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

by

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April 2000
(revised May 2001)

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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 state quality-of-life. The compensating differentials approach derives from early work by Rosen (1979) and Roback (1982), who showed how to extract quality-of-life measures from compensating differentials in wages and house prices. To date, however, empirical analysis of compensating differentials generally has been cross-sectional in nature and provides little insight as to the extent or determinants of changes over time in the quality of life. In this paper, we provide estimates of the evolution in quality-of-life rankings for U.S. states over the 1981- 1990 period. 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 also for other high quality-of-life rural western states. However, we find substantial deterioration in quality-of-life rankings in some states that experienced rapid population growth during the decade. Reduced spending on infrastructure, increased traffic congestion, and air pollution account for the bulk of the deterioration in quality-of-life in these states. As would be expected, improvement in those same factors is shown to result in marked ascension in quality-of-life ranks among other states.

I. Introduction

There has long been widespread interest and popular debate surrounding regional variations in the quality of life. In recent decades, media analysts and others have alleged ongoing and substantial deterioration in quality-of-life in such places as California, New York, and other heavily populated states, in the wake of increasingly evident problems of congestion, school quality, public infrastructure, and the like. At the same time, analysts have extolled the growing agglomeration of produced and natural amenities in Colorado, Arizona, Wisconsin, and areas of the Pacific Northwest. With quality-of-life issues thought to importantly affect the competitiveness of states and hence their growth trajectories, elected representatives, economic development officials, and policy analysts increasingly have focused on the attractiveness of state and local amenity and public finance packages.

In an economic sense, the popular concern regarding quality-of-life can be described in terms of the importance of location-specific amenities to household utility, a view that generally has been confirmed by empirical studies of residential location decisions (see, for example, Graves (1980) or Blomquist and Berger (1992)). Economists have further contributed to the quality-of-life debate by pointing out that pecuniary differences across locations in wages or land rents should compensate for the differences in nonpecuniary characteristics that traditionally are included as elements of the popular rankings, and thus that differences in wage rates and land rents could be used to estimate the utility valuation weights (or implicit prices) of the nonpecuniary characteristics of locations.[1] Although there has been some disagreement about exactly how to specify such models, the standard approach now used in the academic literature is based on the work of Rosen (1979) and Roback (1982, 1988). 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 (1992), and Stover and Leven (1992), among others.

One uniform aspect of this research has been the absence of research on the extent or sources of changes over time in the quality-of-life. In virtually all of the published studies, researchers have focused on relative rankings of the quality-of-life at a single point in time. This focus is not surprising, given the severe limitations associated

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[1]In contrast to this economic approach, the choice of individual amenities included in the popular rankings and the weights assigned to these amenities are often chosen in an ad hoc manner, and thus 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.

with the compilation of time-series data on amenities. However, the results from existing studies provide little insight as to the evolution over time in quality-of-life comparisons, or what amenity changes have been most effective in improving the attractiveness of particular locations.

In this paper, we construct a time-series of quality-of-life rankings by applying the compensating differentials approach employed in the existing literature to a pooled time-series and cross-section of states for the 1981-1990 period. After assembling an extensive dataset of state-specific amenity characteristics, many of which vary over time, we estimate the implicit prices associated with each amenity characteristic and use these prices to construct the first comprehensive time-series of state-level rankings of the quality-of-life. As in other studies, this approach sheds light on the types of amenities that are most important to household location decisions. Further, the analysis also identifies the contribution of specific amenities to the relatively large movements in quality-of-life evidenced in some states.

In addition to our focus on comparisons of the quality of life rankings over time, our paper differs from the existing literature in terms of how we implement the compensating differentials approach. In particular, we extend the standard approach by deriving our estimates from a three-equation system that goes beyond earlier specifications in the literature to explicitly test for the capitalization of amenities in wages, housing rents, and prices of other commodities traded primarily in local markets. Prior empirical analyses failed to adequately specify and test the capitalization of place-specific amenities in the cost-of-living other than housing.

One limitation of our focus on changes in the quality-of-life is that consistent time-series of many location-specific attributes and amenities are available only at the state level. In choosing a location, households and firms may also compare amenities at the city or county level, and in this sense our estimates of quality-of-life differentials across states provide a partial view of the factors affecting the location decision. Nevertheless, many locational attributes are well-proxied at the state level of aggregation, including those associated with climate, proximity to natural amenities (parks, waterways, coastlines), funding of transportation infrastructure, public welfare, and higher education. Also, fiscal, environmental, and public infrastructure policies that affect the quality-of-life are often formulated at the state level and thus vary considerably more across states than across local areas within a state. Thus, while our analysis does not provide a full description of the range of locational opportunities within a state, we can give a sense of the average attractiveness of each state, and in this regard, our study provides the first significant estimates of changes in state-level quality-of-life and its determinants.[2]

Our results indicate that there were significant changes in the quality-of-life rankings of U.S. states over the 1981-1990 period. For example, perceptible improvements in the relative quality-of-life are found for such states as Alaska and Arizona, whereas notable declines are evident for Hawaii, Nevada, and New Hampshire. 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 appear to explain 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. As described below, 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

[2] The analysis also provides a meaningful adjunct to our earlier assessment of the determinants of directional migration among U.S. states (Gabriel, Mattey, and Wascher [1995]). In that analysis, we sought to specify and test the relative importance of economic versus quality-of-life effects in the determination of intertemporal population

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.[3] 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 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

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flows. In the absence of direct estimation of state-level quality-of-life indices, that analysis employed state-level fixed effects to proxy for those effects.

[3]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.

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

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

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

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.[4] As is evident in equation (4), 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 (1992)). Here, we add a 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 cross-sectional in nature, and thus do not provide much insight as to intertemporal evolution in the amenity vector or in quality-of-life rankings across places. However, 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 relatively high, given the level of public good provision, and at

[4]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 that 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 ignored, might explain why they found very different results from the reduced form and structural approaches.

other times the rate of taxation is less onerous.[5] 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 equilibration of household utility across places.6

We use equations (5) and (6) of Gyourko and Tracy (1992) 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

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

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 a_{jt} component of the error term, u_{ijt}. A reduced-form housing cost equation is defined similarly, as

$$\ln n_{ijt} = ?_0 + \mathbf{H}_{it}\, ?_1 + \mathbf{Z}_{jt}\, ?_2 + v_{ijt}, \qquad v_{ijt} = d_{jt} + \eta_{it}, \tag{6}$$

where $\mathbf{H}_{it}$ is a vector describing the characteristics of the housing unit occupied by individual i.[7]

Gyourko and Tracy 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, $?_1$) needed for quality-adjusting the wages and housing costs from the estimation of the parameters that describe capitalization of observed amenities (β_3, $?_2$) and unobserved

[5]For example, tax receipts from petroleum industry activities in Alaska have funded a relatively high 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.

[6] While a fuller dynamic specification might be appropriate for a complete description of the time-series properties of wages and prices, we find that this equilibrium assumption is adequate for capturing changes in quality-of-life and thus we maintain it throughout most of the paper.

[7]In Gyourko and Tracy, 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.

amenity capitalization and average quality differences in workers and housing across states (a_{jt}, d_{jt}). We estimate the quality-adjustment parameters ($\beta_1, \beta_2, ?_1$) for a benchmark year with microdata from the 1990 Census of Population and Housing.[8] The wage and housing cost variables from the Census pertain to 1989, and for these first-stage regressions we collapse the state-specific components of wages and housing costs into fixed effect parameters (? , µ):

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

$$\ln n_{ij89} = ?_0 + \mathbf{H}_{i89}\, ?_1 + \mu_{j89}, \qquad \mu_{j89} = \mathbf{Z}_{j89}\, ?_2 + d_{j89} \tag{8}$$

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 ($\mathbf{X}_{..89}, \mathbf{Y}_{..89}$):

$$\ln w^*_{.j89} = \ln w_{.j89} - ((\mathbf{X}_{.j89} - \mathbf{X}_{..89})\,\beta_1 + (\mathbf{Y}_{.j89} - \mathbf{Y}_{..89})\,\beta_2). \tag{9}$$

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

$$\ln n^*_{.j89} = \ln n_{.j89} - (\mathbf{H}_{.j89} - \mathbf{H}_{..89})\,\gamma_1. \tag{10}$$

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 ($? \ln w^*_{.jt}$) are constructed by estimating the counterpart to equation (7) for each year of the sample separately, 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

[8]We use a 1/1000 subsample of the Public Use Microdata A Sample. The wage variable is defined as total wage and salary earnings in 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 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 using a measure of homeownership user costs.

derived from the CPS.

The changes in state-level quality-adjusted housing costs (? $\ln n^*_{jt}$) 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.[9] In the user cost computation, state-level house price variation over the sample period is based on the Fannie Mae-Freddie Mac repeat sales price indices.[10]

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. Items in the index are chosen to be representative of the expenditure basket for a mid-management standard of living and consist of 59 separate goods and services that are available in most areas. In particular, the ACCRA index includes some widely traded commodities where prices are set in national or international markets and some other commodities and services with limited local trade areas. A more complete description of the index is available in American Chamber of Commerce (1995).[11]

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[9]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 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), 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 (1992) analyses.

[10]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.

[11]Our choice of the ACCRA index as a measure of cost-of-living differentials is largely dictated by the absence of alternative measures for our sample period. It is not, however, an ideal cost-of-living measure for the compensating differentials approach. In particular, the reference group for the market basket underlying the ACCRA index is not

To construct the non-housing cost-of-living variable used in our model, we first compute a benchmark expenditure level by setting non-housing consumption expenditures on locally-traded items in 1989 for the United States as a whole ($c^*_{.89}$) equal to the product of total consumption per household times the share of non-housing services in total personal consumption spending. We then compute an index of the cost-of-living except housing for each city in the ACCRA sample, 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.[12] These state-level aggregates of the ACCRA indices are used to distribute the benchmark level of 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. Similarly, by construction, our quality-adjusted measure of the local cost-of-living excluding housing (c^*_{jt}) is expressed in terms of expenditures in 1989 dollars.

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

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necessarily representative of the median family in each state and is limited to items that are available in all regions; as a result, the ACCRA index may not be reflective of the aggregate consumption bundle of a particular city's population. Koo, Phillips, and Sigalla (2000) find this to be an important reason why the ACCRA measure differs from a one-time experimental regional cost-of-living measure constructed by the Bureau of Labor Statistics for a limited number of Metropolitan Statistical Areas (Kokoski, Cardiff, and Moulton, 1994). In addition, Koo, et al. criticize the ACCRA index because it is computed by comparing the price-level in a particular city to the unweighted average of all the cities in the sample, so that the measure is relative to the city average rather than to the average price paid by consumers in the United States. However, it is not obvious that this is a problem for our analysis given that households and firms are assumed to be comparing average costs across geographic locations. Moreover, to the extent that the biases identified in the ACCRA measure are roughly constant over time, they will be subsumed in the fixed state effects coefficients included in our model. Nonetheless, these considerations lead us to inject a note of caution in interpreting our results using this index, and we subsequently provide some comparisons of our results with a model excluding the non-land cost-of-living equation.

[12] Because the ACCRA indices do not cover a fixed set of cities in each year, the national average to which the regional indices are benchmarked may vary with compositional changes over time in the specific cities included in any given year. This variation does not present a problem for our estimation strategy as these compositional effects will be captured in the fixed year effects included in the second stage of our estimation procedure. More problematic, perhaps, is the fact that the state-level indices we compute will also be affected by such compositional changes. Our sense, however, is that within-state variation in cost-of-living differences is small compared with cross-state differences. Moreover, because the ACCRA indexes are measured independently of the other variables in the

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 year-specific dummy variables T_t :

$$\ln w^*_{jt} = ß_0 + \mathbf{Z}_{jt}ß_3 + T_t\, ß_4 + a_j + e_{.jt} \tag{11}$$

$$\ln n^*_{jt} = ?_0 + \mathbf{Z}_{jt}?_2 + T_t?_3 + d_j + \eta_{.jt} \tag{12}$$

$$\ln c^*_{jt} = ?_0 + \mathbf{Z}_{jt}?_1 + T_t\, ?_2 + ?_j + v_{.jt} \tag{13}$$

As the data have 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 (a_j, d_j , $?_j$) are not separately identified from the coefficients on observed amenities ($ß_3$, $?_2$, $?_1$). 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 $\mathbf{Z}_{jt}$. Note that the set of amenity variables is similar to that in Blomquist et al (1988) and in Gyourko and Tracy. 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 climatic 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 the joint influence 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.[13] In particular, we include measures of income, sales, and

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equation, we can arguably treat these compositional changes simply as measurement error in the dependent variable of the non-land cost-of-living equation.

[13]Two points should be kept in mind when considering the effects 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

property tax rates, as well as 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 time 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 time invariant. 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 the number of 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

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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 (1992) emphasize the possible capture of amenity values by the public sector.

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 (Maine) lowered 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 therein, we cannot ascertain how these developments affected the overall 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 affect measured quality of life at the state level and where, geographically, the most significant changes in 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, γ_2, θ_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:

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

where $n^*_{...}$, $c^*_{...}$, and $w^*_{...}$ 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: 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. In contrast, air quality is estimated to be very important to households' evaluations of the desirability of a location, with higher levels of either ozone or carbon monoxide pollution estimated to be significant disamenities. The results do not lend support to arguments that tough environmental regulations are an inefficient way to achieve desired environmental outcomes; given those outcomes, an additional proxy for the stringency or leniency of environmental regulation has no perceptible compensating differential effect.

As expected, holding constant state public service levels, 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 highways; in addition, congestion as measured by commuting times is undesirable. The estimate of the implicit price for the higher education share of state and local government budgets is not statistically different from zero, although poorer school quality, as proxied by a high student-teacher ratio, is estimated to be a disamenity.

The attributes and implicit prices imply a broad range of capitalized amenity values across the United States. For example, the full sample average for precipitation is 35 inches per year,[14] 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 $18 in compensating differential, and the range of attribute values implies that an Alabama resident's quality-of-life is held down by $1164 (1989 dollars), about $1027 more than the $137 compensating differential for a Nevada resident (final columns of table 2). The widest range of contributions from an amenity characteristic derives from 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 $15,716 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:

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

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 higher price markups on housing and other locally-traded commodities) to live in the higher amenity state.[15]

In table 3, we present the rankings of the states in 1981 and 1990 derived from the estimated quality-of-life indexes, along with the standard errors of the assigned ranks (and changes in ranks) implied by our estimation procedure. The standard errors are computed from Monte Carlo simulations that perturb the coefficient vectors in

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

[15] Although the index is denominated in constant 1989 dollars, we do not make intertemporal comparisons of the index values because we have not included in the indexes any part of the contributions of the time fixed effects, which might partially reflect the evolution of the aggregate national quality-of-life.

accordance with the estimated variance-covariance matrices for the three second-stage equations. In particular, we use 1000 draws from the multivariate normal distribution and report the empirical standard deviations of the ranks (or changes in ranks) for each state. In general, the ranks appear to be estimated fairly precisely, with standard deviations ranging from less than 0.1 for New York to 6.6 for Alaska. Similarly, the standard deviations of the change in ranks range from 0.1 for New York to 8.5 for Alaska.[16,17]

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 both 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. In the metropolitan area results of Gyourko and Tracy (1992), 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 excludes 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 Idaho, 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

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[16] The standard errors associated with our quality of life rankings are well below those estimated by Gyourko and Tracy (1992), who find it difficult to precisely estimate the quality of life across metropolitan areas. While it is difficult to make direct comparisons to Gyourko and Tracy's results, we suspect that the substantially greater number of observations we use to estimate the quality-adjusted wage and housing cost variables, coupled with the presence of time-series variation in our amenity vector, has increased the precision with which we can estimate the quality of life rankings of U.S. states.

[17] Note that the standard deviations of the ranks and change in ranks depend not only on the implied standard errors of the quality-of-life estimates, but also on how the state quality-of-life estimates are clustered. For example, the quality-of-life estimate for New York is well below that for any other state, so that the changes in the quality-of-life valuation in the Monte Carlo simulations have little effect on its rank. In contrast, not only is the quality-of-life estimate for Alaska driven by some amenities for which the implicit prices are estimated less precisely, but the overall value for Alaska is quite close to those for a number of other states; either of these factors may cause the simulated changes to have a sizable affect on its rank.

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.[18] 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, Nevada, and Hawaii are estimated to have deteriorated noticeably in the quality-of-life rankings.

The evolution of particular characteristics, evaluated at their implicit prices, provides some evidence on how changes in amenities may have contributed to an improvement or deterioration in the quality-of-life ranks for individual states. 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, Washington, New Mexico, and Hawaii), other states with high rates of population growth saw significant improvements in estimated quality-of-life (Alaska, Arizona, Florida, and Colorado). Thus, population growth, per se, does not explain the evolution of rankings. Rather, the key to maintaining or increasing quality-of-life appears to have been the management of population growth.

For example, the states that experienced deteriorating quality-of-life tended to cut back on the share of state and local government expenditures devoted to highways and transportation infrastructure, leading to increases in traffic congestion and average commuting times. 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 sometimes not strict enough to prevent an increase in carbon monoxide pollution relative to other states. In contrast, some other fast-growing states with very stringent air quality management regimes--particularly Arizona, California, and Colorado--benefited from improved air quality.

Although these patterns are most evident for particular states, in table 4 we present some summary statistics on the contribution of amenities to the evolution of quality-of-life ranks. For this summary, we have categorized 31 states as having a "stable" quality-of-life ranking that changed less than five places between 1981 and 1990. Among other states, nine experienced a large deterioration in relative quality-of-life (increased ranks) and ten experienced a

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[18]Although California experienced substantial net out-migration of population at the end of the decade, that flow appears to have been fueled largely by a 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)).

large improvement (decreased ranks). On average across the states in the improving group, the rank improved by about nine places, and the average change in rank for the deteriorating group was about eleven places.

In the group with deteriorating quality-of-life ranks, the attribute with the largest 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.[19] Relatively longer commuting times, higher carbon monoxide and ozone levels, and a lower share of state and local spending on welfare also were large sources of deterioration in quality-of-life for these states. For example, New Hampshire, New Mexico, and Nevada all saw a deterioration in their quality-of-life ranks associated with a reduced highway spending share, increased commuting times, and lower air quality.

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 benefited 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.

V. Robustness and Other Results

Despite the extensive set of amenities included in the quality-of-life calculations, we undoubtedly have excluded--usually for data reasons--numerous other location-specific attributes that might influence either overall rankings or the implicit price estimates for particular amenities. Most of our main results appear robust to reasonable alternative specifications of the model, but we did notice some interesting sensitivities. For example, the result 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

[19] 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 that state’s 1990 commuting times, but extrapolated all other California traits from their 1981 values using the average change in traits for the nation as a whole (as shown in the 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.

choices. However, the interpretations of the evolution of some other 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.

In particular, 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 more 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. 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.[20]

More generally, a potential solution to this omitted variables problem is to include group fixed effects in the quality-of-life estimates (see, for example, Gyourko and Tracy (1992)).[21] Of course, the use of fixed effects has its own problems because there could be location-specific attributes that are correlated with the dependent variables but

[20]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.

[21]In addition to the OLS type of estimates provided by Blomquist et al (1988), Gyourko and Tracy 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. Indeed using our data, we rejected the random effects specification in favor of the fixed effects model using the standard Hausman test. Moreover, unlike in Gyourko and Tracy, our use of time-series amenity data permit the use of a fixed effects specification by adding the needed degrees of freedom.

should not be treated as influencing the quality-of-life estimates.[22] Regardless, it is 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 (a_{jt}, d_{jt}, $?_{jt}$) are not separately identified from the coefficients on observed amenities ($ß_3$, $?_2$, $?_1$) 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 $\mathbf{Z}_{jt}$. Alternatively, one can achieve identification by deleting the time-invariant amenities (such as whether a state is located on a coast) from $\mathbf{Z}_{jt}$ and restricting the fixed effect parameters to vary only over location ($a_{jt}=a_j$, $d_{jt} = d_j$, $?_{jt} = ?_j$). Employing these restrictions, we 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 (and standard errors) we obtain with such state-level fixed effects are shown in table 5. As is evident from 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 .84. Changes in rankings are also positively correlated across the models (with a correlation coefficient of .45), although the rankings do not change nearly as much in the fixed effects version as in the observed amenities version of the model. This occurs for two reasons. First, the fixed effects coefficients pick up substantially more of the cross-section variation in the data than do the full vector of amenities. Second, the absence of cross-sectional identifying information reduces the implicit prices of many of the time-varying amenity characteristics. The end result is that the fixed effects tend to dominate the rankings in both years, with the time-varying amenities producing only minor changes in ranks. Whether this is sensible depends on the extent to which the fixed effects are capturing additional amenities that were excluded from the base model, or whether, as we suspect, the fixed effects are picking up additional quality effects in wages and house prices that we were unable to control for in the first stage of the estimation process.

With respect to the individual rankings, some states show different patterns in the fixed effects model than in the full-amenity specification, while others do not. For example, Alaska and Arizona, which show the greatest

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[22] 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

improvement in table 3, instead show a slight deterioration in table 5. Similarly, the deterioration in ranks for both Hawaii and New Hampshire are less severe in the fixed effects model than in the full-amenity specification. For other states, however, the results are fairly similar. Nevada shows a marked deterioration in its relative quality-of-life in both specifications, while mountain states such as Montana and Wyoming consistently rank high, and industrialized states consistently rank low.[23]

Additionally, it is potentially interesting to compare the results from our three-equation version of the model and the two-equation counterpart employed by previous researchers. As we noted earlier, we think that the addition of the non-housing expenditure equation to the model represents a conceptual improvement to the quality-of-life framework. Moreover, from a practical standpoint, the R^2 for that equation is nearly as high as for the wage and housing expenditure equations and many of the coefficients on the amenities are statistically significant, suggesting that amenity values are capitalized in non-housing prices as well. In particular, using only the wage and housing expenditure equations results in a somewhat higher full implicit price for temperature extremes but a somewhat lower implicit price for precipitation and wind speed; the two-equation model also results in a negative full implicit price for being on a coast, contrary to our priors. For the time-varying amenities, the largest differences were for commuting time and air quality, for which the full implicit prices in the two-equation model were more negative than in the three-equation model. In addition, full implicit prices were somewhat more positive in the two-equation version for both state and local expenditures on welfare and highways.

Despite these differences, for both the full-amenity and fixed-effects specifications, the rankings generally were quite similar in the two- and three-equation models, with Spearman rank correlation coefficients over 0.9 for the

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calculating the quality-of-life indices.

[23]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., we found that a migration disequilibrium adjustment has little effect on the implied quality-of-life rankings.

ranks and between 0.8 and 0.9 for the changes in ranks. This suggests that, in our exercise, much of the capitalization of amenity differences in non-land costs are mirrored either in terms of lower wages or in terms of higher housing costs. There were, however, a few states for which extending the specification matters. For example, Iowa shows no deterioration in ranking when only two equations are used, in contrast to the results for the three-equation model. In addition, Alaska, Hawaii, and Rhode Island all exhibit a somewhat lower relative quality of life in the two-equation specification than in the three-equation version of the model. Nonetheless, the overall similarity of the results for the two- and three-equation specifications suggests that researchers unable to obtain prices for locally traded non-housing goods may still be able to compute suitable quality-of-life estimates using only wage and rent data.

VI. Conclusion

In recent years, increasingly evident problems of congestion, school quality, public infrastructure, and the like have led to widespread media discussion and policy debate regarding evolution in the quality-of-life at the city and state level. Yet little is known about changes over time in the distribution of place-specific amenities, their valuation by households and firms, and thus their effects on measured quality-of-life. To address these issues, this research applies the compensating differential paradigm to provide the first empirical estimates of the extent and sources of evolution in the quality-of-life among U.S. states. Results of the analysis suggest that some states recorded a substantial deterioration in estimated quality-of-life over the 1980s, in large part due to limited infrastructure investment in the wake of rapid growth. Estimates from our model suggest that reduced spending on highways, increased traffic congestion, and elevated air pollution have been the most important contributors to the deterioration in the quality-of-life in those states. Elsewhere, the quality-of-life either has remained relatively high or has improved. States ascended in our estimated quality-of-life rankings for a variety of reasons, including improved air quality, increased highway spending, and reduced tax burdens and commute times.

In sum, our findings suggest that state-level mitigation of infrastructure deterioration, congestion, and air pollution serve to perceptibly elevate estimated quality-of-life. Similarly, careful calibration of the tax and expenditure mix, so as to adequately reflect consumer preferences, can go some distance in enhancing measured quality-of-life. While our analysis does indicate the determinants of changes in the attractiveness of states over time, data resources do not permit a full description of within-state variation in the quality-of-life. Data development permitting such an exercise would similarly benefit local government policymakers seeking to assess how changes in city- or county-specific attributes affect the quality-of-life.

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