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Recent years have witnessed widespread media attention and policy debate regarding the causes and consequences of population flight from California. While some analysts' reports link the reversal in California migration flows to cyclical swings in the state economy, other commentaries focus on alleged deterioration in California amenities and quality of life. This paper employs a logistic migration model to evaluate the role of economic and other location-specific effects in the determination of California domestic migration flows. The model is estimated using data for each of the 50 U.S. states for the 1981–1992 period. Various simulations of the model for the California case are then undertaken to indicate the effect of evolution in economic conditions and other location-specific effects on California net migration. A baseline simulation predicated on a reversion in the state's unemployment rate, wage, and house price differentials to average levels observed in the 1981–1992 period suggests a substantial slowing in California out-migration. Further, deterioration in California location-specific fixed effects, as estimated from the otherwise unexplained portion of the acceleration of out-migration in the more recent 1989–1992 period, serves to dampen the simulated improvement in California net migration only modestly. Overall, our research findings suggest that a large part of the unprecedented and sizable domestic out-migration from California is temporary, to be largely reversed in the context of a rebound in the California economy.

In recent years, widespread attention has been paid to the population flight from California. In marked contrast to previous periods, when large migration flows into California were the norm, the state recently has experienced a sizable net out-migration of residents, most notably to neighboring states. The reasons for this flight from California have been the subject of much debate. Some observers link the reversal in California migration flows to changes in the state economy, including the sizable cutbacks in defense and aerospace employment, which adversely affected the job and income prospects of many California residents. Moreover, the relatively high price of housing in California has long been a well-publicized concern for potential in-migrants, and many out-migrants from California are said to be attracted by lower house prices elsewhere. Finally, some analysts allege that there has been a deterioration in the California quality of life, due to such factors as increased congestion, diminished funding of public services and infrastructure, and heightened awareness of problems of public safety, such as violence and earthquakes.

This study evaluates the influence of economic and other location-specific conditions on California migration flows. In particular, the research quantifies the importance of labor market, income, house price, and other location-specific effects in a general model of state-to-state gross migration flows. We then apply the model directly to California's case to assess the proportion of the recent net out-migration that might reflect the alleged deterioration in California amenities and to determine whether a narrowing of the gap in economic activity between California and other states would lead to a reversal of the population flight of recent years.

From a state policy perspective, the likelihood of a reversal in the strong domestic net out-migration from California has important implications. For example, forecasts of California net migration significantly influence the debate over projected California budgetary deficits. In this regard, Bowman, et al., (1994) argue that migration projections from the State Department of Finance, which assume a surge in domestic net migration to California once the state's recession ends, lead to an overstatement of the cyclical portion of the state's budget deficit; Bowman, et al., argue that a sizable portion of the current state budget deficit is structural and will not be significantly ameliorated by a reversal in population flows.

Moreover, an understanding of California migration trends would be helpful for business planning and public policy development throughout the West. The recent exodus from California imparted some spillovers on neighboring states, and decisionmakers would like to know whether the exodus is likely to persist. On the plus side, migrants from California served to spur job growth in destination states, by augmenting the labor force and through the multiplier effects of their spending. Moreover, the financial and human capital brought by ex-Californians boosted local investment in housing and other sectors in the destination states. In contrast, the influx of Californians into neighboring states has generated a host of complaints as well, mainly centered on rising house prices and increased congestion.

The theoretical model we use to explain migration patterns is unique in several senses. First, we explain both inflows and outflows of people from each state—the gross migration patterns—whereas most preceding studies of time-series migration patterns explain only the net flows. Second, we parameterize the unobserved amenity values of residing in particular locations as fixed effects. Greenwood, Hunt, Rickman, and Treyz (1991) also have applied a fixed effects approach to measuring relative amenities, but their work is in a net migration context. The closest precedent to our approach to modeling gross migration flows is Frees (1993), who also uses fixed effects. However, his estimating equations are less directly tied to a theoretical model than our equations are, and he does not focus on the gross migration model's implications for net California migration, as we do.

The model we estimate closely fits the actual California net migration experience over the 1981–1992 period, for which comprehensive estimates of interstate migration are available. Separate data from the U.S. Bureau of the Census and the California Department of Motor Vehicles (DMV) suggest that net out-migration from California accelerated further in 1993 and 1994, which also is consistent with the model's predictions. Looking ahead, simulations of the model suggest that a narrowing of the gap between economic conditions in California and other states would result in a sharp reduction in net migration out of California. We also estimate a variant of the model that allows for a recent shift in the California-specific fixed effects, so as to test for the perceived deterioration in the state's quality of life or other structural shifts. Because the model fits well without such an allowance for structural change, the results suggest that only a relatively small portion of the recent net out-migration is due to structural change.

The organization of the paper is as follows. The next section presents recent trends in migration and economic activity as pertains to California and the West. Section II describes specification, estimation, and findings of a logistic model of the determinants of U.S. state-to-state migration. In Section III, the model is simulated for a number of possible future paths of the California economy, and the implications for the state's migration balance are described. Our brief concluding section reviews the major findings of the paper, and an appendix discusses the robustness of the results to alternative econometric specifications.

I. Migration over the Economic Cycle: California and the Twelfth Federal Reserve District

In the last few years, California has experienced a net out-migration of people for the first time in the post-World War II period (see Figure 1). In focusing on the likelihood that such outflows will persist, this article will address only California domestic migration—that is, the movement of people between California and other U.S. states—which has been the source of the recent net outflow; the article will not explicitly address immigration across national borders which is estimated to have been more stable.[1]

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1. Throughout the 1980s, a net increase of more than 200,000 immigrants from abroad per year was augmented by somewhat less than 100,000 net domestic arrivals. More recently, the flow of immigrants has continued strong, but net domestic migration to California turned negative. Recent Census Bureau estimates place the net outflow of the sum of domestic and international migrants from California at about 70,000 persons in fiscal year 1992–1993 and at about 140,000 people in fiscal year 1993–1994. The IRS data on address changes, which provide the basis for the Census Bureau domestic migration estimates, show a fiscal year 1987-1988 peak in net domestic migration, followed by substantial net outflows in recent years (Figure 2)

Although we do not explicitly model the influence of immigrant arrivals on the domestic migration decision, the effects of immigration are included implicitly, insofar as they affect our measured economic conditions or the amenity values of residing in particular locations, which we parameterize and estimate statistically. However, our study is relatively aggregative and omits potential influences of immigration on domestic migration that could appear in a more disaggregative study.

For example, an important strand of the literature on the economics of immigration—recently reviewed well by Borjas (1994)—focuses on the impact of immigrants on the earnings and employment of original residents (natives). A common argument in this literature is that an influx of lower-skilled, lesser-educated immigrants should have a larger impact on lower-skilled, lesser-educated natives than on higher-skilled, better-educated natives. Although empirical results on this phenomenon are mixed, Frey's (1993, 1994) studies of the 1990 Census results did show that states such as California which received large immigrant inflows in the late 1980s had relatively large outflows of less-skilled, lower-paid native workers, which is consistent with a strong labor-market substitution effect. However, using more recent data from State of California tax returns, Bolton (1994) suggests that there might have been some shift in the income distribution of migrants from California. Although the 1990 Census data show that the median income of those who left California in the 1985–1990 period was less than the statewide median income, Bolton's more recent data show that between 1989 and 1992 domestic out-migrants from California had a higher average income than the statewide average. The prevalence of relatively high wages among recent out-migrants is consistent with either the (contentious) view that the out-migrants seek to avoid the public fiscal distress (higher taxes, lower levels of government service) associated with high rates of immigration or the view that recent adverse shocks to California labor demand have been concentrated in traditionally high-paying industries, such as aerospace.

FIGURE 1

California Net Domestic and International Migration (Thousands of Persons)

Fiscal Year
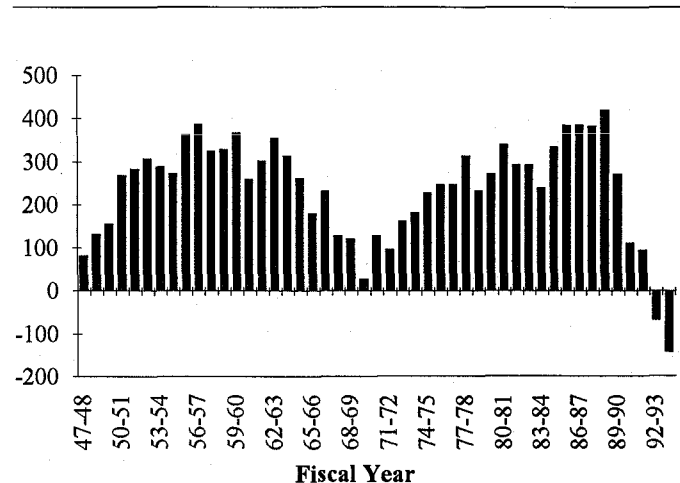
Source: Components of Population Change
U.S. Bureau of the Census and State Department of Finance

The sharp reversal in California net domestic migration flows was spurred in part by state-level variations in the extent and timing of business cycle fluctuations. In the early 1990s, employment opportunities were reduced substantially during the severe downturn in the California economy, with especially sizable layoffs in aerospace and other defense industries. As shown in Table 1, during the 1989–1992 period of weakness the unemployment rate in California averaged 6.9 percent, well in excess of the average rate recorded by other states of the Twelfth District and by the United States as a whole.[2] In contrast, in the earlier periods of net in-migration to California—the 1981–1983 recession and 1984–1988 expansion—the California unemployment rate was a bit below the average unemployment rate in the remainder of the Twelfth District and about the same as the average unemployment rate in the United States.

House prices and other components of the cost of living have continued to run quite a bit higher in California than in other states. To a certain extent, the labor force has been compensated for the higher cost of living by higher nominal wages and salaries. For example, in 1981–1983, per

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2. The Twelfth District of the Federal Reserve System comprises Alaska, Arizona, California, Hawaii, Idaho, Nevada, Oregon, Utah, and Washington.

TABLE 1

STATE CHARACTERISTICS

	UNEMPLOYMENT RATE (percent)			WAGES AND SALARIES PER CAPITA (thousands of $ per year)			HOUSE PRICES (thousands of $)		
	81–83	84–88	89–92	81–83	84–88	89–92	81–83	84–88	89–92
California	9.0	6.5	6.9	7.8	10.1	12.2	121.7	155.6	211.5
12th District	9.5	7.0	5.8	6.8	8.3	10.5	88.3	108.4	137.1
US	9.0	6.7	6.3	6.8	8.8	11.0	79.4	104.6	133.2

NOTES: The statistics for the aggregate groups, the 12th District except California and the United States, are population-weighted averages of the statistics for the individual states. The United States group excludes the District of Columbia.

capita wage and salary income in California averaged about $1,000 per person higher than in other states, and nominal incomes have grown at about the same pace in California as in other states since then. However, with house prices rising much faster in California than in other states, wage rates adjusted for the cost of living became relatively more attractive elsewhere. By 1989–1992, the average of new and existing home prices (on mortgage transactions) in California had increased to $211,500, pushing the ratio of wages to house prices down to 5.8 percent from 6.4 percent in 1981–1983. In contrast, for the Twelfth District except California, the average ratio of nominal wages to house prices has held steady at 7.7 percent since the 1981–1983 period.

States neighboring California were the primary recipients of the sizable and unprecedented net population outflows of population from the state during the early 1990s. Among states of the Twelfth District, the largest net outflows during the 1989–1992 period were to Washington, Oregon, and Nevada. Relative to state population levels as recorded in the 1990 decennial census, Californians relocating in Nevada and Oregon constituted a full 5 and 3 percent of destination state population, respectively.

Shorter distances undoubtedly are an important reason that flows to adjacent states from California were large. However, some of the largest recent net outflows of population from California also have been to other District states at better positions in their regional business cycles, to states with relatively high wages, and to states with more affordable housing (Table 2).[3] The California net migra-

tion flows were particularly large to nearer, more populous District states with relatively large unemployment differentials (computed as California less other state). For example, the unemployment rate differential between California and Nevada widened to 1.3 percentage points in 1989–1992, about ¾ percentage point more than the long-run average differential; during that same period, net migration to Nevada accelerated to more than 16,000 persons per year.

In all regions, nominal wages and salaries per capita and house prices have trended upward, so unadjusted wage or house price differentials between states are not very comparable over time. Accordingly, in Table 2 (and the econometric modeling that follows), we normalize each state's wage and house price by the average for the United States in that year. The entries in the second and third rows of Table 2 show the differentials between California and other states in relative wages and relative house prices. Consistent with our earlier comment that wages in California increased at about the same pace as wages in the average of other states in the District, there also has been little change over time in relative wage differentials on a state-to-state basis. For example, in the 1989–1992 period, California's relative wage differential with Nevada was

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3. A few transformations of the data are useful in illustrating this point. For one thing, states can have markedly different long-run average

unemployment rates, so the raw unemployment rate differential is not always a reliable guide to the states' relative positions in their regional business cycles. For example, the unemployment rate in Alaska has averaged about 1¾ percentage points more than the California unemployment rate, owing, in part, to the geographic segmentation of labor markets within the expansive Alaskan territory. In contrast, Hawaii has tended to have a relatively low unemployment rate. Thus, in interpreting the 1989–1992 unemployment rate differential between California and other District states, it is useful to compare the current differentials with the mean unemployment rate differentials.

TABLE 2

STATE-TO-STATE DIFFERENTIALS BETWEEN CALIFORNIA AND OTHER TWELFTH DISTRICT STATES

	ALASKA	ARIZONA	HAWAII	IDAHO	NEVADA	OREGON	UTAH	WASHINGTON
1981–1992:								
Unemployment rate	−1.78	.80	2.92	.12	.50	−.50	1.55	−.65
Relative wage	−.39	.24	.01	.41	−.03	.25	.30	.12
Relative house price	.31	.51	−.17	.63	.38	.65	.54	.48
Distance	2363	357	2568	671	228	827	580	961
1989–1992:								
Net migration–Actual	−.11	−9.60	−.68	−5.27	−16.36	−20.89	−2.96	−23.48
Unemployment rate	−1.04	.93	3.72	.96	1.30	.67	2.14	.63
Relative wage	−.18	.25	−.08	.38	−.03	.22	.29	.08
Relative house price	.59	.69	−.28	.69	.49	.76	.64	.48
1993–1994:								
Net migration–Implied	−4.38	−46.84	−10.93	−5.55	−18.60	−14.16	−7.30	−29.48
Unemployment rate	1.06	2.85	3.99	3.19	2.49	2.41	5.19	2.09
Relative wage	−.22	.17	−.15	.27	−.11	.12	.17	.01
Relative house price	.32	.66	−.62	.80	.49	.56	.46	.43
At 1981–1992 Mean:								
Net migration–Implied	−1.87	−23.43	−.90	−.94	−9.45	−4.85	.51	−10.65

NOTES: The unemployment rates are measured as a percentage of the labor force. The relative wages and house prices are proportions of the national average wage and house price. Distance is in miles. Net migration is in thousands of address changes per year, with fiscal year data shown as if it pertained to the calendar year at the beginning of the fiscal period.

3 percent of the national wage, the same as the long-run average differential. Over the sample period, the California house price differential widened noticeably with respect to all other District states except Hawaii and Washington.[4]

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4. House prices increased much faster in Hawaii than in California over the sample period. Although Washington house prices were near the national average at the beginning of the sample period, they increased faster than the national average in the late 1980s, leaving the California-Washington house price differential close to 48 percent of the national average throughout the sample period.

II. THE DETERMINANTS OF PLACE-TO-PLACE MIGRATION

The data used in this paper are a combination of Internal Revenue Service data on place-to-place migration flows and other published information on state-level economic, geographic, and population characteristics. State level migration flows for the 1981–1992 period are taken from IRS data covering people filing tax returns or listed as exemptions on others' tax returns for consecutive years. State-to-state empirical migration rates are computed as the number of individuals moving from state i to state j in year

t, as a percentage of the total number of people initially residing in the origin state i in that year.[5] For each of the 12 years of the analysis, the result is a 50×50 contingency table (exclusive of the District of Columbia) for which the off-diagonal elements represent estimated place-to-place regional migration probabilities. Our in-sample analysis focuses on these 29,400 off-diagonal elements.[6]

Following Fields (1979, 1982), Schultz (1982), and Gabriel, Shack-Marquez, and Wascher (1992, 1993), among others, we choose a logistic specification for the multivariate choice problem.[7] Given the choice among a finite number of destinations, individuals are assumed to choose the location yielding the highest expected net discounted return on migration.[8] Specifically, the probability of migration from state i to state j in period t, Π_{ijt}, is assumed to be

(1) $\Pi_{ijt} = \exp(Z_{ijt})/[\Sigma_k \exp(Z_{ikt})]$ $i,j = 1,...,50$; $t = 1,...,\tau$,

where the Z variables are indices of the expected return to moving to particular places. The common normalization factor, $\Sigma_k \exp(Z_{ikt})$, constrains these individual probabilities to sum to unity.[9]

For derivation of an estimating equation, we restate equation (1) in terms of the (logarithm of) the ratio of the probability of migrating from state i to state j in year t (Π_{ijt}) to the probability of remaining in state i (Π_{iit}). Equation (1) implies that this logarithm of the odds ratio is simply the difference between an index of the expected return to moving to that specific state j (Z_{ijt}) and the expected return to remaining in the origin state i (Z_{iit}):

(2) $\ln(\Pi_{ijt}/\Pi_{iit}) = Z_{ijt} - Z_{iit}$ $i,j = 1,...,50$; $i \neq j$ and $t = 1,...,\tau$.

Accordingly, we use the logarithm of the ratio of the empirical migration rate to the rate at which individuals remain in the origin state as the dependent variable, Y_{ijt}.[10]

In specifying the index variables, Z_{ijt}, we assume that the value of migration between two states varies directly with

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5. A clarification of the dating of our variables is in order. The IRS migration flows labeled, for example, with fiscal year 1991–1992 or calendar year 1992 pertain to address changes between the filing periods for the 1991 and 1992 tax years; this reflects address changes from roughly April 1992 to April 1993. The population variable labeled 1992 is an estimate for July 1, 1992, and the economic variables labeled 1992 are annual averages of economic conditions throughout that calendar year.

6. In the regressions reported below, only 29,292 of the 29,400 observations are used. We lose about 100 observations due to nonavailability of the housing price data for one of the states in a single year and a few observations from censoring in the IRS data to honor nondisclosure restrictions.

7. Despite the richness of the IRS dataset, relatively little research has taken advantage of the state-to-state gross migration time series. One exception is Frees (1992, 1993), whose primary interest has been in developing short-term forecasting models of gross migration rates. Frees (1992) explores univariate time-series models and shows that origin-destination-specific fixed effects can explain a substantial portion of the variance in out-migration rates. Frees (1993) also considers the potential usefulness of selected economic variables in short-term forecasting; he shows that the inclusion of changes in income and employment as explanatory variables only marginally improves upon a random walk model for forecasting origin-destination-specific migration rates. However, it is difficult to interpret the role of underlying economic and other forces driving migration in Frees' models with profligately parameterized fixed effects.

8. Implicitly, the choice of remaining in the region is subsumed in this decision process as is the case when the net discounted return on migration is less than or equal to zero for all possible destinations.

9. The logistic specification is quite useful for application to time-series of aggregated place-to-place migration data, such as in this study. However, some weaknesses are likely to become more apparent if the logistic specification were used to model the multi-dimensional dis-

crete choices of individual households. For example, the logistic specification assumes the "independence of irrelevant alternatives" (IIA), which means in the migration context that the economic conditions in areas other than the ultimately chosen destination have no effect on the desirability of residing in the chosen destination. The most likely form of violation of IIA concerns repeat migration. Individuals moving for a second time are much more likely to return to their initial origin state than to move on to other states, other conditions equal. For potential repeat migrants, the evolution of economic conditions in the initial origin state likely affects the evaluation of alternative destinations, and the logistic specification would not incorporate such dependence.

10. The empirical migration rate, or relative frequency of migration, is the number of migrants from state i to state j in year t (Π_{ijt}) as a proportion of the population in the origin state i in year t. The dependent variable, Y_{ijt}, expresses (the logarithm of) this migration rate as a proportion of the migration rate for the diagonal cells, Π_{iit}, which capture the fraction of individuals remaining in the origin state. Unfortunately, the information needed to construct the diagonal cells is directly available only for selected years in the IRS dataset counting exemptions on matched returns. However, the numerator for the diagonal cells, M_{iit}, can be constructed indirectly as $M_{iit} = (POP_{it} - \Sigma_{q \neq i} M_{iqt})$, where POP_{it} is the Census estimate of the noninstitutionalized resident population in the origin state, and the M_{iit} and M_{iqt} variables are migration flows denominated in terms of numbers of people, not migration rates. One drawback of this approach is that the IRS migration data and Census population estimates are not fully consistent; in fact, our constructed estimate of the remaining population likely overestimates the true number of stayers by a noticeable amount. We infer this because over the 1990–1992 period for which the diagonal cells of the IRS dataset are available, the implicit population level of matched exemptions (ΣM_{iqt}) for California is only about 77 percent of the Census Bureau's estimate of the actual population level in California. For all U.S. states, the coverage ratio averages about 80 percent and ranges from about 75 percent to 90 percent, and the state-specific coverage ratios appear to be quite persistent over time.

There are a number of reasons why the count of matched exemptions tends to fall short of the actual population. First, not all individuals are required to file federal taxes, and some required filers fail to comply. Second, as discussed by Frees (1992), typically a small fraction of the returns for a given year have not yet been received when the matching process is executed. Third, the returns are matched based on the Social Security number of the primary filer, and data entry and other errors can

economic opportunity and location-specific amenities in the destination region and inversely with economic conditions and location-specific amenities in the origin region. Also, irrespective of the characteristics of potential destinations, origin region populations are assumed to have different propensities to migrate, depending on, for example, individuals' positions in the overall life-cycle. The empirical literature suggests that the migratory propensity of individuals—the raw tendency of households to move—peaks when such individuals are in their 20s and also is higher for people with a college education; moves are especially common at times of career job choice and marriage. As individuals mature, have families of their own, and develop job, neighborhood, and social ties, mobility declines. We attempt to capture the life-cycle phenomenon concisely by focusing on the population age structure.

More specifically, the index Z_{ijt} is comprised of a linear combination of the relevant economic and location-specific amenity-related characteristics of the origin and destination states (X_{it} and X_{jt}), the transactions costs of moving between i and j (D_{ij}), and a population age-structure proxy for individual traits associated with the propensity to migrate (T_{it}):

$$(3)\ Z_{ijt} \begin{aligned} &= \lambda X_{it} + \Phi X_{jt} + \gamma T_{it} + \delta D_{ij}, \quad i \neq j\ ; \\ &= \lambda X_{it} + \Phi X_{jt} - \gamma T_{it}, \qquad\quad i = j\ ; \end{aligned}$$

$$i,j = 1,\dots,50 \text{ and } t = 1,\dots,\tau.$$

Equation (1) can be rewritten in terms of the difference between the elements of Z_{ijt} and Z_{iit} as[11]

$$(4) \qquad \ln(\Pi_{ijt}/\Pi_{iit}) = \Phi(X_{jt} - X_{it}) + 2\gamma T_{it} + \delta D_{ij}.$$

This particular specification implies that the origin and destination economic and amenity conditions, X, operate symmetrically on migration, with Φ representing the parameter vector associated with differentials between origin and destination state conditions.[12] The symbol γ represents

the parameter vector associated with origin population characteristics proxying the propensity to migrate. Transactions costs associated with migration are assumed to be time-invariant in our specification; δ represents the parameter associated with this effect.

Adequately measuring the relevant economic and amenity conditions is difficult. With regard to measurable economic conditions, in addition to job prospects, the cost-of-living and wage levels in an area clearly are important in the migration decision. However, amenities are capitalized to some extent in house prices and wage rates, so controlling for differences in amenities across areas also is important to modeling the effects of differences in such measurable economic conditions. One approach is to follow Blomquist, Berger, and Hoehn (1988), among others, in trying to identify and measure explicitly the large number of relevant location-specific amenities important to potential migrants.[13] However, such an approach likely would impart problems of omitted and misspecified amenity variables into the analysis. Another problem with including explicitly measured amenities in a migration analysis is that not all amenities are freely provided and fully capitalized in house prices and wage rates. For example, Gyourko and Tracy (1991) point out that fiscal differentials—variations in the extent to which amenities are priced through taxation and other fiscal means—also have a strong influence on quality-of-life index values. Such state fiscal conditions also should be included in any explicit attempt to sort out the pecuniary and nonpecuniary influences on migration.

Absent comprehensive information on amenity-related variables by state, we parameterize amenities using a fixed effects approach, akin to the Greenwood, Hunt, Rickman, and Treyz (1991) use of fixed effects in a net migration model. Specifically, we represent the amenities in destination location j by a single parameter, A_j^d, and amenities in the origin location i by the parameter A_i^o.[14] We further assume that the composite index of the effective attractiveness of an area, ΦX, can be represented as a linear

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inhibit proper matching on this basis. Nevertheless, the IRS data appear to be the best available source for estimates of interstate migration. See, for example, Wetrogan and Long (1990) for a comparison with other sources.

11. Note that both T_{it} and D_{ij} remain in the model after the normalization is made. In the first case, this occurs because elements of T_{it} that are positively associated with the general propensity to migrate have a positive influence on any Z_{ijt} for which $i \neq j$, but a negative influence on any Z_{iit} and hence do not cancel out when the difference is taken. In the second case, we assume that there are no transactions costs associated with not migrating ($D_{ii} = 0$) so that this variable does not appear in Z_{iit}.

12. Alternatively, asymmetries in information flows about economic and other conditions in the origin and destination regions or asymmetries in

the returns to migration may give rise to unequal effects on migration from origin and destination variables. In the empirical investigation that follows, we report on the effect of relaxing the symmetry restriction on the migration coefficients.

13. For example, the degree of air pollution can be approximated by the amount of particulates in the air, the pleasantness of the weather can be represented by data on temperatures and rainfall, and the quality of public schools can be approximated by educational testing results or measures of educational expenditures.

14. In contrast to the directional migration specification contained in this paper, the Greenwood, Hunt, Rickman, and Treyz (1991) paper was limited to an analysis of net migration. Accordingly, the authors did not

combination of the amenity parameter and three measures of economic conditions—the unemployment rate (U), the relative wage rate (W), and the relative house price (P):[15]

(5) $\Phi X_{st} = A^o_s + \beta_1 U_{st} + \beta_2 W_{st} + \beta_3 P_{st}$

for s indexing origins (i)

$\Phi X_{st} = A^d_s + \beta_1 U_{st} + \beta_2 W_{st} + \beta_3 P_{st}$

for s indexing destinations (j).

Accordingly, the state-to-state migration model that we estimate is:

(6) $Y_{ijt} = \alpha + \beta_1 (U_{jt} - U_{it}) + \beta_2(W_{jt} - W_{it}) + \beta_3 (P_{jt} - P_{it})$

$+ 2\gamma T_{it} + \delta D_{ij} + \Sigma_s A^d_s F_{is} - \Sigma_s A^o_s F_{js} + \varepsilon_{ijt}$,

where the index of states (s) in the sum over the state-level dummy variables (F) runs from the first to forty-ninth state; Wyoming is the reference state subsumed in the estimate of the intercept α. Thus, the analysis implicitly includes 100 fixed effects, one for each state as a migration origin and one for each state as a migration destination. We represent the propensity to migrate by a linear combination of a vector, T, of origin state population age structure variables. Transactions costs, D, are proxied by the distance between the most populous city in the origin state and the most populous city in the destination state.[16] The error term in equation (6), ε, captures measurement and specification error.

Estimation results are shown in Table 3. The model explains approximately 80 percent of the variation in the standardized probability of state-to-state migration over the 1981–1992 period. In general, the coefficient estimates are consistent with a priori expectations. A high unemployment rate in the destination state relative to the origin state decreases migration, while relatively high wages in the destination state increase migration.[17] In addition to having the appropriate signs and being precisely estimated, the coefficients on the unemployment and wage differentials appear to be important in explaining interstate population flows. In contrast, the coefficient on state level house price differentials is not significantly different from zero.[18]

The results also suggest higher migratory propensities among younger household heads. As indicated in Table 3, beyond the baseline 0–17 years age category when individuals tend not to be household heads, the estimated coefficients decline monotonically with the age of the origin region population and turn negative and sizable for the 45 years and older age groups. This is not surprising, as the expected economic return to migration is a discounted stream of income over the remaining work years in the destination region, and this discounted value is less likely to exceed the transactions costs of moving if the income stream under consideration is of short duration.

Space limitations preclude presentation and discussion of the full set of state-level fixed effects, although Table 3 does show the estimated fixed effects for California. Relative to the baseline state (Wyoming), California is both

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distinguish between the amenity value for a location as an origin and its amenity value as a destination. We choose not to impose equality of the fixed effects for a place as an origin and as a destination, in part because the local fiscal situation depends on whether a resident is coming or going. For example, according to California's Proposition 13, homeowner property tax assessments vary substantially as a function of the timing of the purchase of the home. In general, our results suggest the inappropriateness of the restriction of equal fixed effects for a place as an origin and a destination.

15. The relative wage and relative house prices are defined as described in the discussion of Table 2, as that state's nominal wage or house price as a proportion of the national average wage or house price in the given year.

16. The distance measures were computed from the longitudinal and latitudinal co-ordinates of the most populous cities in the states, using the information supplied in the U.S. City SAS map dataset. The wage variable is computed from BEA's measures of state-level personal income; the population levels implicit in the published estimates of overall per capita income were used to convert the published estimates of wage and salary income to a per capita basis. The house price variable is the average price of homes involved in selected mortgage transactions in the state as published by the Department of Housing and Urban Development. The estimates of the population age structure were compiled from a variety of Census Bureau sources.

17. Strictly speaking, the statement that a high unemployment rate in the destination state relative to the origin state decreases migration applies in terms of deviations from state-specific mean unemployment rates. For simplicity of exposition, we have motivated the fixed effect parameters solely in terms of amenity values, but, in practice, the estimates of the fixed effects also will control for other state characteristics, such as high average unemployment rates.

18. The lack of significance of house price differentials in the determination of place-to-place migration flows is somewhat at odds with the results in Gabriel, Shack-Marquez, and Wascher (1992), which indicated that relatively high destination house prices deterred migration, whereas high origin region house prices have little effect on migration. The earlier findings were derived from a purely cross-sectional model for the late 1980s, using quality-adjusted new house price data for the nine census divisions; location-specific fixed effects were not included in the analysis. Because the constant-quality house price series are not available for existing homes or at the state level, the more geographically disaggregate, time-series analysis presented here uses non-quality-adjusted prices of a composite of new and existing houses.

In addition, we impose the symmetry restrictions in this study, primarily for expository convenience. If the symmetry restrictions on the unemployment rates, wages, and house prices are relaxed, the overall fit of the model barely changes, and the estimated coefficients on the unemployment rates remain near the estimated values from the restricted version. The house price coefficients remain indistinguishable from zero.

TABLE 3

ESTIMATES OF THE LOGISTIC MODELS OF STATE-TO-STATE MIGRATION

EXPLANATORY VARIABLE	ESTIMATED COEFFICIENTS		
	Without Structural Change	With Structural Change	
Unemployment rate differential (β_1)	−.053 (−20.71)	−.053 (−20.64)	
Relative wage differential (β_2)	.346 (7.13)	.343 (7.04)	
Relative housing price differential (β_3)	.009 (.35)	.011 (.44)	
Distance (δ)	−.001 (−172.65)	−.001 (−172.65)	
Population age structure			
18–24 years ($2\gamma_1$)	4.620 (4.02)	4.496 (3.85)	
25–44 years ($2\gamma_2$)	1.061 (1.15)	1.026 (1.11)	
45–64 years ($2\gamma_3$)	−2.469 (−1.66)	−2.510 (−1.67)	
65+ years ($2\gamma_4$)	−4.115 (−3.26)	−4.152 (−3.26)	
	Full-sample	1981–1988	1989–1992
California-as-origin (A^o_5)	.450 (7.90)	.439 (7.06)	.458 (7.13)
California-as-destination (A^d_5)	3.844 (94.16)	3.872 (87.62)	3.784 (68.94)
Memo:			
Goodness of fit (R^2)	.79	.79	

NOTES: The OLS regression was fitted through 29,292 observations on the dependent variable, y_{ijt}, over the 1981–1992 sample period. Additional explanatory variables included, but not shown above, were an intercept and fixed effects for each other state (except Wyoming) as an origin and as a destination. Conventional t-statistics for the hypothesis of a zero coefficient are shown in parentheses.

more likely to be an origin and more likely to be a destination, other things equal. More important, relative to all other states, the estimated fixed effect for California-as-origin is near the middle of the range. In contrast, the estimate of California-as-destination is quite high and is exceeded only by Florida-as-destination. Correspondingly, the model implies that there would be a strong net in-migration to California if unemployment rates, wages,

house prices, and age structures were equalized across states.

As reported in the second column of Table 3, we also evaluated the temporal stability of the California fixed effects estimates. Specifically, the migration model was re-estimated with interactive terms on the California origin and destination fixed effects, one set for the pre–1988 period of net in-migration to California and another set for

the 1989–1992 period of net out-migration from California. Theoretically, this less restrictive model lets the fixed effects estimates over the more recent period reflect the conjectured deterioration in California amenities, the alleged decline in the "quality of life" that has been the topic of media commentaries. The estimates of the California fixed effects for the more recent period are consistent with the deterioration hypothesis, although the coefficients are not markedly different from the estimates for the 1981–1988 period.

These results are shown more clearly in Figure 2, which plots actual California net migration flows over the 1981–1992 sample period, together with fitted values for California net migration over that same period as derived from the two estimated place-to-place migration models. Both the models "with" and "without" structural change fit the actual pattern of net domestic migration relatively well, particularly in the years since 1988. However, allowing for structural change in the California fixed effect boosts implied net in-migration prior to fiscal year 1989–1990 by about 14,000 persons per year and lowers migration by about 23,000 persons per year for the 1989–1992 period. Thus, the model with structural change is consistent with a swing in the desirability of residing in California sufficient to induce a drop of about 37,000 people per year in net in-migration to California. Actual California migration ranged from a peak net inflow of about 100,000 in fiscal year 1987–1988 to a net outflow of about 170,000 in fiscal year 1992–1993, and so the model with structural change attributes at most only about 20 percent of the swing in net migration to changes in the California "quality of life."

The parameter estimates presented in Table 3 clearly indicate that migration is influenced by economic conditions. However, it is difficult to discern the relative magnitudes of the economic effects on net migration from the parameter estimates alone because the model is specified in terms of logarithms of gross migration flows. For this reason, it is useful to translate the parameter estimates into the effects on net migration. However, once the model is transformed in this manner, it is no longer linear in the explanatory variables, and the standard, additive decomposition of the contributions of the explanatory variables is not applicable. Figure 3 shows an alternative method of illustrating the relative importance of selected explanatory variables.[19] In particular, the contribution of each variable is calculated from comparing the overall fit of the model

FIGURE 2

CALIFORNIA NET DOMESTIC MIGRATION
(THOUSANDS OF ADDRESS CHANGES)
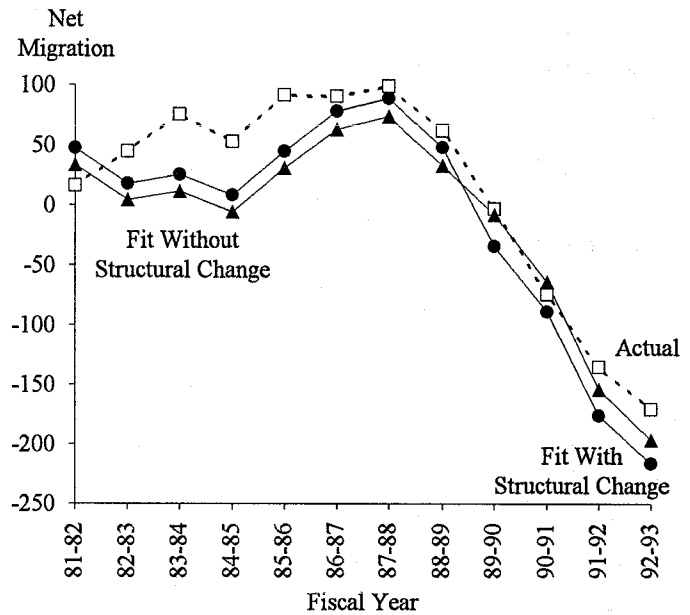
to a simulation that isolates the contribution of the variable in question to the overall fit. The simulation permits the relevant variable to vary as it did over history, but holds the other variables steady at their sample means. In calculating these simulated contributions, the origin state population levels also are allowed to evolve over time.

The upper left panel of the figure shows that changes over time in unemployment rate differentials have dominated the fit of the model for California net in-migration over the sample period and, in particular, explain a large portion of the decline in net migration after 1987. Time-series movements in the relative wage differential induced little change in net in-migration to California over 1982–1987, but relatively slow growth in California wages after that induced some additional migration to other

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19. Figure 3 illustrates the contributions to the fit for net migration. We also describe here the contributions of the variables to the fit for the logarithmic gross migration rates by reporting how the goodness-of-fit of

migration equations such as (6) change as variables are added incrementally to the regression specification. Proceeding in order of importance, the fixed effects contribute 56 percentage points to the R^2. Adding the distance variable boosts the explanatory power about 22 percentage points further, to 0.78 of the variance in the dependent variable. The unemployment differential explains only about 1 percentage point of the variance in the dependent variable, and the remaining economic variables explain even less. These statistics are indicative of the fact that the cross-sectional variance in the dependent variable dwarfs the portion of the variance in the dependent variable due to changes in economic conditions over time.

FIGURE 3

CONTRIBUTIONS TO FIT
FOR CALIFORNIA NET DOMESTIC MIGRATION
(THOUSANDS OF ADDRESS CHANGES)

Unemployment Differential

Relative Wage Differential

Relative Housing Price Differential
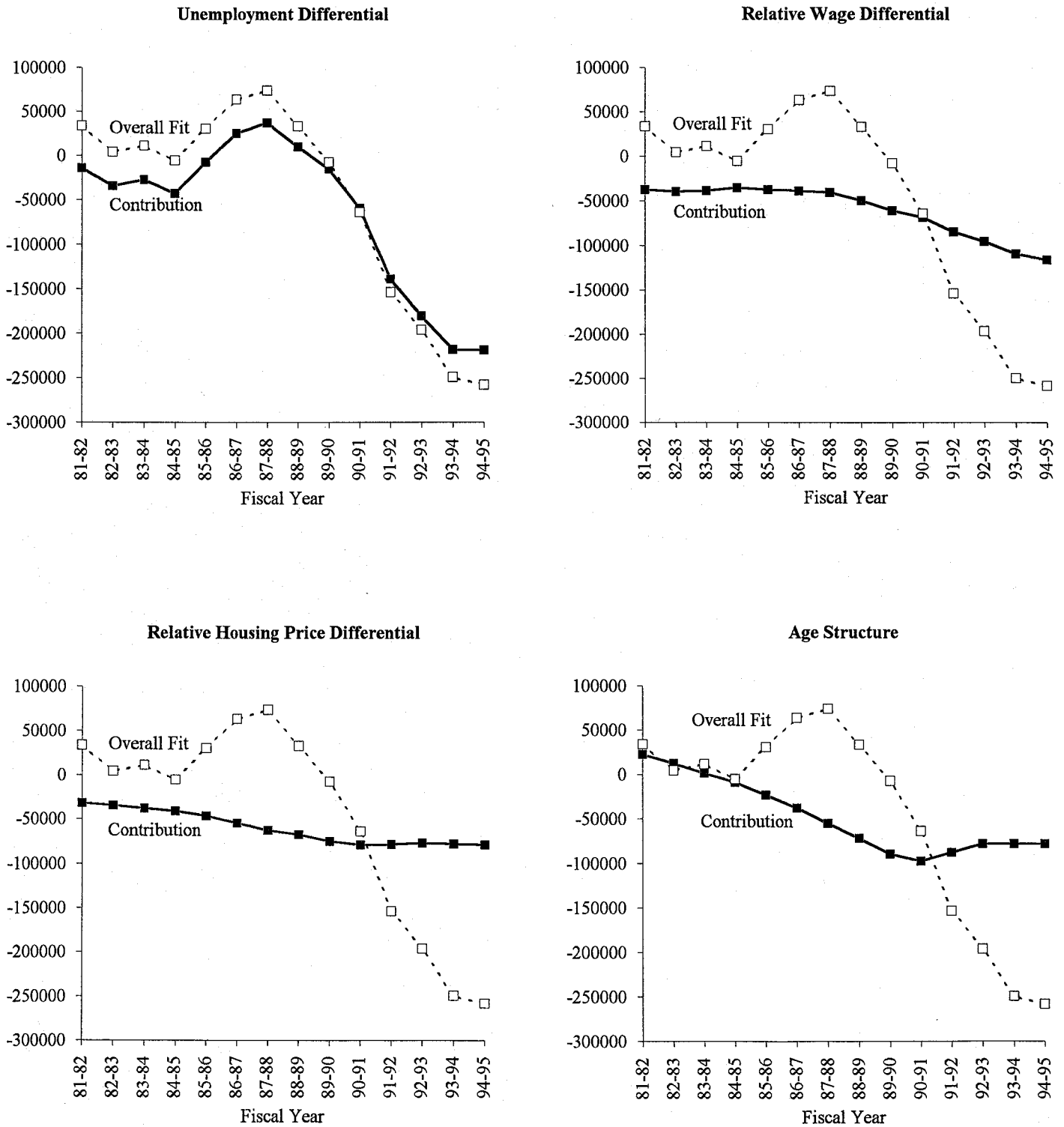
Age Structure

states. Changes over time in relative house price differentials had only a minor effect on the fit, while the evolution of the age structure was a more noticeable component of the net out-migration from California.[20]

III. MODEL SIMULATIONS OUT-OF-SAMPLE

In 1993 and 1994, the divergence between economic opportunities in California and other states widened further, so the model predicts an acceleration of net out-migration from California. The unemployment rate differential between California and other Twelfth District states widened from about 1 percentage point in 1989–1992 to about 2¾ percentage points, on average, in 1993–1994. Relative to California, employment prospects improved in every other Twelfth District state (Table 2). Job opportunities were particularly strong in Idaho and Utah, where the unemployment rate differentials exceeded 3 percentage points in 1993–1994. Recent relative wage developments also have induced further out-migration from California (Table 2 and Figure 3). During 1993–1994, per-capita wage and salary income grew faster in low unemployment rate states —such as Idaho, Nevada, and Utah—than in California.[21]

Specifically, given the recent historical values of the economic variables, the fit of the model ("without structural

FIGURE 4

ALTERNATIVE SIMULATIONS
FOR CALIFORNIA NET DOMESTIC MIGRATION
(THOUSANDS OF ADDRESS CHANGES)
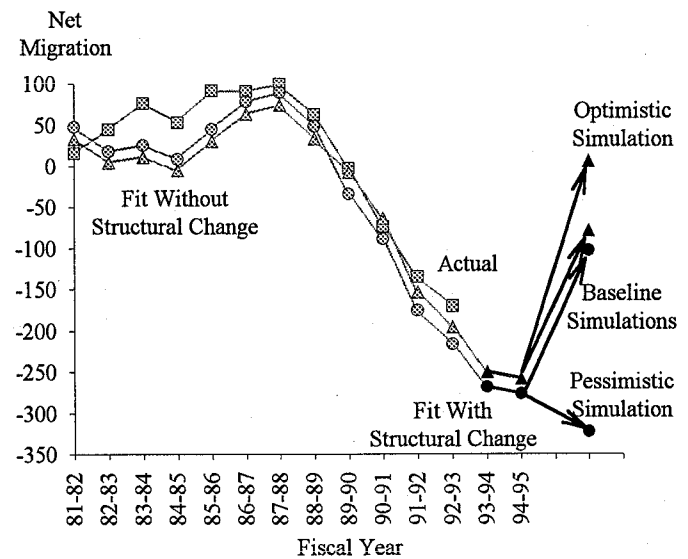
change") implies that net out-migration from California accelerated from almost 200,000 persons in fiscal year 1992–1993 to about 250,000 persons in fiscal year 1993–1994 and 258,000 persons in fiscal year 1994–1995 (Figure 4).[22] Although IRS data on interstate migration is publicly available only through fiscal year 1992–1993 at present, other data sources suggest that actual out-migration for California did accelerate further in recent years. For example, in the recent Census Bureau estimates of the components of population change for the state—which incorporate as yet unpublished IRS data on address changes in fiscal year 1993–1994—net domestic out-migration from California is shown to have increased by about 45,000 persons that year. Data on address changes from the California Department of Motor Vehicles (DMV) also

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20. One notable aspect of the calculated contributions to the levels of net migration is that they all drop at least 60,000 persons over the sample period. For the relative house price differential contribution series, the swing is almost entirely due to the evolving origin state population levels. To get a more specific calculation of the effect of evolving origin state population levels, we computed the difference between the fit of the model for the level of California net migration when dynamic population weights are used in the aggregation of implied migration rates and the fit when static (sample-mean) origin population weights are used in the aggregation. The difference swings about 60,000 persons over the sample period. To better explicate this phenomenon, we note that the states which historically have provided the largest net inflows to California—Ohio, Michigan, New York, and Illinois—had declining population shares over the sample period. Also, states that received the largest average levels of net out-migration from California—Washington, Nevada, Oregon, and Arizona—experienced more rapid population growth than the overall United States.

One possibility, consistent with this latter finding, is that there is a substantial reflective component to the California migration process. Historically, relatively young people from major urban centers in the Northeast and Midwest might have moved to California to take jobs or complete their education. Once they had relocated in California, their next moves might have been to other western states. The pool of potential in-migrants to California from the Northeast and Midwest has declined in relative size over time; the pool of people which may have less interest in migrating to California, as indicated by their residence in other western states, has increased in relative size over time.

21. The per capita wage and salary income estimates for 1994 are based on personal income data through the second quarter of 1994. Also, the

HUD series on housing prices was unavailable after 1993, so the 1994 estimates in this table extrapolate the HUD housing prices on the basis of changes in the median prices of single family homes in selected metropolitan areas, as reported by the National Association of Realtors and aggregated to the state level by the authors.

22. In this and other ex-post simulations of the model, the population age structures of the states are held constant at their 1992 values. Also, the 1992 figures for the remaining population in each state, M_{iit}, are used in the simulations to convert implied migration rates to implied migration levels.

offer a more current perspective on actual California migration and provide state-to-state detail not available in the Census data. During fiscal year 1993–1994, net driver's license address changes from California to other U.S. states tallied about 125,000, an acceleration relative to the roughly 100,000 drivers' license address changes recorded in the previous fiscal year. Using the California State Department of Finance's rule-of-thumb of 1.5 persons per driver's license, this is an acceleration of net domestic out-migration of about 35,000–40,000 persons, roughly consistent with the recent Census figures.[23]

On a state-by-state basis, the recent DMV data suggest that in the fiscal year ending in mid-1994, the relatively large net outflows of people from California to Arizona, Idaho, Oregon, and Washington continued at about the previous year's pace, and the net outflow to Nevada picked up a bit. This state-by-state pattern also is broadly consistent with the predictions of the model for the calendar years 1993 and 1994; for example, the Nevada unemployment rate dropped particularly sharply in 1994, so the model implies a pickup in migration from California to Nevada last year.

Figure 4 also presents simulations of California net migration flows under the assumption of a reversion sometime after 1994 of state level unemployment, wage, and house price differentials to the average levels observed during the 1981–1992 estimation period. These baseline simulations show the path of California net migration in the presence of a moderate rebound in California economic conditions and less rapid economic growth elsewhere. For example, the unemployment rate differential with respect to Idaho is assumed to narrow about 3 percentage points, from 3.19 percentage points in 1993–1994 (Table 2) to 0.12 percentage point in the baseline simula-

tion. Unemployment differentials with respect to other Twelfth District states and the rest of the United States are assumed to narrow by a similar order of magnitude.

In the case of the model "without structural change," the outflow from California is predicted to slow to about 78,000 persons per year. In other words, about 70 percent of the 1994–1995 estimated outflow of 258,000 persons is attributable to deviations of current economic conditions from mean values. Allowing for a one-time structural change only reduces the predicted slowing by about 25,000 persons per year, to a pace of about 103,000 persons per year. Thus, if the moderate rebound in California economic conditions is assumed to occur, but the earlier structural shift in amenity values also is assumed to persist, out-migration from California is predicted to be substantial, but much less than the recent pace.

On a state-by-state basis, the baseline simulation of the model without structural change (last row of Table 2) suggests that the small net outflows to Utah will be reversed, and the net outflows to most other District states will slow substantially. For example, at the 1981–1992 sample mean for the economic variables, implied net migration from California to Washington slows to 10,650 persons from an estimated 29,480 persons per year in 1993-1994. For most other Twelfth District states the model also implies that about two-thirds of the current net out-migration from California was due to deviations of economic conditions from mean values.

Figure 4 shows additional California net migration scenarios. Relative to the baselines, one of these scenarios is more optimistic for California and the other is more pessimistic. In the optimistic scenario, the California unemployment rate drops by an additional 1 percentage point and the relative wage increases by an additional 5 percent of the national wage; all other characteristics of the baseline simulation are maintained. As is evident from the figure, such a relatively moderate resurgence in California labor demand, coupled with long-run stability of estimated state-level fixed effects, implies slightly positive net domestic in-migration to California.

In the pessimistic scenario, not only is some earlier deterioration in the amenity value of residing in California admitted, but some further deterioration is assumed to occur.[24] Moreover, no additional narrowing of unemployment rate differentials or improvement in California relative wages is allowed; the economic conditions are held fixed at their estimated 1994 values. This more pessimistic

23. However, the levels of domestic migration implied by the driver's license address change data, using the 1.5 persons per driver's license rule-of-thumb, are not always consistent with the levels of migration implied by the IRS data, and changes in the two series also are inconsistent in some years. There are a number of reasons why it would be imprudent to take the DMV-based estimates literally. First, there are quite understandable reasons why the ratio of people to driver's licenses might evolve over time, such as changes in the age structure of the population and changes in the proportion of the population that can afford automobiles. Second, there are a few clear time-series breaks in the ratio of migration flows implied by the IRS data to the migration flows implied by the DMV data (using the 1.5 persons rule-of-thumb); one of these breaks coincides with a switch across state agencies in responsibilities for tabulating and processing the driver's license data. Last, the California state agencies which have released the data urge users to be cautious about interpreting them, owing to factors such as lags in the reporting of address changes to state authorities and instances of nonreporting from destination states.

24. Specifically, over the estimation period the "model with structural change" implies that a swing in the desirability of residing in California—owing to factors other than the measured unemployment rate,

scenario implies that net out-migration from California increases by about 37,000 relative to the predicted fiscal year 1994–1995 experience; in the pessimistic scenario, about 314,000 people per year, on net, continue to leave the state.

IV. CONCLUSIONS

This paper evaluates the role of economic and nonpecuniary location-specific factors in the determination of California domestic migration flows over the 1981–1992 period and beyond. Our results indicate that changes over time in state unemployment rate differentials explain most of the changes in California net migration. In contrast, changes over time in relative state wage and house price differentials have only small effects on migration flows. With respect to location-specific attributes and amenities, we find evidence of a relatively strong secular draw to California as a migrant destination for the sample period as a whole. Moreover, we do not find strong evidence in support of the deterioration in California amenities or "quality of life," topics of numerous media commentaries. In particular, because the model fits well without such an allowance for structural change, the results suggest that only a relatively small portion of the recent net out-migration is due to structural change.

Simulations of the model suggest that California net out-migration would slow substantially in the presence of a moderate rebound in California economic conditions—specified as a reversion in state-level unemployment rate, wage, and house price differentials to the average levels observed during the 1981–1992 estimation period—from about 258,000 people in fiscal year 1994–1995 to about 78,000 people in the future. Indeed, an alternative simulation of the model suggests that a somewhat stronger resurgence in California labor demand, coupled with no deterioration in amenities, would result in slightly positive net in-migration to California. In summary, our research findings suggest that much of the recent domestic out-migration from California has been in response to poor economic conditions and would be largely reversed in the context of a rebound in the California economy.

One might be tempted to conclude from our findings that the alleged deterioration in California amenities is of limited importance in the determination of changes in population flows to and from the state, relative to changes in the availability of jobs. This conclusion is consistent

with our results, but we have not definitively proven the case for the limited importance of changes in amenities here, because we have not attempted to evaluate the extent to which recent job losses in the state were driven by deterioration in California amenities. An alternative possibility, which also is consistent with our findings, is that a deterioration in California amenities influenced business location decisions, with a resulting decline in job opportunities. A structural decomposition of the historical changes in the state's unemployment rate would be useful for assessing the plausibility of various near-term economic and demographic scenarios, but we do not attempt such an unemployment rate decomposition here. Whether or not a decline in amenities substantially contributed to the recent weakness in California labor demand, it is clear that efforts to create and retain jobs are critical to any improvement in the California domestic migration balance.

APPENDIX: ROBUSTNESS OF THE RESULTS

This appendix discusses the robustness of the results to alternative econometric specifications. Because we are primarily concerned with projected migration in the scenario of a rebound in California economic conditions, we focus here on the sensitivity of the point estimates of the coefficients on the driving economic variables.

One potential criticism of the way we estimated the baseline specification (Table 3) and described the results concerns the possible presence of correlation between the error term and right-hand side variables in the estimated equation (6). The OLS estimator we use provides consistent estimates of the parameters of the equation when that equation is interpreted as a conditional expectation relationship, where the error term is orthogonal to the regressors by definition; this is the sense in which we use the model in the projection exercise. However, if the equation and error term are given a structural interpretation, a correlation between the error term and the included regressors could lead to inconsistency in the OLS estimator for the structural parameters. To check on the plausibility of the structural interpretation of the OLS estimates, we re-estimated the "without structural change" version of the model by instrumental variables (IV), using lagged values of the unemployment rate differential, wage differential, and housing price differential as instruments for the contemporaneous values of these three variables. The basic pattern in the point estimates of the coefficients on these key variables is invariant to this choice of estimator. Specifically, the coefficients on the relative wage and housing price differentials remain near 0.35 and 0.10, respectively, when IV is used. The coefficient on the unemployment rate differential increases somewhat in absolute

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wage, and house price differentials—induced an increase of about 37,000 people per year in net out-migration from California. In the pessimistic scenario, an equal-sized structural change is assumed to recur, boosting the effect to 74,000 people per year.

value, from –.053 under OLS to –.086 under IV. Thus, the unemployment rate differential remains the most important contributor to the explanation of changes in net migration under this alternative estimator.

A second issue concerns the parsimonious way we have parameterized the fixed effects and whether the IV estimator described above is consistent for structural parameters in the presence of fixed effects. We have implemented the fixed effects estimators with dummy variables, which is numerically equivalent to estimating slope coefficients with variables transformed to deviations from mean form. The corresponding instrumental variables estimators of slope coefficients in fixed effects models generally are not consistent for structural parameters if the instruments are merely pre-determined, but not strictly exogenous, as discussed in Keane and Runkle (1992). In our application of IV above, the instruments fail this strict exogeneity requirement. However, IV with predetermined instruments provides consistent estimates under some alternative transformations that remove the fixed effects, including the simple transformation of taking the first differences of the data over the time dimension. Doing so implicitly allows for one fixed effect for each origin-destination pair, as in Frees (1992, 1993), whereas our equation (6) only allows for one fixed effect for each state as an origin and one fixed effect for each state as a destination. We have re-estimated the model in first-difference form, expressing each origin-destination-specific (i,j) variable as a deviation from its lagged value. IV estimation of the difference specification, using second and third lags of the economic variables as instruments, yields a coefficient of –.062 on (changes in the) unemployment rate differential, still somewhat close to the OLS estimate of this coefficient from the levels specification, reported in Table 3. With the difference specification, the coefficient estimates on the other variables do not change enough to alter our conclusion that the unemployment rate differential is the most important contributor to the explanation of changes in net migration.

We also evaluated the sensitivity of the model results to alternative dynamic specifications. To investigate possible delays in the response of migration to changed economic conditions, we re-specified equation (6) to include lagged (prior year) values of the economic variables instead of current year values. In this case, the OLS estimate of the coefficient on the unemployment rate differential is –0.061, about the same as in the original specification. If, instead, both current and lagged fundamentals are used, then the sum of the coefficients on the unemployment rate differential is –0.067, with a roughly 50–50 split between the contributions.of the current and lagged values. Our projection exercise returns all variables to their sample means at some unspecified date in the future, and we do

not focus on the implied timing of the slowing of net out-migration from California. These additional results on the dynamics suggest that the timing of the response to a narrowed unemployment rate differential would not be fully contemporaneous.

A final econometric issue concerns the presence of heteroskedasticity in the errors of equation (6). In the residuals from the estimation results shown in Table 3, there appears to be some heteroskedasticity as a function of the mean migration rate, as suggested by Frees (1992). To check this, we computed the standard deviation of the residuals for each origin-destination pair and then regressed this dependent variable on the mean logarithmic migration rate for that origin-destination pair. The slope coefficient in this second-stage regression is –.024, which shows that state pairs with higher directional migration rates tend to have less unexplained volatility in this flow, in percentage terms. However, the R^2 in this second-stage regression is only .16, which suggests that only a moderate amount of heteroskedasticity could be explained this way. This feature of the error term affects the efficiency of the estimators and the consistency of estimators of standard errors of coefficients. However, it does not affect the consistency of the coefficient estimators themselves, and we leave the computation of more efficient estimators for future work.

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