, i.e.

$$E[\tau^F_{jm}] = E[\tau^f_{jm}] = 0 \qquad \text{for each } m \tag{11}$$

so that ϕx_m and σx_m are *by construction* the average fixed costs for the $\mathcal{J}_m$ (actual) firms in market m. We further assume there exist instruments h_m such that

$$E[\tau^F_{jm}|h_m] = E[\tau^f_{jm}|h_m] = 0 \qquad \text{for each } m. \tag{12}$$

5 Econometric Assumptions

The model is estimated in two stages: first we estimate the revenue model using demand-side data and standard IV techniques, second we estimate fixed cost parameters using the revenue model and the moment inequalities implied by Nash equilibrium.

At the lower level of the listener choice model we note from (5) that the log of share of listeners s_{jm} may be differenced as follows to give:

$$\ln s_{jm} - \ln s_{km} = \beta\left(x_{jm} - x_{km}\right) + \mu(\ln n_{jm} - \ln n_{km}) + \Delta\xi_{jm} \tag{13}$$

where k is a randomly selected reference firm in market m and $\Delta\xi_{jm} = \xi_{jm} - \xi_{km}$.[11] As n_{jm} and n_{km} are chosen by the firms they may be correlated with $\Delta\xi_{jm}$ so we assume $\Delta\xi_{jm}$ at the true parameters (β_0, μ_0) is mean independent of a set of instruments z_{jm} as follows:

$$E\left[\Delta\xi_{jm}\left(\beta_0, \mu_0\right)|z_{jm}\right] = 0 \tag{14}$$

where the variables included in x_{jm} and z_{jm} are detailed in the next section. At the upper level of the choice model we difference the expression (6) to give:

$$\ln S_m - \ln(1 - S_m) = \alpha x_m + \lambda V_m + \xi_m \tag{15}$$

where V_m is constructed using parameters estimated with (14). As V_m is endogenous we assume ξ_m at the true parameters (α_0, λ_0) is mean independent of instruments z_m[12]

$$E\left[\xi_m(\alpha_0, \lambda_0)|z_m\right] = 0. \tag{16}$$

The final part of the revenue model is the advertising inverse demand curve (8), which can be linearized by taking logs. Here S_m is endogenous and we assume that ω_m at the true parameters (γ_0, η_0) is mean independent of z_m, i.e.

$$E\left[\omega_m(\gamma_0, \eta_0)|\, z_m\right] = 0 \tag{17}$$

where z_m is exogenous data (we use the same z_m as in equation (16)).

The estimation of the fixed cost parameters requires that we compute the revenue each firm j would have earned had it chosen an alternative action a from the set $\{R, L, U, D\}$ defined in Section 3. The simultaneous play framework means that the actions of all rival firms $\mathbf{n}_{-jm}$ are assumed independent of the actions of firm j. Thus we may compute the equilibrium value of r_{jm} for any alternative firm j action by solving for a new (p_m, q_{jm}) holding $\mathbf{n}_{-jm}$ fixed, using the estimated listener choice and advertising demand model. In the following we treat $r_{jm}(n_{jm}, \mathbf{n}_{-jm})$ as being observable in this way for any $(n_{jm}, \mathbf{n}_{-jm})$.

[11] In many logit studies the *outside option* is used for k and its utility is set to zero (see Berry (1994)). In our nesting structure we exclude the outside option from the lower level so we instead use a single firm k as the reference firm for each $j \in \mathcal{J}_m$ and thus lose one observation per market.

[12] We correct standard errors to allow for the presence of of estimated parameters in variable V_m.

In our specification we identify fixed costs using the assumption that the observed market structure $\mathbf{n}_m = (n_{jm}, \mathbf{n}_{-jm})$ is a Nash equilibrium as defined using profit differences in (3), which we rewrite as follows:

$$\Delta_a \pi_{jm}(n_j, \mathbf{n}_{-jm}; (\phi_0, \sigma_0)) \geq 0 \ , \forall j \in \mathcal{J}_m, \ \forall a \in \mathcal{A} \tag{18}$$

for each m where (ϕ_0, σ_0) are the true fixed costs parameters. As unobserved profits are confined to an additively separable part of fixed costs (see (9),(10)) it follows that the profit difference $\Delta_a \pi_{jm}$ can be written as the sum of an observed part $\Delta_a \bar{\pi}_{jm}$ and a part $\Delta_a v_{jm}$ that is unobserved (by the econometrician):

$$\Delta_a \pi_{jm}(n_{jm}, \mathbf{n}_{-jm}; (\phi_0, \sigma_0)) = \Delta_a \bar{\pi}_{jm}(n_{jm}, \mathbf{n}_{-jm}; (\phi_0, \sigma_0)) + \Delta_a v_{jm}. \tag{19}$$

We introduce notation $d_{jm}(a)$ to denote the decision rule of firm j in market m when faced with alternative a, such that $d_{jm}(a) \neq a$ indicates that alternative a is not chosen (i.e. the firm weakly prefers the observed action to the alternative a). To estimate the model we require that the structural error $\Delta_a v_{jm}$ in (19) has nonpositive expectation after conditioning on $d_j(a) \neq a$ and nonnegative instruments h_m, i.e.:[13]

$$E\left[\sum\nolimits_{j \in \mathcal{J}_m} \Delta_a v_{jm} | \ h_m, \ d_{jm}(a) \neq a\right] \leq 0 \qquad \text{for } a \in \{R, L, U, D\}\,. \tag{20}$$

If this condition holds then the Nash condition (18) implies that the expectation of the *observed* profit difference $\Delta_a \bar{\pi}_{jm}(n_{jm}, \mathbf{n}_{-jm}; (\phi_0, \sigma_0))$ will be positive at the true (ϕ_0, σ_0) conditional on alternative a not being chosen, i.e.

$$E\left[\sum\nolimits_{j \in \mathcal{J}_m} \Delta_a \bar{\pi}_{jm}(n_{jm}, \mathbf{n}_{-jm}; (\phi_0, \sigma_0)) | \ h_m, d_{jm}(a) \neq a\right] \geq 0 \qquad \text{for } a \in \{R, L, U, D\} \tag{21}$$

and bounds to (ϕ_0, σ_0) can be identified using the empirical analogues of these inequalities. As the moment conditions are inequalities we are able to identify the parameters up to a set (see Manski (2003)). The variables h_m are helpful as they add additional moment inequality conditions that potentially narrow the identified set.

We can now construct the observed profit difference $\Delta_a \bar{\pi}_{jm}$ for the alternatives $\{R, L, U, D\}$ defined in Section 3:

$$\Delta_a \bar{\pi}_{jm} = \begin{cases} r_{jm}(n_{jm}, \mathbf{n}_{-jm}) - r_{jm}(n_{jm}+1, \mathbf{n}_{-jm}) + \sigma x_m & \text{if } a = R \\ r_{jm}(n_{jm}, \mathbf{n}_{-jm}) - r_{jm}(n_{jm}-1, \mathbf{n}_{-jm}) - \sigma x_m & \text{if } a = L \\ \\ r_{jm}(n_{jm}, \mathbf{n}_{-jm}) - [r_{jm}(n_{jm} - n_{lm}, \mathbf{n}_{-jm}) + r_{lm}(n_{lm}, \mathbf{n}_{-lm})] + \phi x_m & \text{if } a = U \\ r_{jm}(n_{jm}, \mathbf{n}_{-jm}) - (\sigma x_m)\, n_{jm} - \phi x_m & \text{if } a = D \end{cases} \tag{22}$$

where in the case of $a = U$—in which firm j replicates itself into two firms with identical cost and demand characteristics—we use $n_{lm} = \lceil n_{jm}/2 \rceil$, i.e. the smallest integer $\geq n_{jm}/2$, and assume the new firm l has the same fixed costs and mean utility δ_{lm} (in function (4)) as firm j.

[13] This is assumption 3 in PPHI (2006).

For each a we now discuss whether (20) holds, beginning with the two ordered choice alternatives $a = \{R, L\}$.[14] In the case of $a = R$ the unobserved profit difference relative to the alternative of offering one extra station is the unobserved per-station fixed cost saving, i.e. $\Delta_R v_{jm} = \tau^f_{jm}$. Since $d_{jm}(R) \neq R$ for all $j \in \mathcal{J}_m$, the expectation of τ^f_{jm} is unchanged by conditioning on $d_{jm}(R) \neq R$ so that

$$E\left[\sum\nolimits_{j\in\mathcal{J}_m} \tau^f_{jm}\Big| h_m, d_{jm}(R) \neq R\right] = E[\sum\nolimits_{j\in\mathcal{J}_m} \tau^f_{jm}\Big| h_m] = 0$$

for each m where the last equation follows from (12). Thus we satisfy condition (20). In the case of $a = L$ the unobserved profit difference relative to offering one fewer station is the extra unobserved per-station fixed cost i.e. $\Delta_L v_{jm} = -\tau^f_{jm}$. Since $d_{jm}(L) \neq L$ for all $j \in \mathcal{J}_m$, the expectation of $-\tau^f_{jm}$ is unchanged by conditioning on $d_{jm}(L) \neq L$ so that we again satisfy condition (20), i.e.

$$E\left[\sum\nolimits_{j\in\mathcal{J}_m} -\tau^f_{jm}\Big| h_m, d_{jm}(L) \neq L\right] = E[\sum\nolimits_{j\in\mathcal{J}_m} -\tau^f_{jm}\Big| h_m] = 0$$

for each m. For the case of $a = U$ the same argument applies as for $a = R$ except that now the unobserved profit difference (relative to the alternative of two firms) is equal to the unobserved per *market* fixed cost saving, i.e. $\Delta_U v_{jm} = \tau^F_{jm}$. Since $d_{jm}(U) \neq U$ for all $j \in \mathcal{J}_m$, the expectation of τ^F_{jm} is unchanged by conditioning on $d_{jm}(U) \neq U$ so that we again satisfy condition (20), i.e.

$$E\left[\sum\nolimits_{j\in\mathcal{J}_m} \tau^F_{jm}\Big| h_m, d_{jm}(U) \neq U\right] = E[\sum\nolimits_{j\in\mathcal{J}_m} \tau^F_{jm}\Big| h_m] = 0$$

for each m. The case of $a = D$ is slightly different. Relative to the alternative of not entering the unobserved profit difference for firm j is the unobserved per-market fixed costs plus total unobserved per-station fixed costs, i.e. $\Delta_D v_{jm} = -\tau^F_{jm} - n_{jm}\tau^f_{jm}$. Since $d_{jm}(D) \neq D$ for all $j \in \mathcal{J}_m$ it follows again that $E[-\tau^F_{jm}\big| h_m, d_{jm}(D) \neq D] = 0$ so that

$$E\left[\sum\nolimits_{j\in\mathcal{J}_m} -\tau^F_{jm} - n_{jm}\tau^f_{jm}\Big| h_m, d_{jm}(D) \neq D\right] = -E\left[\sum\nolimits_{j\in\mathcal{J}_m} n_{jm}\tau^f_{jm}\Big| h_m\right].$$

However because n_{jm} is endogenous and thus negatively related to τ^f_{jm} we expect a negative correlation between τ^f_{jm} and n_{jm} and it follows that $-E[\sum n_{jm}\tau^f_{jm}|h_m] > 0$, which implies we violate condition (20). To overcome this we substitute the observed lower bound to overall station costs (given by $\Delta_R r_{jm}$) for the observed station costs σx_m, i.e. in equation (22) we now write

$$\Delta_D \bar{\pi}_{jm} = r_{jm}(n_{jm}, \mathbf{n}_{-jm}) - (\Delta_R r_{jm})n_{jm} - \phi x_m,$$

which ensures that the unobserved profits in (19) are now always nonpositive, satisfying condition (20).

[14] The logic here matches that for the ordered choice problems in PPHI (2006).

We assume that each market m is an independent draw from a population satisfying our assumptions. Then the sample analogue of (21) is given by the matrix:

$$G(\sigma, \phi, x) = \frac{1}{M}\sum_{m=1}^{M}\left[\frac{w_m}{J_m}\sum_{j\in\mathcal{J}_m}\left[\begin{array}{c}\Delta_R r_{jm} + \sigma x_m \\ \Delta_L r_{jm} - \sigma x_m \\ \Delta_U r_{jm} + \phi x_m \\ \Delta_D r_{jm} - (\Delta_R r_{jm})n_{jm} - \phi x_m\end{array}\right]\otimes h_m\right] \tag{23}$$

where w_m is a weight. As $\otimes$ denotes Kronecker product (23) contains a total of $4 * H$ where H is the number of instruments in h_m. We discuss w_m and h_m in detail in Section 6. All parameters $\theta = (\phi, \sigma)$ that satisfy the system of sample inequalities in (23) are included in our estimate Θ of the identified set:

$$\Theta = \{\theta : G(\theta, x) \geq 0\} \tag{24}$$

where the system (24) is linear in parameters making it a standard linear programming problem. We present estimates of the maximum and minimum of each element of the parameter vector, e.g. the estimate of the lower bound to the first parameter given by

$$\hat{\theta}_1^{lb} = \arg\min_{\tilde{\theta}\in\Theta}\tilde{\theta}_1.$$

PPHI (2006) show the consistency of the estimates under the assumptions we have made and provide two alternative simulation techniques for constructing confidence intervals, one of which is conservative and the other provides shorter confidence intervals that have better coverage properties for samples that are large enough assuming certain regularity conditions; in Appendix 2 we provide Monte Carlo analysis (assuming these conditions hold) that shows the coverage properties are good .

The ownership regulations (described in Section 2) imply an upper legal limit to the number of stations that any firm j may own in any market. This creates what PPHI call a boundary problem which in our case implies that the assumed inequality relationship (18) may not hold for $a = R$. The option of simply dropping the affected right-bounded observations from the estimation creates a potential violation of the assumption in (20), because the unobserved profit difference from the right is expected to be higher for firms that are not at the right boundary, resulting in an estimated lower bound for per-station fixed costs that is too high. PPHI (2006, p56) point out it is possible to resolve this problem by substituting a random variable that we know results in the condition (20) being satisfied. In our case, we can substitute the average $\Delta_R r_{jm}$ for the firms in market m that are not at the boundary for the $\Delta_R r_{jm}$ of the firm that is at the boundary: this will result in the expectation of the unobserved portion of the profit change being nonpositive as required.[15]

To determine whether a firm in any market is in fact at a boundary we need to know the total number of stations the FCC counted—for the purposes of its ownership regulations—as

[15] In the ATM example used in PPHI (2006), there is a boundary is to the left for those firms that are observed to have no ATM machines so that the change in profits from the left are replaced with the average returns of the first ATM for the stations that do have ATMs. Our procedure is identical except that our bounadry is to the right.

operating in each market. Since 2004 the FCC has used Arbitron's definition of a market and counted the number of stations listed as "home" to the market (as defined by a further market research firm). Before 2004 the FCC used an alternative market definition based on signal-contours, which resulted in roughly similar but somewhat less tight limits than the Arbitron market definition; thus after the new definition was introduced a number of firms violated the new limits that did not violate the old limits (see FCC (2003b), DiCola(2006)). Our study is for the year 2000 so to determine whether a firm is at a legal boundary we should in principle use the signal-contours method. However this method is highly complicated as it is station-centric—i.e. must be computed separately for each station in each market—and more importantly requires detailed technical information to which we do not have access. Therefore we use the Arbitron based market definitions. This results in about 12% of firms being classified as being at (or in violation of) a boundary, which is likely to be an overestimate given that the new market definitions resulted in tighter limits than the old definitions. Fortunately for our estimation method, performing the boundary adjustment (detailed in the last paragraph) in cases when it is not needed (i.e. in cases where the firm was not actually at a boundary) does not create any violations of the assumptions needed to estimate bounds, although it does result in a reduced lower bound to per-station costs. Furthermore, neither the method used to determine market size nor the boundary correction method of the last paragraph has any effect on the bounds estimated for per-*market* fixed costs—the costs that we use to measure efficiency gain in Section 6—as these are not influenced by the two inequality conditions $\{R, L\}$ for ordered choice of n_{jm}.

Finally in this section we note two advantages of using the replication alternative for actual entrants (i.e. using $a = U$) instead of using conditions based on potential entrants to derive inequalities informative about the lower bound to per-market fixed costs: (i) we avoid arbitrary assumptions about the identity or number of potential entrants in each market or the characteristics $\left(x_{jm}, \xi_{jm}\right)$ needed to compute revenues for each potential entrant and (ii) we condition on the same set of firms (i.e. the actual entrants) as the other moment inequalities which means that we can use assumption (12) throughout; this in turn implies clearly that the estimated bounds are for the average per-market fixed costs of the actual entrants.

6 Data and Results

6.1 Data

To estimate the model we use data from two sources: Arbitron and Duncan. Arbitron's (Fall 2000) survey of listeners gives for each Arbitron market (i) demographic information from the US Census, (ii) listening data on all commercial radio stations, and (iii) station characteristics such as power and height of transmitters. To this Duncan (2001) adds station-by-station ownership, revenue and format information. For the year 2000 we have both sources of data for the 140 markets listed in Appendix 3.

To compute market shares (s_{jm}) we use Arbitron's average quarter-hour rating (AQH) audience measure. This gives the number of persons listening to each station for at least five minutes during a quarter hour, averaged over quarter hours throughout the week (Monday-

Table 5: Description of Data

Variable	Units	Mean	Std Dev
A: Firm-Market Data			
n_{jm}		3.254	1.831
s_{jm}		0.152	0.108
FM_{jm}		0.733	0.273
$\text{FM}{*}\text{height}_{jm}$	Km	0.289	0.193
$\text{FM}{*}\text{power}_{jm}$	Megawatt	0.050	0.037
#Obs: 598			
B: Market Data			
Share in-metro	%	11.4	2.4
#Stations		21.3	8.2
#Firms		5.3	1.7
Population	100,000s	9.7	13.6
Ad Price (p_m)	$100	5.5	1.6
Income	$10,000	5.0	1.4
College	%	46.0	8.1
#Markets:140			

Table 6: Station Numbers and Broadcasting Content

n_{jm}	1	2	3	4	5	6	7	8
# (j,m pairs observed)	107	147	110	98	62	45	25	8
Average # Formats	1	1.9	2.81	3.72	4.73	5.69	6.44	7.46

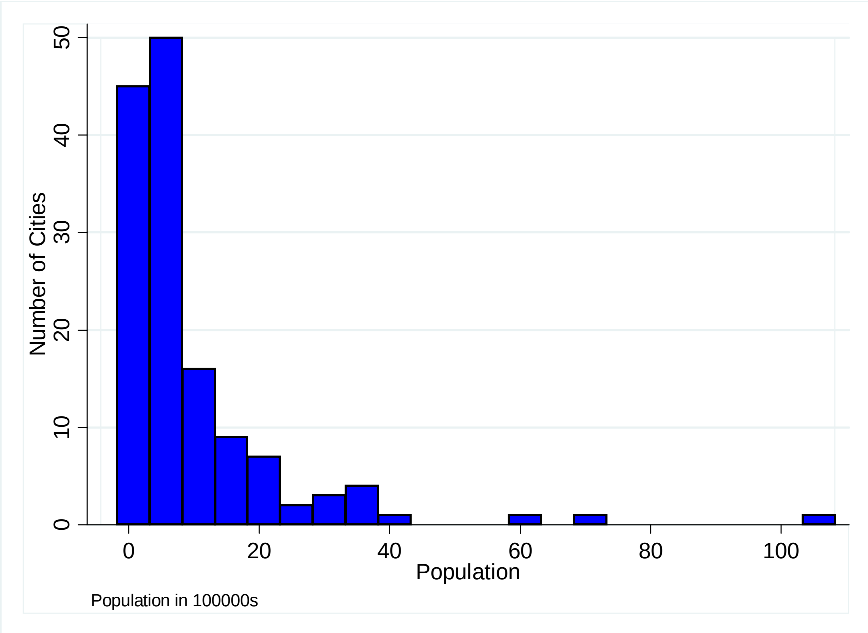

Figure 2: Histogram: City Population

Sunday 6am-midnight). We compute the advertising price p_m for each m by dividing the aggregate revenue of the firms in market m by the total number of listeners to those firms. These variables are as constructed (for 1992 data) in Berry and Waldfogel (1999), where a more detailed discussion of the data can be found.

Table 5 gives summary statistics. Panel A describes the firm-market level data used for the market share regression (13). There is a good level of variation for the key variables s_{jm} and n_{jm}. The other three variables (denoted x_{jm} in the model) are measures of firm j's average station signal quality: FM_{jm} is the proportion of firm j's stations in market m that broadcast in FM; $height_{jm}$ is the average height—and $power_{jm}$ the average power—of these (FM) stations. $height_{jm}$ and $power_{jm}$ are interacted with FM_{jm} as they are only measured for FM stations. Panel B describes market-level data for the 140 markets. The markets vary in terms of population, mean household income, proportion of the population with a college degree, and number of stations and firms. Figure 2 presents a histogram of market population showing that the bulk of the markets have population of up to 2 million. Finally, advertising revenue per listener (p_m) varies around a mean of \$550.

Table 6 shows the frequency of each n_{jm} between 1 and 8. The second row shows that the average number of formats increases almost one-for-one with n_{jm}. Thus by including n_{jm} in the market share regression we pick up the effect of the number of formats.

6.2 Results of Demand Estimation

Table 7 presents estimated parameters for alternative specifications of equation (13). Recall that equation (13) is based on the difference in log market share between firm j and a

Table 7: Utility Parameters—Lower Nest

	(i) OLS	(ii) OLS	(iii) OLS	(iv) OLS	(v) IV
Constant	0.005	0.006	0.021	0.006	0.001
	(0.037)	(0.037)	(0.034)	(0.035)	(0.036)
$n_{jm} - n_{km}$		0.966			
		(0.058)			
$n_{jm}^2 - n_{km}^2$		-0.072			
		(0.007)			
$\ln\left(n_{jm}/n_{km}\right)$ $[\mu]$	1.317		1.245	1.139	0.917
	(0.044)		(0.045)	(0.051)	(0.163)
$\text{FM}_{jm} - \text{FM}_{km}$			0.529	0.417	0.234
			(0.093)	(0.093)	(0.336)
$\text{Height}_{jm} - \text{Height}_{km}$			0.574	0.443	0.684
			(0.194)	(0.188)	(0.322)
$\text{Power}_{jm} - \text{Power}_{km}$			0.297	0.349	0.437
			(0.091)	(0.088)	(0.109)
Firm Dummies	No	No	No	Yes	Yes
R^2	0.660	0.655	0.715	0.764	
R^2 1st stage					0.494
Overidentification Test (P-value)					0.496

Standard Errors in (); #Obs: 458

reference firm k selected at random for each m. We include a constant term in the regressions which we expect to be insignificant since we expect zero mean utility difference between j and a randomly selected firm k. In specification (i) the only variable is (difference in) $\ln n_{jm}$ yet the R^2 is very high, which indicates a very strong relationship between station numbers and market share. In specification (ii) we try an alternative quadratic specification for the effect of n_{jm}. This brings no improvement to the fit, so in remaining regressions we use the $\ln n_{jm}$ specification, which has the advantage of being consistent with the two-level nested logit model of station choice discussed in Section 4.

The estimate of μ (the parameter on $\ln(n_{jm}/n_{km})$) may be biased upward in specification (i) because we have not allowed for the possibility that n_{jm} (which is chosen by firm j) is positively correlated with ξ_{jm}. To address this issue we begin by introducing variables to control for some of the unobserved utility. In specification (iii) we introduce (difference in) the three observable firm-market variables affecting the average quality of the signal as discussed in Section 6.1. These variables are significant, improve the fit of the model, and cause the parameter μ to fall, consistent with reduction of omitted variable bias. To further control for unobserved quality, specification (iv) adds firm dummies (for the largest firms) which causes μ to decline further.[16]

To deal with any remaining bias specification (v) uses IV estimation. We include in the exogenous data z_{jm} in condition (14) the firm-market variables x_{jm}, the firm dummies, and interactions of the firm dummies with market m population. Consistent with the expected direction of bias, μ declines further. The overidentification test rejects the hypothesis that the overidentifying assumptions are invalid. As the estimated μ in the final specification lies towards the upper end of the $[0, 1]$ range permitted in the logit model, the model implies a high level of business stealing (as opposed to self-cannibalization) when a firm adds a new station.

Parameters for the upper level of the nested logit are presented in the first two columns of Table 8 where Reg1-Reg3 are regional dummies ("West" is the base region with no dummy). To allow for correlation between V_m and ξ_m we run both OLS and IV models, though the parameters are not very different. In the the IV model the following data are used for z_m (in (16)): market income, population, population squared, proportion of population that is college-educated, and region dummies. The parameter λ on V_m lies towards the lower end of the $[0, 1]$ range permitted in the nested logit model, which implies that a new radio station attracts most of its audience from people who already listen to radio.

The parameters in the inverse advertising demand function are in the final two columns in Table 8. We obtain significant effects from the regional dummies and college education. To allow for endogeneity of the variable $\log S_m$ we run an IV specification in column (2), where the vector of exogenous data z_m is identical to that used in the upper level of the nested logit model. The overidentification tests in Table 8 reject the hypothesis that the overidentifying assumptions in the two IV regressions are invalid.

We now check the revenue function $r_{jm}(n_{jm}, \mathbf{n}_{-jm})$ our estimates imply, which is important given that $r_{jm}(n_{jm}, \mathbf{n}_{-jm})$ is used to estimate fixed costs. Table 9 presents revenue information using $r_{jm} \equiv q_{jm}p_m$. Panel A aggregates across firms and markets and presents

[16] The largest firms are defined as those present in 12 or more markets. This yields 23 firm dummies. These firms cover 83% of the market share of listeners.

Table 8: Parameters—Total Listeners and Advertising Demand

	Total Listeners ($\ln S_m - \ln(1 - S_m)$)		Advertising Demand $\ln p_m$	
	(i) OLS	(ii) TSLS	(i) OLS	(ii) TSLS
Constant	-1.808	-1.944	1.816	1.136
	(0.061)	(0.087)	(0.246)	(0.731)
V_m	0.037	0.080	–	–
	(0.010)	(0.021)		
$\log S_m$	–	–	-0.427	-0.163
			(0.071)	(0.277)
Income	0.785	0.537	0.035	1.464
	(0.459)	(0.502)	(1.479)	(2.116)
College	-0.204	-0.229	1.562	1.481
	(0.093)	(0.101)	(0.293)	(0.319)
Reg1 (Mid West)	0.007	0.028	0.232	0.244
	(0.021)	(0.024)	(0.065)	(0.069)
Reg2 (North East)	0.033	0.050	0.135	0.166
	(0.026)	(0.029)	(0.081)	(0.091)
Reg3 (South)	-0.019	-0.014	0.169	0.168
	(0.021)	(0.023)	(0.066)	(0.069)
R^2	0.175	–	0.382	–
R^2 1st stg		0.350		0.169
Overidentification Test (P-value)		0.740		0.242

Standard Errors in (). Income and College scaled up by 100; #Obs=140

Table 9: Station Revenue - Predictions of Model

A: Revenue Per Station ($m)			
All Markets (140 markets)		Markets under 2m (122 markets)	
Data: 4.186	Model: 4.138	Data: 2.275	Model: 2.256

B: Revenue Distribution By Firm Scale (All 140 Markets)								
n:	1	2	3	4	5	6	7	8
Data	0.060	0.126	0.180	0.181	0.122	0.137	0.088	0.105
Model	0.056	0.116	0.170	0.175	0.133	0.133	0.097	0.120
#firms	103	147	110	98	62	45	25	11

Table 10: Effect of Scale on Predicted Revenue

	Number of Stations (n):							
	1	2	3	4	5	6	7	8
All 140 Markets								
$r(n)$	4.817	8.669	12.044	15.072	17.828	20.359	22.701	24.879
$r(n)/n$	4.817	4.335	4.015	3.768	3.566	3.393	3.243	3.110
Markets under 2m								
$r(n)$	2.726	4.853	6.679	8.291	9.737	11.047	12.244	13.346
$r(n)/n$	2.726	2.426	2.226	2.073	1.947	1.841	1.749	1.668

Figures in $m; Figures are averages for each j in each m

mean per-station revenue. The model predicts an average revenue of $4.1m per-station. If we exclude the 18 markets with a population over 2 million the figure is $2.2m, which shows the important effect of population.

Panel A reports the corresponding revenue figures computed directly from the data. The near-perfect fit is an implication of the model and data: the price variable p_m in the model is constructed from observed market level revenue per listener data, and the empirical market share model has a perfect fit property that matches predicted and observed S_m. This is thus a check on the implementation of the model rather than on the realism of its assumptions.

The next check is however informative about the realism of the model. It concerns the *distribution* of revenue across firms of different size (in terms of number of stations (n_{jm})). We seek to check whether the the model is predicting this distribution well, without systematically over- or under-predicting the revenues from operating at any particular size. This is informative because we do not use individual firm revenue to estimate of the model—revenue data enters through a market-wide p_m—so there is nothing that matches observed and predicted r_{jm} for each j and m. Panel B of Table 9 compares the predicted and actual revenue distribution for each $n = 1, ..., 8$, and shows that the model matches the data well.

Finally, to get a feel for the shape of a firm's revenue function in n_{jm}, we compute predicted revenue for each firm j and market m at alternative numbers of stations n_{jm}, holding $\mathbf{n}_{-jm}$ constant. Table 10 shows simple averages across the firm-market observations. The total revenue function is concave with diminishing marginal (and average) revenue from each extra station.

6.3 Results of Fixed Cost Estimation

To facilitate a parsimonious parameterization of the fixed cost function we estimate the fixed cost model only on the markets with a population under 2 million. This eliminates only 18 of the 140 markets but allows us to concentrate on a much less heterogeneous group of cities—as Figure 2 shows. (This elimination in fact has very little effect on the results of the paper as shown in a robustness check later in this section). The total number of observations is 486 firm-market observations over 122 markets.

The fixed cost parameters are shown in Table 11.[17] Panel A reports three sets of estimates

[17] The parameters and CIs are not adjusted for the presence of estimated parameters in the revenue

using just the ordered choice inequalities $\{R, L\}$ which identifies bounds on just the per-station costs. In a very simple specification model (i) estimates a single parameter and sets $h_m \equiv 1$ in equation (23) so there are only two moment inequalities, determining the upper and lower bounds respectively. The bounds and the confidence intervals are informative and suggest average annual fixed per-station costs of about \$1.6m.[18]

Hereafter we specify h_m as a 3-vector comprising a constant term, an indicator function for whether market m is above median market population, and another for whether m is below median market population.

Model (ii) introduces a parameter on market population in per-station costs. We find that population is an important determinant of station costs, rising about \$0.2m for every 100,000 of population. This is likely to pick influences such as costs of property and labour and the cost of acquiring a licence in markets with licence scarcity.

Models (i) and (ii) deal with the boundary issue by discarding the observations that we determine are at a legal upper boundary. This results in discarding 12% of the right-bounded observations and has the potential to result in an estimate of the lower bound that is too high. Model (iii) corrects for the boundary problem as described in Section 5 which results in slightly wider bounds.

In panel B the Table reports parameters estimated using all four Nash conditions $\{R, L, U, D\}$ which allows bounds to the per-market costs to be estimated. It is in the per-market costs that we are principally interested. Model (iv) does not correct either for boundary issues or for the endogeneity of n in the viability constraint $a = \{D\}$. Model (v) does correct for the problem of endogenous n. As expected this correction slightly increases the upper bound to the estimates of per-market costs. Model (vi) corrects for the problem of endogenous n and also corrects for boundary problems in the bounds to the right $\{R\}$ in exactly the same manner as model (iii) and with the same effect. Note that in model (vi) the per-station cost estimates are identical to those in model (iii) and the per-market costs are identical to those in model (v); this is because—when the endogeneity of n has been corrected for—the system of moment inequalities can be solved as two independent linear programming problems. In particular this implies that the problem of estimating per-market costs has been separated from the estimation of per-station costs and issues that affect the latter have no effect on the former. To the extent that per-station costs are contaminated by boundary issues, issues from the FCC's definition of the market, and issues relating to licence scarcity, this is helpful as we know that any such contamination does not transfer to the per-market cost estimates, and it is the per-market estimates that we use in the next section to compute efficiency gain.

Note that all parameters are identified up to an interval, which shows that for each model there is a set of parameters such that none of the inequality constraints are violated. This itself is a useful check on the specification of the system.

In addition to presenting the upper and lower bounds to individual parameters we can present estimates of bounds to interesting functions of the parameters. Panel A of Table 12 presents lower and upper bounds to fixed costs for the median-population market, along with 95% inner confidence intervals (for brevity we present inner CIs hereafter). The estimated

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$r_{jm}(n_{jm}, \mathbf{n}_{-jm})$. The effect of this adjustment is expected to be small.

[18] The Inner and Conservative CI's are identical because the number of parameter bounds equals the number of inequalities (see PPHI (2006) for a discussion) .

Table 11: Parameters—Fixed Costs (486 observations in 122 Markets)

A: Ordered Choice Model using $a \in \{R, L\}$

Model		(i)*		(ii)		(iii)	
Boundaries:		$n_{jm} < \bar{n}$		$n_{jm} < \bar{n}$		$n_{jm} \leq \bar{n}$	
Per-Station Costs	Constant σ_1	[15.197	17.169]	[2.009	4.945]	[1.377	5.259]
	Inner CI	12.996	19.296	0.353	6.690	0.036	6.924
	Conservative CI	12.996	19.296	0.068	6.807	-0.304	7.029
	Population σ_2			[1.773	2.363]	[1.663	2.426]
	Inner CI			1.454	2.726	1.363	2.774
	Conservative CI			1.351	2.736	1.249	2.787

B: Full Model using $a \in \{R, L, U, D\}$

Model		(iv)		(v)		(vi)	
Boundaries		$n_{jm} < \bar{n}$		$n_{jm} < \bar{n}$		$n_{jm} \leq \bar{n}$	
Adjust $\Delta_V \bar{\pi}$?		NO		YES		YES	
Per-Station Costs	Constant σ_1	[2.009	4.945]	[2.009	4.945]	[1.377	5.259]
	Inner CI	0.513	6.237	0.442	6.699	-0.029	6.760
	Conservative CI	-5.325	9.141	0.304	6.780	-0.380	7.092
	Population σ_2	[1.773	2.363]	[1.773	2.363]	[1.663	2.426]
	Inner CI	1.470	2.720	1.450	2.694	1.330	2.739
	Conservative CI	0.842	3.356	1.356	2.698	1.243	2.792
Per-Market Costs	Constant ϕ_1	[-4.330	9.342]	[-5.396	13.632]	[-5.396	13.632]
	Inner CI	-7.789	12.248	-9.708	16.245	-9.120	16.124
	Conservative CI	-15.678	17.953	-9.708	17.088	-10.074	18.170
	Population ϕ_2	[0.703	3.940]	[0.278	4.311]	[0.278	4.311]
	Inner CI	0.234	4.764	-0.186	5.443	-0.201	5.400
	Conservative CI	-0.515	6.865	-0.732	5.395	-1.210	5.764

[.] are lower and upper bounds; Units: Population in 100k; Costs in \$100k . *For model (i), h=1.

Table 12: Costs implied by Fixed Cost Parameters

A: Comparison of Per-Market and Per-Station Costs						
	Per-market Cost		Per-Station Cost		Per-Market Cost/ Per-Station Cost	
	Lower	Upper	Lower	Upper	Lower	Upper
Bounds	[9.658	21.134]	[12.221	14.514]	[0.684	1.711]
Inner CI	8.675	23.071	11.504	15.232	$\langle at\ \theta_* \rangle$	$\langle at\ \theta^* \rangle$

B: Total Cost Per-Station at Alternative Scales

n	1	2	3	4	5	6	7	8
Minimal Economies of Scale (at θ_*)								
$\bar{F}/n$	24.041	19.160	17.533	16.719	16.231	15.906	15.673	15.499
Maximal economies of Scale (at θ^*)								
$\bar{F}/n$	33.302	22.794	19.291	17.540	16.489	15.789	15.288	14.913

Units: $100,000; assumes median market population (492k)

per-market costs are significant relative to per-station costs. To get a feel for the extent to which economies of scale are implied Panel A also reports bounds to the *ratio* of per-market to per-station costs and Panel B presents the shape of the average cost schedule at each of these bounds; we find that even at the lower bound per market costs are at least 68% of per-station costs and imply steep declines in the average cost schedule for the first few stations. However for both maximal and minimal economies of scale the schedule flattens out for n_{jm} greater than about four.

6.4 Measures of Efficiency Gain from Deregulation

The deregulation of the 1990s resulted in a substantial reduction in the number of firms per market but little change in the number of stations per market or in the number of listeners to radio (see Section 2). In this section we compute measures of the cost-efficiency gain from deregulation arising from the elimination of per-market fixed cost duplication. We do not conduct a full welfare analysis (given the important unpriced effects on listeners this would be beyond the scope of the paper); our focus is entirely on the effect of deregulation on per-market fixed costs.

Table 13 presents some measures of the efficiency loss from imposing the old ownership restrictions (in Table 1) on the year 2000 markets. In the first two columns we impose the pre 1992 restrictions (in which $n \leq 2$) while in the next two we impose the more relaxed 1996 restrictions (in which $n \leq 4$). We do not solve for a new Nash equilibrium. Instead, we simply keep the total number of stations constant and redistribute the surplus stations of the firms that violate the reimposed ownership limit. We consider two ways to distribute these stations. In the columns marked "Yes" we distribute the surplus stations only to new entrants in a way that minimizes the number of entrants needed (subject to the ownership limit). In the columns marked "No" we initially distribute surplus stations to incumbents unconstrained by the limit and only use new entrants once there are no unconstrained incumbents left. Clearly, "No" represents the minimum possible increase in per-market

Table 13: Effect of Imposing Tighter Ownership Limits

	Ownership Limit			
	n<=2		n<=4	
Redistribution only to new entrants:	Yes	No	Yes	No
Number of firms (initially 486)	857	824	586	518
Number of stations (initially 1575)	1575	1575	1575	1575
Number of stations/firm (initially 3.24)	1.838	1.911	2.687	3.045
Change in fixed cost (as % of revenue)				
Lower Bound	11.952	11.063	3.092	0.840
Upper Bound	28.653	25.113	6.851	1.819
95% Inner CI (lower)	11.363	10.130	2.790	0.748
95% Inner CI (upper)	31.128	27.629	7.510	1.982
Lower Bound (estimated on all 140 markets)	12.480	11.161	3.107	0.860
Upper Bound (estimated on all 140 markets)	28.167	25.075	6.872	1.823

fixed costs needed to accommodate the overall number of stations when the regulations are reimposed.

The above reallocation of stations implies a change in the number of firms ΔJ_m for each m. The aggregate change in per-market fixed costs is given by $\phi_1 \Delta J_m + \phi_2 \sum_m \Delta J_m pop_m$ where pop_m is market population. We compute the lower and upper bounds to this function given the set of inequality constraints in (23) as well as confidence intervals. The results are presented as a proportion of industry revenue in Table 13. We find that the 1992 regulations result in efficiency gains of at least 11% whether or not we reallocate some of the surplus stations to incumbents.

The 1996 regulations (analyzed in the remaining two columns) result in gains of at least 3% if stations are reallocated to new entrants only and 1% is stations are reallocated to incumbents first. The results suggest that the 1992 deregulation brought substantial efficiency gains while this is less clear for the 1996 deregulation. The bottom two rows of Table 13 present a robustness analysis to show the effect of using all 140 markets in (23), instead of just the 122 markets under 2m people. This makes very little difference to the results.

In Table 13 we minimized the number of new entrants needed to accommodate the redistributed surplus stations (subject to ownership limits); given that the market equilibrium may generate smaller firms this implies a conservative estimate for the increase in per-market costs (particularly for columns marked "No"). An alternative measure is constructed in Table 14 which performs the same exercise except that instead of imposing previous ownership limits we impose the average market concentration ratios that prevailed before the deregulation. Using 1996 data we compute the median size of the largest three firms for markets in each population quartile. That gives us firm sizes $\{n_1^q, n_2^q, n_3^q\}_{q=1}^{q=4}$ for each quartile q. We then impose these firm sizes as upper limits to the size of the largest three firms in each market. Specifically, we require that the largest firm in market m in quartile q does not exceed n_1^q, the second largest firm does not exceed n_2^q, and that the third and subsequent firms do not exceed n_3^q. Where violations of this requirement occur we redistribute surplus

Table 14: Effect of Imposing 1996 Concentration Ratios

Stations distributed only to new entrants:	Yes	No
Number of firms (initially 486)	687	654
Number of stations (initially 1575)	1575	1575
Number of stations/firm (initially 3.241)	2.293	2.408
Change in fixed cost (as % of revenue)		
Lower Bound	6.420	5.454
Upper Bound	15.838	12.159
95% Inner CI (lower)	6.221	4.938
95% Inner CI (upper)	17.096	13.403

stations either to entrants only, or to both incumbents and entrants, as we described for the previous table. We then compute upper and lower bounds to the change in per-market fixed costs. As Table 14 shows we obtain more firms than we got in Table 13 when we reimposed the 1996 regulations. The lower bound to the efficiency gain is about 6% of industry revenue.[19]

7 Conclusions

The structure of the profit function estimated in this paper shows that the radio industry is characterized by significant per-market fixed costs, resulting in economies of scale in station numbers. This provides an explanation for the locally clustered industry structure that has emerged since deregulation. Our structural model provides bounds to fixed costs that suggest ownership deregulation brought substantial efficiency gains via reduced duplication of per-market fixed costs; this is particularly true for the pre-1992 restrictions that limited firms to two stations per market. The presence of these efficiency gains was one of the main arguments in favour of the relaxation of ownership restrictions in place at the start of the 1990s. We focus entirely on the measurement of efficiency gains on the cost side from reduced fixed costs. A complete welfare analysis of the deregulation would include the welfare of listeners and advertisers, beyond the scope of this paper.

8 Appendix 1

The Multinomial Test of Agglomeration and Dispersion (MTAD) is based on the likelihood of the market structure $\mathbf{n}_m = (n_{1m}, ..., n_{Jm})$ under the null hypothesis that the stations are allocated to the firms independently with probability $\mathbf{p} = p_1, ..., p_{J_m}$ as in the model in section 2. The average log likelihood of the observed market structure—derived using the

[19] In fact our data on 1996 concentration, from Duncans (1997), is recorded the end of 1996, by which time some consolidation after the 1996 deregulation had already occurred. Thus the lower bound is conservative.

Table 15: Monte Carlo Coverage Ratio Analysis

	Parameter	$\%\{\hat{\theta}_{mc} \in 95\%\ CI^{*}\}$	
		Lower Bound	Upper Bound
Per-Station Costs	Constant	0.928	0.938
	Market Population	0.948	0.924
Per-Market Costs	Constant	0.940	0.957
	Market Population	0.961	0.954

5,000 replications (per point estimate). *Inner Confidence Interval for boundary point estimate

multinomial probability density function—is given by:

$$l\left(\mathbf{n},\mathbf{p}\right)=\frac{1}{M}\sum_{m=1}^{M}\ln\left[\binom{n_m}{n_{1m},...,n_{Jm}}p_1^{n_{1m}}...p_J^{n_{Jm}}\right].$$

The MTAD statistic MT is the difference between $l\left(\mathbf{n},\mathbf{p}\right)$ at the observed market structure $\mathbf{n}^{obs}$ and the value of $l\left(\mathbf{n},\mathbf{p}\right)$ expected under the assumption that $\mathbf{n}$ is generated by the model of independent random allocation, i.e.:

$$MT\left(\mathbf{n},\mathbf{p}\right)=l\left(\mathbf{n}^{obs},\mathbf{p}\right)-E\left[l\left(\mathbf{n},\mathbf{p}\right)\right].$$

To obtain $E\left[l\left(\mathbf{n},\mathbf{p}\right)\right]$ we compute the distribution of $l\left(\mathbf{n},\mathbf{p}\right)$ under independent random allocation by simulating $\mathbf{n}$ in 500 samples for each market. As shown in Greenstein and Rysman (2004) the statistic $l\left(\mathbf{n},\mathbf{p}\right)$ is normally distributed and a value of $MT<0$ indicates local clustering.

9 Appendix 2

PPHI (2006) show that the inner confidence intervals converge to the true limiting confidence intervals if the sample size is large enough assuming that there are exactly K binding population moments, where K is the number of parameters. One way of providing guidance on whether the sample size is large enough is through a Monte Carlo procedure that mimics the data generating process. We generate the moments in (23) from the observed data sample assuming (for each boundary point) the estimated boundary point $\hat{\theta}$ is the true θ using moments that are a random draw from a normal centered at the actual mean and covariance when evaluated at $\hat{\theta}$. We then calculate the distribution of the estimator using 5000 replications and compute the frequency with which it falls within the inner confidence interval. As we estimate bounds to four parameters there are eight $\hat{\theta}$ boundary points so the Monte Carlo procedure is done eight times. The results for fixed cost model (vi) are reported in Table 15 and show that the confidence intervals have good coverage ratios—around 95% of the estimated $\hat{\theta}_{mc}$ fall within the estimated 95% inner confidence interval.

10 Appendix 3

The 140 local markets used in the study were (ranked by market size):

Los Angeles, Chicago, San Francisco, Philadelphia, Detroit, Boston, Washington DC, Houston, Atlanta, Miami-Ft. Lauderdale, Seattle, Phoenix, San Diego, Minneapolis-St. Paul, Nassau-Suffolk, St. Louis, Baltimore, Tampa-St. Petersburg, Pittsburgh, Denver, Cleveland, Portland OR, Cincinnati, Sacramento, San Jose, Riverside-San Bernardino, Kansas City, Milwaukee, San Antonio, Columbus OH, Providence, Salt Lake City, Charlotte, Norfolk, Indianapolis, New Orleans, Greensboro-Winston Salem, Nashville, Memphis, Hartford, Raleigh-Durham, West Palm Beach, Rochester NY, Louisville, Oklahoma City, Dayton, Birmingham, Richmond, Greenville-Spartanburg, Albany-Schenectady-Troy, Tucson, Honolulu, Tulsa, McAllen-Brownsville, Grand Rapids, Fresno, Wilkes Barre-Scranton, Allentown-Bethlehem, Knoxville, Akron, Ft. Myers-Naples FL, El Paso, Albuquerque, Omaha, Monterey-Salinas-Santa Cruz, Syracuse, Harrisburg, Toledo, Springfield MA, Greenville-New Bern-Jacksonville, Baton Rouge, Little Rock, Charleston SC, Wichita, Gainesville-Ocala FL, Mobile, Des Moines, Spokane, Colorado Springs, Johnson City-Kingsport-Bristol, New Haven, Lafayette LA, Ft. Wayne, York, Lexington, Chattanooga, Roanoke, Worcester, Huntsville, Lancaster, Portsmouth-Dover-Rochester, Flint, Jackson, MS, Pensacola, Canton, Saginaw-Bay City-Midland, Reno, Fayetteville NC, Beaumont Port Arthur, Corpus Christi, Shreveport, Appleton-Oshkosh, Peoria, Montgomery, Springfield MO, Huntington WV, Macon, Rockford, Salisbury-Ocean City, Utica-Rome, Evansville, Savannah, Erie, Tallahassee, Portland, ME, Anchorage, Binghamton, Johnstown, Wilmington NC, Odessa-Midland, Lubbock, Asheville, Topeka, Green Bay, Manchester, Terre Haute, Waco, Springfield, IL, Sioux Falls, Fargo, Duluth, Charlottesville VA, Wheeling, Burlington VT, Panama City, Lafayette IN, Waterloo-Cedar Falls, Altoona, Billings, Bismarck.

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