COMPENSATING DIFFERENTIALS AND EVOLUTION OF THE QUALITY-OF-LIFE AMONG U.S. STATES

by

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Abstract

This paper provides the first application of the compensating differential paradigm to the evaluation of the extent and sources of evolution in quality-of-life among U.S. states. In addition to providing estimates of quality-of-life rankings for U.S. states over the 1981- 1990 period, we use estimated implicit prices on place-specific amenities to calculate the contributions of various factors to evolution in the quality-of-life. Our findings indicate that the quality-of-life rankings are relatively stable across model specifications and over time for certain poorly ranked, densely-populated midwestern and eastern industrial states and for many high quality-of-life rural western states. However, we also find evidence of a substantial deterioration in the quality-of-life in some states that experienced rapid population growth during the decade, with reduced spending on highways and increased traffic congestion and air pollution accounting for the bulk of the deterioration in quality-of-life in these states. In contrast, states exhibiting an improvement in the quality-of-life rankings ascended for a variety of reasons, including reduced state and local government income tax burdens, improved air quality, increased highway spending, and reduced commute times.

I. Introduction

There has long been widespread popular interest and media debate over regional differences in the quality of life. This interest implicitly assumes that the set of amenities specific to a household's geographic location is an important element of household utility, a view that generally has been confirmed by empirical studies of residential location decisions (see, for example, Graves (1980) or Berger and Blomquist (1992)). Economists have entered the quality-of-life debate directly only in recent decades, arguing that pecuniary differences across locations in wages or land rents should compensate for the differences in the nonpecuniary characteristics of locations that traditionally were included as elements of the popular rankings.[1] Although there has been some disagreement about exactly how to specify such models, the basic approach consistently now used in the academic literature is based on the work of Rosen (1979) and Roback (1982, 1988), who demonstrated how differences in wage rates and land rents across locations could be used to estimate the utility valuation weights (or implicit prices) of the nonpecuniary characteristics of locations. More recent applications of this compensating differential approach to estimating the quality of life include Hoehn et al (1987), Blomquist et al (1988), Gyourko and Tracy (1991), and Stover and Leven (1992), among others.

One uniform aspect of research on this topic is the static nature of such estimates. In virtually all of the research, quality-of-life estimates have focused on relative rankings at a single point in time. This focus is perhaps not surprising given the severe data limitations associated with the compilation of time-series data on amenities. However, the static nature of existing quality-of-life analyses provides little insight as to the nature or determinants of intertemporal evolution of nonpecuniary changes in local area living conditions.

In this paper, we extend the existing literature in two directions. First, we construct a time-series of quality-of-life rankings by applying the equilibrium compensating differentials approach employed in the static literature to a pooled time-series and cross-section of states for the 1981-1990 period. After computing the implicit prices associated with each amenity characteristic, we interact those prices with an extensive dataset of time-varying amenity characteristics to construct time-series of state-level rankings of the quality-of-life. Second, we derive our estimates of implicit prices from a three-equation system that goes beyond earlier two-equation specifications in the literature by explicitly accounting for the capitalization of amenities in wages, housing rents, *and* prices of other commodities traded

[1] The choice of individual amenities to be included in the popular rankings and the weights assigned to the individual amenities vary considerably across the popular "quality-of-life" indexes. Among the best known of the popular rankings of quality-of-life is Boyer and Savageau's *Places Rated Almanac* (1985, 1989, 1993). In that publication, selected individual amenities are assigned equal weights.

primarily in local markets. Prior empirical analyses failed to adequately address the capitalization of place-specific amenities in the cost-of-living other than housing. Our results show significant evolution in the quality-of-life rankings over the 1980s. For example, perceptible improvements in the relative quality-of-life are found for such states as Alaska and Arizona, while notable declines are evident for Idaho, Nevada, New Hampshire, and New Mexico. In contrast, sparsely populated mountainous Western states such as Montana and Wyoming are highly ranked in the estimated quality-of-life throughout the decade, while more densely-populated midwestern and eastern industrialized states consistently score the lowest in terms of estimated quality of life. With respect to individual amenities, reduced spending on highways and increased traffic congestion and air pollution account for the bulk of the deterioration in states with declining quality-of-life rankings. In contrast, states with improved quality-of-life rankings ascended for a variety of reasons, including reduced state and local government income tax burdens, improved air quality, increased highway spending, and reduced commute times.

Finally, we examine the robustness of our results to numerous modeling issues that have been raised in the literature. For example, like other studies in the economics literature, our list of amenity indicators undoubtedly omits some relevant location-specific characteristics, and it is useful to compare our baseline results with those from a model that includes parameters for unobserved group effects. We find that many of our results are robust across alternative specifications.

II. Theoretical Approach

Rosen (1979) provides the conceptual framework for most economic analyses of quality-of-life differences across locations, and in this regard, our paper is no different. In particular, Rosen assumes that locations can be viewed as interrelated bundles of wages, rents, and amenities, with the specific makeup of the bundles offered differing across locations. Households and firms then compete for a fixed number of sites across those areas, with households seeking to maximize utility and firms attempting to minimize costs through their locational choices. If agents face no informational or mobility-related transactions costs and households have common preferences for amenities, then the Rosen model implies--given a fixed distribution of amenities--that wages and land rents will vary across locations in order to equilibrate household utility. In particular, a spatial equilibrium is attained when moving would neither improve household utility nor reduce firm costs.[2] In Roback's (1982) generalization of this model to include land as a factor of production, land rents are higher in more desirable amenity-rich areas, but the effect of higher amenities on wages also depends on how the amenities affect firm productivity.

Roback (1982) also expands upon the simple model by introducing a sector producing a vector of commodities that are not traded beyond the location's borders. The Roback discussion focuses on housing as the prototypical locally-traded commodity, with rental price n_j in location j. In addition to housing, we consider an additional local commodity, with price c_j .

Irrespective of the effect of amenities on firm productivity, household valuations of alternative locations, indexed by j, can be derived from an indirect utility function V of the form

$$V = V(w_j, n_j, c_j; a_j), \tag{1}$$

where, in addition to the prices mentioned above, w_j is the wage rate in location j, and a_j is a scalar (for expositional purposes) indexing local amenities. The unit price of the traded commodity is suppressed. As with the standard approach, effective labor supply is assumed always to be unity, so the wage rate w_j also indexes the representative consumer's income. Furthermore, we ignore saving, so that current income is also the consumer's level of expenditures.

The total differential of this indirect utility function is zero at maximal utility and is given by

$$dV = V_w dw + V_n dn + V_c dc + V_a da = 0, \tag{2}$$

where the V_c , V_n , V_c , and V_a terms are partial derivatives of the indirect utility function with respect to income, the prices of the local consumption items (n_j and c_j), and the amenity index. Correspondingly, the amount of income needed to compensate a household for a small change in amenities is given by

$$V_a/V_w = - V_n/V_w (dn/da) - V_c/V_w (dc/da) - dw/da. \tag{3}$$

Roy's Identity implies that the effective demand for each of the consumption items will be equal to the ratio of the utility lost by foregoing some of that item to the utility gained by augmenting income slightly; this gives a basis for rewriting $-V_n/V_w$ and $-V_c/V_w$ in terms of quantities consumed. Thus, if housing is consumed in amount h and other local goods in amount y, the compensating differential also can be written as

[2] Insofar as utility is invariant across locations, the term "quality of life" can be misleading. Here, as is standard in the literature, we use the term "quality of life" to mean the household's aggregate valuation of the nonpecuniary characteristics of a location.

$$V_a/V_w = h\ (dn/da) + y\ (dc/da) - dw/da \quad = n\ h\ d(\ln n)/da + c\ y\ d(\ln c)/da - w\ d(\ln w)/da. \tag{4}$$

The main strategy in the recent empirical literature has been to estimate separate (reduced form) equations for the (logarithm of) housing expenditures and (the logarithm of) wage income as functions of amenities, from which d(ln n)/da and d(ln w)/da can be directly computed.[3] As is evident in equation (4) above, failure to account for amenity capitalization in the prices of local non-traded goods other than housing may result in biased estimates of compensating differentials. In the literature, locally-traded goods other than housing either have been ignored (Blomquist et al (1988)) or included as an observed amenity in the wage and housing expenditure equations (Gyourko and Tracy (1991)). Instead, we add the separate reduced form equation for the price of local commodities other than housing, from which we compute the capitalization of amenities in such prices, the d (ln c)/da term.

III. Empirical Specification

Although evolution in the quality-of-life has been the subject of much media commentary and speculation, empirical applications of the Rosen-Roback paradigm to date have been static in nature and, accordingly, have failed to provide much insight as to intertemporal changes in the amenity vector or the evolution in the quality-of-life across places and over time. Yet, it seems perfectly natural to expect that such changes have occurred. For example, environmental attributes have changed in some places, owing to natural disasters or the spillover effects of local economic development. Public goods (such as school services and public safety) vary in quality over time. In some places at some times, household tax burdens are high, given the level of public good provision, and at other times the rate of taxation is less onerous.[4] In the case of efficient markets and low adjustment costs, such changes in place-specific amenities should lead to rapid adjustment in wages and prices of locally-traded goods, so as to maintain the necessary

[3] In an alternative approach, Stover and Leven (1992) compute a quality-of-life index via estimated amenity prices from a single structural equation. In this structural equation, housing expenditures are a function of the endogenous wage premium in the area, in addition to the housing quality and amenity characteristics which are the driving (assumed exogenous) variables in the reduced form approach. One major disadvantage of the structural approach is that instrumental variables are required to obtain consistent estimates of the full implicit amenity prices; this econometric issue, which Stover and Leven (1992) ignored, might explain why they found very different results from the reduced form and structural approaches.

[4] For example, tax receipts from petroleum industry activities in Alaska have funded a relatively high and varying level of public services to households there. In California, the level of public service provision did not fully adjust immediately to the large reductions in local property tax rates that followed the passage of Proposition 13.

equilibration of household utility across places.

We use equations (5) and (6) of Gyourko and Tracy (1991) as the point of departure for our empirical specification. Using their notation, but generalizing to allow dependence of all variables on time (t), the reduced-form wage equation for individual i in state j is

$$\ln w_{ijt} = \beta_0 + \mathbf{X}_{it}\beta_1 + \mathbf{Y}_{it}\beta_2 + \mathbf{Z}_{jt}\beta_3 + u_{ijt}, \qquad u_{ijt} = \alpha_{jt} + \epsilon_{it} \tag{5}$$

Here, $\mathbf{X}_{it}$ is a vector of individual worker traits--such as age and educational attainment--that are correlated with worker productivity, $\mathbf{Y}_{it}$ is a vector of industry and occupational controls for worker quality, and $\mathbf{Z}_{jt}$ is a vector of observed state amenity and fiscal attributes. The contribution of unobserved locational characteristics to state-wide wages (the "group effect") is parameterized as the α_{jt} component of the error term, u_{ijt}. A reduced-form housing cost equation is defined similarly, as

$$\ln n_{ijt} = \gamma_0 + \mathbf{H}_{it}\gamma_1 + \mathbf{Z}_{jt}\gamma_2 + v_{ijt}, \qquad v_{int} = \delta_{jt} + \acute{\eta}_{it}, \tag{6}$$

where $\mathbf{H}_{it}$ is a vector describing the characteristics of the housing unit occupied by individual i.[5]

Gyourko and Tracy (1991) estimate equations (5) and (6) using microdata from the Census of Population and Housing for a single point in time. Given our emphasis on evolution in the quality-of-life and the lack of consistent time-series on individual worker and housing characteristics ($\mathbf{X}_i$, $\mathbf{Y}_{it}$, and $\mathbf{H}_{it}$) from this source, we find it convenient to separate the estimation of the nuisance parameters (β_1, β_2, γ_1) needed for quality-adjusting the wages and housing costs from the estimation of the parameters that describe amenity capitalization (β_3, α_{jt}, γ_2, δ_{jt}). We estimate the quality-adjustment parameters (β_1, β_2, γ_1) for a benchmark year with microdata from the 1990 Census of Population and Housing.[6] The wage and housing cost variables from the Census pertain to the year of 1989, and for these first-stage

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[5]In Gyourko and Tracy (1991), the dependent variable in the housing cost regression actually is housing *expenditures*, not the rental price of owner- and tenant-occupied housing. Similarly, our dependent variable in the benchmark Census year is housing expenditures. However, as explained below, the intertemporal variation in this dependent variable is derived only from intertemporal variation in the rental prices of quality-adjusted owner- and tenant-occupied housing. Therefore, we continue to speak of this dependent variable as if it is a housing rental price variable.

[6]We use a 1/1000 subsample of the Public Use Microdata A Sample. The wage variable is defined as total wage and salary earnings for the year of 1989, with the sample restricted to those persons 16 years or older, working, and reporting nonzero earnings, hours, and weeks worked. These restrictions result in a sample of 671,591 individuals for the wage equation. Following Gyourko and Tracy (1991) and others, the housing expenditure variable was derived from Census information on householder reports of gross rents (for renter-occupied units) and from owners' estimates of house value for 1989. In the latter case, owners' estimates of property value were converted to an annual rental equivalent measure using a measure of homeownership user costs.

regressions we collapse the state-specific components of wages and housing costs into fixed effect parameters (λ , μ):

$$\ln w_{ij89} = \beta_0 + \mathbf{X}_{i89}\beta_1 + \mathbf{Y}_{i89}\,\beta_2 + \lambda_{j89} + \epsilon_{i89}\,, \qquad \lambda_{j89} = \mathbf{Z}_{j89}\beta_3 + \alpha_{j89} \tag{7}$$

$$\ln n_{ij89} = \gamma_0 + \mathbf{H}_{i89}\,\gamma_1 + \mu_{j89}\,, \qquad \mu_{j89} = \mathbf{Z}_{j89}\,\gamma_2 + \delta_{j89} \tag{8}.$$

Results of the estimation of the quality-adjustment nuisance parameters are described in the appendix tables.

Given the first-stage estimation results, we construct quality-adjusted state-level average wages in the benchmark year by adjusting the actual state-level average wages ($\ln w_{.j89}$) by the inferred contributions of the differences between the actual state average worker characteristics ($\mathbf{X}_{.j89}$,$\mathbf{Y}_{.j89}$) and the U.S. national average worker characteristics ($\mathbf{X}_{..89}$,$\mathbf{Y}_{..89}$):

$$\ln w^*_{.j89} = \ln w_{.j89} - ((\mathbf{X}_{.j89} - \mathbf{X}_{..89})\,\hat{\beta}_1 + (\mathbf{Y}_{.j89} - \mathbf{Y}_{..89})\,\hat{\beta}_2\,) \tag{9}.$$

Quality-adjusted housing costs are similarly constructed, by imputing to the state the national average housing unit characteristics:

$$\ln n^*_{.j89} = \ln n_{.j89} - (\mathbf{H}_{.j89} - \mathbf{H}_{..89})\,\hat{\gamma}_1 \tag{10}.$$

Before and after the Census year, we use estimates of changes in quality-adjusted wages and housing costs from sources other than the Census. The changes in state-level quality-adjusted wages ($\Delta \ln w^*_{.\,jt}$) are constructed by estimating the counterpart to equation (7) for each year of the sample, using annual microdata from the Current Population Survey's (CPS) outgoing rotation file. The final time-series of state-level quality-adjusted wages are created by extrapolating the base-year 1989 quality-adjusted wages with these changes in quality-adjusted wages derived from the CPS.

The changes in state-level quality-adjusted housing costs ($\Delta \ln n^*_{.jt}$) are computed as a weighted average of changes in renters' costs and owners' costs, with weights reflecting the base-year share of housing expenditures due to each (tenure) type of housing occupant. For renters, expenditures are extrapolated by changes in the Bureau of Labor Statistics' (BLS) measure of the CPI for residential rent, adjusted so the cumulative growth rate of the rent series for each state matches the rate of change in the state's median rent as reported in the 1980 and 1990 Censuses. For owners, state-specific time-series on the user cost of housing are constructed.[7] In the user cost computation, state-level house price

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[7]Homeownership user costs display substantial cross-state and time-series variation. The user cost of homeownership (for state j in year t) utilized here is defined as
$uc_{jt} = [r_t + d + pt_j + m) - tx_{jt}(r_t + pt_j) - g_{jt}]HP_{jt}$, where r is the average of monthly rates on fixed-rate and conforming

variation over the sample period is based on the Fannie Mae-Freddie Mac repeat sales price indices.[8]

Our non-housing cost-of-living variable is derived from the American Chamber of Commerce Research Associations (ACCRA) publications of survey results on prices for specific, comparable items in more than 200 cities nationwide. For each city, we compute an index of the cost-of-living except housing, normalized so that the national average cost-of-living index is 100. The state-level indices are the averages of the non-housing cost-of-living indices for the cities in each state. The ACCRA index includes some widely traded commodities, where prices are set in national or international markets, and some other commodities with limited local trade areas. Accordingly, we reduce the expenditure weight on the cost-of-living excluding housing to the share of non-housing services in total personal consumption spending. That is, in the benchmark year of 1989, non-housing consumption expenditures on locally-traded items, $c^*_{..89}$, for the United States as a whole are set equal to the product of total consumption per household times this expenditure share. The state-level aggregates of the ACCRA indices are used to distribute such expenditures across localities and over time, giving us a full set of observations on the local cost-of-living except housing, $c^*_{.jt}$.

To facilitate intertemporal comparisons, the nominal wage and housing cost variables are deflated by the overall CPI for the nation. Hereafter, references to wages and housing costs are expressed in constant 1989 dollars. Simlarly, our quality-adjusted measure of the local cost-of-living excluding housing ($c^*_{.jt}$) is indexed to the level of expenditures on such items in 1989.

In the second stage of model estimation, the state-level time-series of quality-adjusted wages, housing costs, and non-housing cost-of-living are regressed on the set of amenity characteristics to determine the response of these variables to changes in amenities. The three-equation reduced form includes aggregates of equations (5) and (6) and a similar equation for the non-housing cost-of-living, generalized to include time fixed-effects through a vector of annual

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conventional mortgages (Freddie Mac), d is the rate of property depreciation (see Poterba [1991]), pt is the average property tax rate in the state on FHA loans, m is the maintenance rate (see Poterba [1991], tx is the marginal combined state and federal income tax rate, and g is the expected rate of capital gains. We let g vary across states and over time, based on the predictions of a univariate autoregressive model for the rate of house price change in each state. Our time- and state-varying user cost measure contrasts with the uniform 7.85 percent estimate of homeownership user costs utilized in the Blomquist et al (1988) and Gyourko and Tracy (1991) analyses.

[8]The Freddie Mac and Fannie Mae Conventional Home Mortgage Home Price Series comprises the only comprehensive and quality-adjusted state-level house price series available over the sample period. The price index is derived based upon a weighted repeat sales methodology applied to approximately 2 million repeat sales transactions occurring between 1975 and the mid-1990s.

dummy variables T_t :

$$\ln w^*_{.jt} = \beta_0 + \mathbf{Z}_{jt}\beta_3 + T_t\,\beta_4 + \alpha_{jt} + \epsilon_{.jt} \tag{11}$$

$$\ln n^*_{.jt} = \gamma_0 + \mathbf{Z}_{jt}\gamma_2 + T_t\gamma_3 + \delta_{jt} + \acute{\eta}_{.jt} \tag{12}$$

$$\ln c^*_{.jt} = \theta_0 + \mathbf{Z}_{jt}\theta_1 + T_t\,\theta_2 + \kappa_{jt} + v_{.jt} \tag{13}$$

As the data has been aggregated across individuals within the state, some of the parameters of interest are not identified unless further restrictions are imposed; in particular, the group effect parameters (α_j, δ_j , κ_j) are not separately identified from the coefficients on observed amenities (β_3, γ_2, θ_1). In our baseline variant of this second-stage analysis, which we term the "observed amenities model", we assume that the group effect is not a component of the quality of life and regress wages, housing costs, and the non-housing cost-of-living on an extensive vector of locational amenities $\mathbf{Z}_j$. By intention, the set of amenity variables is largely similar to that utilized in Blomquist et al (1988) and Gyourko and Tracy (1991). Amenity controls include weather and other climatic variables (precipitation, humidity, heating degree days, cooling degree days, wind speed, and sunshine), and recreation opportunities (a dichotomous variable indicating whether the state borders an ocean, the Gulf of Mexico, or the Great Lakes, a variable indicating percentage of state area covered by inland waterways, the percentage of state area in federal lands, and the number of visitors to state and national parks relative to state population).

In addition, we assume that households desire high environmental quality but also want the environment to be protected efficiently. Aside from the climactic variables, environmental quality is proxied by the number of hazardous waste sites in the state and two measures of air pollution, levels of ozone and carbon monoxide. To allow for joint measurement of environmental outcomes and environmental protection efforts, we also include a measure of the leniency of state environmental regulation--the composite score of the Green Policy Index. The analysis also includes a variety of state and local fiscal measures.[9] In particular, we include measures of income, sales, and property tax rates, as well as

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[9]Two points should be kept in mind when considering the effect of local fiscal conditions on the amenity value of a location. First, tax burdens and the levels of public good provision should be considered jointly, as higher taxes can be a means for financing additional desired public goods. Second, whether the whole menu of fiscal conditions has any net effect on the relative desirability of the location depends on whether the public goods are provided efficiently and priced (taxed) appropriately; given the likely deviations from this ideal, Gyourko and Tracy (1991) emphasize possible capture

estimates of the shares of state and local government expenditures in three government service categories: higher education, public welfare, and highways. Finally, the amenity vector includes average commute times as a measure of traffic congestion in the state, school quality (proxied by the mean student-teacher ratios in the state's public schools), and public safety (the rate of violent crimes per capita). Fuller descriptions and sources for the amenity variables are given in the appendix.

The set of amenity characteristics include some that are invariant over time and others that change in each year. The included controls for weather and climate (precipitation, humidity, heating degree days, cooling degree days, wind speed, and sunshine) as well as the categorical variable indicating proximity to an ocean or inland body of water are essentially time invariant and hence are entered for a single year. Some other locational controls, including number of hazardous waste sites, acreage in federal lands, visitors to state and federal parks, and the index of environmental regulatory leniency, similarly displayed limited intertemporal variation or were unavailable on a time-series basis. Amenity controls that vary both across states and over time include air pollution (the levels of ozone and carbon monoxide), commute times, state and local income, property, and sales tax rates, student-teacher ratios in the public schools, incidence of violent crime, and state and local government budgetary shares in the categories of post-secondary education, welfare, and highways.

U.S. averages of state-level trait values for 1981 and 1990 suggest some deterioration over time in many quality-of-life characteristics, while other amenities registered perceptible improvement (table 1). The average commute time rose somewhat over the sample period, in the wake of increases in population and urban congestion. Similarly, consistent with popular perceptions of declines in public safety, the rate of violent crime moved up by about one-fifth. However, with the coming of age of the baby-boom generation and the decline in school age children during the 1980s, some easing of student-teacher ratios was recorded. For the U. S. as a whole, the 1980s also witnessed some perceptible improvements in air quality. Shares of spending on public welfare programs increased some over the 1980s, as did tax rates; however, income taxes rose less quickly than property taxes and sales and other taxes.

For most of the characteristics, there is substantial heterogeneity across states in terms of how amenities have changed (right-hand columns of table 1). Commute times decreased in some states and increased in others. One state

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of amenity values by the public sector.

(Maine) managed to lower its student-teacher ratio by 6.7 students per teacher, while another (Wisconsin) added 1.8 students per teacher. Some state taxation structures also were modified over this sample period. For example, Massachusetts enacted a large property tax rate reduction that contributed to a drop of $14.56 in property taxes per $1000 of state personal income; in contrast, this measure of property tax rates increased by about $11 in New Hampshire and Texas. States also chose different spending priorities; for example, the share of state and local government expenditures on higher education was reduced 3.8 percentage points in Oklahoma, whereas the higher education spending share increased 4.5 percent in Maryland. Illinois experienced the largest increase in violent crime; in contrast, violent crime dropped sharply in Nevada. Although air pollution levels rose in some eastern seaboard states, there were sizable declines in ozone and carbon monoxide levels in many western states with large metropolitan areas, particularly California and Colorado.

These statistics demonstrate the large temporal and spatial variation in particular amenities. However, without a relative valuation of the amenities and an accounting for the simultaneous evolution in the amenities, we cannot determine how these developments affected the aggregate quality-of-life. Thus, we turn now to the reduced form wage, housing expenditure, and non-land cost-of-living regressions to determine which amenity characteristics most affected measured quality of life and where, geographically, the most significant changes in capitalized amenity values occurred.

IV. Results

The basic results from estimating the three-equation form of the standard amenities model (without group effects) over the 1981 to 1990 period are shown in table 2. The first three columns report the estimated amenity coefficient vectors β_3, $\hat{\gamma}_2, \theta_1$ from the wage equation (11), the housing expenditure equation (12), and the non-housing cost-of-living equation (13), respectively. Overall, the amenities (together with the annual time dummies) explain about 70 to 75 percent of the variation in (quality-adjusted) housing costs, wages, and the non-housing cost-of-living. As is evident from the third column of the table, substantial capitalization of local amenities appears to occur via the non-housing cost-of-living, in contrast to the assumption implicit in most previous representations of this model. In particular, about three-fourths of the estimated coefficients on the included amenity variables are statistically significantly different from zero in the non-housing cost-of-living regression.

For each amenity (indexed by k), the fourth data column of table 2 presents the full implicit price calculated as:

$$FP_k = n^*_{\cdots} \hat{\beta}_{3k} + c^*_{\cdots} \hat{\theta}_{1k} - w^*_{\cdots} \hat{\gamma}_{2k} \ , \tag{14}$$

where $n^*_{\cdots}$, $c^*_{\cdots}$, and $w^*_{\cdots}$ are the full-sample means of the housing expenditure, non-housing cost-of-living, and wage variables.

The general pattern of estimated full implicit prices is consistent with our *a priori* beliefs about whether a characteristic is an amenity or disamenity. In the few instances where an implicit price estimate has an unanticipated sign, the standard errors indicate that this might plausibly be due to imprecision in the estimation. Certain climatic and recreational variables add appreciably to the amenity value of a place: location on a coast, an abundance of inland water area, public stewardship of federal lands, and access to national parks. In contrast, other climatic variables take away from the amenity value of a place: high levels of precipitation, humidity, temperature extremes (heating and cooling degree days), or windy conditions.

Soil pollution, as measured by the number of hazardous waste sites, is a disamenity, but the estimated effect is small. The results do not lend support to arguments that tough environmental regulations are an inefficient way to achieve desired environmental outcomes; given the achieved environmental outcomes, the stringency or leniency of environmental regulation has no perceptible compensating differential effect.

The general pattern of estimated full implicit prices on time-varying state traits also is sensible. Air quality is estimated to be very important to households' evaluations of the desirability of a location; higher levels of either ozone or carbon monoxide pollution are significant disamenities. Congestion, as measured by commuting times, is undesirable. Poorer school quality, as proxied by a high student-teacher ratio, also is a disamenity.

As expected, holding constant state public service levels such as school quality, higher state and local government income tax rates require a compensating differential, and the negative estimated full implicit price is clearly distinguishable from zero. In contrast, the implicit prices on property tax rates and sales and other taxes are estimated less precisely. Households appear to prefer that the composition of state and local government expenditures be tilted towards public welfare and public infrastructure, such as highways. The estimate of the implicit price for the higher education share of state and local government budgets is not statistically different from zero.

The attributes and implicit prices imply a broad range of capitalized amenity values. For example, the full sample average for precipitation is 35 inches per year,[10] but precipitation ranges from a low of 7 inches per year in Nevada to a high of 64 inches per year in Alabama. The full implicit price estimates indicate that each inch of rainfall per year commands $12 in compensating differential, and the range of attribute values implies that an Alabama resident's quality-of-life is held down by $781 (1989 dollars), about $690 more than the $91 compensating differential for the Nevada resident (final columns of table 2). The widest range of contributions from an amenity characteristic is associated with winter temperature extremes; households in the state with the need for the most heating effort (North Dakota at 8968 heating degree days) are estimated to be willing to sacrifice $15,642 in wages if they could avoid all such cold weather extremes, as in Hawaii, which has 0 heating degree days.

To aggregate across amenities, the set of full implicit prices and values of observed amenities are combined in state-level quality-of-life (QOL) indexes for each year t:

$$QOL_j = \Sigma_k FP_k Z_{kjt} \qquad j = 1,...,50;\ t = 1981,\ 1990. \qquad (15)$$

The difference in the value of the index between two states is a measure of the composite premium that the average household pays (through lower wages or price markups on housing and other locally-traded commodities) to live in the higher amenity state. Although the index is denominated in constant 1989 dollars, we do not make intertemporal comparisons of the index values because we have not included any part of the contributions of the time fixed effects--which might partially reflect the evolution of the aggregate national quality-of-life--in the indexes.

The cross-sectional pattern of estimated quality-of-life ranks shows some similarities to those in the preceding literature on compensating differentials. Densely populated industrialized states--including the midwestern states of Indiana, Illinois, and Michigan and the eastern seaboard states of New York, New Jersey, and Maryland--score relatively low in terms of estimated quality-of-life (table 3). In the metropolitan area results of Gyourko and Tracy (1991), a somewhat similar pattern of low-rankings for industrialized midwestern and eastern cities appeared. The Greenwood et.al. (1991) aggregation to the state level of the Blomquist et.al. (1988) metropolitan area results--which excluded the effect of fiscal conditions--also showed midwestern states such as Indiana, Illinois, and Michigan to be low-ranked, although densely populated eastern states generally were ranked toward the middle of the range. In extending the analysis beyond states with large metropolitan areas, we find that less densely populated, rural western states such as

[10]This and other full-sample means are shown in parentheses below the descriptions of the state traits in table 2.

Montana, South Dakota, and Wyoming are ranked highly in the estimated quality-of-life.

In general, the rural western states that were highly ranked in 1981 remained highly ranked throughout the decade. Similarly, the low rankings for industrialized midwestern and eastern states were relatively stable over time. Contrary to much public discussion on the subject, the relative quality-of-life in California also changed little over the decade.[11] Among the states which exhibited a significant evolution in the relative quality-of-life, Alaska and Arizona stand out with large improvements in the rankings, while New Hampshire and a few western states--Idaho, Nevada, and New Mexico--are estimated to have deteriorated noticeably in the quality-of-life rankings.

By studying the evolution of particular characteristics, evaluated at their implicit prices, we have discovered some interesting patterns in how changes in amenities have contributed to large improvement or deterioration in the quality-of-life ranks for individual states. In particular, in contrast to the conventional wisdom, population growth, per se, does not explain the evolution of rankings. Although many of the states with deteriorating quality-of-life ranks faced the pressures of rapid population growth during the 1980s (e.g., Nevada, New Hampshire, Georgia, Utah, Washington, New Mexico, and Hawaii), other states with large improvements in estimated quality-of-life (Alaska, Arizona, Florida, and Colorado) also were among the states with the largest rates of population increase. Rather, the key to maintaining or increasing quality-of-life appears to have been in how governments managed the population growth.

For example, the states which experienced deteriorating quality-of-life rankings tended to cut back on the share of state and local government expenditures devoted to highways and transportation infrastructure, and traffic congestion and average commuting times increased. Furthermore, when this population growth and increased congestion occurred in an area with initially relatively good air quality, air pollution control efforts--such as mandates to use cleaner-burning fuels--were allowed to remain more lax, and carbon monoxide pollution in particular increased, relative to other states. In contrast, some other fast-growing states with very stringent air quality management regimes--particularly Arizona, California, and Colorado--benefitted from improved air quality.

Although these patterns are most evident for individual states, we present some summary statistics on the

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[11] Although California experienced substantial net out-migration of population in recent years, that flow derived largely from the lack of job opportunities in the state, given defense-sector downsizing and other sources of unemployment, not from deterioration in the quality-of-life (Gabriel, Mattey and Wascher (1995)).

contribution of amenities to the evolution of quality-of-life ranks in table 4. For this summary, we have categorized a state as having a "stable" quality-of-life ranking if it changed less than five places in the rankings between 1981 and 1990. Among other states, ten experienced a large deterioration in relative quality-of-life (increased ranks) and eleven experienced a large improvement (decreased ranks). On average across the states in the improving group, the rank decreased a bit over nine places, and the average change in rank for the deteriorating group was about ten places.

In the group with deteriorating quality-of-life ranks, the attribute with the largest average contribution to the deterioration was state and local government expenditures on highways, which accounted for a movement of 2.7 places in the ranks, on average.[12] Increased commuting times, higher carbon monoxide levels, and a lower share of state and local government spending on public welfare also were large sources of deterioration in quality-of-life for these states. Each of the states of New Hampshire, New Mexico, and Nevada exhibited the prototypical pattern of deteriorating quality-of-life ranks from a reduced highway spending share, increased commuting times, and higher carbon monoxide levels.

The patterns among states with improving quality-of-life ranks were more diverse. On average, an increased budget share for public welfare was an important contributor, but this was not a broad-based phenomenon. The large average contribution of income tax reductions was similarly narrow-based; Alaska households benefitted from the elimination of the income tax there, but overall tax revenues and government services were maintained by a shift towards petroleum-related taxes in the sales and other taxes category. Large declines in carbon monoxide levels noticeably boosted the quality-of-life in Arizona and Colorado. In South Carolina, the improved quality-of-life owed to better schools and a broad-based realignment of the state and local government expenditure mix.

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[12]As an illustration of how we calculated the contribution of a specific amenity to the change in a state's rank, consider the example of how much increased commuting times affected California's quality-of-life rank. To compute this, we evaluated California's quality-of-life index at the 1990 commuting times but placed all other California traits at simulated values which move forward from the 1981 California values for each trait by the amount which that trait changed in the U.S. as a whole (third data column of table 1). Similarly, amenity traits (including commuting times) in other states were rolled forward from their 1981 values by the national average changes in the traits.

V. Robustness and Other Results

Most of our main results are robust to reasonable alternative specifications of the model, but we did notice some interesting sensitivities. Our findings that increased commuting times and air pollution in states with rapid population growth and relatively low highway spending were associated with deteriorating quality-of-life rankings were relatively robust to various specification choices. However, we noticed that the interpretations of the evolution of some state traits--such as the violent crime rate--were sensitive to whether or not other closely-related state traits--such as spending on prisons--were included in the model.

For example, although prison spending and crime rates are positively correlated in the cross-section dimension, the state-specific time-series patterns of prison spending and crime rates are somewhat heterogeneous. Some states dealt with the potential for increased crime by increasing law enforcement efforts, thus letting prison populations swell, but preventing the actual crime rates from escalating further. Other state and local governments have managed to maintain relatively low budget shares on prison spending. We found that if the state and local government expenditure share on correctional facilities (prisons) is included in the model, it receives a large implicit price which clearly differs from zero, and the model implies that increased prison spending (or the underlying threat of crime driving the increased spending) is an important contributor to the deterioration in quality-of-life in some states.

In addition, we considered including a measure of the prices paid for electricity by end-users in each state as a locational characteristic, under the reasoning that the relative efficiency of regulated monopolies or relative endowments of resources facilitating low-cost (hydroelectric) or high-cost (nuclear) electric power production could be an important place-specific characteristic. If such an electricity price variable is included in the model, it receives a large, negative full implicit price, and the large increases in relative electricity prices in some states contribute noticeably to a deterioration in their relative quality-of-life. Survey evidence suggests that relative electricity prices are high on the list of most important factors for manufacturing firm location choice, and our results in this regard suggest that such productive disamenities are capitalized in the local wages and prices faced by households. However, given our inability to measure and reflect a full range of factors affecting firm productivity and costs, we chose to omit this variable from the final specification.

These examples illustrate that despite the extensive set of amenities which we (and others) have included in the quality-of-life calculations, we undoubtedly have excluded--usually for data reasons--other location-specific attributes that might influence either overall rankings or the implicit price estimates for particular amenities. One suggested solution to this omitted variables problem is to include group fixed effects in the quality-of-life estimates (see, for example, Gyourko and Tracy (1991)).[13] Of course, as these authors also point out, the use of fixed effects has its own problems because there could be location-specific attributes that are correlated with the dependent variables but should not be treated as influencing the quality-of-life estimates. For example, if the omitted variables affect wages because they pertain to the quality of the workforce (i.e., they are missing elements of the $\mathbf{X}_{it}$ or $\mathbf{Y}_{it}$ variables of equation (5)) , then they should not be treated as amenities in calculating the quality-of-life indices. In any event, it seems useful to examine the sensitivity of our results to the inclusion of fixed effects, both in terms of the rankings and their evolution through time.

As noted above, the fixed effect parameters (α_j, δ_j , κ_j) are not separately identified from the coefficients on observed amenities (β_3, γ_2 , θ_1) without further restrictions, and thus in our baseline model we assumed that the fixed effects were absent and regressed wages, housing costs, and the non-housing cost-of-living on an extensive vector of locational amenities $\mathbf{Z}_j$. Alternatively, one can achieve identification by deleting the time-invariant amenities (such as whether a state is located on a coast) from $\mathbf{Z}_j$ and restricting the fixed effect parameters to vary only over location ($\alpha_j=\alpha_j$, $\delta_j = \delta_j$, $\kappa_j = \kappa_j$) . Employing these restrictions, we have re-estimated equations (11) through (13) and computed quality-of-life indices using an alternative version of the definition of full-implicit prices (14) that includes the fixed effects coefficients as capitalized amenities.

The rankings we obtain with such state-level fixed effects are shown in table 5.[14] As is evident from

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[13]In addition to the OLS type of estimates provided by Blomquist et al, Gyourko and Tracy (1991) use a random effects estimator with and without group effects for explaining wages and housing expenditures of individuals within and between metropolitan areas. Here, we continue to use the terminology "group effect" even though we implement a fixed effects, not a random effects, approach to estimation. The disadvantage of the random effects approach is that the unknown group effects parameters must be orthogonal to the included regressors in order for the parameters of interest to be estimated consistently.

[14]We implement this group (fixed) effect version of the model by subtracting the average quality-adjusted wage level in each state from the sum of average quality-adjusted housing expenditures and expenditures on (local) cost-of-living except housing and treating this as a quality-of-life index. In terms of the implied rankings, this is equivalent to estimating the reduced form equations (11) through (13) with the terms involving observed amenities (Z_j) suppressed,

comparing tables 3 and 5, quality-of-life rankings using the two methods are highly correlated, with a Spearman rank correlation over the entire sample of .85. As in the observed amenities specification, mountain states such as Montana and Wyoming rank high, while industrialized states generally rank low. In contrast, the rankings do not change as much in the fixed effects version as in the observed amenities version of the model. Although changes in rankings are positively correlated across the models (with a correlation coefficient of .35), many fewer states evolve more than 2 places in the ranks. This occurs because there is some sensitivity in the estimated full implicit prices to including fixed effects. In particular, although the estimated implicit prices on commuting times, carbon monoxide, and ozone pollution are similar across the models, most of the estimated full implicit prices on fiscal variables change quite a bit. Accordingly, states where increased congestion and air pollution accounted for a significant portion of the deterioration in quality-of-life (e.g., Nevada and New Hampshire) still show such deterioration in the fixed effects model. However, states where the change in rankings was driven by the evolution of fiscal conditions (e.g., South Carolina) often show a different pattern in the fixed effects model.[15]

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and redefining equation (15) in terms of the group effect parameters.

[15]Another aspect of quality-of-life models that has been criticized in recent years is their reliance on the assumption that labor and non-traded goods markets are in equilibrium. In particular, numerous studies have found that migration can, at least in part, be explained by differences across locations in wages and house prices, suggesting that such differences may not be immediately or fully offset by differences across locations in amenities. Greenwood et. al. (1991) address this issue by adjusting the quality-of-life estimates for the disequilibrium implied by nonzero net migration flows; their results for 1980 suggest that most states were in disequilibrium in that year, but that adjusting for that disequilibrium had little effect on the quality-of-life estimates. We experimented with this issue by entering a net migration variable directly into the wage, housing expenditure, and non-land cost-of-living equations and interpreting the contribution of the net migration variable to the fit of the regressions as a measure of the extent to which the wages and non-traded goods expenditures were in disequilibrium. Similar to Greenwood et. al. (1991), we found that a migration disequilibrium adjustment has little effect on the implied quality-of-life rankings.

VI. Conclusion

State and local government policymakers continually struggle with the question of how to maintain a high quality-of-life in the presence of rapid population growth. Although no single piece of research will be able to establish a definitive answer to this question, this paper illustrates that the compensating differential paradigm is a useful framework for developing empirically-grounded, albeit tentative answers.

In providing the first application of the compensating differential paradigm to the evaluation of the extent and sources of evolution in state quality-of-life, we find that states had mixed results in adapting to the stresses of rapid population growth. In some states with rapid population growth, the quality-of-life has remained generally good. However, in other fast-growing states, the evidence points to a substantial deterioration in relative quality-of-life. Our model estimates suggest reduced spending on highways and increased traffic congestion and air pollution have been the most important contributors to the deterioration of quality-of-life in declining states. In contrast, states with improved quality-of-life ranks ascended for a variety of reasons, including reducing state and local government income tax burdens, improved air quality, increased highway spending, and reduced commute times.

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TABLE 1
Evolution of Amenity Characteristics

	U.S. Average Characteristic			*Maximal Change in State Characteristic*	
Time-Varying State Trait	1981	1990	Change	Decrease	Increase
Commuting Time (minutes)	19.82	20.35	+.53	-2.16	2.25
Violent Crime Rate (crimes per 100,000 population)	447	534	+87	-295	523
Air Quality-Ozone (parts per million)	.12	.11	-.01	-.08	.02
Air Quality-Carbon Monoxide (parts per million)	9.26	6.61	-2.65	-18.5	1.5
Student-teacher ratio (students per teacher)	18.32	17.25	-1.07	-6.70	1.80
State and local taxes ($ per $1000 of personal income)					
on income	23.53	24.90	+1.37	-123.78	13.25
on property	30.04	32.90	+2.86	-14.56	11.49
on sales and other	48.04	50.05	+2.01	-86.83	17.93
State and local expenditures (proportion of general expenditures)					
on higher education	.104	.087	-.014	-.038	.045
on public welfare	.111	.131	+.020	-.037	.055
on highways	.096	.087	-.009	-.041	.017

Source: Calculations by the authors.

TABLE 2

Parameter Estimates, Full Implicit Prices, and Quality-of-Life Index Components

Time-Invariant State Trait	Housing Expenditure Equation	Wage Equation	Nonland Cost of Living Equation	Full Implicit Price	QOL Component Minimum	QOL Component Maximum
Precipitation	-.0028	-.0003	-.0003	-12	-91	-781
(35 inches per year)	(.0015)	(.0007)	(.0004)	(24)		
Humidity	.0083	.0034	.0004	-51	-1888	-4056
(65.5 percent)	(.0022)	(.0011)	(.0006)	(33)		
Heating Degree Days	.00002	.00006	.00002	-1.74	0	-15642
(5091 per year)	(.00001)	(.00000)	(.00000)	(.19)		
Cooling Degree Days	.00002	.00006	.00004	-1.65	0	-7364
(1215 per year)	(.00002)	(.00001)	(.00000)	(.31)		
Wind Speed	.0035	.0063	-.0036	-216	-1361	-2809
(9.36 miles per hour)	(.0061)	(.0031)	(.0016)	(92)		
Sunshine	.0173	.0058	.0016	-54	-1622	-4650
(59.4 percent of possible)	(.0026)	(.0013)	(.0007)	(39)		
Coast	.1216	.0320	.0275	27	0	27
(.52, =1 if state on coast)	(.0254)	(.0128)	(.0068)	(385)		
Inland Water	.0044	-.0048	.0041	226	48	2945
(2.7 percent of land area)	(.0041)	(.0021)	(.0011)	(62)		
Federal Land	.0058	.0052	.0010	34	14	2761
(15.3 percent of land area)	(.0009)	(.0005)	(.0002)	(14)		
Visitors to National Parks	-.0005	-.0002	-.0001	1.6	0	1502
(148 per 100 population)	(.00006)	(.00003)	(.00002)	(.9)		
Visitors to State Parks	.00010	.00005	.00006	-.5	-13	-913
(365 per 100 population)	(.00004)	(.00002)	(.00001)	(.6)		
Number of hazardous waste sites	-.0010	-.0000	.0001	-5	-5	-577
(25.2 sites per state)	(.0005)	(.0002)	(.0001)	(7)		
Environmental Regulation Leniency	-.00006	-.00001	-.00000	.0	18	77
(2200 on Green Policies Index)	(.00003)	(.00001)	(.00000)	(.4)		

TABLE 2 (continued)
Parameter Estimates, Full Implicit Prices, and Quality-of-Life Index Components

Time-Varying State Trait	Housing Expenditure Equation	Wage Equation	Nonland Cost of Living Equation	Full Implicit Price	QOL Component Minimum	QOL Component Maximum
Commuting Time	.0479	.0336	.0068	-736	-9565	-21589
(20.1 minutes)	(.0049)	(.0025)	(.0013)	(74)		
Violent Crime Rate	-.00035	-.00011	-.00009	.4	21	554
(475 per 100,000 population)	(.00006)	(.00003)	(.00002)	(.9)		
Air Quality-Ozone	1.15	1.07	.12	-26809	-804	-7388
(.12 parts per million)	(.33)	(.17)	(.09)	(5002)		
Air Quality-Carbon Monoxide	-.004	.003	-.003	-160	-160	-4486
(8.2 parts per million)	(.003)	(.002)	(.001)	(46)		
Student-teacher ratio	-.0204	-.0000	-.0046	-187	-2453	-4758
(17.6 students per teacher)	(.0060)	(.0030)	(.0016)	(90)		
State and local taxes						
on income	-.0008	.0007	.0001	-28	0	-3916
($24 per $1000 of personal income)	(.0007)	(.0003)	(.0002)	(10)		
on property	.0043	-.0000	.0008	23	226	2040
($32 per $1000 of personal income)	(.0010)	(.0030)	(.0003)	(15)		
on sales and other	.0048	.0014	.0012	-3	-52	-614
($51 per $1000 of personal income)	(.0005)	(.0003)	(.0001)	(8)		
State and local expenditures						
on higher education	-2.84	-.71	-.81	-3059	-132	-491
(.10 of general expenditures)	(.48)	(.24)	(.13)	(7204)		
on public welfare	-1.86	-1.55	-.31	36455	1495	8140
(.12 of general expenditures)	(.40)	(.20)	(.11)	(6063)		
on highways	-1.46	-1.86	-.76	46760	1931	7840
(.09 of general expenditures)	(.55)	(.28)	(.15)	(8379)		
Memo: Goodness-of-fit (R^2)	.73	.74	.70			

Source: Calculations by the authors.
Notes:
a. Conventional standard errors are in parentheses. The coefficient estimates and quality-of-life components are computed from OLS regressions of the logarithms of the real wage, housing expenditures and cost-of-living except housing on the variables shown and time fixed effects by year (not shown).

TABLE 3
Quality-of-Life Index Ranks by State, 1981 and 1990

	Quality-of-Life Rank				*Quality-of-Life Rank*		
State	1981	1990	*Change*	*State*	1981	1990	*Change*
Alabama	29	31	+2	Montana	5	3	-2
Alaska	25	6	-19	Nebraska	10	17	+7
Arizona	35	16	-19	Nevada	12	27	+15
Arkansas	6	5	-1	New Hampshire	20	37	+17
California	41	42	+1	New Jersey	44	46	+2
Colorado	46	40	-6	New Mexico	7	19	+12
Connecticut	30	29	-1	New York	50	50	0
Delaware	37	35	-2	North Carolina	22	21	-1
Florida	21	12	-9	North Dakota	15	8	-7
Georgia	31	36	+5	Ohio	38	33	-5
Hawaii	17	26	+9	Oklahoma	24	23	-1
Idaho	4	4	0	Oregon	19	15	-4
Illinois	47	49	+2	Pennsylvania	39	38	-1
Indiana	45	44	-1	Rhode Island	9	9	0
Iowa	14	18	+4	South Carolina	26	20	-6
Kansas	18	22	+4	South Dakota	1	2	+1
Kentucky	23	24	+1	Tennessee	34	30	-4
Louisiana	13	7	-6	Texas	27	25	-2
Maine	8	10	+2	Utah	36	41	+5
Maryland	48	47	-1	Vermont	11	13	+2
Massachusetts	28	28	0	Virginia	32	32	0
Michigan	49	48	-1	Washington	33	39	+6
Minnesota	42	45	+3	West Virginia	16	14	-2
Mississippi	3	11	+8	Wisconsin	40	34	-6
Missouri	43	43	0	Wyoming	2	1	-1

Source: Calculations by the authors from the results in table 2.

TABLE 4
Contributions of Amenities to the Evolution of Quality-of-Life Ranks, 1981-90

Time-Varying State Trait	——*States with QOL Ranks which are*—— Deteriorating(+)	Improving(-)	Stable
All amenities	10.1	-8.9	-.1
Commuting Time	2.2	-1.5	-.2
Violent Crime Rate	.1	-.2	.0
Air Quality-Ozone	-.3	.1	-.1
Air Quality-Carbon Monoxide	.7	-1.1	.2
Student-teacher ratio	.2	.0	.0
State and local taxes			
on income	.2	-1.5	.2
on property	-.1	-.4	.1
on sales and other	.0	.0	-.0
State and local expenditures			
on higher education	.0	-.1	-.0
on public welfare	2.4	-2.8	.4
on highways	2.7	-.9	-.2

Source: Calculations by the authors.

TABLE 5
Quality-of-Life Index Ranks by State, 1981 and 1990
Fixed Effects Specification

	Quality – of – Life Rank				*Quality – of – Life Rank*		
State	1981	1990	*Change*	*State*	1981	1990	*Change*
Alabama	20	18	-2	Montana	1	1	0
Alaska	30	32	+2	Nebraska	29	29	0
Arizona	19	21	+2	Nevada	23	28	+5
Arkansas	14	14	0	New Hampshire	12	16	+4
California	35	39	+4	New Jersey	49	49	0
Colorado	39	37	-2	New Mexico	9	9	0
Connecticut	44	44	0	New York	45	45	0
Delaware	34	33	-1	North Carolina	21	23	+2
Florida	11	11	0	North Dakota	16	13	-3
Georgia	31	30	-1	Ohio	43	43	0
Hawaii	33	34	+1	Oklahoma	8	10	+2
Idaho	2	3	+1	Oregon	13	12	-1
Illinois	48	48	0	Pennsylvania	42	41	-1
Indiana	38	36	-2	Rhode Island	17	19	+2
Iowa	22	20	-2	South Carolina	15	17	+2
Kansas	26	25	-1	South Dakota	3	4	+1
Kentucky	24	22	-2	Tennessee	28	26	-2
Louisiana	10	8	-2	Texas	25	24	-1
Maine	4	5	+1	Utah	40	40	0
Maryland	50	50	0	Vermont	6	6	0
Massachusetts	46	47	+1	Virginia	41	42	+1
Michigan	47	46	-1	Washington	27	27	0
Minnesota	37	38	+1	West Virginia	18	15	-3
Mississippi	7	7	0	Wisconsin	36	35	-1
Missouri	32	31	-1	Wyoming	5	2	-3

Source: Calculations by the authors.

APPENDIX TABLE A.1
Regression for Housing Quality Adjustments

Variable	Coefficient Estimate	Standard Error
Age of Structure	-.008	.0001
Number of Rooms	.162	.0023
Number of Bedrooms	.033	.0042
D(Mobile Home)	-1.056	.0238
D(One-family house detached)	.007	.0227
D(One-family house attached)	.022	.0247
D(2 Apartments in building)	.093	.0239
D(3-4 Apartments in building)	.050	.0240
D(5-9 Apartments in building)	.025	.0241
D(10-19 Apartments in building)	.081	.0242
D(20-49 Apartments in building)	.075	.0250
D(50 or more apartments in building)	.083	.0248
D(House on less than 1 acre)	-0.048	.0063
D(Business or medical office on property)	.079	.0177
D(Complete plumbing facilities)	.551	.0295
D(Complete kitchen facilities)	.108	.0315
D(House or apartment part of condominium)	.148	.0115
Intercept	6.802	.0610
D(State where located)		
—South Dakota (minimum)	-.323	.0623
—North Dakota	-.266	.0652
—....		
—California	1.108	.0474
—Hawaii (maximum)	1.224	.057
Memo: Goodness-of-fit (R^2)	.50	

Source: Calculations by the authors from the Census of Population and Housing Public Use Microdata A (5 percent) sample, subsetted to a 1 in 1000 (.1 percent) subgroup.

Notes:

a. The dependent variable is the logarithm of the annual amount of housing expenditures per household. The regression uses observations on 82,225 housing units which report either a gross rent figure (for renter-occupied units) or an owner's estimate of value. Owners' estimates of value are converted to an annual rental-equivalent using the .0785 user cost estimate of Peiser and Smith (1985). Top-coded owners' value estimates are imputed at 500,000 dollars.

b. The regression includes 50 state dummy variables; only selected minimal and maximal coefficient estimates on these dummy variables are shown. The reference category subsumed in the intercept is for the state of Wyoming.

APPENDIX TABLE A.2
Regression for Worker Quality Adjustments

Variable	Coefficient Estimate	Standard Error
D(Age of worker)		
20-24 years old	.1175	.0036
25-34 years old	.2999	.0035
35-44 years old	.3891	.0036
45-54 years old	.4460	.0038
55-64 years old	.4397	.0038
65 years or older .3790	.0045	
D(Education of worker)		
High school graduate	.0983	.0022
Some college	.1371	.0024
College degree (bachelors)	.3338	.0032
Master level degree	.5440	.0045
Professional degree	.5834	.0092
Ph.D. degree	.5829	.0119
D(Disabled with work limitations)	-.0542	.0033
D(Current School Enrollment)	-.0571	.0025
D(Does not speak English well)	-.0323	.0029
Intercept	1.6501	.0197
D(State where located)		
— Utah (minimum)	-.3406	.0151
— South Dakota	-.3224	.0178
—		
— Connecticut	.1949	.0149
— Washington, D.C. (maximum)	.4676	.0189
D(Occupation)		
— Farm operators (minimum)	-.6230	.0210
— Sales related occupations	-.3815	.0277
—		
— Physicians and dentists	.4642	.0234
— Optometrists and podiatrists (maximum)	.4897	.0500
D(Industry)		
— Miscellaneous services (minimum)	-.1922	.0191
— Museums, art galleries and zoos	-.1070	.0281
—		
— Oil and gas extraction (maximum)	.7322	.0147
— Railroads (maximum)	.7331	.0204
Memo: Goodness-of-fit (R^2)	.32	

Source: Calculations by the authors from the Census of Population and Housing Public Use Microdata A (5 percent) sample, subsetted to a 1 in 1000 (.1 percent) subgroup.

Notes:

a. The dependent variable is the logarithm of the hourly amount of wage and salary earnings per person. The regression uses observations on 671,591 persons who are 16 or older, worked in the reporting year, and reported nonzero earnings, hours and weeks of work. Annual reported wage and salary income in 1989 is converted to an hourly basis by dividing by the product of weeks worked and usual hours worked per week in 1989.

b. The regression includes 53 dummy variables for occupational categories, 78 dummy variables for industry affiliations, and 50 state dummy variables; only selected minimal and maximal coefficient estimates on these dummy variables are shown. The reference categories subsumed in the intercept are for the youngest workers, with minimal educations, in military occupations in the active duty military, and in the state of Wyoming.

APPENDIX TABLE A.3
Description of Data Sources

Variable Name and Source
Precipitation
U.S. National Oceanic and Atmospheric Administration, *Climatography of the United States*, No. 81
as published in The Statistical Abstract of the United States, 1993, Table No. 387
Humidity
U.S. National Oceanic and Atmospheric Administration, *Comparative Climactic Data*, annual
as published in The Statistical Abstract of the United States, 1994, Table No. 387
Heating Degree Days
U.S. National Oceanic and Atmospheric Administration, *Climatography of the United States*, No. 81
as published in The Statistical Abstract of the United States, 1993, Table No. 391
Cooling Degree Days
U.S. National Oceanic and Atmospheric Administration, *Climatography of the United States*, No. 81
as published in The Statistical Abstract of the United States, 1993, Table No. 392
Wind Speed
U.S. National Oceanic and Atmospheric Administration, *Comparative Climactic Data*, annual
as published in The Statistical Abstract of the United States, 1994, Table No. 387
Sunshine
U.S. National Oceanic and Atmospheric Administration, *Comparative Climactic Data*, annual
as published in The Statistical Abstract of the United States, 1994, Table No. 387
Coast
Specified by the authors to equal one if the state abuts an ocean, the Great Lakes, or the Gulf of Mexico.
Inland Water
Boating Industry magazine, January 1990
as published in Hall and Kerr, *1991-1992 Green Index*, p. 109
Federal Land
U.S. Department of the Interior, Bureau of Land Management, "Public Lands Statistics, 1989"
as published in Hall and Kerr, *1991-1992 Green Index*, p. 109
Visitors to National Parks
U.S. Department of the Interior, National Park Service, "National Park Service Statistical Abstract, 1989"
as published in Hall and Kerr, *1991-1992 Green Index*, p. 109
Visitors to State Parks
National Association of State Park Directors, "Annual Information Exchange", April 1990
as published in Hall and Kerr, *1991-1992 Green Index*, p. 109
Number of hazardous waste sites
U.S. Environmental Protection Agency, press release.
as published in The Statistical Abstract of the United States, 1994, Table No. 372
Environmental Regulation Leniency
composite score on Green Policies (state policy initiatives and leadership in Congress)
as tabulated by and published in Hall and Kerr, *1991-1992 Green Index*, p. 5

APPENDIX TABLE A.3 (continued)
Description of Data Sources

Variable Name and Source
Commuting Time
U.S. Bureau of the Census, *Census of Population and Housing*, 1980 and 1990
as published in the United States Summary volume of General Social and Economic Characteristics for 1980,
table 238 and for 1990 as published in The Statistical Abstract of the United States, 1993, Table No. 1017
Violent Crime Rate
U.S. Federal Bureau of Investigation, *Crime in the United States*, annual, various issues
as published in The Statistical Abstract of the United States, various issues.
Air Quality-Ozone
U.S. Environmental Protection Agency, *National Air Quality and Emissions Trends Report*, annual,
various issues.
Air Quality-Carbon Monoxide
U.S. Environmental Protection Agency, *National Air Quality and Emissions Trends Report*, annual,
various issues.
Student-teacher ratio
The numerator is public elementary and secondary school enrollment from U.S. National Center for
Education Statistics, *Statistics of State School Systems*, biennial and
Digest of Education Statistics, annual
as published in The Statistical Abstract of the United States, various issues
The denominator is number of public elementary and secondary school teachers from National
Education Association, *Estimates of School Statistics*, various issues
as published in The Statistical Abstract of the United States, various issues
State and local taxes
on property, income and sales and other
U.S. Bureau of the Census, *Governmental Finances*, various issues
State and local expenditures
on higher education, public welfare, highways, and corrections
U.S. Bureau of the Census, *Governmental Finances*, various issues
Cost-of-living except housing
American Chamber of Commerce Researchers Association, *ACCRA COST OF LIVING INDEX* serial,
issues for third quarters of 1981 through 1990