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 noticeable 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 and have recently been endorsed by the U.S. National Academy of Sciences (Nordhaus and Kokkelenberg 1999).[1] Less attention has been given to similarly adjusting national price indices. And yet, incorporating public goods into price indices would not only be consistent with the principles advocated in income accounts, it would also bring them closer to their own objective. Since these indices are defined as the relative cost of achieving a given 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 goods in his illustration.[2]

This paper suggests two concrete ways that a green cost-of-living index might be constructed and provides examples for households in Los Angeles from 1989 to 1994. Air

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[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 1994a,b) but was directed by Congress to discontinue the work. For additional discussion and examples, see Lutz (1993) and Repetto et al. (1989).

[2] The notion is also discussed in Boskin et al. (1996) and Abraham et al. (2000). These authors also discuss other ways that the U.S. Consumer Price Index (CPI) can 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 of which parallels the kinds of adjustments that could be made for public goods.

quality, an important public good in the region, dramatically improved over the period covered. Accounting for this improvement in a cost-of-living index reduces the estimated inflation rate by about 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).

Theory of Cost-of-Living Indices and Estimation Strategy

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 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.

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:

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

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

where w_k^t is the expenditure share for good k computed at period t quantities and baseline 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.

The first way to introduce quality into the index, applicable to the widest variety of circumstances, is to add additional commodities to the Laspeyres index and evaluate them at their *implicit* or *virtual* prices. Virtual prices, suggested during World War II as a way to account for the constraints on the consumer due to rationing, 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. The key insight here is that from the consumer's perspective, public goods such as the environment are like rationed goods. For example, households cannot choose their individual level of air quality; they must accept what society and the environment give to them.

Let $\boldsymbol{q}$ be a vector of public good levels. Following Neary and Roberts (1980), the virtual price vector $\overline{\boldsymbol{p}}_q$ associated with a level of these public goods $\overline{\boldsymbol{q}}$ is defined implicitly as

$$\overline{\boldsymbol{q}} \;=\; \arg\min \{\boldsymbol{p}_x \cdot \boldsymbol{x} + \overline{\boldsymbol{p}}_q \cdot \boldsymbol{q}\} \;\; s.t. \;\; u(\boldsymbol{x}, q) \;\geq\; \overline{u} \tag{4}$$

Virtual prices are a function of the prices of other goods, the rationing level, and the utility 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:

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

Equation (5) shows the conventional Laspeyres bounds, with an additional term for the virtual prices and rationed quantities of the public goods and an adjustment for "virtual income." Simply rearranging and canceling some terms, this equation can be rewritten as

$$\frac{\left(\boldsymbol{p}_x^b \cdot \boldsymbol{x}^a\right) - \overline{\boldsymbol{p}}_q^b \cdot \Delta \boldsymbol{q}}{\left(\boldsymbol{p}_x^a \cdot \boldsymbol{x}^a\right)} \geq \mathrm{I}^{ab}\left(\boldsymbol{p}_x^b, \boldsymbol{p}_x^a, \overline{\boldsymbol{p}}_q^b, \overline{\boldsymbol{p}}_q^a, u^a\right). \tag{6}$$

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, this 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.

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 numerator of the right side of Equation (5), $\overline{p}_q^b$ is evaluated at the prices and the level of the public good in the comparison period but when income is hypothetically adjusted to achieve reference-period utility. 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 the advantage of this approach in 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.[3]

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 individual species or collective habitat 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. Then when p^* and p^{**} are defined implicitly by the equations

$$v(p_h^*, \boldsymbol{q}_h^a, \boldsymbol{p}_x^b, m(p_h^b, \boldsymbol{q}_h^b, \boldsymbol{p}_x^b; u^a)) = v(p_h^a, \boldsymbol{q}_h^a, \boldsymbol{p}_x^a, y^a) \tag{7}$$

and

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

it can be shown that the following bounds for the Laspeyres and Paasche indices still hold (Willig 1978):

[3] For more on weak complementarity, see Bockstael and McConnell (1993), Smith and Banzhaf (2001), and Willig (1978).

$$L^{*ab} \equiv \frac{p_h^* h^a + \boldsymbol{p}_x^b \cdot \boldsymbol{x}^a}{p_h^a h^a + \boldsymbol{p}_x^a \cdot \boldsymbol{x}^a} = \geq I^{ab}(\mathbf{p}^b, \mathbf{p}^a, q^b, q^a, u^a) \quad (9)$$

$$P^{**ab} \equiv \frac{p_h^b h^b + \boldsymbol{p}_x^b \cdot \boldsymbol{x}^b}{p_h^{**} h^b + \boldsymbol{p}_x^a \cdot \boldsymbol{x}^b} = \leq I^{ab}(\mathbf{p}^b, \mathbf{p}^a, q^b, q^a, u^b) \quad (10)$$

L^* is the Laspeyres index with p^* replacing p^b, while 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.

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, obtaining both the marginal values for the first index and the quality-adjusted price of housing for the second. These hedonic indices are consistent with the definitions of $p*$ or $p**$ under certain conditions.[4]

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.[5] Similarly, they have

[4] 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.

[5] See Palmquist (forthcoming) for an overview of this context. Smith and Huang (1996) give a metaanalysis and bibliography of applications to air quality.

a long history in quality-adjusted price indices for market goods, as required for the second index, and in the United States are now used to adjust the indices for computers, televisions, and apparel.[6]

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

$$\ln(p_j) = \alpha^a D_j^a + \alpha^b D_j^b + f(\boldsymbol{q}_j) + \varepsilon_j \tag{11}$$

where D_j^a and D_j^b are indicator variables for the base 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 (11) 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 second variation uses the 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 variety *j* with a certain quality in period *a*. Its quality changes so that in period *b,* we designate it as variety *j'*. Then the predicted prices from a hedonic regression $\hat{p}$ can be used to impute the price relatives for each variety as if they had existed in each period, namely $\hat{p}_j^b / p_j^a$ and $p_{j'}^b / \hat{p}_{j'}^a$. These individual price relatives can then be averaged and entered into a

[6] Griliches (1961) is the classic reference. Other examples include Gatzlaff and Ling (1994), Kokoski et al. (1999), Liegey (1994), and Raff and Trajtenberg (1997).

price index in the same way as other goods. With this 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.[7] In the empirical illustration given below, I instead explicitly compute user cost at time t as

$$R^t = (i^t + \tau^t + \delta + \pi^t)\cdot P^t \tag{12}$$

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

$$\pi^t \quad = \quad \frac{1}{5}\sum_{s=t-4}^{t}\frac{P^{s+1}}{P^s} \tag{13}$$

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

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[7] 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).

asset prices determine P^{t+1}/P^t.[8] 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 indices estimated below are equally applicable to either asset prices or rental prices.

After describing the data used in the empirical example, I estimate such indices for the Los Angeles example.

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. 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

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[8] 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.

the period. Air quality in turn is represented by ozone and particulate matter less than 10 microns in diameter (PM_{10}). These are the pollutants most commonly in violation of California's air quality standards and the pollutants found to cause the greatest damages to human health.

Air quality data were obtained from the California Air Resources Board.[9] These data are some of the richest in the world: for ozone and PM_{10}, respectively, an average of 50 and 21 monitors are available each year in the study area, and monitors in neighboring counties 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 (e.g., on the *LA Times* weather page) in the form of ozone alerts. In contrast to ozone, PM_{10} damages tend to follow from chronic exposures. For this reason, I use annual geometric means of PM_{10} to represent particulate pollution. Pollution values at the nearest monitor are imputed to each house. The mean pollution values by county are also included in Table 1. In addition, Figure 1 plots the change in air quality over the period covered. Because air quality improved substantially, price indices that ignore public goods may have overstated inflation.

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 only some of the likely locational variables that might matter to households. Proxy variables for school quality (expenditures per student, teacher-to-student ratios, and test scores) were obtained from the National Center for Education Statistics. Crime rates per 10,000 people were obtained at local jurisdictions from the California Department of Justice. The distance to the coast was calculated for each house, and indicator variables created for houses within one and three miles. 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

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9 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/airsweb.

reported in Table 1. Each of these additional public goods is available only in a single cross section. Thus, for the purposes of this illustration, they are included only as statistical controls and do not enter the index on the same footing as air quality.

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. They also 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 11. Consequently, I use semilog models to estimate direct hedonic regressions. For consistency with the other attributes, I also renormalize the environmental variables and crime variable so that they measure *air quality* and *safety* rather than *pollution* and *crime*.[10]

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 characteristics, as well as school district fixed effects. Model 2 replaces the school district fixed effects with county fixed effects and school quality variables, and drops particulate matter. Model 3 is the same as Model 2 with the addition of local demographic variables. A case can be made for each specification, and each has drawbacks. The fixed effects in Model 1 have the advantage of controlling for unobserved spatial goods but depend on intracommunity variation to estimate the parameters for public goods, which may be too sensitive to the imputation scheme. The demographic variables in Model 3 clearly have an effect

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[10] In particular, I measure days without an ozone exceedence, PM_{10} quality as $100\text{-}PM_{10}$, and SAFETY=3000-(CRIME RATE). The latter two normalizations are convenient round numbers sufficient to cover the maximum values of particulate matter and crime rates, respectively.

on housing prices and are included in most empirical work, but they are unlikely to be included in an official price index. Model 2 emerges as a preferable compromise.

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 the fact that square feet is held constant. The estimated coefficients on air quality, interpreted as marginal values, imply an annual willingness to pay of about $18 for a one-unit improvement in annual average PM_{10} (Model 1) and $8–$28 for one fewer violation of the daily ozone standard. Although most other studies have not used the same measures of pollution, these estimates are roughly comparable to those in the previous literature.[11] These marginal values imply a total (linearized) annual value of $232–$816 for the 29 fewer ozone alerts experienced on average in 1994 relative to 1989.

Implementing the first method discussed above, these values are used to construct an extra term in the BLS's Los Angeles CPI as shown in Equation (6). The resulting index is displayed in Figure 2 along with the BLS index. Incorporating reductions in ozone into the index lowers the inflation rate by an average of about 0.1 percentage points per year. The difference is noticeable in 1989–90 and 1992–93, the years with the largest improvements (see Figure 1). 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.

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 3 and 4 report annual housing subindices 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

[11] Brucato et al. (1990) find a $66 value for one fewer violation of the daily ozone standard in a hedonic regression with San Francisco data. Other studies have focused on total suspended particulates (TSP). In their metaanalysis, Smith and Huang (1995) find a mean value of $14 for a one-unit improvement in TSP. Chay and Greenstone (2000) find values on the order of $30 when instrumenting for air quality variables. However, Chattopadhyay (1999) reports a range of marginal value of $21–$363 from his hedonic models, and Beron et al. (2000) report values of $29–$151 for Los Angeles. To convert such TSP values to PM_{10}, a common rule of thumb is to double them.

approach the annual fixed effects are omitted. In terms of absolute levels, the hedonic subindices, as expected, show prices peaking in 1990 and then steadily falling for the next two years. Although the asset prices continue to fall, rising interest rates increase the index for the 1991–94 period. Comparisons of models with and without public goods imply that inflation in housing prices would again be overestimated when public goods are omitted. Depending on the model used, housing inflation is overestimated by an average of 0.1 to 0.6 percentage points per year when public goods are omitted, with a range of -0.3% to 1.5% per year. The largest statistically significant differences are again in the years with the greatest improvement in air quality (1989–90 and 1992–93).[12]

With a weight of 0.28 on housing in the U.S. Consumer Price Index, these differences again correspond to a 0.1 percentage point difference in overall inflation, with the biggest statistically significant differences again occurring in 1989–90 and 1992–93. Thus, this second approach to the index is comparable to the first. Figure 3 again compares the overall inflation measures, substituting the estimated subindices from Model 2 for the BLS housing subindex for LA.

Summary and Conclusions

This paper demonstrates how the environment and other public goods can be incorporated consistently into a 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-

[12] A comparison of Tables 3 and 4 indicates that the estimated difference made by public goods is generally not sensitive to the procedure for calculating the subindex. These results are robust in other ways as well. For example, they are not sensitive to the specification of the hedonic price regression, such as the inclusion of interaction effects among variables. In addition, the direct hedonic indices are also robust to the assumption of constant parameter values across years, and 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).

living index can make a substantial difference. For this test case, cost-of-living indices that ignore dramatically improving public goods can be significantly overstated—here, by about 0.1 percentage points per year. Although this is a seemingly small figure, it is commensurate with the findings of Boskin et al. (1996), who estimated an adjustment of 0.6 percentage points for 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.

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)? Second, how are some public goods to be objectively measured in "physical" terms? Third, 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? Fourth, what sampling techniques should be used for gathering quality as well as price data for such an index? Many other questions could also be asked.

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 not only better achieve the stated objective of measuring the true cost of living, it would also provide more balanced information to policymakers regarding the true state of national welfare. It would also 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$)	40.5	35.6	49.3	53.7	27.5
Expenditures per pupil	5,440	4,800	4,940	4,850	4,720
Teachers per pupil	0.041	0.041	0.040	0.040	0.040
Test score	4.79	5.68	4.80	4.82	5.04
Crime rate	625.5	556.9	717.6	669.3	396.3
1 mile of coast (0/1)	0.03	0.06	0	0	0.04
3 mile of coast (0/1)	0.09	0.17	0	0	0.14
Median income ($)	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	0.0379	0.0469	0.0321	0.0416	0.0328	0.0473
Year 1991	0.0247	0.0333	0.0203	0.0341	0.0136	0.0279
Year 1992	-0.0111	0.0023†	-0.0119	0.0027†	-0.0197	-0.0053
Year 1993	-0.0879	-0.0696	-0.0902	-0.0681	-0.1042	-0.0801
Year 1994	-0.1286	-0.1109	-0.1314	-0.1097	-0.1457	-0.1222
Orange County	—	—	-0.0853	0.0310	-0.0987	-0.0435
San Bernadino County	—	—	-0.3981	-0.4580	-0.3346	-0.3887
Riverside County	—	—	-0.4954	-0.5465	-0.4430	-0.4724
Ventura County	—	—	-0.1681	-0.1207	-0.1402	-0.1113
1 mile of coast	0.1741	0.1750	0.2106	0.2367	0.1081	0.1159
3 miles of coast	0.1519	0.1534	0.1706	0.2314	0.1020	0.1471
Ozone-free days	0.0003	—	7.7e-4	—	0.0011	—
Air quality (PM_{10})	0.0007	—	—	—	—	—
Test score	—	—	0.1185	—	0.0429	—
Teachers per pupil	—	—	20.4063	—	14.4013	—
Safety	2.7e-5	—	1.4e-4	—	2.7e-5	—
Bathrooms	0.0319	0.0317	0.0431	0.0484	0.0234	0.0239
Bedrooms	-0.0185	-0.0183	-0.0261	-0.0293	0.0027	0.0017 *
House (sq. feet)	3.4e-4	3.4e-4	3.8e-4	4.0e-4	2.8e-4	2.9e-4
Lot (sq. feet)	6.3e-6	6.4e-6	5.7e-6	5.8e-6	5.3e-6	4.8e-6
Fireplace	0.0830	0.0830	0.0993	0.1098	0.0559	0.0559
Swimming pool	0.0560	0.0560	0.0553	0.0575	0.0387	0.0359
Age	-0.0038	-0.0038	-0.0011	-0.0016	-0.0027	-0.0029
Age^2	1.1e-5	1.1e-5	-5.4e-7†	1.3e-6 †	2.8e-5	3.3e-5
Percentage black	—	—	—	—	-0.3361	-0.3455
Percentage Hispanic	—	—	—	—	-0.0610	-0.1001
Percentage college graduate	—	—	—	—	1.2221	1.2611
Percentage married	—	—	—	—	0.0819	0.0792
Constant	11.4825	11.6965	9.6096	11.5629	10.3301	11.5341
School fixed effects	Yes	Yes	No	No	No	No
N	295,227	295,227	295,227	295,227	292,697	292,697
R^2	0.75	0.75	0.70	0.68	0.77	0.76

Dependent variable: log of sales price in current dollars. All variables significant at 1% level unless noted.

*Significant at 5% level. †Not significant at 10% level.

Table 3. Annual Direct Hedonic Housing Price 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	0.956 (0.953–0.959)	0.964 (0.961–0.967)	0.950 (0.947–0.953)	0.959 (0.956–0.962)	0.951 (0.948–0.953)	0.965 (0.962–0.967)
1990–91	0.949 (0.947–0.952)	0.949 (0.946–0.951)	0.951 (0.948–0.953)	0.955 (0.952–0.958)	0.944 (0.941–0.946)	0.944 (0.941–0.946)
1991–92	1.003 (1.001–1.006)	1.008 (1.006–1.011)	1.007 (1.004–1.01)	1.008 (1.005–1.011)	1.006 (1.004–1.008)	1.006 (1.004–1.008)
1992–93	1.033 (1.03–1.035)	1.038 (1.035–1.04)	1.031 (1.028–1.034)	1.039 (1.036–1.041)	1.025 (1.022–1.027)	1.035 (1.032–1.037)
1993–94	1.241 (1.239–1.244)	1.241 (1.238–1.243)	1.241 (1.238–1.244)	1.240 (1.237–1.243)	1.240 (1.238–1.243)	1.240 (1.237–1.242)
Average difference	0.004		0.004		0.005	

90% confidence intervals shown in parentheses, based on White-corrected standard errors.

Table 4. Annual Weighted Imputed Hedonic Housing Indices, 1989–1994

	Model 4		*Model 5*		*Model 6*	
Year	*With public goods*	*Without public goods*	*With public goods*	*Without public goods*	*With public goods*	*Without public goods*
1989–90	0.993 (0.989–0.997)	0.994 (0.991–0.996)	0.989 (0.985–0.993)	0.998 (0.995–1.001)	0.978 (0.975–0.982)	0.993 (0.99–0.996)
1990–91	0.978 (0.975–0.98)	0.978 (0.975–0.981)	0.986 (0.982–0.989)	0.993 (0.99–0.996)	0.971 (0.968–0.974)	0.972 (0.969–0.975)
1991–92	1.028 (1.023–1.032)	1.030 (1.026–1.032)	1.033 (1.031–1.036)	1.036 (1.034–1.039)	1.028 (1.024–1.03)	1.029 (1.025–1.032)
1992–93	1.059 (1.057–1.064)	1.064 (1.06–1.066)	1.059 (1.056–1.063)	1.072 (1.068–1.074)	1.050 (1.047–1.053)	1.061 (1.059–1.065)
1993–94	1.275 (1.271–1.277)	1.272 (1.268–1.275)	1.283 (1.279–1.285)	1.283 (1.28–1.287)	1.275 (1.271–1.277)	1.272 (1.27–1.276)
Average difference	0.001		0.006		0.005	

90% confidence intervals shown in parentheses, based on a bootstrap with 300 draws.

FIGURE 1. Ozone Exceedences in the Los Angeles Market (Weighted by Population)

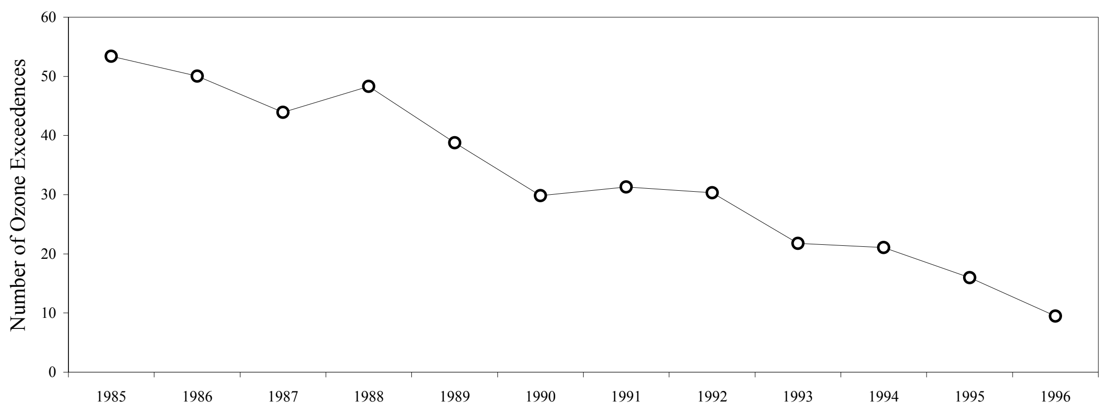
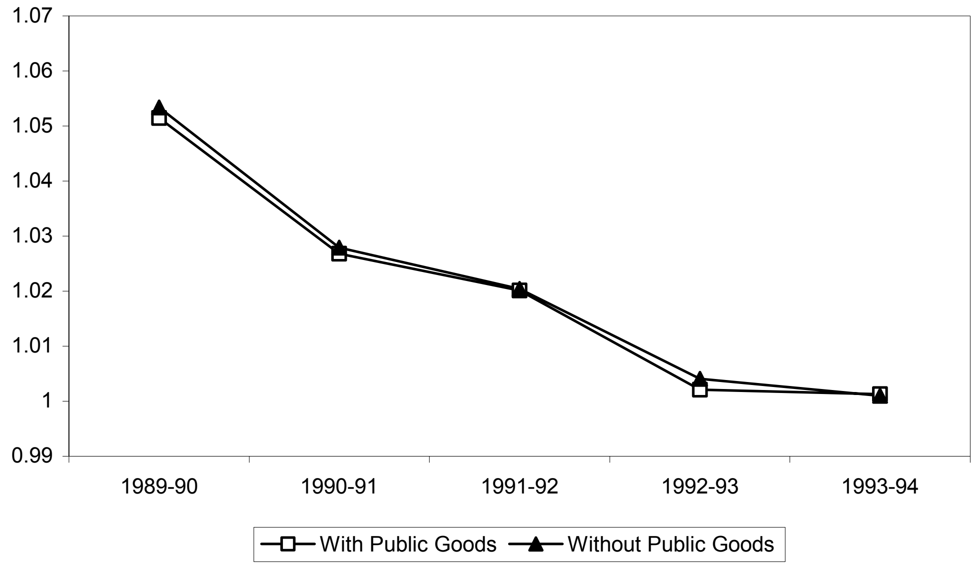


Figure 2. Annual Total Inflation with and without Public Goods, Estimated by Adding a Class for Public Goods

Figure 3. Annual Total Inflation with and without Public Goods, Estimated from Adjustment to Housing Subindex



Based on direct hedonic Model 2, Table 2.

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