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ABSTRACT

It is well documented that, in the presence of substantial fixed costs, markets offer preference majorities more variety than preference minorities. This fact alone, however, does not demonstrate the market outcome is in any way biased against preference minorities. In this paper, we clarify the sense in which the market outcome may in fact be biased against preference minorities, and we provide some conditions for such bias to occur. We then estimate the degree of bias in a particular industry using an empirical model of entry into radio broadcasting with two types of listeners, a preference majority and a minority, and the two types of stations targeting those respective listeners. Listening model estimates are used to infer fixed costs, which can then be used to find optimal station configurations as well as the welfare weights on different groups that rationalize the current configuration. The ensuing estimates reveal welfare weights that are 2-3 times higher for whites than blacks, and 1.5-2 times higher for non-Hispanic than Hispanic, listeners. The difference between the black and Hispanic results arises from the different patterns of importing and exporting: Hispanics listen to non-Hispanic-targeted stations more than blacks listen to white-targeted stations; and whites listen to black-targeted stations more than non-Hispanics listen to Spanish-language stations. Researchers and policy makers might add product markets to labor markets and other contexts that warrant attention for disparate treatment of minorities.

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1 Introduction

In differentiated product markets with high fixed costs, the product choice set is determined, in part, by the sizes of the populations of consumers with different tastes. In such situations, preference majorities may find relatively many products that cater to their tastes, while preference minorities find relatively few. In light of this well-documented fact, one might be tempted to conclude that the market favors preference majorities. On the other hand, if high fixed costs are needed in order to create a product that caters to a relatively small group of consumers, failing to supply this product may not be viewed as evidence of bias against this group. This paper seeks to clarify and measure the sense in which markets might be found to favor preference majorities. Our approach is to analyze the market outcome as if it were generated by a social planner. The question we address is whether the implicit social planner places more weight on the preferences of one group as opposed to another.

The notion of modeling market outcomes as though they were generated by a social planner is familiar. We teach undergraduates that the perfectly competitive outcome in a particular market matches the preferences of a social planner who cares equally about consumer and producer surplus. Less commonly, we could think of the classic monopoly quantity as the choice of a planner who cares only about producer surplus. In public economics, a particular system of redistributive taxes may be described via the preferences of a social planner who places differential weight on different groups. In regulatory economics, the complicated political process that generates regulated prices is often treated as the reflection of a single planner's relative preferences over different groups.[1] The task of this paper is to explicitly estimate the welfare weights that a differentiated product market's free entry equilibrium implicitly attaches to different groups of consumers.

A simplified example makes the exercise clear. Suppose that there are two distinct and different-sized groups with different preferences, a preference majority with 1,000 members, and a preference minority with 500 members. Preferences are distinct in the sense that each group consumes only the product targeted at that group. Assume that a first product targeted at each group attracts one quarter of the group to consumption, a second attracts an additional eighth, a third an additional sixteenth, and so on. Fixed costs of products targeting each group are identical at 100, marginal costs are zero, and the price of output for each group is constant at 1. Then, total revenue to products targeting the majority grows with entry: 250, 375, 437.5, 468.8, 484.4. With revenue split equally across products, average revenue per product declines: 250, 187.5, 145.8, 117.2, 96.9. Revenues, overall and per firm, are half as large for the firms targeting the minority. Given fixed costs of 100, the majority gets four products, while the minority gets one.

[1] See Ahmad and Stern (1984); Ross (1984); Armstrong, Cowan, and Vickers, (1994); Seim and Waldfogel (2013).

More products targeting one group does not indicate bias, however, because a planner who cares equally about groups would still choose to offer fewer products to the preference minority. Instead, possible bias manifests itself as inter-group differences in costs per incremental consumer. The marginal entering product serving the majority draws 31 individuals to consumption (469-438), while the marginal entrant serving the minority draws 125. Thus, the market entry process - our fictitious planner - incurs the same incremental fixed cost to serve an additional 31 members of the majority as it spends to serve an additional 125 members of the minority; and the planner rejects a reallocation that would serve 62.5 additional members of the preference minority at the cost of 31 majority listeners. Hence, the planner must attach a higher weight to majority than minority consumption.

We should emphasize that the foregoing example merely illustrates a possibility, albeit a plausible one. An appendix, below, demonstrates that the result is reasonably general when preferences are distinct in the sense that each group consumes only its own product (and when the taste for own-variety is the same across groups). A key point is that just as nothing about market entry in the presence of fixed costs guarantees an efficient number of products (Mankiw and Whintson, 1984), no force compels equalization of the planner's welfare weights for different groups of consumers in the presence of fixed costs and heterogeneous consumers. The appendix also shows simple conditions under which distinct preferences lead to implicit welfare weights that favor the preference majority.

The argument becomes more theoretically complicated when preferences are not perfectly distinct, so that (for example) members of the preference majority prefer their own product, but will also consume some of the product aimed at the preference minority. This is the more general, and most common, case. We refer below to the importing and exporting of group consumption across differently targeted products and show how it is reflected in implicit welfare weights. In general, empirical analysis is required to determine the sign, as well as the magnitude, of any bias toward the preference minority. Cross-group differences in prices and costs provide further nuance to the analysis (again, see the appendix), reinforcing the need for empirical analysis.

While effects of this type can operate whenever preferences differ across groups, their detection requires not only data on entry and consumption by group but also observations on multiple markets. Moreover, we require variation in relative group size across market areas. Such variation is not always readily available: while preferences for some products vary sharply across age or gender, there is often little variation across geographic markets in age and, especially, gender distributions.

We use the radio broadcast market - and listening patterns by race and Hispanic status - as our empirical context, for three reasons. First, fixed costs are important in this market. Indeed, all costs are fixed. Second, as Waldfogel (2003) documents, preferences differ sharply between groups of consumers. Blacks and whites - and Hispanics and non-Hispanics - listen largely to stations offering different programming formats. Third, terrestrial radio markets are local, and the share of population that is black or Hispanic varies substantially across US metropolitan areas. These features together give rise to strategies for measuring implicit planner weights on different types of consumers revealed by radio listening and market entry patterns.

Documenting the well-being that markets deliver to different groups of market participants has a long history in some areas of economics. A large volume of research in labor economics seeks to measure disparate outcomes in labor markets by race, gender, and other characteristics and, moreover, whether these disparate outcomes are motivated by animus or other causes. Other work documents differential patterns of residential location, criminal sentencing, health status, and other outcomes by group status.[2] This work is a first attempt to quantify potentially differential impacts on well-being operating through product markets. We should note that the mechanism posited here is not animus but rather simply the way that differentiated product markets function in the presence of high fixed costs and heterogeneous preferences.

Literature. This work is closely related to both the empirical work documenting preference externalities and the structural entry literature. A series of papers on preference externalities provides descriptive evidence about the operation of differentiated products markets where consumers have heterogeneous preferences, and fixed costs are substantial.[3] For example, Waldfogel (2003) documents that blacks and whites - and Hispanics - have sharply different preferences in radio programming. Because of fixed costs, groups face more products in markets with more consumers sharing their preferences. As a result, individuals derive more satisfaction in markets where they are more numerous. Blacks and Hispanics listen to radio more in markets with larger black and Hispanic population shares, respectively.

The empirical entry literature is vast.[4] Most relevant to our exercise is Berry and Waldfogel (1999), which uses data on the number of products available, along with detailed information on prices and quantities to estimate demand and, by extension, fixed costs. Using these estimates, they estimate the extent of excessive entry in US radio broadcasting in 1993. They find that free entry produces three times the number of stations that would be socially optimal for the purpose of producing advertising. They also point out that the observed entry pattern can be rationalized as socially optimal by a sufficiently high benefit of programming to users. Said another way, the model reveals the welfare weight that the hypothetical social planner implicitly attaches to the listeners at the marginal station. Formally, the current exercise extends this analysis to two types of products along with two types of listeners which, in turn, allows us to infer welfare weights

[2]See Klepper, Tierney, and Nagin (1983) for a discussion of evidence on criminal justice outcomes by race, Altonji and Blank (1999) and Neal and Johnson (1996) for evidence on pay differentials by race, and Bayer, Casey, Ferreira, and McMillan (2012) for evidence on racial patterns of prices paid in housing markets.

[3]See also George and Waldfogel (2003).

[4]Bresnahan and Reiss (1990) and Berry (1992) use data on the number of firms/products per market to draw inferences about post-entry competition. Mazzeo (2002) extends the framework to differentiated products, again relying only on information on the number of products to draw inferences about post-entry competition.

specific to each group. Our strategy for estimating fixed costs in the presence of heterogeneous stations follows Berry, Eizenberg and Waldfogel (2013) who study the optimality of product variety in radio markets from the point of view of a social planner who places equal weights on the welfare of stations and advertisers.

The paper proceeds in 4 sections. Section 2 presents a theoretical model linking the welfare weights that the planner attaches to various consumer groups to observable data. Section 3 describes the data used in the study. Section 4 presents estimates of a structural model of listening which, in turn, allow estimates of fixed costs. Section 5 employs the model for three tasks. First, we compare marginal entry conditions by race and Hispanic status, finding that marginal entry adds more minority listeners per dollar of fixed cost. This points to a bias against minority listeners. Second, we directly estimate the welfare weights that free entry implicitly attaches to black vs. white, and Hispanic vs. non-Hispanic, listeners. With listening demand estimates allowing for reasonable substitution patterns (i.e. nested logit), we find that the planner attaches 2-3 times higher weight to white vs black listeners and 1.5-2 higher weights for non-Hispanic relative to Hispanic listeners. These disparities in treatment are substantial compared with economic disparities by group attracting the attention of researchers and policy makers in other areas of economics. Third, we show that the difference between the black and Hispanic results arises from differences in the structure of preferences.

2 Model

The goal of this paper is to estimate a model of entry with two types of products and two types of consumers and to infer the value that - through free entry - the planner implicitly attaches to the two consumer types. To this end we estimate models of demand that the two types of consumers have for two groups of products. Given an equilibrium condition, these models allow us to estimate fixed costs for the two groups of products. Our estimates of fixed costs in conjunction with demand allow us to infer the planner's welfare weights.

2.1 Free Entry Equilibrium

There are two types of consumers, a preference minority (which we term blacks) and preference majority (whites), and there are two types of products, labeled B and W.[5] As the targeting implies, stations aim for their respective groups but also attract listeners from the other group. We can summarize a market configuration with the pair (N_W, N_B), which shows the number of W-targeted products and the number of B-targeted products in the local market. We assume that within each type, stations are symmetrically differentiated. Define $\ell_W(N_W, N_B)$ as the

[5] Two comments are immediately in order. First, whites include all non-blacks. Second, despite the notation, the model will be used for both blacks and whites as well as for Hispanics and non-Hispanics.

per-station listeners to a white-targeted station, and define $\ell_B(N_W, N_B)$ analogously. If a_W is the per-listener ad revenue at a white-targeted station, then $r_W(N_W, N_B) = a_W \ell_W(N_W, N_B)$ is per-station revenue at a white-targeted station, with $r_B(N_W, N_B)$ defined analogously. Finally, define FC_B and FC_W as the (exclusively) fixed costs of operating stations of the respective types.[6] Additional entry reduces per-station revenue, both within and across types.

There is free entry, so a station enters as long as the revenue it expects exceeds its cost. Equilibrium is achieved when there are no profitable opportunities for entry, and yet every entering firm is profitable. Thus the equilibrium conditions are

$$r_B(N_W, N_B) - FC_B > 0 > r_B(N_W, N_B + 1) - FC_B \quad (1)$$

$$r_W(N_W, N_B) - FC_W > 0 > r_W(N_W + 1, N_B) - FC_W \quad (2)$$

Intuitively, these conditions require that per-station profit be close to zero for both types in equilibrium. These conditions give rise to the observed configuration (N_W, N_B).

2.2 Free Entry as a Solution to the Planner's Problem

We can also view the problem from a planning perspective. Although there is no planner, we can use the planning framework in conjunction with free entry to infer the welfare weights that the hypothetical planner implicitly attaches to different groups.

First, we allow free entry to determine the number of stations operating and, given the costs of operating stations, the entire amount of resources to be spent operating stations (K). Then the planner chooses how many white and black targeted stations (N_W and N_B) to operate to maximize welfare, subject to the constraint that total station operation costs do not exceed K. That is, the planner maximizes $\mathcal{L} = W(L_B(N_B, N_W), L_W(N_B, N_W)) + \lambda[K - FC_B N_B - FC_W N_W]$, where $L_W = \ell_W N_W$ is the total number of listeners to white-targeted stations, L_B is defined analogously, and λ is a Lagrange multiplier. Ignoring the fact that stations are discrete, the first two first order conditions for the problem are:

$$\frac{\partial \mathcal{L}}{\partial N_B} = 0 \Rightarrow \frac{\partial W}{\partial L_B}\frac{\partial L_B}{\partial N_B} + \frac{\partial W}{\partial L_W}\frac{\partial L_W}{\partial N_B} = \lambda FC_B, \text{ and} \quad (3)$$

$$\frac{\partial \mathcal{L}}{\partial N_W} = 0 \Rightarrow \frac{\partial W}{\partial L_B}\frac{\partial L_B}{\partial N_W} + \frac{\partial W}{\partial L_W}\frac{\partial L_W}{\partial N_W} = \lambda FC_W \quad (4)$$

The first of these conditions indicates that the planner chooses a number of black-targeted stations such that the planner's welfare weight for black listeners $(\partial W/\partial L_B)$ times the change in

[6] Modeling radio stations as having no variable costs follows Berry and Waldfogel (1999) and is justified by the non-rival and non-excludable nature of the radio signal: there is no marginal cost associated with serving a marginal listener.

black listening with an additional black station ($\partial L_B/\partial N_B$), plus ($\partial W/\partial L_W$) times ($\partial L_W/\partial N_B$) equals the Lagrange multiplier times the fixed cost of operating a black station. The second condition is defined analogously. We will develop a way to observe everything but the welfare derivatives and the Lagrange multiplier below, which will allow us to infer the welfare weights up to a scalar. That is, we have a system of two equations with two unknowns (plus the Lagrange multiplier), which we can solve directly for the relative welfare weights attached to the two groups.

We can write these first order conditions in matrix notation as $Ab = f$, where the elements of A are the change in (white or black) listeners with a one-station increase in the number of white or black targeted stations, b contains the two welfare derivatives divided by the Lagrange multiplier, and F contains the fixed costs of operating the two kinds of stations. That is,

$$A = \begin{bmatrix} \frac{\partial L_B}{\partial N_B} & \frac{\partial L_W}{\partial N_B} \\ \frac{\partial L_B}{\partial N_W} & \frac{\partial L_W}{\partial N_W} \end{bmatrix}, b = \begin{bmatrix} \frac{\partial W}{\partial L_B}/\lambda \\ \frac{\partial W}{\partial L_W}/\lambda \end{bmatrix}, \text{ and } f = \begin{bmatrix} FC_B \\ FC_W \end{bmatrix}$$

We can compute the relative welfare weights b as $A^{-1}f$.

In addition to inferring welfare weights at the existing number of stations, we can solve for alternative values of N_W and N_B that correspond to different planning problems. First, we can maximize total listening subject to the constraint that costs not exceed their current level K. The resulting maximand is then:

$$\mathcal{L} = L_B(N_B, N_W) + L_W(N_B, N_W) + \lambda[K - FC_B N_B - FC_W N_W],$$

and the first order conditions for this problem are:

$$\frac{\partial L_B}{\partial N_W} + \frac{\partial L_W}{\partial N_W} = \lambda FC_W \text{ and } \frac{\partial L_B}{\partial N_B} + \frac{\partial L_W}{\partial N_B} = \lambda FC_B,$$

or, invoking the equal-marginal rule, that

$$\frac{\frac{\partial L_B}{\partial N_B} + \frac{\partial L_W}{\partial N_B}}{FC_B} = \frac{\frac{\partial L_B}{\partial N_W} + \frac{\partial L_W}{\partial N_W}}{FC_W}$$

This has the simple interpretation that if the system is maximizing listening the number of marginal listeners per dollar is equated across groups of stations.

Implementing this approach requires us to have estimates of $L_B(\cdot,\cdot)$ and $L_W(\cdot,\cdot)$, FC_W, and FC_B, to which we now turn.

3 Data

The data for this study include station-level listening by blacks and non-blacks (whom we term whites) for all of the commercial radio stations in 127 major US markets in 2005 (and 83 US markets for Hispanics and non-Hispanics). The listening data are drawn from Aribtron's Fall 2005 listening survey. Arbitron collects listening data on each station in the US. For markets with large black populations, they report station-level listening data separately for blacks and non-blacks. Similarly, for markets with substantial Hispanic populations, Arbitron separately reports listening data for Hispanics and non-Hispanics.

The basic listening measure is AQH listening, the share of population listening to a radio station during an average quarter hour. We treat this as a discrete choice. Because the vast majority of population (roughly 95 percent) listens to radio at some point, we treat population as the potential market.

We observe one important product characteristic, the format of the radio station. We have two sources of data on formats. BIA classifies each station into one of 20 broad formats (which they term format categories).[7] One of these categories is urban, which we treat, along with jazz and religious, as black-targeted. Another is Spanish which we treat as Hispanic-targeted. By tallying stations, we can calculate the number of black- and white-targeted stations - and the number of Hispanic- and non-Hispanic-targeted stations - in each market. We focus in this study on the stations broadcasting from inside the metro area. Thus, we treat listening to stations broadcasting from outside the market as part of the outside good.

We supplement the listening data with the following information at the market level: black and white population and the average ad price per listener, which we calculate as the market-level advertising revenue divided by the number of listeners (from BIA). We have data on black and white listening and ad prices for 127 markets. These markets average 1.2 million in population overall, with 0.189 million blacks. On average 9.77 percent of population listens to the radio for at least 5 minutes during an average quarter hour. Blacks listen more, averaging 10.94 percent while the white average is 9.37 percent. The annual ad price per AQH listener is $649.8. See Table 1a.

We have the analogous data for Hispanics and non-Hispanics for 83 markets, which average 1.5 million in total population, with 0.334 million Hispanics. See Table 1b. These markets have an average of 30.8 stations overall. Overall listening averages 9.69 percent in these markets. Hispanic listening is somewhat higher than non-Hispanic listening: 10.76 percent versus 9.18 percent.

One well-known fact about radio listening (see Waldfogel 2003) is that preferences differ sharply across groups. As Table 2 shows, the Urban format stations collectively attract 4 percent

[7] BIAKelseys Media Access Pro product (see http://www.biakelsey.com/Broadcast-Media/Media-Access-Pro/Radio/).

of white listening and 55 percent of black listening.[8] Two other formats, Religion and Jazz/New Age have audiences that are over a third black. We treat these three formats as black-targeted; and based on this designation, there are 7.7 black-targeted stations per market. As Table 3 shows, listening is similarly segregated by Hispanic status. Spanish-language stations attract 53 percent of Hispanic listening but less than a percent of non-Hispanic listening. We treat only Spanish-language stations as Hispanic-targeted, and the markets with Hispanic listening data average 4.5 Hispanic-targeted stations.

The relationship between the number of products available, by type, and the share of various population groups listening to radio is important for this study, and we document it descriptively as a step toward building a model. Figures 1a-c show how the share of population listening, overall and by group, varies with the number of stations N, overall, targeting whites, and targeting blacks. Figures 2a-c repeat the exercise by Hispanic status. In all figures we see a positive relationship: markets with more stations attract a higher share of population to listening. The apparent slope is steeper for black - and Hispanic - listening, meaning that an additional minority-targeted station raises the share of the minority group listening more than an additional majority-targeted station raises the share of the majority group listening.

4 Empirical Implementation and Results

To estimate welfare weights we need a model of how black and white listening to black and white targeted stations vary with the number of black and white stations (the configuration), as well as the fixed costs of operating each type of station. To estimate these quantities, we need to develop a listening model that allows us to model how listening to each type of station varies with the configuration of available stations.

4.1 Listening Model

The utility that consumer i derives from station j is:

$$u_{ij} = x_j \beta_i + \xi_j + \epsilon_{ij}$$

where x includes observable characteristics of station, β_i is potentially a person-specific coefficient, ξ_j is unobserved quality of the product, and ϵ_{ij} is extreme value and iid across consumers and products. When $\beta_i = \beta$, then this results in the logit model. In our context, products have one observable characteristic, format. For the purpose of entry modeling we will treat stations as symmetric within format. Individual i is a member of one of two groups (B and W).

[8] We verify below that our results are robust to a narrower measure of black stations that includes only the urban format.

Here, with the logit model, if there are N_B stations in the black format and N_W stations in the white format, the share of group g listening to a particular station in format f may be written:

$$s_f^g = \frac{e^{\delta_f^g}}{1 + N_B \cdot e^{\delta_B^g} + N_W \cdot e^{\delta_W^g}}$$

and the share of group g's population listening to stations in format f, denoted by a capital S, is given by

$$S_f^g(N_B, N_W) = \frac{N_f \cdot e^{\delta_f^g}}{1 + N_B \cdot e^{\delta_B^g} + N_W \cdot e^{\delta_W^g}}$$

where $f = B$ or W and $g = B$ or W.

Following Berry (1994), we can estimate the parameters of the logit model by regressing a station's mean utility for listeners in a group on a black station type dummy, separately for black and white listening.

$$ln(s_j) - ln(s_0) = x_j\beta + \xi_j$$

where s_j is the share of consumers choosing product j, s_0 is the share choosing the outside good, x contains a constant and a black-targeted station dummy, and ξ is the characteristics of product j not observed by the econometrician. This specification allows groups to have different preferences for differently-targeted stations, and the results will confirm this. While the logit model is easy to estimate and implement, it incorporates potentially unrealistic substitution patterns, which we will give the data the opportunity to reject.

The nested logit model is a generalization of logit that allows pre-determined groups of products to have different levels of substitutability within the nest vs outside the nest. Logit allows each product a different mean utility - and therefore share - but products are otherwise equally substitutable. The simplest nested structure that makes sense in this context follows Berry and Waldfogel (1999). All radio stations are in the nest. The outside good (non-listening) is separate. This model has the following expression for the share of group g's population listening to each station in format f:

$$s_f^g(N_B, N_W) = \frac{e^{\delta_f^g/(1-\sigma_g)}}{N_B e^{\delta_B^g/(1-\sigma_g)} + N_W e^{\delta_W^g/(1-\sigma_g)})} \cdot \frac{\left(N_B e^{\delta_B^g/(1-\sigma_g)} + N_W e^{\delta_W^g/(1-\sigma_g)}\right)^{(1-\sigma_g)}}{1 + \left(N_B e^{\delta_B^g/(1-\sigma_g)} + N_W e^{\delta_W^g/(1-\sigma_g)}\right)^{(1-\sigma_g)}}$$

Simplifying, the share of group g individuals in the market listening to stations in format f overall is:

$$S_f^g(N_B, N_W) = \frac{e^{\delta_f^g/(1-\sigma_g)}}{\left(N_B e^{\delta_B^g/(1-\sigma_g)} + N_W e^{\delta_W^g/(1-\sigma_g)}\right)^{\sigma_g} \left[1 + \left(N_B e^{\delta_B^g/(1-\sigma_g)} + N_W e^{\delta_W^g/(1-\sigma_g)}\right)^{(1-\sigma_g)}\right]}$$

The parameter σ_g describes the substitutability among products in the radio station nest. If $\sigma_g = 0$, then this model reverts to the plain logit. in this case, each station adds substantial variety and attracts additional consumers to the market. Higher values of σ_g reflect greater substitutability among stations. If $\sigma_g = 1$, then additional stations add nothing to the value of offerings to consumers.

Again following Berry (1994) and Berry and Waldfogel (1999), we can estimate this model with the regression:

$$ln(s_j^g) - ln(s_0^g) = \delta_j^g + \sigma_g ln\left(s_j^g/(1 - s_0^g)\right) + \xi_j.$$

That is, we regress the log share of the product less the log share of the outside good on a dummy for format as well as the product's share of inside products. The latter term is endogenous and requires an instrument. Because the term is essentially the reciprocal of the number of products, measures of market size (black and white population) provide natural instruments. We estimate the equation separately by group. Thus, as with logit, we allow blacks and whites to have different preferences for black and white targeted stations. We also allow blacks and whites to view stations as differently substitutable for one another.

Table 4 reports results. The first two columns report simple logit estimates, showing that blacks prefer black stations, and whites prefer white stations. Columns (3)-(6) report nested-logit estimates following the single-level nesting specification described above. The σ_g parameters are significant - and closer to 1 than 0 - rejecting the logit model and indicating strong substitutability of stations. The σ_g parameter is higher for whites than blacks, indicating that whites view radio stations as closer substitutes for one another than do blacks. Table 5 presents first-stage regressions, in columns (3) and (6). The endogenous variable, which is the station's share of group listening, is lower as the group is larger, as expected. Markets with more people have more stations and lower listening shares for each station. Own-race relationships are especially strong. Columns (5)-(6) in Table 4 report IV estimates from the one-level nest model. Apparent substitutability is higher in OLS than in IV estimates.

We also explore a model with nests at two levels. First, consumers choose whether to listen to radio. If they listen to radio, they make a second choice of whether to listen to black or white targeted radio stations. This is analogous to the model Verboven (1996) employs to analyze automobile choices in European markets. This model has two substitution parameters, one that reflects the substitutability of stations within each group (σ_1) and a second that reflects the substitutability of the nests for one another (σ_2). As before, we can estimate these parameters and the differential mean utility by station format separately for each group, via the following estimating equation:

$$ln(s_j) - ln(s_0) = x_j\beta + \sigma_1 ln\big(s_j/s_f\big) + \sigma_2 ln\big(s_f/(1-s_0)\big) + \xi_j$$

This regression has two endogenous terms. The first term is the station's listening as a share of listening to the stations in the format f (black or white). The second share on the right-hand side is the format's share of total listening (e.g. all black station listening as a share of total listening). The two shares on the right hand side are endogenous, and we instrument them using measures of the size and composition of the population (white and black population). The model with one nest is a special case of the model with 2 levels, with $\sigma_1 = \sigma_2$. Columns (7)-(10) of Table 4 report estimates from these two-level nested logit models. Using OLS, the two substitution terms are quite similar for white listeners, while the second term is lower for blacks (indicating that stations are better substitutes within than across formats). In the IV estimates, the substitutability across formats falls substantially for whites.

Tables 6 and 7 re-estimate these models using Hispanics and non-Hispanics rather than blacks and whites. Many patterns are similar. We reject the simple logit. Instrumental variables estimates of the one-level model indicate less substitutability of stations than OLS estimates. The two-level model yields IV estimates for some correlation coefficients (σ) that exceed the value 1, implying a rejection of the model as these coefficients are restricted to lie in the $[0,1)$ interval. We therefore rely on the 1-level IV model in the remainder of our analysis.

4.2 Station Fixed Costs

We assume that there is free entry, and we assume that the configuration we observe is a Nash equilibrium. Recall that $r_B(N_B, N_W) = a_B\ell_B(N_B, N_W)$ is the revenue per black-targeted station, with an analogous term defined for white-targeted stations. Then, following Berry, Eizenberg and Waldfgoel (2013), we can infer bounds on fixed costs from the equilibrium conditions

$$r_B(N_B+1, N_W) < FC_B < r_B(N_B, N_W) \text{ and } r_W(N_B, N_W+1) < FC_W < r_W(N_B, N_W).$$

We can estimate revenue with our listening model in conjunction with our estimates of market size (population of the respective black and white groups Pop_B and Pop_W), along with our data on the per-listener annual advertising revenue.

Implementing this requires estimates of the ad prices at minority and majority-targeted stations (a_W and a_B). Unfortunately, we do not have access to ad prices at the station level (only at the market level). There is some reason to think that ad revenue per listener differs between minority and majority-targeted stations. For example, in a study commissioned for the FCC, Ofori (1999) presents evidence that minority-targeted stations generated less revenue per listener than white-targeted stations. Napoli (2002) shows that power ratios (estimates of station revenue per listener) were about 25 percent lower for minority-targeted stations.

Given this finding, one approach would be to constrain the ad price to be 25 percent lower at minority-targeted stations relative to majority-targeted stations. Yet, a problem with this approach is that it does not allow the price per listener to vary with the group mix listening at the stations. This is a problem because minorities listen to majority-targeted radio stations when minority-targeted options are absent. If we assume that majority-targeted stations get higher prices per listener regardless of listener composition, we get the perverse effect that a first minority-targeted station reduces market ad revenue as minorities shift from the majority-targeted station where by assumption all listeners receive a high price per listener to minority-targeted station where listeners receive low prices.

A natural solution to this problem is to price listeners, rather than stations, by type. We can infer prices per listener type from Napoli's estimate that average prices per listener are 25 percent lower at minority-targeted stations. Define b_B as the black share of listeners to black-targeted stations, and define b_W as the black share of listeners at white-targeted stations. Then if the prices of black and white listeners are p_B and p_W, respectively, then the price per listener at a black-targeted station is $p_W(1 - b_B) + p_B b_B$. The price per listener at a white-targeted stations is analogously defined as $p_W(1 - b_W) + p_B b_W$. Finally, define $p_B = \phi p_W$. Then Napoli's estimate implies that $(1 - b_B) + \phi b_B = .75[(1 - b_W) + \phi b_W]$. The b terms are available as data (see Tables 2 and 3), so we can solve for ϕ directly. We find that $\phi = 0.635$ for blacks relative to whites and $\phi = 0.708$ for Hispanics relative to Hispanics. We observe an average ad price per listener in each market (a), which we can express as $a = p_W(1 - b) + \phi p_W b$. Thus, $p_W = a/[1 - b + b\phi]$, and $p_B = \phi p_W$. We estimate our models in two ways. First, we estimate the models with equal prices per listener across groups ($\phi = 1$). Second, we estimate the models with $\phi = 0.65$.

Our revenue model, along with some of the conditions for Nash equilibrium, gives us bounds on fixed costs. For markets with at least one station of each type, we observe both lower and upper bounds on fixed costs. There is one complication: a few markets (two in the black sample and eleven in the Hispanic sample) lack a station in the minority-targeted format. Hence, for those markets we have only a lower-bound for fixed costs.

One approach to inferring fixed costs is to make a distributional assumption, then to estimate via maximum likelihood. This approach works when the model has a unique equilibrium station configuration so that any particular fixed cost pair gives rise to a particular observed number of stations. In our context we cannot rule out multiple equilibria, so we adopt a different approach to estimating fixed costs.

Following Berry, Eizenberg, and Waldfogel (2013), we estimate fixed costs from the bounds directly. In our context the upper and lower bounds on per-station fixed costs turn out to be close, on average within 5 percent. Even so, we have to decide, within the bounds, what values of fixed costs to use. We make a set of decisions that are designed to work against a finding of bias again blacks and Hispanics. First, we estimate minority fixed costs based on the upper bound of costs, while we estimate non-minority fixed costs based on the lower bound. Given the structure of the model this will tend to place higher weights on minority listeners, all things equal. Next, we have to deal with the markets without minority stations. There are 2 for the race model and 11 for the Hispanic status model. These numbers are small enough that they will not much influence our overall results. In these cases, we set minority fixed costs equal to the estimated non-minority fixed costs.

Table 8 reports averages of the estimated bounds on fixed costs over markets (recalling that upper bounds are used for minority-targeted stations, and that lower bounds are used for majority-targeted stations). When we assume equal per-listener ad prices across groups, the OLS (IV) estimates indicate that per-station fixed costs are roughly \$2.2 (\$2.1) million per year for white stations and \$2.0 (\$1.6) million for black stations (noting that the OLS/IV definition applies to the estimates of the listening equation reported in Tables 4 and 6 above). When we allow for group-specific prices, we find slightly lower fixed costs for black stations of \$1.5 to \$1.8 million. Patterns are quite similar for non-Hispanic vs Hispanic stations: roughly \$2.5 million for non-Hispanic stations and \$1.5 million for Hispanic-targeted stations.

5 Using the Model

5.1 Does the Equimarginal Rule Hold?

Before turning to welfare weights, we first ask whether the current station configurations obey the equimarginal rule. In particular we can ask how the change in total listening with the dollars spent on additional majority-targeted station compares with the analogous figure for minority-targeted stations. That is, how does $\left(\frac{\partial L_B}{\partial N_W} + \frac{\partial L_W}{\partial N_W}\right)/FC_W$ compare with $\left(\frac{\partial L_B}{\partial N_B} + \frac{\partial L_W}{\partial N_B}\right)/FC_B$ across markets? To this end, we calculate the average of the log ratio, for blacks/whites and Hispanics/non-Hispanics. These log ratios average 0.33 (standard error=0.008) and 0.32 (.035), respectively. That is, the number of additional listeners generated by a dollar spent on the marginal minority-targeted station exceeds the number of additional listeners generated by a dollar spent on the marginal majority-targeted station. This suggests that the planner attaches higher welfare weights to majority listeners.

5.2 Welfare Weights

Using the empirical estimates of the model, we can of course calculate welfare weights directly. Tables 9 and 10 present estimates of black and white listener welfare weights from each of the models. With equal ad prices the nested logit OLS model with one nest (inside goods) yields a white welfare weight that is 2 times the black weight. Using the IV estimates, the white weight is 1.56 times the black weight. When we allow for different ad prices across groups, the white weight is roughly three times the black weight using OLS and over twice the black weight using IV.

The welfare weights - and measures comparing groups' - are random variables. Performing hypothesis tests requires measure of their standard errors, which we calculate by bootstrapping. We perform 100 replications (sampling metropolitan areas with replacement), using the resulting coefficients to calculate the welfare weights in each market. We can compare welfare weights in multiple ways. If we define, for example, β_B as the weight attached to black listeners in a metro area and $\overline{\beta}_B$ as the average of the welfare weights across markets, then some methods for comparing welfare weights across groups include: $\overline{\beta}_W/\overline{\beta}_B$, $\overline{\beta}_W - \overline{\beta}_B$, and $\overline{\beta_W/\beta_B}$. The standard errors show that all of the bias measures from the different-price model are statistically significant.

Table 10 compares non-Hispanic and Hispanic welfare weights. Using equal prices we see relatively little evidence of differential welfare weights across groups. Neither of the ratios $\overline{\beta}_{NH}/\overline{\beta}_H$ nor $\overline{\beta_{NH}/\beta_H}$ is significantly different from 1, nor are the differences $\overline{\beta}_{NH} - \overline{\beta}_H$ different from zero. With different ad prices across groups, however, the non-Hispanic weight is 1.5 to 2 times the Hispanic weight (and statistically significantly different from 1), and the difference in the weights is statistically significantly positive. We conclude that entry patterns reveal biases against minority consumers, particularly for blacks relative to whites and less so for Hispanics relative to non-Hispanics. We note here that we also estimated the black-white model with a narrower set of black targeted stations (including only urban). All substantive results were similar.

5.3 Exploring the Difference between the Black and Hispanic Results

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One aspect of the results is rather puzzling: why does entry in this market give rise to a smaller relative weight for blacks (relative to whites) than to Hispanics relative to non-Hispanics? Both groups are of similar relative size, and both groups attract similar amounts of entry. One possibility is sample composition. The 83 sample markets with Hispanic data are, on average, larger than the 127 markets in the black sample; and the Hispanic shares average 30 percent, compared with only 20 percent for blacks in the other sample (see Tables 1a and 1b). Yet, results for blacks do not change when we restrict attention to the markets whose average black share matches the average Hispanic share in the Hispanic sample.

While the preferences differences - between blacks and whites and between Hispanics and non-Hispanics - are similar in that both minorities have preferences that differ sharply from their majority complements, there are some important differences. Table 2 shows how black and white radio listening is divided among black and white-targeted stations (summarized in the last two rows above the total row). Table 3 shows the analogous distribution for Hispanics and non-Hispanic listening to Hispanic-targeted and non-Hispanic-targeted stations. A few interesting differences emerge from this comparison.

First, consider "exporting" by the respective minority groups, the extent of majority listening to minority stations: whites listen to black stations, but non-Hispanics do not listen to Hispanic (Spanish-language) stations. About ten percent of white listening is to black-targeted stations, while virtually no non-Hispanics listen to Hispanic stations. Second, consider "importing" by minority groups. Here, Hispanics listen to non-Hispanic stations far more than blacks listen to white stations. Roughly half of Hispanic listening is to non-Hispanic targeted stations, while only a quarter of black listening is to white-targeted stations.

Welfare weights are functions of the matrix of partial derivatives of listening with respect to entry, as well as station fixed costs. Average listening partials based on the 1-nest IV models are shown in table 11a and b. As with the raw data, whites listen to black stations more than non-Hispanics listen to Hispanic stations; and Hispanics listen to non-Hispanic stations more than blacks listen to white stations.

These patterns affect calculation of the welfare weights directly. With two groups, i and j, the welfare weight for group i is given by: $\beta_i = \frac{1}{\Delta}\left(\frac{\partial L_j}{\partial N_j}FC_i - \frac{\partial L_j}{\partial N_i}FC_j\right)$, where Δ is the determinant of the matrix A:

$$\Delta = \frac{\partial L_i}{\partial N_i}\frac{\partial L_j}{\partial N_j} - \frac{\partial L_i}{\partial N_j}\frac{\partial L_j}{\partial N_i}$$

The relative welfare weight (e.g black relative to white) is therefore given by

$$\frac{\beta_i}{\beta_j} = \frac{\frac{\partial L_j}{\partial N_j}FC_i - \frac{\partial L_j}{\partial N_i}FC_j}{\frac{\partial L_i}{\partial N_i}FC_j - \frac{\partial L_i}{\partial N_j}FC_i}$$

Assuming, as is plausible, that both numerator and denominator are positive, it is easy to see that β_i/β_j is higher as $\frac{\partial L_i}{\partial N_j}$ is higher; and β_i/β_j is lower as $\frac{\partial L_j}{\partial N_i}$ is higher.

If we treat i as the minority and j as the complementary majority, then β_i/β_j is more favorable to the minority as $\frac{\partial L_i}{\partial N_j}$ is higher and as $\frac{\partial L_j}{\partial N_i}$ is lower. That is, the ratio is more favorable to the minority group, the more that minorities import (majority entry promotes minority listening) and the less that minorities export (minority entry promotes majority listening). This is intuitive. Suppose the minority group did not export at all, meaning that the majority did not enjoy the programming targeted at the minority. Then whatever level of entry targeting the minority would arise simply because of the planner's weight on the minority group. If the majority also liked the programming, then we would rationalize the same level of minority-targeted programming with a smaller minority weight. Hence, for any entry configuration, a greater exporting by a group is consistent with a lower welfare weight on that group.

When we compare blacks and Hispanics, we see two patterns in the data (and model analogues of data): $\frac{\partial L_i}{\partial N_j}$ is higher for Hispanics than for blacks; and $\frac{\partial L_j}{\partial N_i}$ is lower for Hispanics than for blacks. These differences work together to make β_i/β_j more favorable for Hispanics (relative to non-Hispanics) than for blacks (relative to whites).

How much of the difference between β_B/β_W and β_H/β_{NH} do these differences in preference patterns explain? We can calculate β_B/β_W using the average values of the cross partials for blacks/whites in Table 11a, along with the different-price fixed costs from Table 8 ($2.12 million for whites, $1.49 million for blacks). The baseline ratio is 0.47, indicating that the market attaches half the welfare weight to blacks that it attaches to whites. When we use the Hispanic value of $\frac{\partial L_i}{\partial N_j}$ (177 rather than 104), the ratio rises to 0.58. When we instead use the Hispanic value of $\frac{\partial L_j}{\partial N_i}$ (68 rather than 216), holding $\frac{\partial L_i}{\partial N_j}$ at its baseline value, the ratio is 1.06. When we use Hispanic values for both cross partials, the ratio is 1.32. The cross partials clearly exert a large impact on the inferred welfare weights. Changing both cross partials changes the result from a 2:1 preference in favor of whites over blacks to 30 percent preference for blacks over whites.

These results show that, while entry in radio broadcasting implicitly favors whites over blacks, the magnitude of the finding arises from a specific pattern of preferences that is absent, for example, for Hispanics vs non-Hispanics.

6 Conclusion

We have shown that with heterogeneous consumers, free entry in markets with fixed costs and imperfect substitutes can reveal substantially different implicit valuations of different types of consumers. Free entry results in station configurations that reflect a planner's implicit valuation of white listeners that is two to three times of the planner's valuation of a black listener, on a per-listener basis; and the non-Hispanic-Hispanic differential is 1.5 to two. The difference between the black and Hispanic results arises from different preference patterns, that relative to Hispanics, blacks export more and import less. The resulting differences in welfare weights are substantial racial disparities, in relation to disparities such as the wage differential, that occupy the attention of policy makers and researchers. Beyond these specific results, this study points to differentiated product markets as contexts where free entry can, contrary to the conventional understanding, give rise to disparate treatment of preference minorities. More research is needed to determine whether this a widespread phenomenon or a feature specific to particular groups of consumers in this industry.

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A More General Conditions for Bias

Here we extend the numeric example of the introduction to show that, in the absence of "importing and exporting" of demand across groups of products, the market's implicit social planner favors the larger group. In particular, in a model that generalizes the example, we show that the implicit welfare weight equals price times a measure of the degree of business stealing. Since business stealing naturally increases in N, this implies that the planner favors the larger group, or at least the group that receives the greater degree of variety as measured by N_i.

We assume that the product-share functions are the same across groups, so that differences in the "taste for variety" do not drive results. Fixed costs can differ across groups, however. Differences in prices will only overturn the result if equilibrium prices are higher for the minority group.

Consider a model with 2 groups that have distinct mutually exclusive preferences: group i only buys product type i. There are N_i varieties of product i and we follow Mankiw and Whinston (1986) in treating the equilibrium quantities and prices as functions of N_i. Each product sells for an equilibrium price $p_i(N_i)$. Because of competition, price may (or may not) decline in the number of products. The population size is M_i, of which a share $S_i(N_i)$ buys one of the symmetric products, so that each product has share $s_i(N_i) = \frac{S_i(N_i)}{N_i}$. We begin by assuming that the only costs are fixed costs, F_i, and then introduce variable costs later. As is natural, we assume that $S(N_i)$ is increasing in N_i while product share $s(N_i)$ is decreasing in N_i.

The degree of business stealing The market share of a firm entering into an N_i product equilibrium can be decomposed as the sum of a "business expansion" effect and a "business stealing" effect. That is, as we move from a $N_i - 1$ product equilibrium to a N_i product equilibrium, the share of the "new" product is the sum of increase in total market share plus the decrease in the market share of the "old" products:[9]

$$s(N_i) = [S(N_i) - S(N_i - 1)] + \big[(N_i - 1)\big(s(N_i - 1) - s(N_i)\big)\big] \tag{5}$$

Clearly, the first term in brackets is business expansion: the increase in total output. The second term is business stealing; it is the absolute value of the decline in the production of the "existing" $N_i - 1$ firms when the N_i-th firm enters. We define the "degree" of business expansion as being the fraction of the new product share that comes from increased total output:

$$\frac{S(N_i) - S(N_i - 1)}{s(N_i)}. \tag{6}$$

[9]Note that $S(N_i) = s(N_i) + (N_i - 1)s(N_i)$ and $S(N_i - 1) = (N_i - 1)s(N_i - 1)$ so that $S(N_i) - S(N_i - 1) = s(N_i) - (N_i - 1)\big(s(N_i - 1) - s(N_i)\big)$. Solving for $s(N_i)$ gives (5)

This fraction varies between zero and one; zero is no business expansion (complete business stealing) and one no business stealing.

We assume a *decreasing degree of business expansion*: the fraction in (6) is strictly decreasing in N_i. This is easy to verify for standard functional forms like the logit and is consistent with a decreasing marginal taste for variety.

In what follows it is useful to define the inverse of this fraction as the "degree of business stealing":

$$\frac{s(N_i)}{S(N_i) - S(N_i - 1)}.$$

This is the ratio of the share of the "new" product to the total market expansion caused by the move from $N_i - 1$ to N_i products; it varies between one and infinity, with infinity being perfect business stealing. Ignoring the discreteness in N_i, we can also define the degree of business stealing as

$$\frac{s(N_i)}{\partial S(N_i)/\partial N_i}. \tag{7}$$

Following the assumption of decreasing business expansion, the degree of business stealing, as measured by (7) is increasing in N_i.

Free entry equilibrium In a free entry equilibrium, ignoring the discreteness of N_i, the free entry number of firms, N_i^e satisfies

$$M_i p_i(N_i^e) s_i(N_i^e) - F_i = 0,$$

or

$$p_i(N_i^e) s_i(N_i^e) = \frac{F_i}{M_i}. \tag{8}$$

This just says that per-capita revenue equals per-capita fixed costs.

Planner's Problem We can compare this to the case where a social planner cares about the consumption of group i, $Q_i(N) = M_i S(N_i)$, according to the function $W_i\big(Q_i(N_i)\big)$. The social planner solves the problem

$$\max_{N_i} \left[W_i\big(M_i S_i(N_i)\big) - N_i F_i\right].$$

Again ignoring discreteness, the social planner's first order condition is

$$M_i \frac{\partial W_i}{\partial Q_i}\frac{\partial S_i}{\partial N_i} - F_i = 0,$$

or

$$\frac{\partial W_i}{\partial Q_i}\frac{\partial S_i}{\partial N_i} = \frac{F_i}{M_i}.$$

Plugging in from the free entry condition (8) at the observed $N_i = N_i^e$

$$\frac{\partial W_i}{\partial Q_i}\frac{\partial S_i}{\partial N_i} = p_i(N_i)s_i(N_i),$$

or

$$\frac{\partial W_i}{\partial Q_i} = p_i(N_i)\frac{s_i(N_i)}{\partial S_i/\partial N_i}. \tag{9}$$

This says that *the implied social welfare weight is equal to price times the degree of business stealing*. With no business stealing, the right hand side equals price, implying that the planner places the "correct" marginal weight of p_i on the good.[10]

Since we assume that the degree of business stealing is increasing in N_i, if prices and fixed costs are equal, then we know that *whichever group has the larger population (and therefore the larger N_i) is implicitly favored by the social planner.* As noted in the text, this implies that, holding total costs constant, total consumption, $Q_1 + Q_2$, could be higher. If fixed costs are different, then we still know that the social planner favors the group that receives more variety. Differences in prices can only potentially overturn that result if prices are higher for the group with less variety.

Note that the result does not apply to the empirical model of the paper, with shares given by $s_i(N_i, N_j)$, because the tastes are not equal and exclusive. Therefore, empirical work is necessary to establish any result for the real industry.

Extension to Variable Cost Consider the case where costs depend on product output, $q_i(N_i) = M_i s(N_i)$. In this case the social welfare problem is

$$\max_{N_i} \left[W_i\big(M_i S_i(N_i)\big) - N_i C_i(M_i s(N_i))\right].$$

The first-order condition is

$$M_i \frac{\partial W_i}{\partial Q_i}\frac{\partial S_i}{\partial N_i} = C_i + N_i mc_i(q_i) M_i \frac{\partial s_i}{\partial N_i}.$$

[10]This is correct only in the traditional case, unlike radio, where there are no consumption externalities. In radio, it would make sense for the social planner to place weight on the unpriced value of listening.

This implies that the planner's marginal value, in excess of marginal cost, is equal to markup times the degree of business stealing:

$$\begin{aligned}
\frac{\partial W_i}{\partial Q_i}\frac{\partial S_i}{\partial N_i} &= \frac{C_i}{M_i} + N_i\frac{\partial s_i}{\partial N_i}mc_i \\
&= p_i s_i + N_i\frac{\partial s_i}{\partial N_i}mc_i \\
&= p_i s_i + \left[\frac{\partial S_i}{\partial N_i} - s_i\right]mc_i \\
&= (p_i - mc_i)s_i + mc_i\frac{\partial S_i}{\partial N_i} \\
\frac{\partial W_i}{\partial Q_i} &= (p_i - mc_i)\frac{s_i}{\partial S_i/\partial N_i} + mc_i \\
\frac{\partial W_i}{\partial Q_i} - mc_i &= (p_i - mc_i)\frac{s_i}{\partial S_i/\partial N_i}. \qquad (10)
\end{aligned}$$

The first line divides the first order condition by M_i and the second line substitutes in from the free-entry equilibrium condition $p_i M_i s_i = C_i$, implying $C_i/M_i = p_i s_i$. The third line uses the fact that

$$\frac{\partial S_i}{\partial N_i} = \frac{\partial(N_i s_i)}{\partial N_i} = N_i\frac{\partial s_i}{\partial N_i} + s_i.$$

The math here is quite similar to the math in Mankiw and Whinston's (1986) analysis of business stealing, except that they set the planner's marginal value to the "correct" level of p and derive the result that entry is excessive, whereas we solve for the marginal planner's value that justifies the observed outcome.

The result in (10) shows, again, that the implicit planner values the group receiving more variety, at least if p_i and mc_i are the same for both groups.

B Figures and Tables

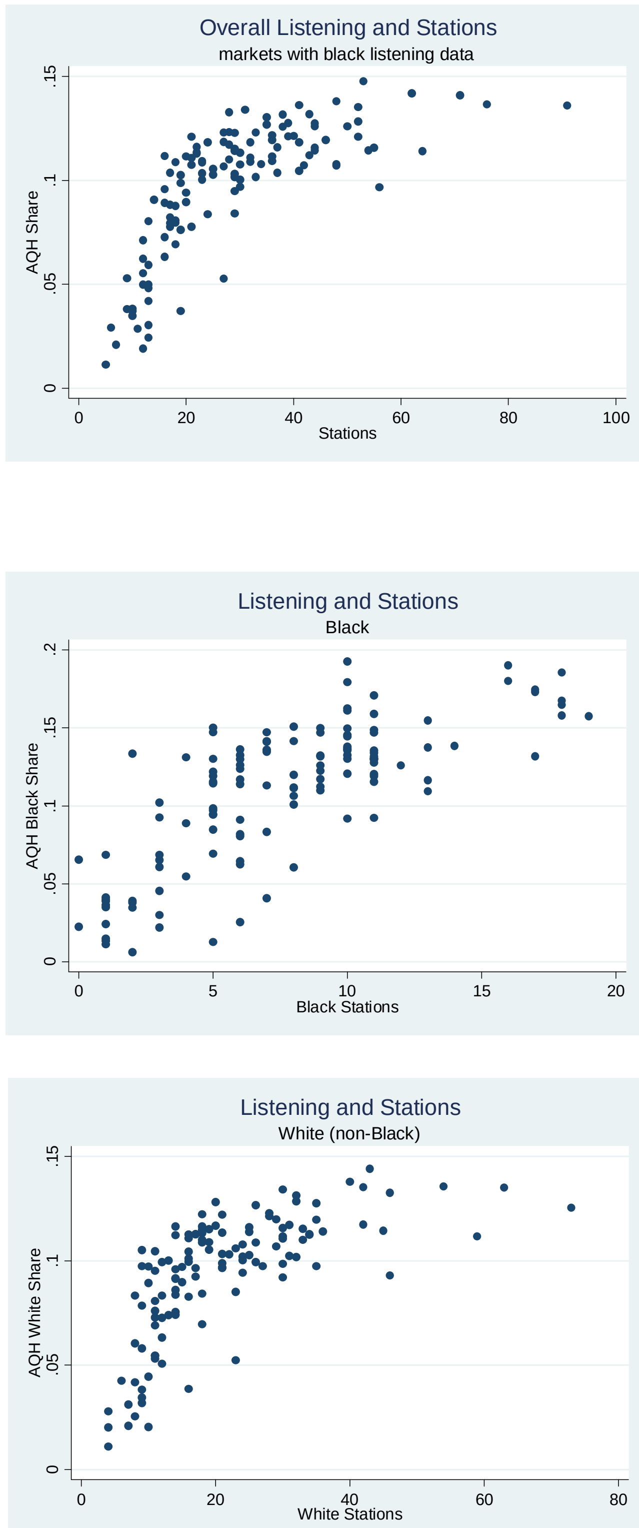

Figure 1: Listening and Stations (Race, inclusive format definition)

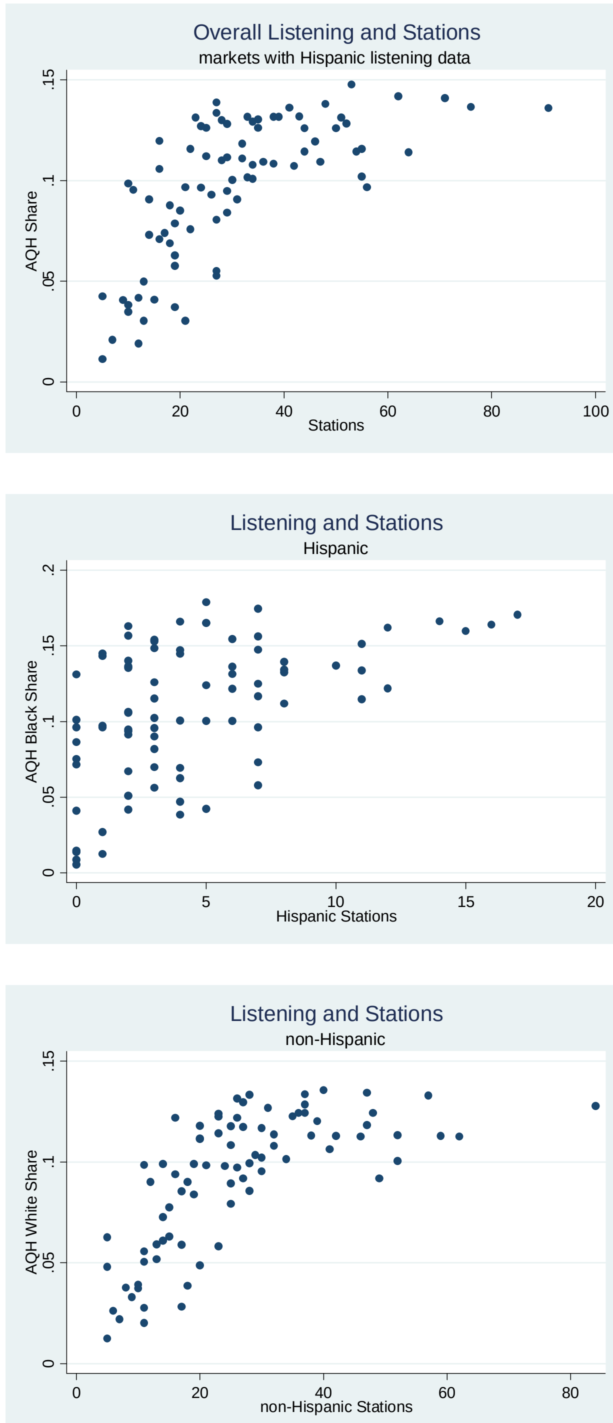

Figure 2: Listening and Stations (Hispanic Status)

Table 1: Summary Statistics

1a: Race

	mean
Overall Listening share	9.77%
white share	9.37%
black share	10.94%
Stations	29
White Stations	21.3
Black Stations	7.7
Ad price per listener	649.8
White Pop	1,013,414
Black Pop	188,842
N	127

1b: Hispanic Status

	mean
Overall Listening share	9.69%
Non-Hispanic share	9.18%
Hispanic share	10.76%
Stations	30.8
Non-Hispanic Stations	26.3
Hispanic Stations	4.5
Ad price per listener	636.6
Non-Hispanic Pop	1,166,220
Hispanic Pop	334,454
N	83

Table 2: Black and White Listening by Format

Format	Black	White	Black Share	Format Share among Blacks	Format Share among Whites
Adult Contemporary	1920	22413	7.89%	5.51%	15.91%
Album Oriented Rock/Classic Rock	200	8258	2.36%	0.57%	5.86%
Classical	243	4279	5.37%	0.70%	3.04%
Contemporary Hit Radio/Top 40	2728	10726	20.28%	7.84%	7.61%
Country	344	14011	2.40%	0.99%	9.95%
Easy Listening/Beautiful Music	18	685	2.56%	0.05%	0.49%
Ethnic	101	477	17.47%	0.29%	0.34%
Jazz/New Age	2295	3722	38.14%	6.59%	2.64%
Middle of the Road	5	165	2.94%	0.01%	0.12%
Miscellaneous	233	1272	15.48%	0.67%	0.90%
News	1440	15251	8.63%	4.14%	10.83%
Nostalgia/Big Band	74	1364	5.15%	0.21%	0.97%
Oldies	446	6605	6.33%	1.28%	4.69%
Public/Educational	97	1178	7.61%	0.28%	0.84%
Religion	3819	5142	42.62%	10.97%	3.65%
Rock	406	11564	3.39%	1.17%	8.21%
Spanish	171	17462	0.97%	0.49%	12.40%
Sports	537	4096	11.59%	1.54%	2.91%
Talk	578	6235	8.48%	1.66%	4.43%
Urban	19160	5970	76.24%	55.03%	4.24%
Black-targeted total	25274	14834	63.00%	72.60%	10.50%
White-targeted total	9541	126041	7.00%	27.40%	89.50%
Total	34815	140875			

Note: Columns 1 and 2, respectively, show the number of black and white listeners to each format, using Arbitron listening data for Fall 2005 linked by station with the formats reported in BIA Kelsey.

Table 3: Hispanic and non-Hispanic Listening by Format

	Hispanic	Non-Hispanic	Hispanic Share	Format Share among Hispanics	Format Share among non-Hispanics
Adult Contemporary	3265	16182	16.79%	8.50%	15.44%
AOR/Classic Rock	960	5800	14.20%	2.50%	5.53%
Classical	259	3425	7.03%	0.67%	3.27%
CHR/Top 40	3671	7958	31.57%	9.56%	7.59%
Country	923	8808	9.49%	2.40%	8.40%
Easy Listening/BM	56	647	7.97%	0.15%	0.62%
Ethnic	130	475	21.49%	0.34%	0.45%
Jazz/New Age	684	4761	12.56%	1.78%	4.54%
Middle of the Road	2	136	1.45%	0.01%	0.13%
Miscellaneous	218	1289	14.47%	0.57%	1.23%
News	955	12377	7.16%	2.49%	11.81%
Nostalgia/Big Band	74	1073	6.45%	0.19%	1.02%
Oldies	1076	4445	19.49%	2.80%	4.24%
Public/Educational	92	1020	8.27%	0.24%	0.97%
Religion	781	4747	14.13%	2.03%	4.53%
Rock	1258	8044	13.52%	3.28%	7.67%
Spanish	20238	674	96.78%	52.71%	0.64%
Sports	366	3560	9.32%	0.95%	3.40%
Talk	456	5148	8.14%	1.19%	4.91%
Urban	2934	14257	17.07%	7.64%	13.60%
Hisp-targeted total	20238	674	96.80%	52.70%	0.60%
Non-Hisp-targeted total	18160	104,152	14.80%	47.30%	99.40%
Total	38398	104826			

Note: Columns 1 and 2, respectively, show the number of Hispanic and non-Hispanic listeners to each format, using Arbitron listening data for Fall 2005 linked by station with the formats reported in BIA Kelsey.

Table 4: Logit Listening Model Estimates for Blacks and Whites

	(1) Black	(2) white	(3) Black	(4) white	(5) Black	(6) white	(7) Black	(8) white	(9) Black	(10) white
Black Format	1.2385	-0.9252	0.2237	-0.0464	0.4552	-0.1436	0.3356	0.1029	0.5509	2.9719
	(0.0591)**	(0.0586)**	(0.0194)**	(0.0133)**	(0.0315)**	(0.0180)**	(0.0233)**	-0.0541	(0.0549)**	(0.8623)**
sigma			0.8443	0.9312	0.6517	0.8282				
			(0.0059)**	(0.0036)**	(0.0186)**	(0.0114)**				
sigma 1							0.8578	0.9304	0.666	0.8152
							(0.0060)**	(0.0036)**	(0.0195)**	(0.0155)**
sigma 2							0.714	0.9985	0.5393	2.2347
							(0.0166)**	(0.0239)**	(0.0563)**	(0.3894)**
Constant	-6.3187	-6.1063	-2.6373	-2.4492	-3.4769	-2.8537	-2.7561	-2.4452	-3.5713	-2.7565
	(0.0347)**	(0.0285)**	(0.0277)**	(0.0155)**	(0.0822)**	(0.0454)**	(0.0308)**	(0.0156)**	(0.0925)**	(0.0658)**
	logit	logit	Nested (1)	Nested (1)	Nested (1)	Nested (1)	Nested (2)	Nested (2)	Nested (2)	Nested (2)
			OLS	OLS	IV	IV	OLS	OLS	IV	IV
Obs.	2138	3372	2138	3372	2138	3372	2138	3372	2138	3372
R^2	0.17	0.07	0.92	0.96			0.92	0.96		

Standard errors in parentheses. * significant at 5%; ** significant at 1%. Inclusive black format definition.

Table 5: Instruments for Black and White Listening Estimates

	(1) sigma 1 white	(2) sigma 2 white	(3) Sigma white	(4) sigma 1 black	(5) sigma 2 black	(6) Sigma black
Black pop (mil)	0.2018	-0.2848	-0.0338	-0.5065	-0.3932	-0.8158
	-0.115	(0.0746)**	-0.1129	(0.1188)**	(0.0502)**	(0.1318)**
White pop (mil)	-0.2567	0.0757	-0.2043	-0.1305	0.0621	-0.0906
	(0.0212)**	(0.0137)**	(0.0208)**	(0.0224)**	(0.0092)**	(0.0248)**
Constant	-3.1485	-0.7364	-3.7912	-2.6045	-0.9877	-3.5026
	(0.0325)**	(0.0206)**	(0.0319)**	(0.0358)**	(0.0139)**	(0.0397)**
F-value	206.53	16.54	184.67	180.86	30.69	172.48
Observations	3372	3704	3372	2138	3689	2138
R-squared	0.11	0.01	0.1	0.14	0.02	0.14

Notes: Standard errors in parentheses. * significant at 5%; ** significant at 1%

Table 6: Logit Listening Model Estimates for Hispanics and Non-Hispanics

	(1) Hisp	(2) Non-Hisp	(3) Hisp	(4) Non-Hisp	(5) Hisp	(6) Non-Hisp	(7) Hisp	(8) Non-Hisp	(9) Hisp	(10) Non-Hisp
Hisp Format	1.0511	-2.4062	0.1458	-0.1121	0.2884	-0.3743	0.1139	-0.1506	0.8083	1.2681
	(0.0750)**	(0.1230)**	(0.0255)**	(0.0336)**	(0.0325)**	(0.0495)**	(0.0264)**	-0.1429	(0.2735)**	-0.945
Sigma			0.8648	0.9232	0.7286	0.8177				
			(0.0070)**	(0.0051)**	(0.0174)**	(0.0145)**				
sigma 1							0.8743	0.9234	0.5032	0.807
							(0.0073)**	(0.0051)**	(0.1200)**	(0.0162)**
sigma 2							0.7376	0.9152	2.2646	1.1578
							(0.0293)**	(0.0294)**	(0.7907)**	(0.1960)**
Constant	-6.1468	-6.2278	-2.5462	-2.5054	-3.1134	-2.9309	-2.5893	-2.5049	-2.9893	-2.9719
	(0.0341)**	(0.0326)**	(0.0313)**	(0.0222)**	(0.0736)**	(0.0591)**	(0.0326)**	(0.0223)**	(0.1447)**	(0.0655)**
	logit	logit	Nested (1)	Nested (1)	Nested (1)	Nested (1)	Nested (2)	Nested (2)	Nested (2)	Nested (2)
			OLS	OLS	IV	IV	OLS	OLS	IV	IV
Observations	1794	2229	1794	2229	1794	2229	1794	2229	1794	2229
R-squared	0.1	0.15	0.9	0.95			0.91	0.95		

Notes: Standard errors in parentheses. * significant at 5%; ** significant at 1%

Table 7: First-Stage Regressions for Hispanic Listening

	(1) sigma 1	(2) sigma 2	(3) sigma	(4) sigma 1	(5) sigma 2	(6) sigma
Hispanic Pop (mil)	0.2225	-0.2539	0.0114	-0.2033	-0.094	-0.2781
	(0.0578)**	(0.0565)**	-0.0575	(0.0526)**	(0.0122)**	(0.0529)**
Non-Hispanic Pop (mil)	-0.2856	0.0599	-0.2675	-0.1558	-0.0023	-0.1578
	(0.0221)**	(0.0218)**	(0.0220)**	(0.0206)**	-0.0047	(0.0207)**
Constant	-3.4284	-0.5734	-3.6867	-2.878	-0.5562	-3.4637
	(0.0431)**	(0.0417)**	(0.0429)**	(0.0403)**	(0.0088)**	(0.0405)**
Observations	2229	2523	2229	1794	2587	1794
R-squared	0.11	0.01	0.14	0.14	0.06	0.17

Notes: Standard errors in parentheses. * significant at 5%; ** significant at 1%

Table 8: Fixed Costs of Station Operation ($mil)

Model	White	Black	Non-Hispanic	Hispanic
1-level OLS, equal prices	2.17	2.01	2.6	2.14
	(0.03)	(0.16)	(0.11)	(0.31)
1-level IV, equal prices	2.06	1.66	2.46	1.75
	(0.05)	(0.08)	(0.10)	(0.12)
1-level OLS, different ad prices	2.24	1.8	2.65	1.85
	(0.08)	(0.16)	(0.11)	(0.29)
1-level IV, different ad prices	2.12	1.49	2.51	1.46
	(0.05)	(0.07)	(0.10)	(0.10)
N	127		83	

Notes: Bootstrap standard errors in parentheses.

Table 9: Relative Welfare Weights by Race

	White	Black	Ratio	Ratio	Difference	Ad price
	$\overline{\beta}_W/\lambda$	$\overline{\beta}_B/\lambda$	$\overline{\beta}_W/\overline{\beta}_B$	$\overline{\beta_W/\beta_B}$	$\overline{\beta}_W/\lambda - \overline{\beta}_B/\lambda$	
1 nest OLS	10,049**	5388**	1.87**	1.93**	4661**	equal
	(1310.10)	(645.00)	(0.24)	(0.24)	(1140.00)	
1 nest IV	3598**	2468**	1.46	1.56	1130	equal
	(945.00)	(172.00)	(0.37)	(0.40)	(931.00)	
1 nest OLS	10,858**	3990**	2.72**	2.91**	6868**	different
	(1416.00)	(483.00)	(0.35)	(0.35)	-1,252	
1 nest IV	3886**	1846**	2.10**	2.36**	2040**	different
	(1021.00)	(128.00)	(0.52)	(0.60)	(1002.00)	

Bootstrap standard errors in parentheses based on 100 replications. * significant at 5%; ** significant at 1%. Null hypothesis for ratio columns is that ratio=1. Weights based on fixed costs computed using bounds approach.

Table 10: Relative Welfare Weights by Hispanic Status

	Non-Hisp	Hisp	Ratio	Ratio	Difference	Ad price
	$\overline{\beta}_{NH}/\lambda$	$\overline{\beta}_H/\lambda$	$\overline{\beta}_{NH}/\overline{\beta}_H$	$\overline{\beta_{NH}/\beta_H}$	$\overline{\beta}_{NH}/\lambda - \overline{\beta}_H/\lambda$	
1 nest OLS	8947**	6824**	1.31	1.49**	2123	equal
	(1589.00)	(1295.00)	(0.23)	(0.22)	(1434.00)	
1 nest IV	3499**	3364**	1.04	1.25	134	equal
	(632.00)	(566.00)	(0.16)	(0.20)	(631.00)	
1 nest OLS	9827**	5511**	1.78**	2.24**	4316**	different
	(1744.00)	(1119.00)	(0.34)	(0.29)	(1545.00)	
1 nest IV	3835**	2751**	1.39*	1.90**	1084*	different
	(695.00)	(471.00)	(0.22)	(0.30)	(639.00)	

Bootstrap standard errors in parentheses based on 100 replications. * significant at 5%; ** significant at 1%. Null hypothesis for ratio columns is that ratio=1. Weights based on fixed costs computed using bounds approach.

Table 11: Model Estimates of the Change in Group Listening with an Additional Station

(a) black/white

...with an additional	Change in black listening	Change in white listening
black station	317	216
white station	104	500

(b) Hispanic/non-Hispanic

...with an additional	Change in Hisp listening	Change in non-Hisp listening
Hispanic station	509	68
Non-Hisp. station	177	516

Notes: average values of the partials across all sample markets.
N=127 for blacks and whites, while N=83 for Hispanics and non-Hispanics.