DISCUSSION PAPERS IN ECONOMICS

Working Paper No. 10-09

The Effect of Amenities on Local Wage Distributions

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November 2010

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November 5, 2010

Abstract

This paper imposes structure on the model presented in Roback (1988) by employing Elickson's (1971) single-crossing condition and predicts that local wage distributions will contract with an improvement in amenities. The range of estimated amenity-wage gradients across the wage distribution reveals the misleading nature of the average amenity-wage gradient, which is generally estimated in the literature. Workers at the lower end of the wage distribution are shown to earn more in locations with better amenities while those in the higher end are shown to earn less. In addition, both the implicit price paid for amenities and the implicit share of income spent on amenities are shown to increase substantially with wage level. The latter provides the first empirical evidence of an assumption that it commonly employed in urban models, namely, that amenities are luxury goods.

JEL classification: D31; J31; R13; Q51. Keywords: Amenities; Amenity Capitalization; Wage Distribution;

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1 Introduction

This paper imposes structure on the model presented in Roback (1988) by employing Elickson's (1971) single-crossing condition on worker utility. The prediction derived from the model is that local wage distributions will contract with better amenities. Using a difference-in-differences approach to identify the effect of amenities on wage quantiles, wage distributions are found to contract with an improvement in amenities as workers at the lower end of the wage distribution earn more while those in the higher end earn less. Since amenity-wage gradients are found to differ in sign and magnitude systematically with wage level, the average wage gradient generally estimated in the literature will be greatly misleading for most workers.

The simultaneous capitalization of amenity value into rents and wages was first modeled in the study of Roback (1982), which shows that rent and wage gradients are determined by the relative value of local amenities to firms and workers. Roback (1982) notes that the estimated wage gradient is an average of all workers in the sample, who do not have identical incomes and may have different tastes. Thus, the model is extended to allow for two types of workers in Roback (1988) and many types of workers with differing levels of education in Beeson (1991). Both authors state that when tastes and productivities are allowed to differ across workers, the signs and relative magnitudes of wage and rent gradients are ambiguous. Therefore, more structure must be imposed on the model to derive concrete predictions.

The model presented here incorporates a common assumption used in sorting models, the single-crossing assumption described in Elickson (1971), to impose structure on Roback's (1982) model and derive predictions on how amenities are capitalized into wages differently across workers. If preferences exhibit the single-crossing condition, then all price gradients may be signed conditional on the amenities' effect on productivity. In addition, if housing is a necessary good and amenities are luxury goods, then the wage gradient will be decreasing with wage whenever the rents are increasing with amenities. That is, high-wage workers will pay more of a wage penalty to reside in a location with better amenities than low-wage workers. In the empirical analysis presented in Section 3, rents are shown to be increasing with amenities, housing is shown to be a necessary good, and amenities are shown to be luxury goods. Thus, the three conditions of the model that are sufficient for wage distributions to contract with better amenities are shown to hold in the data.

Estimating wage gradients separately at different points in the wage distribution is important for at least three reasons. First, average wage gradients are misleading, especially for the workers whose actual wage gradients are of opposite sign to the average. Second, it allows us to observe amenity valuation separately by wage income, which may be more of interest to policy makers. For example, these estimates may be used to make more precise predictions on changes in the tax base due to a change in amenities, particularly when income taxes are progressive. Lastly, these estimates allow the estimated implicit price paid for amenities to vary across the wage distribution. The implicit price estimates in turn enable us to show that the share of income spent on amenities increases with wage level. This result confirms what has generally been assumed but never empirically shown, namely, that amenities are luxury goods.

The empirical results in Section 3 tie together previous results that separate workers by education (Black et al 2009) or occupation (Lee 2010). Black et al (2009) show that the return to a college education is lower in cities with higher housing price residuals, which the authors use as a measure for local amenity values. Lee (2020) uses city population as a proxy for amenities and estimates wage gradients that decrease with the skill of occupation within the medical sector.[1]

However, since both education and skill are positively correlated with wage income, it is not possible to tell if these results reflect a difference in preferences across education groups, occupations, or wage levels. Thus, the problem in considering previous approaches to separating workers in the data is that they do not explicitly ties wage gradients to relative wage levels. The advantage of the approach taken here is that wage gradients are allowed to vary across wage level *within* education groups and occupations. Therefore, the differences in

[1]He was unable to show the same pattern for workers outside of the medical sector. Here, wage distributions are shown to contract with amenities for multiple industry-occupation groups.

wage gradients reported here are due to differences in wage levels, not differences in education or occupation.

The majority of previous studies on amenity valuation have estimated price gradients from cross sections of micro level data. Since it is difficult to control for, let alone identify, all the local amenities that are capitalized into rents and wages, these studies almost surely suffer from omitted variable bias. For example, tropical storms and hurricanes are correlated with proximity to the ocean, which is itself an amenity. Thus, failure to control for ocean proximity will bias workers' value of tropical storms upward and, in the worst case scenario, make it appear that hazardous weather is an amenity. In the data used here, some omitted amenities such as ocean proximity, national monuments, etc do not vary across time. Therefore, I follow Bayer et al (2009) by using a difference-in-differences approach to alleviate the possible omitted variable bias present in cross-sectional estimates.

The paper is organized as follows. Section 2 briefly reviews the model presented in Roback (1988) appended to include the single-crossing condition of worker utility and derives the prediction on wage capitalization across workers. The empirical approach taken to test the prediction on wages and results are described in Section 3 and the data sets are described in Section 4. Results are presented in Section 5. Section 6 contains robustness checks and tests of the model's assumptions. Finally, Section 7 concludes.

2 Theoretical Model

Consider a continuum of locations, each with its own unique bundle of amenities and supply of land. Denote $\Phi_j \in [\underline{\Phi}, \overline{\Phi}]$ as the continuous amenity index that describes the overall level of amenities, which is exogenously set, in location j. It is assumed that every worker assigns the same subjective weights to the individual amenities. Since workers and firms do not explicitly pay for the amenities they consume, the value of these amenities is capitalized in rents and wages.

2.1 Workers

There are 2 types of workers indexed by $k \in \{A, B\}$. Workers differ across type with respect to how their labor input enters the production function. The reader may think of these types as being high-skilled and low-skilled respectively although it is not necessary for workers to differ in skill. The important aspect of worker differences is that they enter the production function as separate inputs that are imperfectly substitutable. Thus, their wages are allowed to differ within a location.

Each worker selects a single location in which to reside and work. A worker of type k residing in location j earns wage w_j^k and pays rent r_j. The worker selects a consumption basket of land l and a numeraire consumption good x to maximize his utility in location j and has an indirect utility function $V(r_j, w_j^k; \Phi_j)$.[2]

Migration is assumed to be costless and thus utility must be equated across location in equilibrium for each type of worker. Otherwise, some workers would have incentive to move. This condition is described by

$$V(r_j, w_j^A; \Phi_j) = V^A \tag{2.1}$$

for type A workers and symmetrically

$$V(r_j, w_j^B; \Phi_j) = V^B \tag{2.2}$$

for type B workers. Roback (1988) allows preferences to differ across worker types. Here, I find it useful and plausible to assume that preferences are identical but that a worker's willingness to accept higher rents in locations with better amenities increases with his wage similar to Elickson (1971).[3] That is, I assume workers only differ exogenously in their place in the production function. This may be thought of as assuming that workers have different preferences over occupations or differences in innate ability.

[2] It is assumed that the utility function of the workers is such that a positive amount of land l is always purchased.

[3] A utility function that satsfies this condition is provided in Elickson (1971) as the nested CES function $U(x, l; \Phi) = \{[a_1\Phi]^{\frac{1}{\rho}} + [(a_3X)^{\frac{1}{\omega}} + (a_2l)^{\frac{1}{\omega}}]^{\frac{\omega}{\rho}}\}^{\rho}$ where the a's, ρ, and ω are constants, $\rho = \frac{\sigma}{\sigma-1}$, and $\sigma < 1$.

2.2 Firms

The market for the consumption good is perfectly competitive and the good's price, which is taken as the numeraire, is set by global markets. Firms produce the consumption good according to a CRS production function $x_j = f(n_j^A, n_j^B, l_j; \Phi_{j9})$ using land and both types of labor, n_A and n_B, which are assumed to be imperfect substitutes. Both types of labor are necessary for production and amenities may effect the productivity of all inputs. In equilibrium, the unit cost function is equal to the price of the numeraire in equilibrium

$$C(w_j^A, w_j^B, r_j; \Phi_j) = 1. \tag{2.3}$$

Otherwise, firms would have incentive to enter or exit a particular market where unit cost is not equal to unity.

To solve for the effect of amenities on local prices, equations 2.1 to 2.3 are totally differentiated and expressed in percentage changes.[4] Following Roback's (1989) notation, C_p is defined as the quantity of the factor whose price is p necessary to produce one unit of output. The system of equations is

$$\theta_{w^A}\hat{w}^A + \theta_{w^B}\hat{w}^B + \theta_r\hat{r} + \eta\hat{\Phi} = 0 \tag{2.4}$$

$$\hat{w}^A - s_l^A\hat{r} + s_\Phi^A\hat{\Phi} = 0 \tag{2.5}$$

$$\hat{w}^B - s_l^B\hat{r} + s_\Phi^B\hat{\Phi} = 0, \tag{2.6}$$

where $\theta_p = C_p p$ is the cost share of factor price p,[5] hats denote percentage change $\hat{z} = \mathrm{d}z/z$, $\eta = \theta_{w^A}\eta_w^A + \theta_{w^B}\eta_w^B + \theta_r\eta_r$ is the share weighted sum of the effect of amenities on the productivity of each factor, and $\eta_p = C_{p\Phi}(\Phi/C_p)$ is the amenity elasticity of factor demand

[4] Roy's identity is used here to convert preference ratios into land consumption $-V_l^k/V_w^k = l_k$.

[5] Recall the unit cost of x is equal to one.

for the factor whose price is p.[6] Note that if an amenity is productive, it will lower unit costs of the firm ($\eta < 0$) and if it is unproductive, it will raise costs ($\eta > 0$).

Simultaneously solving equations 2.4 through 2.6 yields the factor price gradients in percentage changes.

$$\hat{w}^A/\hat{\Phi} = \Delta^{-1}[-s_\Phi^A(\theta_r + \theta_{w^B}s_l^B) + s_l^A(s_\Phi^B\theta_{w^B} - \eta)] \tag{2.7}$$

$$\hat{w}^B/\hat{\Phi} = \Delta^{-1}[-s_\Phi^B(\theta_r + \theta_{w^A}s_l^A) + s_l^B(s_\Phi^A\theta_{w^A} - \eta)] \tag{2.8}$$

$$\hat{r}/\hat{\Phi} = \Delta^{-1}(s_\Phi^A\theta_{w^A} + s_\Phi^B\theta_{w^B} - \eta) \tag{2.9}$$

where $\Delta = \theta_r + s_l^A\theta_{w^A} + s_l^B\theta_{w^B} > 0$ is the percentage of revenue from each unit of x that accrues to land.[7] As this term increases, the magnitude of all gradients decrease. That is, as more of the total revenue generated by production is soaked up by land costs, all prices are less able to respond to changes in amenities.

The first term in the numerator of equation 2.7 is more negative the larger the implicit budget share of amenities (s_Φ^A) and the larger is the amount spent on land by other agents ($\theta_r + \theta_{w^B}s_l^B$). The value of s_Φ^A captures the direct effect of type A's demand for amenities while that of $\theta_r + \theta_{w^B}s_l^B$ captures the fact that as more revenue from x production is directed toward land, less is available for type A's wages. The second term can only be signed if the productivity effect of amenities is known. So long as amenities are not unproductive ($\eta \leq 0$), the second term will be unambiguously positive and increasing with the budget share of land (s_l^A) and the amount of revenue from production that the other workers spend on amenities ($s_\Phi^B\theta_{w^B}$). The intuition here is that the more revenue is spent on land, the less willing workers will be to trade off amenities for wages and the more cost savings the firm enjoys from type

[6] To see this derivation, write $C(w^A, w^B, r; \Phi) = C_{w^A}(\Phi)w^A + C_{w^B}(\Phi)w^B + C_r(\Phi)r$ before taking the total derivative in equation 2.3.

[7] This is because the firm pays θ_r of each unit of revenue to land and θ_w^k to worker k, who spends s_l^k of θ_w^k on land.

B's willingness to accept lower wages for better amenities, the more revenue is available for type A's wages. Equation 2.8 is symmetric to equation 2.7.

All price gradients are increasing with the productivity of amenities. If amenities are productive, then more productive locations will have a higher demand for land *and* labor, increasing the price of both. The derivatives of these gradients with respect to parameters are difficult to sign without considerable added structure, which will not be attempted here.

2.3 Appended Model

It is not possible to sign equations 2.7 to 2.9 without imposing more structure on the model. The single-crossing assumption first described in Elickson (1971) imposes enough structure on the model to derive clear predictions on the signs and relative magnitudes of the wage gradients.[8] The testable prediction that arises from the appended model is that wage distribution will contract with better amenities.

Assumption 1 $\frac{V_\Phi}{V_r}$ *is decreasing in* w.

This assumption requires wealthier workers to be willing to pay a greater rent premium for an increase in amenities. Figure 1 displays a graphical interpretation of the assumption by plotting worker indifference curves in $\{\Phi, r\}$ space. Holding utility constant, the rents that a worker is willing to pay will be monotonically increasing with amenities according to the implicit function theorem.[9]

Consider the location with rent and amenity level defined by the crossing of the two indifference curves displayed in Figure 1. Assume without loss of generality that in the location under question $w^A > w^B$. According to Assumption 1, Type A workers will then be more willing to accept an increase in rents in exchange for better amenities for any amenity-rent point than type B workers. This will ensure that the workers' indifference

[8]This assumption is commonly employed in sorting models to achieve stratification of households by income. See Epple et al (1984) and (1993), Epple and Sieg (1999), Walsh (2007), and Banzhaf and Walsh (2008) for examples.

[9]Too see this, apply the implicit function theorem to $V(r_j, w_j^k; \Phi_j) = \overline{V}^k$ and note that $V_\Phi > 0$ and $V_r < 0$.

curve cross only once in $\{\Phi, r\}$ space and is depicted in Figure 1 by the steeper slope of type A's indifference curve.

Elickson (1971) first used the single-crossing assumption to explain why jurisdiction boundaries were so persistent and why wealthy jurisdictions were only willing to enter cooperative agreements on public good provision with those similar to them in wealth (Williams et al 1965). The single-crossing condition ties willingness to pay for local public goods or amenities to income, giving rise to income stratification similar to that found in club goods when the only difference across members is income (Jaramillo et al 2001). This condition has also found empirical support. Evidence of income stratification was observed as early as 1958 in Wood's (1958) *Suburbia* and Epple and Sieg (1999) provide an empirical test for the single-crossing condition using crime and school quality measures as amenities. Epple and Sieg (1999) find that the condition holds in 1980 Census data for the Boston Metropolitan Area. Furthermore, a direct test of the assumption is carried out here in Section 6.1.

In what follows, it is useful to reinterpret Assumption 1 in terms of budget shares of land and amenities. By suppressing location subscripts and assuming without loss of generality that $w^A > w^B$, Assumption 1 can be rewritten as

$$\frac{V_\Phi^A}{V_r^A} < \frac{V_\Phi^B}{V_r^B}.$$

Multiplying both sides by $(-\Phi/r)$ and applying Roy's identity yields

$$\frac{s_\Phi^A}{s_l^A} > \frac{s_\Phi^B}{s_l^B},$$

where $s_l^k = (rl_k/w_k)$ is the budget share of land and $s_\Phi^k = (V_\Phi/V_w^k)(\Phi/w^k)$ is interpreted as the implicit budget share of amenities for type k. Assumption 1 allows for the signing all price gradients conditional on amenities' effects on productivity as discussed below.

To completely characterize the possible price gradient signs when preferences satisfy the single-crossing condition, the values of η at which each price gradient switches sign are first solved for. From equation 2.7 it can be shown that $(\hat{w}^A/\hat{\Phi}) \lesseqqgtr 0$ as

$$\eta \gtreqqless s_\Phi^B \theta_{w^B} - (s_\Phi^A/s_l^A)(\theta_r + \theta_{w^B} s_l^B) \equiv \eta_{w^A}^*. \tag{2.10}$$

Likewise, from equation 2.8 it can be shown that $(\hat{w}^B/\hat{\Phi}) \lesseqgtr 0$ as

$$\eta \gtreqless s_\Phi^A \theta_{w^A} - (s_\Phi^B/s_l^B)(\theta_r + \theta_{w^A} s_l^A) \equiv \eta_{w^B}^*. \tag{2.11}$$

Finally, equation 2.9 implies that $\hat{r}/\hat{\Phi} \lesseqgtr 0$ as

$$\eta \gtreqless s_\Phi^A \theta_{w^A} + s_\Phi^B \theta_{w^B} \equiv \eta_r^*. \tag{2.12}$$

It is relatively straightforward to show that $\eta_{w^A}^* < \eta_{w^B}^* < \eta_r^*$ when preferences exhibit single-crossing.[10] However, additional information on preferences is required in order to determine the difference in wage gradients,

$$\left(\frac{\hat{w}^A}{\hat{\Phi}} - \frac{\hat{w}^B}{\hat{\Phi}}\right) = \Delta^{-1}(-\theta_r(s_\Phi^A - s_\Phi^B) - (1-\theta_r)s_l^B s_l^A[(s_\Phi^A/s_l^A) - (s_\Phi^B/s_l^B)] + (s_l^B - s_l^A)\eta).^{11} \tag{2.13}$$

Equation 2.13 reveals that knowledge of relative budget shares is required to determine the difference in wage gradients. Three possible cases of relative land and amenity shares exist under Assumption 1 but only the empirically plausible case, that when $s_\Phi^A > s_\Phi^B$ and $s_l^A < s_l^B$, is considered here. See Kerr (2011) for a theoretical treatment of the remaining two cases.

Figure 2 displays the relative values of all price gradients over the range η when amenities have a Hick's neutral effect on productivity.[12] Notice that it is possible for all prices to be increasing with amenities if amenities are sufficiently productive (η sufficiently negative). This would be the case if amenities are productive enough for the firms to value land more than both workers. Under this scenario, both types of workers must be compensated for rent premiums that are excessive according to their preferences.

For expositional purposes, imagine a situation where the Hick's neutral productivity effect of amenities changes in a particular location where all price gradients are positive and all else is held constant. As amenities become less productive (η increases), the profitability of firms and thus the demand for all inputs at that location decreases, putting downward

[10]See appendix for proof.

[11]Note that the magnitude of this difference is decreasing in Δ, the percentage of revenue from x production that accrues to land. A similar theoretical prediction in Black et al (2009) states that the more amenities are capitalized into rents, the less the return to education will be.

[12]Under single-crossing, $\eta_{w^B}^* \gtrless 0$ is also possible. All other thresholds ($\eta_{w^A}^*$, η_r^*, and $\tilde{\eta}$) are guaranteed to be the sign they are shown to be in Figure 2 and $\eta_{w^A}^* < \eta_{w^B}^* < \eta_r^*$ will always hold.

pressure on all prices. As rents fall, $\hat{w}^A/\hat{\Phi}$ is the first gradient to turn negative since type A's willingness to pay higher rents for amenities is more responsive to the change in amenity levels, followed by $\hat{w}^B/\hat{\Phi}$, whose responsiveness is not as great.

If amenities become unproductive enough (η sufficiently positive), the decrease in demand for land from the firms and workers will cause the rent gradient to become negative. In this situation, amenities are so unproductive that if they were to improve (Φ increase), the decrease in demand for land from lowered productivity and wages would outweigh the increase in demand from workers due to migration.

Finally, note that at a sufficiently positive value of η, the wage gradients cross and type B workers will suffer a greater wage penalty than type A workers. This occurs at $\tilde{\eta}$, which is defined as the value of η for which equation 2.13 equals zero. Intuitively, if amenities are sufficiently counterproductive so that rents are actually decreasing with amenities, type B workers, who are assumed to spend a greater portion of their income on land, benefit more from lower rents and thus be willing to give up a greater percentage of their income, rendering $\hat{w}^B/\hat{\Phi} < \hat{w}^A/\hat{\Phi} < 0$ possible.[13]

The prediction on the relative magnitudes of the wage gradients is added to the results, which are summarized by Figure 3. All results listed above the η axis hold under Assumption 1. The results shown below the axis are specific to the case where $s_\Phi^A > s_\Phi^B$ and $s_l^A < s_l^B$. Given the empirical support of this case in the literature and the evidence presented in Section 6.1, we expect equation 2.13 to be negative implying local wage distributions that contract with better amenities.

As can be seen from Figures 2 and 3 and derived in the appendix, type A's wage gradient will be less than type B's whenever the rent gradient is positive. Therefore, the three conditions which together are sufficient for the prediction of contracting wage distributions are that housing shares be decreasing with wage, amenity shares be increasing with wage, and rents be increasing with amenities.

Proposition 1 $\frac{\partial(\hat{w}/\hat{\Phi})}{\partial w} < 0$ *if* $\frac{\partial s_l}{\partial w} < 0$, $\frac{\partial s_\Phi}{\partial w} > 0$, *and* $(\hat{r}/\hat{\Phi}) > 0$.

[13] See appendix for proof.

The intuition behind Proposition 1 is as follows. Both worker types pay the same rent premium for amenities in any location yet type A workers are more willing to do so. Therefore, type B workers must be compensated for rent premiums that are excessive according to their willingness to pay by forgoing less wages than type A.

There exists a plethora of evidence that housing shares decrease with income and that rents increase with amenities.[14] However, to my knowledge there does not yet exist empirical evidence that amenity shares increase with income, or that amenities are luxury goods.[15] In Section 6.1, amenity shares are estimated across the wage distribution and found to increase with wage income. Decreasing housing shares and a positive rent gradient are also estimated. Thus, all conditions for wage distributions to contract with better amenities hold in the data.

3 Empirical Strategy

In this section, the prediction that local wage distributions contract with amenities is empirically tested. There are three empirical challenges to overcome in the process. First, the theoretical model does not provide a clear cutoff point that defines when a worker is high-wage and when he is low-wage. Second, the wage income observed in the data is top-coded, artificially restricting the variation in wages observed at the very top of the wage distribution. Lastly, estimates are likely to suffer omitted variable bias from unobserved amenities that are correlated with observed amenities. The manner in which each of these issues is dealt with is described below.

To avoid arbitrarily setting a cutoff for when a worker is considered high-wage or low-wage, quantile regressions are employed to analyze the impact of an increase in amenities across the distribution of wages (Koenker and Basset 1978). This not only sidesteps the possible selectivity bias in separating the sample on the dependent variable, but provides a

[14] See Mayo (1981) for a review of early studies and Hansen et al (1998) for evidence from Lorenz curve approach as well as the traditional approach. The general conclusion from this line of research is that income elasticities of housing is less than unity, even when elasticities are allowed to vary by income.

[15] Chen and Rosenthal (2008) find that the responsiveness in migratory behavior to amenities increases with education. Since income generally increases with education, the authors take this as suggestive evidence that amenities are a normal good, a condition which is necessary but not sufficient for amenities to be luxury goods.

richer description of amenities' effect on wages. The estimated coefficients on the amenity index are predicted to decrease with wage quantile. This would yield evidence that the low-wage workers are being penalized less (or compensated) for residing in high-amenity (high-rent) locations than high-wage workers.

A well-known problem with using Census data is that the income variable is top-coded to protect the privacy of individuals.[16] For this reason, coefficients are estimated via the censored quantile regressions first proposed by Powell (1984, 1986a) by using the simple algorithm described in Buckinsky (1994).[17] To determine the effect of the control variables on the q^{th} quantile, the censored quantile regression finds the vector $\beta^{(q)}$ and constant $\gamma^{(q)}$ that solves

$$\min_{\beta^{(q)},\gamma^{(q)}} \frac{1}{N} \sum_{ijt} \rho_{(q)}(w_{ijt} - \min\{w^0_{jt}, \mathbf{x}'_{ijt}\beta^{(q)} + \gamma^{(q)}\Phi_{jt}\}) \tag{3.1}$$

where $\rho_{(q)}(\lambda) \equiv (q - I(\lambda < 0))\lambda$ is the tilted absolute value function or the "check function", w_{ijt} is the natural log of wage income, w^0_{jt} is the censoring value, and $\mathbf{x}_{ijt}$ is a vector of wage controls.[18] Subscripts i, j, and t denote individual, location, and time respectively. Location here is defined as the Metropolitan Statistical Area (MSA). The minimization problem is then iteratively resolved using only observations whose estimated conditional quantile $\hat{w}^{(q)}_{ijt}$ is less than the censoring value w^0_{jt} until convergence is reached.

This procedure is applied to the 10^{th}, 25^{th}, 50^{th}, 75^{th}, and 90^{th} wage quantiles. The coefficients of a quantile regression are interpreted analogously to those of a least-squares regression. For example, $\hat{\gamma}^{(10)}$ is interpreted as the effect of a one unit increase in the amenity index on the 10^{th} quantile of the wage distribution. If local wage distributions contract with amenities, we would expect to see $\gamma^{(q)} < \gamma^{(q')}$ for all $q > q'$.

Empirical studies that estimate quantile regressions over the distribution of wages generally estimate bootstrapped standard errors from 100 repetitions (Buchinsky 1994; Machado

[16] 1.98% of the workers in the sample used here are top-coded.

[17] An alternative estimator is suggested in Buchinsky and Hahn (1998) but was shown to take more than twice the amount of computing time to estimate. Given the large sample size used here, this estimator was not used.

[18] The censoring values differ across states and years in the Census data. Thus, it is written with a time and location subscript.

and Mata 2005; Poterba and Rueben 1994[19]). Due to the large data set used in this study, it is not practical to do so here. Therefore, I estimate bootstrapped standard errors from 100 repetitions only for the coefficients estimated from regressions where censoring is an issue. For the remaining coefficients, the estimated standard errors are those reported by STATA, which are estimated using the method suggested by Koenker and Bassett (1982) with the density of the residuals at zero estimated by the method described in Rogers (1993). Finally, for comparison with previous literature, the effect of amenities on average wages is also estimated using a least squares regression.[20]

To alleviate omitted variable bias present in cross-section estimates, I follow Bayer et al (2009) in using a difference-in-differences approach. The coefficients on the amenity index are identified off its variation across time while controlling for other MSA factors via fixed effects. If the coefficients of interest were estimated using only a cross section of data, the amenity index and the MSA fixed effects would have a one-to-one relationship, prompting us to use only the amenity index. The index would then be correlated with any MSA-specific factor, which would likely bias the estimates.

However, by using MSA fixed effects and variation of amenity scores across time, we are able to alleviate this potential bias. The estimated relationship for each conditional quantile is

$$w_{ijt}^{(q)} = \mathbf{x}_{ijt}'\beta^{(q)} + \gamma^{(q)}\Phi_{jt} + \varphi_j^{(q)}\text{MSA}_j + \epsilon_{ijt}^{(q)}, \tag{3.2}$$

where x_{ijt} contains a year indicator and the difference between the wage of observation 2 in location 1 in year 2 and observation 1 in location 1 in year 1 identifies $\gamma^{(q)}$,

$$w_{212}^{(q)} - w_{111}^{(q)} = (\mathbf{x}_{212} - \mathbf{x}_{111})'\beta^{(q)} + \gamma^{(q)}(\Phi_{12} - \Phi_{11}) + \epsilon_{212}^{(q)} - \epsilon_{111}^{(q)}. \tag{3.3}$$

The MSA terms drop out here as both observations reside in the same location.

[19]Poterba and Rueben (1994) estimated both bootstrapped and analytical standard errors and found the differences between them to be small. The same result is found here where using either standard error would not change the confidence level with which any of the coefficients are estimated.

[20]The symmetrically censored least squares estimate is more appropriate here as a measure of the average effect of amenities on wages. However, in this particular case, the censoring did not pose an issue and thus the estimation reduces to OLS.

Only unobservables that are fixed over time will be controlled for by the MSA indicators. Any MSA-specific effect that is variable across time and correlated with the index may still effect the estimates. In particular, there may be unmeasured productivity shocks whose change across time is correlated with the change in the amenity index. These omitted variables could bias the results towards showing wage distributions that contract with the amenity index if the change in the omitted productivity shocks disproportionately increase wages at the bottom of the wage distribution or disproportionately decrease wages at the top of the distribution.

The coefficients on the amenity index are estimated on the full sample as well as separately by aggregate industry-occupation group as defined by the Census.[21] While Lee (2010) was unable to find wage gradients that decreased with income for occupations outside of the medical sector, the methodology employed here is able to do so.

4 Data

Data for this study come primarily from the US Census 1990 and 2000 5% microdata samples (available at usa.ipums.org). From this sample I use full-time[22] non-hispanic white males that claim to be the "head of household", are not students or in the military, and are between the ages of 25 and 55 to ensure that location decisions were driven by those that are not looking for locations in which to retire and may be expected to have amenities capitalized in their wages.[23] Furthermore, only individuals that could be identified with an MSA for which amenity values were available were kept in the sample.[24] This left me with 689,714 observations in 266 MSA's for 1990 and 764,585 observations in 293 MSA's for 2000.

[21]Aggregate ocupations used here are Managerial and Professional, Technical-Sales-Administrative, and Precision Production-Craft-Repairers. Aggregate industries used are Construction, Manufacturing, Retail Trade, Professional and related services.

[22]A worker is designated full-time if he reported working at least 35 hours a week and 48 weeks a year.

[23]Graves and Waldman (1991) find that retirees sort to locations with low rates of wage capitalization.

[24]Some MSA definitions in the Census data differed slightly from those in the *Places Rated Almenac*, which is used to construct the amenity index here. Those found to have differences were dropped from the sample.

In the wage regressions, each individual's wage and salary income is used as the dependent variable. The control variables are total hours worked in the last year (calculated from average hours worked in a week and total weeks worked), sex, age, potential experience,[25] categorical dummies for education, the ability to speak English, work disabilities, marital status, MSA of residence, and a year dummy. Squared terms for age and potential experience are also included.

In Section 6.1, rental rates and housing values are regressed on housing characteristics and the amenity index to test the model's assumption on the rent gradient. The dependent variable used is reported rental rates for renters and the reported value of the house for owners. Descriptive variables include the housing unit's age, acreage, categorical variables for type of heating used, number of other units physically attached, indicators for trailers and units that are not houses (ex. boats), existence of complete plumbing, number of bedrooms, phone availability, MSA location, and indicators for having a kitchen and ownership of the unit.[26]

The amenity data is obtained from the *Places Rated Almanac*, a semi-regular publication of location specific amenity measures used to construct rankings of metropolitan areas. MSA's in the United States (and Canada) are scored in several categories and then ranked according to their overall score. For the amenity index Φ_{jt}, the scores of all categories that did not include data on wages or housing costs were obtained from the *Places Rated Almanac* 1989 and 2000 editions: climate, crime, arts, recreation, healthcare, education, and transportation.

In the 1989 edition, *Places Rated* describes the methodology of scoring each category and lists both the scores and the ranks of MSA's for each category and overall. However, in the 2000 edition, only percentiles are reported for each category. For example, the MSA

[25] Potential experience was calculated by $exp = age - agework$ where $agework$ is the youngest age at which an individual could have been employed. Most papers use $agework$ = years in school + 5 However, this yields $agework < 16$ for individuals who graduated early, which was not legal in the US for the time period under study. For these cases, I assume $agework = 16$.

[26] This indicator is used in place of multiplying housing values by a constant to convert them to monthly rents. In this way, I let the data decide the most appropriate adjustment rather than enforce an arbitrary value.

ranked first in an amenity category has a score (percentile) of 100 for that category. For this reason, once all MSA's that did not match the Census sample were dropped, I recalculated the percentiles of each category separately for both years and used these in place of the scores. The amenity index for any MSA in a particular year is then constructed as the sum of scores across all categories.

Although the data sources are mostly consistent across years, one caveat to the *Places Rated* data is that the methodology to scoring the MSA's for each category is altered slightly across years. These alterations range from small to relatively significant. For example, the crime category scoring methodology is identical across years but uses 5 year averages in 1989 and 8 year averages in 2000. However, the climate category's scoring is altered more significantly. In the 2000 edition, information is added to the scoring procedure (hazardous weather measures, seasonal effects, etc) that was not included in the 1989 edition.

To address the above listed data issues, three strategies were implemented. First, the quantile wage regressions were estimated with the individual categories of amenities taking place of the amenity index. In this way, the pattern of wage gradients may be observed by category and if a single category is deemed to be flawed in measurement, it need not affect the others. This approach also allows the categories to have unequal weights in worker preferences opposed to the amenity index. Second, the wage regressions were estimated seven additional times, removing a single category form the index one at a time. Lastly, the effect of an improvement in the precision of the amenity measure on the estimates is tested using a Monte Carlo expirement.

Results from these tests are discussed in Section 6.2 and no evidence is found of the data artificially forcing the results presented here. Note that in order for the measurement error of the amenity index to *cause* the results, it must be that the change in measurement across years favors the amenity measures of MSA's whose wage distributions contracted relatively more than the others over the sample time period. No reason is found here to believe that this is occurring.

Descriptive statistics for the change in amenity measures are presented in Table 1 and a geographic distribution of the change in the amenity index across time is displayed in Figure 4. The bins in Figure 4 were constructed so that each contain 20% of the MSA's. Correlations of the change in amenities across time are presented in Table 2. The changes in category percentiles are mostly uncorrelated as the largest of these, the correlation of the change in recreation and the change in health care, takes a value of 0.263.

5 Results

5.1 Amenities' Effect Across the Wage Distribution

The results from the quantile wage regressions are presented in Table 3. In addition to the full sample results, results from quantile regressions run on aggregate industry-occupation groups with more than 75,000 total observations are also displayed. The coefficients are scaled up so that they may be interpreted as the percentage change in wages due to a 100 point increase in the amenity index (this is equivalent to increasing a single amenity category from worst to best). Notice the significantly positive coefficients on all of the lower quantiles. This is evidence that low wage workers are compensated with higher wages in the presence of better amenities. Likewise, negative signs on the coefficients for the higher quantiles is evidence of high-wage workers trading off wages for amenities.

The results confirm that wage distributions contract with amenities. All groups have coefficients that are positive for the lowest quantile and generally monotonically decrease as quantile increases. Only three coefficients, that from the 75^{th} quantile regression for administrators and sales representatives in retail (row 4) and managers in professional services (row 5) and that from the median regression for managers in manufacturing (row 6), do not follow the monotonically decreasing pattern. However, these differences are not likely significant.[27] In all but one group, administrators and sales representatives in retail (row 4), the coefficient eventually becomes negative as quantile increases. The lack of negative

[27] I am unable to statistically test that the coefficients are monotonically decreasing with quantile. To do so requires simultaneous estimation of all censored quantile regressions. Due to the length of computing time necessary for such an estimation, the approach was not taken here.

coefficients for this group indicates that these workers are likely not near the top of their local wage distributions and thus do not earn enough to give up wages for better amenities.

The results in Table 3 conform to those in Black et al (2009) and Lee (2010). The comparable regression in Black et al (2009) estimates the effect of a housing price index that proxies for amenity levels on the returns to education, measured by the ratio of income of college graduates to high-school graduates. The authors estimate a negative relationship, implying that the spread of income decreases as housing prices (amenities) increase. Lee (2010) assumes city size is correlated with consumption variety, which is treated as a consumption amenity. The author estimates city size-wage gradients that decrease with the skill of occupation for workers in the medical sector. In this paper, a more direct approach is taken by using actual measures of amenities and estimating their effect across the wage distribution while controlling for education and occupation (as well as industry). Furthermore, here we are able to observe contracting wage distributions for multiple industry/occupation groups outside of the medical sector.

5.2 Average and Median Wage Gradients

To compare the methodology used here to the previous literature and illustrate the importance of the former (particularly when equity is valued), the average wage gradient is estimated by OLS to compare to the quantile regression results. The standard errors are clustered by MSA-year and the results are presented in Table 4. The median regression from Table 3 is also presented for a separate measure of central tendency.

The coefficients from both regressions are positive but not significant. Using either regressions to test the model in Roback (1982) with one type of worker, the insignificant coefficients would imply that amenities have no effect on wages. This may be true for the worker with the average or median wage but is certainly not true for all workers as observed by the coefficients in Table 3. When taken as representative of all workers, the OLS and median regression results in Table 4 overestimate the wage gradient for high-wage workers and understate that of low-wage workers.

5.3 Estimated Implicit Prices Across the Wage Distribution

Since workers do not explicitly pay for local amenities, they will pay for them implicitly through rent premiums and forgone wages. The implicit price of an amenity is the additional income spent on land plus the forgone wages due to the presence of the amenity. Implicit prices are often constructed from average wage and rent gradients to value local amenities, environmental goods, and public goods and/or to measure the quality of life. If wage gradients differ across workers, then the implicit price paid for local amenities will likely differ as well. If these differences are great, the standard approach of estimating mean implicit prices may not be useful, particularly when equity is valued or when a particular income group is the subject of concern. The implicit price paid for amenities can be derived From equations 2.1 and 2.2 for each worker as

$$P^k \equiv \left(\frac{V_\Phi^k}{V_w^k}\right) = l^k \frac{\mathrm{d}r}{\mathrm{d}\Phi} - \frac{\mathrm{d}w_k}{\mathrm{d}\Phi} \tag{5.1}$$

or in percentage terms as

$$(P^k/w^k) = s_l^k \frac{\mathrm{d}\log r}{\mathrm{d}\Phi} - \frac{\mathrm{d}\log w^k}{\mathrm{d}\Phi}. \tag{5.2}$$

Recall that all workers are assumed to have the same rent gradient in each location but are allowed to have different wage gradients and budget shares for land. Here each worker earning the q^{th} quantile wage is treated as a separate worker type and the budget share for housing is substituted for the budget share for land.

To compare the implicit shares of income spent on amenities, we must hold constant both the amenity index and the price of housing that each quantile worker faces so that only wage differs across the workers. Thus, this analysis must be carried out at the MSA level rather than the national level where housing prices and the amenity index will differ across the workers.

The implicit price is estimated for each quantile for Chicago, the MSA with the largest number of observations.[28] An estimate of the share of income devoted to housing is used as a proxy for the share of income devoted to land. This is estimated for both years and for each quantile by finding the q^{th} quantile wage earner and taking his housing share value as the estimate.[29] The rent and wage gradients are taken from Table 3. The pattern of implicit prices across wage quantile are qualitatively the same across years. To conserve space, only those reports for the year 2000 are displayed here in the first row of Table 5.

The numbers reported in the first row of Table 5 are interpreted as the percentage of income a worker is willing to pay for a 100 point increase in the amenity index (equal to an improvement from worst to best in a single category). For example, the worker whose wage is equal to the 90^{th} wage quantile in Chicago is willing to give up approximately 2.3 percent of his income for a 100 point increase in the amenity index. The monotonicity in the implicit prices mimics that observed in the coefficients on the amenity index in the wage regressions; larger quantiles pay more for amenities. Interestingly, some workers have negative estimated implicit prices. This implies that their wage gradient is positive and dominating the rent gradient in their implicit price. That is, they must be compensated above and beyond what is necessary for them to purchase an equal amount of housing when amenities improve.

This estimated implicit prices demonstrate the importance of thinking beyond mean valuations for amenities when equity is of concern. The estimates vary greatly over the wage distribution of workers in Chicago.[30] These differences hold important implications for policy makers. For example, this suggests that policy makers that subsidize low-income housing in cities with high rent premiums may be sacrificing more efficiency than is necessary since low-income workers are compensated with wage premiums. However, this is only suggestive.

[28]Implicit prices were also estimated for all industry/occupation groups with qualitatively similar results and not presented here for brevity.

[29]Monthly rent and yearly wage divided by 12 was used to calculate the housing share. If the observation(s) owned their housing unit, its value is converted into monthly rent by subtracting the estimated coefficient on ownership from the censored median regression on rents/home values. If more than one worker earned wages equal to the q^{th} quantile wage earner, the median of their housing shares was used.

[30]Implicit prices are also shown to vary within the wage distributions of aggregate industry/occupation groups. These results are not shown here but available upon request.

A full analysis would entail formally modeling local governments that take into consideration the migratory responses to changes in local amenities and prices.[31]

6 Assumption and Robustness Checks

6.1 Assumption Checks

The three assumptions of the model that lead to the prediction of contracting wage distributions are that rents increase with amenities, land shares are decreasing with wages, and implicit amenity shares are increasing with wages. Each of these assumptions are tested below and found to hold in the data.

To test the requirement that rents are increasing with amenities, a symmetrically censored least squares (SCLS) regression (Powell 1986b) and a censored median regression of the rental rate or housing values on housing characteristics and amenities are estimated.[32] The SCLS estimation equation is

$$P_{ijt} = \min\{(P_{jt}^{0,own}), \mathbf{h}'_{ijt}\beta_h + \gamma_h \Phi_{jt} + \epsilon_{ijt}\}, \tag{6.1}$$

where P_{ijt} is the natural log of monthly rental rate (reported value of unit if owned), $P_{jt}^{0,own}$ is the censoring value conditional on ownership status,[33] $\mathbf{h}_{ijt}$ is a vector of housing unit characteristics, and Φ_{jt} is the amenity index. The algorithm for SCLS "recensores" the data from below to restore symmetry in the errors and estimates OLS coefficients on observations with predicted values below the top-code. The analogous conditional median is estimated from a censored median regression and the coefficient is also reported.

We expect to see positive coefficients from both the SCLS and the censored median regressions on rents and housing values. Since this result is standard in the literature, the sign and significance of the coefficients serve as a test of how noisy of a measure the amenity index is. If the signs of these coefficients are not positive or significant, then the amenity

[31] The model presented here does not consider taxes, local government, or publicly provided goods. For a model that deals with such issues in a Roback (1982) framework with one type of worker, see Albouy (2009)

[32] 2.5% of owners and 4.0% of renters are top-coded in the sample.

[33] Rents and housing values have separate top-codes.

measure may not be precise enough to accurately measure amenities' effect on the wage distribution. The results for rents are presented in Table 6. Reported standard errors are those estimated from the last iteration of either algorithm clustered by MSA-year.[34]

The signs of both coefficients are positive and significant with 99% confidence. This is evidence that the amenity index constructed here is sufficient for empirically measuring local amenities. Mean and median rental rates and housing values would increase by approximately 3.7% and 4% respectively with a 100 point increase in the amenity index.

The estimated implicit prices can be used to check the assumption that the share of wages spent on amenities is increasing with wage ($\frac{\partial s_\Phi}{\partial w} > 0$). The assumption that land shares decrease with wage ($\frac{\partial s_l}{\partial w} < 0$) is also verified by the data. Table 5 displays the estimates of housing and amenity shares as well as their ratio for each quantile of the full sample using estimated shares for the year 2000. The same pattern is observed for each industry-occupation group and across years but not presented here for brevity. Housing shares are decreasing with income and amenity shares are increasing with income. Thus, the assumptions presented in Section 2 for the prediction on relative wage gradients to hold are validated by the data. Furthermore, the increasing shares of wages spent on amenities constitute the first empirical evidence that amenities are luxury goods.

6.2 Robustness Checks

One potential issue with the analysis above is that one cannot tell how much variation across time is due to the change in scoring methods and how much is due to actual variation in amenities. Although weather patterns do change over time, the variation in the climate category displayed in Table 1 may seem too high to be strictly a result of actual variation in climate conditions. Some outliers are due to the addition of hazardous conditions in the 2000 edition. As a robustness check, I removed each category one at a time from the amenity index and reran the quantile regressions for the industry/occupation groups. My analytical results were qualitatively the same across all specifications.[35]

[34]Bootstrapped standard errors were not estimated due to practical reasons.
[35]These results are not presented here but available upon request.

To allow amenities to have differential effects on wage quantiles, the amenity index was split into its composite categories and the empirical analysis was repeated. To conserve space, only the estimated coefficients on the individual amenity categories for the largest group, Managers in the Professional services industry, are presented in Table 7. For the most part, the pattern of decreasing coefficients holds. The biggest exception is the crime category, for which the pattern did not hold for any group. This may be due to crime's differential effect across wage level. Those with high wages that live in high-crime cities may reside in neighborhoods with private security and/or removed from the troubled neighborhoods. If this is the case, it violates the assumption that all amenities are viewed the same by all workers and we would not expect the results to hold for this category.

Finally, to test if refinements of the measurements used in The Places Rated Almanac across time could be driving my statistical results, a Monte Carlo experiment was conducted using simulated data with an amenity measure that improved in precision over time. A total of 10,000 simulations were run using a model where amenities were constructed to effect all wage quantiles the same.[36] If the increase in precision of amenity measure is forcing the results, we would expect to see a pattern in the coefficients similar to the results presented here. The experiment did not suggest any bias of the amenity coefficients in either direction. In particular, estimated coefficients were not systematically decreasing with quantile. Details are located in the appendix.

7 Conclusion

This paper imposed Ellickson's (1971) single-crossing condition on Roback's (1989) model of amenity capitalization, resulting in the prediction that local wage distributions will contract with better amenities. The assumption is not only theoretically popular in both intra-urban and inter-urban models, but it finds empirical support in Epple and Sieg (1999). Wage distributions were shown to contract empirically as high-wage workers were found to earn less, and low-wage workers were found to earn more in locations with better amenities.

[36] A constant positive, negative, and zero effect of amenities on wages were all tested.

Three empirical challenges were encountered in measuring the effect of amenities on wage distributions: deciding which workers were high-wage and which were low-wage, the presence of top-coded wages, and omitted variable bias. Quantile regressions were employed to estimate amenities' effect on wage distributions and sidestep the issue of choosing a high-wage/low-wage cutoff point. The problem of top-coded wages were dealt with by estimating censored quantile regressions using Buchinsky's (1994) algorithm. Finally, The omitted variables bias was alleviated by using a difference-in-differences strategy.

The assumptions of the model that are sufficient for wage gradients to decrease with wage level (ie wage distributions to contract) as amenities improve were shown to be satisfied in the data. Rents were found to increase with amenities, housing shares were found to decrease with wages, and amenity shares were found to increase with amenities. To my knowledge, this last result serves as the first empirical evidence that amenities are indeed luxury goods as many have assumed.

The results of the paper tied together previous studies whose results may also imply that local wage distributions contract with better amenities. The results of Black et al (2010) indicate that the returns to education decrease with amenities while Lee (2010) finds similar results with the return to skill in the medical industry. Since skill and education are both positively correlated with wage levels, the theory and results presented here, which control for education and skill via industry-occupation controls, suggest that previous results may be largely due to the manner in which amenities are capitalized across wage levels.

Although local governments are not formally modeled here, the results suggest interesting tax implications. If households are willing to pay more to live in higher amenity locations, we might expect a local government to be able to extract more taxes from the residents. However, it may be possible for an improvement in amenities to decrease tax revenue. The contraction of the wage distribution may result in lower income tax revenue overall, especially if taxes are progressive. This may or may not be offset by an increase in property tax revenue. To solve this problem, a local government must be formally introduced to the model. This is left for the topic of future research.

Table 1: Descriptive Statistics of the Change in Amenity Scores

Category	Mean	Standard Deviation	Min	Max
ΔClimate	-0.27	31.43	-58.41	86.81
ΔHealthcare	-0.01	18.92	-61.72	52.34
ΔCrime	0.59	14.64	-50.20	54.33
ΔTransportation	0.23	24.82	-82.08	65.49
ΔEducation	0.50	21.77	-65.47	66.17
ΔArt	0.24	18.17	-69.90	56.68
ΔRecreation	0.05	23.92	-81.71	76.65

Table 2: Correlation of Amenity Changes Across Time

Variables	Δhealth	Δcrime	Δtransport	Δedu	Δart	Δrec	Δclimate
Δhealth	1.000						
Δcrime	-0.092	1.000					
Δtransport	0.162	-0.123	1.000				
Δedu	0.059	-0.127	0.068	1.000			
Δart	0.059	0.046	-0.014	0.254	1.000		
Δrec	0.263	-0.055	0.151	0.014	0.021	1.000	
Δclimate	0.041	-0.057	0.158	-0.150	-0.009	-0.111	1.000

Table 3: Coefficients on the Amenity Index

Dependent Variable: ln(wage)	quantile					
Group	90	75	50	25	10	N
IND: Manufacturing OCC: Labor	-0.233	0.335	1.128***	2.141***	2.714***	88,994
	(0.67)	(0.483)	(0.539)	(0.64)	(0.898)	
IND: Manufactruing OCC: Production	-2.016***	-0.809***	-0.33	0.285	1.551	82,208
	(0.71)	(0.482)	(0.493)	(0.687)	(1.026)	
IND:Construction OCC: Production	-1.294	0.216	1.607***	2.644***	3.585***	75,600
	(0.795)	(0.648)	(0.643)	(0.815)	(1.186)	
IND: Retail OCC: Admin	0.157	2.168***	0.897***	2.404***	4.01***	77,909
	$(1.380)_a$	(0.883)	(0.742)	(0.836)	(1.205)	
IND: Professional OCC: Management	-4.214***	-1.215	-1.297***	0.197	1.425	156,145
	$(1.466)^a$	$(0.924)^a$	(0.579)	(0.638)	(1.022)	
IND: Manufacturing OCC: Management	-2.097***	0.0128	0.617	0.401	2.633***	115,822
	$(1.202)^a$	$(0.593)^a$	(0.428)	(0.524)	(0.912)	
Full Sample	-1.711***	-0.440**	0.097	1.173***	2.780***	1,437,957
	$(0.284)^b$	$(0.183)^b$	(0.164)	(0.199)	(0.294)	

standard errors are in parentheses

a bootstrapped standard errors

b analytical standard errors from last iteration of censored quantile regression

Table 4: OLS and Median Wage Regressions

Regression	Amenity Index Coefficient	N
OLS	0.472 (0.560)	1,454,299
Median	0.097 (0.164)[a]	1,454,299

standard errors are in parentheses

a analytical standard errors reported by STATA in the last iteration

Table 5: Housing Shares, Amenity Shares, and Implicit Prices in Chicago (1999)

	quantile				
	90	75	50	25	10
Implicit Price of Amenity Index ($P^{(q)}/w^{(q)}$)	2.326	1.058	0.652	-0.281	-1.590
Amenity Share (s_Φ^q)	0.983	0.465	0.296	-0.131	-0.765
Housing Share (s_l^q)	0.124	0.125	0.151	0.180	0.240
Ratio (s_Φ^q/s_l^q)	7.911	3.723	1.957	-0.729	-3.183

Implicit Price is percent of wage income implicitly paid for a 100 unit increase in Φ_j

Table 6: SCLS and Censored Median Housing Regressions

Regression	Amenity Index Coefficient	N
SCLS	3.738*** (1.252)	1,454,299
Median Regression	4.046*** (0.000)	1,454,299

standard errors are in parentheses.

Table 7: Amenity Coefficients for IND: Professional OCC: Management

	quantile				
Amenity	90	75	50	25	10
Climate	-7.344**	-2.055	-2.177*	1.055	3.531*
	(3.125)[a]	(1.494)[a]	(1.192)	(1.295)	(1.999)
Health Care	-3.137	2.88	-0.448	0.959	0.317
	(5.006)[a]	(2.411)[a]	(1.929)	(2.096)	(3.244)
Education	-11.525 **	-6.84 **	-2.811	0.594	5.353
	(5.851)[a]	(2.835)[a]	(2.261)	(2.453)	(3.806)
Recreation	0.107	0.948	3.13*	3.921**	5.137*
	(4.604)[a]	(2.223)[a]	(1.777)	(1.927)	(2.979)
Arts	5.893	5.652	0.712	1.477	-5.701
	(7.367)[a]	(3.633)[a]	(2.916)	(3.170)	(4.893)
Transport	-5.159	-6.038***	-5.347***	-5.537***	-1.839
	(4.319)[a]	(2.081)[a]	(1.664)	(1.808)	(2.787)
Crime	1.176**	-1.946***	-1.747	-2.498	-5.606
	(7.088)[a]	(3.44)[a]	(2.753)	(3.002)	(4.649)

standard errors are in parentheses

a analytical standard errors from last iteration

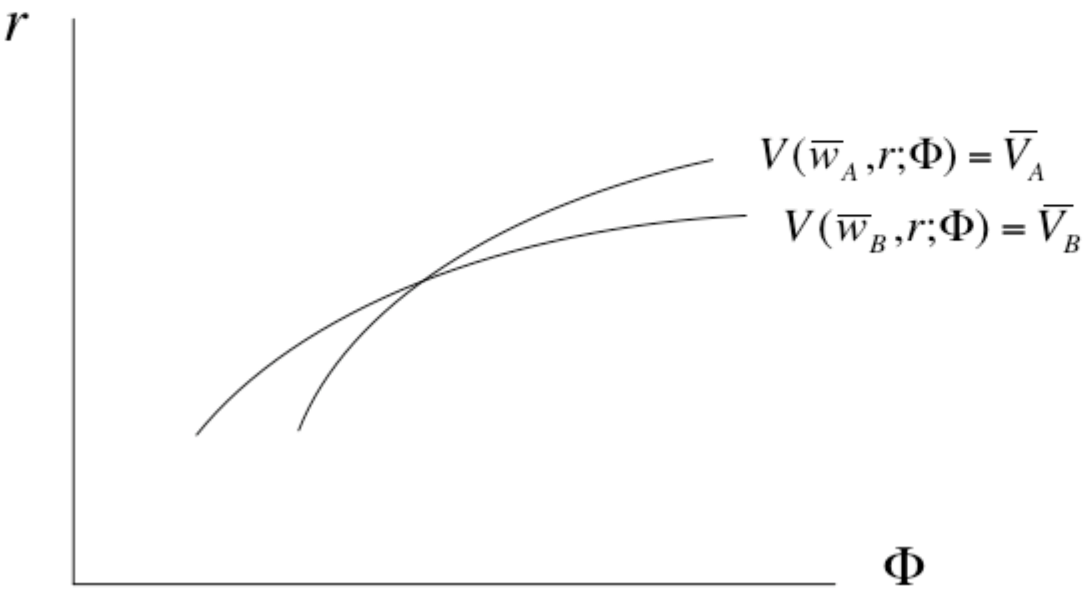

Figure 1: Single-Crossing Condition

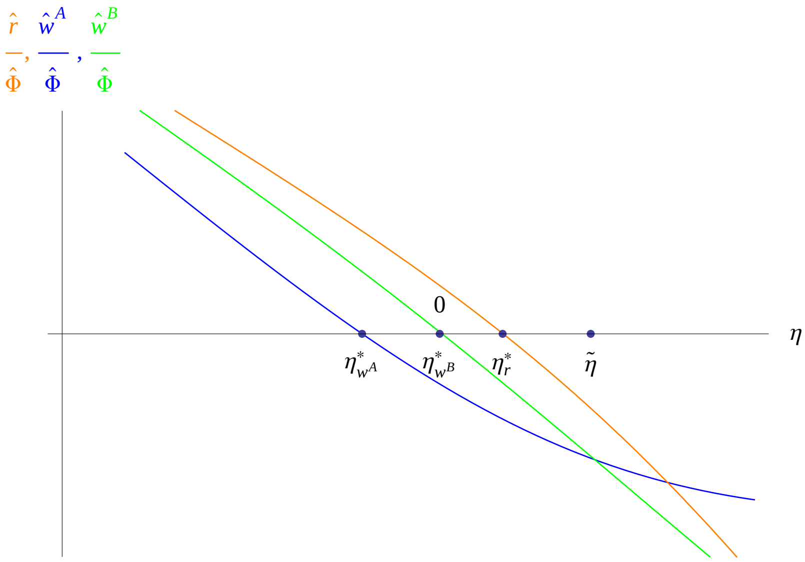

Figure 2: Price Gradients as Functions of a Hick's Neutral Change in η

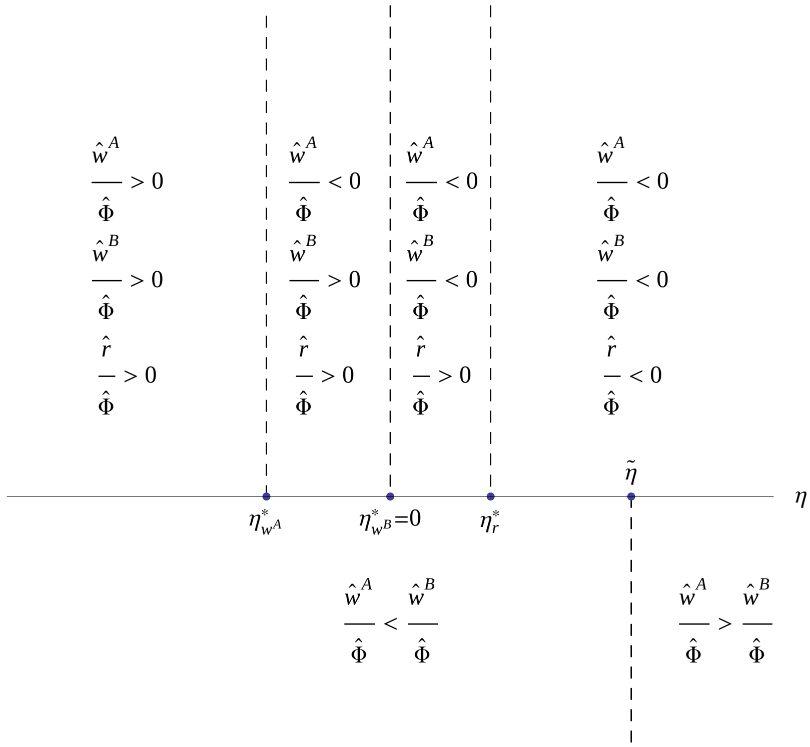

Figure 3: Summary of Gradient Signs and Relative Magnitudes

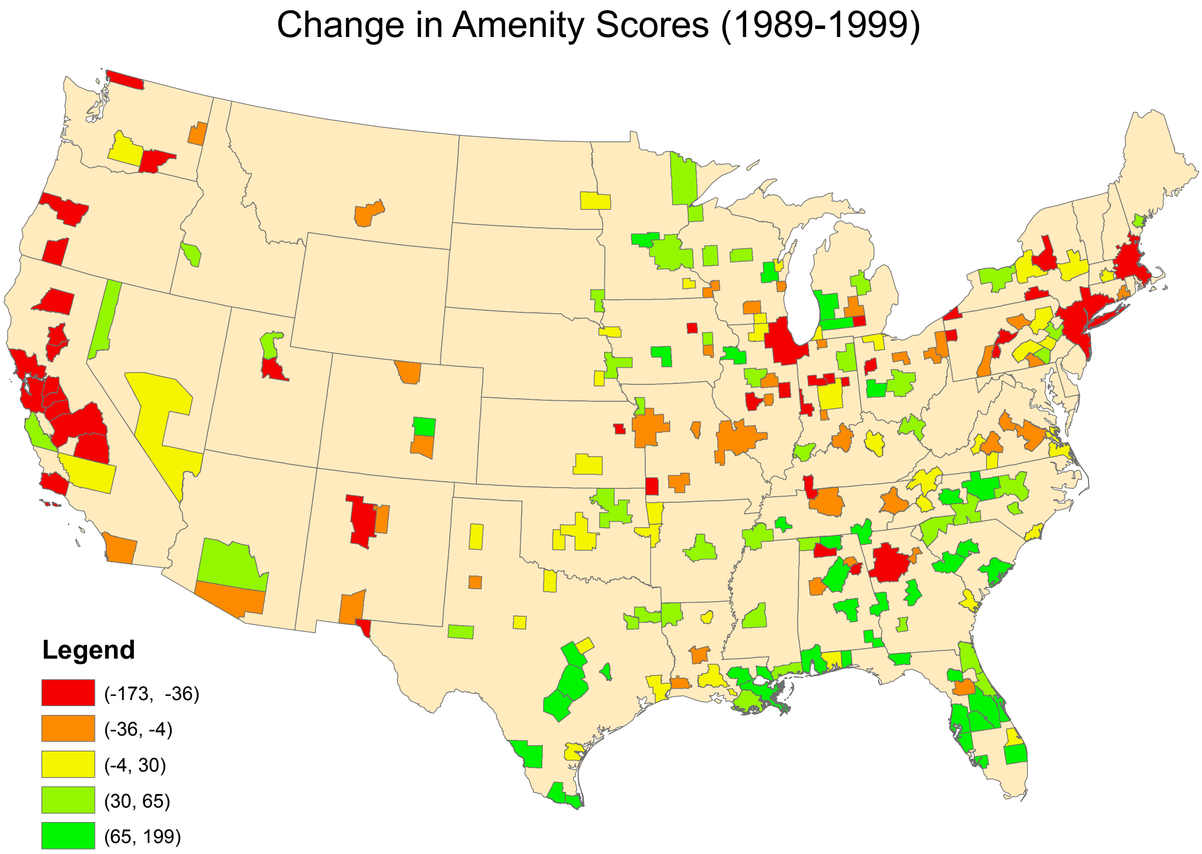

Figure 4: Change in Amenity Index (1999-1989)

A Appendix

Proof that $\eta^*_{w^A} < \eta^*_{w^B} < \eta^*_r$

$$\eta^*_{w^A} < \eta^*_{w^B}$$

when

$$0 < (\theta_r + \theta_{w^A} s_l^A + \theta_{w^B} s_l^B)[(s_\Phi^A/s_l^A) - (s_\Phi^B/s_l^B)].$$

The single-crossing assumption guarantees that the right hand side is positive. Likewise,

$$\eta^*_{w^B} < \eta^*_r$$

when

$$0 < (s_\Phi^B/s_l^B)(s_l^B\theta_{w^A} + s_l^A\theta_{w^B} + \theta_r),$$

where all terms on the ride hand side are positive. By transitivity, the proof is complete.

Proof that $(\hat{w}^A/\hat{\Phi}) - (\hat{w}^B/\hat{\Phi}) < 0$ **when** $s_l^A < s_l^B$**,** $s_\Phi^B < s_\Phi^A$ **and** $\eta < \tilde{\eta}$**.**

Signing equation 2.13 is requires solving for

$$-\theta_r(s_\Phi^A - s_\Phi^B) - (1-\theta_r)s_l^B s_l^A[(s_\Phi^A/s_l^A) - (s_\Phi^B/s_l^B)] + (s_l^B - s_l^A)\eta \lesseqgtr 0$$

or

$$\tilde{\eta} \equiv \frac{\theta_r(s_\Phi^A - s_\Phi^B) + (1-\theta_r)(s_\Phi^A s_l^B - s_\Phi^B s_l^A)}{s_l^B - s_l^A} \gtreqless \eta$$

Therfore, $\tilde{\eta}$ is the value of η for which $(\hat{w}^A/\hat{\Phi}) = (\hat{w}^B/\hat{\Phi})$. What remains is to show where this value is relative to the η_p's for $p \in \{A, B, l\}$.

Here, the case where $s_\Phi^A > s_\Phi^B$ and $s_l^A < s_l^B$ is considered. In this case, $\eta^*_r < \tilde{\eta}$. This inequality holds when

$$s_\Phi^A\theta_{w^A} + s_\Phi^B\theta_{w^B} < \frac{\theta_r(s_\Phi^A - s_\Phi^B) + (1-\theta_r)(s_\Phi^A s_l^B - s_\Phi^B s_l^A)}{s_l^B - s_l^A},$$

which can be rearranged as

$$0 < \Delta(s_\Phi^A - s_\Phi^B).$$

Clearly, this holds when $s_\Phi^A > s_\Phi^B$. The assumption $s_l^A < s_l^B$ dictates the direction of the inequality. Thus, under the assumptinos $s_l^A < s_l^B$ and $s_\Phi^B < s_\Phi^A$, $(\hat{w}^A/\hat{\Phi}) - (\hat{w}^B/\hat{\Phi}) < 0$ when $\eta < \eta^*_r$.

A.1 Monte Carlo Experiment

To test whether the additional information used in the 1999 categories of the amenity index has a bias that makes it more likely to find conditional quantile coefficients that decrease with quantile, a monte carlo experiment was carried out. First, 100 draws of x were taken where $x \sim U(0,1)$. The dependent variable y was constructed such that

$$y = a + bx + e \tag{A.1}$$

for half the observations and

$$y = a + b(x + u) + e, \tag{A.2}$$

for the other half, where $e \sim N(0,1)$ and $u \sim N(0,1)$. The u was applied to half of the observations to model roughly half of the observations in the data set used in Section 3 that had amenity measures from 1989, which utilized less information and are assumed to be noisier measures of the true amenity levels compared to the measures of 1999.

Quantile regressions were then run for quantiles 10, 25, 50, 75, and 90. The process was repeated 1,000 times for all $b \in \{-5, 0, 5\}$. The estimated coefficients for the higher quantile regressions were not observed to be systematically smaller that those of the lower quantiles. More specifically, $\hat{b}^q > \hat{b}^{q'}$ for $q < q'$was found to hold approximatley 50% of the time while $\hat{b}^q < \hat{b}^{q'}$ approximately 50% of the time across all simulations suggesting that the effect was purely random in the simulation. These differences were very small and it is likely that the hypothesis that $\hat{b}^q = \hat{b}^{q'}$ could not be rejected in any of the trials. Thus, no evidence was found that a more precise x variable would force estimated quantile coefficients to be decreasing with quantile.

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