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ABSTRACT

This paper reports the first comprehensive approach for measuring the general equilibrium willingness to pay for large changes in air quality. It is based on a well defined locational equilibrium model. The approach allows estimation of households' indirect utility function and the underlying distribution of household types. With these estimates it is possible to compute a new locational equilibrium and the resulting housing prices in response to exogenous changes in air quality. This permits construction of 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 scope of more traditional approaches. The empirical approach of this paper provides, for the first time, an internally consistent framework for estimation and applied general equilibrium welfare analysis. We compute the general equilibrium willingness to pay for the changes in air quality between 1990 and 1995. We implement our empirical framework using data from Southern California, an area which has experienced dramatic 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 can be especially important.

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1 Introduction

Economic regulation of the environment arguably began with the 1990 Clean Air Act Amendments. Among their provisions, those establishing the tradeable permit program for sulfur dioxide emissions have received the most attention by economists (Joskow, Schmalensee, and Bailey, 1998). However, Section 812's requirement that EPA evaluate the benefits and costs of the Act's implementation may, over time, be recognized as equally important. This mandate is likely to foster a more direct role for economic analysis in both the evaluation and the design of environmental policy. The results of this paper suggest that it will be essential to incorporate general equilibrium effects into benefit measures of large policy changes.

To date, the literature has been unbalanced in its treatment of general equilibrium effects. Hazilla and Kopp (1990) provide the first estimates of the importance of general equilibrium effects for the costs of environmental regulations, but do not address benefits.[1] The costs of complying with environmental regulations are estimated to be three to four percent of U.S. gross domestic product (GDP). EPA's Retrospective Analysis (1997) of the benefits of the air quality legislation from 1970 to 1990 reports a staggering 42 trillion (1990) dollars in benefits as their mid-range estimate. This is about five and a half times the current estimates for 1999 real GDP. However, the partial equilibrium analysis used by

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[1] Smith and Espinosa (1996) do consider the environmental benefits of air quality improvement using a calibrated CGE model for countries in the European Union and its major trading partners. While their approach allowed air pollution to have a non-separable effect on preferences, their calibration procedure assumed that there was a private good serving as a perfect substitute for the roles assumed for air quality. Our approach estimates the role of air pollution in household choice and relaxes the perfect substitution assumption.

EPA to determine the benefits of large air quality changes realized over the twenty year period can hardly be considered appropriate.[2]

Our research adapts the Epple and Sieg (1999) locational equilibrium estimator to measure household preferences for air quality. This approach offers a behaviorally consistent framework to estimate the effects of spatially delineated non-market goods for household preferences. Our benefit estimates control for households' adjustments to large policy changes, treating them as they should be, namely as composites of air quality and prices changes. We apply the locational equilibrium approach to study improvements in air quality between 1990 and 1995 in Southern California. We observe ozone reductions ranging between 4 and 36 percent across the different locations in the five counties in our sample. We find that general equilibrium estimates of Hicksian willingness to pay for these ozone reductions are between 2 and 4 percent of annual household income. In contrast, the estimates of annual benefits in the EPA's Retrospective Analysis for a single year (1990) were about $10,000 for each household - over 25 percent of average income.

In addition to demonstrating the importance of general equilibrium effects, this research extends past efforts to measure consumer preferences for environmental amenities in two ways. First, most revealed preference approaches for measuring consumer values for amenities rely on models that either assume preferences satisfy both weak complementarity (Mäler, 1974) and the Willig (1978) condition or require a spatial distribution of ameni-

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[2] In addition, the peer reviewers for the EPA Retrospective Analysis recommend to re-evaluate the measurement of individual's willingness to pay for changes in the risk of premature mortality. The majority of the benefits in the EPA analysis stem from changes in mortality risks for older adults. The committee argues that the values estimated for statistical lives may be too large for this group. See Schmalensee and Cropper (1997).

ties that offers a continuum of choices.[3] The locational equilibrium framework does not require either. In fact, the single crossing property (used to assure an equilibrium sorting of households) can be seen as a relaxation of the Willig condition. Second, even in situations where a combination of theory and data allow estimation of marginal willingness to pay functions from reduced form hedonic models, these functions are partial equilibrium descriptions of behavior. They do not incorporate the adjustments that reflect general equilibrium responses to large scale environmental changes. Moreover, there are complex self selection and endogeneity issues that arise in these equilibrium matchings that are difficult to incorporate in reduced form models.

The remainder of the paper is organized into seven sections. In the next section we describe the highlights of our findings. Section three contrasts our approach with the past literature measuring the benefits of air quality improvements. Section four reviews the locational equilibrium estimator and describes how it can be adapted to use data on housing expenditures to improve the resolution in estimating household preferences. Section five describes the unique data set on housing prices and characteristics, air quality, and public education that are available for Southern California. Section six summarizes our estimation results for the model. Section seven outlines the general equilibrium benefit measurement. Finally, Section eight discusses conclusions and implications.

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[3] Hedonic methods rely on the assumption of a continuum of air quality alternatives. There has been no systematic analysis of the implications of departures from this assumption. This issue is especially important to environmental amenities, where it is likely that within the market area assumed for a hedonic there would be a discrete number of alternative values for the available amenity services.

2 Overview of Results

Figures 1 and 2 summarize the features and implications of the general equilibrium benefit estimates computed using our estimates of household preferences for housing, air quality, and public education. Figure 1 plots our estimates of the general equilibrium willingness to pay (WTP) for the improvement in ozone that occurred in Southern California between 1990 and 1995. Each point is an average, computed for the households estimated to be members of the ninety-two school districts we use as the communities in the locational equilibrium model. For each community, average WTP is plotted against the estimated average income in that community. Figure 1 displays the diversity in gains from the air quality improvements between 1990 and 1995. Average gains range from approximately \$200 to over \$2000.

INCLUDE FIGURE 1

The differences in WTP by income should not be taken as a gauge of the responsiveness of the Hurwicz and Uzawa (1971) income compensation function to income. Because these differences are means for general equilibrium solutions, they are distinguished by five characteristics. Households in each community are assumed to have different (unobserved) tastes for public goods and different incomes. They also experience different changes in ozone concentrations, housing prices, and public education. Our averages of general equilibrium WTP for each school district reflect the effects of all of these differences on the average attributed to each community.

To understand how this happens, consider a stylized description of the general equilibrium solution and WTP computation. We alter community attributes from the initial 1990 ozone conditions to the 1995 ozone readings for each of the 92 communities. At first, households are not in equilibrium; some of them have opportunities to increase their well-being by moving. The same conditions assuring an initial equilibrium sorting allow us to compute the new distribution of households among communities and resulting housing prices. This process implies households that change their community will experience changes in three variables affecting our measures of their general equilibrium WTP – the ozone concentration, the price they face for a homogeneous unit of housing, and their quality of public education. For those households that do not change community, their WTP reflects only the changes in ozone and the new general equilibrium housing prices for that community.

INCLUDE FIGURE 2

Figure 2 depicts the extent of the ozone changes and the proportionate changes in prices to sustain the new equilibrium sorting. We find that there is a strong positive correlation between ozone changes and predicted price changes. Our model thus predicts that some of the households that used to live in the high amenity communities will relocate to take advantage of the relative larger air quality improvements and more attractive housing prices (in relative terms) in the less affluent communities. The communities with the largest improvements in air quality experience increases in housing prices of up to 7.8 percent. In contrast, housing prices in communities with small improvements of air quality decrease in equilibrium up to 6.5 percent.

3 The Previous Literature

Following 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 rely on hedonic models and use measures of particulate matter to characterize air quality conditions.[4] Several methodological lessons emerge from this literature. This section summarizes a few of the insights from the first round of hedonic research in this area and then outlines the findings and potential limitations of the newest studies.

Early work establishes that it is reasonable to expect that home-buyers take account of air pollution as a site specific amenity and that the significant negative relationship between housing values and air pollution is not a statistical "artifact". The conditions for identifying the function describing the marginal rate of substitution between air quality and a numeraire are 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 determine households' locational choices; and the supply responses of housing producers to households' demands for dwelling and site characteristics. Data limitations also have important effects on the results from these early hedonic models. Nonetheless, Smith and Huang (1995) conclude that the estimated marginal values for reductions in particulate matter do consistently respond to

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[4] For a review and a meta-analysis of these early studies see Smith and Huang (1995).

income, increases in air pollution (as measured by Total Suspended Particulates), and local housing market conditions. They also find that use of estimates of these marginal values without adjustment for how they are developed leads to dramatic differences in the benefits attributed to air quality improvement.

Recent studies extend the literature in a number of ways ways. Chattopadhyay (1999) focuses on identification of the second stage marginal willingness to pay function (MWTP) based on maintained preference restrictions. He uses a selected sample for Chicago acquired through FHA mortgage applications and assumes a single, specific preference function to identify the MWTP function using housing sales for this one area. Beron, Murdoch, and Thayer (1999) is another study which attempts to implement the second stage of the hedonic analysis. They achieve identification through variation in marginal price functions for air quality improvements over time, assuming that the market is changing during the period spanned by their sample of housing sales. However, the use of different hedonic price functions alone may not be sufficient to identify the second stage. Palmquist (1991) notes that the multiple markets must differ significantly to identify the marginal rate of substitution equations. Beron et al. (1999) also face the problem that variation in important socio-economic characteristics cannot be measured for years other than 1980 and 1990.

Chay and Greenstone (1998) attempt to relax exogeneity assumptions in their analysis and propose the use of the attainment status of each county as an instrument for air pollution improvements. They argue that if this instrument can be considered an exogenous gauge of air quality improvement then it avoids problems associated with departures from random assignment of pollution readings across units of observations used in a hedonic regression.

However, the attainment/non-attainment status of a county as an instrument to capture regulation induced changes in particulate matter may not be as unambiguous as Chay and Greenstone (1998) imply.[5]

Overall these studies identify new methods and data for overcoming some of the limitations of the early research attempting to measure the willingness to pay for improvements in air quality. They also document a set of shortcomings with reduced form frameworks for measuring consumer preferences. By design these reduced form approaches tend to be exceptionally demanding in the data required to recover sufficient information to evaluate consumer preferences for non-market amenities from observed price responses alone. As such they motivate this effort to consider an alternative type of revealed preference method based on a well-defined locational equilibrium model.

Our approach differs from the previous literature by supplementing the available detailed micro data on actual housing prices with aggregate information on housing markets and income as well as for school and environmental quality. The locational equilibrium model used in this paper imposes restrictions on the observable distributions of households given the diversity in local public goods and housing prices. This permits us to estimate the underlying structural parameters of the model and to construct welfare measures that take into account the equilibrating adjustments of households to non-marginal changes in air quality.

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[5] Crandall (1983) and Melnick (1983) describe a regulatory system over the first decade used in their analysis, 1970-1980, that would not be a reliable indicator of air pollution. As a result both of the authors' evaluations suggest that the monitoring networks for air quality and for point source compliance were inadequately developed to allow attainment to be consistently linked to air quality. For the latter period, 1980-1990, the argument may be more plausible because the air quality monitoring and compliance networks were established.

4 Locational Decisions and the Valuation of Environmental Amenities

4.1 A Locational Equilibrium Model

The starting point for our work 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, and 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 preference-income 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, $f(\alpha, y)$.

Each household has preferences defined over a composite of local public goods. We assume the local public good has two components. The first component corresponds to local public education which is provided by each community. The second is our exogenous environmental amenity. It is not affected by local community decisions and is determined by the mix of point and mobile sources of air pollution emissions that are outside the choice process described by the model. Each community has 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

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[6] This approach follows a long tradition in urban economics and local public finance beginning with Tiebout (1956). 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, Romer, and Sieg (2000). Berry, Levinsohn, and Pakes (1995) provide an alternative approach for estimating differentiated product models. For a recent application of that approach to study school choice in California see Bayer (1999).

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 properties. Households maximize utility with respect to a budget constraint:

$$\begin{aligned} &\max_{(h_j, b_j)} \; U(\alpha, h_j, b_j, g_j) \\ &s.t. \;\; p_j \, h_j \;=\; y \;-\; b_j \end{aligned} \tag{4.1}$$

Household preferences are represented in our structural model using the indirect utility function given in (4.2). In principle, this can be derived by solving the optimization problem given in (4.1).

$$V(\alpha, y, g_j, p_j) \;=\; \left\{ \alpha \, g_j^{\rho} \;+\; \left[e^{\frac{y^{1-\nu}-1}{1-\nu}} \; e^{-\frac{B p_j^{\eta+1}-1}{1+\eta}} \right]^{\rho} \right\}^{\frac{1}{\rho}} \tag{4.2}$$

The necessary conditions for an equilibrium sorting of households among communities can easily be characterized when individual preferences satisfy the "single crossing" property. This condition requires that the slope of an "indirect indifference curve" in the (g, p) plane

$$M(\alpha, y, g_j, p_j) \;=\; -\frac{V_g}{V_p} \tag{4.3}$$

be monotonic in y, given α (and α, given y).

This restriction contrasts with the dual requirements of weak complementarity and the Willig [1978] condition for some revealed preference approaches to non-market valuation.[7] Weak complementarity requires that there exist a non-essential private good, linked to the

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[7] The notable exceptions would be hedonic models and discrete choice random utility models.

environmental amenity, so that the value of improvements in that amenity is zero if none of the private good is consumed. For example, it assumes that improvements in water quality at a lake are only of value to those who use the lake for recreation. In terms of the indirect utility function, this condition implies $V_g = 0$ when the weak complement's price is at the choke price. This condition restricts the same slope used to define the single crossing property to be independent of income. The Willig condition is required to develop Hicksian welfare measures for changes in amenities from market demands (Bockstael and McConnell, 1993). By characterizing households with indirect utility functions, the locational equilibrium framework not only allows the necessary conditions for an equilibrium to be used in recovering preference estimates, it also relaxes the requirements for consistent revealed preference estimates.[8]

The specification of the household decision process and our indirect utility function in (4.2) implies that households consider public goods and housing prices at the extensive margin as a choice among a finite set of alternative communities. Then conditional on this decision, they select the optimal amount of housing which is independent of g_j. For a given location, the specification of preferences over g_j parallels what would be found in definitions for public goods giving rise to non-use value (Hanemann, 1988; Randall, 1991). In our case the necessary conditions for an equilibrium sorting of households, heterogeneous in their preferences for g_j, is what allows us to identify this separable component. Similar strategies

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[8] Note that

$$\frac{\partial M}{\partial y} \quad = \quad \frac{V_y}{V_p^2}[h^* \, V_{gy} \; + \; \pi_g \, V_{py}] \tag{4.4}$$

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

may have promise in measuring this rather specialized conception of the non-use values for other public goods.

Our model is also somewhat similar to a hedonic property value framework in that it is important to control for the heterogeneity in housing characteristics when measuring the separate community specific housing prices, the $p_j's$. To do so, we draw a distinction between housing characteristics and public goods conveyed by location. We retain the hedonic model's assumption of a continuum of choice in the structural characteristics and lot size. Thus, within each community the "law of one price" is realized through a hedonic price function controlling for differences in housing characteristics. Public goods are available in a finite set of choice alternatives.[9]

Preference and income heterogeneity implies different individuals will prefer the public good provision and gross-of-tax housing price in different communities. If preferences satisfy the single crossing properties, a locational equilibrium satisfies three properties: boundary indifference, stratification, and ascending bundles (Epple and Platt, 1998). The boundary indifference conditions (a set of (α, y) such that $V(\alpha, y, g_j, p_j) = V(\alpha, y, g_{j+1}, p_{j+1})$) define household membership in each community. With preferences given by (4.2) the set of agents in community C_j is defined as:

$$C_j \;=\; \left\{(\alpha, y) \;\middle|\; K_{j-1} \;\leq\; \ln(\alpha) \;-\; \rho\,\left(\frac{y^{1-\nu}-1}{1-\nu}\right) \;\leq\; K_j\right\} \tag{4.5}$$

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[9] As a result it is possible, as Bockstael and McConnell (1999) suggested recently, to have changes in environmental quality that might not lead to a behavioral response.

where the intercept K_j is given by:

$$K_j \quad = \quad \ln(e^{-\rho \frac{B p_{j+1}^{\eta+1} - 1}{1+\eta}} - e^{-\rho \frac{B p_{j}^{\eta+1} - 1}{1+\eta}}) \; - \; \ln(g_j^\rho - g_{j+1}^\rho) \tag{4.6}$$

The single crossing properties imply that, holding tastes constant, households will stratify by income among the set of communities. Likewise holding income constant, those households with greater taste for public goods will tend to live in the high public good communities. The ascending bundles condition implies rankings of communities by housing prices and public good levels will be the same.

Integrating $f(\alpha, y)$ over the set C_j defined by equation (4.5) provides a measure of the households in each community, denoted $P(C_j)$. Quantities for housing expenditures can be derived in a similar fashion. More specifically, Roy's identity implies that the logarithm of housing expenditures is given by the following equation:

$$\ln(p_j h_j) \quad = \quad \ln(B) \; + \; (\eta + 1) \; \ln(p_j) \; + \; \nu \; \ln(y) \tag{4.7}$$

Furthermore it is easy to show that marginal distribution for $\ln(y)$ in community j is given by

$$f_j(\ln(y)) \quad = \quad \int_{-\infty}^{\infty} 1\{(y, \alpha) \in C_j\} \quad f(\ln(y), \ln(\alpha)) \;\; d\ln(\alpha) \; \Big/ \; P(C_j) \tag{4.8}$$

where $1\{(y, \alpha) \in C_j\}$ is an indicator function associated with community j. Equations (4.7) and (4.8) completely describe the marginal distribution for $\ln(p_j h_j)$ in each community. In the estimation procedure, we match quantiles of housing expenditure distributions. The r-th quantile of the logarithm of housing expenditures in community is given by $\ln(B) + (\eta+1)\ \ln(p_j) + \nu\ q_{\ln(y)}(r)$ where $q_{\ln(y)}(r)$ is the r-th quantile of the log-income distribution in community j.

The estimation procedure outlined in the next section exploits only the necessary conditions for residential choices in equilibrium to define identifying restrictions. However, we need to compute the locational equilibrium of the model to compute welfare estimates for non-marginal changes in the ozone concentrations. To close the model, we assume that housing stock is owned by absentee landlords and that the housing supply is given:

$$H_j \quad = \quad l_j\, p_j^{\tau} \tag{4.9}$$

where l_j is a community specific constant, which reflects the differences in land endowments and other fixed factors, and τ is the constant supply elasticity. In the CGE analysis, we calibrate the housing supply function such that the predicted initial housing demand equals the housing supply for our estimated model. To solve for the new equilibrium prices after an (exogenous) change in the level of public good provision, we solve a system of nonlinear equations given by the J local housing market clearing conditions. Given the hierarchical structure of the model, we can start with an initial guess for the new housing price in the first community. We then compute the housing prices for all other communities such that the first $J - 1$ housing markets clear. We keep adjusting the housing price of the first community until the last housing market also clears.

4.2 Estimation

Estimation is implemented 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 vector. The boundary indifference equation implies that quantiles of the housing expenditure distribution of community j depend on g_j only through the community specific intercepts K_j. We treat the K_j's as unknown parameters and constrain them to replicate the characteristics observed for each community's population. A subset of the parameters of the 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]

Matching the predicted housing distributions with those derived from panel data of housing transactions in each community simplifies identification and estimation of the pa-

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[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.

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

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. Solving equation (4.6), which characterizes the community specific intercept, for the levels of public good provision yields the following equation:

$$g_j \;=\; \left\{ g_1^{\rho} \;-\; \sum_{i=2}^{j} \left(Q_i - Q_{i-1}\right)\ \exp(-K_{i-1}) \right\}^{1/\rho} \tag{4.10}$$

where $Q_i = e^{-\,\rho\,\frac{B p_i^{\eta+1} - 1}{1+\eta}}$. Note that all variables on the right hand side of equation (4.10) are either observed or have been estimated in the previous step of the estimation procedure. We then replace the (unobserved) level of public good provision, g_j, with a linear index $x_j'\gamma + \omega_j$, where ω_j reflects unobserved characteristics of each community, x_j is the vector of local public goods and amenities. γ is a parameter vector to be estimated. This yields a well defined nonlinear regression model. We construct an instrumental variable estimator for the remaining structural parameters of the model as outlined in detail in Epple and Sieg (1999). The second step completes the estimation procedures and allows us to recover almost all structural parameters of the underlying model.

4.3 General Equilibrium Welfare Measurement

The conventional partial equilibrium Hicksian willingness to pay, WTP_{PE}, for a change in air quality conditions is defined as:

$$V(\alpha, y - WTP_{PE}, \bar{g}_j, p_j) \quad = \quad V(\alpha, y, g_j, p_j) \qquad (4.11)$$

It is the reduction to income required to equalize utility realized with the local public goods index adjusted to reflect the improved air quality, $\bar{g}_j$, with that of the original conditions, g_j. However, households will adjust their community locations in response to the changes in air quality. The new equilibrium should involve new housing prices as well. Thus, the evaluation of the policy change should reflect the general equilibrium market adjustment induced by the improvement in non-market environmental amenities. This general equilibrium willingness to pay, WTP_{GE}, is given in (4.12)

$$V(\alpha, y - WTP_{GE}, \bar{g}_k, \bar{p}_k) \quad = \quad V(\alpha, y, g_j, p_j) \qquad (4.12)$$

Notice that both g and p are assumed to change and that the definition allows for households to adjust their location, e.g., the subscripts for $(\bar{g}_k, \bar{p}_k)$ need not match (g_j, p_j). Thus, the general equilibrium willingness to pay requires a solution for the new general equilibrium distribution of households and associated price vector. With these new housing prices and amenities households remaining in community j can be expected to experience new conditions because the level of amenities and prices in j change from (g_j, p_j) to $(\bar{g}_j, \bar{p}_j)$. Households moving from j to a new community k would have their amenities and price change from (g_j, p_j) to $(\bar{g}_k, \bar{p}_k)$.

One final aspect of the welfare measurement concerns the distinction between owners and renters. Our definitions of WTP_{GE} to this point assume all households are renters and ignore the differences in rental payments, ΔR_j:

$$\Delta R_j \quad = \quad (\bar{p}_j - p_j)\, H_j \tag{4.13}$$

The difference offers an approximate gauge of the importance of the owner/renter distinction.

In summary, the framework allows a measure of mean willingness to pay after taking account of the re-location and price adjustment associated with a large exogenous change in local public goods. Because these computations are developed by numerical simulation, it is also possible to consider how this mean gain is distributed across communities and income groups by computing the conditional expectations based on the initial conditions for households in the region.

5 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.

5.1 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 (in acres)	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.

One potential drawback associated with using California data relates to Proposition 13, which in 1978 limited property taxes to one percent of assessed value and limited growth in assessments. It has been argued in the literature that Proposition 13 created a lock-in effect on homeowners. A household faces a tax on mobility because property taxes are based on the market value at the time of the last sale. If the current market value exceeds the assessed value, the revaluation creates additional mobility costs. O'Sullivan, Sexton, and Sheffrin (1995)'s detailed quantitative analysis of these lock-in effects indicated that they are small – "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 are questions relating to turn-over in housing markets and the representativeness of our measures for the distribution of housing expenditures. It is important to gauge whether our sample is representative of the underlying housing stock of each of the school districts. The pattern of housing expenditures across communities is quite close between the US Census and our estimates, 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). 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 over-sampling of new homes (relative to what would be desired for measures of the overall housing stock) is not surprising, because newly built homes will automatically show up in housing transactions, whereas older homes will only show up when they turn over. According to the 1990 US Census, 70 to 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 houses.

5.2 Housing Prices

An important aspect of the empirical analysis of a model of households' locational and housing choices involves constructing inter-community housing price indices that control for the observed differences in the quantity and quality of housing consumed within and across communities. We assume that 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 using an assumed continuum of choices for the structural characteristics of housing to develop a price index for a homogeneous unit of housing in each community. The market value of a specific house located in a given community can be converted into the imputed rent using the approach outlined in Poterba (1992). We also observe a vector of housing specific characteristics denoted by z_{jn}. Let u_{jn} denote the unobserved housing characteristics. We assume that the quality adjusted units of housing is given by:

$$h_{jn} \quad = \quad e^{\delta' \, z_{jn} \, + \, u_{jn}} \tag{5.1}$$

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. Using our specification for h_{jn} in equation (5.1) in this definition and taking logarithms provides a well-defined regression model, that we use to construct housing price indices 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) discuss the specific details of these regressions. This comparison of alternative approaches suggests that housing price estimates are robust across different econometric specifications and index formulas. 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 index 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.

5.3 Air Quality

Data for observed concentrations 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}.

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 compromises created by this interpolation. In Riverside County,

Table 2: Descriptive Statistics for the 92 School Districts

Variable	Mean	Std. Deviation	Minimum	Maximum
Population size	50473	149169	4559	1433477
Total expenditures per student	4852	583	3936	6705
Instructional expenditures per student	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
Index of Housing Prices	2.25	0.95	1.00	6.12
Ozone Level[a]	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

These descriptive statistics are the averages summarized for the 92 school districts in our sample.

[a] average of the top 30 one hour daily maximum readings.

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. We assign a three year centered average to each house in the sample that sold during the temporal window (1988 - 1992). The community measure is then computed by averaging over the houses in the community.

Three measures of air pollution – ozone concentration, ozone exceedances and particulate matter are 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 (Smith and Huang [1995], Beron et al.[1999]). 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.

5.4 School Districts and School Quality

Table 2 reports some descriptive statistics for 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 addition, Proposition 13 limited the growth rates of expenditures over the last twenty years. As a result of these events, 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. First, 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 cash and non-cash voluntary contributions to school districts.

With small differences in educational expenditures, a better measure of school quality would 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 report average writing, reading and math scores of the school districts in our sample. All three test scores significantly differ among school districts. Math scores range from 1.46 to 3.33. Reading scores range from 2.29 to 3.77.

6 Estimation Results

6.1 Results from the Locational Equilibrium Estimator

The parameters of the locational equilibrium model are estimated using a two step procedure. The first matches the distributions of housing expenditures of the 92 school districts observed in our sample with those predicted by our equilibrium model. The second stage matches the observed indices 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

parameter	$\mu_{\ln(y)}$	$\sigma_{\ln(y)}$	$\lambda_{\ln(y),\ln(\alpha)}$	$\rho/\sigma_{\ln(\alpha)}$	ν	η	B
estimate	10.52	0.34	-0.31	-0.26	0.86	-0.17	1.19
std. error	—–	(0.01)	(0.04)	(0.03)	(0.02)	(0.06)	(0.28)

Function value: 0.0028. Degrees of freedom: 270.

In general all parameters have the expected signs and are estimated reasonably accurately. Permanent income is treated as a latent variable with 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 0.34 which is much lower than 0.75, the estimated standard deviation of current income from the census. 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 local public goods is negative and equal -0.31, which suggests a limited amount of stratification by income among communities in the sample.

We can also identify and estimate the income elasticity of demand for housing, ν. The point estimate of ν is equal to 0.86 with an estimated standard error of 0.02. This estimate is consistent with Polinsky (1977)'s early summary of the income elasticity estimates for consistent micro models. These estimates 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. This is not as close to the early estimates reported by Polinsky. While this estimate is about one-fifth the average he reports, it can be expected to be sensitive to the procedures used to adjust prices (or quantities) to measure the demand for a homogeneous bundle of housing services. The estimates Polinsky selected as best addressed this issue by using the Bureau of Labor Statistics' index of the annual cost of a standardized package of housing. This index does not have the same resolution as our individualized hedonic price index. Moreover, he reports estimates (by Carliner (1973) and others) on either side of his central measure with some as low as -.10.

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 estimate is at the lower 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.

INCLUDE FIGURE 3

Figure 3 offers another way of evaluating our estimation strategy by plotting the empirical and the predicted quantiles of the housing expenditure distribution. Our model fits the data reasonably well. We match 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).

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

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 and has a parameter normalized to one.

which uses functions of the rank of the community as instruments. Table 4 reports the findings of this analysis. All estimated models use the average math score in the community to measure school quality. Three different measures of air quality are considered. 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 model considers 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 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 local public good. If we include multiple measures of air pollution, we continue to estimate a negative sign for ozone, but 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 the effect of ozone is estimated 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. As a result of these findings and the dramatic improvements in ozone concentration in the area during the nineties we focus our policy scenarios on the model in column II.

7 The Benefits for Ozone Changes – 1990 to 1995

7.1 Context for the Policy Analysis

Between 1990 and 1995, Southern California experienced significant air quality improvements. As we noted at the outset (and plotted in Figure 2), the decline in ozone concentrations ranged from 3 to 36 percent across the school districts in our sample. In Los Angeles County the number of days exceeding the federal one hour ozone standard dropped by 27 percent from 120 to 88 days. Such large changes have been frequent topics of local newspaper articles and home buying guides and are widely recognized by home buyers.[11] These changes were similar in magnitude to nationwide changes in ozone that were attributed to

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[11] For example, the LA Times' Sunday section (September 27, 1998) included an article "Nothing to Sneeze At" describing where to "buy" air free of smog.

the air quality regulations in EPA's Retrospective Analysis.[12] They estimated that ozone concentrations in 1990 were about 90 percent of what they would have been without the regulations. Thus, a comparison of our estimates of the general equilibrium willingness to pay for improvements in Southern California (as one time increments) in comparison to partial equilibrium measures for the same change offers a convenient basis for gauging how large the errors in partial equilibrium welfare estimates can be.

Our description of these results is divided into four subsections. First we discuss the general equilibrium WTP estimates by county, highlighting differences in the gains across communities as identified in Figure 1. In the next subsection we describe how relaxing our assumption of an inelastic supply function affects our results and consider the change in housing expenditures before and after the reduction in ozone concentration. The next subsection develops the Bartik-Kanemoto upper and lower hedonic bounds and describes how these estimated bounds compare to the general equilibrium estimates by county. In the last subsection we comment on the role of the Willig condition for non-market valuation.

7.2 General Equilibrium Benefit Estimates

The first five columns of Table 5 summarize the key features of the general equilibrium welfare estimates in 1990 dollars for the communities in each county. Column I reports the average (across the school districts in each county) proportionate change in ozone. Column II reports the partial equilibrium estimate, WTP_{PE}, for the ozone change with prices held

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[12] The EPA Retrospective Analysis [1997] considered changes in all the criterion air pollutants that could be attributed to federal air pollution regulations since 1970. The ratio of control to non-control in 1990 for ozone measured by the one-hour peak concentration ranged from a low of .60 to a high (indicating increases) of 1.20.

at their original value in each school district. Column III reports the general equilibrium estimate, WTP_{GE}, for the ozone change which incorporates adjustment of households in response to the change in air quality. Column IV provides the proportionate change in public good provision in equilibrium. Column V summarizes the corresponding average proportionate price change.

Table 5: Comparative Evaluation of Benefit Measures

	No Supply Response								Supply Response	
	I	II	III	IV	V	VI	VII	VIII	IX	X
	Δozone	WTP_{PE}	WTP_{GE}	Δg	Δp	$\frac{WTP_{GE}}{y}$	$\frac{WTP_{GE}}{UB}$	$\frac{WTP_{GE}}{LB}$	WTP^1_{GE}	WTP^2_{GE}
LA	.213	1724	1773	.026	.002	.045	.431	.603	1770	1773
O	.184	1159	1728	.018	-.021	.044	.469	.765	1728	1732
R	.207	1391	832	.038	.043	.022	.198	.297	898	954
SB	.168	1278	846	.035	.033	.022	.223	.341	906	958
V	.062	286	1294	.006	-.047	.034	.362	.547	1301	1310

Dollar estimates for benefits and income measures in 1990 dollars. The supply responses are computed assuming an elasticity of 1 (WTP^1_{GE}) and 2 (WTP^2_{GE}) in computing general equilibrium price responses.

We find that average partial equilibrium estimates of the welfare gains range from \$286 in Ventura County to \$1724 in LA County. Orange, Riverside and San Bernadino county have average gains of approximately \$1200. In contrast, the general equilibrium estimates range from \$832 in Riverside county to \$1773 in LA county. A comparison between Column 1 and Column 2 shows that the partial and general equilibrium estimates differ by as much as \$1008 dollars.

Within each county there is also diversity in the WTP estimates by community (school district). In LA, for example, the WTP_{GE} estimates range from \$994 to \$3004 across school districts. It is important to reiterate that this heterogeneity stems from three characteristics of our general equilibrium analysis. First, the "commodity" used in defining the WTP is a composite. It allows for an exogenous change in ozone, with households permitted to move and alter the air quality they actually experience. The definition given in equation (4.11) captures this effect in distribution between g_j and $\bar{g}_k$. While our analysis assumes that education is invariant in each community, household re-location can cause the final education experienced by a household to be different after the re-location.

Second, households are allowed to have different tastes for public goods. Thus, it is possible–and our estimates confirm–that a few households would actually experience welfare losses that would, with conventional methods, be treated as welfare gains. These losses stem from general equilibrium adjustments. These households experience air quality improvements along with increases in housing prices. If their taste parameter for the public goods is low the composite can yield a loss.

Finally, households have different incomes. Thus the initial level of utility realized under baseline conditions will be different. This outcome is central to the stratification across communities. It also provides an important motivation for relaxing the Willig conditions which would restrict gains due to improvements in non-market resources respond to income.

The differences between partial and general equilibrium estimates also have a strong impact on evaluating the distributional consequences of the improvements in air quality. Comparing Column II and III, we find that the *GE* benefit estimates for Riverside and San Bernadino counties are much lower than *PE* estimates, while the opposite is true for Orange and Ventura county. Columns IV and V provide some explanations for the gap between partial and general equilibrium measures.

As discussed above, the difference between WTP_{GE} and WTP_{PE} is driven by the failure of partial equilibrium measure to account for price changes and household adjustments. In the post improvement equilibrium, consumers face a new locus of community choices ((p, g) pairs). The level of public good provision, g, increases in all 92 school districts. Prices decrease in districts with small improvements and increase in communities with larger improvements. In general, the new equilibrium is characterized by a reduction in the price of achieving a given level of g. The differences between general and partial equilibrium welfare measures mainly arise from the failure to account for the varying price changes associated with public goods changes. In counties with small changes in g (Orange and Ventura) prices drop in the new equilibrium. The partial equilibrium (p, g) pairs in these counties provide a much lower level of utility, at all levels of y and α than do those that are available under the GE prices. Hence, the general equilibrium welfare measures exceed the partial equilibrium measures. Exactly the opposite is the case in the counties that experienced the largest improvements in air quality (Orange and San Bernardino).

To further understand the effects of adjustment, it is useful to consider stylized descriptions of behavior at the disaggregate level. Because the largest air quality improvements occurred in the lower ranked communities, households adjust their locational choices by migrating toward these communities. This consumer response to the increased supply of higher quality air leads to decreased prices in the higher ranked communities and increased prices in the lower ranked communities. Ignoring ownership effects, households in high amenity communities will not only gain due to any further improvement of air quality, but

Table 6: Mean Willingness to Pay Measures By Permanent Income

Income Group	Mean Income	WTP$_{PE}$	WTP$_{GE}$	WTP$_{GE}$/y
1st quartile	24442	994	1039	0.043
2nd quartile	33384	1312	1382	0.041
3rd quartile	41643	1599	1697	0.041
4th quartile	58249	2156	2320	0.040

will also gain due to lower housing prices under the new equilibrium. As we noted earlier some, although not many, households with low tastes for the public good are worse off in the new equilibrium because of price increases in the low g communities.

Table 6 provides some additional insights into the distribution of WTP estimates by income groups. We find that WTP_{GE} estimates are consistently larger than WTP_{PE} estimates when aggregated across all communities in the sample. WTP_{GE} range from \$1039 for the 1st quartile to \$2320 for the 4th quartile. These contrast with the results in Table 5 computed for households in each county separately.

The general equilibrium WTP measures are especially important for large scale changes in non-market amenities like those that have taken place in California. In these circumstances the analysis must incorporate the price changes resulting from adjustment to the change. Of course, it is also important to acknowledge that the merit of including these induced price effects depends on our ability to reliable predict them.

Several indirect gauges of the plausibility of our estimates are possible. First, we compute the community specific price index in 1995 based on actual sale prices and compare these prices with our computed GE measures. This is not an ideal standard. Our analysis holds incomes at the estimates implied by the model and assumes that school quality is constant while the hedonic function for 1995 reflects actual incomes of home buyers as well as all changes in local conditions. Nevertheless, we find that there is a significant positive correlation between the computed prices and the price indices.

The model's qualitative predictions about mobility are also consistent with observed patterns among the five counties in our sample. Of the five counties Ventura, San Bernardino, and Riverside counties experienced the largest population growth during the time period. These are also the counties which experienced the largest improvement in air quality. These findings are consistent with the patterns discussed by Kahn (1997) in his analysis of housing market changes based on the 1980 and 1990 census.

Finally, a further gauge of the overall plausibility is the size of the average willingness to pay in comparison to income. Two types of comparisons are reported – using averages by by community in Table 5 and averages by income quartile in Table 6. Both types of averages yield estimates that fall in the range from 2.2 to 4.5 percent.

7.3 Supply Response and Owner/Renter Distinctions

The last two columns in Table 5 recompute the general equilibrium willingness to pay measures for supply elasticities of 1 and 2. The differences among the three GE estimates are much smaller than those compared with the PE estimates. For example, comparing Column III with Column X, we find that the largest differences in the average WTP_{GE} estimates by county is \$112 dollars for San Bernadino county. For all other counties the differences are much smaller. In contrast the differences between the Column II and Column

III are up to \$1008. Hence we conclude that the average WTP_{GE} estimates are not sensitive to the specification of the housing supply functions.

Table 7: Rent Changes Due to General Equilibrium Adjustments

	Δ Rent		
County	$\tau = 0$	$\tau = 1$	$\tau = 2$
Los Angeles	-7.98	4.64	8.76
Orange	-601.99	-590.18	-586.55
Riverside	581.08	520.95	469.01
San Bernadino	415.71	389.99	341.54
Ventura	-1106.73	-1107.61	-1113.50

As Palmquist (1988) noted, in developing measures of willingness to pay for large scale amenity changes we should note that changes in rents (or expenditures or housing) affect property owners differently from renters. Table 7 summarizes this aspect of our model's implications. Rent changes which are defined as the expenditures for the initial amount of housing, vary substantially across counties with Los Angeles experiencing the smallest changes (in absolute magnitude) and Ventura the largest. Introducing a supply response to price does alter the distribution of gains between owners and renters as illustrated by the last two columns in Tables 5 and 7.

7.4 Comparison to Hedonic Upper and Lower Bounds

In the absence of information that permits identification and estimation of second stage marginal willingness to pay models, conventional hedonic property value models have been used to approximate incremental willingness to pay estimates. An alternative proposed independently by Bartik (1988) and Kanemoto (1988) is to use hedonic models to bound the willingness to pay. Their proposed upper and lower bounds for the WTP are defined as:

$$UB = p_{90}(x, g_{95}) \; - \; p_{90} \tag{7.1}$$

$$LB = p_{95} \; - \; p_{95}(x, g_{90}) \tag{7.2}$$

where x is a vector of housing characteristics, $p_t(\cdot)$ is the hedonic price function estimated for transactions in year t, p_t is housing price in year t (measured in 1990 dollars) and g_t ozone readings for year t for properties sold in that year.

Equation (7.1) defines the upper bound. It uses an estimate of the hedonic property value model with actual sales in 1990 to compute what the housing prices would be with 1995 ozone conditions. The lower bound, given in equation (7.2), uses the sales in 1995 to estimate a hedonic price function and computes the prices in 1990 if these houses had the 1990 ozone concentrations. All sales prices are adjusted to 1990 dollars and measured in annualized terms using Poterba (1992)'s tax corrected annualization factor.

Comparison of the upper and lower bounds with our WTP estimates provides another gauge of the importance of general equilibrium adjustments for benefit estimates derived from large scale environmental changes. We compare the bounds with our WTP estimates by county. The results are given in columns VII and VIII of Table 5. Both the upper and the lower bounds exceed our estimates for WTP_{GE}. The lower bound is generally closer to our estimates but the smallest difference is over 20 percent. The hedonic functions underlying these estimates were specified to be linear to avoid issues associated with the bias in estimates for the predicted prices from the hedonic price function.[13] While there is no reason to expect the WTP estimates for a specific preference function will fall between these bounds, our results provide further indication of the overstatement in partial equilibrium approximations of general equilibrium welfare gains.

7.5 Welfare Estimates and the Willig Condition

It is also possible to use our estimates to offer some perspective on the potential importance of relaxing the Willig condition. Recall when weak complementarity is used to link an amenity to the demand of a private good, we must also assume $(-V_g//V_p)$ is independent of income. In our case, the WTP_{GE} includes a change in g and p. This condition is equivalent to requiring that the WTP per unit of the complementary good is independent of income. A gauge of the importance of this assumption can be developed for a situation where the "commodity" providing the basis for the WTP is a composite (i.e. involving changes in both g and p) using:

$$\frac{\partial(WTP/h)}{\partial y} = \frac{1}{h}\left[\frac{\partial WTP}{\partial y} - \frac{WTP}{y}\,\nu\right] \tag{7.3}$$

Given our estimates, the second term in this expression will be small and departures from the assumption that the expression is zero will stem from the first term inside the brackets.

The responsiveness of WTP_{GE} to income, $\partial WTP_{GE}/\partial y$, is evaluated at a constant level of utility. This expression reduces to a gauge of the sensitivity of the marginal utility of

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[13] We also considered a semi-log (i.e. log of price) and the discrepancies were somewhat smaller but both bounds remained larger than our estimates for WTP_{GE}. A summary with estimates of the hedonic functions is available from the authors.

income to the policy change.

$$\frac{\partial WTP_{GE}}{\partial y} = \frac{V_y(\alpha, y - WTP_{GE}, \bar{g}_k, \bar{p}_k) \; - \; V_y(\alpha, y, g_j, p_j)}{V_y(\alpha, y - WTP_{GE}, \bar{g}_k, \bar{p}_k)} \tag{7.4}$$

Table 8 summarizes our estimates for the range of average values (across school districts in each county). It also reports the elasticity for WTP_{GE} with respect to income, a concept proposed by McFadden and Leonard (1993) to gauge the plausibility of WTP estimates.[14]

Table 8: Change of WTP with Income and the Elasticity of WTP with respect to Income by County

County	$\partial WTP/\partial y$ (range)	ϵ_y^{WTP} (mean)
Los Angeles	.116 - .624	4.23
Orange	.141 - .313	4.16
Riverside	.006 - .126	4.68
San Bernadino	.046 - .130	4.61
Ventura	.128 - .153	4.26

The marginal utility of income is as much as 62 percent greater with the policy than without. For most school districts, the differences are between 10 and 20 percent. Even in smallest cases, the circumstances describing each household's taste for public goods, housing prices and income suggest the demand responsiveness required to offset these changes in WTP with income would need to be substantial.

These findings are also important to the early proposals to use changes in land prices

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[14] There is little direct intuition about what to expect for the elasticity of WTP_{GE} with respect to income (Hanemann, 1999).

to measure the effects of spatially delineated changes in public goods (Lind, 1973; Starrett, 1981). Capitalization relies on two key assumptions: (a) no direct effects of the change being evaluated at the boundary of the area and (b) no sorting within the community. These estimates provide a gauge of the importance of sorting by describing its potential impact on the incentives of different households to sort, given the changes in the public goods and housing prices.

8 Conclusion

We have shown that is is possible to develop general equilibrium benefit estimates for large policy changes using as an example ozone improvements in Southern California between 1990 and 1995. Locational equilibrium models allow estimation of households' preferences, and offer the opportunity to compute the general equilibrium price effects of these large, spatially delineated, changes in air quality. The framework incorporates observed income and unobserved preference heterogeneity. As a result important additional insights can be drawn about: measures of the distribution of gains by location and by income group, estimates of the allocation of gains between owners and renters, and analysis of the importance of assumptions underlying conventional revealed preference methods. Many of these effects have never been estimated before. This application is the first to be conducted at this scale. Further applications should evaluate the importance of preference specification, the incorporation of other spatially delineated public goods and transaction costs for our findings. Nonetheless, our comparison of general equilibrium approximations suggests that this may be an important area for further research essential to improving the benefit analyses of large scale policy interventions.

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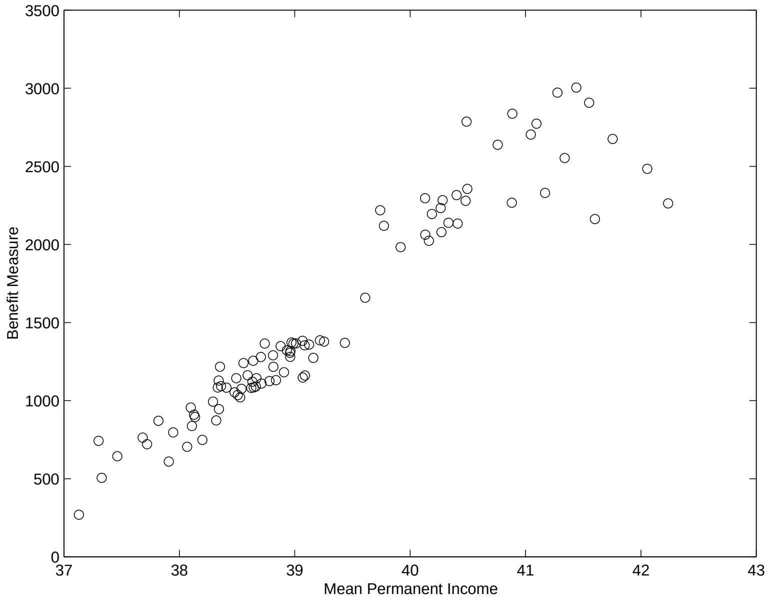
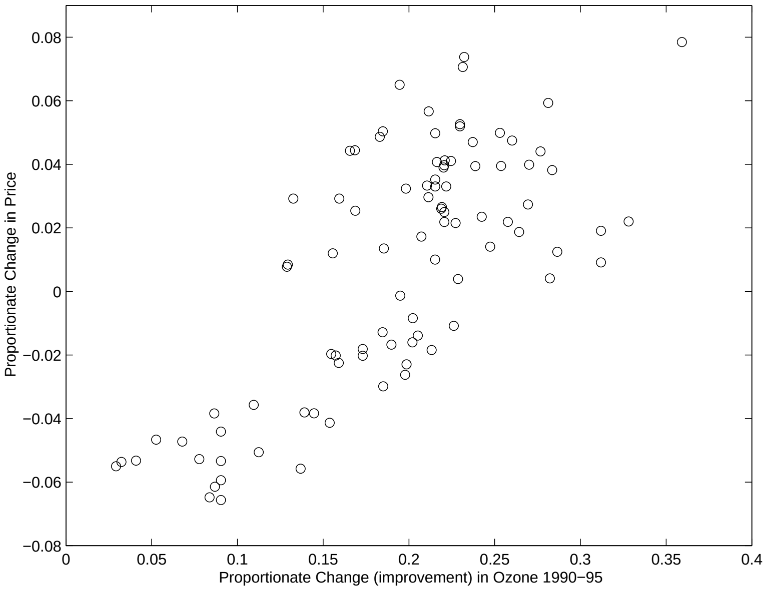
Figure 1:



3500
3000
2500
2000
1500
1000
500
0
37
38
39
40
41
42
43
Benefit Measure
Mean Permanent Income

Figure 2:



0.08
0.06
0.04
0.02
0
−0.02
−0.04
−0.06
−0.08
0
0.05
0.1
0.15
0.2
0.25
0.3
0.35
0.4
Proportionate Change in Price
Proportionate Change (improvement) in Ozone 1990−95

Figure 3: Empirical and Predicted Housing Expenditure Quantiles.