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ABSTRACT

This paper investigates if the location choices made by immigrants when they arrive in the United States are influenced by the interstate dispersion in welfare benefits. Income-maximizing behavior implies that foreign-born welfare recipients, unlike their native-born counterparts, may be clustered in the states that offer the highest benefits. The empirical analysis indicates that immigrant welfare recipients are indeed more heavily clustered in high-benefit states than the immigrants who do not receive welfare, or than natives. As a result, the welfare participation rate of immigrants is much more sensitive to changes in welfare benefits than that of natives.

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I. Introduction

There is widespread concern that the resurgence of immigration in the United States has had an adverse impact on the cost of maintaining the many programs that make up the welfare state. This anxiety played a major role in the recent debate over welfare reform and, in fact, key provisions in the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 deny non-citizens the right to receive most types of public assistance.

The debate over the link between immigration and welfare focuses on two related issues. The first is the perception that there has been a rapid rise in the number of immigrants who receive public assistance. Although early studies of immigrant participation in welfare programs concluded that immigrant households had a lower probability of receiving public assistance than U.S.-born households, more recent studies have shown that this conclusion no longer holds—immigrant households are now more likely to receive welfare than native households.[1] Borjas and Hilton (1996) report that when one includes both cash and non-cash benefits (such as Medicaid and Food Stamps) in the definition of welfare, nearly 21 percent of immigrant households received some type of assistance in the early 1990s, as compared to only 14 percent of native households. The increasing participation of immigrants in welfare programs has

[1] See, for example, Blau (1984), Tienda and Jensen (1986), and Borjas and Trejo (1991).

spawned a rapidly growing literature that attempts to determine if immigrants "pay their way" in the welfare state.[2]

There is also some concern over the possibility that the generous welfare programs offered by many U.S. states have become a "magnet" for immigrants. The magnet hypothesis has several facets. It is possible, for example, that welfare programs attract immigrants who otherwise would not have migrated to the United States; or that the safety net discourages immigrants who "fail" in the United States from returning to their source countries; or that the huge interstate dispersion in welfare benefits affects the residential location choices of immigrants in the United States and places a heavy fiscal burden on relatively generous states. Despite their potential importance, there has been little systematic study of these magnetic effects, and there is little empirical evidence that either supports or refutes the conjecture that welfare programs have affected the size, composition, or geographic location of the immigrant flow.[3]

This paper begins to document the link between immigrant welfare use and some of the potential magnetic effects of welfare benefits. In particular, I investigate whether the residential choices made by immigrants in the United States are influenced by the interstate dispersion in benefits. It turns out that these magnetic effects can lead to striking and easily observable

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[2] See, for example, Huddle (1993) and Passel and Clark (1993). Smith and Edmonston (1997, Chapters 6 and 7) present a very careful accounting of the fiscal impact of immigration both in the short run and in the long run.

[3] The recent work of Olsen and Reagan (1996) investigates the out-migration decision of foreign-born persons surveyed by the National Longitudinal Surveys of Youth, and finds that young immigrants who receive welfare are less likely to leave the United States. Blank (1988), Gramlich and Laren (1988), Meyer (1998) and Walker (1994) analyze the impact of welfare programs on location decisions for the entire population, but reach somewhat conflicting conclusions. The Blank, Gramlich-Laren, and Meyer studies report evidence that women eligible for welfare are less likely to migrate out of (or more likely to migrate into) states with high benefit levels, while Walker does not find any evidence that low-income households migrate in search of higher welfare benefits.

outcomes as long as immigration is motivated by income-maximizing behavior. In particular, foreign-born welfare recipients, *unlike* native welfare recipients, should be clustered in the state that offers the highest benefits. As a result of this geographic clustering, the sensitivity of welfare participation rates to differences in state benefit levels should be greater in the immigrant population than in the native population.

The empirical analysis presented in this paper uses the 1980 and 1990 Public Use Microdata Samples (PUMS) of the decennial Census to test the theoretical implications. The data reveal a great deal of dispersion in the welfare participation rate of immigrants across states, and indicate that less-skilled immigrants—and, more specifically, immigrant welfare recipients—are much more heavily clustered in high-benefit states than immigrants who do not receive welfare, or than natives. The evidence, therefore, is consistent with the hypothesis that the generous welfare benefits offered by some states have magnetic effects and alter the geographic sorting of immigrants in the United States.

II. Theory

The intuition underlying the hypothesis developed in this paper is easy to explain. Persons born in the United States and living in a particular state often find it difficult (i.e., expensive) to move across states. Suppose that migration costs are, for the most part, fixed costs—*and* that these fixed costs are relatively high. The existing differences in welfare benefits across states, therefore, may not motivate large numbers of natives to move because the interstate benefit differentials might be swamped by the migration costs. In contrast, immigrants arriving in the United States are a self-selected sample of persons who have chosen to bear the fixed costs of the geographic move. Suppose that once the costs of moving to the United States are incurred, it costs little to choose one particular state over another. The sample of newly arrived immigrants will then tend to live in the "right" state.

Income-maximizing behavior on the part of immigrants and natives thus generates two interesting and empirically testable propositions. First, while welfare recipients in the native population are "stuck" in the state where they were born, welfare recipients among new immigrants should be clustered in the states that offer the highest welfare benefits. Second, the probability that a newly arrived immigrant receives welfare should exhibit "excess sensitivity" (relative to natives) to the level of welfare benefits.

To develop these ideas formally, consider first how natives in the United States decide where they wish to reside.[4] Suppose there are two states in the country, states 1 and 2. Initially, the native population is randomly distributed across states. The relationship between log wages and skills in state j is given by:

$$\text{(1)} \qquad \log w_j = \mu_j + \eta_j v,$$

where w_j gives the worker's earnings in state j; μ_j gives the mean log earnings in the state; and the random variable v measures deviations from mean log earnings and has finite variance.[5] It is useful to interpret v as a measure of relative ability or skills that are perfectly transferable across regions, so that the parameter η_j gives the rate of return to skills in state j. Without loss of

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[4] The theoretical framework builds on the multi-region extension of the Roy model developed by Borjas, Bronars, and Trejo (1992).

[5] A more general model would derive the wage-skill curves in equation (1) from a more primitive framework that takes into account interstate differences in natural resources, amenities, and other forms of physical

generality, rank the states so that $\eta_2 > \eta_1$. Finally, assume that natives (and immigrants) are income-maximizers.

I introduce the welfare program offered by each state in a very simple fashion. Each state guarantees a minimum level of log income $\bar{w}_j$ to *all* its residents, regardless of whether the person was born in the state. For simplicity, I assume that $\bar{w}_j$ is exogenously determined, and I ignore the issues related to the funding of the welfare program.[6]

Suppose initially that it is costless to move from one state to the other. Panel (*a*) of Figure 1 illustrates the optimal geographic sorting that occurs when state 1 offers a more generous program, or $\bar{w}_1 > \bar{w}_2$. The figure indicates that all persons who have skills below v_A choose to enter the welfare system offered by state 1; all those with skills between v_A and v_B chose to work in state 1; and all those with skills exceeding v_B end up working in state 2.[7] Note that the assumption of perfect mobility generates a clustering of the least skilled workers in the state that offers the highest welfare benefits (state 1). Panel (*b*) shows that a similar clustering, but in a different state, occurs when $\bar{w}_2 > \bar{w}_1$. All persons below the threshold skill level v_A then choose to enter the welfare system in state 2. The key implication of the analysis when there is perfect mobility is clear: Native welfare recipients will be clustered in only one state, the state that offers the highest welfare benefits.

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capital. This extension is important because it would help us understand why regions differ in the wage offers they make to otherwise identical workers.

[6] The exogeneity assumption is invalid in the long run. A state's welfare policy is probably very sensitive to the in-migration of potential welfare recipients either from other states or from abroad. The empirical analysis reported below uses data from the 1980-1990 period, a period of relative stability in the "fundamentals" of welfare policy.

[7] The most skilled workers move to state 2 because that state offers the highest rate of return to skills. Borjas (1987) presents a detailed discussion of how regional differences in the rate of return to skills determine the type of selection that characterizes the migrant sample.

Let's now suppose that it is costly to move across states, and that these migration costs are relatively high. Define the migration costs π as a fraction of a person's income in his or her state of birth. In particular, a person born in state j finds that it costs πw_j dollars to move to state k.[8] For simplicity, I assume that the time-equivalent measure of migration costs π is fixed across persons.

Panel (*a*) of Figure 2 illustrates how persons born in state 1 sort themselves across states. To illustrate the main insight, I assume that the alternative destination offers higher welfare benefits than the state of birth. Migration costs effectively raise the wage-skill curve in state 1. As a result, the persons who would have potentially migrated to a "welfare magnet" now stay put in their native state. As drawn, the migration costs are not sufficiently high to deter all migration, but they do deter the migration of all potential welfare recipients. Panel (*b*) illustrates the sorting of persons born in state 2. Before migration costs come into play, the migrant flow leaving state 2 is negatively selected (i.e., the migrant flow is composed of the least skilled persons). As drawn, the high level of migration costs stops the internal migration flow altogether. For sufficiently high migration costs, therefore, the least-skilled persons in the population will collect welfare benefits in the state where they were born.

We can now apply this framework to immigrants who originate in source country 0. Persons residing in this country also face the log wage distributions given in (1) for each state. Potential immigrants, however, have an additional option, the wage offer made by the country of origin. The source country's log wage distribution is given by:

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[8] Consider a person born in state 1. Migration occurs when $w_2 - C > w_1$, where C gives the migration costs

(2) $$\log w_0 = \mu_0 + \eta_0 v,$$

where μ_0 is the source country's mean log income; and η_0 is the source country's rate of return to skills. For simplicity, I assume that the source country does not offer a welfare system.

The potential immigrant knows that it is costly to move from the source country to the United States. As before, I represent the migration costs as a fraction of the income available to the workers in the source country. It costs πw_0 dollars to move to the United States, and π is fixed across persons. I also assume that the potential migrant does not incur any additional costs in choosing state *j* over state *k*. In other words, it is equally costly to move to any region of the United States.

Figure 3 illustrates how the population of persons born in the source country will be sorted geographically. Note that the fixed migration costs do *not* shift the wage-skills curves offered by states 1 and 2, but do raise (by a constant π) the wage-skills curve for the source country. Panels (*a*) and (*b*) of the figure illustrate the case where the rate of return to skills in the source country is higher than that available in the United States ($\eta_0 > \eta_2 > \eta_1$), but differ in their assumption about which state offers the highest welfare benefits. Panels (*c*) and (*d*) illustrate the case where the rate of return to skills is lower in the source country ($\eta_2 > \eta_1 > \eta_0$), and again differ on their assumption of which state offers the highest benefits.

The main insight of the model is easily grasped by working through a couple of the possible sorting equilibria. Suppose that the rate of return to skills in the source country is higher than in the United States, and that state 1 offers more generous welfare benefits than state 2 (the

(in dollars). An approximately equivalent condition is that $\log w_2 > \log w_1 + \pi$, where $\pi = C/w_1$.

case illustrated in panel (*a*) of Figure 3). Because migration costs shift the wage-skills curve *only* for the source country, the presence of these costs does not change the relative position of the wage-skills curves offered by states 1 and 2. The figure shows that persons who have skills below v_A move to state 1 and receive welfare; persons with skills between v_A and v_B also move to state 1, but enter the labor force; persons with skills between v_B and v_C move to state 2, and work there; and persons with skills exceeding v_C remain in the country of origin.[9]

The analysis, therefore, indicates that welfare recipients in the immigrant population are clustered in the state that offers the highest welfare benefits—even where there are fixed costs of migration. In other words, the geographic sorting of immigrants across the United States looks qualitatively similar to the geographic sorting that would have been observed among natives *in the absence of migration costs*.[10]

Consider now the case where the source country has a low rate of return to skills, and state 2 offers more generous welfare benefits (see panel (*d*) of Figure 3). All persons with skills below v_A move to state 2 and enter the welfare system; those with skills between v_A and v_B stay in the source country; those with skills between v_B and v_C move to state 1 and work; while the most skilled persons move to state 2 and work there. Again, the optimal sorting of persons across states suggests that welfare recipients should be clustered in the state that offers the highest welfare benefits.

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[9] Panels (*a*) and (*b*) of Figure 3 indicate that if countries that offer a relatively high rate of return to skills export *any* immigrants to the United States, they will certainly have to export some welfare recipients. Since many immigrants do, in fact, originate in such countries, the fixed costs of migration probably do not swamp the income differential between the source country and the welfare programs offered by some of the states. This fact is crucial for understanding why there are some welfare recipients in the immigrant population, but there may be no welfare recipients in the respective subsample of native migrants.

[10] Compare, for example, panel (*a*) in Figure 3 with panel (*a*) in Figure 1.

In sum, the income-maximization hypothesis, combined with the assumption that there are relatively high fixed costs of migration, generates a very strong theoretical prediction. The sample of immigrant welfare recipients will be clustered in the state that offers the highest welfare benefits, while the sample of native welfare recipients will be much more dispersed across the states. In effect, the "magnetic" effects of welfare lead to a different geographic sorting of immigrant and native welfare recipients.[11]

This equilibrium sorting implies that the welfare participation rate of immigrants should be much more sensitive to interstate differences in benefits than the welfare participation rate of natives. As a result, the "benefit elasticity"—i.e., the change in the welfare participation rate induced by a given percentage change in the benefit level—should be greater in the immigrant population than in the native population. The empirical analysis presented below will attempt to determine if the data on immigrant welfare participation exhibit both the clustering effect as well as the excess sensitivity of welfare participation rates to benefit levels.

These strong predictions are obviously derived by making very strong assumptions. The model ignores many other factors that determine location decisions. For example, the residential segregation of ethnic groups in a small number of states promotes the formation of ethnic networks that provide information about labor market opportunities *and* welfare benefits to potential migrants in the source countries.[12] These informational flows effectively reduce the

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[11] The model also illustrates the existence of a different type of magnetic effect: some persons who would not have migrated in the absence of welfare programs will now choose to move to the United States. In panels (*c*) and (*d*) of Figure 3, the immigrants who become welfare recipients would have stayed in the source country if the welfare programs were not available. Borjas and Trejo (1993) present a more detailed discussion of how welfare programs alter the type of selection that characterizes the immigrant population.

[12] Borjas and Hilton (1996) conclude that there is some transmission of information about welfare programs from the older immigrant waves to new arrivals *within* an ethnic group.

costs of migrating to specific states for particular ethnic groups, and might lead to a different geographic sorting than the one predicted by the model. Any empirical analysis of the magnetic effects induced by interstate differences in welfare benefits, therefore, must take into account the information networks that might exist within ethnic groups.

III. Geographic Clustering

The empirical analysis uses data drawn from the 1980 and 1990 PUMS. The household is the unit of observation. A household is classified as an immigrant household if the household head was born outside the United States and is either an alien or a naturalized citizen. All other households are classified as native households. The immigrant extract drawn from each census consists of a 5 percent random sample from the population, while each native extract consists of a .5 percent random sample. The empirical analysis is restricted to households that do not reside in group quarters and are headed by persons who are at least 18 years of age.

I classify a household as receiving public assistance if any member of the household received public assistance income in the calendar year prior to the Census. The cash benefit programs for which the Census reports public assistance income include Aid to Families with Dependent Children (AFDC), Supplemental Security Income (SSI), and general assistance. The data do not contain any information on the household's participation in non-cash programs, such as Food Stamps and Medicaid.

Table 1 reports the welfare participation rates (i.e., the percent of households receiving cash benefits) for the various states and the District of Columbia. Overall, the fraction of immigrants who received public assistance rose between 1980 and 1990—at the same time that the fraction of natives who received assistance declined. In 1990, 9.1 percent of immigrant households received benefits, as compared to 7.4 percent of native households.

There is a lot of interstate variation in both the direction and magnitude of the welfare gap between immigrants and natives. The data for California, the state with the largest immigrant population and some of the highest benefits, mirror the national results. The fraction of native households that received public assistance declined from 9.1 to 8.6 percent during the decade, while the fraction of immigrant households receiving assistance rose from 10.9 to 12.0 percent. Similarly, in New York, the state with the second largest immigrant population, the native welfare participation rate fell from 9.7 to 8.6 percent, while the immigrant participation rate rose from 9.3 to 10.0 percent.[13]

Other immigrant-receiving states experienced different trends. In Texas, the welfare participation rate of natives rose from 5.8 to 6.4 percent, while that of immigrants fell from 10.8 to 10.0 percent. In Florida, the welfare participation rate of natives fell from 5.9 to 5.5 percent, but that of immigrants fell even faster, from 10.2 to 8.5 percent.

The theoretical analysis presented earlier suggests that immigrant welfare recipients should flock to the state that offers the highest level of welfare benefits—much more so than native welfare recipients. It is well known that there is a great deal of dispersion in AFDC benefit levels across states.[14] Figure 4 summarizes the trends in AFDC benefits between 1970

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[13] In 1990, 28.5 percent of immigrant households lived in California, 15.6 percent in New York, 9.2 percent in Florida, 7.5 percent in Texas, 5.2 percent in New Jersey, and 5.1 percent in Illinois.

[14] In principle, what matters is the interstate variation in AFDC benefit levels after adjusting for cost-of-living differences across states. Unfortunately, federal statistical agencies do not report state-specific cost-of-living indices. The empirical analysis reported below partially controls for this problem by including state-specific fixed effects in the regressions, so that the variation that identifies many of the key parameters is generated by within-state changes in welfare benefits.

and 1990 for the main immigrant receiving states—relative to the benefits provided by the median state. In 1970, California was the median state. In 1990, California's benefits were almost twice as high as those offered by the median state. In fact, by 1990, California's AFDC benefit package was (almost) the most generous in the nation: 20 percent larger than New York's, 89 percent larger than the one in Illinois, and almost 280 percent greater than that offered by Texas.[15]

The theory thus suggests that we should find a clustering of immigrant welfare recipients in a small number of states—particularly in California. Table 2 shows that immigrants on welfare do indeed cluster in California, and that this clustering became more pronounced as California's benefit level rose relative to that of other states. In particular, the table reports the fraction of households (by welfare recipiency status) that lived in California. In 1990, California was home to 9.6 percent of the natives who do not receive welfare and to 11.5 percent of the natives who do. At the same time, California is home to 27.6 percent of the immigrant households that do not receive welfare and 37.6 percent of the immigrant households that do. The "difference-in-difference" estimator suggests that there may indeed exist a purposive clustering of less-skilled immigrants in California. Moreover, the comparison of the 1980 and 1990 data suggest that the clustering of immigrant recipients in California became more pronounced during the 1980s.[16]

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[15] Only Alaska offered more generous AFDC benefits in 1990.

[16] Beginning in 1980, a "deeming" requirement was instituted for some of the immigrants who applied for Supplemental Security Income (SSI). The income of the immigrants' sponsors was "deemed" to be part of the immigrant's resources during the immigrant's first three years in the United States, reducing the chances that newly arrived immigrants could qualify for SSI. It is unlikely that the increased clustering of immigrants in California can be attributed to the partial introduction of *nationwide* deeming requirements.

Table 2 also shows that there is more clustering when the data are restricted to immigrant households that arrived in the United States in the five-year period prior to the Census.[17] The theoretical discussion suggests that the clustering decisions made by new immigrants yield the "best" estimates of the magnetic effect of welfare benefits. The geographic sorting observed in the sample of recent arrivals reflects the impact of economic conditions at the time of migration, including the existing interstate differences in welfare benefits. Over time, the states change their welfare benefits and neither the immigrants who live in the United States nor natives can fully respond to such changes. Table 2 indicates that the geographic clustering of welfare recipients in California is much more pronounced for recent immigrants than for earlier arrivals. In particular, 45.4 percent of new welfare recipients live in California, as compared to only 28.9 percent of those who do not receive welfare.

The data also show that the demographic group most closely linked with the AFDC program—namely, female-headed households with children under 18 years of age—also exhibits the same type of geographic clustering. In particular, 9.6 percent of native female-headed households that receive welfare live in California, as compared to 11.1 percent of the native households that do not. Among new immigrants, 41.0 percent of the female-headed households that received welfare lived in California, as compared to 31.3 percent of the ones that do not.

There are a number of issues that might affect the interpretation of the clustering evidence. For instance, California is home to a large refugee population. These refugees (mainly from Southeast Asia) have very high welfare participation rates. One could argue that their choice of California as a final destination has little to do with purposive clustering, but might

[17] The year of immigration of the household is determined by the household head's year of arrival.

instead be attributed to political decisions, the location (and geographic networks) of the charitable agencies that sponsor their entry, and a host of other factors. Table 2, however, shows that refugee groups alone cannot account for the clustering effect of recent immigrant welfare recipients into California. Although the Census does not provide information on the type of visa used by an immigrant to enter the United States, we can reasonably classify all immigrants who originate in the main refugee-sending countries as refugees.[18] In 1990, 43.7 percent of the newly arrived non-refugees who were welfare recipients lived in California, as compared to 30.0 percent of the non-refugees who did not receive welfare.

California is also the destination of a large number of Mexican immigrants. Their location decision is probably dominated by California's proximity to Mexico and by the extensive links that are introduced through family ties and ethnic networks (Massey and España, 1987). Table 2 shows that the clustering effect remains even if we focus on the non-Mexican population. Among recent arrivals, 44.2 percent of the non-Mexicans who receive welfare live in California, as compared to 23.6 percent of those who do not.

Finally, the clustering effect documented in Table 2 can be interpreted in a very different way. In particular, the higher welfare benefits offered by California sweep further into the distribution of reservation incomes of households already living in California, and thus attract a larger number of those households into the welfare system. As a result, the relatively higher welfare participation rate in California may have little to do with a behavioral clustering effect,

[18] Thirteen countries accounted for over 90 percent of the refugees admitted during the 1980s. These countries are Afghanistan, Bulgaria, Cambodia, Cuba, Czechoslovakia, Ethiopia, Hungary, Laos, Poland, Romania, Thailand, the former U.S.S.R., and Vietnam.

but is instead the mechanical result of "sweeping" further into the distribution of economic alternatives as welfare benefits increase.

It is important, therefore, to present the "difference-in-differences" calculation implicit in the discussion *after* controlling for the demographic and socioeconomic factors that determine a household's eligibility and propensity to participate in welfare programs. In particular, consider the descriptive regression model:

$$C_{it} = X_{it}\,\alpha_0 + \alpha_1\, I_{it} + \alpha_2\, B_{it} + \alpha_3\,(I_{it} \times B_{it}) + \varepsilon_{it}, \tag{3}$$

where C_{it} is a dummy variable indicating if household i lives in California in year t; X is a vector of socioeconomic characteristics; I_{it} is a dummy variable indicating if the household is foreign-born; and B_{jt} is a dummy variable indicating if the household receives welfare benefits. I use the linear probability model to estimate the regression in (3) separately for each Census year.

The coefficient α_3 provides the difference-in-differences estimator of the "clustering gap" in California between immigrant welfare recipients and non-welfare recipients *relative* to the same gap in the native population. Table 3 reports the estimated coefficient from regressions estimated in a variety of samples and using a number of different specifications. The unadjusted coefficients reported in the first column can, of course, be obtained by differencing the relevant probabilities in Table 2. Column 2 presents the adjusted difference-in-differences from regressions that include the educational attainment, gender, and age of the household head, the number of persons living in the household, the number of children (under age 18), the number of elderly persons (over age 65), and a vector of race dummies (black, Hispanic, or Asian). The adjusted coefficients show that the clustering gap in the immigrant welfare population is much larger than that observed among observationally equivalent native welfare recipients.

As I noted earlier, there are many factors—beyond those captured by economic opportunities and welfare benefits—that motivate some immigrant groups to cluster in certain states. The presence of ethnic enclaves implies that an immigrant group may have better information about conditions in a subset of the states. I capture these network effects by adding a vector of national origin fixed effects to the regression specification in equation (3).[19] These fixed effects capture the impact of factors that are specific to the national origin group in terms of their location decision in California (such as the share of the ethnic group that resides in that state). Column (3) of the table shows that inclusion of the national origin fixed effects has little impact on the estimated coefficients. There is "excess" clustering among immigrant welfare recipients—even when we consider immigrants from within specific national origin groups.

Although the regression coefficients reported in Table 3 help us describe the pattern of clustering present in the data, they should not be interpreted as structural parameters. To obtain a deeper understanding into why such clustering takes place, it is useful to investigate how California's immigrants differ from immigrants who choose to reside in other states. In particular, does the clustering arise partly because less-skilled immigrants are more likely to live in California? Consider the regression model:

$$(4) \qquad C_{it} = X_{it}\,\beta_0 + \beta_1\, I_{it} + \beta_2\,(I_{it} \times X_{it}) + n_{i\ell} + \varepsilon_{it},$$

[19] This vector includes 92 dummy variables, 91 national origin variables for immigrants (including an "other" category for those immigrants who do not belong to one of the 90 largest groups) plus a dummy variable indicating native status. The "other" immigrant group contains about 3 percent of the immigrant sample.

where $n_{i\ell}$ gives a vector of national origin fixed effects indicating if person i was born in country ℓ, and the coefficient β_2 estimates the differential impact of various socioeconomic characteristics on the probability that immigrants reside in California.

The regressions reported in Table 4 suggest that there is *relative* negative selection of immigrants into California—in the sense that less-educated immigrants are much more likely to live in California than less-educated natives. Note also that the regressions include a vector of national origin fixed effects, so that the negative selection in education is observed even within national origin groups. The evidence, however, is less conclusive about how immigrants self-select on the basis of other socioeconomic characteristics. Even though immigrants with larger households or with a large number of elderly members are relatively more likely to reside in California (and these characteristics are positively correlated with welfare recipiency), female-headed households or households with a larger numbers of children are relatively less likely to live there.[20]

Of course, it is the "average" selection that occurs over all of these socioeconomic characteristics that determine the extent of geographic clustering. And the data summarized in this section indicated that this selection process directs welfare-prone immigrant households to California, the state that offered some of the highest welfare benefits in the country in 1990.

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[20] The correlation between the presence of elderly persons in the household and the geographic clustering of immigrants suggests a promising avenue for further research: an examination of how interstate differences in SSI benefits affects the location decision of elderly immigrants. There is, however, one potential problem that would have to be addressed in an analysis of the magnetic effects of SSI. A large fraction of elderly immigrants might be tied movers; their working-age children (who probably respond to differences in labor market opportunities) may determine the residential location of the entire household.

IV. The Excess Sensitivity of the Benefit Elasticity

The clustering of immigrant welfare recipients into a very small number of states implies that the welfare participation of immigrant households should be more responsive to changes in state benefit levels than that of native households. If new immigrants were fully informed about alternative economic opportunities—*and* if one could control for all the other factors that influence the location decision—the immigrants who eventually end up as welfare recipients have an infinite supply elasticity to the state that offers the highest benefits. In contrast, an increase in benefits by state j does not attract natives from other states if migration costs are sufficiently high, but simply moves some of the natives already living in that state from the labor force to the welfare rolls. The benefit elasticity, therefore, should exhibit excess sensitivity in the immigrant sample.[21]

We can test this theoretical implication by pooling the 1980 and 1990 data and estimating the following regression model separately in the native and immigrant samples:

$$P_{ijt} = X_{ijt}\,\beta + \delta\,\overline{w}_{jt} + \alpha\,\overline{y}_{jt} + s_j + d_t + \varepsilon_{ijt}, \tag{5}$$

where P_{ijt} is a dummy variable that indicates if household i in state j receives cash benefits in Census year t; X is a vector of socioeconomic characteristics; $\overline{w}_{jt}$ is a measure of the welfare benefit in the state; $\overline{y}_{jt}$ is a vector of other state-specific variables which might influence welfare participation; s_j is a state fixed effect; and d_t is a period fixed effect. The coefficient δ measures

[21] Note that the labor supply adjustments induced by higher welfare benefits for persons already living within a state will also be observed in the immigrant sample—so that the parameter of interest is the *difference* in the benefit elasticities between the immigrant and native populations.

the benefit elasticity, the impact of *within-state* changes in benefit levels on the welfare participation rates of household in that state. I estimate equation (5) using the linear probability model. All standard errors are adjusted to account for the clustering of observations within a state.

As in the previous section, the vector X includes the education, gender, and age of the household head, the number of persons living in the household, the number of children (under age 18), the number of elderly persons (over age 65), and a vector of race dummies (black, Hispanic, or Asian). The regressions in the immigrant sample also include a vector of dummy variables indicating the year of migration.[22] The measure of welfare benefits $\bar{w}_{jt}$ gives the maximum AFDC monthly benefit (in logs) offered to a family of three living in state j (in either 1980 or 1990). Finally, the vector $\bar{y}$ includes two measures of economic activity in the state: the log of per-capita disposable income and the unemployment rate.[23]

Table 5 reports the relevant coefficients. If we restrict our attention to the sample of female-headed households with children, the regressions indicate that an increase in the state's welfare benefit level raises the welfare participation rate for both natives and immigrants.[24] The

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[22] These dummy variables indicate if the immigrant household migrated between 1985 and 1990, 1980 and 1984, 1975 and 1979, 1970 and 1974, 1965 and 1969, 1960 and 1964, 1950 and 1959, and prior to 1950.

[23] The state-specific data for welfare benefits are drawn from U.S. House of Representatives (1993); the other state-specific variables are drawn from U.S. Bureau of the Census (various issues). Note that the data on welfare benefits refers to the benefits offered as of 1980 or 1990. One could argue that these data should reflect the welfare offers made by the various states prior to the immigrants entering the United States (say as of 1975 or 1985). The timing of the welfare benefit data that should enter the regressions in the native sample, however, is much less obvious. I reestimated all the regressions reported in this paper using the 1975-1985 welfare benefits data. If anything, the results suggest a wider gap between the benefit elasticities of immigrants and natives. The benefit elasticity in the sample of newly arrived immigrant female-headed households is .31 (with a standard error of .09). The corresponding elasticity for natives is .03 (.02).

[24] Moffitt (1986) also presents evidence that changes in a state's AFDC benefit level increase the welfare participation rate of female-headed households.

size of the benefit elasticity, moreover, is about the same for the two groups—a doubling of benefits raises the welfare participation rate by about 10 percentage points. Note, however, that the benefit elasticity for immigrants increases substantially when the sample is restricted to recent immigrant arrivals. In this subsample of immigrants, a doubling of welfare benefits raises the probability of participation by 20.8 percentage points. The fact that the elasticity is much stronger for this group seems to confirm a key implication of the theoretical framework: Immigrants who have just arrived in the United States should exhibit excess sensitivity to interstate differences in welfare benefits.[25]

As noted earlier, a positive benefit elasticity in either the native or immigrant sample may be a "mechanical" result. Higher welfare benefits sweep further into the distribution of reservation incomes, and attract a larger number of those households *already in the state* into the welfare system. Although it might seem possible to isolate the behavioral effect of purposive clustering by controlling for the variables that determine welfare participation and by looking at the difference between the immigrant and the native benefit elasticities, the inclusion of the vector X in equation (5) does not solve the problem completely. After all, the possibility remains that an increase in benefit levels—even when measuring the impact on an immigrant who is observationally equivalent to a native—is integrating over different distributions of reservation incomes. Ideally, we would like to adjust the distributions so that the integration takes place over relatively similar areas. Obviously, it is difficult to control for these distributional

[25] I also estimated regressions that included an interaction between the AFDC benefit level and the educational attainment of the household head. This variable had a negative and significant impact, suggesting that the benefit elasticity is larger for the groups that we would expect to be most responsive to changes in state benefit levels (such as the less skilled). In the sample of newly arrived female-headed households, the coefficient of the interaction variable was -.004, with a standard error of .002. The benefit elasticity estimated at zero years of schooling was .271, with a standard error of .109.

differences (particularly those that are unobserved) unless one is willing to build in much more structure into the estimation procedure.

One relatively simple approach, and one that minimizes the amount of structure that is imposed on the data, is to equalize the distribution of *observable* socioeconomic characteristics in the two populations. Consider the following procedure.[26] First, classify the (newly arrived) immigrant population into groups that are defined by skill and other socioeconomic characteristics. Second, calculate the relative frequency of each of these groups in the immigrant population, and use these frequencies to reweigh the native sample so that the distribution of observable characteristics in the native sample is identical to that observed in the immigrant sample.

I defined cells according to the sex, education, race, and sex of the household head, the number of persons in the household, the number of persons under the age of 18, and the number of persons over the age of 65.[27] I calculated the relative frequency observed in each of these cells in the sample of newly arrived immigrants in the 1980 and 1990 Censuses, and then applied these weights to the native population. Table 5 shows that the re-weighting of the native sample has little impact on the point estimate of the benefit elasticity but increases the standard error substantially.

Although the benefit elasticity estimated in the sample of new immigrant arrivals is almost three times the size of that estimated in the native sample, the benefit elasticities tend to

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[26] This approach is similar to the weighting schemes used by Card and Sullivan (1988) and Imbens and Hellerstein (1996).

[27] There are 2 categories for sex, 5 for education, 6 for age, 4 for race, 5 for the number of persons in the household, 6 for the number of persons under the age of 18, and 6 for the number of persons over the age of 65.

have relatively large standard errors. As a result, the t-statistic for the test that these two elasticities are equal is about 1.1, regardless of whether the native sample is weighted to resemble the immigrant sample. Therefore, I cannot reject the hypothesis that the two elasticities are the same.

The difference between these the native and immigrant elasticities widens if the regression includes variables that measure the importance of ethnic enclaves in the state. I use two variables to measure these ethnic effects. The first (q_1) gives the fraction of the state's population that belongs to the same national origin group as the immigrant household, while the second (q_2) gives the fraction of the national origin group that lives in the particular state. These variables should help capture the importance of ethnic enclaves in terms of both the absolute and relative size of the group.[28]

The inclusion of these ethnic enclave variables raises the estimated elasticity in the sample of newly arrived female-headed households to about .23. As a result, the t-statistic for the test of the equality of this elasticity with that estimated in the native sample increases to 1.4 (when the native sample is not weighted) and to 1.3 (when the native sample is weighted).[29]

There are 43,200 potential cells, but most of these cells are empty. There are only 3,215 valid cells in the 1980 data, and 3,628 in 1990.

[28] The ethnic enclave variables are constructed from the 1980 and 1990 census files.

[29] The fraction of the state's population that belongs to the particular ethnic group has a negative impact on welfare participation, while the fraction of the ethnic group that lives in the state has a positive impact on welfare participation. An increase in the supply of a particular group of workers would presumably reduce the economic opportunities available to that group, and may even provide more information about welfare programs, so that one might expect to find a positive correlation between ethnic concentration and welfare propensities. At the same time, however, a larger ethnic concentration might generate "network effects" in job opportunities as well as reduce the economic penalty associated with not being proficient in the English language (McManus, 1990; Lazear, 1998). It would be of interest to develop a more detailed test of these alternative hypotheses.

It is well known that there are systematic differences in skills and welfare participation across national origin groups (Borjas, 1987; Borjas and Trejo, 1993). It is instructive to determine if the excess sensitivity of the benefit elasticity in the immigrant sample remains even after controlling for national origin. I reestimated the key regressions after including the vector of country-of-origin fixed effects defined in the previous section. As Table 5 shows, controlling for national origin increases the size of the immigrant benefit elasticity (to .27) and reduces its standard error. The *t*-statistic for the test that this elasticity is greater than the native benefit elasticity is 1.93 (when the native sample is not weighted) and 1.71 (when the native sample is weighted).[30]

The bottom panel of Table 5 reestimates the regression in the sample that contains all households (not just those that are female-headed with children). Within-state changes in AFDC benefits have no impact on welfare participation in the native population. They do, however, have an impact in the sample of newly arrived immigrants. If we control for the share of the household's ethnic group in the state, the data suggest that a doubling of AFDC benefits in the state raises the welfare probability by about 4 percentage points. The *t*-statistic for the test of equality between this elasticity and that found in the native sample lies around 1.4, regardless of whether the native sample is weighted.

The relatively large benefit elasticity found in the immigrant sample can be interpreted as a behavioral effect only if the difference-in-difference estimator has *completely* netted out the

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[30] I also estimated the regression models on the samples of immigrants who are not refugees or who are not Mexicans, thus removing from the analysis the key groups that tend to cluster in California and that might be "driving" the results. The benefit elasticity remains around .25 in the newly arrived female-headed population even among non-refugees or non-Mexicans.

"mechanical" impact of higher AFDC benefits on welfare propensities.[31] Although I standardized the native data so that the distribution of native reservation incomes is roughly similar to that found in the immigrant sample, there remain unobserved differences in income and economic opportunities between the two samples. An alternative approach is to calculate the mechanical elasticity directly by using the eligibility rules in each state's AFDC program (in both 1980 and 1990) to estimate the fraction of immigrants who are eligible to receive welfare benefits. The regression of these predicted probabilities on the within-state change in AFDC benefits gives the benefit elasticity that arises simply because higher benefits sweep further into the distribution of reservation incomes of persons already residing in a particular state.

In 1980, a household was eligible for AFDC if the household's earned income (minus deductions mainly for work expenses and childcare) was less than 150 percent of the state-set "pay standard." In 1990, a household was eligible if household income was less than 185 percent of the state-set "need standard" *and* if household income (minus deductions) was less than 150 percent of the pay standard.[32] Consider the sample of female-headed immigrant households who arrived between 1985 and 1990. I use the actual distribution of household income and household composition in this sample to calculate the fraction of households that would be eligible for

[31] Note, however, that the so-called mechanical effect induced by sweeping further over the distribution of reservation incomes of persons already residing in a particular state is itself a behavioral response (although not the one implied by geographic clustering). As the AFDC benefit level increases, a larger number of households find that the welfare benefits exceed their reservation price. The increase in the welfare recipiency rate arises because of this labor supply effect.

[32] A more detailed discussion of the eligibility rules is given in U.S. House of Representatives (1996, pp. 389-391). The income measures available in the data refer to the year prior to the Census. The eligibility rules described in the text, therefore, actually refer to 1979 and 1989. To avoid confusion, the discussion refers to data points in terms of the Census year. There is also an asset test for AFDC eligibility. Because of data constraints, I ignored this test in the simulation. Blank and Ruggles (1996) show that the asset requirement does not play a major role in determining eligibility. I am grateful to Aaron Yelowitz for providing the state-level data on need and pay standards.

AFDC in *each* state under the eligibility rules existing in both 1980 and 1990. Adjusting the eligibility probabilities by a state-specific "take-up" rate yields the predicted welfare recipiency rates. The take-up rates were calculated by comparing the welfare recipiency rate of households that actually live in a particular state to the eligibility rate of households in that state. This simulation exercise yields Δp_j, the 1980-1990 change in the predicted probability of welfare recipiency in state j attributable to changes in the benefit structure (and net of any purposive clustering). I then estimated the regression:

$$\Delta p_j = a + b\,\Delta \bar{w}_j + e_j, \tag{6}$$

where $\Delta \bar{w}_j$ gives the change in (log) AFDC benefits. The coefficient b captures the mechanical impact of higher AFDC benefits on the probability of welfare recipiency in state j. I also conducted the simulation and estimated the regression for the sample of native female-headed households.

One difficult conceptual problem arises in the calculation of eligibility rates. Some previous studies use actual household income to determine the household's eligibility (e.g., Blank and Ruggles, 1996). Actual income, however, incorporates the household's labor supply response to the presence of the welfare system.[33] The simulation requires the household income that would have been observed in the absence of the welfare system. I initially use actual household income to calculate eligibility rates. I also regressed household income on a vector of

[33] The use of lagged household income would not solve the problem because labor supply responses (if they exist) would probably occur before the household applies for AFDC benefits. Yelowitz (1995) and Currie and Gruber (1997) discuss alternative methods of avoiding the endogeneity problem.

socioeconomic characteristics in the subsample of households that did not receive welfare, and used this regression to predict household incomes for the households that did.[34] The results presented below are not sensitive to the alternative income measure used.

Table 6 reports the coefficients estimated from the regression model in (6). The mechanical benefit elasticity hovers around .11 or .12 in both the native and immigrant samples. As a result, the impact of changing AFDC benefits on the welfare recipiency rate of native households can be explained by the fact that higher benefits sweep more native households (already living in the state) into the welfare system. Therefore, there is no evidence that natives exhibit a migration response to interstate difference in welfare benefits. As we saw in Table 5, however, the immigrant benefit elasticity is on the order of .27, almost three times as large as the mechanical "distribution-sweeping" effect in the immigrant sample. Nevertheless, the hypothesis that the estimated benefit elasticity equals the mechanical elasticity cannot be rejected (the *t*-statistic is 1.46).

One can also use the simulation to calculate the fraction of welfare recipients who would be expected to live in California. Because California has high benefits, it should have a disproportionately large fraction of welfare recipients. The exercise requires that we specify how welfare recipients would be distributed across the country in the absence of interstate differences in benefits. I assume that this "null" distribution is equal to that of immigrant households that do not receive welfare.[35] As we saw earlier, 31.3 percent of the immigrant households that do not

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[34] The variables included in the regression are the education, gender, and age of the household head, the number of persons in the household, the number of persons under the age of 18, and the number of persons over the age of 65.

[35] It is possible that California's generous welfare system induces many less-skilled or risk-averse immigrants might choose to move there for insurance; they will enroll in the welfare system if the job opportunities

receive welfare live in California. Table 6 reports that if there were no behavioral response to California's high benefit levels (in the sense of purposive clustering), we would expect around 34 percent of the immigrant welfare recipients to live in California. In fact, 41 percent of the welfare recipients live in California. The distribution-sweeping effect, therefore, explains about a third of the excess clustering exhibited by immigrant welfare recipients. Moreover, the data reject the hypothesis that the fraction of welfare recipients who actually live in California equals the fraction predicted by the simulation (the t-statistic is 3.23).

Overall, the results suggest that the purposive clustering that occurs in the immigrant population does lead to excess sensitivity in the benefit elasticity. An additional test of the model's validity can be obtained by estimating equation (5) in a sample of *native* households that have moved across states. It would seem that this sample should be roughly equivalent to the immigrant sample—they are both self-selected and contain persons who found it worthwhile to incur the costs of moving.

Table 7 reports the benefit elasticities obtained when the model is estimated in the sample of native households that can be classified as "movers," in the sense that their state of residence changed in the five-year period prior to the Census. Although one might have expected to find sizable benefit elasticities, the regressions that do not reweigh the native sample show that there is no correlation between welfare participation rates and state benefit levels. It would seem, therefore, that the analysis of this migrant sample rejects a key implication of the model.

Note, however, that the benefit elasticity in the native mover sample increases significantly, to .26, when the native sample is weighted so that its demographic characteristics

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do not pan out. The fraction of non-welfare recipients who live in California may then not truly reflect the geographic sorting of immigrants that would have taken place if all states had offered identical welfare benefits.

resemble those of the recent immigrant population. This result, therefore, suggests that the benefit elasticity is roughly the same for movers (whether from abroad or from other states) when we equalize the skill composition of the two samples.

The importance of reweighting the native mover sample arises because native movers differ significantly from native stayers and from immigrants. The theoretical discussion showed that if migration costs are sufficiently large, few natives may move across states in search of higher welfare benefits. Instead, most of the migrants may be the workers who have the most to gain from geographic differences in labor market opportunities. In fact, the welfare participation rate for native movers in 1990 is 5.2 percent (24.3 percent for female-headed households). Native stayers had *higher* welfare participation rates: 7.6 percent in the entire sample, and 26.5 percent in the sample of female-headed households. Table 7 indicates that if one adjusts the distribution of skills in the native migrant sample (giving greater weight to the less skilled workers who make up the bulk of the immigrant population), an increase in benefit levels increases the welfare recipiency rate substantially.

Finally, Table 8 reports the benefit elasticities when equation (5) is estimated in the sample of households that live *outside* California. It turns out that the benefit elasticity estimated in the (female-headed) immigrant sample falls substantially when California is omitted from the analysis, while the elasticity estimated in the native sample is about the same. In particular, the immigrant elasticity is .15 (with a standard error of .13), and the native elasticity is .08. Table 8 also reports the sensitivity of the benefit elasticity to the omission of another major immigrant-receiving state, New York. If New York is omitted from the data, the immigrant elasticity is over .3, while the (unweighted) native elasticity remains at about .10. The evidence, therefore, indicates that it is not the omission of a major immigrant-receiving state that is driving the results; rather, it is the omission of *the* immigrant receiving state that has the highest welfare benefits.

It might seem disturbing that a single observation is driving the results of the study. After all, Table 8 shows that the excess sensitivity of the benefit elasticity in the immigrant sample is attributable almost entirely to immigrant households living in California. However, this is *precisely* what the theory predicts should happen. Welfare recipients in the immigrant population will cluster in the state that offers the highest benefits, and it is this clustering that creates a large positive correlation between welfare participation rates and state benefit levels. Removing that single state from the analysis should greatly weaken the correlation—and this is, in fact, what happens.

V. Summary

There are sizable differences in welfare benefits across states. If migration decisions are guided by income-maximizing behavior, these interstate differences in welfare benefits will lead to a very different geographic sorting of welfare recipients in the immigrant and native populations. Suppose, in particular, that all migrants—regardless of whether the move is internal within the United States or across international borders—incur relatively high fixed costs. These costs deter the migration of many potential native welfare recipients to states that offer higher benefits. As a result, native welfare recipients will be (more or less) randomly distributed over the United States. In contrast, immigrants are a self-selected sample of persons who have chosen to incur the fixed costs of migration. If the marginal cost of choosing the "right" state is small once the immigration decision is made, immigrant welfare recipients will cluster in the state that offers the highest benefits. The differential geographic clustering of welfare recipients between immigrants and natives also suggests that the correlation between welfare participation rates and welfare benefit levels should be larger among immigrants.

The empirical analysis used the 1980 and 1990 PUMS of the U.S. Census to test these theoretical implications. The data indicated that immigrant welfare recipients are much more likely to be geographically clustered than immigrants who do not receive welfare, and are also much more clustered than natives. In 1990, for example, 29 percent of newly arrived immigrants who did not receive welfare lived in California (a state that offered some of the highest welfare benefit levels). In contrast, 45 percent of newly arrived welfare recipients lived there. Much of this "clustering gap" arises because less-skilled immigrants are disproportionately drawn to California. The analysis also revealed that changes in a state's welfare benefits have a much larger impact on the welfare participation rate of immigrants than of natives.

The empirical evidence presented in this paper is consistent with the hypothesis that interstate differences in welfare benefits generate strong magnetic effects on the immigrant population. Because of the potential policy significance of these findings, it is important to emphasize that much of the empirical evidence presented in this paper is relatively weak (in the sense that the statistical significance of the results is often marginal). Moreover, there may well be alternative stories that explain the evidence. Nevertheless, the analysis does suggest that the wealth-maximization hypothesis generates a number of interesting and empirically testable implications of welfare magnets. The continued application of these theoretical insights to the study of magnetic effects may help resolve many of the unanswered questions about the behavioral and economic impact of the many programs that make up the welfare state.

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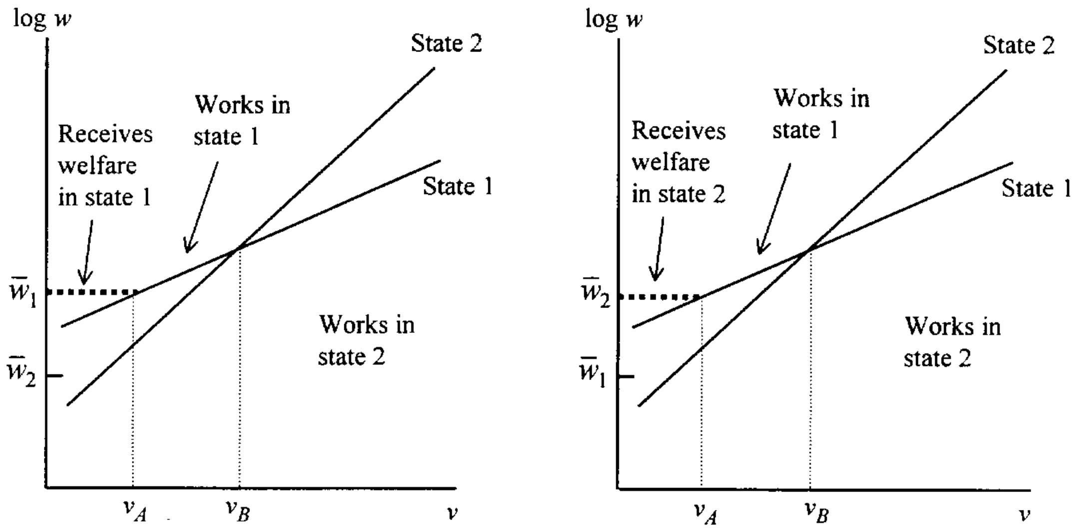
FIGURE 1

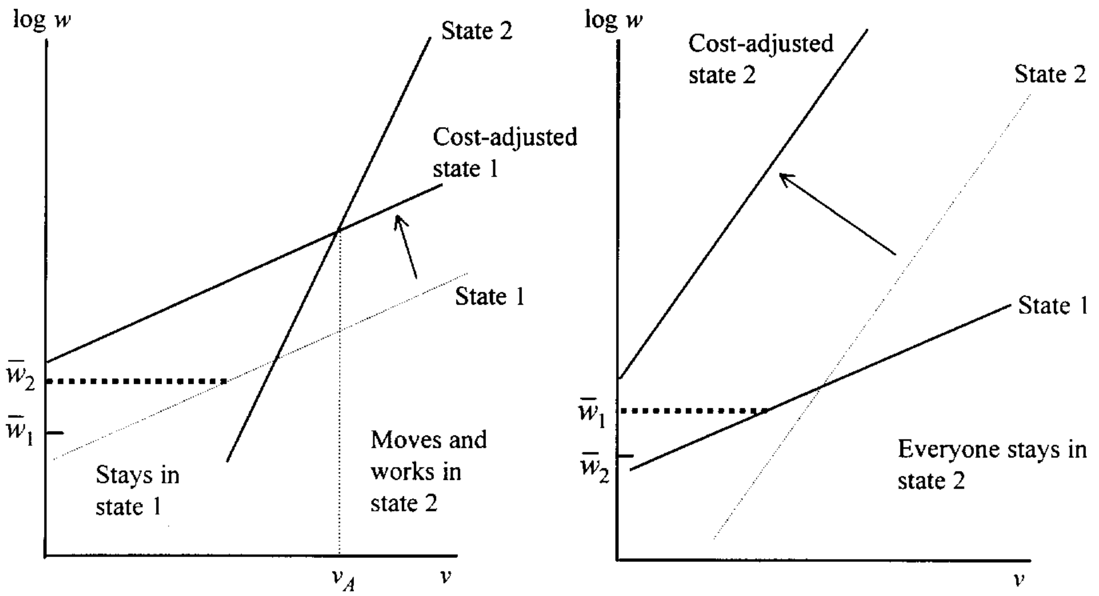
GEOGRAPHIC SORTING OF NATIVES WITH COSTLESS MIGRATION



(a) $\bar{w}_1 > \bar{w}_2$ (b) $\bar{w}_2 > \bar{w}_1$

FIGURE 2

GEOGRAPHIC SORTING OF NATIVES WITH FIXED MIGRATION COSTS

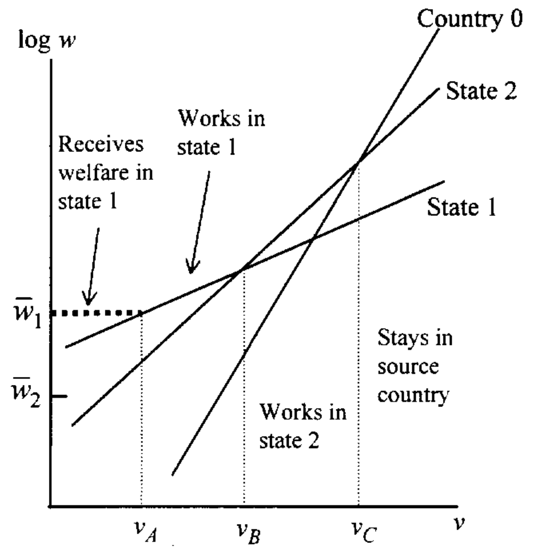


(a) Persons Born in State 1, and $\bar{w}_2 > \bar{w}_1$

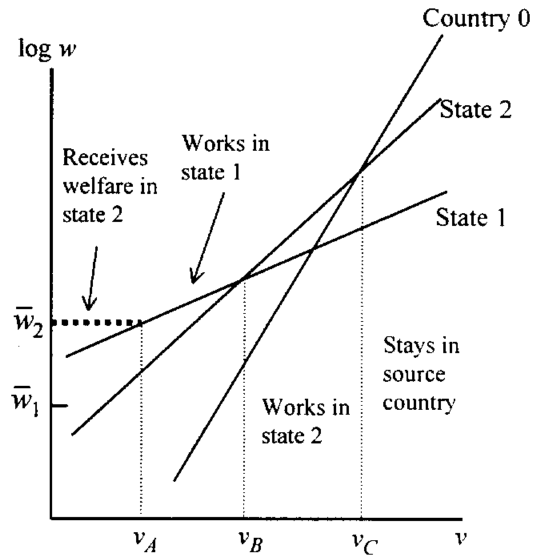
(b) Persons Born in State 2, and $\bar{w}_1 > \bar{w}_2$

FIGURE 3

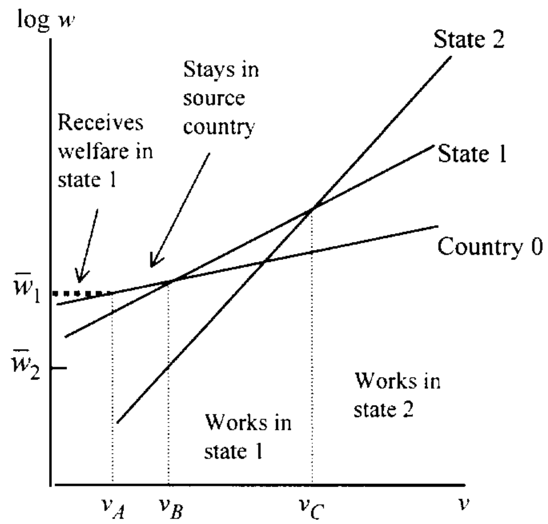
GEOGRAPHIC SORTING OF IMMIGRANTS WITH FIXED COSTS OF MIGRATION
(Wage-Skills Curve in Source Country is Adjusted for Migration Costs)



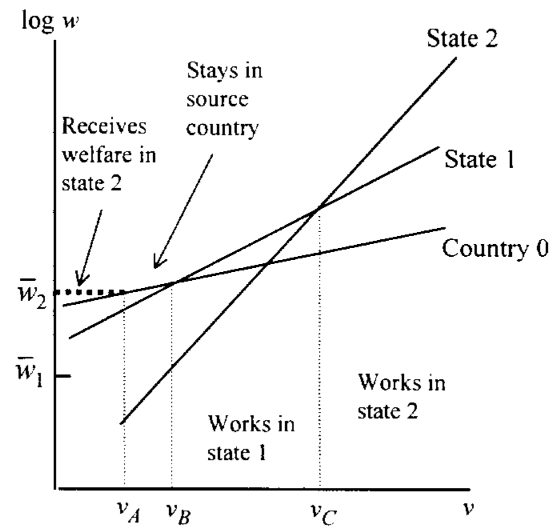
(a) $\eta_0 > \eta_2 > \eta_1$, and $\bar{w}_1 > \bar{w}_2$



(b) $\eta_0 > \eta_2 > \eta_1$, and $\bar{w}_2 > \bar{w}_1$

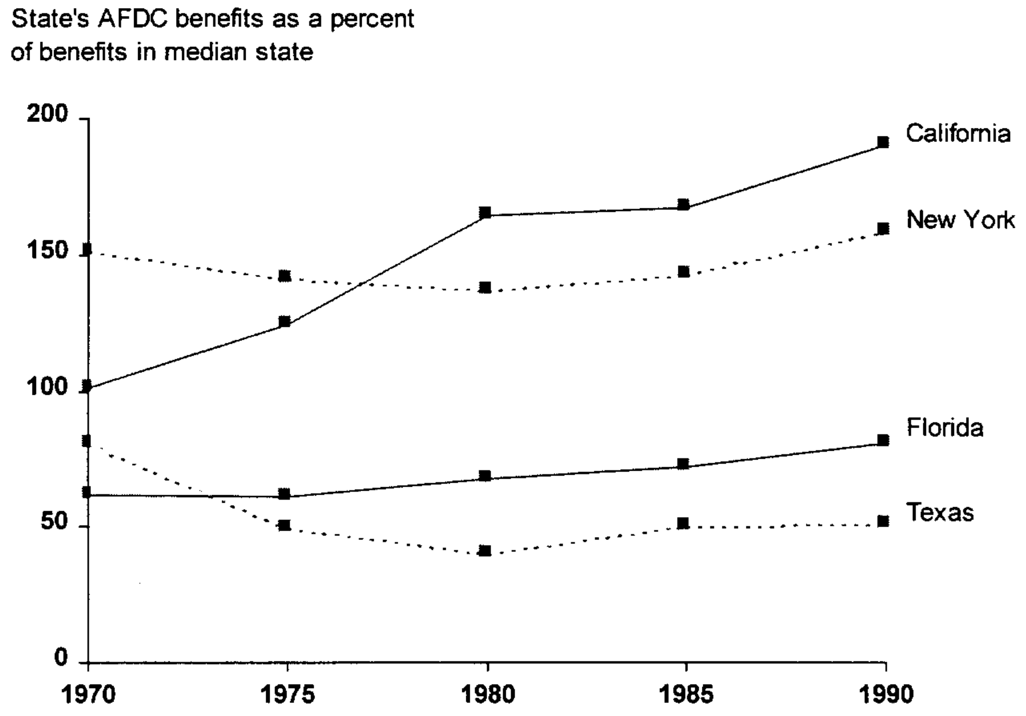


(c) $\eta_2 > \eta_1 > \eta_0$, and $\bar{w}_1 > \bar{w}_2$



(d) $\eta_2 > \eta_1 > \eta_0$, and $\bar{w}_2 > \bar{w}_1$

FIGURE 4. AFDC BENEFITS IN MAIN IMMIGRANT-RECEIVING STATES, 1970-1990



Source: U.S. House of Representatives (1993), pp. 666-667. The data refer to the maximum benefits for a three-person family.

TABLE 1. WELFARE PARTICIPATION RATES OF NATIVES AND IMMIGRANTS, BY STATE

	1980		1990	
State	Natives	Immigrants	Natives	Immigrants
Alabama	11.4	5.0	8.5	3.4
Alaska	7.6	5.0	8.8	8.3
Arizona	5.0	7.8	6.1	7.6
Arkansas	10.9	6.3	9.3	5.6
California	9.1	10.9	8.6	12.0
Colorado	5.5	7.8	5.4	7.6
Connecticut	6.0	5.5	5.1	5.4
Delaware	6.8	5.4	6.0	2.7
District of Columbia	12.5	6.1	9.6	3.6
Florida	5.9	10.2	5.5	8.5
Georgia	9.9	6.3	8.3	3.8
Hawaii	7.2	11.3	8.5	9.3
Idaho	4.5	3.9	5.7	4.5
Illinois	7.4	5.6	7.0	5.8
Indiana	5.2	5.0	5.1	4.1
Iowa	5.0	7.8	5.8	7.5
Kansas	4.9	5.3	5.5	7.0
Kentucky	9.3	6.9	9.3	4.3
Louisiana	11.2	8.2	10.8	7.2
Maine	9.5	9.5	7.6	8.0
Maryland	7.0	4.7	6.4	4.1
Massachusetts	9.2	11.4	7.2	10.2
Michigan	9.3	6.9	9.5	7.1
Minnesota	5.8	7.7	5.6	13.9
Mississippi	15.3	10.7	12.4	8.4
Missouri	7.4	6.0	7.2	3.8
Montana	4.5	8.9	6.3	6.0
Nebraska	4.6	6.1	4.6	5.6
Nevada	3.3	4.2	5.0	4.5
New Hampshire	5.7	6.8	4.6	4.1
New Jersey	8.1	6.3	5.4	5.8
New Mexico	8.6	8.2	7.7	9.9
New York	9.7	9.3	8.6	10.0
North Carolina	8.3	4.5	7.0	3.2
North Dakota	4.3	4.9	6.2	8.5
Ohio	7.7	5.3	8.6	5.4
Oklahoma	8.1	6.3	7.6	5.0
Oregon	6.4	7.2	5.9	7.2
Pennsylvania	8.6	7.3	7.3	6.6
Rhode Island	8.0	9.4	6.8	11.1
South Carolina	9.2	6.2	8.5	3.8
South Dakota	5.2	7.4	6.4	6.3
Tennessee	9.7	6.1	8.9	3.8
Texas	5.8	10.8	6.4	10.0
Utah	4.3	6.1	5.7	5.4
Vermont	7.8	8.4	6.8	4.2
Virginia	6.7	4.4	5.9	4.3
Washington	6.8	7.5	6.4	8.5
West Virginia	9.1	7.7	9.8	3.3
Wisconsin	7.2	7.2	7.2	10.5
Wyoming	2.2	3.8	5.7	7.0
United States	7.9	8.7	7.4	9.1

TABLE 2

GEOGRAPHIC CLUSTERING OF WELFARE RECIPIENTS IN CALIFORNIA
(Percent of households living in California)

Group	1980 Not on welfare	1980 On welfare	1990 Not on welfare	1990 On welfare
Natives:				
All households	9.7	11.2	9.6	11.5
Female-headed households with children	10.4	9.6	9.6	11.1
Immigrant households:				
All households	22.4	28.6	27.6	37.6
Non-refugee households	24.6	31.4	29.6	37.4
Non-Mexican households	17.8	22.4	22.0	33.1
Newly arrived immigrants (in U.S. less than 5 years):				
All households	30.1	36.9	28.9	45.4
Non-refugee households	31.2	37.4	30.0	43.7
Non-Mexican households	24.9	34.4	23.6	44.2
Immigrants; female-headed households with children:				
All households	26.1	27.6	30.6	36.2
Non-refugee households	27.4	28.6	31.9	34.3
Non-Mexican households	19.3	17.3	22.4	26.5
Newly arrived immigrants				
All households	33.4	32.7	31.3	41.0
Non-refugee households	34.3	30.9	32.5	37.0
Non-Mexican households	25.6	26.6	23.3	35.7

TABLE 3

DIFFERENCE-IN-DIFFERENCES ESTIMATE OF CLUSTERING GAP
(Dependent variable: Probability of residing in California)

Group	1980 (1)	1980 (2)	1980 (3)	1990 (1)	1990 (2)	1990 (3)
Newly arrived immigrants:						
All	.053	.039	.058	.147	.153	.122
	(.006)	(.006)	(.006)	(.015)	(.015)	(.016)
Non-refugees	.047	.032	.033	.118	.126	.080
	(.008)	(.007)	(.007)	(.020)	(.020)	(.020)
Non-Mexicans	.079	.055	.069	.187	.192	.155
	(.006)	(.006)	(.006)	(.016)	(.016)	(.018)
Newly arrived female-headed households:						
All	.002	-.015	.011	.081	.073	.079
	(.014)	(.014)	(.015)	(.033)	(.032)	(.036)
Non-refugees	-.026	-.042	-.002	.030	.038	.044
	(.016)	(.016)	(.016)	(.041)	(.040)	(.041)
Non-Mexicans	.018	-.011	.011	.108	.091	.095
	(.015)	(.015)	(.016)	(.037)	(.036)	(.042)
Includes socioeconomic characteristics	No	Yes	Yes	No	Yes	Yes
Includes national origin fixed effects	No	No	Yes	No	No	Yes

Notes: Standard errors are reported in parentheses. The coefficients estimate the difference in the probability of residing in California between welfare and non-welfare recipients in the respective immigrant sample *relative* to the difference between welfare and non-welfare recipients in the native population. The vector of socioeconomic characteristics includes the education, gender, and age of the household head, the number of persons living in the household, the number of children under age 18, the number of persons over age 65, and a vector of race dummies (black, Hispanic, or Asian). The regressions that include a vector of national origin fixed effects differentiate among 91 immigrant groups.

TABLE 4

IMPACT OF SELECTION IN OBSERVABLE CHARACTERISTICS
ON RESIDENCE IN CALIFORNIA
(Dependent variable: Probability of residing in California)

	Newly arrived immigrants		Female-headed, newly arrived immigrants	
	1980	1990	1980	1990
Interaction of immigrant indicator with:				
Sex	.018	.005	---	---
	(.004)	(.010)		
Education	-.010	-.012	-.009	-.010
	(.000)	(.001)	(.002)	(.003)
Age	.000	.000	.000	-.002
	(.000)	(.000)	(.001)	(.001)
No. of children < 18	-.016	-.011	-.012	-.005
	(.002)	(.005)	(.008)	(.016)
No. of persons ≥ 65	.019	.043	.047	.085
	(.006)	(.014)	(.026)	(.051)
No. of persons in household	.025	.032	.025	.024
	(.002)	(.004)	(.006)	(.010)
No. of observations	414,165	469,022	32,469	43,130

Notes: Standard errors are reported in parentheses. All the regressions also include the education, gender, and age of the household head, the number of persons living in the household, the number of children under age 18, the number of persons over age 65, a vector of race dummies (black, Hispanic, or Asian), and a vector of 92 national origin fixed effects (where one of the variables in this vector indicates native-born status).

TABLE 5. DETERMINANTS OF PROBABILITY OF RECEIVING WELFARE, USING POOLED 1980 AND 1990 CENSUS

Group	Log AFDC benefit level	Log per capita income	Unemployment rate	Measures of ethnic enclave	
				q_1	q_2
Female-headed households with children					
Natives (N=67,775)	.088	-.334	.010	---	---
	(.031)	(.082)	(.003)		
Natives; weighted sample	.062	-.258	.021	---	---
	(.082)	(.174)	(.006)		
Recent immigrants (N=7,824)	.208	-.366	.014	---	---
	(.103)	(.265)	(.008)		
Recent immigrants	.234	-.418	.012	-1.487	.187
	(.100)	(.250)	(.007)	(.431)	(.047)
Recent immigrants, with national	.271	-.261	.019	-1.129	.122
origin fixed effects	(.090)	(.188)	(.005)	(.432)	(.056)
All immigrants (N=53,285)	.112	-.122	.003	---	---
	(.054)	(.072)	(.004)		
All immigrants	.119	-.156	.003	-1.109	.184
	(.060)	(.072)	(.004)	(.497)	(.075)
All households					
Natives (N=796,074)	.001	-.043	.002	---	---
	(.009)	(.025)	(.001)		
Natives; weighted sample	-.027	-.062	.007	---	---
	(.039)	(.079)	(.003)		
Recent immigrants (N=87,113)	.032	-.121	.000	---	---
	(.026)	(.076)	(.003)		
Recent immigrants	.042	-.162	-.001	-1.256	.099
	(.029)	(.084)	(.003)	(.184)	(.031)
Recent immigrants, with national	.019	-.155	.001	-.990	.066
origin fixed effects	(.022)	(.062)	(.002)	(.337)	(.027)
All immigrants (N=655,328)	-.019	-.040	-.003	---	---
	(.015)	(.025)	(.001)		
All immigrants	-.016	-.060	-.002	-.868	.105
	(.020)	(.024)	(.001)	(.225)	(.032)

Notes: Standard errors are reported in parentheses. The standard errors are adjusted for the grouping of observations within a state. The variable q_1 gives the fraction of the state's population that belongs to the same national origin group as the household, and the variable q_2 gives the fraction of the immigrant group that lives in the particular state. The regressions hold constant the household head's education, age, and gender, the number of persons residing in the household, the number of persons under age 18, the number of persons over age 65, a vector of race dummies, a dummy variable indicating if the observation was drawn from the 1990 Census, and a vector of state fixed effects. The regressions in the immigrant samples also include a vector of dummy variables indicating the year of migration. The national origin fixed effects differentiate among 91 national origin groups in the immigrant population.

TABLE 6. SIMULATED PROBABILITIES OF RECEIVING ASSISTANCE AND THE LEVEL OF AFDC BENEFITS
(Sample of female-headed households with children)

	Definition of household income:	
	Observed income	Predicted income
Coefficient of within-state change in log AFDC benefits for:		
Newly arrived immigrant households	.105	.109
	(.070)	(.075)
Native households	.126	.120
	(.051)	(.056)
Percent of welfare recipients predicted to live in California in 1990:		
Newly arrived immigrant households	33.8	32.3
Native households	11.1	11.4
Percent of welfare recipients who actually lived in California in 1990:		
Newly arrived immigrant households	41.0	41.0
Native households	11.1	11.1

Notes: Standard errors are reported in parentheses. The "simulated" welfare probabilities are obtained by determining if a particular household in a national sample of immigrant (or native) households qualifies for AFDC benefits in each of the states (in both 1980 and 1990), and by adjusting these eligibility probabilities by state-specific take-up rates.

TABLE 7

ESTIMATED BENEFIT ELASTICITIES IN SAMPLE OF NATIVE MOVERS

	Female-headed with children	All households
Native movers	-.009	-.010
	(.082)	(.010)
Native movers, weighted to resemble immigrant sample	.257	.123
	(.200)	(.061)
Sample size	6,276	75,427

Notes: Standard errors are reported in parentheses. The standard errors are adjusted for the grouping of observations within a state. The regressions hold constant the household head's education, age, and gender, the number of persons residing in the household, the number of persons under age 18, the number of persons over age, the state's log per-capita income and unemployment rate, and a vector of state fixed effects.

TABLE 8

ESTIMATED BENEFIT ELASTICITIES FOR FEMALE-HEADED HOUSEHOLDS RESIDING OUTSIDE IMMIGRANT-RECEIVING STATES

	Living outside California		Living outside New York	
	Benefit elasticity	Sample size	Benefit elasticity	Sample size
Natives	.092	61,024	.098	62,476
	(.037)		(.031)	
Natives, weighted to resemble immigrant sample	.084	61,024	-.008	62,476
	(.092)		(.067)	
Recently arrived immigrants	.148	5,102	.305	6,520
	(.128)		(.132)	

Notes: Standard errors are reported in parentheses. The standard errors are adjusted for the grouping of observations within a state. The regressions hold constant the household head's education, age, and gender, the number of persons residing in the household, the number of persons under age 18, the number of persons over age 65, a vector of race dummies, a dummy variable indicating if the observation was drawn from the 1990 Census, the state's log per-capita income and unemployment rate; the fraction of the state's population that belongs to the same national origin group as the household, the fraction of the immigrant group that lives in the particular state. and a vector of state fixed effects.