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SPECIAL ISSUE PAPER

# International immigration and domestic out-migrants: are domestic migrants moving to new jobs or away from immigrants?

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**Abstract** Domestic migrant responses to geographically concentrated immigration flows play central roles in determining the local economic impacts of immigration and the geography of the ethnic composition of the population. Possible motivations for domestic migrant responses include: increased labor market competition associated with new immigrants and ethnic or cultural avoidance. We use US annual state-to-state migration flows from the Internal Revenue Service to assess the existence and nature of the link between geographically concentrated immigration and domestic migration. We find some evidence of a domestic migrant response to immigrants, particularly to greater cumulative shares of the foreign born, which we interpret as providing some support of the ethnic or cultural avoidance hypothesis.

**JEL Classification** R23 · R11 · F22 · R12

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## 1 Introduction

The domestic migration response to geographically concentrated flows of immigrants has become a central issue in the immigration debate. Geographic areas serving as gateways for immigrants in the United States often experience significant net domestic out-migration. The issue is of importance because the out-migration response affects the geographic spread of labor market effects of immigration (Borjas et al. 1996) and the potential for geographic ethnic segregation (Frey 1995a), as well as related sorting by class. Likewise, it is a key determinant of how immigration affects local and regional net population growth.

From 1985 to 1990, among the high immigration states of California, New York, Texas, New Jersey, Illinois, and Massachusetts, all but California experienced significant domestic out-migration, with California subsequently experiencing net domestic out-migration after 1990 (Frey 1995a). The top six gateway cities for immigrants from 1990 to 1996 (Los Angeles, New York, San Francisco, Chicago, Miami, and Washington, D.C.) collectively received more than 2.8 million immigrants, while also losing 3.4 million net domestic migrants (Frey 1999a). At the US county level, Partridge et al. (2008b, 2009) report significant domestic out-migration as associated with higher rates of immigration for both nonmetropolitan and metropolitan areas for the period 2000–2005. From 2000 to 2009, four of the six states with the most number of immigrants (California, Texas, Florida, New York, Illinois, and New Jersey) also had the most domestic out-migrants. Yet when considered as population shares rather than absolute numbers, Nevada and Arizona become among the top immigration states, replacing New York and Illinois, and hence, only two of the six top immigration states have the highest rates of domestic out-migration.[1] Explanations in the literature for the potential connection between net domestic migration and immigration have been varied.

A primary explanation focuses on the potential displacement of domestic residents by immigrants in the local labor market. According to what we term the Borjas (2003, 2005) "bathtub" model, immigrants represent an increase in the local labor supply, which reduces the local wage rate. To the extent immigrants and domestic workers are perfect substitutes, the reduction in the wage rate induces domestic out-migration, only ceasing when the wage rate climbs back to the spatial equilibrium level. Net out-migration can manifest itself either through out-migration of existing residents (Card 1990, 2000; Card and DiNardo 2000) or by attracting less domestic residents from elsewhere (Filer 1992; Keeton and Newton 2005). Labor market effects then are dispersed across the nation, much as water added to a bathtub spreads out.

Frey (1995b,c) finds rates of out-migration among local natives in high immigration areas to be greatest for those with a high school education or less. This presumably occurs because they primarily compete with low-skilled immigrants for jobs. Walker et al. (1992) similarly find a net loss of one blue-collar worker for every seven

[1] The correlation across states (including the District of Columbia) between the numbers of immigrants and domestic migrants is −0.32, while that between the flows when taken as shares of state population is 0.15. All figures are based on calculations by the authors using US Census Bureau data accessed at http://www.census.gov/popest/states/NST-comp-chg.html on December 23, 2009.

immigrants into a metropolitan area. Partridge et al. (2008b) report that the displacement of domestic residents by new immigrants in nonmetropolitan US counties most likely occurred because they have a relatively low-skilled employment distribution (e.g., food processing, agriculture).

Yet other forces may produce the opposite result or no relationship whatsoever. Complementarity in production between low-skilled immigrants and both high-skilled domestic residents and capital can produce a positive relationship between high-skilled domestic migration and immigration (Walker et al. 1992). If immigrants locate in areas where there are shortages for their skill type, rather than domestic out-migration, population of the area will grow (Saiz 2003). Immigrants may fill jobs that domestic residents do not want and even create jobs at a scale that exceeds what would happen in the absence of immigration (Linton 2002). To be sure, Partridge et al. (2008a) find heterogeneous effects in which immigration appears to have its strongest (net-negative) impacts on county employment in the Western United States.

Wright et al. (1997) attributes a statistically positive relationship between immigration and the number of natives with high levels of education to US metropolitan areas serving as immigration gateways while also undergoing labor-market restructuring from globalization. The demand for high-skilled labor increases, while "deskilling and downgrading in both manufacturing and service sectors" (p. 239) reduces low-skilled wages, inducing their out-migration, with immigrants filling the newly created low-wage jobs.[2] Frey (1996) contends, however, that the decline in urban manufacturing jobs was greatest in the 1970s and 1980s, not the 1990s when the negative relationship between immigration and domestic out-migration was strongest.

The use of terms such as "balkanization" (Frey 1996) and "white flight" (Frey and Liaw 1998) to describe demographic trends in immigration and domestic migration suggests that cultural or ethnic differences underlie domestic out-migration from gateway cities (Ellis and Wright 1998). Admittedly, labor market competition arising from concentrated immigration affects age and education groups differentially, also often falling along ethnic lines. However, in what (Ley 2007, p. 232) classifies as "cultural avoidance" in his taxonomy of explanations for the nexus between immigration and domestic migration, white domestic residents may be reluctant to have neighbors of differing cultures and ethnicity. This reluctance includes the possibility that immigrants are associated with increased social costs, leading to domestic out-migration aside from labor market considerations. For example, Alesina et al. (1999) report lower levels of local public services in urban areas with more diverse populations.

More recent evidence on the composition of domestic out-migrants casts doubts on the cultural avoidance explanation. Relative to their population shares, Frey (2003) reports that whites were underrepresented and nonwhites were overrepresented among domestic out-migrants from New York City and Los Angeles dating from 1995 to 2000. Suro and Singer (2002) find greater out-migration and less in-migration of the lesser-educated across all races and ethnicity in states with high levels of foreign-born

[2] Trejo (1997) provides evidence that Mexican immigrants have particularly low levels of average education, which accounts for their lower wages. In addition, Rivera-Batiz (1999) reports that undocumented workers have lower skills than documented workers, though their wages are still lower than what would be expected given their skills.

education. Similarly, Kritz and Gurak (2001) find that in only five states was a net gain in working-age foreign-born men accompanied by a net loss of native-born non-Hispanic men from 1985–1990, with only one state (Hawaii) among the five high-immigrant states.

Along with potential complementarity between immigrants and natives in production, Ottaviano and Peri (2006) suggest that the variety of urban consumption amenities such as ethnic restaurants accompanying a high rate of immigration is attractive to native households. They provide evidence in the form of increased native-born wages and housing rents between 1970 and 1990 in metropolitan areas where the foreign-born share increased. Similarly, wage evidence is provided by Greenwood et al. (1996), who find current immigrants only having significant adverse wage effects on recent immigrants, but not on natives or longer established immigrants, suggesting that current immigrants are less substitutable with natives or established immigrants. Cortes (2008) reports similar results, suggesting the degree of substitutability relates to common language difficulties among current and recent immigrants.[3]

Ottaviano and Peri (2006, 2008) note that considering the labor market in the aggregate versus focusing solely on narrow groups (e.g., low-educated whites) is necessary to capture the full range of general equilibrium outcomes that may occur. For instance, a new Thai immigrant restaurant may displace an existing American style diner, producing no net change in restaurant employment, though abundant low-skilled immigrants and cultural diversity alternatively may attract highly educated domestic workers. Another key factor is how the capital stock adjusts to influxes of immigrants (Ottaviano and Peri 2008).

Therefore, in this study, we examine the nexus between domestic migration and immigration for US states using Internal Revenue Service place-to-place migration data for 1990–2007. We find evidence of domestic migration responses to immigration. Primarily, however, support is found for the ethnic or cultural avoidance hypothesis rather than the existence of labor market competition effects.

## 2 Empirical implementation

The theoretical model begins with Borjas (2003, 2005) bathtub model as implemented by Partridge et al. (2008b, 2009). It is a basic labor demand and labor supply model with domestic and immigrant labor serving as perfect substitutes. In equilibrium, real wages are equalized across regions. In disequilibrium, workers migrate to reestablish equilibrium wage levels across all labor markets. The addition of an immigrant worker in the region implies that one domestic worker will out-migrate to another region. This follows because new immigrants increase local labor supply and reduce wages. Domestic workers then out-migrate until wages are equalized across all local labor markets. The effects of immigration are dispersed across the country, leading to a "bathtub" effect. What we add to the bathtub model is other nonlabor market motivations that affect utility beyond just wage levels—in particular, factors such as cultural avoidance.

[3] See Longhi et al. (2005) for further discussion of how immigrants affect native-born wages.

Like the past literature, we do not focus on government policies, which unlike economic conditions, are more likely to change relatively slowly between state pairs.[4]

In contrast to Partridge et al. (2008b, 2009) examination of aggregate net migration flows, we examine place-to-place domestic migration flows using US Internal Revenue Service (IRS) data. This allows for more direct analysis of the type of areas selected by domestic migrants. Further it allows for consideration of the role of distance in influencing domestic migration between regions. A disadvantage of using place-to-place flows is the extensive number of zeros between pairs of counties, which leads to our use of states as the units of analysis. Although states are not "local" labor markets, they provide a lower-bound estimate of "displacement" because native residents may relocate within their given state in response to recent immigrants locating in their community (Borjas et al. 1996; Cortes 2008).

There are 1,176 state-to-state migration flows for the lower 48 states plus the District of Columbia. This provides considerably more information than the standard approach, which would be analogous to only estimating 49 net migration rates on immigration rates and other control variables (e.g., Borjas 2005). Moreover, the state-to-state data allow us to consider whether the domestic out-migrants are moving to states with relatively greater (or lower) shares of the foreign born than the origin state, which is an issue that has not been considered in past research. So we can examine whether domestic out-migrants are primarily driven by labor market effects or by possible aversion to states with greater shares of the foreign born, not just new immigrants.

The IRS migration data are based on personal income tax returns. Tax return data typically form the core of US Census Bureau estimates of domestic migration. A state-to-state migration occurs when the address of the filer has changed states between tax years. The gross in- and out-flow numbers are then based on the number of exemptions on individual tax returns. Not every internal US migrant files a tax return, but the underlying assumption is the unreported domestic moves are in proportion to the IRS migration estimates. Likewise, immigration figures also are derived from the tax return data, where new immigrants reflect the number of exemptions on tax returns in which the filer lived abroad in the past year. Immigrants are defined as the net of those directly arriving from a foreign country into a state versus those moving out of a state to a foreign country. Domestic migration includes individuals already residing in the United States who change their state of residence, which include both US natives and the foreign-born who have resided in the United States.

Much of the previous literature considered decade-long periods because of their reliance on decennial census data (Card 2000; Borjas et al. 2008; Ottaviano and Peri 2008). While data availability was a key factor, this also has the advantage that year-to-year idiosyncrasies are smoothed over longer periods, and it may take time for domestic migration flows to respond to changes in immigration patterns and economic

[4] Given our focus on economic conditions and immigration, another reason for not controlling for government policies is that they could be affected by economic conditions or by immigration flows—e.g., states and localities may reduce benefits in the face of large numbers of immigrants or raise taxes during weak economic times. Thus, controlling for government policies would potentially be endogenous with migration, and it could 'steal' some of the effects of the economic and immigration variables that we desire to measure. For a survey of how regional government policies affect growth, see Brown and Taylor (2006) and Dalenberg and Partridge (1995).

conditions. In our case, because the IRS data uses tax returns and other data such as employment are annual averages, the data are not perfectly aligned on an annual basis, so considering a long period helps to smooth over these differences. We use 1993–2007 data for our analysis. The beginning period is chosen to correspond to when the 1990s economic expansion was firmly in place, whereas the ending period is just before the December 2007–June 2009 economic recession (in which domestic migration is measured early in each calendar year). We replicate regression specifications used in the previous literature to assess whether this time period affects the general patterns and conduct sensitivity analysis by splitting the time period into halves.

We alternately specify four different domestic migration outcome measures as our dependent variable (including District of Columbia but excluding Alaska and Hawaii). First, we begin with the aggregate net migration measure for each state *i* that is typical in this literature:

$$[\text{NETMIG}_i/\text{POP}_i] \times 1000000, \tag{1}$$

which is defined as the average annual net migration into a given state over the 1993–2007 period divided by the beginning 1993 population. The advantage of the overall net migration rate is that it replicates past research (e.g., Borjas 2005; Partridge et al. 2008b, 2009). Likewise, using aggregate net migration is consistent with Borjas' model of aggregate labor-market demand and supply—it is not an individual state-to-state model of migration. However, there are three problems with using aggregate net migration. First, there are only 49 observations per period using state data. Second, when using aggregate net flows, we are unable to identify the types of states in-migrants are choosing. Third, it ignores the role of distance because for each state the flows between it and all other states are weighted equally. For example, a large out-flow in one state may produce greater migration flows to nearby states than suggested by their characteristics because of close proximity (Douglas 1997).

We next consider state-to-state migration flows in an attempt to uncover the types of states domestic migrants favor when they exit a given state. Our first measure considers every state-to-state net migration pair (e.g., Alabama has net migration flows with each of the other states). Denoting the gross in-migrants moving to state $i$ from state $j$ as $M_{ij}$, we employ the following two state-to-state net migration measures:

$$[(M_{ij} - M_{ji})/((\text{Pop}_i + \text{Pop}_j) \times 0.5)] \times 1000000 \tag{2}$$

$$[(M_{ij} - M_{ji})/(\text{Pop}_i \times \text{Pop}_j)] \times 1000000 \times 1000000, \tag{3}$$

where the annual migration figures are averaged over the 1993–2007 time period, while the population numbers are measured at the beginning of the period. Our preference is the measure in Eq. 2 because it is closest to the net migration definition in Eq. 1, in which the respective two state populations are averaged in the denominator.[5]

[5] The measure in Eq. (3) is used by Douglas (1997) to measure state-to-state migration, whereas Tabuchi and Kentaro (forthcoming) note that the dependent variable used in assessing state-to-state migration patterns may matter. By considering both migration measures, we assess whether our conclusions are robust.

The advantage of the net migration measure is that it captures the relative growth of a particular place due to differential economic or quality-of-life reasons—i.e., households voting with their feet. Conversely, gross migration flow data are considerably more noisy because many migrants move for "random" reasons that are not easily accounted for by regression controls—e.g., family and other personal reasons. Thus, state-to-state net flows balance out "random" migration flows to reflect perceived utility differentials.

We next consider in-migration rates into destination state $i$ from origin state $j$ as the dependent variable.

$$M_{ij}/\mathrm{POP}_j \tag{4}$$

The trade-off with using this measure is that our theoretical immigration model is not based on gross migration flows. While the state-to-state net migration models are preferred because of their consistency with past literature, the resulting empirical results help determine which particular destination state characteristics are associated with in-migration. Another advantage is that the destination state fixed effects account for other unmeasured factors such as industrial restructuring, age of infrastructure, taxes, quality of public services.

All of the explanatory variable groups are measured as the destination state characteristics minus the origin-state characteristics. For example, job growth is measured as employment growth in the destination state minus that in the origin state. This reveals whether net migration between state pairs are affected by which states have relatively higher rates of immigration or initial immigrant population shares.[6] Using the four migration outcomes ($\mathrm{OUTCOMES}_{ij}$) as the dependent variable, the base regression model can be written as:

$$\begin{aligned}\mathrm{OUTCOMES}_{ij} = {} & \beta_0 + \beta_1(\mathrm{IMG}_i - \mathrm{IMG}_j) + \beta_2(\mathrm{IMGSH}_i - \mathrm{IMGSH}_j) \\ & + \beta_3(\mathrm{EMP}_i - \mathrm{EMP}_j) + \beta_4(\mathrm{WAGE}_i - \mathrm{WAGE}_j) \\ & + \beta_5(\mathrm{GEOG}_i - \mathrm{GEOG}_j) + \sigma_s + e_{ij},\end{aligned} \tag{5}$$

where IMG is the average annual number of international immigrants that moved to the state over the period of interest divided by the initial 1993 state population. IMGSH measures the initial (1990) share of the state's population that is foreign born. EMP is state employment growth over the period, whereas WAGE is the initial 1993 wage level. The geography measures include the state average of the US Department of Agriculture's amenity index score, which ranges from 1 (lowest) to 7 (highest). Another geography measure is an indicator variable for whether the state borders the Atlantic Ocean, the Pacific Ocean, or the Gulf Coast. Finally, when the in-migration rate is the dependent variable, the models include the log of the 1993 origin-state population,

[6] One could also assess whether these effects are stronger in the origin or destination state by separately including the origin and destination state variables in the model, which is a topic we leave to future research. Specifically, we are not asking whether the marginal effects of origin or destination states are different, but rather does having a greater immigration share in one state versus another state affect migration patterns between the two.

distance between the origin and destination states (measured from population-weighted centroid), and state fixed effects ($\sigma_s$).[7] The residuals are denoted as $e_{ij}$.

Regarding the contemporaneous immigration flow measure, a 'Borjas' bathtub (2003, 2005) model suggests that all else equal, states with greater contemporaneous immigration rates experience greater rates of domestic out-migration to states with lower immigration rates.

Borjas' base model does not directly consider how the *initial* stock of immigrants affects domestic migration because presumably past flows of immigrants would have already affected wages levels and *past* domestic migration flows. Yet if a given share of domestic households has a preference to live in a place with a low share of immigrants—perhaps due to 'noneconomic' concerns—then the initial population share of immigrants in a given state would be positively related to domestic out-migration and negatively related to domestic in-migration.

The tax return data do not differentiate natives from long-term immigrants for domestic moves. So both groups are treated as "natives" in the immigrant flow variable. There are two offsetting effects for longer-term (established) immigrants in terms of their migration propensity relative to natives. First, they are likely less mobile than native-born residents.[8] Yet if new immigrants are a close substitute to somewhat less recent immigrants, they are more prone to migrate, which would be consistent with the Borjas bathtub model. Nonetheless, because our goal is to understand relative population rates in response to immigration, not being able to differentiate native-born *domestic* migrants from long-term immigrant *domestic* migrants is not crucial, i.e., we are examining how immigration affects net growth differentials.

The IRS data likely misses many undocumented workers, which is a common problem in past research. Typically, documented and undocumented workers are assumed to be highly correlated because they likely would be attracted to the same locations due to either local economic growth or local immigrant network effects. A high correlation between documented and undocumented immigrants creates a simple scaling issue that is not of serious consequence. Namely, because the immigrant share is scaled down, the estimated immigrant coefficient would be correspondingly scaled up. Yet if the correlation between the documented and undocumented migration flows is low, this would create measurement error bias or omitted variable bias, in which the latter would apply if undocumented immigrant flows respond to socioeconomic factors that are not controlled for in the empirical model.

The Census Bureau attempts to measure both documented and undocumented workers and likely misses fewer undocumented workers than does the IRS. Again, the scaling issue likely works to increase the size of the immigrant share coefficient. Any measurement error would likely bias the *t*-statistic to zero. Because we doubt cultural

[7] Distance is not included in state-to-state net migration models because the distance effects that reduce migration gross flows from (say) Alabama to Arizona would likewise depress gross migration flows from Arizona to Alabama—producing little net impact.

[8] Keeton and Newton (2005) report that between 1995 and 2000, new immigrants accounted for about three times greater influence on metropolitan area population growth differentials than net migration flows of *established* immigrants, though the latter group is much larger. Likewise, they show that established immigrant net migration flows are much less responsive to local employment growth than either new immigrants or domestic migrants.

avoidance strongly affects past immigrants, the results more likely reflect the responses of natives.

Relative employment growth should attract in-migrants and deter out-migrants (Partridge and Rickman 2003). Initial wage levels have a more ambiguous impact on net domestic migration. For firms, it may deter start-ups and expansion, which would reduce demand for workers and dampen net migration flows. Yet from the household perspective, higher initial wage levels would attract greater net migration flows. Controlling for initial wage levels accounts for the possibility that the initial immigrant share affects initial wage levels through composition effects (presumably lowering it on average), and thus, the initial immigrant share variable will more cleanly control for the noneconomic effects of immigration. The state wage and employment data are taken from the Bureau of Economic Analysis Web site (www.bea.gov).

We expect that net migration flows are positively related to the amenity index and to whether the state borders an ocean. Also, in-migrant flows should be positively related to origin-state population, while distance between origin and destination states would depress gross in-migration flows. The in-migration models also include the destination state fixed effects ($\sigma_s$) because they account for other characteristics that in-migrants consider when choosing a state—e.g., the quality of housing, government services, or unmeasured industrial restructuring.

## Endogeneity and instrumental variables

A primary empirical concern is that immigration may be endogenous because positive economic shocks can jointly increase immigration and net migration. Following the literature (Card 1990, 2000; Card and DiNardo 2000), we use deep lags of past immigrant stocks as instruments for current immigrant flows (which assumes that long-established immigrant networks attract new immigrant flows). Using Census data from the Geolytics company, the identifying instrumental variables for immigration are the 1970 share of the population that is either foreign born, or have one foreign-born parent, and the 1980 population share that is foreign born, which follows Partridge et al. (2008b, 2009). To account for the key role of Mexican immigrants, another instrument is an indicator for states bordering Mexico. If these factors are not associated with current immigration flows, they would be weak instruments in the first-stage model, though we do not find this to be the case below.

Employment growth also could be endogenous to the same demand shocks that affect immigration. As our job growth instrument, we use the state industry mix employment growth rate, which has been widely used as an exogenous instrument (Bartik 1991; Blanchard and Katz 1992). It is defined as the initial year's state employment shares in each (one-digit) industry multiplied by the national growth rate in each industry and then summed across all industries, forming the hypothetical employment growth rate if the state's industries grew at the national average over the sample period. Thus, changes in *national industry* demand are the exogenous shifters.

Demand shocks also could affect the initial-year wage level. Analogous to the industry mix variable, our identifying instrument for wage levels is a "wage mix" variable defined as the initial-year industry employment shares in each of state's industries

multiplied by the national wage level in each industry, summing this across all industries. This value forms the hypothetical state wage rate if each of its industries paid the corresponding national average wage. National wage differences across industries then are the exogenous shifters.

## 3 Results

Table 1 reports the unweighted means and standard deviations for the variables across the samples we employ. The first set of empirical results is reported in Table 2. Columns 1 and 2 report the results using the standard net migration model employed by Card (2000) and Borjas (2005) with the aim of replicating the county-level results of Partridge et al. (2008b, 2009), though we employ a different time period.[9] Column 1 contains the OLS results, while Column 2 shows 2SLS results treating 1993 wage level, 1993–2007 employment growth, and 1993–2007 immigration rate as endogenous. Both the OLS and IV results show the same pattern at the state level as prior studies found for counties that higher immigration rates are associated with lower net migration rates.[10]

Column 3 reports 2SLS estimates from a model that substitutes 1993–2007 Census population growth for the (domestic) net-migration-dependent variable. The annual average net immigration coefficient is statistically insignificant, which again suggests that immigration flows are offset by out-migration of native domestic residents. Thus, the aggregate state-level findings are consistent with previous county-level findings despite the use of a different time period and a higher-level of geographic aggregation that obscures migration at the county level.

The primary aim of this study is to assess whether migrants mostly move to better economic opportunities or whether they are influenced by "noneconomic motivations" related to high levels of past immigration. Thus, we assess whether controlling for contemporaneous immigrant flows, the initial stock of immigrants also is negatively associated with domestic net migration, because the initial stock may be associated with noneconomic motivations.

Column 4 adds the 1990 foreign-born share, measured three years before the starting period of the immigrant flow variable, to the net migration model. As can be seen, higher contemporaneous immigrant flows are associated with less domestic net migration, in which the coefficient is a slightly smaller than when the 1990 immigrant share is *not* included. Moreover, the 1990 immigrant share variable also is negative and

[9] The definition of the net migration variables slightly differ, though it does not affect the conclusions.

[10] The first-stage results suggest that the instruments for the 1993 wage level variable are strong ($F = 12.0$ in the first stage) and moderately strong for the immigrant variable (F nearly equals 8.7 in the first stage), though the instruments appear to be weak for the employment growth variable ($F = 2.7$ in the first stage). However, the mixed performance of the instruments in this model is likely a function of the limited sample size of 49 in this set of regressions. We are not concerned with the performance of our instruments though because the IV immigrant coefficient differed from the OLS coefficient in the expected manner, and this set of net migration results are more for replication of past research. Below, in our key, state-to-state regressions with a sample size of well over 1,000, the identifying instruments are consistently very strong.

**Table 1** Means and (standard deviations)

| | (1) Aggregate net migration rate/ %population change | (2) State-by-state net migration rate | (3) State-by-state in-migration rate |
|---|---|---|---|
| *Dependent variables* | | | |
| Avg. annual net migration rate at the state level (1993–2007)[a] | 839.1 (5402.7) | na | na |
| Avg. annual population growth rate at the state level (1993–2007)[a] | 0.95 (0.80) | na | na |
| Avg. state-by-state net migration rate EQN 2[b] | na | −0.18 (204.07) | na |
| Avg. state-by-state net migration rate EQN 3[c] | na | −1.12 (99.64) | na |
| Avg. state-by-state in-migration rate[d] | na | na | 529.6 (1126.9) |
| *Independent variables* | | | |
| Diff in immigration rate 1993–2007 | 417.05 (330.36) | 53.93 (464.28) | −4.31e-10 (467.3) |
| Diff in % foreign born 1990 | 4.72 (4.39) | 1.02 (6.12) | 0.0 (6.21) |
| Diff in employment growth 1993–2007 | 0.26 (0.16) | −0.001 (0.22) | 3.72e-11 (0.22) |
| Diff in ocean border dummy | 0.41 (0.5) | 0.01 (0.70) | 0.0 (0.70) |
| Diff in amenity rank | 3.75 (1.02) | 0.09 (1.44) | 1.58e-12 (1.44) |
| Diff in ln(1993 wage) | 10.08 (.15) | 0.04 (.21) | 2.73e-11 (.21) |
| ln(origin pop93) | na | na | 15.00 (1.01) |
| Distance (km) | na | na | 1658.73 (974.71) |
| Sample size | 49 | 1176 | 2,352 |

Unweighted means and standard deviations are only reported when the variable is used in a specification with that sample. See the text for more details of the variable definitions

[a] Net migration rate is defined as the state's net migration between 1993 and 2007 divided by 1993 population, multiplied by one million (Eq. 1). The population growth rate is defined as the percentage change in population between 1993 and 2007. The sample statistics are for those used in for the models in Table 2

[b] The dependent variable defined as $[(M_{ij} - M_{ji})/((\text{Pop}_i + \text{Pop}_j) \times 0.5)] \times 1000000$. The dependent variable is used in column 1 of Table 3

[c] The dependent variable defined as $[(M_{ij} - M_{ji})/(\text{Pop}_i \times \text{Pop}_j)] \times 1000000 \times 1000000$. The dependent variable is used in column 2 of Table 3

[d] The dependent variable defined as $M_{ij}/\text{POP}_j$ in Eq. 4. The dependent variable is used in columns 3 and 4 of Table 3

statistically significant, illustrating that on balance, domestic migration is negatively related to the initial stock of immigrants.

The negative relationship between 1993 and 2007 domestic net migration and the initial 1990 immigrant share is consistent with the notion that domestic migrants avoid places with more immigrants cumulatively and are not just crowded out by contemporaneous migration flows in which immigrants are substitutes for domestic

labor supply.[11] Likewise, because initial wages are accounted for, the initial immigrant stock influence is presumably not working through wage effects. Taken together, these results are consistent with the cultural/ethnic avoidance hypothesis.

While the aggregate state-level results in Table 2 suggest that domestic migrants are leaving states with higher immigrant flows and initial stocks, they do not directly address our hypothesis of whether these domestic migrants are avoiding states with higher immigrant flows and (especially) stocks. To consider this, columns 1 and 2 of Table 3 report the IV results of the state-to-state net migration models to directly examine whether on a state-by-state basis, domestic migrants tend to locate in particular states with different initial immigrant stocks.

These models use the state-to-state-dependent variables defined in Eqs. 2 and 3. As noted above, we prefer the measure in column 1 because it is most directly comparable to the net migration and population growth models that we employ. Also, the $R^2$ statistic is much higher for the model reported in column 1, which also suggests that using this dependent variable fits the data better. Before turning to the regression results, note that the joint Cragg-Donald $F$-statistic for the strength of the instruments is over 14, suggesting the instruments are strong.

The results suggest that state-to-state net migration rates are not statistically related to most of the variables, with most being insignificant. However, in both models, the 1990 foreign-born share is negative and statistically significant, suggesting that states with higher initial foreign-born shares receive fewer domestic net migrants. Conversely, the difference in the respective state 1993–2007 immigration rates is statistically insignificant. Thus, while the aggregate state results in Table 2 suggest that contemporaneous immigration flows matter, when considering disaggregate state-to-state migration flows, migrants appear to be more influenced by the initial immigration shares. The results suggest that some domestic migrants are avoiding states with high initial shares of immigrants. Finally, there is evidence that migration flows are positively related to being near an ocean and to states with high amenities (though the latter variable is only very marginally statistically significant in column (1)).[12]

To assess threshold effects, we consider whether contemporaneous immigration flows have a larger negative association with domestic net migration when there is higher initial stock of immigrants—i.e., a greater cultural avoidance effect. We consider this possibility by adding an interaction variable between the initial 1990 foreign-born share and the 1993–2007 rate of international immigration and estimating the model using IV (not shown). Using the preferred dependent variable from Eq. (2), we

[11] The correlation between the 1990 foreign-born share and the 1993–2007 immigration rate is only 0.17—which is not particularly high, suggesting that multicollinearity is not behind the results. Also note that after including the 1990 foreign-born share, the immigrant flow variable remains well identified with the first-stage $F$-statistic on the identifying instruments equaling nearly 11.

[12] Differential total employment growth is also statistically insignificant when considering state-to-state flows. A possible reason is that when considering state-by-state migration flows, overall differences in *total* job growth are too noisy of a measure. On a state-by-state basis, migration flows may be more directly related to the particular industries that are faring well in each individual state due to work force composition, while in the aggregate, overall job growth averages out these individual state industry-composition effects. This is akin to how the current account balance between (say) Argentina and the United States is not necessarily reflective of the total US current account balance.

**Table 2** Net migration and population growth regressions

| | (1) Ols net migration w/o foreign-born share | (2) IV net migration w/o foreign-born share | (3) IV population growth w/o foreign-born share | (4) IV net migration with foreign-born share |
|---|---|---|---|---|
| Diff in immigration rate 1993–2007 | −4.03** (−2.69) | −5.54** (−2.40) | −7.7E-5 (−0.31) | −4.37* (−1.85) |
| Diff in % foreign born 1990 | | | | −1088.7** (−2.40) |
| Diff in employment growth 1993–2007 | 29027.3 ** (6.81) | 31425.0*** (3.02) | 3.65*** (3.20) | 13734.0 (1.10) |
| Diff in ocean border dummy | 1528.8 (1.62) | 2079.6* (1.95) | 0.15 (1.26) | 2051.4* (1.92) |
| Diff in amenity rank | −696.3 (−1.08) | −940.4 (−0.76) | 0.12 (0.88) | 2783.3 (1.42) |
| Diff in ln(1993 wage) | −4299.3 (−1.20) | −9952.6* (−1.92) | −0.13 (−0.22) | 17822.1 (1.41) |
| Adj./uncent. $R^2$ | 0.68 | 0.67 | 0.92 | 0.67 |
| Sample size | 49 | 49 | 49 | 49 |
| first-stage *F*-statistic diff immigration[a] | | 8.72 ($p = .0000$) | 8.72 ($p = .0000$) | 10.65 ($p = .0000$) |
| first-stage *F*-statistic diff employ. growth[b] | | 2.66 ($p = .036$) | 2.66 ($p = .036$) | 2.06 ($p = 0.091$) |
| first-stage *F*-statistic diff 1993 wage level[c] | | 12.04 ($p = .0000$) | 12.04 ($p = .0000$) | 1.51 ($p = .2095$) |
| Cragg-Donald Wald *F* statistic[d] | | 1.67 | 1.67 | 0.57 |

Robust *t* and *z* statistics are in parentheses. ***, **, and * indicate significant at 1, 5, and 10%, respectively

[a] The joint *F*-statistic for the five identifying instruments (borders Mexico; 1970 share that is either foreign born, or one foreign-born parent; 1980 foreign-born share, 1993 wage mix, 1993–2007 industry mix job growth) in the first-stage regression for differential 1993–2007 average annual immigration growth

[b] The joint *F*-statistic for the five identifying instruments (borders Mexico; 1970 share that is either foreign born, or one foreign-born parent; 1980 foreign-born share, 1993 wage mix, 1993–2007 industry mix job growth) in the first-stage regression for differential 1993–2007 employment growth

[c] The joint *F*-statistic for the five identifying instruments (borders Mexico; 1970 share that is either foreign born, or one foreign-born parent; 1980 foreign-born share, 1993 wage mix, 1993–2007 industry mix job growth) in the first-stage regression for differential 1993 wage level

[d] Cragg-Donald *F*-statistic for the strength of the instruments across all endogenous variables

find a negative and weakly statistically significant interaction coefficient ($t = -1.74$) (similar results follow using the dependent variable in Eq. (3)). Thus, there is evidence that domestic migrants are more influenced by contemporaneous immigration flows when there are higher initial stocks of immigrants.

We then split the sample period into two, 1993–2000 and 2000–2007, and reestimated the models shown in columns (1) and (2) of Table 3 (not shown for brevity). These results are less precisely estimated than when considering the entire

**Table 3** Net migration and in-migration regressions

| Variables | (1) IV: net migration state by state Eq. 2[a] | (2) IV: net migration state by state Eq. 3[b] | (3) IV: net migration state by state Eq. 2[a] | (4) IV: in-migration state by state[c] | (5) IV: in-migration state by state other than 8 high immig. states[c] |
|---|---|---|---|---|---|
| Diff in immigration rate 1993–2007 | 0.012 (0.55) | −0.008 (−0.71) | 0.009 (0.41) | −0.03 (−0.27) | 0.26 (1.42) |
| Diff in % foreign born 1990 | −14.17*** (−3.64) | −4.05** (−2.03) | −17.43*** (4.73) | −139.0*** (−7.24) | −272.7*** (−5.91) |
| Diff in employment growth 1993–2007 | 126.1 (1.07) | 43.8 (0.72) | 38.2 (0.33) | −2476.1*** (−4.30) | −5604.3*** (−4.84) |
| Diff in ocean border dummy | 36.8*** (3.91) | 12.9*** (2.66) | 35.2*** (3.68) | −127.8*** (−2.74) | −195.2*** (−2.78) |
| Diff in amenity rank | 27.4 (1.5) | 6.4 (0.68) | 35.1** (2.19) | 179.7** (2.06) | 696.8*** (3.64) |
| Diff in ln(1993 wage) | 99.6 (0.94) | 55.4 (1.02) | 229.9* (1.87) | 2497.7*** (4.61) | 6212.8*** (4.91) |
| Diff state employment growth relative to BEA region, 1993–2007[d] | | | −102.9 (1.25) | | |
| ln(origin pop93) | | | | 267.5*** (10.88) | 301.8*** (7.99) |
| Distance (km) | | | | −0.57*** (−23.3) | −0.67*** (−17.17) |

**Table 3** Continued

| Variables | (1) IV: net migration state by state Eq. 2[a] | (2) IV: net migration state by state by Eq. 3[b] | (3) IV: net migration state by state by Eq. 2[a] | (4) IV: in-migration state by state[c] | (5) IV: in-migration state by state other than 8 high immig. states[c] |
|---|---|---|---|---|---|
| Destination state fixed effects | N | N | N | Y | Y |
| $R^2$/cent. $R^2$ | 0.128 | 0.035 | 0.090 | 0.265 | 0.259 |
| Sample size | 1176 | 1176 | 1176 | 2352 | 1968 |
| Cragg-Donald Wald $F$ statistic[e] | 14.79 | 14.79 | 15.40 | 29.56 | 11.30 |

Robust $t$ and $z$ statistics are in parentheses. ***, ** and * indicate significant at 1, 5, and 10%, respectively

[a] The dependent variable defined as $[(M_{ij} - M_{ji})/((\text{Pop}_i + \text{Pop}_j)*0.5)]*1000000$

[b] The dependent variable defined as $[(M_{ij} - M_{ji})/(\text{Pop}_i * \text{Pop}_j)]*1000000*1000000]$

[c] The dependent variable defined as $M_{ij}/\text{POP}_j$ in Eq. 4

[d] The variable is defined as taking the (destination state's employment growth minus the destination state's BEA-region employment growth) minus the (origin state's employment growth minus the origin state's BEA-region employment growth)

[e] Cragg-Donald $F$-statistic for the strength of the instruments across all endogenous variables

time period, which as we described above, is expected given that the underlying data is noisier when using shorter time periods. The results suggest though that the initial immigration level had a stronger negative relationship with domestic migration in the post-2000 period. The other notable pattern is that employment growth differentials appear to have a stronger positive link to net migration in the earlier period, whereas natural amenities have a stronger positive influence in the latter period.

One other possible concern is there could be economic spillovers across labor markets. Such spillovers have not been considered in the past literature, perhaps because states are such large geographic units that these spillovers are likely to be less meaningful. Yet if spillovers are economically consequential, migration behavior would be primarily affected by employment growth differentials among neighboring states—e.g., relative job growth in the entire Great Lakes region may influence whether someone moves to Michigan versus another Midwest state. To account for this possibility, we added the relative BEA regional employment growth differential to the model: i.e., employment growth in the destination state minus the employment growth in its BEA region minus the corresponding figure for the origin state.[13] Using the preferred dependent variable from Eq. (2), the results shown in Column 3 of Table 3 show that the key results are essentially unaffected by including the relative regional employment growth variable, while the relative employment growth variable is statistically insignificant. Thus, we conclude that economic spillovers have an inconsequential effect on the results.

Column 4 of Table 3 reports the results when the dependent variable is in-migration to state $i$ from state $j$ using the dependent variable shown in Eq. 4.[14] This model more directly considers the types of states that domestic migrants are moving to. In this case, contemporaneous 1993–2007 immigration flows remain statistically insignificant in determining the in-migration rates on a state-by-state basis, which could be due to the state fixed effects accounting for migration flows in general. Likewise, the net immigration variable may be noisy on a state-by-state basis because it does not account for the type of immigrant flows that could affect domestic in-migration patterns. Yet the contemporaneous migration flow results are not a refutation of Borjas' model because

[13] We do not employ the spatial lag model to attempt to capture spillovers. Foremost, besides being inconsistent with the past immigration literature, the spatial lag model is inconsistent with standard migration theory. Following the example above, we are unaware of a theory where people choosing to migrate to Ohio would weigh whether other migrants are contemporaneously choosing to migrate to Michigan (which is the assumption of the spatial lag model). Rather such migrants would care about relative employment growth between the two states—i.e., economic migration theory is about relative economic opportunities in the two locales. Of course, migration rates between Ohio and Michigan would be highly correlated, but that is due to Ohio and Michigan having similar economic outcomes, industry structures, and climates, which is again not the same causal structure suggested by the spatial lag model. An alternative reason for using the spatial lag model is it supposedly helps account for omitted variable bias. Yet as pointed out by McMillen (2010), spatial lag models are typically estimated using maximum likelihood, which assumes the researcher perfectly knows the model, implying there are no omitted variables. Thus, using a spatial lag model to overcome omitted variable bias replaces one econometric problem with another. See Overman and Gibbons (2010) and Pinske and Slade (2010) for a related discussion of other identification problems when using spatial lag models without strong theoretical justification.

[14] Note that the Cragg-Donald $F$-statistic is nearly 30, again suggesting that the instruments are strong.

his aggregate *net*-migration model was not developed to describe migration between individual state pairs and it is not aimed at describing *gross* in-migration.

In contrast to the immigration flow results, the initial share of foreign born remains negative and highly statistically significant ($t = 7.24$). A one standard deviation increase in the percent foreign born is associated with a 0.77 standard deviation decrease in in-migration rates. Again, it appears that domestic residents are avoiding states that have high initial shares of foreign born, indicating other factors are at work rather than contemporaneous shifts in labor demand and supply. To be sure, this model accounts for destination state fixed effects to control for destination types and the model also accounts for the initial wage level, meaning that the initial foreign-born share influence does *not* occur through its possible composition effect on initial wage levels.

Because domestic migrants who are not locating in "gateway" immigration states may be especially sensitive to the share of immigrants, Column 5 reports the results from omitting in-migration rates into the eight high-immigrant "gateway" states of California, Arizona, New Mexico, Texas, Florida, New York, and New Jersey. These results show that the initial immigrant share coefficient is almost twice the magnitude in this regression, further suggesting that domestic migrants are avoiding states with initial immigrant shares, especially when choosing *not* to locate in a gateway state. Although these results may not be definitive, they are consistent with the need for economists to consider the effects of noneconomic factors such as cultural or ethnic avoidance in their models.

## 4 Conclusion

Using annual IRS state-to-state migration data, this paper estimated the domestic migrant response both to new immigrants and to the existing foreign-born share of the population. While analysis of aggregate migration flows suggest net domestic out-migrant responses to immigration flows occur, the result did not hold in analysis of state-to-state flows. Nevertheless, examination of the state-to-state migration flows revealed a negative net domestic migration response to the foreign-born share, suggesting domestic migrants may be motivated in part by ethnic or cultural avoidance. Further analysis revealed domestic in-migration as significantly negatively affected by the foreign-born share but not by recent immigration flows.

Overall, the results point to the need for economic models to include noneconomic factors in examining the nexus between domestic migration and immigration. Likewise, more attention should be given to the cumulative effects of past immigration and not just to the effects of current immigration flows. The dynamics of the relationship are further complicated by the increased rates of natural population growth in traditionally high-immigrant areas and the increasing share of previous immigrants among domestic migrants. Finally, more research should be conducted to see whether the origin of the immigrants matters in terms of the cultural avoidance. Does it matter whether the immigrants are from Mexico versus the Caribbean? More expansive models and micro-data appear to be needed to better capture these complexities.

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