Green Price Indices*

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Abstract: This paper suggests two theoretically consistent and empirically tractable ways that a cost-of-living index can be expanded to include the environment and other public goods. In addition, it presents an empirical illustration of such an index for Los Angeles, California, incorporating air quality and other spatially varying public goods using a hedonic model. The results indicate that the required information can be recovered and that including public goods can make a substantial difference in the index.

Key Words: air quality, green accounting, hedonic regression, nonmarket valuation, price index.

JEL Classification Numbers: E31, H40, I00, Q25, R10.

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Introduction

“Green” national income accounts that reflect the quality of the environment and other public goods have received increasing interest internationally. Japan, Germany, and Sweden currently compute some form of green economic account, and the U.S. National Academy of Sciences has endorsed the concept (Nordhaus and Kokkelenberg 1999).[1] Less attention has been given to similarly adjusting national price indices. And yet, green price indices would be necessary for a dual price deflator consistent with Green GDP. More importantly, greening the indices would bring them closer to their own objective. Since the indices are defined as the relative cost of achieving a fixed standard of living, insofar as public goods affect this standard of living they have a theoretical place in the index. In one call for such a program, Nordhaus (1999) has advocated an “augmented cost-of-living index” but only nets out the cost of producing such

[1] In 1993, the United Nations endorsed a system of Environmental and Economic Accounting as a framework for satellite national accounts. At about the same time, the U.S. Bureau of Economic Analysis (BEA) began tabulating such an index (U.S. BEA 1994) but was directed by Congress to discontinue the work. For general discussion and examples, see Lange (2003), Lutz (1993) and Repetto et al. (1989).

goods in his illustration.[2] Others have discussed adjustments based on measures of benefits, but have not been precise or consistent about the point at which benefits would be measured.

This paper suggests two concrete ways that a green cost-of-living index might be constructed. Both suggestions clarify the welfare information required for the context of price indices, which work to date has not noticed is different than that usually used in benefit-cost applications. The paper also provides examples for households in Los Angeles from 1989 to 1994. Air quality, an important public good in the region, dramatically improved over the period covered, but education quality, measured by teachers per student, first improved and then declined. Accounting for these changes reduces the cost of living in the first part of the period, but has offsetting effects in the second part. The average annual effect overall is to lower the cost of living by 0.1 percentage points per year, a level consistent with recent estimates of the bias from quality change in market goods (Boskin et al. 1996, Lebow and Rudd 2003).

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[2] The notion is also discussed in Boskin et al. (1996) and Schultze and Mackie (2002). While being ambivalent toward public goods, in general the Boskin commission endorsed other ways that the U.S. Consumer Price Index (CPI) could be brought closer to the theoretical ideal of a cost-of-living index, including better allowance for substitution and adjustments for quality change in market goods. The latter parallels the kinds of adjustments that could be made for public goods. In contrast, the Schultze committee has been less sanguine about the entire interpretation of price indices as cost-of-living indices, but still confirmed the usefulness of an exploratory green satellite account.

Theory of Cost-of-Living Indices and Estimation Strategy

Background

Following the classic definition of Konüs [1924] (1939), a cost-of-living index is defined as the ratio of money expenditures needed in two scenarios to hold a representative household's utility constant. Abstracting for now from public goods, let utility in period t be given as $u^t = u(\mathbf{x}^t)$, where $\mathbf{x}^t$ is a vector of consumption goods. The household maximizes utility subject to prices and its income, giving rise to the indirect utility function $v(\boldsymbol{p}_x,y)$. Let $m(\boldsymbol{p}_x,u)$ denote the minimal cost of achieving a given utility level at a given vector of prices. Finally, denote period a as the reference period and b as the comparison period. Then the cost-of-living index at reference-period utility is

$$\mathrm{I}^{ab}\left(\boldsymbol{p}^b, \boldsymbol{p}^a, u^a\right) = \frac{m(\boldsymbol{p}_x^b, u^a)}{m(\boldsymbol{p}_x^a, u^a)}. \tag{1}$$

It measures the proportionate increase (or decrease) in expenditures required to maintain the utility level of the reference period at new prices. We may extend this concept to include a vector of public goods $\bar{\boldsymbol{q}}$ by using the conditional expenditure function $\tilde{m}(\boldsymbol{p}_x^t, \bar{\boldsymbol{q}}^t, u^a)$, in which the levels of the public goods are held constant in the expenditure minimization problem.

Even ignoring public goods, the cost-of-living index cannot be computed in practice because of its reliance on unobserved information about preferences. However, Konüs showed that the following empirically tractable compromises bound a true cost-of-living index:

$$\mathscr{L}^{ab} \equiv \frac{\sum_k p_k^b x_k^a}{\sum_k p_k^a x_k^a} = \sum_k \frac{p_k^b}{p_k^a} w_k^a \geq I^{ab}(\boldsymbol{p}^b, \boldsymbol{p}^a, u^a) \tag{2}$$

$$\mathscr{P}^{ab} \equiv \frac{\sum_k p_k^b x_k^b}{\sum_k p_k^a x_k^b} = \sum_k \frac{p_k^b}{p_k^a} w_k^b \le I^{ab}(\boldsymbol{p}^b, \boldsymbol{p}^a, u^b), \tag{3}$$

where w_k^t is the expenditure share for good *k* computed at period *t* quantities and reference prices. The indices are known as the Laspeyres and Paasche indices respectively; they are an average of the relative prices of each commodity, weighted by their share of total expenditures in either the reference or the comparison scenario. The Laspeyres index is an upper bound on the true cost-of-living index at reference utility; the Paasche index is a lower bound on the true cost-of-living index at comparison utility. Because it incorporates historical data, the Laspeyres index is used in most national price statistics.

Green Cost of Living Indices

The question is how to introduce public goods into these indices. The first way to do so, applicable to the widest variety of circumstances, is to add additional commodities to the Laspeyres or Paasche indices and evaluate them at their *implicit* or *virtual* prices (Neary and Roberts 1980). Virtual prices, first suggested for price indices during World War II as a way to account for rationing (Rothbart 1940-1), are the price at which consumers would choose to consume the quantity of goods that they are forced to consume by rationing when income is also adjusted to cover these expenditures. Virtual prices are equally applicable to public goods, since, like rationed goods, public good levels are not freely chosen by individual households.

The Marshallian virtual price vector $\bar{\boldsymbol{p}}_q$ associated with a level of these public goods $\bar{\boldsymbol{q}}$ is defined implicitly as

$$\bar{\boldsymbol{q}} = \arg\max\ u(\boldsymbol{x},\boldsymbol{q}) \quad s.t.\ \ y + \bar{\boldsymbol{p}}_q \cdot \bar{\boldsymbol{q}} \ge \boldsymbol{p}_x \cdot \boldsymbol{x} + \bar{\boldsymbol{p}}_q \cdot \boldsymbol{q} \tag{4}$$

Virtual prices are a function of the prices of other goods, the rationing level, and the budget constraint. They are the price at which households would (just) demand the rationed quantity if they were free to choose its level. Accordingly, the virtual price can be interpreted as the marginal willingness to pay for the rationed good.

Because it treats changes in public goods as changes in (virtual) prices, this approach is readily consistent with standard formulae, such as the Laspeyres index, that are functions of prices. To work from this perspective, we need only introduce into the index additional goods that are evaluated at their virtual prices and weighted at hypothetical expenditures—just like normal market goods, but with an adjustment to total expenditures to account for the hypothetical expenditures. Borrowing a term from Nordhaus (1999), we might call this an "augmented cost-of-living index," since it adds new terms into the standard price index.

More specifically, by analogy to the usual Laspeyres argument, the following bound holds:

$$\frac{\left(\boldsymbol{p}_x^b \cdot \boldsymbol{x}^a + \overline{\boldsymbol{p}}_q^b \cdot \overline{\boldsymbol{q}}^a\right)}{\left(\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a + \overline{\boldsymbol{p}}_q^a \cdot \overline{\boldsymbol{q}}^a\right)} \geq \frac{m\left(\boldsymbol{p}_x^b, \overline{\boldsymbol{p}}_q^b, u^a\right)}{m\left(\boldsymbol{p}_x^a, \overline{\boldsymbol{p}}_q^a, u^a\right)}. \tag{5}$$

This is the usual Laspeyres upper bound, with additional terms for the virtual prices. The expression can be re-written in the form

$$\frac{\left(\boldsymbol{p}_x^b \cdot \boldsymbol{x}^a + \overline{\boldsymbol{p}}_q^b \cdot \overline{\boldsymbol{q}}^a\right) - \overline{\boldsymbol{p}}_q^b \cdot \overline{\boldsymbol{q}}^b}{\left(\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a + \overline{\boldsymbol{p}}_q^a \cdot \overline{\boldsymbol{q}}^a\right) - \overline{\boldsymbol{p}}_q^a \cdot \overline{\boldsymbol{q}}^a} \geq \frac{m\left(\boldsymbol{p}_x^b, \overline{\boldsymbol{p}}_q^b, u^a\right) - \overline{\boldsymbol{p}}_q^b \cdot \overline{\boldsymbol{q}}^b}{m\left(\boldsymbol{p}_x^a, \overline{\boldsymbol{p}}_q^a, u^a\right) - \overline{\boldsymbol{p}}_q^a \cdot \overline{\boldsymbol{q}}^a}, \tag{6}$$

which shows the Laspeyres bound with an explicit adjustment for the difference between "virtual income" and actual income.

Simply rearranging and canceling some terms, this equation can be rewritten further as[3]

$$\frac{(\boldsymbol{p}_x^b \cdot \boldsymbol{x}^a) - \bar{\boldsymbol{p}}_q^b \cdot \Delta \boldsymbol{q}}{(\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a)} \geq \frac{\tilde{m}(\boldsymbol{p}_x^b, \boldsymbol{q}^b, u^a)}{\tilde{m}(\boldsymbol{p}_x^a, \boldsymbol{q}^a, u^a)}. \tag{7}$$

This is a Laspeyres index for market goods, with an adjustment in the numerator equal to the marginal willingness to pay for the public goods times the change in their levels. In one sense, it is a natural extension of the usual Laspeyres concept to public goods: it converts the welfare information of quantity changes into price changes and proceeds with the usual definition of the price index.[4]

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[3] To derive equation (6) from (5), multiply both sides by $(\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a + \bar{\boldsymbol{p}}_q^a \cdot \bar{\boldsymbol{q}}^a)/\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a$ and, for the right-hand side, use $\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a + \bar{\boldsymbol{p}}_q^a \cdot \bar{\boldsymbol{q}}^a = m(\bar{\boldsymbol{p}}_q^a, \boldsymbol{p}_x^a, u^a)$. To derive (7), use $\tilde{m}(\boldsymbol{p}_x^t, \bar{\boldsymbol{q}}^t, u^a) = m(\boldsymbol{p}_x^t, \bar{\boldsymbol{p}}_q^t, u^a) - \bar{\boldsymbol{p}}_q^t \cdot \bar{\boldsymbol{q}}^t$.

[4] This approach loosely mirrors the quality-of-life index of Blomquist, Berger, and Hoehn (1988), essentially a quantity index weighted with marginal prices instead of a price index, and the constant-climate consumption index of Cragg and Kahn (1999). However, those papers did not attempt to incorporate the public goods into indices for market goods and so were not precise about the point at which to evaluate marginal willingness to pay.

For some public goods, these implicit prices may be obtained from non-market valuation tools based on revealed preferences, such as hedonic price regressions, random utility models, averting behavior methods, or sorting models. Values for other, more pure public goods may be obtainable only from survey-based methods. Regardless of the valuation method used, this approach has the advantage of requiring only marginal values rather than the total values that frequently impose greater measurement difficulties. Unfortunately, the marginal values would have to be evaluated at an unusual point from an empirical perspective. As suggested by the constraint in Equation (4), or by the numerator of the right side of Equation (6), $\bar{p}_q^b$ *is evaluated at the prices and the level of the public good in the comparison period but when income is hypothetically adjusted to cover the virtual expenditures*. Since this point is never observed, it can only be estimated with nonmarket valuation techniques that are capable of recovering an entire willingness-to-pay function, thus undermining this approach's advantage of using only marginal values.

An alternative approach is more restrictive in the circumstances under which it is applicable but also overcomes the latter difficulty. This approach models public goods as a *weak complement* to a non-essential market good. Weak complements are goods that are enjoyed only when their associated complements are consumed in positive quantities. For example, air quality in community *j* is enjoyed only when living in community *j*. Similarly, quality characteristics of market goods are usually weak complements: the number of bedrooms in a house, for example, has no value unless one lives in the house. In this way, spatially delineated public goods may be modeled as additional qualitative differences among houses or apartments.[5]

The restrictiveness of weak complementarity should be acknowledged at the outset. For example, under this assumption a wetland would be valued only to the extent that it is tied to observable activities, such as outdoor recreation; the existence values of the ecosystem or the individual species are overlooked. Nevertheless, weak complementarity has two advantages. First, it allows preference information about the public goods to be recovered from observable activity in the linked market. Second, as shown by Willig (1978), weak complementarity allows an exact adjustment for the public good to enter the index via the preexisting Laspeyres or Paasche subindex for the linked market good. In particular, the price p^* that compensates the household for forgoing the quality change when income is already adjusted to obtain reference utility can be substituted for the market good's p^b in the Laspeyres index. An analogous p^{**} can be defined for the Paasche index.

Consider the case of spatially delineated public goods, such as air quality, crime, and education, which are weak complements to housing in a given location. Denote p_h as the price of housing and $\mathbf{q}_h$ as a vector of characteristics measuring the quality of housing and weakly complementary public goods. Define p^* and p^{**} implicitly as

$$v(p_h^*, \boldsymbol{q}_h^a, \boldsymbol{p}_x^b, m(p_h^b, \boldsymbol{q}_h^b, \boldsymbol{p}_x^b; u^a)) = v(p_h^a, \boldsymbol{q}_h^a, \boldsymbol{p}_x^a, y^a) \tag{8}$$

[5] For more on weak complementarity, see Bockstael and McConnell (1993), Palmquist (2002), Smith and Banzhaf (2003), and Willig (1978).

and

$$v(p_h^{**}, \boldsymbol{q}_h^b, \boldsymbol{p}_x^a, m(p_h^a, \boldsymbol{q}_h^a, \boldsymbol{p}_x^a; u^b)) = v(p_h^b, \boldsymbol{q}_h^b, \boldsymbol{p}_x^b, y^b)\,. \tag{9}$$

These are the price adjustments, in lieu of the more familiar income adjustments, that compensate for quality changes. Note that the price adjustments are defined at a point where income is adjusted to maintain utility at the appropriate level for each index.

Using p^* and p^{**}, it can be shown that the following bounds for the Laspeyres and Paasche indices still hold (Willig 1978):

$$\mathscr{L}^{*ab} \equiv \frac{p_h^* h^a + \boldsymbol{p}_x^b \cdot \boldsymbol{x}^a}{p_h^a h^a + \boldsymbol{p}_x^a \cdot \boldsymbol{x}^a} \geq I^{ab}(\boldsymbol{p}^b, \boldsymbol{p}^a, \boldsymbol{q}^b, \boldsymbol{q}^a, u^a) \tag{10}$$

$$\mathscr{P}^{**ab} \equiv \frac{p_h^b h^b + \boldsymbol{p}_x^b \cdot \boldsymbol{x}^b}{p_h^{**} h^b + \boldsymbol{p}_x^a \cdot \boldsymbol{x}^b} \leq I^{ab}(\boldsymbol{p}^b, \boldsymbol{p}^a, \boldsymbol{q}^b, \boldsymbol{q}^a, u^b) \tag{11}$$

$\mathscr{L}^*$ is the Laspeyres index with p^* replacing p^b, while $\mathscr{P}^{**}$ is the Paasche index with p^{**} replacing p^a. They are the usual Laspeyres and Paasche concepts with a subindex defined in cost-of-living terms replacing the usual price relative for the good with quality change. In this case, the adjusted subindex may be considered a group index of housing and its weakly complementary public goods. Since it uses an adjustment to prices of an existing component (here, housing) instead of adding components, I call this approach an “adjusted cost-of-living index.”

Both the augmented and the adjusted approaches to the green indices start with the Könus bounds to a true cost-of-living index. This framework clarifies the precise welfare information required for constructing an index consistent practices, particularly the income level where prices must be evaluated. This point that has been missed so far in discussions of green index numbers (e.g. Nordhaus 1999, Blomquist, Berger, and Hoehn 1988) as well as in related work on adjusted cost-of-living indices for market goods (e.g. Nevo 2003), which proceed from the measures commonly used for benefit-cost analysis.

Hedonic Estimation

In this paper, I illustrate both approaches to incorporating public goods, using as an example a cost-of-living index for Los Angeles (LA), California, for the years 1989–1994, when an important public good of concern, air quality, underwent substantial improvement. To recover the information required for these indices, I estimate hedonic regressions of housing prices that adjust for air quality and other public goods.

In the hedonic framework, prices of quality-differentiated goods are modeled as a continuous function of the underlying characteristics, yielding the relationship $p_j = p(\boldsymbol{q}_j)$, which can be estimated with standard regression techniques. Hedonic models have a long history of being used to recover marginal values for environmental and other public goods, as required for the first index, with air pollution being a particularly common application.[6] Similarly, they have a long history in quality-adjusted price indices for market goods, as required for the second

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6 See Palmquist (2003) for an overview of this context. Smith and Huang (1996) give a metaanalysis and bibliography of applications to air quality.

index, and in the United States are now used to adjust the indices for computers, televisions, and apparel.[7]

For the augmented cost-of-living index, a hedonic price regression can be used to estimate the marginal values of public goods, which enter the index as shown in Equations (5)-(7). Because the hedonic equilibrium conditions obtain at baseline conditions, these marginal values ignore the income effects discussed above. Thus, the results presented here are only an approximation, valid to the extent that income effects are small. For the adjusted cost-of-living index, hedonic price regressions can be used to estimate quality-adjusted price indices. Under certain conditions, these hedonic indices are consistent with the definitions of p^* or p^{**} in the adjusted cost-of-living index (Equations 10 and 11).[8]

Such hedonic price indices come in two primary variations. The first variation, often called the "direct" hedonic price index, decomposes the hedonic price equation into two pieces, a quantity index based on characteristics and a temporal (or spatial) price index capturing shifts in the quantity index. For example, a direct hedonic price index for two periods, *a* and *b*, could be estimated with a hedonic equation of the form

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[7] Griliches (1961) is the classic reference. Other examples include Gatzlaff and Ling (1994), Liegey (1994), and Raff and Trajtenberg (1997).

[8] See Fisher and Shell (1972), Muellbauer (1974), and Trajtenberg (1990) for a discussion of these conditions. Banzhaf (2002) offers an alternative estimation approach that relaxes them, using discrete-choice random utility models.

$$\ln(p_j) = \alpha^a D_j^a + \alpha^b D_j^b + f(\boldsymbol{q}_j) + \varepsilon_j \qquad (12)$$

where D_j^a and D_j^b are indicator variables for the reference and comparison years, respectively (taking a value of one for the year the house is sold and zero the other year) and the α's are the associated intercepts. Taking the exponent of both sides, it is clear that $e^{\alpha^b}/e^{\alpha^a}$ is a price index corresponding to the quantity index $e^{f(\,)}$. In this approach, omitting pubic goods from $f(\,)$ causes a standard omitted variable bias in the fixed effects. Equation (12) may be estimated by assuming $f(\,)$ remains constant for all time periods, or by chaining the index so that it is constant for any two adjacent years. Intuitively, if public goods are positively (negatively) correlated with time, omitting them should bias the estimated index upward (downward).

The direct hedonic index has the advantage of a simple interpretation, but restricts the shape of the hedonic function to remain constant over at least two time periods, even when chained. The second, less structural, variation uses period-by-period hedonic regression to impute what the price of a good would have been if its quality had remained constant. Then, the predicted price can be used in a price relative for the commodity and entered into a price index in the natural way. For example, suppose there is a house j with characteristics $\boldsymbol{q}_j^a$ in period a. Some of its characteristics such as the quality of local public goods and, of course, its age change in period b so that we designate its characteristics $\boldsymbol{q}_j^b$. The predicted prices from hedonic regressions can be used to impute the price relative for the house as if its quality had remained constant. For example, a Laspeyres index would hold the bundle constant at reference period qualities and use the ratio $\hat{p}^b(q_j^a)/\hat{p}^a(q_j^a)$, where $\hat{p}^t$ is the predicted value from the regression in period t. A Paasche index would hold the bundle constant at comparison period qualities and use the ratio $\hat{p}^b(q_j^b)/\hat{p}^a(q_j^b)$. These individual price relatives can then be averaged and entered into a price index in the same way as other goods. With this imputation approach, the effect of omitting public goods appears in errors in the predictions.

To be consistent with an annual cost-of-living index, the housing "prices" in the above discussion should properly be annual user costs or rental rates. For the Consumer Price Index (CPI), the Bureau of Labor Statistics (BLS) currently imputes the user cost for owner-occupied housing by matching houses to observed rents at similar units. This approach has the advantage of allowing the market to determine the annualization rate but the disadvantage of failing to directly observe the owner-occupied market.[9] In the empirical illustration given below, I instead explicitly compute user cost at time t as

$$R^t = (i^t + \tau^t + \delta - \pi^t)\cdot P^t \tag{13}$$

where R^t is the user cost, P^t is the asset price, i^t is the rate of return forgone by holding the house, τ^t is the property tax rate, δ is the depreciation rate, and π^t is expected asset appreciation at time t.

Following Poterba (1992), I assume that δ=0.04 and that the opportunity cost of holding the housing asset is the prevailing 30-year conventional fixed-rate mortgage for each year, obtained from the Federal Home Mortgage Corporation, plus a risk premium of 0.04. Based on work by O'Sullivan et al. (1995), I assume a constant effective property tax rate of τ=0.0055.

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[9] Perhaps for this reason, the LA CPI subindex for owner-occupied housing shows continuing inflation over the period covered, in contrast to empirical work showing marked deflation led by the highest tier of housing (Case and Shiller 1994).

Finally, I assume that the expected asset appreciation is a 5-year moving average of realized appreciation, with the final year forward-looking:

$$\pi^t = \frac{1}{5}\sum_{s=t-4}^{t}\frac{P^{s+1}}{P^s} \tag{14}$$

That is, at time t expected changes in asset price are an average of the actual change that occurs from time t to t+1 and the changes in the previous four years. Hedonic price regressions on the asset prices determine P^{t+1}/P^t.[10] The final annualization factors range from 12 to 16% over the period. This approach has the advantages of using owner-occupied data for the index of owner-occupied housing but the disadvantage of being sensitive to the particular assumptions used, especially about the discount rate (see Gillingham 1983). Nevertheless, the relative difference made by public goods in this illustration is not affected by this assumption. Moreover, the hedonic methodology used below is equally applicable to either asset prices or rental prices.

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[10] For the purposes of computing user cost, π^t is estimated from the imputation hedonic models reported below but does *not* hold the quality of public goods constant. Rather, it is the forecasted joint effect on asset values of both prices and quality changes. That is, it is the estimate of $p^b(\boldsymbol{q}^b)/p^a(\boldsymbol{q}^a)$, not $p^b(\boldsymbol{q}^a)/p^a(\boldsymbol{q}^a)$. Thus, like a change in prices, an exogenous change in quality of the house is a wealth transfer that alters the opportunity cost of owning. Expected future increases in quality, inasmuch as they increase future asset values, decrease the opportunity cost of home ownership.

Both versions of the adjusted cost-of-living index, as well as the augmented cost-of-living index, are illustrated in this paper. The application requires information on housing prices and attributes over time, and public goods over time. These data are discussed in the following section.

Data

This study uses a large set of actual housing prices and quality characteristics for Los Angeles from 1989 to 1994 obtained from Transamerica Intellitech, a market research firm. After deleting about 10% of the observations as outliers or because of inconsistent values, there were 319,641 observations remaining for analysis, each for a unique house. The data indicate that unadjusted housing prices began falling after 1991, an observation consistent with previous analysis of the LA real estate market and with the overall recession in the LA economy during that period (Case and Shiller 1994). Additional variables include the size of the lot in acres, the area of the house in square feet, the number of bathrooms and bedrooms, the presence of a fireplace, the presence of a swimming pool, and the age of the house. Table 1 summarizes the means of these variables by county in the study area. In addition, the location of each house in latitude and longitude is known, allowing them to be linked to separate data on locational public goods.

As noted previously, for the sake of this illustration the “environment” is represented by air quality, the medium of most concern in LA and the one that saw the most rapid change over the period. Air quality in turn is represented by ozone. This is the pollutant most commonly in violation of California's air quality standards and the one that receives the most attention in public health communications.[11]

Air quality data were obtained from the California Air Resources Board.[12] These data are some of the richest in the world: for ozone, an average of 50 monitors are available each year in the study area, plus monitors in neighboring counties that can aid in interpolation. Because epidemiology and toxicology studies have consistently found that acute episodes of high ozone concentrations cause the most damages, I measure ozone with the expected number of exceedences of the U.S. one-hour standard. This measure has the additional advantage of coinciding with the information communicated to residents in the form of ozone alerts (e.g., on the *LA Times* weather page). The mean pollution values by county are also included in Table 1.

Air quality, or even the environment generally, is not the only public good that might belong in cost-of-living indices. School quality, crime, and proximity to the coast are the

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[11] Particulate matter is also an important pollutant and one generally found to have the most important health effects. The results in this paper are qualitatively unchanged if the annual average of particulate matter under 10 microns (PM_{10}) is used instead of ozone. However, because of their high correlation, including both variables typically leads to a switch of sign in one.

[12] Information on these data can be obtained from the California Air Resources Board's Web page at www.arb.ca.gov/aqd/aqd.htm. Information is also available from the U.S. Environmental Protection Agency at www.epa.gov/air/data.

locational variables most likely to matter to households. Proximity to the coast was available from the housing location data, and enters through indicator variables for being within 1 mile and 3 miles from the coast. Such static locational variables will of course not affect the cost-of-living index over time, though it would affect locational comparisons.

Proxy variables for school quality were obtained from the National Center for Education Statistics. The main variable used is the teacher-student ratio, which was obtained for each school district and year.[13] In addition, achievement test scores (the sum of math and reading scores from the California Learning Assessment System test) were obtained for 1993 and appear as a cross-sectional control, but do not affect the index over time. Crime rates per 10,000 people were obtained at local jurisdictions from the California Department of Justice in 1990. For other years, only county-level crime rates were available, so the annual percent changes at the county level were applied to each jurisdiction within that county starting from its 1990 baseline. Finally, since the population in a neighborhood may also affect housing prices or proxy for omitted characteristics, demographic data for local neighborhoods (on average, 800 residents per neighborhood) were obtained from the U.S. Census. The means of all these variables are also reported in Table 1.

The average levels of the three main public goods (air quality, teacher-student ratios, and public safety), indexed to 1989, are shown in Figure 1 expressed as goods. The figure indicates that air quality consistently improved from 1989 to 1994, with the average number of days

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[13]For some school district-years, missing values were interpolated from adjacent years.

without an ozone exceedence increasing from 312 to 339. In contrast, teacher-student ratios at first improved in the first year and then declined. (The average number of students per teacher declined from 24.6 to 24.3, then increased to 25.8 at the end of the period.) Relative to the other two goods, public safety remained fairly flat throughout the period, improving very slightly overall, with most of the improvement toward the end of the period. Given these changes, one would expect that adjusting for public goods would tend to reduce the measure of the cost of living for the first year or two of the sample (1989-91), when air quality and teacher-student ratios both improved. In later years (1991-94), the adjustment would have an uncertain affect depending on the relative weights given to air quality and education.

Estimation and Results

Because the hedonic function is an equilibrium relationship arising from unknown preferences and cost functions, the choice of functional form of hedonic regressions is an empirical matter. Relative to an unrestricted Box-Cox specification, semilog specifications with these data have the lowest mean squared error of any common specification after converting to price levels. And too, they have been found to be one of the better specifications using a criterion of absolute value of errors in marginal values in experiments with simulated data (Cropper et al. 1988). Finally, semilog models have the advantage of being readily consistent with direct hedonic equations defined by Equation (12). Consequently, I use semilog models to estimate direct hedonic regressions. For consistency with the other attributes and ease of interpretation, I also renormalize the environmental variable and crime variable so that they measure *air quality* and *safety* rather than *pollution* and *crime* (hence, all attributes are goods).[14]

Table 2 reports the results of three model specifications. To test the effect of including public goods, each model is estimated with and without public goods. Model 1 includes a complete list of structural characteristics, as well as school district fixed effects. Model 2 replaces the school district fixed effects with county fixed effects and the test-score measure of school quality. Model 3 is the same as Model 2 with the addition of local demographic variables. A case could be made for each specification. The fixed effects in Model 1 have the advantage of controlling for unobserved spatial goods but require intracommunity variation to identify the parameters for air quality and public safety, and temporal variation to identify the parameter for class size, which may be insufficient. The demographic variables in Model 3 clearly have an effect on housing prices and are included in most empirical work, but they may not be included in an official price index for political reasons. While Model 2 may thus be the best political compromise, Model 3 is probably preferable from an analytical point of view.

All the models perform well by the usual statistical criteria. Almost all the attributes, including public goods, have positive signs and are significant. The negative estimated coefficient on bedrooms in most models is at first surprising but must be interpreted in light of

[14] In particular, I measure days without an ozone exceedence and SAFETY=3000-(CRIME RATE).

the fact that square feet is held constant. The exponent of the annual intercept is the measure of housing price inflation from the pervious year, while the other coefficients represent the percentage change in housing prices corresponding to a marginal change in the attribute. To facilitate the interpretation of the coefficients, Table 3 reports mean marginal values for each public good from each model, normalized to a 1% change in the public good and averaged over households. The table shows that values for observable public goods are generally lowest in Model 1, which controls for inter-community differences with school-district fixed effects. The table shows that the annual willingness to pay for one fewer violation of the daily ozone standard is \$10–\$29, or \$32 to \$91 for a 1 percent change in ozone-free days.[15] In comparison, the willingness to pay for a 1 percent change in educational quality is about twice is high, at \$52 to \$171, while the value for a 1 percent change in public safety is somewhat lower at \$26 to \$30.

To implement the first method discussed above, the augmented cost-of-living index, these values are used to construct an extra term in the BLS's Los Angeles CPI as shown in

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[15] Although most other studies have used other measures of pollution, these estimates are roughly comparable to those in the previous literature. Brucato et al. (1990) find a \$126 value for one fewer violation of the daily ozone standard in a hedonic regression with San Francisco data, when inflated to 1990 dollars. Other studies have focused on total suspended particulates (TSP). In their meta-analysis of some of these studies, Smith and Huang (1995) find a value that for a 1% change in their particulate measure would be \$181 in 1990 dollars. Thus, the measures reported here likely understate the importance of air quality if not all public goods.

Equation (7). Specifically, the average 1989 household income in this area according to the US Census ($69,700) was used for $\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a$ in the denominator of the expression on the left of Equation (7). The first term in the numerator of this expression, $\boldsymbol{p}_x^b \cdot \boldsymbol{x}^a$, was inflated using the BLS's Los Angeles CPI for each year, while the second term, $\overline{\boldsymbol{p}}_q^b \cdot \Delta\boldsymbol{q}$, was estimated with the reported marginal values and quantity data.

The resulting index, using marginal values from Model 3 (the "preferred model") is displayed in Figure 2 along with the BLS index. Incorporating public goods reduces the estimated cost of living by 0.4 and 0.3 percentage points in the first two years respectively, when both air quality and education are largely improving (see Fig. 1). In the remaining three years, the improving air quality and declining education approximately offset one another, with public goods having no overall affect the third and fifth years, and increasing the estimated cost of living by 0.1 percentage points in the fourth year. The average effect over the five years is to lower the cost-of-living index by about 0.1 percentage points per year. Although the adjustment seems small, it is on the same order of magnitude as the adjustments for quality in market goods found by the Boskin Commission and Lebow and Rudd (2003).

As discussed above, an alternative to using marginal values is to estimate an adjusted price index for the weakly complementary market good—in this case housing. Tables 4 and 5 report annual housing sub-indices for the direct and imputation approaches, respectively. These indices use the same hedonic regressions reported in Table 2, but in the case of the imputation approach the annual fixed effects are omitted and the index is estimated separately each year and chained (as discussed above). In terms of absolute levels, the hedonic sub-indices, as expected, show prices peaking in 1990 and then steadily falling. In terms of the relative difference made by public goods, the adjusted indices again generally lower the cost of living in the earlier years (when air quality and education improve) and tend to raise it in the later years (when declining education offsets improved air quality). The over-all affect differs across models, with air quality more important in Model 3 and education more important in Model 2 (consistent with the marginal values shown in Table 3).

These estimated housing sub-indices can replace the BLS housing price index in the CPI to create an adjusted cost-of-living index that accounts for public goods. Figure 3 shows such an adjusted cost-of-living index, again using the preferred Model 3. The relative difference made by public goods to the over-all cost of living is remarkably similar to the case of the augmented index. Accounting for public goods reduces the estimated change in the cost of living by 0.3 and 0.2 percentage points respectively in the first two years, and again increases it by a cumulative 0.1 percentage points in the remaining three years. Again, the average impact is roughly to reduce the estimated cost-of-living index by 0.1 percentage points per year.

These results are generally sensitive to which of the three functional forms is used, as they imply different relative weights for the different public goods. However, there are good reasons to prefer model 3 and in general the results are robust to other assumptions. For example, the imputation indices are robust to the use of weighted versus unweighted averages of homes (to adjust the sample of sales to the overall stock of housing) and the use of geometric versus arithmetic averages.

A particularly important issue is the choice of pooled versus chained indices. Pooled models offer the advantage of more statistical power and of imposing stability of the virtual prices (or virtual price functions) over time. However, this stability may be perceived as a disadvantage once the issue of taste change is raised. Taste change, of course, is a standard hobgoblin of neoclassical economics generally and of demand estimation in particular. For index numbers, it raises troubling questions about the fundamental meaning of comparing bundles over time (or space) if the tastes of those consuming the bundles differ. The problem can be minimized by chaining the index so that Laspeyres reference bundles and virtual prices are used only for comparing adjacent years. The results presented here are robust to chaining the augmented cost-of-living index and the adjusted cost-of-living index with the direct hedonic, while the imputation hedonic is already chained by its nature.[16]

Summary and Conclusions

This paper demonstrates how the environment and other public goods can be incorporated consistently into a green cost-of-living index. It demonstrates this process with an application to a regional index in Los Angeles, where air quality—the primary environmental focus—has rapidly improved over the past quarter-century.

The estimated models perform well statistically, suggesting it is possible to recover the required information. Moreover, the results indicate that including public goods in the cost-of-living index can make a substantial difference. For this test case, cost-of-living indices that ignore improving public goods can be significantly overstated—here, by about 0.1 percentage

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[16] Chaining the augmented index implies year-by-year marginal values estimated from separate hedonic regressions; chaining the direct index implies an overlapping series of hedonic regressions on two years of data, with an intercept shifter for the second year.

points per year. Although this is a seemingly small figure, it is commensurate with the findings of Boskin et al. (1996) and Lebow and Rudd (2003), who estimated an adjustment of 0.6 and 0.4 percentage points respectively for all market goods. And as with the market goods case, such a difference could have large impacts on the distribution of resources, realigning cost-of-living adjustments to wages, pensions, bond coupons, and income tax payments by billions of dollars, especially when compounded over time. Of course, the direction of the adjustment would differ in other contexts, as the quality of public goods might change in any direction. Certainly, over a long time horizon goods like education and public safety are generally perceived to have worsened, as have some environmental goods such as open space.

The indices presented here describe the change in cost of living over time. In addition to this application, the approached used would be equally valid for cost-of-living comparisons over space. Such an application would be closer in spirit to earlier work on quality-of-life indices by Blomquist, Berger, and Hoehn (1988) and recent work by Timmins (2003) which focuses on time-invariant regional differences. However, such comparisons raise questions about the choice of reference, or base, regions, and even stronger questions about the stability of preferences across regions—both old dilemmas in index number theory (see, e.g., Mudget 1951).

This empirical demonstration is for only a small handful of public goods in one region. To fully implement this approach in national indices, additional thought would have to be given to several questions. First, what additional public goods could be incorporated into the index and what criteria would be used to identify them (e.g., importance, degree of change, measurability)? Here, we must consider both additional local public goods, such as open space, and more pure public goods such as national defense or ecosystem diversity. In addressing this question, we should bear in mind that some public goods may be as difficult to measure in "physical" quantities as it is to price them. A trade-off will inevitably arise between the comprehensiveness of the index and the precision (or defensibility) of the estimate—a trade-off now implicitly made very conservatively by omitting all public goods.

Second, if there are multiple public goods that are highly correlated, making separate valuations difficult to identify, could a composite bundle (index) of those goods be used instead of individual measures? Third, what sampling techniques should be used for gathering quality as well as price data for such an index? As a related issue, to what extent could information in the virtual price functions be transferred from one region of the country to others? The issue is important since, in some areas, even local public goods like air quality may not vary sufficiently to identify virtual prices. Many other questions could also be asked, and many are the same questions being asked of other environmental indicators (e.g. The H. John Heinz III Center 2002).

Although perhaps not all types of public goods could ever be included with confidence in the cost-of-living index, even an index of spatially differentiated goods would be a step forward in our accounting of the environment and other public goods. Such an index would better achieve the stated objective of measuring the true cost of living, and so would provide more balanced information to policymakers regarding the true state of national welfare. It also would have the potential to induce major shifts in the political landscape of environmental policy. Favorable environmental policies would be reflected in the national accounts that command popular attention. The economic incentives of other players would also change as industry would no longer internalize only the costs of environmental policy but also some of the benefits, through lower wage adjustments. In these ways, an augmented or adjusted cost-of-living index could place the environment in a new seat at the political table.

Table 1. Observations and Means of Housing Variables by County, 1989–1994

Variable	*Los Angeles*	*Orange*	*Riverside*	*San Bernardino*	*Ventura*
N	144,731	71,147	43,588	37,686	22,489
Price (nominal $)	234,674	262,891	138,025	150,236	235,151
Lot size (acres)	0.18	0.16	0.24	0.21	0.22
House size (square feet)	1,568	1,766	1,629	1,619	1,831
Bathrooms	1.92	2.17	2.07	2.11	2.24
Bedrooms	3.03	3.33	3.26	3.29	3.47
Stories	1.12	1.41	1.30	N/A	N/A
Air conditioning (0/1)	0.16	0.18	0.84	0.80	N/A
Heat (0/1)	0.88	0.99	0.89	0.99	N/A
Gas (0/1)	N/A	N/A	0.95	0.88	N/A
Fireplace (0/1)	0.54	0.20	0.84	0.80	0.80
Swimming pool (0/1)	0.16	0.13	0.12	0.13	0.15
Garage (0/1)	0.83	0.59	0.95	N/A	N/A
Age of house	39.2	25.7	10.9	17.3	19.1
Ozone exceedences	30.4	8.6	52.6	78.0	5.1
PM_{10} ($\mu g/m^3$)	40.5	35.6	49.3	53.7	27.5
Expenditures per pupil	5,440	4,800	4,940	4,850	4,720
Teachers per pupil	0.040	0.040	0.040	0.040	0.039
Test score	4.79	5.68	4.80	4.82	5.04
Crime rate	612	530	715	651	401
1 mile of coast (0/1)	0.03	0.06	0	0	0.04
3 mile of coast (0/1)	0.09	0.17	0	0	0.14
Median income ($)	47,027	59,166	40,444	44,054	51,325
Percentage white	0.65	0.82	0.81	0.74	0.84
Percentage black	0.09	0.01	0.05	0.07	0.02
Percentage Hispanic	0.28	0.14	0.19	0.25	0.18
Percentage married	0.60	0.68	0.68	0.67	0.70
Percentage with children	0.38	0.40	0.43	0.46	0.43
Percentage over age 65	0.08	0.06	0.09	0.05	0.06
Percentage college graduate	0.17	0.22	0.10	0.12	0.17

Table 2. Direct Hedonic Price Regressions

	Model 1		Model 2		Model 3	
Variable	With public goods	Without public goods	With public goods	Without public goods	With public goods	Without public goods
Year 1990	0.0405	0.0452	0.0290	0.0395	0.0337	0.0447
	(0.0018)	(0.0017)	(0.0019)	(0.0020)	(0.0017)	(0.0017)
Year 1991	0.0256	0.0314	0.0166	0.0324	0.0081	0.0259
	(0.0018)	(0.0017)	(0.0019)	(0.0019)	(0.0016)	(0.0016)
Year 1992	-0.0050	-0.0009†	-0.0094	-0.0008†	-0.0257	-0.0091
	(0.0018)	(0.0016)	(0.0018)	(0.0019)	(0.0016)	(0.0016)
Year 1993	-0.0744	-0.0733	-0.0700	-0.0722	-0.0973	-0.0841
	(0.0019)	(0.0016)	(0.0018)	(0.0018)	(0.0016)	(0.0016)
Year 1994	-0.1123	-0.1124	-0.1046	-0.1114	-0.1367	-0.1242
	(0.0021)	(0.0016)	(0.0019)	(0.0018)	(0.0017)	(0.0015)
Orange Co.			-0.0706	0.0307	-0.0930	-0.0436
			(0.0017)	(0.0014)	(0.0016)	(0.0013)
Sanbern Co.			-0.4040	-0.4583	-0.3340	-0.3890
			(0.0018)	(0.0017)	(0.0017)	(0.0016)
Riversd. Co.			-0.4904	-0.5472	-0.4389	-0.4731
			(0.0018)	(0.0018)	(0.0018)	(0.0018)
Ventura Co.			-0.1701	-0.1213	-0.1512	-0.1118
			(0.0021)	(0.0020)	(0.0019)	(0.0018)
1 Mi. of Coast	0.1737	0.1747	0.1951	0.2365	0.1063	0.1160
	(0.0052)	(0.0052)	(0.0054)	(0.0054)	(0.0047)	(0.0047)
3 Mi. of Coast	0.1526	0.1536	0.1844	0.2314	0.1160	0.1472
	(0.0029)	(0.0029)	(0.0027)	(0.0028)	(0.0023)	(0.0023)
Ozone Free	0.0004		0.0007		0.0011	
Days	(3.81E-5)		(1.92E-5)		(1.69E-5)	
Test Score			0.1180		0.0424	
			(0.0012)		(0.0011)	
Teachers to	0.0524		0.1731		0.0815	
Students x100	(0.0049)		(0.0030)		(0.0028)	
Safety	4.40E-5		0.0002		5.09E-5	
	(5.15E-6)		(3.92E-6)		(3.56E-6)	
Bathrooms	0.0322	0.0317	0.0454	0.0484	0.0257	0.0240
	(0.0013)	(0.0013)	(0.0014)	(0.0015)	(0.0012)	(0.0013)
Bedrooms	-0.0185	-0.0183	-0.0263	-0.0292	0.0025	0.0017**
	(0.0008)	(0.0008)	(0.0009)	(0.0009)	(0.0008)	(0.0008)
Bldg Size (sqft)	0.0003	0.0003	0.0004	0.0004	0.0003	0.0003
	(1.73E-6)	(1.73E-6)	(1.81E-6)	(1.87E-6)	(1.68E-6)	(1.70E-6)
Lot Size (sqft)	6.34E-6	6.36E-6	5.66E-6	5.79E-6	5.26E-6	4.83E-6
	(1.14E-7)	(1.14E-7)	(1.16E-7)	(1.18E-7)	(1.07E-7)	(1.06E-7)
Fireplace	0.0829	0.0828	0.1018	0.1095	0.0582	0.0557
	(0.0011)	(0.0011)	(0.0011)	(0.0012)	(0.0010)	(0.0010)

Continued

Table 2 Continued

Variable	Model 1 With public goods	Model 1 Without public goods	Model 2 With public goods	Model 2 Without public goods	Model 3 With public goods	Model 3 Without public goods
Swimming Pool	0.0564 (0.0013)	0.0560 (0.0013)	0.0557 (0.0014)	0.0575 (0.0014)	0.0397 (0.0012)	0.0358 (0.0013)
Age	-0.0038 (9.26E-5)	-0.0038 (9.27E-5)	-0.0011 (9.27E-5)	-0.0016 (9.59E-5)	-0.0027 (0.0001)	-0.0029 (8.48E-5)
Age^2	1.06E-5 (1.24E-6)	1.07E-5 (1.23E-6)	5.36E-7† (1.28E-6)	1.09E-6† (1.31E-6)	2.96E-5 (1.15E-6)	0.0000 (1.15E-6)
Pct Black					-0.3360 (0.0037)	-0.3446 (0.0036)
Pct Hispanic					-0.0611 (0.0031)	-0.0995 (0.0030)
Pct College					1.2319 (0.0077)	1.2604 (0.0077)
Pct Married					0.0758 (0.0038)	0.0797 (0.0038)
Constant	9.3803 (0.0256)	9.8211 (0.0042)	7.8179 (0.0152)	9.6881 (0.0040)	8.6421 (0.0152)	9.6588 (0.0044)
School Fixed Effects	Yes	Yes	No	No	No	No
N	294,683	294,683	294,683	294,683	292,153	292,153
R^2	0.75	0.75	0.70	0.68	0.77	0.76

Dependent variable: log of sales price in current dollars.
Standard Errors in Parentheses.
All variables significant at 1% level unless otherwise noted.
**Significant at 5 percent level. †Not significant at 10 percent level.

Table 3. Marginal WTP normalized to 1 percent change (Average nominal dollars 1989-94)

	Model 1	Model 2	Model 3
Ozone	32.32	58.48	91.23
Teachers per Student	51.55	170.18	80.14
Public Safety	25.95	98.27	30.03

Table 4. Annual Direct Hedonic Housing Price Indices, 1989–1994

Year	Model 1 With public goods	Model 1 Without public goods	Model 2 With public goods	Model 2 Without public goods	Model 3 With public goods	Model 3 Without public goods
1989-90	1.041	1.046	1.029	1.040	1.034	1.046
	(1.038-1.044)	(1.043-1.049)	(1.026-1.033)	(1.037-1.044)	(1.031-1.037)	(1.043-1.049)
1990-91	0.985	0.986	0.988	0.993	0.975	0.981
	(0.983-0.988)	(0.984-0.989)	(0.985-0.991)	(0.99-0.996)	(0.972-0.977)	(0.979-0.984)
1991-92	0.970	0.968	0.974	0.967	0.967	0.966
	(0.967-0.972)	(0.966-0.971)	(0.972-0.977)	(0.965-0.97)	(0.964-0.969)	(0.963-0.968)
1992-93	0.933	0.930	0.941	0.931	0.931	0.928
	(0.931-0.935)	(0.928-0.932)	(0.939-0.944)	(0.929-0.933)	(0.929-0.933)	(0.926-0.930)
1993-94	0.963	0.962	0.966	0.962	0.961	0.961
	(0.961-0.965)	(0.960-0.964)	(0.964-0.968)	(0.959-0.964)	(0.959-0.963)	(0.959-0.963)

Ninety percent confidence intervals shown in parentheses, based on White-corrected standard errors of the annual coefficients in Table 2.

Table 5. Annual Weighted Imputed Hedonic Housing Indices, 1989–1994

	Model 1'		Model 2'		Model 3'	
Year	With public goods	Without public goods	With public goods	Without public goods	With public goods	Without public goods
1989-90	1.047	1.077	1.069	1.080	1.063	1.074
	(1.042-1.052)	(1.074-1.080)	(1.066-1.072)	(1.076-1.083)	(1.061-1.066)	(1.071-1.077)
1990-91	1.027	1.013	1.019	1.027	0.998	1.006
	(1.023-1.030)	(1.011-1.016)	(1.016-1.022)	(1.024-1.030)	(0.995-1.000)	(1.003-1.008)
1991-92	1.029	0.991	0.999	0.997	0.985	0.988
	(1.025-1.033)	(0.988-0.933)	(0.996-1.002)	(0.994-1.000)	(0.982-0.987)	(0.985-0.990)
1992-93	1.002	0.954	0.969	0.961	0.951	0.951
	(0.997-1.007)	(0.952-0.956)	(0.966-0.972)	(0.959-0.964)	(0.949-0.954)	(0.949-0.954)
1993-94	1.015	0.986	0.994	0.994	0.981	0.985
	(1.013-1.018)	(0.984-0.988)	(0.991-0.997)	(0.992-0.996)	(0.979-0.984)	(0.983-0.987)

Imputation hedonics based on Laspeyres bundles (i.e., holding quality constant at baseline levels).

Ninety percent confidence intervals shown in parentheses, based on bootstrap with 300 draws. (Each represents a random re-drawing from the data with replacement, regressions in each year on those data, and the computation of the average price-relative.)

Figure 1. Plot of Public Goods overTime
(three-year moving average over households, 1989=100)

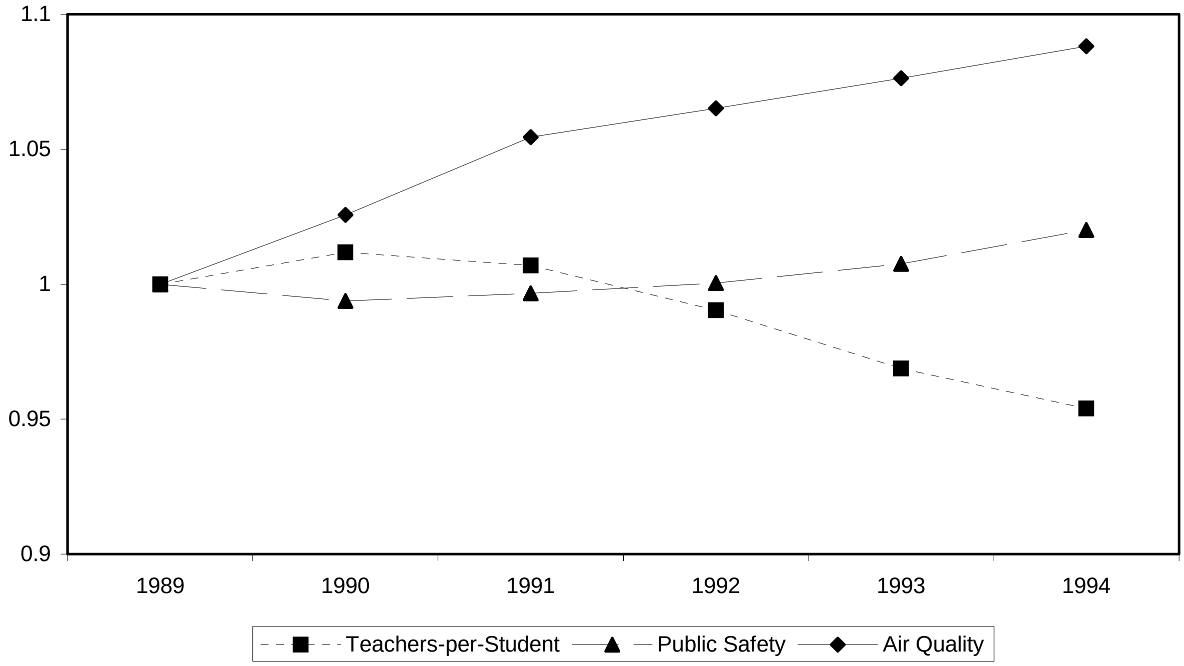
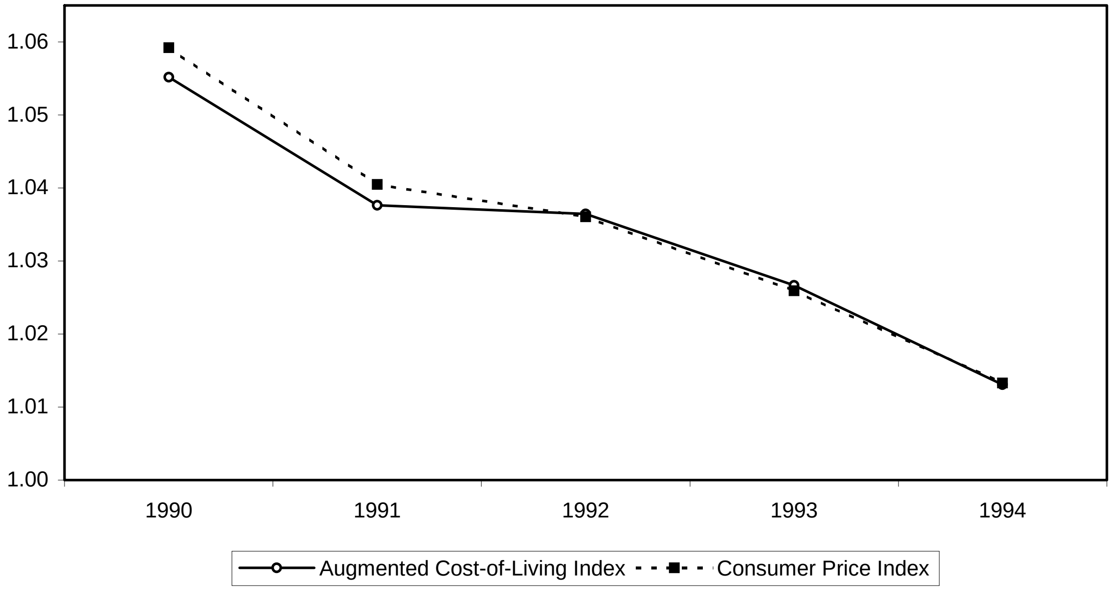
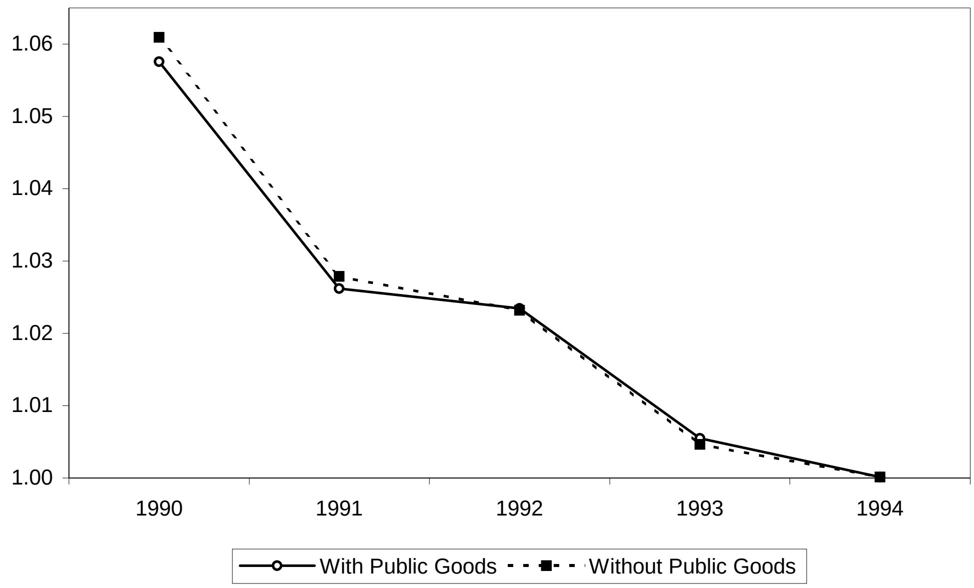


Figure 2. Annual Cost-of-Living Index with and without Public Goods, Augmented Cost-of-Living Index



Based on direct hedonic Model 3, Table 2.

Figure 3. Annual Cost-of-Living Index with and without Public Goods, Adjusted Cost-of-Living Index



Based on direct hedonic Model 3, Table 2.

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