Measuring the Impact of Air Quality on Housing Markets and Residential Choices in Southern California

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Abstract

The purpose of this paper is to provide a comprehensive approach for measuring the economic gains to households from air quality improvements based on a well defined locational equilibrium model. The approach taken in this paper allows us to estimate the indirect utility function of households and the underlying distribution of household types. We can, therefore, construct welfare measures which properly take into consideration the adjustments of households in equilibrium to non-marginal changes in air quality. These types of measures are outside the limited scope of more traditional approaches. The empirical approach of this paper, hence, provides for the first time an internally consistent framework for estimation and applied general equilibrium welfare analysis. We also compute welfare measures for marginal changes in air quality and compare them with those obtained from hedonic regressions. We implement our empirical framework using data from Southern California, an area which has experienced drastic improvements in air quality during the past 20 years. Our findings are by and large supportive for our approach and suggest that accounting for general equilibrium effects in applied welfare analysis is important.

1 Introduction

Over the past twenty years there have been dramatic improvements in air quality in Southern California. This area was widely regarded as one of the most severely polluted regions in the United States. Ozone concentrations (based on measures used in EPA's standard setting) have been cut by fifty percent since 1980. The number of days in Los Angeles County exceeding the Federal one hour standard in 1998 was one fourth that experienced in 1980.[1] These changes are among the largest reductions to result from federal air quality regulations. Unfortunately, we know very little about the economic gains to households that resulted from these air quality improvements.[2]

Locational equilibrium models imply that households respond to changes in local conditions. As a result, evaluations of large new policy changes must be based on *different* reduced form relationships (i.e. hedonic price and wage regressions) to estimate the benefits of the changes.[3] Alternatively, empirical approaches must recover the structural parameters of the underlying equilibrium model, which then can be used to evaluate new policies. However, there have been no attempts to recognize the restrictions implied by an equilibrium sorting in the estimation of the parameters of the structural model. These neglected adjustments are part of a larger problem. Econometric research based on reduced form relationships often fails to account for the complex self-selection and endogeneity problems which arise in equilibrium.

This paper responds to these challenges in four ways. First, we use the properties of a locational equilibrium model to estimate consistently household preferences for air quality, local public goods (principally public education), and housing. Second, because the model

[1]The Air Resources Board's ozone data summaries for Los Angeles County indicate that the member of days exceeding the Federal one-hour standard in 1980 was 156. In 1998 the number was 36. Ozone concentrations, measured using the maximum one hour readings were 0.49 ppm in 1980 and only 0.22 in 1998.

[2]While the recent prospective analysis of air pollution reductions does attempt to take account of the beneficial effects of these changes, it is done in terms of physical effects at an aggregate level, with little ability to disaggregate to regional changes.

[3]At a conceptual level, this point is widely appreciated (Bartik, 1988; Palmquist, 1988).

allows households to have heterogeneous preferences for air quality and other public goods it is possible to describe how an exogenous change in air quality will induce a new equilibrium sorting of households including the associated price adjustment. Third, we use the model to compute estimates of the marginal willingness to pay for air quality improvements and compare them to more approximate estimates derived from the most current hedonic property value models developed for ozone pollution. Finally, we use the preference estimates to compute the general equilibrium sorting and housing prices that would result from the dramatic reductions in air pollution that have taken place in this region. These types of measures are outside the limited scope of more traditional approaches. The empirical approach of this paper, thus, provides for the first time an internally consistent framework for estimation and applied general equilibrium welfare analysis.

Our findings suggest that consistent estimates of preferences can be developed and that partial equilibrium measures of the benefits from air pollution reductions from these models fall within the bounds we would expect, based on the hedonic models. The estimates from the equilibrium location model are smaller and can be more consistently linked to differences in local incomes, housing costs, and the quality of public schools. In contrast to the hedonic framework these estimates also provide measures of the parameters of housing demand. Moreover they also provide the information necessary to compute a new general equilibrium sorting of households after large air quality changes. We find that large changes in air quality are likely to have significant general equilibrium effects. They affect both the composition of populations within communities and the resulting equilibrium in the local housing markets. In particular, we find that an intra-marginal improvement of air quality increases population sizes and housing prices in lower amenity communities, thus partially off-setting the gains from better air quality. In contrast, high amenity communities experience population declines and lower housing prices, which increases the gains to households living there. The distribution of the willingness to pay which controls for the general equilibrium effects is more dispersed than the one based on partial equilibrium measures.

The remainder of the paper is organized as follows: Section 2 reviews the previous literature on measuring the benefits of air quality improvements. Section 3 reviews and extends the conceptual empirical work of Epple and Sieg (1999) in developing the locational equilibrium estimator. We apply this framework along with conventional hedonic property value regressions using a unique data set for the Los Angeles metropolitan area. Section 4 presents an extended discussion of the data set that we assembled from various sources for this analysis. Section 5 reports the estimation results. Section 6 discusses how to construct benefit estimates from each method. It also shows how to compute the new locational equilibrium which arises after an intra-marginal change of air quality and the associated welfare measures. Section 7 offers some conclusions.

2 The Previous Literature

After early efforts to estimate the incremental willingness to pay for ozone reductions in Los Angeles (Brookshire, Thayer, Schulze, and d'Arge, 1982), the majority of the revealed preference estimates for the benefits from air quality changes have relied on hedonic models and used measures of particulate matter to characterize air quality conditions (Smith and Huang, 1993, 1995). Several methodological lessons emerged from this early literature, but none of the studies was able to consistently estimate the marginal willingness to pay function for air quality improvements. Four recent studies claim to offer substantial advances over these initial works. This section describes the general lessons from the first round of hedonic research in this area and, then, outlines the findings and potential limitations for each of the newest studies.

The early work established that it was reasonable to expect that home buyers took account of air pollution as a site specific amenity and that the significant negative relationship between housing values and air pollution was not a statistical "artifact". The conditions for identifying the function describing the marginal rate of substitution between air quality and a numeraire were acknowledged to be more complex than initially outlined by Rosen

(1974) because of: the nonlinearity in household's budget constraints (and related joint determination of the marginal prices for housing and site characteristics); the importance of additional, correlated but unobservable, characteristics that could determine households' locational choices; and the supply responses of housing producers to households' demands for dwelling and site characteristics.

Finally, it was found that early data limitations had important effects on the results from hedonic models. Use of census data with owners' beliefs about the sales price for their home (in contrast to a market price) and reliance on selected samples (screening on criteria for federal mortgage subsidies) were acknowledged to be important limitations. Nonetheless, Smith and Huang (1995) conclude that the estimated marginal values for reductions in particulate matter did consistently respond to income, increases in air pollution (as measured by TSP), and local housing market conditions. They do find that use of estimates of these marginal values without adjustment for how they were developed leads to dramatic differences, as much as three times larger estimates, in the benefits attributed to air quality improvement.

The newest work considers two extensions to the literature: (a) identification of the second stage marginal willingness to pay function based on either maintained preference restrictions (Chattopadhyay, 1998) or an assumption of adequate variation in marginal price functions thru the pooling of separate marginal estimates for air quality improvements across markets (Beron, Murdoch, and Thayer, 1999); and (b) development of an instrument for air pollution that allows for departures from a random assignment of pollution readings across the cross-sectional units used in a hedonic property model (Chay and Greenstone, 1998).

Only one of the four studies overcomes the problems in the earlier literature. Kahn (1997) and Chay and Greennestone (1998) rely on owners' reports of their housing values from the Census and not actual selling prices. Chattopadhyay (1998) relied on a selected sample for Chicago available through FHA mortgage applications and had to assume a single, specific preference function to identify the MWTP function using housing sales for this one area. Beron et al. (1999) use housing samples for Los Angeles, Orange, Riverside, and San Bernardino counties over the period 1980 to 1995. They include as pollution measures the annual average of daily maximum ozone readings and the annual average for total suspended particulates along with their primary focus – annual average visibility. Separate hedonic models are reported for each year, with the marginal willingness to pay estimates averaged for a census tract, pooled across years, and specified as functions of census tract characteristics. Use of different hedonic price functions alone is not sufficient to identify the second stage. As Palmquist (1991) suggests, and Ohsfeldt and Smith (1985) confirm, the multiple markets must differ significantly to identify the marginal rate of substitution equations. In this application appreciation in the level of housing price seems to be the primary source of the differences in marginal prices for visibility over time. The estimated parameters for visibility, ozone, and TSP seem quite stable, with the proportionate increase in price due to a mile increase in visibility ranging from a low of .03 to a high of .08 over the sixteen years models were estimated. There is a further problem with their second stage model in that variation in other socio-economic characteristics cannot be measured for years other than 1980 and 1990. Apparently the authors specified that the values of these variables for each census tract measures in each year between 1980 and 1995 corresponded to the nearest census year.[4]

Thus, despite a considerable re-awakening of interest in using revealed preference models to estimate the benefits from air quality improvements, all the evidence available to date has been compromised by limitations in the data used to investigate how air quality influences behavior.

[4]The authors do not explain how these decisions were made. Private correspondence suggests that they assumed a 9.5 percent rate to annualize the marginal prices.

3 Locational Decisions and the Valuation of Environmental Amenities

3.1 Background

Most theoretical locational equilibrium models in urban economics and local public finance assume that households are mobile and make locational choices based on the mix of site specific public goods and the housing market conditions in different communities (Tiebout, 1956). If these models provide a good description of observed economic behavior, we should expect that the models' predictions of the distribution of households among communities would correspond with distributions observed in the data. Additionally, we should be able to match the levels of public good provision and the estimated housing prices with those predicted by our equilibrium models. These ideas provide the background for the locational equilibrium estimator recently developed by Epple and Sieg (1999). This method thus uses a specification of a locational equilibrium model in the estimation process, and thereby, takes account of the endogeneity of local public good provision, prices, and the underlying distribution of households across communities.[5]

Two aspects of these locational equilibrium models are important to highlight from most of the earlier models used to measure households' values for site specific amenities. First, the hedonic property model assumes there exists a continuum of alternatives in the attributes of a house and its location that might contribute to its worth to a consumer. These models draw a distinction between housing characteristics and the public goods conveyed by location. We retain the assumption of a continuum of choice in structural characteristics and lot size. So the "law of one price", as realized in the specification of the geographic domain for the hedonic price function, determines how we adjust for

[5]This framework was originally developed by Epple and Sieg (1999) and Epple, Romer, and Sieg (1999) to study the equilibrium properties of model describing the interaction between individual choices, housing markets, and collective choices that determine tax and spending decisions within a system of local jurisdictions. Parts of this approach are similar in logic to the work by Berry, Levinsohn, and Pakes (1995) which was recently applied to study school choice in California by Bayer (1999).

heterogeneity in housing characteristics. However, we do not require that this continuum be present for aspects of a location that are important to household's choices. Indeed, the second feature of the model maintains that public goods are available in a finite number of alternatives. Each community a household might choose offers a different vector of public goods. In this setting, households are assumed to select a best community and then to purchase an "amount" of housing within it. Thus, the set of communities comprising the individual's choice set determines the extent of the market for the public goods available. The locational equilibrium reveals the willingness to pay for public goods. With sufficient restrictions these sorting conditions provide sufficient information to identify the parameters of the indirect utility function.

3.2 A Locational Equilibrium Model

The starting point is a model of residential decisions in a system of multiple communities.[6] The metropolitan area consists of a finite number of communities which in our case are school districts. Each is assumed to have fixed boundaries. There is a continuum of households living in the metropolitan area. Households differ with respect to income, y. Moreover, households with similar income can differ in their valuation of public goods. This unobserved heterogeneity is captured by a taste parameters for public goods, α. A household is fully characterized by its endowment-preference tuple (α, y). The continuum of households in the metropolitan area is implicitly described by the joint distribution of (α, y), which is represented by the density, $f(\alpha, y)$.

A household living in community j has preferences defined over a composite of local public goods and environmental amenities. We will assume the local public good has two components. One of them is at least partially financed by a property tax and for our application will correspond to local public education.[7] The second is our exogenous environmental

[6] Our model is a natural extension of earlier work by Epple, Filimon, and Romer (1984) and Epple and Romer (1991) which was first used in empirical work by Epple and Sieg (1999) and Epple et al. (1999).

[7] We could also extend our model and include a state wide income tax into the analysis.

amenity that varies across the communities. It is not affected by local community decisions and is determined by the mix of point and mobile sources of air pollution determined outside the choice process described in our analysis. Each community also provides a local housing good, h_j, and a composite private good, b_j. Denote with p_j the gross-of-tax price of a unit of housing services in community j. The preferences of a household are represented by a utility function, $U(\alpha, h_j, b_j, g_j)$ that satisfies the standard assumptions. Households maximize utility with respect to a budget constraint:

$$\begin{aligned} &\max_{(h_j, b_j)} \; U(\alpha, h_j, b_j, g_j) \\ &s.t. \;\; p_j \, h_j \; = \; y \; - \; b_j \end{aligned} \tag{3.1}$$

Household preferences will be represented in our structural model using the indirect utility function derived by solving the optimization problem given in equation (3.1). We assume that the indirect utility function is given by the following specification:

$$V(\alpha, y, g_j, p_j) \; = \; \left\{ \alpha \; g_j^\rho \; + \; \left[e^{\frac{y^{1-\nu}-1}{1-\nu}} \; e^{-\frac{Bp_j^{\eta+1}-1}{1+\eta}} \right]^\rho \right\}^{\frac{1}{\rho}} \tag{3.2}$$

Next, consider the slope of an "indirect indifference curve" in the (g_j, p_j)-plane:

$$\begin{aligned} M(\alpha, y, g_j, p_j) \; &= \; \frac{dp_j}{dg_j}\Big|_{V=\bar{V}} \\ &= \; \frac{\alpha \; g^{\rho-1} \; [e^{\frac{y^{1-\nu}-1}{1-\nu}}]^{-\rho} [e^{-\frac{Bp^{\eta+1}-1}{1+\eta}}]^{-\rho}}{B \; p^\eta} \end{aligned} \tag{3.3}$$

$M(\cdot)$ is monotonic in y given α (and vice versa), hence indifference curves in the (g_j, p_j)-plane satisfy the "single-crossing" properties. These single crossing properties permit a structure in which necessary conditions for an equilibrium can easily be characterized.

It is also important to consider their implications for the restrictions used to measure the value of changes in environmental quality. The Epple-Sieg model describes a mixed discrete/continuous demand process. It also allows prices to be determined through the equilibrium sorting of households among communities. This framework contrasts with the three common revealed preference approaches to non-market valuation. Conventional hedonic models reveal marginal willingness to pay for site amenities because we assume there exists a continuum of choice. In the case of random utility models we assume there exists a discrete array of alternatives and generally describe how policies affect the expected value of the maximum utility that each individual can realize. Willingness to pay in this setting is defined to hold the expected value of that maximum utility constant and the error is assumed to reflect unobserved individual heterogeneity (known by individuals being modeled but unknown to the analyst). In this context a change in quality for any choice alternative does have welfare implications. Finally, with other private goods, where weak complementarity is used to provide the link between their consumption and environmental quality, we must also assume what is referred to as the Willig (1978) condition. This restriction assures it is possible to consistently recover Hicksian welfare measures (Bockstael and McConnell, 1993) from the weakly complimentary commodity. In the context of equation (3) it requires that $\frac{\partial M}{\partial y} = 0$.

Use of the properties of the locational equilibrium relaxes assumptions of all three models. The single crossing replaces the Willig condition that $\frac{\partial M}{\partial y}$ must equal zero with the requirement that it must be positive. This condition restricts the relative size of the change in the marginal utility of income with changes in g in comparison to p.[8] It does not require a continuum of choice in public goods. As a result it is possible, as Bockstael and McConnell (1999) suggested recently, to have changes in environmental quality and not observe a behavioral response. As these authors explain:

> "Individuals will not change their behavior if they cannot adjust at the margin and if their next best alternative generates less utility than their current choice, even with environmental degradation. This is another way of saying that a

[8] Note that

$$\frac{\partial M}{\partial y} = \frac{V_y}{V_p^2}[h^* V_{gy} + \pi_g V_{py}] \tag{3.4}$$

where h^* is the Marshallian demand for housing, $\pi_g = V_g/V_y$ is the virtual price for g, $V_{gy} > 0$ and $V_{py} < 0$.

continuum of environmental quality alternatives rarely exists for the individual. ...An individual may not incur the transactions of relocating as a result of air quality degradation, unless the deterioration in air quality is quite large. In each case the individual may, instead, suffer in (behavioral) silence." (Pg. 26)

Prior to the Epple-Sieg framework revealed preference models had no specific means to accommodate these types of non-responses. Their model allows these cases to be considered by using the properties of the equilibrium sorting of heterogeneous households among a discrete set of alternatives, along with the restrictions incorporated in the indirect utility function to estimate the model.

Let (g_i, p_i) and (g_j, p_j) be the level of public good provision and gross-of-tax housing price in community i and j, respectively, and suppose that some individuals prefer (g_j, p_j) and others prefer (g_i, p_i). Then locational choices in equilibrium will satisfy three properties: (1) boundary indifference, (2) stratification, and (3) ascending bundles. Boundary indifference implies that the set of individuals indifferent between the two communities is given by the set of (α, y)'s such that:

$$V(\alpha, y, g_j, p_j) = V(\alpha, y, g_{j+1}, p_{j+1}) \tag{3.5}$$

For the given parameterization of the indirect utility function the indifference locus between community j and $j+1$ satisfies:

$$\begin{aligned} \ln(\alpha) - \rho\left(\frac{y^{1-\nu}-1}{1-\nu}\right) &= \ln(Q_{j+1} - Q_j) - \ln(g_j^{\rho} - g_{j+1}^{\rho}) \\ &= K_j \end{aligned} \tag{3.6}$$

where K_j denotes the community specific intercept and Q_j is given by:

$$Q_j = e^{-\rho\frac{Bp_j^{\eta+1}-1}{1+\eta}} \tag{3.7}$$

For this parameterization, the boundary indifference conditions in equation (3.6) imply a set of non-intersecting planes in the $(\ln(y), \ln(\alpha))$ space. Define the set of agents, C_j, living in community j:

$$C_j \;=\; \left\{ (\alpha, y) \;\middle|\; K_{j-1} \;\leq\; \ln(\alpha) \;-\; \rho \left(\frac{y^{1-\nu} - 1}{1-\nu} \right) \;\leq\; K_j \right\} \tag{3.8}$$

The single-crossing properties imply that, holding tastes constant, households will stratify according to income levels into different communities. More affluent households live in communities with high amenities, poorer households prefer communities with lower levels of public good provision. Similarly, holding income constant, households will stratify according to tastes for the public goods. Households with high valuations of public goods will tend to live in the high amenity communities, while households with low valuations of public goods will prefer the cheaper communities with lower amenity levels.

The ascending bundles implies that the ranking of communities according to housing prices must be the same as the ranking according to the public good vector available in these communities. A measure of households living in community j is given by:

$$P(C_j) \;=\; \int_{C_j} f(\alpha, y) \; d\alpha \; dy \tag{3.9}$$

Based on these results, we can characterize the distributions for income, housing expenditures, and tastes in each community j. For example, mean housing expenditures in community j are given by:

$$\mu_j(r) \;=\; \int_{C_j} B \; p_j^{\eta+1} \; y^{\nu} \; f(\alpha, y) \; d\alpha \; dy \tag{3.10}$$

Similarly we can characterize higher moments of the housing expenditure distributions as well as their quantiles and order statistics.

Public good provision is likely to consist of a number of factors including both school and environmental quality. To capture the impact of the different amenities on public good provision, it is useful to assume that that g_j satisfies the following index assumption:

$$g_j \;=\; s_j \;+\; \gamma \; e_j \;+\; \omega_j \tag{3.11}$$

where s_j denotes school quality, e_j is a measure of environmental quality and ω_j captures the unobserved characteristics of the community.

The estimation procedure outlined in the next section exploits only the necessary conditions for residential choices in equilibrium in defining identifying restrictions. As a result, we do not need to solve numerically for the equilibrium allocations as part of the estimation algorithm. Thus, our estimator does not require that the housing markets are completely specified. However, we need to compute the equilibrium of the model as part of the computation of welfare measures for intra-marginal changes in the provision of public goods. We then close the locational equilibrium model by assuming that there exists a housing supply function in each community with constant price elasticity.[9]

3.3 Computation and Estimation

The structure of the model suggests implementing the estimation procedure in two steps. In the first step, quantiles of the distributions of housing expenditures are matched with their empirical counterparts along the lines suggested by Epple and Sieg (1999) treating the community specific intercepts as fixed effects. For every parameterization of the joint distribution of income and tastes for the population of the metropolitan area and the indirect utility function of the households, the model determines a joint distribution of income, housing expenditures and taste parameters for each community. The estimation strategy is based on the idea that the difference between the empirical quantiles of the distributions of housing expenditures observed in the data and the quantiles predicted by the model should be small, if the model is evaluated at the true parameter value. Equation (3.6) implies that quantiles of the housing expenditure distribution of community j depend on g_j only through the community specific intercepts K_j. Following Epple and Sieg (1999), we can treat the K_j's as unknown parameters or (preferably) constrain them to replicate the characteristics observed for each community's population. A subset of the parameters of the

[9] Since we only focus on locational equilibrium, we do not need to specify the collective choice mechanism which determines public good provision. See Epple et al. (1999) for such an analysis.

model can then be estimated using a Minimum Distance Estimator. One of the advantages of matching housing expenditure distributions instead of income distributions is that we can additionally identify and estimate the parameters of the indirect utility function which related to the housing demand of households.[10]

In the second step, the levels of public good provision implied by the first round estimates are matched with those observed in the data, conditional on differences in housing prices and other amenities. The model implies that the observed levels of public good provision. Given estimates for the community specific intercepts, $K = (K_1, ..., K_J)$ and the first stage parameters, the values of the indexes implied by the model for the public goods provided in each community, g_j can be computed. That is, solving the equation of the community specific intercept for the levels of public good provision, yields a recursive representation for these levels of public good provision. For the parametric specification of the indirect utility function we obtain:

$$g_{j+1} = \left\{ g_1^\rho - \sum_{i=2}^{j+1} (Q_i - Q_{i-1}) \exp(-K_i) \right\}^{1/\rho} \tag{3.12}$$

This restriction together with equation (3.11) provides the orthogonality conditions which can be exploited in a second stage estimator. Following the approach as outlined in Epple and Sieg (1999) we construct an instrumental variable estimator for the remaining structural parameters of the model. The second step completes the estimation procedures and allows us to recover almost all structural parameters of the underlying model.[11]

In summary, the main methodological innovation in the estimation procedure of this paper is that we match predicted housing distributions with those derived from panel data of housing transactions in each community. This feature distinguishes this paper from Epple and Sieg (1999) who focus instead on income distributions in the first stage of the estimation procedure. This approach simplifies identification and estimation of the parameters which

[10] In order to implement these estimators, we also need to estimate prices per unit of housing independently of the structural model. We will discuss this point in detail in the next section of the paper.

[11] Note that this approach is similar in spirit to work in the differentiated products literature by Berry (1994) and Berry et al. (1995) although the actual implementation differs significantly.

govern the demand for housing. It also allows us to treat income as a latent variable in the estimation procedure which is particularly useful if one believes that residential choices and housing demand are based on a more comprehensive measure of income than current income.

4 Data

Our analysis focuses on the Los Angeles Metropolitan Area which consists of the area west of the San Gabriel Mountains and includes parts of five counties: Los Angeles, Orange, Riverside, San Bernardino and Ventura. We assume that the school district corresponds to the community a household selects in making its locational decisions. To implement the model household characteristics by school district, quality measures for local public education, data on housing markets, and air quality measures are required. We discuss the data sources and the main empirical features of these data in this section. Each school district offers a different level of local public education. Air quality can also differ by school district, but these differences are determined by past locational decisions of the point sources of pollution, the overall level and patterns of automobile traffic, the air diffusion system (linking point and mobile emissions to ambient conditions), and the topography and weather conditions in the area. The community or school district realizes a level of air quality based on these exogenous conditions.

4.1 Housing Markets

A comprehensive data base on housing markets in the LA metropolitan area was assembled based on housing transactions collected by Transamerica Intellitech. These data contain housing characteristics and transaction prices for virtually all housing transactions in Southern California between 1988 and 1992. Table 1 reports means of the main variables in the housing sample by counties for these years.

Table 1: Descriptive Statistics of the Housing Sample

Variable	Orange	Riverside	San Bernardino	Ventura	Los Angeles
Number of Observations	40894	33132	24493	14817	109529
Market Value of House	253315	139771	151313	244888	243889
Number of Bathrooms	2.16	2.07	2.10	2.24	1.94
Number of Bedrooms	3.33	3.26	3.27	3.49	3.05
Lot Size	0.16	0.24	0.21	0.22	0.19
Square foot Building	1748	1627	1615	1838	1591
Pool	0.16	0.12	0.13	0.15	0.17
Fireplaces	0.26	0.84	0.79	0.79	0.54
Age	23.8	9.7	16.8	17.4	37.0

Means of housing values and structural characteristics by county.

In general, these data are quite consistent across counties. Significant differences across counties are present for housing values, which is one of the main variable to be explained in our research. Mean values are higher for the LA, Orange, and Ventura counties. Unlike the other two counties, these are coastal counties. In addition, two of the counties (LA and Orange) are more urban than the others. Being further from the central city, it is not surprising that Ventura, San Bernardino, and Riverside counties would have the youngest houses on average. Of the five counties, these three experienced the largest population growth from 1980 to 1990, with Riverside and San Bernardino having the largest population growth in the state at 67 and 54 percent respectively (California Department of Finance 1990). Consequently, it is not surprising that they would have younger houses on average.

One potential drawback associated with using California data relates to Proposition 13. It has been argued in the literature that Proposition 13 created a lock-in effect on house owners. A household faces a tax on mobility because property taxes on newly purchased property are based on the market value. If the market value exceeds the assessed value, the revaluation creates additional mobility costs. O'Sullivan, Sexton, and Sheffrin (1995) provide a detailed analysis of the quantitative importance of these lock-in effects. They conclude that "the effects [of Proposition 13] on mobility and loss of economic welfare from this lock-in effect are small ... We estimate that for the average household a 13 percent inflation rate will lengthen the average time between moves by only approximately two months" (Pg. 138).

Closely related to the lock-in effect is the question relating to turn-over in housing markets and the representativeness of our housing transaction sample. As mentioned above, our sample contains a large number of housing transactions for each school district. So there is no question that there is a large degree of turn-over in the housing market, which is also indirect evidence for the fact that mobility costs are likely to be small. Nevertheless, we need to establish that our sample is a representative sample of the underlying housing stock of each school districts. Some information pertinent to this question is available from the 1990 US Census. In particular, the US census reports histograms on housing expenditures, the number of bedrooms, and the age of houses which we compare to our sample.

The pattern of housing expenditures across communities is quite close in the two data sets, with a correlation of 0.99. However, prices tend to be uniformly higher in the US Census. Across our 92 school districts, prices are 6 to 12 percent higher in the census (inter-quartile range). This might suggest smaller homes are over-sampled in the transactions data, but the data do not confirm this hypothesis. In fact, 34 percent of houses in the US census have two or fewer bedrooms, compared to only 19 percent in our sample. In addition, homes are much younger in our data set. Fifteen percent of our houses are younger than 1 year, whereas in the 1990 census only three percent of homes were built in 1989-1990. The difference is particularly striking in Ventura, San Bernardino, and Riverside counties. There, 36 percent of our sample is younger than 1 year, compared to 6 percent built in 1989-1990 in the census. The over-sampling of new homes is not surprising, because newly built homes will automatically show up in a data set of housing transactions, whereas older homes will only show up when they turn over.

Additional information is provided by the empirical distribution of housing tenure for the five counties of interest based on data collected by the 1990 U.S. Census. We find that between 70 and 80 percent of all households change houses within 10 years. Given the scope of our housing data, we expect to capture most of these housing transactions in our sample. Approximately 20 to 30 percent of the houses in the US Census have a housing tenure which is greater than 10 years. By construction our sample only contains a fraction of these house.

4.2 Inter-community Housing Prices

An important aspect of the empirical analysis of a model of households' locational and housing choices involves constructing reliable inter-community housing price indexes. These price indexes should control for the observed differences in the quantity and quality of housing consumed within and across communities.

There is a substantial literature on the development of quality adjusted quantity indexes for heterogeneous commodities. Indeed, one interpretation of the hedonic model has been in these terms (Triplett, 1990). By separating the effects of a continuum of choices for structural characteristics from decisions to obtain the local public goods and amenities we assume it is possible to unbundle the local public goods and amenities from the effects of structural and lot characteristics. This approach focuses attention on adjusting for the heterogeneity in houses by developing a price index for a 'homogeneous' (e.g. adjusted for structural characteristics and lot size) unit of housing in each community. We express this price as an annualized rent. The logic of this adjustment uses the procedure outlined in Poterba (1992). The market value of a house n located in community j at time t can be converted into the imputed rent which measures the annual flow and is denoted by r_{jnt}. Since California's tax reform implies that tax rates are for all practical purposes equal among the school districts the change from housing values to imputed rents is only a rescaling of the dependent variables. We also observe a vector of housing specific characteristics denoted by z_{jnt}. Let u_{jnt} denote the unobserved housing characteristics.

Thus the quality adjusted units of housing is given by the following equation:

$$h_{jnt} \quad = \quad e^{\delta' z_{jnt} + u_{jnt}} \tag{4.13}$$

By definition, rent measured for a quality adjusted unit is the product of the adjusted housing price and the number of quality adjusted housing units, i.e. $r_{jnt} = p_{jt}\, h_{jnt}$. Substituting equation (4.14) into the identity above and taking logarithms, we obtain the following regression model:

$$\ln(r_{jnt}) \quad = \quad \ln(p_{jt}) \; + \; \delta' \, z_{jnt} \; + \; u_{jnt} \tag{4.14}$$

which can be used to construct housing price indexes for each community in the sample.

We estimate a large number of different regression models of the type described above. Sieg, Smith, Banzhaf, and Walsh (1999) provide a comprehensive discussion of all those regressions. This comparison of alternative approaches suggests that housing price estimates are robust across different econometric specifications. They do not depend significantly on spatial and temporal aggregation schemes used in constructing the data set and the community choice set. Summary statistics for housing price estimates based on the simple fixed effects regression are reported in Table 2. The estimates indicate that relative prices differ by as much as six to one across communities, although the large majority of the housing prices only differ by small amounts.

4.3 Air Quality

Data for observed levels of ozone and particulate matter less than 10 microns diameter (PM_{10}) were obtained from the California Air Resources Board monitoring records. Southern California provides some of the most extensive air quality monitoring in the world. In the five counties of interest no fewer than 45 monitors were measuring ozone each year from 1987 to 1992 (after eliminating monitors active on less than 50 days), and, beginning in 1987, no fewer than 19 were measuring PM_{10}.

Table 2: Descriptive Statistics for Community Attributes for the 92 School Districts

Variable	Mean	Std. Deviation	Minimum	Maximum
Population size	50473	149169	4559	1433477
Total expenditures	4852	583	3936	6705
Instructional expenditures	2619	320	2066	3952
Student Teacher Ratio	24.77	1.52	17.32	28.48
Math score	2.05	0.35	1.46	3.33
Reading score	3.06	0.33	2.29	3.77
Writing score	3.35	0.50	2.40	4.12
Housing Prices	2.25	0.95	1.00	6.12
Ozone Level	0.15	0.04	0.09	0.24
Ozone Exceedances	41.1	34.2	1.00	105.0
PM_{10}	47.0	10.2	30.3	71.2

The Table contains descriptive statistics of the most interesting variables in our sample of 92 school districts in the LA metropolitan area.

Two issues arise in using these data in our models of community and housing choice. First, there is not a one-to-one correspondence between air quality monitors and school districts. Thus, an interpolation problem must be addressed in associating school districts with air quality levels based on spatially discrete measures. Fortunately, the large number of monitors minimizes the compromise created by this interpolation. In Riverside County, for example, half of all houses are within 4.3 miles of at least one monitor, and 90 percent are within 8.3 miles. Air pollution is measured using a centered three-year average (about the sales year) of pollution readings for the nearest monitor to each house. We investigate the effects of distance weighed pollution measures and found no significant difference in the conclusions derived with these measures compared to this temporal average of nearest monitor's readings. The community measure is the average over the homes that sold during the temporal window (1988 - 1992) for each community.

Three measures of air pollution – ozone concentration, ozone exceedances and particulate matter have been considered in evaluating the effects of air quality. Ozone is measured in parts per million (ppm) as the average of the top 30 one hour daily maximum readings at a given monitor during a year. We also consider the observed exceedances of the one hour federal standard for ozone. Particulate matter was measured by the annual geometric mean (in micro grams per cubic meter for particulate matter of 10 microns or less in size). Both pollutants have been well documented to impact health status and have been found to influence housing prices in hedonic studies (Bartik and Smith, 1987; Smith and Huang, 1995). In general the effects of particulate matter is through impacts on increased mortality rates and effects on materials. These impacts have not been shown to have any threshold, so annual mean for particulates is often used in epidemiological and economic analyzes of its effects. Table 2 reports the annual geometric mean of PM_{10} levels in our study area. In the case of ozone however, human health effects are more likely to be triggered at higher levels. The focus on maximum concentrations provides the rationale for considering the average of the 30 highest ozone readings or the number of days violating the federal standard, which is also expressed in terms of an order statistic.

4.4 School Districts and School Quality

Table 2 reports some descriptive statistics of the main variables characterizing 92 school districts in the sample. We find that school districts differ significantly in their size. The smallest districts only contain a few thousand households. The largest school district in the sample is LA unified with more than 1.4 million households, more than 30 percent of the total population in the LA metropolitan area.

The current school finance system in California was mainly shaped by two events: the 1971 decision of the state supreme court in *Serrano vs. Priest* and the approval of Proposition 13 in 1978 by voters in California. Before *Serrano vs. Priest* local school districts had fiscal autonomy. After this ruling the state imposed limits on spending and taxation of local districts and allocated aid in order to off-set inequalities in local spending. The basic idea was to achieve convergence of expenditures per student by increasing the aid to poorer districts and capping the amount of expenditure growth in the richer districts. In 1978 voters in California approved Proposition 13 which limited property taxes to one percent of assessed value and limited the growth of assessments. This rate was much lower than the average rate across school districts in the state at that time. As a result Proposition 13 limited the growth rates of expenditures over the last twenty years. Today most school districts in California have lower per capita expenditures per students than school districts elsewhere in the United States.

Despite the general trend towards equalizing expenditures per student, differences in expenditures per student arise for several reasons. The state equalization formula does not completely equalize expenditures. School districts with large tax bases can generate higher expenditures even under the existing set of rules determining transfers. In addition, school districts can obtain additional funds for special education programs. While these funds are aimed at covering the additional costs for teaching disadvantaged students, there is at least some evidence that these funds are also used to improve the overall quality of education (Cullen and Figlio (1999)). Finally, special fees and voluntary contributions have been important in supplementing local expenditures where school districts have been especially constrained by the reform. Brunner and Sonstelie (1996), for example, report that in 1992 nine of the twelve school districts that raised over \$500 in voluntary contributions per student experienced a decline in revenue limit funding. Table 2 reports some descriptive statistics of the instructional expenditures per student in our sample. In our sample, the mean educational expenditures are \$4852. While the range of educational expenditures is more than \$2500, most observations are within a few hundred dollars of each other. Instructional expenditures show even less variation. The sample mean is \$2619 with an estimated standard deviation of \$320. Both measures do not reflect the cash and non-cash voluntary contributions to school districts.

With small differences in educational expenditures, a better measure of school quality may be based on outcomes rather than state formula spending. There exist substantial differences in school quality among districts, measured by test scores. There are a number of comprehensive tests which have been designed and implemented by the California Department of Education. The primary purpose of these state wide testing programs is to monitor the performance of schools and provide some information to parents. We construct measures of performance from standardized test scores for each school district using the 1992-93 California Learning Assessment System Grade Level Performance Assessment test. Each student taking this exam is assessed at one of six performance levels (with six the highest level).

In Table 2 we also report average writing, reading and math scores of the school districts in our sample. All three test scores significantly differ among school districts. For example, math scores range from 1.46 to 3.33. Reading scores range from 2.29 to 3.77. These differences appear to persist across time. Thus, despite the equalization of public expenditures there remains considerable heterogeneity among school districts.

5 Estimation Results

5.1 Results from the Locational Equilibrium Estimator

The parameters of the locational equilibrium model are estimated using a two step procedure. In the first stage we match the distributions of housing expenditures of the 92 school districts observed in our sample with those predicted by our equilibrium model. In the second stage we match the observed indexes of public good provision with those predicted by our first stage estimates. Table 3 reports the parameter estimates and the estimated standard errors of the parameters which are identified in the first stage. The sample size is 222,865.

Table 3: First Stage Estimation Results

$\mu_{\ln(y)}$	10.52
	—–
$\sigma_{\ln(y)}$	0.34
	(0.01)
$\lambda_{\ln(y),\ln(\alpha)}$	-0.31
	(0.04)
$\rho/\sigma_{\ln(\alpha)}$	-0.26
	(0.03)
ν	0.86
	(0.02)
η	-0.17
	(0.06)
B	1.19
	(0.28)
function value	0.0028
degrees of freedom	270

NOTE: Estimated standard errors are given in parentheses.

In general we find that all parameters have the correct signs and are estimated reasonably accurately. In the estimation procedure, we treat the distribution of permanent income as latent. We only assume that it has the same mean as the distribution of current income which is equal to 10.52. Hence we can interpret the estimate of $\sigma_{\ln(y)}$ as an estimate of the standard deviation of the distribution of permanent income. The estimate of $\sigma_{\ln(y)}$ is 0.30 which is much lower than 0.75, the estimated standard deviation of current income. This finding suggests that the distribution of permanent income in the metropolitan area has a smaller variance than the distribution of current income, i.e. current income contains a significant transitory component. The point estimate for the correlation between income and tastes for public is negative and equal -0.31. This indicates that there is only a limited amount of stratification by income among communities in the sample.

We can also identify and estimate the income elasticity of housing, ν. The point estimate of ν is equal to 0.86 with an estimated standard error of 0.02. This estimate is quite consistent with what Polinsky (1977) reported in his early summary of the income elasticity estimates for consistent micro models allowing for a measure of permanent income. These estimates bracket our estimate of the income elasticity and range from .75 to .90 depending on the other variables included in what he describes as the "correctly specified metro equation". The estimated price elasticity of housing, η, is -0.16 with an estimated standard error of 0.06. It is not as close to the early estimates reported by Polinsky. Indeed this estimate is about one-fifth the average he reports.

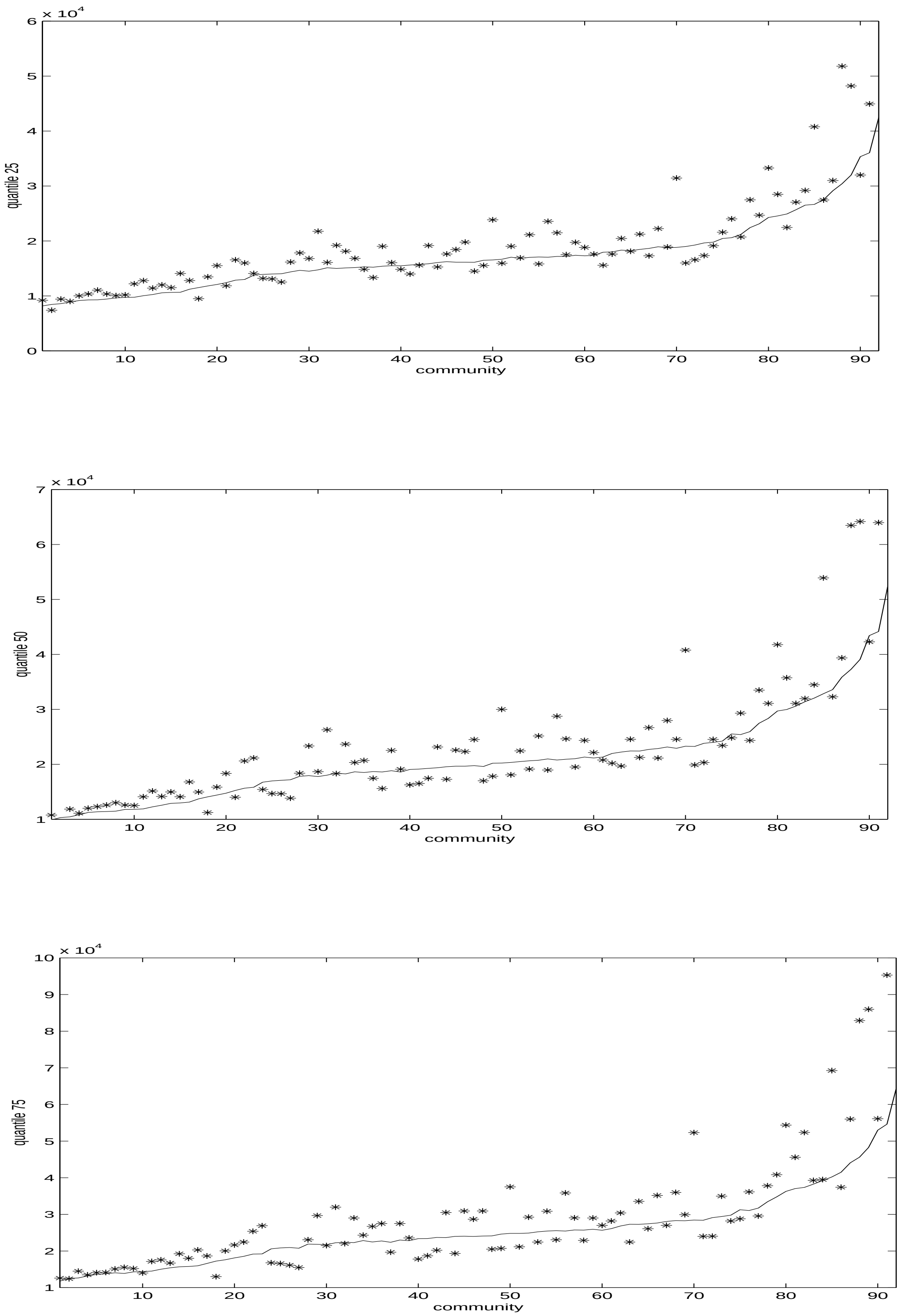
We therefore conclude that the estimated income elasticity of demand for housing corresponds to past estimates and incorporates recognition of the importance of distinguishing the permanent and transitory components of income, as noted in this earlier literature. Our price elasticity estimates are at the lowest end of the range of estimates (in absolute magnitude). However, these earlier studies have either used selected micro samples (i.e. the FHA data bases), aggregate measures without the ability to adjust for the heterogeneity in housing, or assumed site amenities should be bundled with the structural and lot characteristics. It may well be that after adjustment for the distinctions in site and structural attributes were made that the existing price elasticity estimates would be more closely aligned with our findings.

Figure 1 offers another way of evaluating our estimation strategy by plotting the empirical and the predicted quantiles of the housing expenditure distribution. Our model seems to fits the data reasonably well. Our strategy estimates 92 distributions of housing expenditures in the first stage with a model that has only 7 parameters. Given this tight specification of the model, the fit of the model is remarkably good. The correlation between the estimated and the predicted 25th (50, 75) quantile is 0.90 (0.86, 0.82) which is remarkably high.

The rank of the community in the hierarchical equilibrium is determined by its housing price. In our sample, we find that the ranking according to housing prices is quite similar as the ranking according to the different quantiles of the housing expenditure distributions. To quantify the similarity between these two rankings we do a pairwise comparison of communities and determine how often the rank predictions contradict each other. There are a total number of 4186 pairwise comparisons in a sample of 92 communities. Comparing the ranking according to prices and median housing expenditures, we find a total number 585 rank violations. Hence the rank predictions coincide in more than 86 percent of all cases.

In the second stage of the estimation procedure we can identify and estimate the remaining parameters of the indirect utility function and the underlying distribution of tastes for public goods. We estimate the parameters in the second stage using a GMM estimator which uses functions of the rank of the community as instruments. Table 4 reports the findings of this analysis. All estimated models use average math score in the community as a measure of school quality in a school districts, but use three different measures of air quality. The first approach uses average of the 30 highest ozone levels observed in a given year. The second uses the number of ozone hourly exceedances. Finally, the third measure is a geometric mean of PM_{10}. Column I reports the baseline model which ignores air quality. Columns II through IV report estimates obtained when we add one of the three measures to our index of public good provision. Finally, column V uses a specification which included

Figure 1: First Stage Fit



Empirical and predicted housing expenditure quantiles.

Table 4: Second Stage Estimation Results: GMM

	I	II	III	IV	V
ozone level	—	-1.59	—	—	-2.51
		(0.95)			(1.07)
ozone exceedances	—	—	-0.0020	—	—
			(0.0012)		
particulate matter	—	—	—	-0.0017	0.0057
				(0.0036)	(0.0035)
$\mu_{\ln(\alpha)}$	1.16	-0.39	-0.66	0.04	1.11
	(0.39)	(0.94)	(0.84)	(1.15)	(0.30)
$\sigma_{\ln(\alpha)}$	0.19	0.42	0.48	0.37	0.16
	(0.02)	(0.03)	(0.03)	(0.04)	(0.01)

Estimated standard errors are given in parentheses. The sample size is 92. The index is linear. The first component of the index is the average math score.

multiple measures air quality.

In general, the parameter of the air quality has the expected negative sign when one measure of air quality is included as in columns II through IV. Higher air quality increases the level of public good. If we include multiple measures of air pollution, we continue to estimate a negative sign for ozone, however the sign of PM_{10} reverses and is now positive. This finding suggests that here too, as in the case of hedonic models (Palmquist, 1991) it is hard to identify the separate effects of different air quality measures in our sample. Analyzing the estimated standard errors, we find that we estimate the effect of ozone with much greater precision than the effect of PM_{10}. The coefficient of ozone is typically significantly different from zero at reasonable levels of confidence while the opposite is true for the estimated coefficient for PM_{10}.

6 Welfare Analysis and Policy Implication

A primary motivation for this analysis was that it is possible to relax the assumptions that there exists a continuum of air quality choices or that air quality is necessarily a weak complement to some private good. Instead, we demonstrate that it is possible to recover sufficient information to identify the elements in an indirect utility function for heterogeneous households, provided we use the locational equilibrium, implied by the single crossing property and structure of the CES preference function to identify the parameters of the function. This strategy advances the revealed preferences literature by allowing consumers to accept the best of a set of discrete choices. That is, it acknowledges the constraints imposed on people's ultimate behavior by the public goods (community choices) available to them. In this section, we first compute welfare measures for intra-marginal changes in air quality based on our structural model which ignore general equilibrium effects, and compare them to the point estimates based on hedonic regressions which are the standard workhorse of applied welfare economics. We then show how to account for the general equilibrium effects in a behaviorally consistent way and implement the welfare measures based on our parameter estimates.

6.1 Benefits and Welfare Measurement: Partial Equilibrium Analysis

To develop a measure for the Marginal Willingness To Pay (MWTP) we invert the indirect utility function (with $\bar{u} = V$) as in (3.2) and compute the incremental change in E with a change in the quasi-fixed public goods designated as g in equation (6.1):

$$E \;=\; \left\{(1-\nu)\left[\frac{1}{\rho}\,\ln(\bar{u}^{\rho} - \alpha\, g^{\rho}) \;+\; \frac{B\;p^{\eta+1} - 1}{\eta + 1}\right] \;+\; 1\right\}^{\frac{1}{1-\nu}} \tag{6.1}$$

The marginal willingness to pay for a change in the level of public good provision is then given by the following equation:

$$P^{*} \;=\; -\,\frac{\partial E}{\partial g} \tag{6.2}$$

$$= \frac{E^{\nu}\ \alpha\ g^{\rho-1}}{\bar{u}^{\rho} - \alpha\ g^{\rho}}$$

The virtual price or marginal willingness to pay for a change in a component of g scales P^* by the estimated effect of ozone concentration and other public goods on g. In our example the effect of ozone on g is assumed to be constant and is given by γ in the case of the linear index given earlier in equation (3.11).

We simulate the locational equilibrium by drawing 500,000 realizations form the underlying joint distribution of income and tastes based on the parameter estimates reported in Table 3 and Column II of Table 4. We then compute the MWTP for each household for a ten percent reduction in average ozone levels over all 92 school districts as measured by the average of the 30 daily maximum readings. Finally we compute sample averages and standard deviations of the MWTP across the households in each community. These results are reported in the first column of Table 5. We find that the community averages of the MWTP (in 1990 dollars) range from \$372 in the poorest community to \$1509 in the richest community. Most communities have average MWTPs of around \$700. One of the advantages of our approach is that we can also study the dispersion of the MWTPs within each community. We find that standard deviation of the MWTP ranges from \$111 in the poorest community to \$506 in the richest community.

To provide some perspective on the importance of relaxing the assumption of a continuum of choice for air quality conditions we used the residential housing sales for these five counties, over the period 1988 to 1992, to estimate a hedonic price function. Fixed effects for the sale year and the county are included along with the structural characteristics and lot size measures that were part of the price indexes developed to apply the equilibrium location model. The three year, centered (at the sale year) average, for the average of the thirty maximum sendings in each year for ozone concentration at the closest monitor was used to capture the effects of air quality. A semi-log specification was used for the price equation.

This form was found to be most robust in Cropper, Dreck, and McConnell (1988)'s analysis of hedonic models as approximations for the equilibrium prices to match a distribution of buyers to a finite set of houses. Their criteria evaluated the model based on the properties of the estimates of marginal willingness to pay for attributes, and allowed for a variety of specification errors in the hedonic price equation. The semi-log offered the "best" specification when the price function was assumed to omit some structural or site amenities.

Table 5: Marginal Willingness To Pay Estimates

	Locational Equilibrium Estimator		Pooled Hedonic Estimator	
County	Mean	Range	Mean	Range
LA	1010	713 - 1509	683	214 - 4092
ORANGE	949	731 - 1250	643	399 - 1280
RIVERSIDE	678	372 - 727	314	215 - 431
SAN BERNARDINO	663	533 - 732	543	390 - 752
VENTURA	740	663 - 813	341	189 - 520

Table 6 reports the estimates for the model used for our marginal willingness to pay measures in the last column, labeled 6, along with separate estimates by county. The only surprising result is the positive but insignificant effect for the ozone measure for the sub-sample confined to Riverside County. This result contrasts with significant and negative effects for the other five sub-samples. Moreover, if we were to use an average of the exceedances of the federal standard instead of the average of the top readings, the results would be completely comparable across counties, with negative and significant effects in all five counties.

Our comparison uses the hedonic function for the full sample (column 6) and computes the average value of the marginal willingness to pay. These are scaled by the value for a ten percent reduction in the average ozone concentration and converted to 1990 dollars (using the fixed effect parameters of the model). The Poterba annualization was also used to present these approximations for the incremental willingness to pay in annual terms.

Table 6: Table Hedonic Property Value Models: 1988-1992

	1	2	3	4	5	6
sample	LA County	Orange	Riverside	Ventura	San Bernardino	LA Metro
intercept	11.717	11.581	10.805	11.995	12.174	11.170
	(1082.18)	(421.79)	(696.63)	(514.56)	(263.77)	(1316.61)
ozone average3	-1.450	-0.498	0.035	-2.062	-6.003	-1.208
	(-41.43)	(-2.98)	(0.64)	(-16.41)	(-29.78)	(-40.27)
baths	0.285	0.327	0.241	0.177	0.245	0.289
	(108.34)	(54.05)	(40.87)	(21.22)	(36.12)	(137.22)
bedrooms	0.003	0.002	0.095	0.026	0.078	0.025
	(1.40)	(0.47)	(20.83)	(6.09)	(20.12)	(16.86)
age	-0.001	-0.002	0.000	-0.006	-0.005	-0.002
	(-10.23)	(-9.94)	(-1.35)	(-21.40)	(-21.17)	(-20.65)
square foot lot	0.000	0.000	0.000	0.000	0.000	0.000
	(40.61)	(22.63)	(27.58)	(26.96)	(17.93)	(61.61)
fireplace	0.201	0.040	0.202	0.157	0.139	0.172
	(73.28)	(6.94)	(33.13)	(24.27)	(18.84)	(81.57)
pool	0.134	0.057	0.095	0.126	0.117	0.115
	(36.55)	(9.23)	(20.36)	(23.30)	(17.55)	(47.12)
sale88	-0.196	-0.133	-0.233	-0.158	-0.330	-0.194
	(-45.38)	(-17.68)	(-52.07)	(-25.57)	(-22.19)	(-67.41)
sale89	-0.020	0.032	-0.104	0.033	-0.104	-0.038
	(-4.40)	(3.58)	(-24.40)	(4.94)	(-15.35)	(-13.16)
sale91	0.022	0.017	-0.037	-0.055	-0.002	0.003
	(5.11)	(2.00)	(-6.77)	(-8.62)	(-0.39)	(1.09)
sale92	-0.020	-0.018	-0.066	-0.089	-0.050	-0.033
	(-4.89)	(-2.12)	(-14.12)	(-13.93)	(-8.35)	(-1.00)
county LA						0.519
						(164.84)
county Or						0.539
						(135.75)
county Rv						-0.078
						(-26.30)
county Va						0.394
						(96.67)
R^2	0.411	0.255	0.473	0.576	0.511	0.47

NOTE: t-statistics in parentheses.

The second column in Table 5 reports average values for these estimates by county. The averages are computed across housing sales in the five year period. The range of values for each county is across the school districts in each county. Because the computations are summarized for a school district /year we included both dimensions in developing the range of estimates. The numbers in the third column report the observations (school district/years) included in the range of averages summarized by county. Our estimates are on average smaller than those developed from the equilibrium location model, but they do display a wider range due to the scaling of the parameter for the effect of the ozone concentration by the highest priced homes in the area.[12]

6.2 Benefits and Welfare Measurement: General Equilibrium Analysis

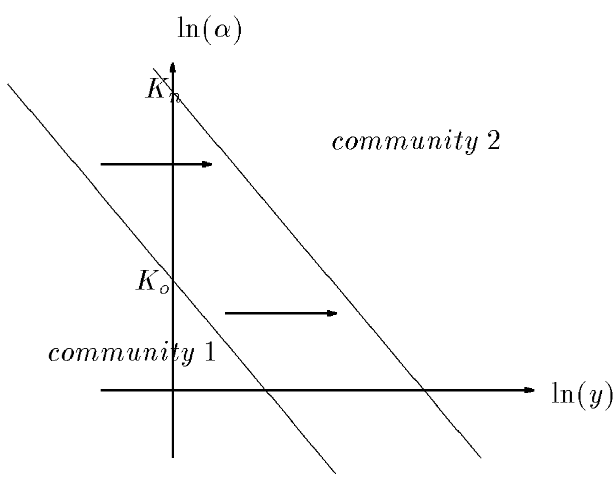
All welfare measures computed in the previous section are based on a partial equilibrium analysis. They ignore that an intra-marginal change in public good provision may have a large impact on the equilibrium allocation of agents across communities. An analysis which ignores these general equilibrium effects induced by the mobility of the households will ultimately lead to biased estimates. The main advantage of our empirical approach is that it allows us to compute the Willingness To Pay (WTP) for intra-marginal changes in the air pollution under a few additional assumptions on the housing stock. Hence it provides an internally consistent framework for estimation and applied general equilibrium welfare analysis.

To illustrate the problems encountered in general equilibrium welfare analysis, consider the following stylized example with two communities. We would like to know how an increase of the level of public good provision in community one affects the distribution of welfare in the economy. Note that an increase in g_1 will shift the boundary indifference locus to the right and hence affects all households in the economy, even those in community

[12]It is difficult to compare these results with the values reported by Smith and Huang (1995) because they considered only particulate matter. The results are substantially smaller than the Beron et al. (1999) estimates for a 20 percent improvement in visibility. It is not clear how the annualization of the change in the asset value was computed to develop their measures of incremental benefits in annual terms.

2 that do not experience a change in public good provision. To compute willingness to pay for an improvement in the level of public good provision (compensating variations), one must distinguish between three possible cases. First, a household used to live in community 1 in the old equilibrium and also lives in community 1 in the new equilibrium. Second, a household used to live in community 2 in the old equilibrium and lives in community 1 in the new equilibrium. Finally, a household lives in community 2 in both cases. This point is illustrated in Figure 2. Given the equilibrium allocations before and after the policy

Figure 2: Evaluating Public Programs



intervention, the willingness to pay (WTP) for a type 1 household is given by the following expression:

$$V(\alpha, g_1^o, p_1^o, y) \quad = \quad V(\alpha, g_1^n, p_1^n, y + WTP_1(y, \alpha)) \tag{6.3}$$

Similarly, one can compute $WTP_2(y, \alpha)$ and $WTP_3(y, \alpha)$ which in contrast to the partial equilibrium approach will not be equal to zero. In particular, households in community 2 who do not move in response to the policy, will have a positive willingness to pay if they are experience lower prices in the new equilibrium. The mean compensating variation is then given by:

$$\widehat{WTP} \quad = \quad \int_{-\infty}^{\infty} \int_{-\infty}^{K_o + \rho \ln(y)} WTP_1(y, \alpha)\, f(\ln(\alpha)\ \ln(y))\ d\ln(\alpha)\, d\ln(y)$$

$$
\begin{aligned}
&+ \int_{-\infty}^{\infty} \int_{K_o+\rho \ln(y)}^{K_n+\rho \ln(y)} WTP_2(y,\alpha)\, f(\ln(\alpha)\ \ln(y))\ d\ln(\alpha)\, d\ln(y) \\
&+ \int_{-\infty}^{\infty} \int_{K_n+\rho \ln(y)}^{\infty} WTP_3(y,\alpha)\, f(\ln(\alpha)\ \ln(y))\ d\ln(\alpha)\, d\ln(y) \qquad (6.4)
\end{aligned}
$$

Finally, one must compute the mean changes in producer surplus in the housing markets. Adding these measures provides an approximation of the benefits of the policy intervention.

In order to implement the procedure above, we close the model and introduce a housing supply function for each community to our model. The simplest way to do this is to assume that the housing stock is constant and equals to the housing demand predicted by our model. We also used specifications in which there is a positive housing supply elasticity and found that the results of the general equilibrium effects are even larger than in the benchmark case above.

Given an assumption on housing supply, we can compute the new locational equilibrium under the modified policy. We consider the same policy change as discussed in the previous section to distinguish between partial and general equilibrium measures. We find that the change in public good provision has significant effects on prices in all communities. In particular, the low amenity communities are relatively more attractive. Households move to these communities. Higher amenity communities experience a decrease in population. We find that approximately 1.6 percent of the population moves to a different community in response to the changed allocation of public goods.

As a consequence of this, housing prices adjust in all communities. Low amenity communities experience price increases of up to two percent while the the high amenity communities see price decreases of up to one percent. These price changes have quite significant effects on the estimated distribution of the WTP in the economy. Households in the high amenity communities not only benefit from the increase in public good provision in their preferred community, but also from the lower housing prices. Hence their WTP increases relatively to the WTP predicted by the partial equilibrium measure. The mean WTP in the richest community is \$1867 which is approximately 20 percent higher than the WTP based on the partial equilibrium measure. In contrast, households living in the low amenity communities face higher housing prices in the new equilibrium than in the old equilibrium. Therefore, their willingness to pay decreases. We find that the community averages of the WTP is $186 in the poorest community which is almost 50 percent lower than the one which ignores the price adjustments. We conclude that that adjusting for the general equilibrium effects is important.

7 Conclusion

Our results confirm the feasibility of incorporating the conditions for an equilibrium sorting of heterogeneous households, into the estimation of their preferences for air quality and other local public goods. Our application to Southern California provides plausible estimates of both the housing demand and the incremental benefits from air quality improvements. The marginal willingness to pay estimates are comparable to those we derived using hedonic models. Finally, we find that the estimates of the willingness to pay for intra-marginal changes should account for the general equilibrium effects which are induced by the change in public goods. We find that controlling for the general equilibrium effects increases the dispersion of the distribution of the willingness to pay estimates.

The important contribution of the method presented in this paper lies in its ability to describe behavior *without* a continuum of choice for environmental quality and to measure the benefits arising from improvements in environmental quality even though households may choose not to adjust their locations. As such it is responsive to Bockstael and McConnell's[1999] arguments that revealed preference methods must acknowledge the potential for behavioral silence. Equally important, it offers the potential to quantify the early bounding conditions used to describe "large" environmental changes in a given region. Originally introduced by Lind (1973) and subsequently extended by Freeman (1974a, 1974b) and Starret (1981), these arguments suggested that aggregates of changes in land prices would be able to approximate the incremental benefits of a policy change, provided the size of the change was not "too" large. By developing a model that recovers preferences from the equilibrium conditions for household sorting among a finite set of alternatives, the our method provides the basis for computing a new locational equilibrium.

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