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# Interjurisdictional housing prices in locational equilibrium [☆]

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**Abstract**

In this study, we discuss how to construct interjurisdictional housing price indexes that are consistent with locational equilibrium theory. In theoretical models, housing is assumed to be homogeneous, which provides problems in empirical analysis. We provide conditions that allow us to treat heterogeneous housing as if it were homogeneous. The same conditions that allow us to develop a quantity index for housing, also imply that we can estimate interjurisdictional housing price indexes using hedonic price regressions. Locational equilibrium models impose a number of restrictions regarding the co-movement of housing price indexes, local public goods, and mean income levels. We propose to use these properties of locational equilibrium models to evaluate the different price index estimates. We estimate a variety of price indexes using a unique panel data set of housing transactions in Southern California. Our empirical results, by and large, support our approach. © 2002 Elsevier Science (USA). All rights reserved.

*Keywords:* Housing prices; Locational equilibrium; Interjurisdictional sorting; Price index; Empirical analysis; Aggregation

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## 1. Introduction

Equilibrium models are increasingly used in urban and public economics to study the sorting of households within a system of local jurisdictions as well as to model the provision of local public goods. Communities compete for households by offering different combinations of (property) taxes and local public goods. Households are mobile and "purchase" their desired level of local public goods by moving to their preferred communities. Housing prices determine the implicit costs of a bundle of public goods and adjust so that households with similar preferences for local public goods reside in the same community. In equilibrium, communities with high levels of public good provision will also have high housing prices.[1] Housing prices play, therefore, a key role in locational equilibrium models. An empirical evaluation of these models rests on our ability to estimate interjurisdictional housing prices.

Estimating housing prices so that they are consistent with the underlying theory is a challenging task for a number of reasons. Most locational equilibrium models assume that housing is a homogeneous good that can be consumed in continuous quantities. Under these assumptions, price differences arise from differences in community specific public goods and amenities. In practice, housing is rarely homogeneous. It differs along many observable dimensions. In this study, we explore restrictions on the set of household preferences that allow us to treat heterogeneous housing as if it were homogeneous. Given the heterogeneity in houses, the sales prices that we observe in the data should be interpreted as measures of expenditures for different mixes and amounts of physical characteristics describing these houses.[2] To compute housing prices that are consistent with locational equilibrium theory, we need to develop a framework that decomposes the observed expenditures for heterogeneous houses into a price index specific to each community and a quantity index that controls for the differences in physical attributes. The main insight of our analysis is that the same set of conditions that allow us to treat housing as a homogeneous good, also imply that we can estimate interjurisdictional housing price indexes using hedonic price regressions.

Price index theory alone does not offer a clear basis for selecting the "best" method for constructing interjurisdictional price indexes. Some criteria are needed to evaluate different housing price index estimates. Our proposal is to evaluate these price index estimates based on their consistency with properties predicted

[1] Tiebout [1] first highlighted the importance of interjurisdictional mobility. This paper gave rise to a number of theoretical models including Ellickson [2], Westhoff [3], Epple et al. [4], Goodspeed [5], Epple and Romer [6], Fernandez and Rogerson [7], Nechyba [8,9], Brueckner et al. [10], Brueckner [11], and Henderson and Thisse [12].

[2] As we explain below, there is also a need to convert sales prices of houses into (imputed) rental expenditures.

by locational equilibrium models. If household preferences satisfy single-crossing conditions, locational equilibria satisfy stratification and ascending bundles properties (Epple et al. [4]). These conditions imply that housing prices determine the rank of a community in the hierarchical system of local jurisdictions. Moreover, the price rank of a community equals the rank implied by public good provision. Broadly speaking, locational equilibrium theory implies that any estimates of indexes for housing prices across different jurisdictions must be positively correlated with observed levels of locally provided public goods, community specific amenities, and mean income levels. We can thus evaluate price index estimates based on their consistency with these predictions.

Our empirical analysis reveals some attractive results. The rank predictions of our estimates for interjurisdictional housing price indexes are robust to the econometric specifications for hedonic price functions. Furthermore, we find that the alternative price index estimates are highly correlated with observed amenities such as educational performance and environmental quality. Finally, our overall conclusions do not depend on spatial or temporal aggregations used to define jurisdictions comprising the choice set for the locational equilibrium model.

The remainder of the paper is organized as follows. We review the main insights of locational equilibrium models and develop the framework for evaluating price indexes in Section 2. Section 3 describes our data, a unique data set of the housing transactions in Southern California. Empirical results are reported in Section 4. Section 5 concludes the analysis.

## 2. Housing markets, housing prices, and locational equilibrium

Our analysis begins by using a standard locational equilibrium model with homogeneous housing to describe how sorting of households in response to differences in local public goods and housing market conditions restricts interjurisdictional housing prices. Homogeneous housing is a key feature of this model. In practice, houses differ along many observed dimensions. As a result, we need to devise a test of locational equilibrium models that deals with heterogeneity in housing. Our theoretical analysis suggests that the assumptions required to aggregate heterogeneous housing units so that they can be described as a homogeneous aggregate also imply that we can estimate interjurisdictional price indexes using hedonic price regressions. However, it is not the only approach possible. Price indexes can also be constructed as statistical indexes that are not linked to economic models of choice. We include one adaptation to a pure statistical price index for comparison.

### 2.1. A basic locational equilibrium model with homogeneous housing

Consider the following interjurisdictional equilibrium model discussed in detail in Epple et al. [4]. The economy consists of a continuum of households living in a metropolitan area. The homogeneous land in the metropolitan area is divided among $J$ communities, each of which has fixed boundaries. Jurisdictions may differ in the amount of land contained within their boundaries.

Communities offer a public good $g$ that may itself be thought of as a composite aggregate incorporating locally provided public goods, environmental amenities, and other community specific attributes. For expositional convenience we will refer to $g$ as a public good, but it is important to recognize that our approach is consistent with the interpretation that $g$ is an index of several community specific amenities. Communities differ in the gross-of-tax price of housing $p$ and assess an *ad valorem* tax $t$ on the value of housing services. Thus, the gross-of-tax price $p$ is related to the net-of-tax price $p^h$ by the identity $p = (1 + t)p^h$.

Households have preferences defined over the local public good $g$, a homogeneous local housing good $h$, and a composite private good $b$. For expositional simplicity, consider the case in which households differ only in their endowed income $y$. The continuum of households is then implicitly described by the distribution of $y$. We assume that this distribution has a continuous density $f(y)$.[3]

The preferences of a household are represented by a utility function $U(g, h, b)$, that satisfies standard assumptions. Each household maximizes its utility subject to its budget constraint

$$\max_{(h,b)} U(g, h, b) \quad \text{s.t. } ph = y - b. \tag{2.1}$$

In this basic model, housing is a homogeneous good that can be consumed in continuous quantities.[4] We can also represent the preferences of a household by the indirect utility function derived from solving the optimization problem in Eq. (2.1). The indirect utility function is defined as

$$V(g, p, y) = U(g, h(p, y), y - ph(p, y)). \tag{2.2}$$

The necessary conditions for an interjurisdictional equilibrium can be characterized if we assume that households' preferences are consistent with the single-crossing property. Define

$$M(g, p, y) = -\frac{\partial V(g, p, y)/\partial g}{\partial V(g, p, y)/\partial p}. \tag{2.3}$$

[3] One can extend these models and allow for heterogeneity in preferences Epple and Sieg [13].

[4] Zoning requirements may force individuals to consume a minimum amount of housing, denoted $\bar{h}_j$, when locating in a community. One could add this inequality constraint to the model specification and require that $h \geqslant \bar{h}_j$.

If $M(\cdot)$ is monotonic in $y$ then indifference curves in the $(g, p)$-plane satisfy "single-crossing" in $y$. This condition guarantees that households will sort by income into communities.[5]

To close the model, we assume that public goods and tax rates are chosen by majority rule. Households behave as price takers and mobility between communities is costless. In equilibrium, every household lives in its preferred community, housing markets clear in each community, and public goods and tax rates reflect the preferences of the median voter in each community. Epple et al. [4] provide conditions that guarantee existence of equilibrium for these types of models.

Housing prices play a key role in locational equilibrium models. A household can only "purchase" public goods by moving to a community. Hence, housing prices measure the implicit costs of consuming public goods. Moreover, locational equilibrium theory predicts that there is strong relationship between housing prices, public goods, and mean income levels. Any equilibrium allocation for communities that differ in housing prices and public goods must satisfy the following three properties:

*Boundary indifference:* The set of "border" individuals between any two adjacent communities are indifferent between the two communities.

*Stratification:* The residents of community $j$ consist of those with income $y$ given by the interval

$$y_{j-1} < y < y_j, \tag{2.4}$$

where the $y_j$s are implicitly define by the boundary indifference conditions.

*Ascending bundles:* Consider two communities $i$ and $j$. If $p_i > p_j$, then $g_i > g_j$ as well as $y_i > y_j$.

Stratification implies that individuals will sort by income in equilibrium. The ascending bundles property states that community rankings by housing prices, public good provision and mean incomes are identical in locational equilibrium. If a community has a higher housing price in equilibrium than another community, it must also have a higher level of public good provision and a higher mean income. Broadly speaking, the ascending bundles property also implies that interjurisdictional price indexes for housing and observed community

[5] Brueckner et al. [10] use a similar argument to describe tendencies of households with different income levels to locate in different neighborhoods in an urban area. They define a bid/amenity function with amenities indexed by distance. Assuming that these functions cross once, it is possible to provide a theory of location by income. Thus, we could apply the same logic within a community provided there is an exogenous distribution of locationally delineated amenities.

specific amenities should display high positive correlation.[6] As we noted earlier, there is an important difference between the sales price for a house and an interjurisdictional price index. The former is not an appropriate index of interjurisdictional housing prices because it does not separate the effects of the structural and locational attributes of a house.

In practice, housing price indexes are not observed and need to be estimated. We discuss next how to construct housing price estimates that are consistent with the theoretical framework presented above. There are a number of different ways to construct interjurisdictional housing price indexes. Some of these approaches can be derived from the economic theory of price indexes. Others are convenient statistical indexes. Locational equilibrium theory is not used in the construction of any of these price indexes. However, it implies a strong restriction on the estimates for any interjurisdictional price index, the community specific public goods, and the mean incomes observed in a system of local jurisdictions. We propose to use this relationship as a new basis for evaluating the potential price indexes.

### *2.2. Interjurisdictional price indexes with heterogeneous housing*

We can construct a consistent price index for heterogeneous housing if we impose additional assumptions on household preferences. To illustrate the basic ideas, consider the following example in which housing is heterogeneous and a function of two characteristics $z_1$ and $z_2$. We assume that these characteristics can be purchased in the market at prices $q_1$ and $q_2$. In general terms, preferences can be written as $u(b, h(z), g)$. We selected a CES specification with nested Cobb–Douglas specifications describing the contributions of $g$, $h$, and $b$ to utility and $z_1$ and $z_2$ to $h$ as given in Eqs. (2.5) and (2.6):

$$u(b, h(z), g) = \left\{\left[h(z)^d b^{1-d}\right]^\rho + g^\rho\right\}^{1/\rho}, \tag{2.5}$$

$$h(z) = z_1^a z_2^{1-a}, \tag{2.6}$$

where $a$, $d$, and $\rho$ are parameters of the utility function that are identical among households. The indirect utility function of a household is then given by

$$V = \left\{\left[d^d (1-d)^{1-d}\left[A q_1^{-a} q_2^{-(1-a)}\right]^d y\right]^\rho + g^\rho\right\}^{1/\rho}, \tag{2.7}$$

where $A = a^{-a}(1-a)^{a-1}$. The price index for housing is in this case defined as $p = A q_1^a q_2^{1-a}$. Thus, the indirect utility function of a household is given by

$$V = \left\{\left[d^d (1-d)^{1-d} p^{-d} y\right]^\rho + g^\rho\right\}^{1/\rho}. \tag{2.8}$$

[6] Similar predictions also arise in a number of alternative models in urban economics. For example, bid-rent models with homogeneous households yield similar results for the sales price of housing. Our analysis contrast with this by focusing on the empirical properties of interjurisdictional housing prices indexes.

Under the assumptions made above, an individual's indirect utility function only depends on the price index $p$, income $y$ and public good provision $g$. Even if housing is heterogeneous, we can still derive indirect utility functions that are similar to those obtained when housing is homogeneous.

More generally, consistent aggregation requires that the attributes defining the heterogeneous housing enter preferences in a separable function that is homogeneous of degree one. Samuelson and Swamy [14] summarize this condition as the homogeneity price theorem for an invariant price index. Thus, price index construction would be straightforward if we were prepared to make these assumptions about housing preferences and observed prices of the housing characteristics, $q_1$ and $q_2$.[7] In practice the task is more complex because we only observe the attributes of housing and the sales price.

## 2.3. Estimating interjurisdictional housing price indexes

The primary insight of this section is that the same conditions allowing us to develop a price index for heterogeneous housing imply that we can estimate housing price indexes using hedonic price regressions. To see how that is done, define the sub-expenditure function $e(q_1, q_2, h)$ as the minimum expenditure necessary to obtain $h$ units of housing services. In our example, there is a closed-form solution for the sub-expenditure function

$$e(q_1, q_2, h) = Aq_1^a q_2^{1-a} h = ph = pz_1^a z_2^{(1-a)}. \tag{2.9}$$

Taking logarithms yields

$$\ln(e) = \ln(p) + a\ln(z_1) + (1-a)\ln(z_2). \tag{2.10}$$

Suppose we observe $e_{jn}$ and $z_{1jn}$, but not $z_{2jn}$ for a sample of $N$ housing units in $J$ communities. Define $v_{jn} = (1-a)\ln(z_{2jn})$. We then obtain the following regression model:

$$\ln(e_{jn}) = \ln(p_j) + a\ln(z_{1jn}) + v_{jn}. \tag{2.11}$$

The housing prices $p_j$ are identified (up to scale) by the fixed effects in the regression model. More generally, the vector $z_{jn}$ of housing characteristics should be defined as being composed of observable attributes $x_{jn}$ and unobservables $u_{jn}$. In practice, most housing units in the United States are owner-occupied. We observe market values or subjective beliefs of property values, but not rental expenditures as implied by the model to this point. We therefore need to convert housing values into imputed rental expenditures. Following Poterba [15] we convert the sales or asset price of a house $n$ located in community $j$ into rental

[7] Locational equilibrium is typically viewed as a long-run equilibrium. The specification of the model discussed above does not rule out the possibility that the prices for tradeable housing characteristics would be equalized in equilibrium.

expenditures $r_{jn}$.[8] By construction we have $r_{jn} = e_{jn}$. We then obtain the well-defined regression model in Eq. (2.12)

$$\ln(r_{jn}) = \ln(p_j) + \beta' \ln(x_{jn}) + u_{jn}, \tag{2.12}$$

where $\beta$ is the component of $\delta$ that corresponds to the observed housing characteristics. Estimating interjurisdictional housing price indexes, thus, fits directly into a conventional hedonic regression framework.[9]

As noted at the outset, our example illustrates a well-known aggregation result. If housing is separable from the other goods in the utility function, then the minimum expenditure function of housing does not depend on the other commodities consumed by the household. Furthermore, if the housing function is homogeneous of degree one, the minimum expenditure function for housing services can be factored into two components, a quantity index and a price index.[10] We can, therefore, write the minimum expenditure function of housing as follows:

$$e(q, h(z)) = e(q, z) = p(q) \cdot h(z), \tag{2.13}$$

where $p(q)$ is our housing price index and $h(z)$ is the quantity index. As a result, we do not need to restrict our attention to log–log specifications of hedonic regressions. Other functional specifications of the fixed-effects hedonic regression, including log–linear models, can also be justified as long as they are consistent with homogeneity of degree one of $h(z)$. We investigate how sensitive our results are to functional specifications of the hedonic regression model.[11]

This sensitivity analysis also allows us to investigate the assumption that the demand for housing does not not depend on the level of public good provision, once we condition on the community choice. If it does, we would expect to find that our empirical findings will be sensitive to the price index used to take account of differences in private consumption of housing. If empirical findings are robust across price index estimates, we view this evidence as supporting the assumption that private housing is separable in household preferences from the public goods distinguishing the communities.

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[8] Note that the tax rate is the only variable in the Poterba formula that depends on the location. In our application, we study the Los Angeles metropolitan area. California school finance laws imply that tax rates are for all practical purposes constant across school districts in the LA metropolitan area. Hence the conversion factor does not depend on the school district. Thus converting housing values into imputed rents is equivalent to rescaling the dependent variable. Since price estimates are only identified up to scale, there is therefore no need to apply the procedure in our applications. If tax rates would differ among school districts, housing price conversions would affect the price index estimates because of differences in property tax rates.

[9] Eq. (2.12) can also be interpreted in the spirit of Griliches [16] as a price function. This interpretation contrasts to Rosen's [17] interpretation of (2.12) as an equilibrium price locus.

[10] For a detailed discussion of this property, see Samuelson and Swamy [14].

[11] We also investigate linear specifications that are harder to justify on theoretical grounds, but commonly used in practice.

### 2.4. A statistical approach to constructing interjurisdictional housing price indexes

Price indexes can be constructed based on economic or axiomatic frameworks. The economic approach, that has been used in the previous section, derives the price index from restrictions on preferences. The axiomatic approach derives indexes to satisfy a predefined set of properties apart from an economic model of behavior.[12] Using this perspective, a price index can also be constructed from hedonic price regressions as the average price of Leontief [19]–Theil [20] (LT) statistical composite commodities.

The intuition for the LT strategy is as follows: Consider a randomly sampled set of housing types, denoted by $\{x_1, \ldots, x_S\}$. We predict the prices of the different types in each community using hedonic regressions. Let $\hat{r}_j(\cdot)$ denote the estimated hedonic price function in community $j$. The predicted price of housing type $x_s$ in community $j$ is then given by $\hat{r}_j(x_s)$. The average predicted price for the set of houses is

$$r_j = \sum_{s=1}^{S} w_s \hat{r}_j(x_s), \tag{2.14}$$

where $w_s$ is the share of houses of type $s$. The average predicted price for this set of houses is a price index since it holds the quantity of housing constant across communities. We estimate these price indexes as an alternative benchmark for the fixed-effect prices discussed above.[13]

### 2.5. Summary

This section illustrated how to construct interjurisdictional housing price indexes that are consistent with locational equilibrium models. Our analysis is motivated by a special class of locational equilibrium models with homogeneous housing goods. However, the issues that arise in developing price indexes also apply to a large number of different models, such as monocentric city models or simple bid-rent models of housing markets. Empirical analysis of the properties of most equilibrium models in urban economics ultimately requires the development of housing price indexes because housing is usually treated as a homogeneous commodity. As we have discussed, this is a challenging, but not insurmountable task given the available data. Nonetheless, it is reasonable to ask the question whether it would be more fruitful to pursue alternative theories that acknowledge

[12] See Diewert [18] for a survey.

[13] Alternative approaches for constructing intertemporal price indexes based on hedonic regressions are the repeat sales proposed by Bailey et al. [21] and a hybrid model proposed by Case and Quigley [22]. For a general discussion of these methods see also Palmquist [23] and Zabel [24].

the fact that housing units are heterogeneous. For example, Nechyba [9] provides a theoretical model with a small number of heterogeneous housing types. Bayer [25] has recently estimated a differentiated product model similar to the basic Berry et al. [26] model. While these models clearly provide new insights into the relationship between housing markets and public good provision, they suffer from some of the same problems that we discussed. For these models to be computationally tractable, one needs to restrict the set of admissible housing types to a relatively small set—at least small compared to the heterogeneity in housing alternatives observed in the data. Thus empirical applications ultimately face similar aggregation problems and need to estimate multiple interjurisdictional housing price indexes: one set of interjurisdictional price indexes for each type of housing good.

## 3. Data

To illustrate the role of locational equilibrium models in discriminating between interjurisdictional price indexes we analyze interjurisdictional sorting in the Los Angeles metropolitan area. For the implementation of our empirical model, communities are defined as school districts. The geographic domain of our study consists of most of the districts in the five counties: Los Angeles, Orange, Riverside, San Bernardino, and Ventura. We use housing transactions assembled by Transamerica Intellitech, a market research firm. Our data set contains housing characteristics and transaction prices for virtually all housing transactions in Southern California between 1985 and 1995.[14]

The data for the current analysis are restricted to the years 1988–1992 (a five-year window surrounding 1990). The housing sales take place inside a ring surrounding the city of LA roughly defined by a series of national forests (specifically, Los Padres National Forest, Angeles National Forest North, Angeles National Forest, and San Bernardino National Forest). It is approximately equal to the Air Resource Board's definition of the South Coast and Ventura Air Basins. Much of the omitted area in the east is sparsely populated desert. This definition of the LA metropolitan area implies that we do not include all school districts in the five counties. Instead, we omit one school district in northern Ventura County, one in northern Los Angeles County, all but nine in the southwestern corner of San Bernardino County, and all but fourteen in western Riverside County. Communities in the selected area are part of the same air basin. Our

[14] In contrast, other work has relied on home-owners' subjective assessments of their homes' value (Zabel and Kiel [27]) or on samples subject to selection effects due to eligibility criteria such as the FHA houses (Chattopadhyay [28]). An appendix that describes in detail how our sample is constructed from the raw data is available from the authors.

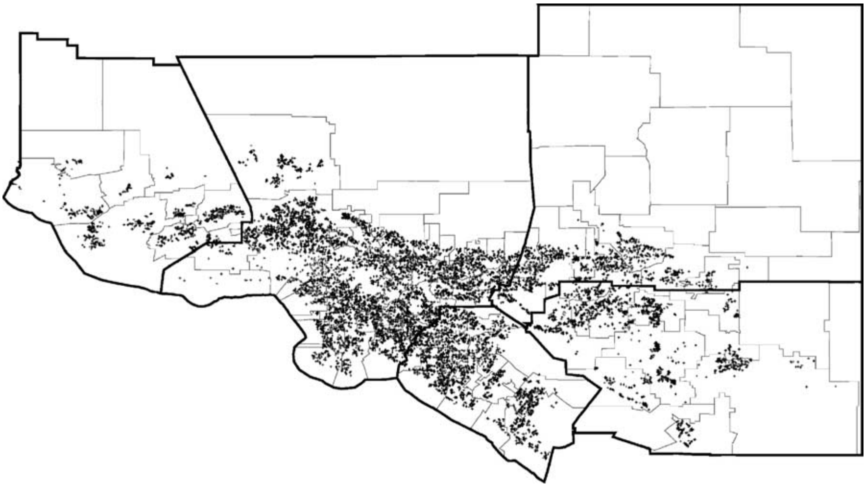

Fig. 1. A map of the LA metropolitan area and our housing sample.

Table 1
Descriptive statistics of the housing sample

| Variable | Orange | Riverside | San Bernardino | Ventura | Los Angeles |
|---|---|---|---|---|---|
| Number of observations | 40,894 | 33,132 | 24,493 | 14,817 | 109,529 |
| Price | 253,315 | 139,771 | 151,313 | 244,888 | 243,889 |
| 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.

complete sample contains 92 school districts. A map plotting our sample of housing transactions for these counties is given in Fig. 1.

The data contain a rich set of independent variables, including the age of the house, the number of bedrooms and bathrooms, the size of the building and lot, as well as qualitative variables for the presence of a swimming pool and a fireplace. Table 1 reports means of the main variables in our sample by county.

We find that significant differences across counties are present only for price, the variable to be explained in our research, in age of the house, and in the average percentage of fireplaces. Mean prices are higher for the LA, Orange, and Ventura counties. These are coastal counties. In addition, two of the counties (LA and Orange) are more urban than the others.

To investigate how well our sample reflects the properties of the underlying housing stock in each community, we compared our sample to the aggregate statistics based on the 1990 US Census. Even after adjusting for inflation, prices tend to be uniformly higher in the US Census. Across our 92 school districts, prices are 6–12% higher in the census (inter-quartile range). This might suggest smaller homes are over-sampled in our data set, but the data do not confirm this hypothesis. In fact, 34% of houses in the US Census have two or fewer bedrooms, compared to only 19% in our sample. In addition, homes are much newer in our data set. 15% of our houses were built less than 1 year ago, whereas in the 1990 census only three percent of homes were built in 1989–1990.

It is not clear why homes appear more expensive in the US census despite being older and having fewer bedrooms on average. Four possibilities suggest themselves: (1) the US Census, which does not rely on transactions data, systematically overestimates housing prices; (2) other characteristics that affect housing prices, included in our hedonic regressions but not available in the US census for comparison, are systematically different in the stock of houses and the set of housing transactions; (3) unobserved variables in our hedonic regression are systematically higher in the stock of houses than in the set of housing transactions; and (4) less valuable houses turn over more often. We can not determine whether one or combination of these potential explanations is driving what we observe.[15]

We supplement the housing data with several measures of local public goods, environmental amenities and other school district specific characteristics. Ozone and PM10 measures are obtained from the California Air Resource Board. Air pollution measures are matched to each house based on the nearest monitoring station. The process used in this study takes account of the year of the sale and the availability of air pollution data for each monitoring station. The measures for ozone and PM10 are a five-year centered average of the readings from the closest monitors for each house based on the year the house was sold. Ozone is measured in parts per million (ppm) as the average of the top 30 one-hour daily maximum readings. Particulate matter is measured by the annual geometric mean (in $\mu g/m^3$ for particulate matter of 10 µm or less in size).

Local school quality is measured by standardized test scores using the 1992–1993 California Learning Assessment System Grade Level Performance Assessment test. Each student taking this exam is assessed at one of 6 performance levels (with 6 the highest level). School district income distributions and other measures of school quality such as student teacher ratios and expenditures per student were taken from estimates developed by the National Center for Education Statistics and School District Data book. Finally, public safety is measured by crimes per 10,000 households taken from the FBI Uniform Crime Reports.

[15] The price difference of 10 percent is consistent across counties, and this result is encouraging.

## 4. Empirical results

### 4.1. Constructing interjurisdictional price indexes

Table 2 reports the estimation results from the regressions for five specifications of the hedonic price equations using the imputed rental expenditures as the dependent variable. Columns 1–3 report estimates using fixed-effects (FE) regressions. Column 1 (2, 3) uses a log–log (log–linear, linear–linear) specification.[16] Column 4 is the same fixed-effects regression as column 1, but estimates the model with quantile regression. Finally, columns 5–9 report separate regressions by county. We find that most of the estimated coefficients have the expected sign and have plausible magnitudes. The $R^2$s of most regressions are well over 0.6 indicating that the structural characteristics observed in our sample explain a large fraction of the observed variation in prices.[17] The sample sizes of all regressions, including those for each county, are quite large. With this large sample size, single observations have an almost negligible impact on the parameter estimates.

Table 3 reports summary statistics of the estimates of the interjurisdictional price indexes obtained from five different hedonic regression approaches. Columns 1 and 2 are based on fixed-effects specifications and correspond to the same columns in Table 2. Column 3 reports our adaptation of the Leontief–Theil (LE) estimates based on the specification in column 3 in Table 2. Column 4 is based on the same column in Table 2. Finally, column 5 reports Leontief–Theil estimates based on columns 5–9 in Table 2.

In general, we find that the summary statistics reported in Table 3 are robust to the choice of functional form and estimation technique. The highest priced school districts have price levels between 6 and 6.5 times that of the lowest. The districts yielding the highest and the lowest values are robust across methods and functional specifications. The only exception is the specification in column 3 that produces slightly larger housing price estimates for the high amenity communities. Housing prices are less dispersed than one might expect based on the range of the estimates. The estimated standard deviation of housing prices is approximately 0.45.

The high degree of similarity among the different regressions is also reflected in the correlations among the pairs of these indexes. Generally, these correlations exceed 0.90. The largest differences across price index models result from the inability of the linear model to account for the extremely high prices in the Beverly Hills school district. If we remove Beverly Hills from the sample, the correlation

[16] log–log refers to a specification using the log of the imputed rent as the dependent variable and the logs of all explanatory variables that are not dummy variables.

[17] Of courses, differences in the $R^2$ are partially due to functional form assumptions and cannot be used for model selection since the models are not nested.

Table 2
Hedonic price regression results

| | OLS Full Sample ln–ln | OLS Full Sample ln–lin | OLS Full Sample lin–lin | Quantile Full Sample ln–ln | OLS Los Angeles ln–ln | OLS Orange ln–ln | OLS Riverside ln–ln | OLS San Bernardino ln–ln | OLS Ventura ln–ln |
|---|---|---|---|---|---|---|---|---|---|
| Bathrooms | 0.0366 | 0.0372 | 19765.03 | 0.0221 | 0.0362 | 0.0500 | 0.0278 | 0.0187 | 0.0168 |
| | (0.0018) | (0.0018) | (476.48) | (0.0010) | (0.0025) | (0.0058) | (0.0043) | (0.0056) | (0.0043) |
| Bedrooms | −0.0390 | −0.0384 | −15034.13 | −0.0203 | −0.0534 | −0.0460 | −0.0107 | −0.0014 | −0.0363 |
| | (0.0012) | (0.0018) | (319.05) | (0.0006) | (0.0018) | (0.0035) | (0.0023) | (0.0032) | (0.0028) |
| Pool | 0.0656 | 0.0637 | 16484.14 | 0.0512 | 0.0830 | 0.0430 | 0.0533 | 0.0612 | 0.0616 |
| | (0.0021) | (0.0021) | (557.9) | (0.0011) | (0.0032) | (0.0057) | (0.0041) | (0.0055) | (0.0047) |
| Fireplace | 0.0747 | 0.0758 | 8998.52 | 0.0542 | 0.0714 | 0.0257 | 0.1040 | 0.0662 | 0.0735 |
| | (0.0018) | (0.0018) | (461.5) | (0.0009) | (0.0026) | (0.0049) | (0.0045) | (0.0058) | (0.0049) |
| Building (SqFt) | −2.0506 | 0.0006 | 34.4877 | −2.66 | −1.99 | −2.23 | −1.35 | −0.9142 | −0.7210 |
| | (0.068) | (4.91E−06) | (1.29) | (0.035) | (0.10) | (0.23) | (0.17) | (0.23) | (0.20) |
| Lot (SqFt) | 0.1424 | 1.29E−05 | 2.18 | −0.0756 | 0.0880 | 0.2845 | 0.2845 | −0.6506 | −0.6177 |
| | (0.033) | (3.89E−07) | (0.102) | (0.017) | (0.061) | (0.094) | (0.064) | (0.11) | (0.091) |
| Age | −0.3120 | −0.0054 | −1334.61 | −0.2076 | −0.2735 | 0.3196 | −0.2260 | −0.0031 | 0.6886 |
| | (0.017) | (0.0002) | (51.88) | (0.0090) | (0.032) | (0.098) | (0.032) | (0.048) | (0.042) |
| $Age^2$ | −0.0170 | 1.06E−05 | 0.2501 | −0.0243 | −0.0244 | −0.0167 | −0.0149 | −0.0241 | −0.0406 |
| | (0.0006) | (1.83E−06) | (0.479) | (0.0003) | (0.0011) | (0.0035) | (0.0011) | (0.0015) | (0.0015) |
| $Lot^2$ | −0.0096 | −1.45E−10 | −4.53E−05 | −0.0184 | 0.0027 | 1.67E−04 | −0.0042 | −0.0090 | 0.0503 |
| | (0.0014) | (4.18E−12) | (1.1E−06) | (0.0007) | (0.0034) | (0.0041) | (0.0024) | (0.0053) | (0.0041) |
| $Building^2$ | 0.1749 | −4.38E−08 | 0.0135 | 0.1863 | 0.1785 | 0.2170 | 0.1499 | 0.0419 | 0.1286 |
| | (0.0050) | (8.62E−10) | (0.0002) | (0.0026) | (0.0075) | (0.017) | (0.013) | (0.017) | (0.015) |
| Age × lot | 0.0326 | 1.23E−07 | 0.0556 | 0.0401 | 0.0187 | 0.0089 | 0.0404 | 0.0512 | 0.0245 |
| | (0.0014) | (5.99E−09) | (0.0016) | (0.0007) | (0.0027) | (0.0076) | (0.0022) | (0.0043) | (0.0032) |
| Age × building | 0.0050 | 2.35E−07 | 0.3820 | −0.0153 | 0.0231 | −0.0505 | −0.0181 | −0.0595 | −0.1069 |
| | (0.0023) | (6.25E−08) | (0.016) | (0.0012) | (0.0043) | (0.014) | (0.0049) | (0.0069) | (0.0061) |
| Lot × building | 0.0098 | −3.82E−10 | 4.88E−04 | 0.0612 | −0.0064 | −0.0210 | −0.0265 | 0.1079 | −0.0380 |
| | (0.0046) | (1.19E−10) | (0.00003) | (0.0024) | (0.0084) | (0.015) | (0.0091) | (0.014) | (0.012) |
| $N$ | 221,296 | 221,296 | 221,296 | 221,296 | 109,076 | 39,856 | 32,948 | 24,633 | 14,753 |
| $R^2$ | 0.6295 | 0.6285 | 0.6690 | 0.5365[a] | 0.5953 | 0.4166 | 0.6294 | 0.6727 | 0.7146 |

Estimated standard errors are given in parentheses. All regressions control for fixed effects.

[a] Pseudo $R^2$ obtained from quantile regression.

Table 3
Descriptive statistics of price index estimates by school district

| | FE ln–ln | FE ln–lin | LT FE lin–lin | Quantile FE ln–ln | LT by county ln–ln |
|---|---|---|---|---|---|
| Mean | 0.99 | 1.01 | 1.07 | 1.01 | 0.99 |
| Std. deviation | 0.42 | 0.43 | 0.59 | 0.44 | 0.41 |
| Min | 0.44 | 0.46 | 0.49 | 0.44 | 0.43 |
| Max | 2.70 | 2.96 | 4.97 | 2.90 | 2.65 |
| Correlation | | | | | |
| Mean income | 0.683 | 0.687 | 0.685 | 0.675 | 0.682 |
| Median income | 0.512 | 0.503 | 0.456 | 0.496 | 0.513 |

The table above contains summary statistics of the housing price index estimates for the 92 school districts in our sample.

between the linear model's price predictions and those of the other models all exceed 0.97.

Specifying the temporal and geographical boundaries delineating the communities assumed to be part of local housing markets is another important decision in empirical analysis. We compare the effects of changing the geographic and temporal dimensions of the price index estimates using three different market definitions over the greater Los Angeles area. Based on the results in Table 3 and our comparison of pairwise correlations of the price indexes, we confine this evaluation to the fixed-effect log–log specification of the price index.

The first market definition includes all residential housing sales between 1988 and 1992 for 92 school districts in Ventura, Los Angeles, Orange, Riverside, and San Bernardino counties. This definition includes the entire greater LA metropolitan area west of the San Gabriel Mountains. The second market definition includes the same years (1988–1992), but is restricted to 36 school districts west of the San Gabriel Mountains in Orange, Riverside and San Bernardino counties. The third market definition includes the same 36 school districts as the second, but uses a longer time period (1985–1995). We computed correlations between price indexes under the three different market definitions. We find that the pairwise correlations all exceed 0.99, suggesting that for the greater Los Angeles area the price indexes provide a comparable assessment of the relative prices attributed to communities regardless of the geographic or temporal definitions used in this area.

The first step in the evaluation of these price indexes suggests that interjurisdictional indexes are robust with respect to: (a) the rationale (i.e., economic–axiomatic) for the price measures, (b) functional form assumptions, (c) different sampling schemes used in the construction of the indexes, (d) spatial aggregation, and (e) temporal aggregation.

Table 4
Rank differences across price indexes

| | Fixed effect ln–ln | | | | Fixed effect ln–ln | | | | Leontief–Theil fixed effect lin–lin | | | | Quantile fixed effect ln–ln | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | # com. w/dif. | | | mean | # com. w/dif. | | | mean | # com. w/dif. | | | mean | # com. w/dif. | | | mean |
| | >0 | >3 | >5 | diff. | >0 | >3 | >5 | diff. | >0 | >3 | >5 | diff. | >0 | >3 | >5 | diff. |
| FE ln–ln | 0 | 0 | 0 | 0.00 | – | – | – | – | – | – | – | – | – | – | – | – |
| FE ln–lin | 44 | 2 | 0 | 0.78 | 0 | 0 | 0 | 0.00 | – | – | – | – | – | – | – | – |
| LT FE lin–lin | 79 | 21 | 8 | 2.50 | 81 | 21 | 7 | 2.57 | 0 | 0 | 0 | 0.00 | – | – | – | – |
| Qantile FE ln–ln | 68 | 15 | 7 | 2.00 | 70 | 19 | 7 | 2.20 | 82 | 26 | 15 | 2.93 | 0 | 0 | 0 | 0.00 |
| LT by county ln–ln | 42 | 2 | 1 | 0.72 | 48 | 6 | 3 | 1.00 | 78 | 20 | 9 | 2.50 | 64 | 17 | 9 | 2.04 |

The table reports summary statistics of rank comparisons based on our price index estimates.

### *4.2. Evaluating interjurisdictional price indexes*

We have seen in the previous section that we can estimate interjurisdictional housing price indexes in a number of different ways. To evaluate the different methods, we analyze whether our estimates are consistent with the predictions of locational equilibrium models. At the bottom of Table 3 we report the first evaluation of the price indexes. This is the correlation between each estimate of the housing price index and a measure of average community income. The correlations are clustered near 0.68. The largest discrepancies are with the linear specification reported in column 3. The primary source of this discrepancy is the Beverly Hills School District that is the highest priced school district and an outlier in all models. The next highest price under the linear specification is 2.95, a difference of almost 4 standard deviations from the level of 4.97 estimated for Beverly Hills.

Housing prices also determine the rank of a community in the hierarchy of communities. We analyze whether the rank predictions are stable across price index estimates. Table 4 reports the number of communities whose ranking by price index differs by more than 0, 3, or 5 places when compared with each of the other price index models. It also reports the mean difference for each pairwise comparison. We find that the rankings are quite stable across price indexes. In all but one of the comparisons, less than 10% of the communities' ranking varies by more than 5 places relative to a total of 92 communities. We also compute Kendall's coefficient of concordance—a nonparametric test statistic that measures the degree of similarity between different rankings. We find that Kendall's coefficient is larger than 0.99 for all pairwise comparisons of rankings as well as the comparison of all five rankings. The different price index approaches primarily lead to small perturbations of the rankings of the communities in the sample.

The next component of our evaluation uses the "ascending bundles" property. Community specific housing price indexes and levels of public goods should monotonically increase in the rank of the community determined by the housing price indexes. Since the level of public good provision determines the desirability of a community, higher amenity communities must have higher housing prices, which gives rise to a strict hierarchy of communities. One problem with analyzing the empirical content of the ascending bundles property is that the level of public good provision is a function of many amenities, and some important components of them may be unobserved to the econometrician. It would be surprising if estimated housing prices and a single observed amenity satisfied this condition perfectly. We address this issue in several ways.

First, we consider the simple correlation structure between housing price indexes along with measures of public education, public safety, and ambient air quality. We compute correlations between each of the price indexes and each of several measures of the public goods. Public education is measured by expenditures per student, students per teacher, and average 10th grade scores for reading, writing, and math. Measures for air quality are based on ambient concentrations of particulate matter and ozone using five-year averages centered at the year of the sale of the house. Means for each school district are developed by first attaching the readings from monitors closest to the house. These readings are then averaged at the school district level. The results are shown in Table 5.

We find that the correlation is negative for bads (students per teacher, crime, and air pollution) and positive for goods (expenditures per student and average test scores). Housing prices are most strongly correlated with output measures of school quality. The correlation coefficients often exceed 0.5. Correlations are also strong with the number of students per teacher, which is an input measure of school quality. Correlations are small with expenditure measures for public education. This should not be surprising given the implications of Proposition 13. Passed in 1978 in California this proposition, along with the 1971 decision of the state supreme court in *Serrano vs. Priest*, caused equalization of school expenditures among school districts. Despite these efforts to equalize expenditures our analysis suggests, that there are significant differences in school quality as measured using test scores among the school districts in the sample. These differences are captured in estimates of interjurisdictional price indexes. Housing price indexes are also negatively correlated with measures of air quality. Correlations are negative as expected when prices are related to the concentration of pollutants. They range from $-0.18$ to $-0.32$. As the equilibrium model implies, communities with lower levels of ozone (higher air quality) tend to have higher housing prices in our sample. Taking each measure of the two public goods independently, the correlation analysis is consistent with the predictions of the ascending bundles property.

An alternative test of the ascending bundles property focuses on the rankings implied by the price and public good measures. Let $R_i^p$ be the rank of community $i$ based on the estimated price index $p$ and $R_i^g$ the rank of community $i$ based on the public good measure $g$. If prices and public goods exactly satisfy the ascending bundles property, we would have an exact correspondence in the ranks by the two criteria, i.e., $R_i^p = R_i^g$ for all communities in the sample. However, our

Table 5
Correlations of price indexes and local public goods

| | Crimes per 10,000 | Expenditure per student | Students per teacher | 5 Year PM average | 5 Year ozone average | Reading test average | Writing test average | Math test average |
|---|---|---|---|---|---|---|---|---|
| FE ln–ln | −0.2236 | 0.2050 | −0.5602 | −0.2592 | −0.3106 | 0.5033 | 0.4065 | 0.6017 |
| FE ln–lin | −0.2147 | 0.2182 | −0.5780 | −0.2419 | −0.2929 | 0.4856 | 0.4014 | 0.5906 |
| LT FE lin–lin | −0.1256 | 0.2939 | −0.6345 | −0.1815 | −0.2443 | 0.3852 | 0.3583 | 0.5267 |
| Quantile FE ln–ln | −0.2083 | 0.2257 | −0.5733 | −0.2393 | −0.2895 | 0.4865 | 0.3975 | 0.5934 |
| LT by county ln–ln | −0.2254 | 0.2057 | −0.5516 | −0.2691 | −0.3183 | 0.5092 | 0.4081 | 0.6031 |
| Crimes per 10,000 | 1.0000 | – | – | – | – | – | – | – |
| Expenditure per student | 0.0658 | 1.0000 | – | – | – | – | – | – |
| Students per teacher | 0.1380 | −0.2922 | 1.0000 | – | – | – | – | – |
| 5 Year PM average | 0.1723 | −0.0187 | 0.1111 | 1.0000 | – | – | – | – |
| 5 Year ozone average | 0.0808 | −0.1201 | 0.1296 | 0.7457 | 1.0000 | – | – | – |
| Reading test average | −0.3757 | −0.0947 | −0.3794 | −0.2570 | −0.1702 | 1.0000 | – | – |
| Writing test average | −0.3073 | −0.0734 | −0.3390 | −0.0055 | −0.0712 | 0.6899 | 1.0000 | – |
| Math test average | −0.4902 | −0.0837 | −0.4657 | −0.1458 | −0.0600 | 0.7804 | 0.6911 | 1.0000 |

The table reports correlations of price index estimates and community specific amenities based on our sample of 92 communities.

ability to measure public goods is limited. It is, therefore, unlikely that these rank predictions will hold perfectly in the data. As a result, we use a rank test statistic to assess the discrepancy between the two rankings, by counting the number of rank violations

$$S_J(q) = \frac{2}{J(J-1)} \sum_{i=1}^{J} \sum_{j=1}^{i-1} 1\{R_j^g - R_i^p > q\}, \tag{4.1}$$

where $1\{\cdot\}$ denotes the indicator function, that is equal to one if the event in brackets is true and zero otherwise. $J$ is the number of communities in the sample. We compute the test statistic above by setting $q = 0,\ 5,\ 10$, i.e., by counting rank violations larger than 0, 5, or 10. There are $J(J-1)/2$ pairwise comparisons between communities. This number is therefore used to normalize the test statistic. If both rankings are identical, we have $S_J(q) = 0$. If the ranking are exactly the opposite, we have $S_J(q) = 1$. Hence

$$0 \leqslant S_J(q) \leqslant 1. \tag{4.2}$$

Thus values of $S_J(q)$ close to zero indicate that the two rankings are similar, values near one indicate that the rankings are very different. We compute our test statistic for different choices of public good measures. The first approach uses the median math score within the school district as a measure for school quality. The results are shown in the first three columns of Table 6.

As we noted earlier, we can also view $g$ as an index of local public goods that allows us to model heterogeneity in public good provision. The index assumption is clearly restrictive and rules out the case that households have heterogeneous preferences for multiple types of public goods. Relaxing this assumption is difficult since equilibria in models with that type of heterogeneity will in general not be hierarchical. The lack of hierarchy raises a large number of difficulties both

Table 6
Rank violations of price indexes

| | Rank violations math score | | | Rank violations public good index | | |
|---|---|---|---|---|---|---|
| | $S_J(0)$ | $S_J(5)$ | $S_J(10)$ | $S_J(0)$ | $S_J(5)$ | $S_J(10)$ |
| FE ln–ln | 0.309 | 0.264 | 0.221 | 0.250 | 0.204 | 0.157 |
| FE ln–lin | 0.312 | 0.266 | 0.224 | 0.253 | 0.203 | 0.161 |
| LT FE lin–lin | 0.318 | 0.272 | 0.230 | 0.249 | 0.200 | 0.157 |
| Quantile FE ln–ln | 0.312 | 0.267 | 0.224 | 0.258 | 0.208 | 0.165 |
| LT by county ln–ln | 0.307 | 0.260 | 0.218 | 0.250 | 0.201 | 0.158 |

The table reports summary statistics of rank violations based on our price index estimates.

from a theoretical and an empirical perspective. Given these difficulties, we do not venture outside the single-index model.

The second approach for measuring rank violations thus uses a linear index of school quality $s$, air quality $e$, and crime $c$. Since the index is linear, it can be written as

$$g = s + \gamma_e e + \gamma_c c, \tag{4.3}$$

where $\gamma_e$ ($\gamma_c$) is a coefficient measuring the relative importance of air quality (public safety).[18] We compute the number of rank violations for each definitions of $g$ as we vary the values of $\gamma_e$ and $\gamma_c$. In the last three columns of Table 6 we report the results for the index that minimizes rank violations. For each measure of public goods we compute three versions of the test statistic discussed earlier.

Table 6 shows that rank violations occur in 31% of all pairwise comparisons, if we use the math score as measure of the community specific public good. If we only count rank violations larger than 5 (10)—if we ignore small perturbations of the rankings—we find that the test statistic drops to approximately 0.27 (0.22). The number of rank violation is markedly smaller for linear combinations of amenities and the math score than those obtained for the math score alone. Rank violations occur in only 25% of all pairwise comparisons if we use the more comprehensive measure of public good provision. The statistic drops to 0.200 (0.157) if we count only rank violations larger than 5 (10). Overall, these results suggest that the price rank of a community is closely related to its local public good rank. Using combinations of amenities and local public goods reinforces our finding that price index estimates yield rankings of communities that are similar to those implied by observed measures of a single local public good. These findings hold regardless of the selection of our price index model.

Comparing the five different price index estimates, we find that the simple Leontief–Theil specification estimates greater price differences between the low income and high income school districts relative to other specifications. The analysis of the rank differences suggests that the fixed-effects indexes and the Leontief–Theil by county specification yield almost identical rankings. Choosing among these three specifications can therefore be done based on computational ease. The fixed-effects specifications are much easier to compute than the Leontief–Theil index.

## 5. Conclusions

In this paper, we have discussed how to construct interjurisdictional price indexes that are consistent with locational equilibrium theory. In theoretical

[18] Since the unit of measurement of the public good index is arbitrary, we have normalized the coefficient of school quality to be equal to one.

models, housing is assumed to be homogeneous, which provides problems in empirical analysis. We have discussed conditions that allow us to treat heterogeneous housing as if it were homogeneous. Furthermore, we show that the same conditions that allow us to aggregate housing also imply that we can estimate interjurisdictional housing price indexes using hedonic price regressions.

There are a number of interjurisdictional price index estimates that are consistent with locational equilibrium theory. We therefore need to find additional criteria to evaluate different index models. If household preferences are consistent with single-crossing conditions, then equilibrium allocations will satisfy stratification and ascending bundles properties. We have shown that these properties impose a number of restrictions regarding the comovement of housing price indexes, local public goods, and mean income levels. We propose to use these properties of locational equilibrium models to evaluate the different price index estimates.

Our application focuses on the LA metropolitan area. We find that the estimated price indexes are robust with respect to the rank predictions for the communities in the choice set. Differences for rank predictions are small. They do not markedly alter the hierarchy defined for the communities. We also consider whether our conclusions are impacted by the choice set used in implementing the model. We find that price indexes are robust to different spatial market definitions and temporal definitions in price measurement. Because our analysis is based on proxies of public goods, we would not expect the ascending bundles prediction to hold perfectly in data. Nonetheless all correlations for individual measures of public goods and amenities have the expected sign and are often large. The analysis of rank violations measured for median math scores or a simple linear index of local public goods reinforce our findings.

Constructing and evaluating interjurisdictional housing price estimates is only the first step towards a general empirical evaluation of locational equilibrium models.[19] Our approach for evaluating different price indexes, however, has relevance beyond applications to locational equilibrium models. By exploiting the conditionality of a test of the equilibrium to the community specific price index used for housing, it is possible to use the consistency of the model's predictions to gauge the importance of the assumption that household preferences can be treated as separable in the private characteristics of housing and the community specific public goods. The robustness of our results confirms a strategy that first adjusts for heterogeneity in housing before using housing price indexes and public goods to evaluate locational equilibrium models.

[19] These models are an increasingly important part of current empirical research in urban economics, see Epple and Sieg [13], Epple et al. [29], Bayer [25], Sieg et al. [30], and Bajari and Kahn [31].

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