Migration, Regional Equilibrium, and the Estimation of Compensating Differentials

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Following Sherwin Rosen's (1979) paper on wage-based indexes of urban quality of life, a number of recent studies have used the level of regional wages or rents to measure regional environmental quality (including the quality of the climate) (see e.g., Jennifer Roback, 1982, 1988; John P. Hoehn et al., 1987; Glenn C. Blomquist et al., 1988). The assumption underlying these studies is that the interregional system is in equilibrium so that wage and rent differentials are compensated differentials and thus serve as accurate proxies for differentials in environmental quality.[1] Under this equilibrium approach, regional differentials in wages and prices do not necessarily reflect utility differences that can be arbitraged through household migration. Only such regional differentials as remain after controlling for amenity differentials across regions should represent utility differentials that would induce migration. A properly specified migration equation should thus include both regional amenity and regional wage and price variables.

Another aspect of the equilibrium assumption in quality-of-life studies is that regional markets are efficient, so that regional prices quickly realign to clear them subsequent to any disequilibrating exogenous changes in demand or supply conditions. The equilibrium theorists believe that at any point in time regional wages and prices have adjusted to their equilibrium values (see Philip E. Graves and Thomas Knapp, 1988 p. 3). If this belief were in fact true, regional differences in wages and prices would indeed represent compensating differentials and could be used to measure the value of the location-specific attributes. However, if regional markets do not tend to clear quickly on a continuous basis, the erroneous assumption that they are in equilibrium at any point in time will lead to biased estimates of amenity valuations and, in general, to biased valuations of the entire bundle of the location-specific characteristics associated with each region, as suggested by Evans (1990). Several shortcomings of recent quality-of-life measures are addressed in the present paper.

The paper is organized as follows. Section I develops the theoretical approach. Because much effort has been expended to measure appropriately the variables that come out of the theoretical model, Section II provides detail on various measurement issues. Section III discusses the econometric approach employed in the study and presents empirical results. Section IV discusses the implications of the results for estimating compensating differentials, and Section V presents a summary and conclusions.

I. The Theoretical Approach

We begin with the hypothesis that an area's rate of population growth due to internal migration during any given year is a function of how well off households expect they will be in the area (a) compared to elsewhere in the United States (u):

$$(1)\quad [(\mathrm{NLF}_{a,t-1}+\mathrm{ECM}_{a,t})/\mathrm{NLF}_{a,t-1}] = h\{[\mathrm{NPV}(\mathrm{EY}_a)/\mathrm{NPV}(\mathrm{EY}_u)],[A_a/A_u]\}$$

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[1]Alan W. Evans (1990) discusses the assumption of equilibrium in analyses of migration and interregional differences.

where NLF refers to natural civilian labor force (exclusive of military personnel and their dependents, persons 65 and over, and immigrants), ECM refers to economic migration (including dependents of the actual economic migrants), NPV refers to net discounted present value at time t_0, EY is expected income at time t_0, and A denotes location-specific amenities, broadly defined. If area a is sufficiently small relative to the United States, conditions in the United States provide a close approximation to those everywhere other than a. Equation (1) includes both the influence of spatial economic opportunity and amenity differentials on migration. We can show that

$$(2)\quad \mathrm{NPV}_{k,t_0} = \mathrm{EY}_{k,t_0}\left\{\left[\sum_{t=t_0}^{T} \frac{(1+g_k)^{t-t_0}}{(1+r_t)^{t-t_0}}\right] - \phi\right\}$$

where r is the periodic discount rate, g is the expected growth rate, and ϕ is the fixed proportion of income required for moving.

Assuming further that relative expected income (RY) can be estimated as the relative wage bill divided by the natural labor force (NLF), and that $g_a = g_u$,[2] we can substitute into (1) and obtain

$$(3)\quad [(\mathrm{NLF}_{a,t-1} + \mathrm{ECM}_{a,t})/\mathrm{NLF}_{a,t-1}] = h[\mathrm{RY},(A_a/A_u)].$$

Using a log-linear functional form and an error term, we obtain

$$(4)\quad \ln[(\mathrm{NLF}_{a,t-1} + \mathrm{ECM}_{a,t})/\mathrm{NLF}_{a,t-1}] = \ln\lambda_a + \lambda_1 \ln \mathrm{RY}_{a,t} + e_{a,t}.$$

The previous t_0 subscript is replaced with t because the model is empirically implemented with observations that span a number of years. We expect that $\lambda_1 > 0$. Individual effects ($\ln\lambda_a$) represent the relative effects of amenities in area a versus area u on the net migration rate for area a.[3] The value of $\ln\lambda_a$ will be negative for amenity-poor areas and positive for amenity-rich areas.

II. Data and Measurement Issues

Annual data are used for the period 1971–1988 for 50 states plus Washington, DC (918 observations). Because many of the required data are not routinely available, they must be constructed. Economic migration is defined as the difference between actual civilian population under age 65 and estimated civilian population under age 65 in the absence of migration. Annual actual population is reported by the Census Bureau, with annual estimates anchored by the 1970 and 1980 decennial censuses. Because the Census Bureau distributes the 1970 census undercount over intercensal years, intercensal-year estimates must be adjusted. Moreover, military personnel and their dependents by age and sex are subtracted from the reported populations.

The demographic procedure used to generate an estimate of population in the absence of migration begins by applying a cohort algorithm to the adjusted 1970 population. Natality and survival rates applied to the cohorts produce estimated births and deaths which are then proportionately adjusted to obtain gross reported birth and death rates in each state.[4] International im-

[2]The assumption that $g_a = g_u$ is made for convenience. If $g_a \neq g_u$ is used instead, the interpretation of the model is slightly changed. Equation (2) becomes $\mathrm{NPV}_{a,t_0}/\mathrm{NPV}_{u,t_0} = (g\mathrm{EY}_{a,t_0}/\mathrm{EY}_{u,t_0})$, where $g = g_a/g_u$. In footnote 3 and the associated text, the implication of the assumption $g_a \neq g_u$ is explained.

[3]If it is assumed that $g_a \neq g_u$, the individual-effects terms will include as a component this differential expectation, as well as the relative effects of amenities. Under our assumption that $g_a = g_u$, the individual effects reflect the influence of relative amenity differentials, broadly defined. Moreover, state-specific moving costs and systematic differences across states in personal characteristics would also act like amenities and be reflected in the $\ln\lambda_a$ term. Our analysis assumes that such differences are small relative to differences in environmental quality.

[4]Natality rates for each state's cohorts are drawn from a 1980 births count divided by the relevant 1980 female age cohorts. Survival rates are estimated from reported deaths and the relevant population cohort in the 1980 census. Natality and survival rates are trended forward and backward from 1980 based on projected and historical U.S. trends.

migrants are then added by cohort to the cohort survival calculations to obtain estimated population in the absence of economic migration.[5]

Annual nominal industry wage rates are defined as average annual total wages and salaries divided by the average annual employment level, as reported by the Bureau of Economic Analysis (BEA). Nominal wage rates are then adjusted for state differences in the cost of living and rate of taxation. The rate of taxation is the sum of income and sales taxes (by state of collection) divided by personal income. Income taxes and personal income are reported by BEA in *Local Area Personal Income*; sales taxes are reported by the Bureau of the Census in *Government Finances*.

State-specific personal-consumption-expenditure deflators are constructed to adjust wage rates for cost-of-living differences. Industry outputs are priced either at the average national price or local production costs, depending on whether the industry is classified as one that primarily fulfills export demand (national industry) or local demand (regional industry).[6] The industry weights for each consumer category are given by a national consumption vector. Cost functions for regional industries are given by Cobb-Douglas substitutability between primary inputs (labor, capital, and fuel) and fixed proportions of intermediate goods. State-specific data are used to calculate all primary input costs.[7]

Housing prices are calculated as a separate series benchmarked to the 1970 and 1980 housing censuses. Changes in housing prices for intercensal years are based on changes in major-region housing prices reported by the Census Bureau. Housing prices for 1981–1988 are projected forward by a weighted average of metropolitan areas within the state or major region, as available from the National Association of Realtors.

Natural labor force is the sum of cohort-specific participation rates times the appropriate cohorts. Natural participation rates are obtained from a cross-section regression for 1980 across the 51 regions for 40 cohorts. Natural rates are desired, instead of actual participation rates, to account for discouraged-worker effects on participation rates. The dependent variable in the regression is the actual participation rate, while the independent variable of special interest is the regional unemployment rate. We obtain natural participation rates by substituting full-employment unemployment rates into the estimated regression equation.[8] The rates are then trended forward and backward from 1980 based on U.S. labor-force participation trends, as reported by the Bureau of Labor Statistics. Finally, the census-undercount adjustment applied to the population is also applied to the labor force.

III. Econometric Approach and Empirical Results

Two basic econometric issues arise in the estimation of equation (4). First, individual effects are associated with each area in our model. The second issue is simultaneous-equations bias. The right-hand-side variable in equation (4) (i.e., RY) is part of a larger model and is endogenous. Consequently, the component of $e_{a,t}$ not associated with individual effects is not independent of the right-hand-side variables. To avoid inconsistency from this feature of the model and to account for the individual effects, we use an instrumental-variables version of the fixed-effects model.[9]

[5]International migrants were allocated to states based on the 1980 state-specific relative distribution of international migrants.

[6]A good is assumed to be sold in a nationally competitive market if the sum of all states' exports exceeds 50 percent of total production (Treyz and Benjamin H. Stevens, 1985).

[7]Primary and intermediate input shares of output are derived from a linear interpolation between the 1977 BEA input–output table and a projected table for the year 2000. Relative capital costs are determined according to the Jorgensonian user-cost-of-capital approach. Relative fuel costs are from the U.S. Department of Energy.

[8]Regression results are available from the authors upon request.

[9]The instrumental variables are from a larger model that is described in Treyz et al. (1988). Based on the

TABLE 1—INSTRUMENTAL-VARIABLES ESTIMATES OF THE FIXED-EFFECTS MODEL OF ECONOMIC MIGRATION

Variable	Parameter estimate	Asymptotic standard error
ln(RY)	0.2147	0.0138
Wald test statistic (H_0: no fixed effects) $= 1{,}202 \sim \chi^2_{[50]}$, $P < 0.0001$		

Table 1 reports the results associated with the instrumental-variables fixed-effects model. The estimated individual effects are not reported. The Wald test for the joint significance of the individual effects reported in Table 1 ($W = 1{,}202 \sim \chi^2_{[50]}$, $P < 0.0001$) rejects the null hypothesis of no such effects.

Given the joint significance of the fixed effects, we use the estimates of these area-specific effects to compute the equilibrium values of RY for each area. To get the equilibrium values for RY, we first note that in equilibrium our measure of net economic migration for area a takes on the value of unity. The equilibrium value of RY for each area therefore is that value generating no net migration. It is the value that just offsets the impact of the estimated individual effect for each area.

Point estimates for the equilibrium values of RY computed in this fashion are shown in Table 2 along with the actual 1980 RY values; 90-percent confidence-interval estimates are also presented. An equilibrium value of RY of less than unity implies area characteristics that are attractive to individuals such that they are willing to accept lower earnings or pay higher local prices, or both, in order to "consume" the area's characteristics. Areas with equilibrium RY's that exceed unity have less attractive characteristics, and therefore individuals require a premium in earnings or lower local prices, or both, to be in equilibrium.

The point-estimate equilibrium values presented in Table 2 indicate that, of the 13 areas in the West, 12 have values less than unity. Eleven of the 17 areas in the South have such values. The comparable figures for the Northeast and Midwest are four of nine and five of 12, respectively. These results indicate that the West and South have more attractive characteristics than the Northeast and Midwest. These findings are consistent with other studies (e.g., Greenwood and Hunt, 1984).

When the 90-percent confidence intervals of equilibrium RY values are used, the West is still the most amenity-rich region, with seven areas. Now, however, the South and Northeast are much closer, with the South having four amenity-rich areas and the Northeast having three. The Midwest now contains only one amenity-rich area. In each of these 15 amenity-rich areas, the upper 90-percent confidence-interval bound is less than unity.

IV. Estimating Compensating Differentials

Several recent studies have estimated regional amenity values by computing compensating differentials in labor and land markets under the assumption of regional equilibrium. The two most recent studies of this type, Blomquist et al. (1988) and Hoehn et al. (1987), make such estimates for 1980. Table 2 indicates that, for a number of states, regional equilibrium did not hold in 1980. The measure of "disequilibrium" by state is the difference between the state's actual relative real after-tax earnings in 1980 (RY) and its corresponding point-estimate equilibrium value (RY*).

In amenity-rich states, where RY* < 1, amenity valuations based on the assumption of regional equilibrium will overstate (understate) the compensating differential when the actual RY value is less (more) than the estimated equilibrium RY* value. For the states where understatement occurs, more net in-migration, ceteris paribus, would be expected until wages fall and local prices and rents rise sufficiently to drive RY down

value of the J. A. Hausman (1978) test statistic ($m = 69 \sim \chi^2_{[1]}$, $P < 0.0001$), we reject the null hypothesis that the random-effects model is correct and settle on the results obtained with the instrumental-variables fixed-effects estimator.

TABLE 2—ACTUAL (RY), POINT, AND 90-PERCENT CONFIDENCE-INTERVAL ESTIMATES OF EQUILIBRIUM (RY*) VALUES OF RELATIVE AFTER-TAX REAL EARNINGS AND COMPENSATING DIFFERENTIALS (QLI), BY STATE, FOR 1980

					QLI	
Region and state	RY	Lower-bound RY*	RY*	Upper-bound RY*	Scaled	Original
Northeast:						
Maine	0.8606	0.7091	0.8160	0.9229	NA	NA
New Hampshire	1.0180	0.8046	0.9018	0.9990	NA	NA
Vermont	0.8595	0.7110	0.8181	0.9252	NA	NA
Massachusetts	1.0065	0.9223	1.0347	1.1471	0.9618	792.13
Rhode Island	0.9233	0.8134	0.9235	1.0336	NA	NA
Connecticut	1.0685	0.9646	1.0770	1.1894	NA	NA
New York	0.9859	0.9348	1.0521	1.1694	0.9902	203.62
New Jersey	1.0907	1.0034	1.1156	1.2278	0.9994	12.69
Pennsylvania	1.0181	0.9200	1.0322	1.1444	0.9923	159.13
Midwest:						
Ohio	1.0401	0.9780	1.0938	1.2096	1.0040	−83.65
Indiana	1.0347	0.9682	1.0824	1.1966	1.0006	−13.27
Illinois	1.0498	1.0195	1.1363	1.2531	1.0330	−683.67
Michigan	1.0608	1.0240	1.1403	1.2566	1.0400	−829.37
Wisconsin	0.9809	0.8831	0.9946	1.1061	0.9907	193.00
Minnesota	0.9986	0.8988	1.0095	1.1202	1.0118	−243.59
Iowa	0.9332	0.8527	0.9685	1.0843	0.9882	245.49
Missouri	1.0177	0.9355	1.0467	1.1579	1.0533	−1,105.23
North Dakota	0.8771	0.7936	0.9097	1.0258	NA	NA
South Dakota	0.8350	0.7413	0.8571	0.9729	0.9612	803.73
Nebraska	0.9586	0.8746	0.9878	1.1010	0.9932	141.09
Kansas	1.0521	0.9164	1.0283	1.1402	0.9875	259.31
South:						
Delaware	1.0563	0.9708	1.0785	1.1862	NA	NA
Maryland	1.0668	0.9748	1.0847	1.1946	1.0107	−223.69
District of Columbia	1.1077	1.3204	1.4576	1.5948	0.9998	3.35
Virginia	1.0460	0.9107	1.0161	1.1215	0.9416	1,210.55
West Virginia	0.9339	0.8596	0.9782	1.0968	0.9996	8.32
North Carolina	0.9357	0.8095	0.9151	1.0207	0.9644	737.41
South Carolina	0.9364	0.7859	0.8909	0.9959	0.9648	728.83
Georgia	0.9895	0.8541	0.9577	1.0613	0.9686	651.59
Florida	0.8816	0.6420	0.7343	0.8266	0.9705	611.38
Kentucky	0.9282	0.8179	0.9289	1.0399	0.9999	2.14
Tennessee	0.9383	0.8077	0.9135	1.0193	0.9933	139.76
Alabama	0.9265	0.8136	0.9228	1.0320	1.0268	−554.80
Mississippi	0.8519	0.7364	0.8499	0.9634	0.9662	699.85
Arkansas	0.8426	0.6987	0.8066	0.9145	0.9646	733.41
Louisiana	1.0580	0.9164	1.0296	1.1428	1.0012	−25.35
Oklahoma	1.0419	0.8301	0.9377	1.0453	0.9959	84.37
Texas	1.1276	0.9121	1.0149	1.1177	1.0199	−412.90
West:						
Montana	0.8762	0.7501	0.8629	0.9757	1.0084	−174.55
Idaho	0.8746	0.7249	0.8322	0.9395	0.9743	533.39
Wyoming	1.1371	0.8564	0.9597	1.0630	NA	NA
Colorado	0.9723	0.8061	0.9086	1.0111	0.9362	1,322.75
New Mexico	0.9190	0.7630	0.8693	0.9756	NA	NA
Arizona	0.9017	0.6654	0.7587	0.8520	0.9521	993.41
Utah	0.9603	0.8521	0.9603	1.0685	0.9888	232.82
Nevada	0.9919	0.6864	0.7762	0.8660	0.9592	845.46
Washington	0.9794	0.8038	0.9078	1.0118	1.0015	−30.64
Oregon	0.9040	0.7615	0.8671	0.9727	0.9821	370.29

TABLE 2—CONTINUED

Region and state	RY	Lower-bound RY*	RY*	Upper-bound RY*	QLI Scaled	QLI Original
California	0.9008	0.7959	0.9028	1.0097	0.9718	584.10
Alaska	1.3140	1.1288	1.2303	1.3318	NA	NA
Hawaii	0.8098	0.7521	0.8658	0.9795	NA	NA

Notes: "Lower-bound RY*" is the lower bound of the 90-percent confidence interval around the point estimate, RY*; "upper-bound RY*" is the upper bound of the 90-percent confidence-interval around the point estimate, RY*. The 90-percent confidence interval estimate is: $\hat{RY}^* \pm 1.645\ (SE_{RY^*})$. The standard error of RY*, SE_{RY^*}, is computed as follows: by definition, $RY_a^* = (1 - \hat{\lambda}_a / \hat{\lambda}_1)$; therefore, an approximate expression for the variance of RY_a^* is: $V(RY_a^*) = \mathbf{J}'\mathbf{V}(\hat{\lambda}_a, \hat{\lambda}_1)\mathbf{J}$, where $\mathbf{J}'$ is the transposed Jacobian matrix,

$$\frac{\partial RY_a^*}{\partial \hat{\lambda}_a} \quad \frac{\partial RY_a^*}{\partial \hat{\lambda}_1}$$

and $\mathbf{V}(\hat{\lambda}_a, \hat{\lambda}_1)$ is the variance–covariance matrix of the estimated parameters $\hat{\lambda}_a$ and $\hat{\lambda}_1$. SE_{RY^*} is therefore the square root of $V(RY_a^*)$.

"QLI original" is the Blomquist et al. (1988) measure of compensating differentials weighted by population to correspond to states; "QLI scaled" is defined as $(\bar{Y}_{US} - \text{QLI original})/\bar{Y}_{US}$ where $\bar{Y}_{US}$ represents mean household earnings in the United States from the 1980 census. This transformation provides a convenient scaling with which to compare the two sets of estimates. However, we do not emphasize such cardinal-based comparisons because of the differences in spatial coverage between our state-based estimates and our extrapolations of the county-based estimates of Blomquist et al.

TABLE 3—CONTINGENCY TABLE FOR DISEQUILIBRIUM GAP (RY − RY*) AND (EQUILIBRIUM) COMPENSATING DIFFERENTIAL (RY*)

RY − RY*	RY* < 1	RY* > 1
Positive	21	5
Negative	10	14
Chi-square contingency-test statistic (H_0: no relationship) = 6.5, $P = 0.0106$		

Note: One area was in equilibrium and therefore had (RY − RY*) = 0.

to the equilibrium value given by RY* in Table 2. For the states where overstatement occurs, additional net out-migration would be required to raise wages and lower local prices and rents sufficiently to drive RY up to the equilibrium value given by RY*.

In amenity-poor states, where RY* > 1, disamenity valuations based on the assumption of equilibrium will overstate (understate) the compensating differential when RY values exceed (fall short of) RY*. Additional net out-migration would be expected in states for which RY < RY* so that wages would rise and local prices and rents would fall, ceteris paribus. Conversely, additional net in-migration would be expected in states where RY > RY* so that wages would fall and local prices and rents would rise, ceterus paribus.

Table 3 presents the 1980 "disequilibrium gaps" cross-classified by the point estimates of the equilibrium values, RY*, for each area. The 19 areas in the right-most column are amenity-poor, requiring a value of RY* > 1. The 31 areas in the left-most column are amenity-rich, requiring a value of RY* < 1. The predominant pattern for 1980 is that amenity-rich areas have values of RY > RY*, while amenity-poor areas have RY < RY*.

The chi-square goodness-of-fit contingency table presented in Table 3 supports this conclusion ($X_{[1]}^2 = 6.5$, $P = 0.0106$). Consequently, estimates of compensating differentials based on the erroneous assumption of regional equilibrium would generally lead to undervaluations of compensating differentials in 1980. Notwithstanding this predominant direction of bias for 1980, ten of the amenity-rich states and five of the amenity-poor states would have had overestimates of compensating differentials under the equilibrium assumption. Our estimates indicate that one area was in equilibrium in 1980, and therefore only 50 observations appear in Table 3.

Based on the confidence-interval estimates of RY*, only seven areas have statistically significant disequilibrium gaps, with six of these having RY values significantly larger than RY* and one having an RY significantly smaller. Of these seven areas, five also have RY* values that are statistically significant: four are significantly amenity-rich, and one is significantly amenity-poor. Of these five areas, a pattern emerges that is consistent with that found in the contingency analysis of Table 3. The four areas that are amenity-rich have RY's that are greater than the confidence interval of RY*; the amenity-poor area has an RY less than the confidence interval of RY*. Therefore, the confidence-interval estimates indicate that the assumption of equilibrium in 1980 would tend to lead to an underestimation of compensating differentials, although the extent of the problem in terms of the number of affected areas is relatively small.

In addition to the potential quantitative bias of erroneously assuming regional equilibrium, the potential exists for a qualitative bias, wherein areas would be misclassified as amenity-rich when they are in fact amenity-poor, and vice versa. To analyze this issue, we took the 1980 county estimates of compensating differentials reported by Blomquist et al. (1988 pp. 98–102 [table 2]) and computed a corresponding state-by-state compensating differential using 1980 county population totals as weights (U.S. Bureau of the Census, 1983). Given the counties reported by Blomquist et al., we were able to compute such estimates for 39 states and Washington, DC. If a state's compensating differential is negative, it is amenity-poor; if it is positive, it is amenity-rich.

TABLE 4—CONTINGENCY TABLE FOR AMENITY/DISAMENITY STATE CLASSIFICATION

QLI	$RY^* < 1$	$RY^* > 1$
Positive	21	7
Negative	3	9

Chi-square contingency-test statistic (H_0: no relationship) = 6.8, $P = 0.0092$

Note: QLI is the Blomquist et al.'s (1988) measure of compensating differentials weighted by population to correspond to states.

Table 4 presents a contingency table for our classification of states, based on point estimates of RY*, as amenity-rich or amenity-poor compared to our population-weighted extrapolations of Blomquist et al.'s county classifications to a state level for the set of 40 areas represented in their study. The results indicate that the two sets of classifications are significantly similar $X^2_{[1]} = 6.8$, $P = 0.0092$). A rank-order correlation test on these 40 observations produces a very significant Spearman's ρ value of -0.57 ($P = 0.0002$), which provides further evidence of the qualitative similarity of the results of the two studies.[10] We utilize these nonparametric tests because of the differences in spatial coverage between our state-based estimates and our extrapolations of the county-based estimates of Blomquist et al.

Table 4 indicates that the similarity in classification is much stronger for the 24 states classified as amenity-rich. For these 24 states, Blomquist et al. have the same classification as in the present paper for 21. For the 16 states we classify as amenity-poor, Blomquist et al. have the same classification for nine. For the ten states that are classified differently, the apparent difference

[10]Spearman's ρ is negative because, in our study, values of RY* below (above) unity indicate amenity-rich (amenity-poor) areas, whereas in the Blomquist et al. study, positive (negative) compensating differentials indicate amenity-rich (amenity-poor) areas. The data definitions therefore lead to a negative correlation.

could plausibly be due to sampling error. To account for this possibility, we performed the analysis with the 20 observations for which our interval estimates indicate that RY* is significantly different than unity. Of these 20 observations, 14 have corresponding values from the Blomquist et al. study. Of these 14, we identify ten as being amenity-rich. The Blomquist et al. results agree in nine instances. Of the four states we identify as being amenity-poor, the Blomquist et al. results agree in two cases. A difference in classification now occurs in only three instances; however, more agreement still exists for amenity-rich areas. Nevertheless, the extent of classification differences appears to be relatively minor.

What we find in our point-estimate-based comparative analysis is that, based on our model and data, the erroneous assumption of regional equilibrium in 1980 leads generally to an undervaluation of compensating differentials and only marginally to a qualitatively incorrect classification of regions. These results are robust with respect to estimates that adjust RY and RY* for industry-mix differences across areas at both the two-digit and four-digit standard-industrial-classification (SIC) code levels.[11] Our interval-estimate-based analysis is consistent with this but generally suggests that the extent of quantitative and qualitative errors stemming from the erroneous assumption of regional equilibrium is relatively minor, at least for 1980.

V. Summary and Conclusions

In this study, we have developed an improved model of net migration that encompasses both equilibrium and disequilibrium components. Instrumental-variables fixed-effects estimates of the model with time-series data for 51 areas over the period 1971–1988 support the importance of both equilibrium and disequilibrium factors in migration.

Solving our estimated model for the equilibrium condition of zero net migration and using point estimates, we find that 12 of 13 states in the West and 10 of 17 states in the South are amenity-rich. Corresponding figures for the Northeast and Midwest regions are four of nine and five of 12. Using confidence-interval estimates, we find that 15 areas can be classified as amenity-rich: seven in the West, four in the South, three in the Northeast, and one in the Midwest. These results indicate that western and southern areas generally have more attractive characteristics for individuals.

Several previous studies have estimated compensating differentials assuming regional equilibrium. We demonstrate that some states are not in equilibrium during the relevant period of these studies. By solving for the equilibrium condition of zero net migration and computing the corresponding compensating differentials, we find that the erroneous assumption of equilibrium causes estimates of compensating differentials to be understated for most areas. However, this can only be demonstrated at a statistically significant level for a few areas. We also find that the erroneous assumption of equilibrium leads to the misclassification of areas as amenity-poor (amenity-rich) when they are amenity-rich (amenity-poor) in only a few cases. Errors generated in the estimation of compensating differentials by erroneously assuming regional equilibrium therefore appear to be relatively minor, both quantitatively and qualitatively.

[11]The adjustment for industry mix removes variations in RY that arise because some areas have higher-paying industries than other areas. Such differentials may in part reflect compensations for industry characteristics. Our treatment of such variations follows, in spirit, the specification of industry dummies by Roback (1982, 1988) and the specification of occupational dummies by Blomquist et al. (1988).

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