Native Competition and Low-Skilled Immigrant Inflows*

Brian C. Cadena †
University of Colorado - Boulder

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Abstract

This paper demonstrates that immigration decisions depend on local labor market conditions by documenting the change in low-skilled immigrant inflows in response to supply increases among the US-born. Using pre-reform welfare participation rates as an instrument for changes in native labor supply, I find that immigrants competing with native entrants systematically prefer cities with smaller supply shocks. The extent of the response is substantial: for each native woman working due to reform, 0.5 fewer female immigrants enter the local labor force. These results provide direct evidence that international migration flows tend to equilibrate returns across US local labor markets.

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†E-mail: brian.cadena@colorado.edu

Nearly twenty percent of working-age adults living in the United States with no more than a high school education were born elsewhere.[1] These low-skilled immigrants come to the United States for a variety of reasons, but many choose to leave their home country in search of better labor market opportunities.[2] This paper investigates how geographic differences in expected earnings within the United States influence newly arriving low-skilled immigrants' location decisions. To do so, I develop a novel estimation strategy that leverages geographically disparate increases in native labor supply resulting from policy changes to the federal welfare system. Previous empirical treatments of this question have examined whether immigration flows are drawn disproportionately to areas with higher wage rates for their skill type.[3] There are, however, two primary challenges to interpreting results based on this approach, and they have proven difficult to overcome. First, measured geographic differences in wage rates among natives or previous migrants are unlikely to capture exogenous differences in the wages immigrants can expect. Instead, they may proxy for unobserved quality among existing workers, or they may reflect unobserved local public goods and other locational amenities (Roback 1982).[4] Second, immigrants are new labor market entrants who value both the probability of finding employment and the expected wage conditional on securing a job. Focusing exclusively on wages ignores other measures of differential labor market prospects, which are also likely to enter into a new immigrant's decision.

In this paper, I overcome both of these difficulties by exploiting differently sized labor market shocks that lead to within-city changes in labor market prospects, regardless of the relative importance of employment probabilities or wages. As a result of a series of policy

[1] Author's tabulations of the 2000 five percent PUMS.

[2] Economists have long considered migration as an investment driven by differences in expected earnings (Sjaastad 1962). Borjas (1987) first developed the hypothesis that US labor markets should be especially attractive to immigrants at the low end of the skill distribution from countries with larger returns to skill. Although empirical evidence suggests that immigrants are more likely to be drawn from the middle of the skill distribution in their home country, they nevertheless tend to fall in the bottom portion of the U.S. skill distribution (Chiquiar and Hanson 2005).

[3] Examples include Bartel (1989), Borjas (2001), Kaushal (2005), and Jaeger (2007).

[4] Also, see Shimer (2001) for a critique of the method employed in Borjas (2001).

changes to the former Aid to Families with Dependent Children (AFDC) program, the labor market participation rate among low-skilled native women in the target population increased by nearly fifteen percentage points. The women entering the labor market due to welfare reform tend to work in the same types of jobs as do newly arriving low-skilled immigrants. Thus, the increase in labor supply among natives can be viewed as an outward shift of the supply curve of a close substitute for immigrant labor.

To determine what influence these supply shocks have on immigrants' location choices, I use MSA-level welfare participation rates in 1990 as an instrument for changes in labor market participation among native workers from 1990 to 2000. Although many of the policy changes were implemented at the federal level, local increases in labor market attachment were strongly related to the share of the population affected by reform. I find that the geographic distribution of immigrants arriving in the 1990s was dramatically different than in the previous decade, with a systematic shift away from cities with more welfare leavers and toward cities with fewer leavers.

In order to interpret this result as the causal effect of native labor supply increases on immigrant inflows, two conditions must be satisfied: The share of a city's population affected by reform must be uncorrelated with other unobserved changes in the value of locating there, and changes to the welfare system must not affect immigrants' preferences over destinations through any mechanism other than the resulting increase in labor supply among natives. I present multiple pieces of empirical evidence suggesting that each of these conditions holds.

My empirical strategy uses long differences in an MSA's share of the immigrant population, which eliminates the influence of fixed amenities like local public goods. Thus, an alternative interpretation of the results must explain *changes* in the locations immigrants select. I consider and rule out several possibilities. First, the changes in immigrant share observed from 1990 to 2000 are not the continuation of pre-existing trends. Second, I control directly for additional factors that are likely to affect where immigrants choose to locate, including controls to address coincident demand shifts and a secular decline in the value of enclaves. Third, I control indirectly for multiple unobserved changes in a city's value by including the change in a city's share of higher-skilled immigrants as an additional covariate. In sensitivity analysis, I further allow for the possibility that cities offer amenities specific to immigrants from different source regions. Across all of these specifications, I continue to find strong evidence that immigrants avoided locating in cities with relatively larger increases in native labor supply.

Finally, I consider the possibility that welfare reform affected immigrants' choices directly, rather than through increased labor market competition. If immigrants were initially attracted to cities offering higher benefits, then reform could have "turned off" these welfare magnets, which provides a potential alternative interpretation of the results. I address this concern directly by allowing for differential inflow trends based on pre-reform benefit levels, changes in immigrant welfare receipt, or state-level policy choices that affected immigrants' eligibility. Consistent with Kaushal (2005), I find no evidence that immigrant inflows respond to changes in benefits, and controlling for this possibility does not qualitatively alter the main results.

After examining each of these alternatives, I present back of the envelope calculations to determine the extent to which differential immigrant inflows offset the number of new native workers due to welfare reform. The results suggest that immigration inflows substantially equalized the reform-induced supply shifts across geography. For each additional native woman working in a labor market due to welfare reform, 0.5 fewer female immigrants choose to enter the local labor force.

The finding that local low-skilled immigration inflows are quite sensitive to labor market conditions has important implications across several strands of the literature. This paper contributes most directly to the literature concerned with the extent to which migration flows mitigate geographic labor market inequality as selective location decisions among international migrants will tend to smooth out local shocks.[5] This dynamic becomes particularly important given that minority workers and the less-educated have significantly lower mobility rates, and are thus disproportionately affected by local demand shocks (Bound and Holzer 2000). As suggested by the framework presented by Borjas (2001), therefore, the ability of earnings-sensitive immigrants to diffuse local shocks throughout the country provides an often overlooked benefit of consistently large inflows of low-skilled immigrants.

Additionally, a large literature has relied on geographic variation in the settlement pattern of immigrants to determine the effect of immigration on native labor market outcomes.[6] On the whole, these studies reveal very similar changes in wage and employment outcomes for native workers, regardless of the extent to which immigration altered the skill mix of a local market (Smith and Edmonston 1997). These results contrast with studies that treat the labor market as nationally integrated, which find much more substantial wage effects (Borjas, Freeman and Katz 1997, Borjas 2003). Although several mechanisms have been proposed to explain these different results, none has found empirical support.[7] The strong response of immigrant inflows to exogenous changes in expected earnings that I identify in this paper suggests that this mechanism may help explain this discrepancy.[8]

Finally, many studies use cross-geography comparisons to evaluate the effect of labor market policies. The results of this paper imply that these types of research designs will tend to underestimate the effect of any policy susceptible to arbitrage by highly mobile

[5]Sjaastad (1962) first proposed that earnings-motivated internal migration decisions may create this positive externality. Topel (1986) provided a formal treatment of the relationship between mobility costs and the persistence of geographic inequality. Barro and Sala-I-Martin (1991) and Blanchard and Katz (1992) provide additional evidence that internal migration tends to diffuse local labor market shocks.

[6]Card's (1990) influential paper found virtually no effect of the Mariel Boatlift on Miami's labor market. Additional examples of this general methodology include Altonji and Card (1991), Lalonde and Topel (1991) and Schoeni (1997).

[7]The literature has examined possibilities including immigrant inflows leading to native outflows (Card and DiNardo 2000, Card 2001) and local economies adjusting to their changing endowments by altering the mix of products they produce (Lewis 2003).

[8]Using the results of this type of study to more precisely quantify the implied bias in a wage regression requires additional assumptions, and I provide a more complete discussion of this possibility in Section 4.5 and in the appendix.

low-skilled immigrants.

The remainder of the paper is organized as follows: the next section provides a conceptual framework for evaluating the role of labor market conditions in immigrants' location decisions and presents descriptive evidence detailing welfare reform's disparate geographic effect on native labor supply; Section 3 presents a discrete choice model and motivates the appropriate empirical methodology for estimation; Section 4 provides the main empirical results and additional robustness checks; the final section further discusses the implications of these findings for previous research and for future policy decisions.

2 Welfare Reform and Native Labor Supply

Over the 1990s, several policy changes were implemented, each of which was designed to increase labor market participation among native women previously eligible for cash welfare benefits. I begin by discussing these changes and how their combined effect created incentives for many low-skilled native women to enter or remain in the workforce. I then provide empirical evidence that these reforms succeeded in substantially increasing the labor market attachment of the target population and that local welfare participation rates prior to reform reliably predict local increases in native female employment over the reform period.

The federal cash welfare system, first implemented in 1935, was originally designed to provide for the material needs of widows with dependent children. By the late 1970s, the demographic makeup of the welfare rolls had changed dramatically: Widows were covered by social security and rising rates of divorce and non-marital childbearing meant that most recipients were in families headed by divorced and never-married mothers. As greater numbers of married women worked, there was political pressure to increase employment among mothers on welfare. In the early 1990s, states were given expanded authority to secure federal waivers from AFDC program rules and in 1996 the Personal Responsibility and Work

Opportunity Reconciliation Act (PRWORA) ended the AFDC program.

After reform, cash assistance was no longer a federal entitlement program; seeking or participating in employment became a pre-condition for benefit receipt. Welfare offices implemented work support and "work first" programs to move women into the workforce. Welfare recipients are now subject to a sixty month lifetime limit (fewer at state discretion), giving potential recipients an incentive to delay benefit receipt and search more intensely for employment opportunities before applying.[9] Most states also reduced the rate at which benefits are taxed away as a recipient earns income (Blank and Matsudaira 2008), and concurrent expansions in the federal Earned Income Tax Credit provided additional financial incentives to work. Each of these policies was designed to increase employment and raise the return to work among women who, in the absence of reform, might have relied mainly on public assistance (Ellwood 2000). These reforms therefore created the textbook definition of a labor supply shift - policy changes resulting in more low-skilled women willing to work at any given wage.

The empirical literature evaluating welfare reform supports the conclusion that these reforms increased employment among the target population. While any credible study of the effect of these changes includes essential controls for the role of the strong macroeconomy over the period in which reform was implemented, most studies (and especially evaluations of demonstration projects using random assignment) find that the policy changes had a significant effect on the labor force attachment of low-skilled women.[10] Figure 1 uses data from the annual March Supplement to the Current Population Survey (CPS) and shows trends in labor market attachment, working positive weeks during the previous year, for women with at most a high school degree between the ages of eighteen and fifty-four living in large

[9] Grogger, Haider and Klerman (2003) find that as much as half of the decline in caseloads resulted from a decrease in the entry rate.

[10] For a careful review of the employment effects of welfare reform, see Blank (2002, pp.1139-1142).

Metropolitan Statistical Areas (MSAs).[11] They are classified according to marital status and parenthood. Labor market participation among single mothers increased dramatically from 70 percent in 1993 to 84 percent in 2000. There was no similar increase among women in either of two comparison groups - married mothers or single women without children. Additionally, the trends become quite similar after 2000, suggesting that the reforms created a roughly permanent increase in supply.

Even though many of the policy changes were implemented at the national level, local labor markets where welfare recipients represented a greater fraction of potential low-skilled workers experienced larger increases in low-skilled labor supply. As evidence of this relationship, consider Figure 2, constructed with the same employment measure as in Figure 1. I first rank each MSA based on the fraction of all low-skilled women receiving cash welfare from 1988 to 1992, and I then calculate annual employment averages for women living in MSAs in the top and bottom quartiles.[12] Although the levels are different, the time pattern of employment in both quartiles is quite similar prior to the mid-1990s. After that point, however, employment increased significantly for women in high participation cities. Employment among women in low participation cities, in contrast, remained roughly flat. By the end of the decade, the employment gap (on average eight to nine percentage points prior to reform) had essentially disappeared. Thus, the level of welfare participation prior to reform reliably predicts increases in native labor supply over the reform period, with overall dynamics consistent with the timing of reform.

How then, should these differential supply shifts affect the expected labor market returns

[11] These data were obtained from the IPUMS-CPS project (King, Ruggles, Alexander, Flood, Genadek, Schroeder, Trampe and Vick 2010). As closely as possible, this set of cities matches the consistent geographic areas described in Section 4.1. The graphs end at 2003 because a subsequent change to MSA definitions makes it difficult to construct a consistent sample after that year.

[12] This figure takes the average over all women living in any city within a quartile. Each quartile contains 33 MSAs, and captures a roughly equal fraction of the sample population. A similar, though somewhat noisier pattern emerges when each MSA average contributes one observation to the quartile mean. The cutoffs for the bottom and top quintiles are around 5 and 10 percentage points, respectively.

of newly arriving immigrants? In the short run, requiring women to work or search for work will increase the competition for each vacancy, making it less likely that any new entrant will find employment in a given period. Eventually, the market will reach a new medium-run equilibrium with more employed low-skilled workers working at lower wages.[13] Welfare reform, therefore, provides exactly the type of shock that substantially alters the labor market returns of newly arriving immigrants.

Of course, immigrants are only likely to respond to these supply increases if welfare leavers represent a meaningful increase in competition for jobs that immigrants are likely to search for. Several descriptive facts suggest that this is the case. First, the population affected by welfare reform and the flow of new immigrants are of similar magnitudes. Welfare rolls fell by 2.3 million adults (from 3.8 to 1.5 million) between 1990 and 2000 (US Department of Health and Human Services 2007).[14] Over that same time period 4.4 million low-skilled immigrants (male and female) entered the country (US Bureau of the Census 2007).

In addition, newly arriving immigrants and welfare leavers tend to work in similar jobs. Altonji and Card (1991) provide a means of calculating the degree to which an influx of one type of worker creates an increase in competition for various other groups of workers. Specifically, the measure calculates the effective increase in supply experienced by each type of worker due to entry by a single type that results in a one percent increase in the overall labor force.[15] Table 1 presents this measure comparing single native women who are currently working but received welfare in the past year to eight other groups of workers based on

[13]If immigration flows fail to offset the differential impact of welfare reform, one might expect that, in the long run, owners of capital will bring the market back into geographic equilibrium by directing investment toward areas with a higher return. If, on the other hand, newly arriving immigrants are sufficiently sensitive to near-term differences in expected earnings, they may manage to smooth out these disparate labor market shocks even without any change in behavior among native workers or capital owners.

[14]Note that these figures may understate the number of women affected by reform as the caseload reached a peak of 4.4 million adults in 1994.

[15] The exact formula is $\sum_j \frac{S_{gj} S_{Lj}}{S_j}$, where S_{gj} is the share of workers in group g in industry or occupation j, S_{Lj} is the share of welfare leavers in j, and S_j is the share of all workers in j.

gender, education and nativity.[16] These results suggest that women entering the labor force in response to welfare reform are especially close substitutes for newly arriving female immigrants. The index for occupation is 2.63 while the index for industry is 1.94, and I therefore focus the majority of the location choice analysis on this group. Low-skilled immigrant men do overlap substantially with welfare leavers (occupation and industry values of 1.22 and 1.15 respectively), however, and I present results examining their location choices for completeness.

On the whole, the descriptive evidence presented in this section suggests that welfare leavers represent a substantial exogenous increase in labor market competition for newly arriving low-skilled immigrants. Further, geographic variation in pre-reform welfare participation induced substantial differences in the degree to which local labor markets were affected. These facts lead to a clear prediction: if newly arriving immigrants are sufficiently sensitive to geographic variation in the expected earnings a labor market offers, they will tend to select locations with fewer women entering the labor market as a result of welfare reform. The next section provides an empirical framework for evaluating this hypothesis.

3 Expected Earnings and Location Choice

Suppose that each metropolitan area in the United States offers an immigrant a level of utility from settling there: U_{isdt}. In this notation, i indexes individuals, s indexes the source region, d denotes destinations (e.g. MSAs) and t indexes time periods (e.g. census decades).[17] The immigrant's decision rule can be expressed in a straightforward way: she chooses to move to

[16] These data come primarily from the 2000 census, although I calculate the shares for welfare leavers using both the 1990 and 2000 censuses to account for the fact that reform may have affected the types of jobs observed among those with benefit receipt. I classify both occupations and industries at the most disaggregate level possible.

[17] Because the data I will use do not provide information on potential immigrants who choose not to migrate, I model only the decision of where to locate conditional on deciding to move to the US. This simplification implicitly assumes that the relative utility of cities within the US is independent of the value of remaining at home.

location j if and only if $U_{isjt} > U_{isdt} \quad \forall d \neq j$.

The central question is the extent to which relative labor market prospects affect the relative utility that each location offers, and thus the likelihood that an immigrant selects a particular location. To begin, I define the total labor market returns as the present discounted value of the stream of expected future earnings that a location offers. For concreteness, suppose that each worker inelastically supplies one unit of labor when employed. Then the expected earnings k periods after migrating depend on the probability of being employed (p^k) and the expected wage conditional on being employed: $\mathbf{E}[w^k|p^k = 1]$. The total value of the stream of future income an immigrant with discount factor $\delta < 1$ expects to earn if she moves to location d in year t is thus:[18]

$$PDV_{dt} = \sum_{k=0}^{T} \delta^k (p_{dt}^k \cdot \mathbf{E}[w_{dt}^k | p_{dt}^k = 1]). \tag{1}$$

Notice that a location's wage rate is only one component of the total labor market returns it offers (Topel 1986). Several authors have used the wages of similarly-skilled workers as a measure of immigrants' expected earnings; yet very little attention has been paid to the probability of finding and maintaining employment (c.f. Bartel 1989, Borjas 2001, Kaushal 2005, Jaeger 2007). This frequent simplification is likely driven by data availability. In the absence of reliable data on the probability of finding employment or the length of initial unemployment spells, replacing expected earnings with prevailing wage rates may provide a reasonable approximation.

Yet, there are two key reasons why focusing exclusively on wages is likely to mismeasure the actual returns as perceived by new immigrants. First, immigrants are, by definition, new entrants into a labor market. Although some may move to a city having already secured employment, most will begin their search upon arrival. Additionally, many immigrants

[18] By omitting the subscript s, I am assuming that similarly-skilled workers from different source countries are perfect substitutes for each other.

(especially those who are unauthorized) enter the country with a relatively short expected time horizon (i.e. low T).[19]

Despite these concerns, the correlation between wage rates and expected labor market returns is likely still positive. The above discussion suggests, however, that variation in labor market tightness may reduce the strength of this relationship, especially if wages and employment probabilities are negatively correlated. Using native supply increases rather than changes in measured wages ensures a clear prediction of how the location decisions of expected earnings-maximizing immigrants should respond.

One remaining concern is whether immigrants are able to gather sufficient information about labor market prospects in order to make an optimal location decision. McLaren (2006) provides time-series evidence that border apprehensions are a reliable leading indicator of US economic growth, suggesting that unauthorized immigrants have access to information about the labor market and do not undertake the risky venture of crossing the border unless they are reasonably confident that they will find work. Selection across geography requires only a minor extension to this finding where potential immigrants have network contacts in more than one city, each providing this type of information. Should that fail, there is also room for trial and error. Once immigrants have paid the fixed costs of moving to the US, they face substantially smaller marginal costs of acquiring more information on where it is easier to find work in addition to lower costs of actually moving again. Of course, many low-skilled legal immigrants enter on family re-unification visas, and most of these likely never consider a location other than where their family members live. Yet immigration flows can still help equilibrate local labor markets as long as there is a sufficiently large group of sufficiently mobile immigrants. In fact, one might expect unauthorized migrants to be especially earnings-sensitive given their willingness to risk apprehension in order to gain

[19]Munshi (2003) details the importance of network connections in facilitating these searches as well as the cyclical nature of migration, wherein the duration of most U.S. stays is relatively short.

access to US labor markets, and I address this hypothesis in the empirical analysis.

3.1 Empirical Specification

In this subsection, I provide additional structure and derive an appropriate empirical specification to evaluate the influence of labor market returns on location decisions. Suppose that the overall utility a city provides can be expressed as a linear function of its expected labor market returns, other observable characteristics X_{sdt}, and an individual-specific unobserved error term. Then the total utility of a location is given by

$$U_{isdt} = \gamma PDV_{dt} + X_{sdt}\beta + u_{isdt}. \tag{2}$$

McFadden (1974) demonstrates that if each u_{isdt} is independently and identically distributed Type I extreme value, γ and β can be estimated through conditional logit models using individual-level data. This approach is commonly adopted by other authors in studies of the location choices of new immigrants (c.f. Bartel 1989, Kaushal 2005, Jaeger 2007). Yet the required assumption almost surely fails. In particular, there are most likely unobserved city attributes that have similar value to all immigrants from the same source region ($u_{isdt} = \eta_{sdt} + \epsilon_{isdt}$). These common error components present two challenges to estimation. First, assuming i.i.d. errors in the presence of these grouped unobserved components will vastly understate the standard errors and lead to incorrect inference. More importantly, these unobserved factors are likely correlated with expected earnings.

My empirical approach improves upon previous studies by explicitly modeling these unobserved components of the error term and taking steps to remove their influence. I employ an estimation strategy new to the immigration literature based on a method previously developed to examine workers' choices among health insurance options (Scanlon, Chernew, McLaughlin and Solon 2002), and my exposition of the econometric model closely follows the original. I begin by deriving an expression relating the observed share of immigrants selecting a particular city to the observed and unobserved components of utility in any given time period. This approach acts as a non-linear analogue to the "group-level regression" solution to the Moulton problem of common error components (Moulton 1990).

Allowing for common unobserved city attributes yields a new representation of the utility offered by a city:

$$U_{isdt} = \gamma PDV_{dt} + X_{sdt}\beta + \eta_{sdt} + \epsilon_{isdt}. \tag{3}$$

Note that this general framework nests the possibility that $\eta_{sdt} = \eta_{dt} \quad \forall s$, i.e. that the unobserved city attributes have similar value to immigrants from all source regions. I estimate models under both assumptions, but I use the most general form for exposition. Rather than assuming that the u terms are i.i.d., I make the much less restrictive assumption that the ϵ_{isdt} terms are distributed i.i.d. Type I extreme value. In other words, conditional on the observed attributes *and* any fixed amenities, the remaining individual-level errors are well-behaved. Given this assumption, the probability that an immigrant selects a given destination in time period t is

$$\pi_{sdt} = \frac{e^{\gamma PDV_{dt} + X_{sdt}\beta + \eta_{sdt}}}{D_{st}} \tag{4}$$

with $D_{st} = \sum_j e^{\gamma PDV_{jt} + X_{sjt}\beta + \eta_{sjt}}$.

This expression closely parallels the probability arising in a conditional logit model with the addition of the unobserved group effects in both the numerator and denominator. In expectation, the share of newly arriving immigrants who select each destination will be equal to these choice probabilities. In practice, the observed shares will differ from the actual choice probabilities due to random sampling error. Let S_{sdt} represent the observed share of immigrants from source s selection location d in year t. Then

$$S_{sdt} = \frac{e^{\gamma PDV_{dt} + X_{sdt}\beta + \eta_{sdt}}}{D_{st}} + \nu_{sdt} \tag{5}$$

Here ν_{sdt} is a mean-zero error term with variance that is inversely proportional to the number of observations within an st cell. Taking logs of both sides yields

$$\ln(S_{sdt}) = \ln(e^{\gamma PDV_{dt} + X_{sdt}\beta + \eta_{sdt}} + D_{st}\nu_{sdt}) - \ln(D_{st}). \tag{6}$$

Taking a first-order Taylor Series approximation around $\nu_{sdt} = 0$ gives

$$\ln(S_{sdt}) \approx \gamma PDV_{dt} + X_{sdt}\beta - \ln(D_{st}) + \eta_{sdt} + \frac{\nu_{sdt}}{\pi_{sdt}}. \tag{7}$$

An appropriately transformed version of the share of immigrants selecting a city will thus be approximately linear in the observed and unobserved attributes, and a regression with $s \times d$ cells as observations will have much better inference properties than will an individual-level conditional logit.

A cross-sectional version of this model, however, is unlikely to identify a causal relationship as any reasonable proxy for expected earnings is likely correlated with the error term. High wage areas likely offer better amenities, and one cannot use measures of competition for jobs, such as native participation rates, as these are endogenous to immigrant inflows. Time differencing provides a partial solution by removing the influence of any amenities that are fixed over time. The grouped error components can be partitioned into factors fixed over time ϕ_{sd} and factors specific to each time period ψ_{sdt}, i.e. $\eta_{sdt} = \phi_{sd} + \psi_{sdt}$.

The differenced specification is therefore:

$$\Delta \ln(S_{sd}) \approx \gamma \Delta PDV_d + (\Delta X_{sd})\beta - \Delta \ln(D_s) + \Delta\psi_{sd} + \Delta\frac{\nu_{sd}}{\pi_{sd}} \tag{8}$$

I estimate a version of Equation (8) by instrumental variables, using changes in native female labor force participation as a proxy for changes in total expected labor market returns and the welfare participation rate prior to reform as the excluded instrument. I use labor market participation (working positive weeks over the prior year) because it most closely captures the size of the exogenous supply shift, i.e. the degree to which natives began competing for low-skilled jobs in a local market. I avoid using alternatives such as wages or instantaneous employment rates, as these are equilibrium outcomes that depend on both the number of natives and the number of new immigrants who enter the labor market. I begin with specifications using shares measured across all immigrants, and, in sensitivity analysis, I include source-specific intercepts to account for the $\Delta \ln(D_s)$ terms.

Using this empirical strategy, I cannot estimate the effect of attributes of a destination or source-destination pair that are fixed over time, including factors commonly considered such as distance and climate similarities. Additionally, parameter estimates for attributes with little variation over time are not well-identified. Many other covariates used routinely in the literature fall into this latter category, including the location of previously-arriving immigrants and the geographic distribution of potential network contacts. The inability to include these variables should not be considered a limitation of the model. Instead, this approach removes the influence of any observed or unobserved aspect of a destination or source-destination pair that is roughly constant across time.

I have motivated this estimation procedure as the appropriate methodology under the assumptions of a particular discrete choice model. It is worth noting, however, that previous work has used a similar reduced-form specification without any structural derivation. Borjas (2001) used the ratio of the share of newly arriving immigrants to the share of previously arriving immigrants as the dependent variable in his analysis of whether new immigrants respond to state differences in wages. Both his dependent variable and the one suggested by the discrete choice model roughly represent proportional differences in a location's immigrant share. Thus, even if the assumptions underlying this precise discrete choice model are violated, this specification has an intuitive reduced form interpretation and continues to improve on previous work that used wage levels to proxy for expected earnings.

4 Data and Results

4.1 Data

The five percent Public Use Microdata Samples of the 1980-2000 decennial censuses provide the majority of the data for the analysis. I consider the location of newly arriving adult immigrants ages 18-54, with at most a high school degree, not living in group quarters. I classify a respondent as an immigrant if he/she is foreign-born and is either a non-citizen or a naturalized citizen.[20] New immigrants are those who arrived in the US during the ten years prior to survey.[21] I restrict the analysis to immigrants from the eleven source regions listed in Table 2.[22] This table shows the distribution of sources across all three waves of the census. This distribution has remained somewhat stable over the sample period with two exceptions: immigration from Mexico increased, while immigration from European countries decreased.[23]

Table 3 provides some basic descriptive statistics for this population. In each census year, the total number of new immigrants is split almost evenly between women and men. Most new immigrants are married and very few live alone as household heads. These values are

[20] I obtained the data from the IPUMS project (Ruggles, Alexander, Genadek, Goeken, Schroeder and Sobek 2010), and I use the IPUMS coding of educational attainment that is designed to be consistent across census surveys.

[21] These immigrants may have previously lived in the United States, but the census question asks when the respondent arrived in the US "to stay".

[22] This restriction eliminates less than three percent of the sample of new immigrants.

[23] This pattern highlights the potential importations of accounting for source-specific amenities as the changing composition of immigrant inflows may have altered the geographical distribution even in the absence of any local labor market shocks. Empirically, accounting for this possibility only slightly alters the results.

quite similar across the different waves of the census, suggesting that changes in the locations these immigrants choose are unlikely due to household composition changes.

I consider all MSAs within the continental US with a nonzero immigrant population in all three census years and an adult population (18-54) of at least 150,000 in 1990 as potential locations for newly-arriving immigrants. These selection criteria result in 157 destinations in each year.[24] For the basic results, I treat the η_{sdt} terms as constant across all source regions. The dependent variable in these specifications is the natural logarithm of the share of all new immigrants living in each city, calculated separately for each census decade. I use person-level weights to calculate these shares, which I calculate separately by gender.

The primary explanatory variable of interest is the change in native female labor market participation, which I quantify using the fraction of all women working positive weeks over the past year.[25] The excluded instrument in the IV specifications is the welfare participation rate in 1990: the fraction of all women who received positive welfare benefits during the year prior to the survey. I also construct variables to measure a number of additional attributes (listed in Table 4) that immigrants may consider when deciding where to locate. I include information from external data sources, as well as other variables directly calculated from the PUMS. I discuss the data sources for covariates from non-census sources as they are introduced to the location choice models.

4.2 Immigrants Avoid Larger Supply Increases

Figure 3 displays the first-stage and reduced form results of an instrumental variables version of Equation 8. Each city contributes one equally weighted observation. The left panel plots

[24]The geographic boundaries of the MSAs change somewhat across waves of the census. I follow Card and Lewis (2007) and use state and county group codes to create consistent areas across the three census years. Ethan Lewis graciously provided programs to do so.

[25]I select this definition to remove the influence of temporary variations in local labor market conditions and to capture the long-run change in the share of the population attached to the labor market. I obtain qualitatively similar results using changes in the current labor force participation rate.

the data used to fit the first stage regression, along with the fitted values. As hypothesized, cities with higher welfare participation prior to reform experienced greater increases in native female labor market attachment over the decade. Each percentage point increase in program participation prior to reform led to a 0.44 percentage point increase in employment. The second panel shows the reduced form, and provides the central finding: Female immigrants arriving during the 1990s were less likely to select cities with larger native populations entering the workforce as a result of welfare reform, as measured relative to the choices of immigrants arriving over the 1980s. Figures 4 and 5 show this relationship geographically. These maps demonstrate that the relationship is not driven by any particular region; instead, the pattern holds broadly across the entire country.[26] In each figure, darker areas represent MSAs with values above than the median, and lighter areas represent areas with values below the median. Areas of the country not included in large MSAs are represented as white. The negative relationship is apparent when looking from map to map as cities turn from light to dark and vice versa.

The parameter estimates from this specification are given in the second column of Table 5. As expected given the figures, the first-stage is strongly significant (the F-statistic on the excluded instrument is well in excess of 10), and the resulting IV estimate is significantly negative. To contrast the IV results, the first column of the table shows the results from estimating this same equation without an instrument. This coefficient is substantially more positive, consistent with an omitted variable such as local demand shocks increasing native employment and attracting newly arriving immigrants.

Interpreting the sign and statistical significance of these coefficients is straightforward. The magnitude can be interpreted as roughly the percentage change in the probability that

[26] These maps appear to suggest that California is especially important in this analysis, but this impression is largely due to the increased visual weight suggested by the larger land area of California MSAs. In results not reported here, I have run additional specifications including fixed effects for each census region and/or excluding California and the results are qualitatively unchanged. All results discussed in notes and not presented in tables are available from the author upon request.

an immigrant selects a given city.[27] The coefficient in column 2 thus implies that a city experiencing a one-percentage point larger than average welfare-reform-induced increase in native female labor supply saw roughly a seventeen percent decrease in the probability that a female immigrant chose to locate there. This difference in supply increases would result from slightly less than a one standard deviation difference in pre-reform program participation.[28]

As an initial falsification test, Figure 6 presents analogous results using data from one decade prior. Importantly, neither the first stage nor the reduced form relationship holds in a time period without a dramatic change to welfare policy.[29] The lack of a first-stage relationship over this period rules out certain alternative interpretations of the employment increases over the 1990s. For example, suppose that high welfare participation were indicative of poor labor market conditions and that the subsequent increases in labor supply were the result of negatively serially correlated shocks. The first-stage results over the 1980s provide no support for this hypothesis. Similarly, the lack of a reduced form relationship rules out the possibility of pre-existing trends away from cities with high welfare participation.[30] This pair of results strengthens the credibility of interpreting the relationships shown in Figure 3 as resulting from immigrants avoiding labor market competition with welfare leavers.

[27] The percentage change in the choice probability resulting from a one unit change in the independent variable is $\frac{e^{\beta \Delta X} - 1}{\Delta X}$, or approximately β for small changes in X. This interpretation provides the reduced form interpretation. Based on the discrete choice model, the change in the odds that a city is selected resulting from a one unit change in X is $p(1-p)\beta$. With small probabilities (the mean is 1/157), the difference between these two interpretations is minimal.

[28] The standard deviation is 3.05 percentage points, and the first stage coefficient implies that each percentage point difference in pre-reform participation leads to a 0.44 percentage point increase in employment.

[29] Both point estimates are slightly negative, and neither is statistically significant.

[30] The welfare participation rates are remarkably stable across cities from 1980 to 1990. Additional analysis (not shown, but available upon request) using the difference between the 1990-2000 change in share and the 1980-1990 change in share as the dependent variable also shows a statistically significant break in trends in favor of cities with fewer welfare leavers in the 1990s.

4.3 Results Robust to Additional Controls

The remainder of Table 5 adds additional control variables to address alternative interpretations of the basic set of results. First, suppose that local labor demand shocks over the 1990s were negatively correlated with welfare participation rates at the start of the decade, even though this appears not to have been the case in the 1980s. Under this scenario, these cities may have lost immigrant share even if immigrants did not react to the increases in native labor supply. Column 3 includes the change in the decade average annual employment growth rate as a means of controlling for this potentially omitted factor.[31] This variable enters the model with the expected sign: the distribution of immigrants shifted away from cities with slowing employment growth and toward cities with improving growth. The parameter estimate for native female supply is not substantially affected, however, suggesting that the supply shocks created by welfare reform were reasonably uncorrelated with the size of any concurrent demand shocks.

Alternatively, suppose that high welfare participation cities also tended to be traditional locations for immigrants. If traditional locations became less popular for reasons unrelated to welfare reform then these cities would have lost immigrant share even in the absence of the policy-driven labor supply increases. The specification in Column 4 addresses this possibility. An immigrant arriving in the 1990s faced a very similar geographic distribution of previously arriving immigrants as did an immigrant arriving in the 1980s. In this differenced specification, therefore, including the fraction of a city that was foreign-born in 1990 as a covariate allows for differential inflow trends based on a city's status as a traditional destination.[32] The negative coefficient estimate is consistent with a diffusion of immigrants

[31] These data come from the County Business Patterns, and I aggregate to the same consistent geographic boundaries used for variables based on the PUMS. The variable is calculated as the average annual growth rate for the counties comprising each MSA.

[32] A specification that includes the change in this variable from 1980 to 1990 yields an imprecisely estimated zero, and including this variable does not qualitatively alter any other results.

across the country, with traditional destinations tending to lose immigrant share. Yet the coefficient on native labor supply remains negative and significant. In fact, even though the point estimate decreases in magnitude, the inclusion of this variable increases the precision of the estimates substantially.

One may also be concerned that changes in other labor market policies affected the desirability of choosing each location. In particular, differential increases in the minimum wage may have affected the returns to entering each local labor market (Cadena 2011). The specification in column 5 includes a measure of minimum wage policy as an additional control to address this concern. I begin by calculating, for each MSA in each month, the percentage change in the minimum wage relative to January at the start of the decade. I then average these changes over each decade and calculate decadal changes. Thus, this variable measures the extent to which an MSA's minimum wage policy became more or less active from the 1980s to the 1990s.[33] The inclusion of this variable has a negligible influence on the other coefficients, suggesting that differences in minimum wage policy are not confounding the analysis.

The time differencing strategy effectively removes the influence of any unobserved city amenities that are fixed over time, and the specifications in Columns 1-5 control directly for multiple time-varying reasons that immigrants selected new locations over the 1990s. Yet there may still be changes in unobserved city-level characteristics that are correlated with the reform-induced supply increases. One way to address this potential source of bias is to include a city's change in immigrant share among a group whose expected earnings should be relatively less affected by welfare reform. To accomplish this, I include the change in the city's share of female immigrants with at least some college education. This variable enters

[33] Because the minimum wage is a state-level policy, I initially calculated this measure for each state. For MSAs that cross state lines, I calculated a population-weighted average of the state-level measure. I use a monthly state panel with exact dates of policy implementation (initially compiled for another of my papers) to construct this measure.

with a positive sign and strong significance (column 6), but the coefficient of interest remains strongly negative.

The final column addresses the so-called "welfare magnets" hypothesis, which provides an alternative mechanism through which welfare reform may have affected immigrants' preferences over cities. Previous research contends that states with more generous welfare benefits attract larger inflows of eligible immigrants (Borjas 1999). Suppose that welfare reform essentially "turned off" these magnets, and, as a result, cities in generous states were no longer especially attractive to immigrants. In fact, immigrants arriving after reform were required to wait five years until receiving benefits, and unauthorized immigrants were barred altogether. To address this possibility, I include both the maximum benefit level for a family of three in 1990, which allows for initially generous cities to become less popular as reform equalizes potential benefits across locations, and a direct measure of the change in the share of low-skilled immigrants receiving welfare benefits. The positive coefficient on the maximum benefit variable and the statistically insignificant coefficient on the change in benefit receipt fail to provide support for this alternative interpretation. I also include dummy variables for whether a state restored each of four programs to post-reform legal immigrants using its own funds.[34] The resulting coefficients are variable and mostly insignificant, and they are, on the whole, consistent with previous work finding no direct effect of policy reforms on immigrants' location choices (Kaushal 2005).

4.4 Robustness and Response Heterogeneity

The set of results above are consistent with the central hypothesis that newly arriving immigrants tended to avoid locations with more welfare leavers, and that they did so to avoid the increase in labor market competition that those supply increases represented. The remainder

[34] These data come from Zimmermann and Tumlin (1999), and I construct population-weighted averages for MSAs crossing state lines.

of the results present robustness checks and address additional hypotheses suggested by this interpretation of these findings.

I first relax the restriction that unobserved locational amenities are equally valuable to all immigrants, with results reported in Table 6. As discussed in Section 3, these attributes are likely different depending on the immigrant's source region. A changing mix of immigrant sources could lead to a different distribution of settlement patterns, even in the absence of immigrants responding to labor market incentives. This set of regressions addresses that concern by explicitly allowing the unobserved city attributes to vary for each source region and changing the dependent variable to source-specific immigrant shares.[35] Each city may have as many as eleven observations, one for each source region identified in Table 2.[36] In columns (4) through (6), I replace the generic immigrant concentration variable from Table 5 with an analogous measure of whether the city was a traditional location for immigrants from the specific source region. The results through all specifications are quite similar to the main results in Table 5. Although somewhat smaller in magnitude and less precisely estimated, these specifications continue to support the conclusion that immigrants chose cities with smaller native supply increases.

The discussion in Section 3 suggested that immigrants arriving without legal authorization are likely to be the most earnings-sensitive. Although the census data do not include an immigrant's visa status, examining the heterogeneity in responsiveness across source regions presents an opportunity to address this question. Table 7 lists the coefficient and standard error from running the specification in column (6) of Table 6 separately for each source popu-

[35]Because the share of immigrants selecting each city will be more precisely estimated for regions with more observations, I weight each source-destination pair by the square root of the total number of observations from each source country. I report standard errors clustered by MSA in all specifications with multiple observations per city.

[36]I eliminate from the entire panel any source-destination pair that contains no immigrants in any of the census years. In order to include the change in high-skilled immigrant share as a control, these shares must also be non-empty. These restrictions explain why the total number of observations is less than $11 \times 156 = 1716$.

lation. The estimates are somewhat noisy, which limits the degree to which these differences allow for sharp conclusions. Nevertheless, the pattern of the point estimates suggests that the most earnings-sensitive migrants are the most likely to be unauthorized (i.e. those from Mexico and Central America).

Table 8 addresses an additional hypothesis implied by the timing of the supply shocks. Although some of the reforms to the welfare system were implemented in the first part of the decade, the descriptive results in Figure 2 suggest that the supply increases were largest in the latter half. If the labor market competition explanation is correct, women arriving early in the decade should be less affected, provided that their initial decision creates some inertia. To create this table, I estimate the specification from column (5) of Table 5 separately for three different subgroups groups. The first group consists of women who arrived prior to 1995 and who are currently living in the same MSA as in 1995. The second group includes only women arriving in the US after 1995. The final group consists of those who arrived in the first half of the decade but who have moved across MSAs since 1995.[37] Because I do not have a measure of native female employment changes at this five-year interval, I report the reduced form coefficients. For reference, the reduced form coefficient from Table 5, column (5) that uses women arriving over the entire sample period is reported in the first column. These results demonstrate that the negative relationship is stronger among later arrivers and internal movers, which is consistent with the timing of the implementation of reform.

For completeness, I consider the responses of low-skilled male immigrants in Table 9. Given the different degree of overlap in occupation and industry seen in Table 1, it is reasonable to expect that these supply shocks represented a larger effective increase in labor market competition for female immigrants than for male immigrants. After the inclusion of important controls (e.g. in column 6), the results suggest that men responded to a lesser

[37] Ideally, one would examine those who chose their current location prior to any reforms (pre-1993), but the census only asks an immigrant's arrival year and location in 1995; so this group cannot be precisely identified.

degree than did female immigrants. The male elasticity is about 75 percent of the female elasticity, which is roughly as anticipated, given that the male indices measuring the effective supply shock of native welfare leavers were 45-60 percent as large as the female indices (Table 1). The somewhat higher than expected coefficient may result from a number of factors including joint location decisions with a spouse or other female family member, or it may be the case that men are simply more earnings-sensitive in deciding where to locate, which leads to a stronger male response to any given expected earnings shock.

4.5 Magnitude of Crowdout and Implications for Previous Studies

Taken as a whole, therefore, the results in this section support interpreting the changing distribution of immigrants' locations as an earnings-maximizing response to changes in local labor market opportunities. One final question concerns the extent to which these changing location patterns effectively "undid" the labor supply shocks created by welfare reform. Figure 7 presents a back of the envelope calculation in response to this question, based on the IV regression results in Table 5, column 6. The x-axis measures the predicted increase in native female labor supply based on the first stage regression, expressed as a fraction of the low-skilled female population (both native-born and immigrants) in 1990.[38] This measure thus provides the percentage growth in the supply of low-skilled native labor due to welfare reform. The y-axis displays the "extra" female immigrants in the local labor force as predicted by the model, also measured as a fraction of the low-skilled female population in 1990. The "extra" immigrants variable is the difference between the predicted number of immigrants joining a city's labor force using actual pre-reform welfare participation rates and the predicted number who would have entered if all cities had the mean participation

[38]The numerator in this calculation is the predicted change in native labor force participation multiplied by the number of natives in 1990.

rate.[39] The slope of the linear regression line is -0.53 with a standard error of 0.07, which implies that when natives representing a one percentage point increase in the local supply of low-skilled female labor entered the labor market, immigrants equivalent to 0.53 percent of the previous workforce chose alternative locations.

This calculation suggests that changing immigration patterns substantially diffused the local supply shocks created by welfare reform throughout the country, which has important implications across multiple literatures. First, the magnitude of the response provides a compelling reason why other research examining the effect of reform-induced supply shocks on other native-born groups have tended to find little to no effect. For example, Blank and Gelbach (2006) find little evidence that welfare leavers had any detrimental effect on the employment and wages of low-skilled men.

In addition, this result presents a potential alternative explanation for the consistently observed similarities in native labor market outcomes between cities receiving large immigrant inflows and comparison cities. In contrast to tests of the so-called "skating rink hypothesis", which find that immigrant inflows are not offset by native outflows (Card and DiNardo 2000, Card 2001), this paper finds substantial displacement of immigrants in response to competing workers entering a local labor market. The key difference is this study's focus on the location decisions of a highly mobile factor, and the results suggest that endogenous immigrant inflows are an important equilibrating force. As a concrete example of how this mechanism may explain otherwise surprising findings, Card and Lewis (2007) find that natives in cities with larger "unexpected" new inflows of Mexican immigrants during

[39]Specifically, this variable is $\frac{(\hat{S}_j^{shock} - \hat{S}_j^{noshock})(Imm_{2000})}{Pop_{j1990}}$, where Imm_{2000} is the total number of new female immigrants who are in the labor force nationwide (census definition), Pop_{1990} is the total size of the low-education female population in 1990, and $\hat{S}$ represents the share of new immigrants predicted by the regression results. To calculate $\hat{S}_{shock}$ I take $\exp(\Delta \widehat{log(S_{2000})} + log(S_{1990}))$, i.e. the exponential of the sum of the log share from 1990 and the predicted change in log share from the regression. $S_{noshock}$ is calculated similarly, but using the average pre-reform participation rate rather than the true rate in forming the predicted change in log share. I rescale the predicted shares so that they add to 1 across all locations (the unadjusted shares add to 0.95 and 1.01).

the 1990s fared no worse than did natives in other cities. Yet these new destinations tended to have smaller relative numbers of former welfare recipients entering the labor market.[40]

A more difficult question is whether the degree of endogeneity identified in this paper implies a substantial bias in a city-level regression of native wages or employment on immigrant inflows more generally. It is certainly true that earnings-sensitive entry will tend to make the estimated coefficient in such a regression more positive (assuming that the true elasticity is negative). In the appendix, however, I show that quantifying the magnitude of this bias requires additional modeling assumptions because knowing how flows change in response to changes in labor market prospects provides only a partial answer. One also needs to know what share of the variation in immigrant inflows is due to factors other than earnings-sensitive entry as well as whether inflows are correlated with unobserved wage changes for merely coincidental reasons.

The total amount of bias can vary substantially depending on these additional factors, even holding constant the degree to which immigrants choose locations based on labor market conditions. For example, if the only source of bias in a wage regression was the response to reform-induced native supply shocks, there would be relatively little bias because the shift away from these supply shocks explains a fairly small portion of the total variation in immigrant inflows.[41] On the other hand, if immigrant flows consistently responded as strongly to other unobserved shocks as they did to the supply increases studied in this paper, this endogeneity could produce a substantial bias, even one large enough to produce an estimated coefficient that was the opposite sign of the truth.[42] In any case, this paper

[40]The cities the authors mention in the text as surprising are Houston, Dallas, Atlanta, Phoenix, Las Vegas, New York, Denver, Portland, Salt Lake City, Washington, DC, Seattle, Raleigh-Durham, Greensboro and Charlotte. Only New York, Seattle and Raleigh-Durham had pre-reform participation rates above the median (see Figure 4).

[41]Calculations described in the appendix suggest that such coefficient would be attenuated by roughly ten percent.

[42]The size of the bias increases as the share of total variation in immigrant inflows that is due to responses to unobserved wage shocks grows, and this share is not generally observable.

provides direct evidence of the type of endogeneity that motivated the use of an instrumental variables approach to isolate exogenous variation in immigrant inflows. Assessing the validity of the most common instrument - predicting inflows based on the location pattern of previous migrants from the same sending country - is thus a valuable topic for future research.

5 Conclusion

This paper provides evidence that immigrants serve as labor market arbitrageurs, differentially selecting areas with better employment prospects. Welfare reform substantially increased the labor market participation of previous recipients, and immigrants tended to favor areas with smaller relative increases in the supply of competing workers. Additional evidence helps rule out a number of alternative explanations for this pattern, including pre-existing trends away from these cities, concurrent demand increases, a secular decline in the value of traditional locations, and other unobserved amenity changes valued similarly by immigrants of all skill levels similarly. Further, the pattern of the heterogeneity in responses among subsamples provides additional support for this interpretation. Overall, the data support interpreting these changing location patterns as evidence that newly arriving low-skilled immigrants are sensitive to expected earnings when they choose where to locate.

This finding has important implications for the low-skilled labor market. Selective immigration flows tend to reduce geographic earnings inequality by helping to create a national labor market. This equilibrating function of substantial low-skilled immigration flows is especially valuable given the large barriers to moving among native-born low-skilled workers. This benefit of immigration is seldom discussed in the policy debate, and future research providing a precise estimate of its magnitude would be quite valuable.

Finally, many authors use geographic variation to determine the effect of policies or shocks on labor market outcomes. In general, future geography-based labor market research should take account of selective immigration inflows as a potential confounding factor, even in the absence of significant internal migration. This study reveals a substantial response of immigration flows to labor market incentives, and thus provides a potential alternative explanation for any surprising result in which the resulting variation in outcomes across locations was smaller than expected.

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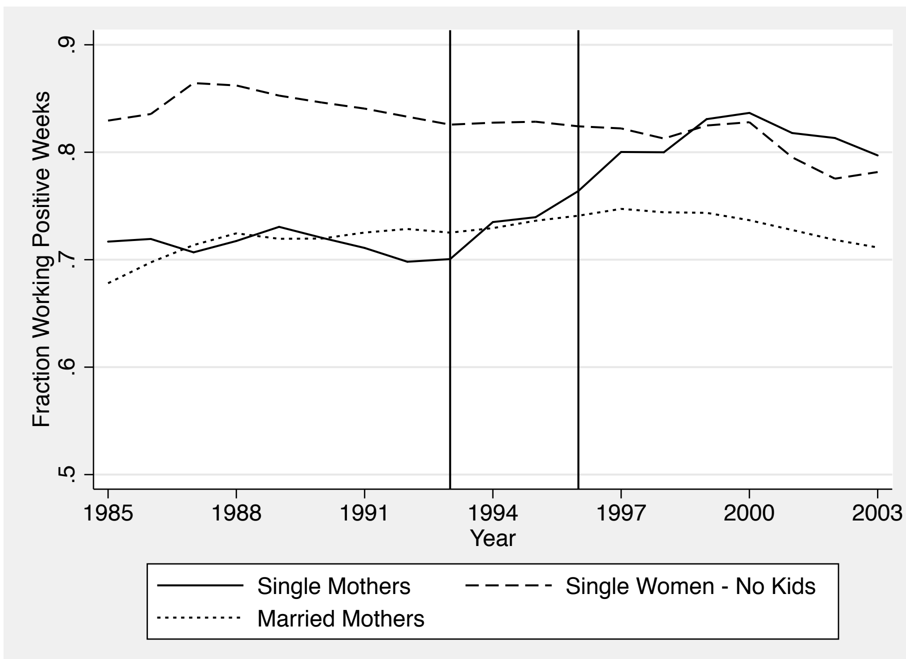

Figure 1: Female Employment Rates 1979-2003, Age 18-54, HS Degree or Less.
Notes: Source: Author's Calculations from the 1986-2004 March CPS. The sample is limited to women living in MSAs with an adult population of 150,000 in 1990.

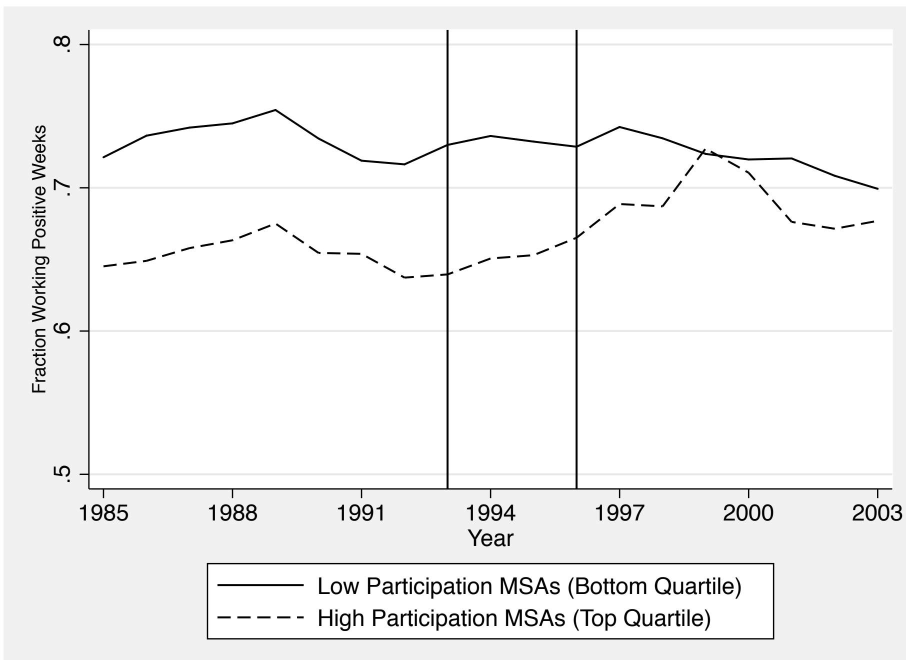

Figure 2: Female Employment Rates by Pre-Reform Welfare Participation
Notes: Source: Author's Calculations from the 1986-2004 March CPS. Selection criteria maintained from Figure 1. Participation rankings based on 1988-1992 participation rates. Low participation is less than 5.4 percent. High participation is greater than 10.7 percent.

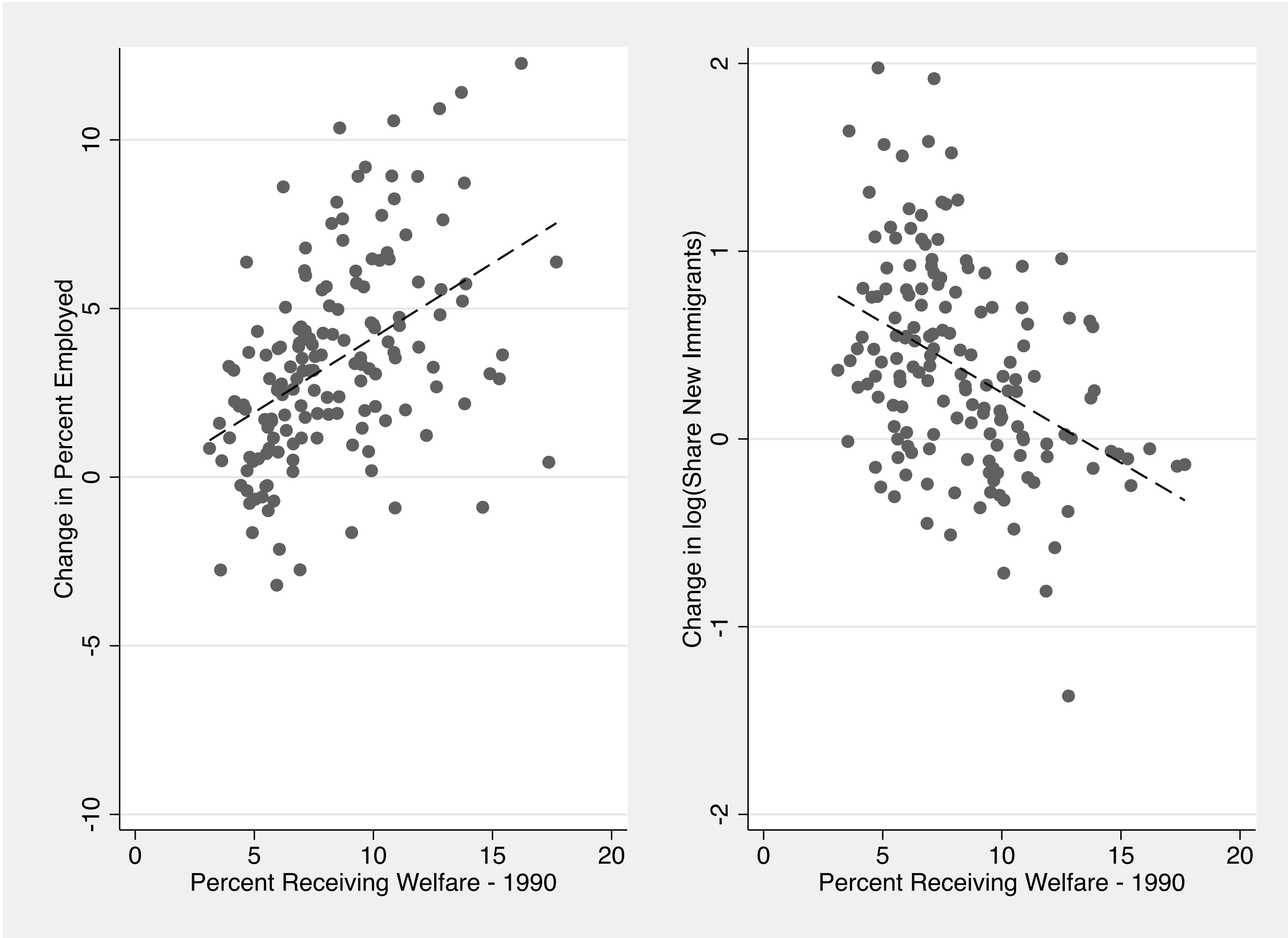

Figure 3: Changes in Native Female Employment and New Immigrant Locations 1990-2000 By 1990 Welfare Participation

Notes: Source: Author's Calculations from 1990 and 2000 PUMS. Analysis includes MSAs with at least 150,000 adult residents. Immigrant shares and welfare participation rates calculated among women 18-54 with a HS Degree or Less.

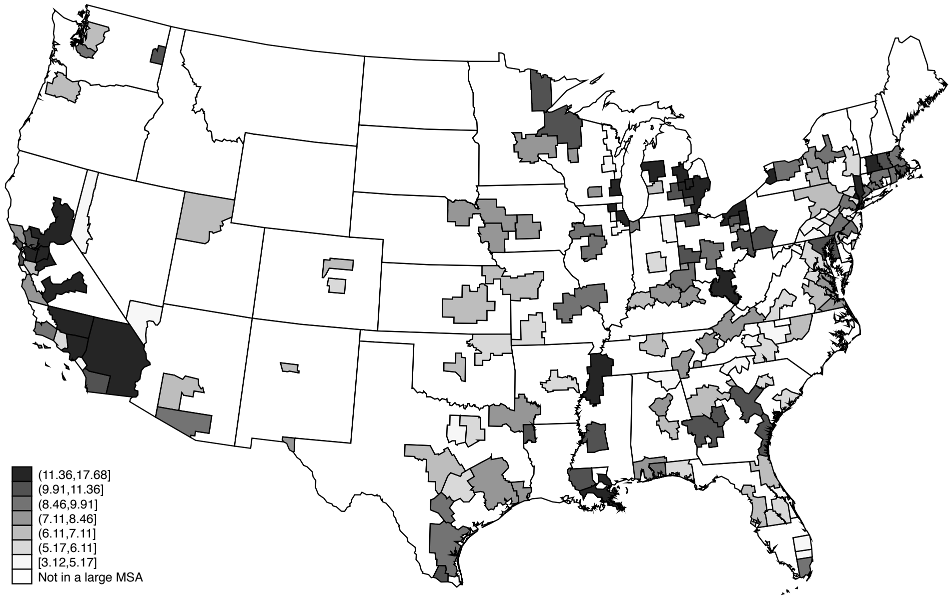

Figure 4: Fraction of Low-Skilled Female Population Using Welfare 1990

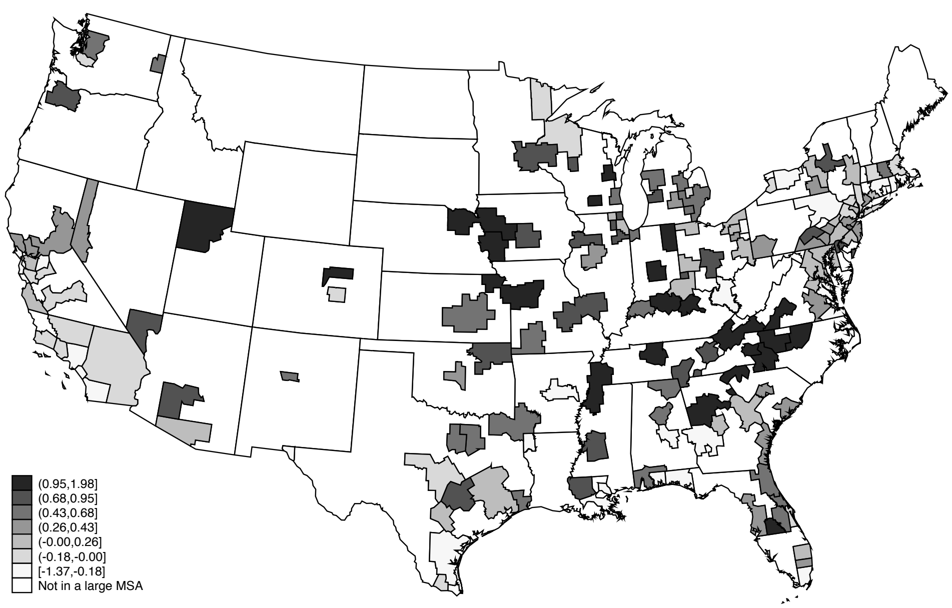

Figure 5: Change in log(Low-Skilled Female Immigrant Share) 1990-2000

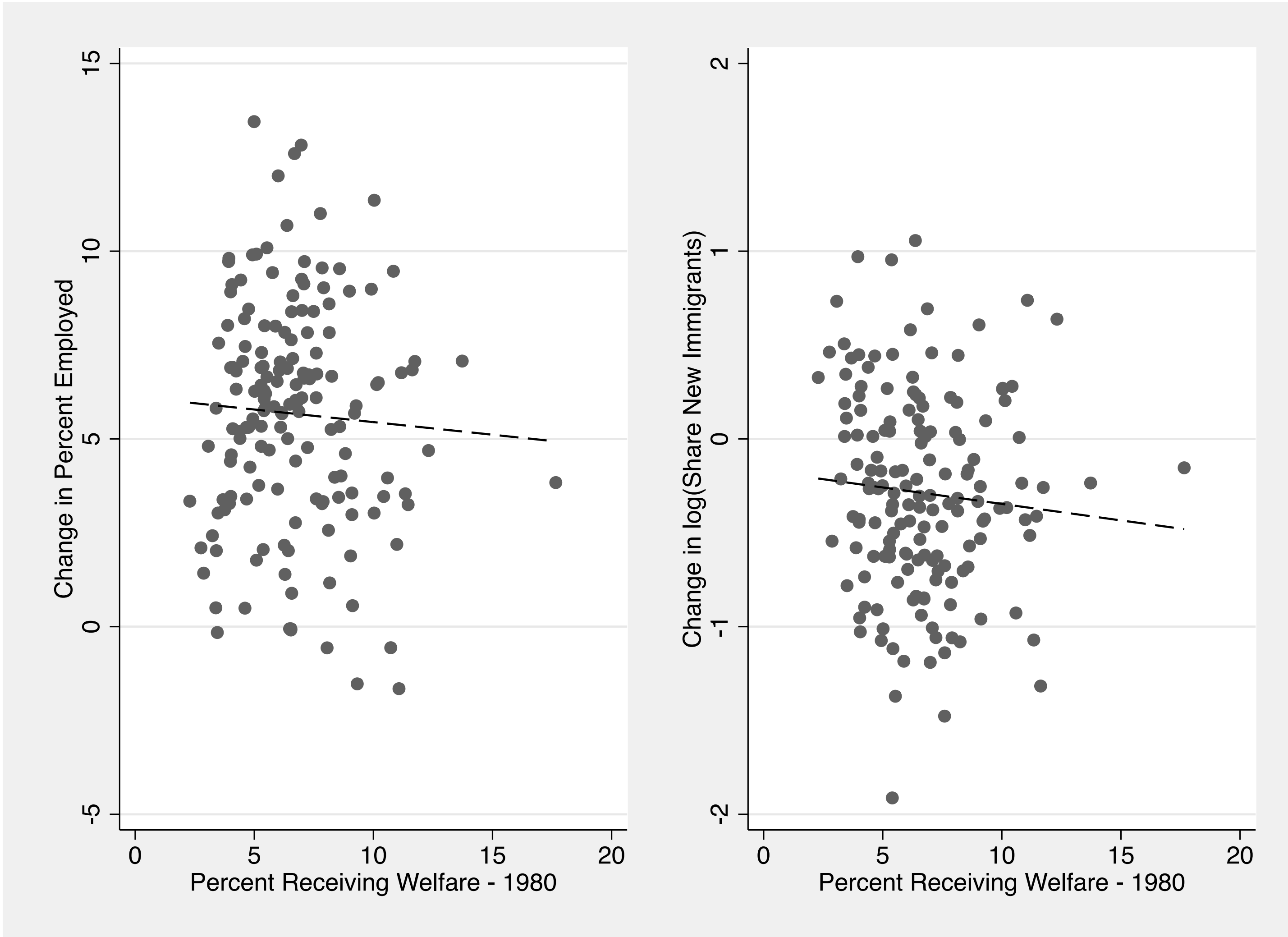

Figure 6: Falsification Test - Changes in Native Female Employment and New Immigrant Locations 1980-1990 By 1980 Welfare Participation

Notes: Source: Author's Calculations from 1980 and 1990 PUMS. Analysis includes MSAs with at least 150,000 adult residents. Immigrant shares and welfare participation rates calculated among women 18-54 with a HS Degree or Less.

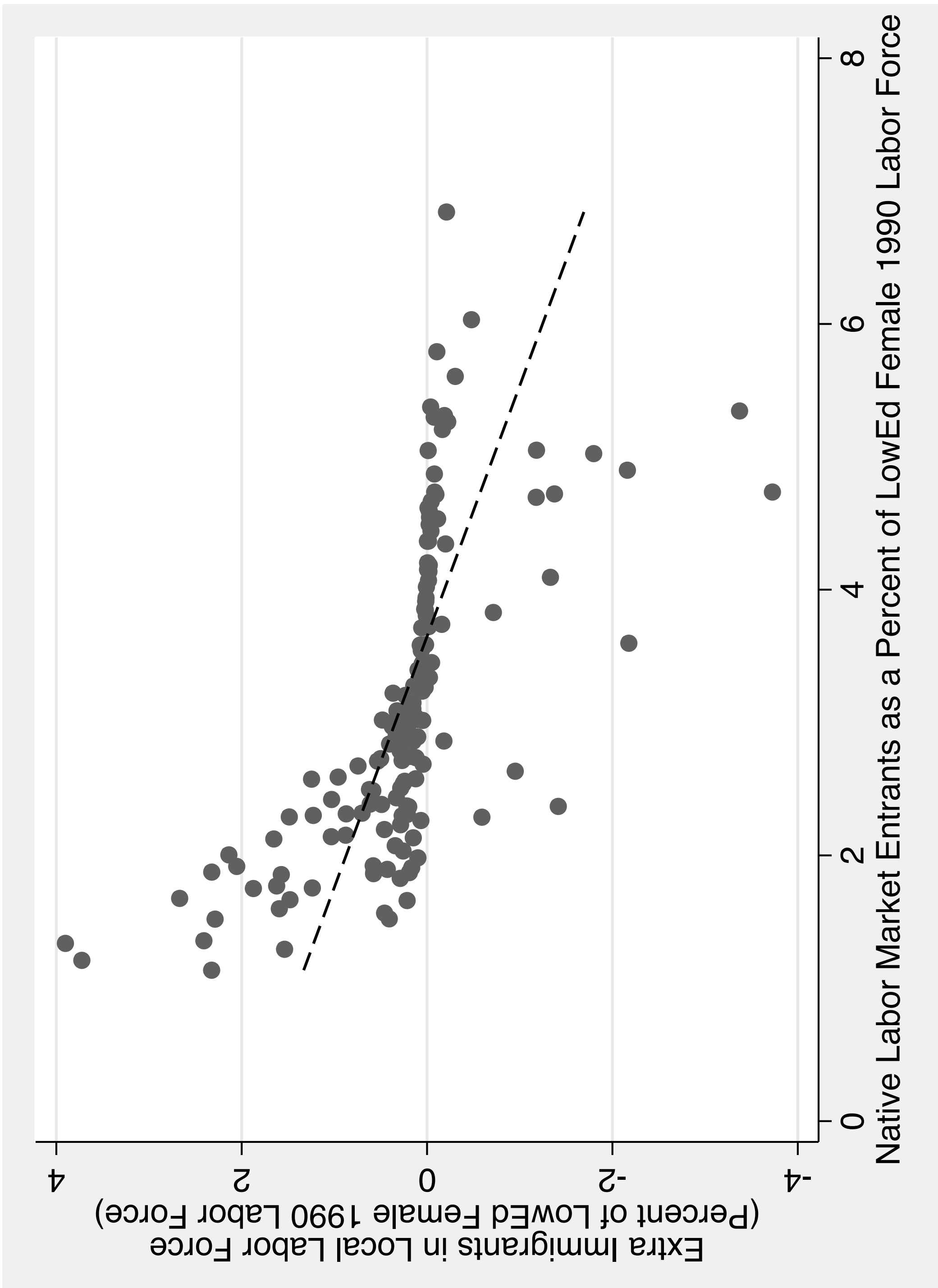

Figure 7: Predicted Additional Immigrants Selecting a City By Predicted Natives Entering Labor Force
Notes: Predictions based on Table 5, column 5 and are described in the text. Regression Coefficient: -0.53 (0.07)

Table 1: Relative Supply Increases due to Entry by Welfare Leavers

Group	Occupation	Industry
Low Ed Native Women - received welfare	4.16	2.42
Low Ed Immigrant Women	2.63	1.94
Low Ed Native Women - did not receive welfare	1.74	1.37
Low Ed Immigrant Men	1.22	1.15
High Ed Immigrant Women	1.12	1.17
Low Ed Native Men	0.85	0.86
High Ed Native Women	0.77	0.95
High Ed Immigrant Men	0.62	0.84
High Ed Native Men	0.45	0.70

Notes: Author's calculations from the 1990 and 2000 PUMS. The sample is limited to individuals living in MSAs with a population of at least 150,000 in 1990. Definition of relative supply increase found in the text. Industry and occupation codes are the 1990-basis time-consistent codes from IPUMS. Welfare leavers' industries and occupations measured in the 1990 and 2000 PUMS - all others in 2000 only. The immigrant sample is restricted to persons arriving after 1990. Low Ed means no more than a high school degree. High Ed means at least some college.

Table 2: Distribution of Source Regions - New Immigrants 18-54, HS Degree or Less

	1980	1990	2000
Mexico	34.0%	36.5%	46.9%
Central America	5.8%	13.5%	10.3%
Caribbean	12.9%	11.6%	9.1%
Europe	16.0%	7.2%	7.6%
South America	6.7%	6.8%	6.9%
Southeast Asia	8.4%	10.2%	6.5%
East Asia	7.9%	7.6%	4.7%
Southwest Asia	2.8%	3.2%	3.2%
Africa	1.3%	1.3%	2.9%
Middle East	3.1%	1.6%	1.4%
Canada	1.2%	0.6%	0.5%
Observations	98,831	144,375	230,457
Total (Weighted)	1,976,620	3,169,712	5,139,692

Source: Author's Calculations from 1980-2000 PUMS. New immigrants are those arriving in the US fewer than ten years prior to the survey. Person-level census weights used. Sample limited to individuals born in the source countries listed.

Table 3: Descriptive Statistics of Immigrants Arriving Over the Previous Decade, Age 18-54, HS Degree or Less

	1980	
	Men (47.7%)	Women (52.3%)
Unmarried	40.3%	30.5%
No spouse in household	47.1%	34.3%
Household head, no spouse	11.1%	10.1%
Any children in household	44.5%	62.9%
No HS Degree	66.4%	62.8%
	1990	
	Men (52.4%)	Women (47.6%)
Unmarried	50.8%	38.9%
No spouse in household	62.2%	44.7%
Household head, no spouse	10.9%	9.8%
Any children in household	33.6%	56.2%
No HS Degree	58.4%	54.4%
	2000	
	Men (52.8%)	Women (47.2%)
Unmarried	52.1%	37.4%
No spouse in household	65.2%	43.6%
Household head, no spouse	10.8%	9.2%
Any children in household	31.0%	56.4%
No HS Degree	52.7%	48.2%

Source: Author's Calculations from the 1980-2000 PUMS. Sample selection criteria are the same as for Table 2. The numbers in parentheses give the percent of all immigrants who were of each gender. Person-level census weights used.

Table 4: MSA Characteristics 1980-2000

	1980		1990		2000	
	Mean	SD	Mean	SD	Mean	SD
Native Female Employment Rate, HS Degree or Less	66.72	5.77	72.36	5.98	75.74	5.22
Percent Receiving Welfare - Low Education Women	6.67	2.44	8.29	3.05	4.89	2.02
Average annual employment growth rate - over past decade	--	--	2.62	1.62	2.08	1.21
Immigrant Share of MSA Population, previous census	--	--	6.05	6.27	8.07	8.99
Average Increase in Log(Current Minimum Wage) relative to decade start	--	--	-0.2128	0.0131	0.0785	0.0766
Maximum cash benefit, family of three (hundreds of dollars)	--	--	3.90	1.64	--	--
Change in percent of immigrants receiving welfare (from previous decade)					-1.11	3.19
MSA in state that restored Food Stamp benefits	--	--	--	--	0.58	0.48
MSA in state that restored TANF benefits	--	--	--	--	0.36	0.47
MSA in state that restored Medicaid benefits	--	--	--	--	0.32	0.46
MSA in state that restored SSI benefits	--	--	--	--	0.15	0.36

Source: Author's Calculations from the 1980-2000 PUMS and 1980-2000 County Business Patterns. Welfare policy data from Zimmerman and Tumlin (1999). Selection criteria maintained from Table 2. Person-level census weights used.

Table 5: Percentage Change in Share of Female Immigrants Selecting a City

	OLS	IV	IV	IV	IV	IV	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Change in Native Female Employment Rate	-0.0286*	-0.169***	-0.189***	-0.131***	-0.132***	-0.104***	-0.117***
	(0.0158)	(0.0428)	(0.0432)	(0.0263)	(0.0288)	(0.0240)	(0.0271)
Change in decade average employment growth rate			0.249***	0.138***	0.138***	0.0992***	0.128***
			(0.0496)	(0.0324)	(0.0342)	(0.0332)	(0.0387)
Immigrant Share of MSA Population, 1990				-0.0337***	-0.0338***	-0.0262***	-0.0349***
				(0.00334)	(0.00377)	(0.00411)	(0.00607)
Change in Average Increase in Log(Real Minimum Wage)					-0.0146	-0.0851	0.551
					(0.424)	(0.400)	(0.753)
Change in log(Share of Immigrants - College Degree)						0.371**	0.451***
						(0.157)	(0.155)
Maximum cash benefit, family of three							0.0753**
							(0.0380)
Change in Percent of Low-Skilled Female Immigrants Receiving Welfare							0.0178
							(0.0121)
MSA in state that restored Food Stamp benefits							0.0190
							(0.0949)
MSA in state that restored TANF benefits							-0.0222
							(0.121)
MSA in state that restored Medicaid benefits							-0.213*
							(0.116)
MSA in state that restored SSI benefits							0.490***
							(0.134)
Constant	0.470***	0.944***	1.144***	1.163***	1.169***	0.935***	0.581*
	(0.0746)	(0.144)	(0.158)	(0.115)	(0.220)	(0.211)	(0.349)
Number of Cities	157	157	157	157	157	157	157
First Stage Coefficient		0.442	0.413	0.487	0.474	0.429	0.435
First Stage F-stat		25.51	27.44	50.05	41.98	27.28	30.03

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Notes: Authors Calculations from the 1990 and 2000 PUMS. Selection criteria maintained from Table 2. The excluded instrument is the female welfare participation rate in 1990. Immigrant shares are calculated using person-level weights. The regressions are unweighted.

Table 6: Percentage Change in Share of Female Immigrants Selecting a City. Source-Destination Differences.

	OLS	IV	IV	IV	IV	IV	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Change in Native Female Employment Rate	-0.0208	-0.155***	-0.159***	-0.120***	-0.0989**	-0.0785**	-0.0820*
	(0.0190)	(0.0532)	(0.0507)	(0.0394)	(0.0441)	(0.0388)	(0.0445)
Change in decade average employment growth rate			0.173***	0.120***	0.0902**	0.0529	0.0456
			(0.0415)	(0.0297)	(0.0372)	(0.0324)	(0.0361)
Ethnic group members as pct of MSA pop - decade start				-0.236***	-0.222***	-0.200***	-0.176***
				(0.0270)	(0.0289)	(0.0264)	(0.0259)
Change in Average Increase in Log(Real Minimum Wage)					0.738	0.510	0.923
					(0.508)	(0.440)	(0.758)
Change in log(Share of Immigrants - College Degree)						0.305***	0.275***
						(0.0479)	(0.0444)
Maximum cash benefit, family of three							0.0312
							(0.0399)
Change in Percent of Low-Skilled Female Immigrants Receiving Welfare							-0.0180
							(0.0186)
MSA in state that restored Food Stamp benefits							-0.142
							(0.107)
MSA in state that restored TANF benefits							0.193
							(0.143)
MSA in state that restored Medicaid benefits							-0.00741
							(0.120)
MSA in state that restored SSI benefits							-0.293***
							(0.111)
Constant	0.0664	0.430**	0.601***	0.466***	0.177	0.147	-0.0414
	(0.115)	(0.181)	(0.181)	(0.149)	(0.277)	(0.246)	(0.373)
Observations (Source-Destination Pairs)	1,136	1,136	1,136	1,136	1,136	1,136	1,136
First Stage Coefficient		0.357	0.352	0.369	0.364	0.360	0.377
First Stage F-stat		20.70	22.42	26.04	18.91	18.34	21.76

Standard Errors clustered by MSA in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Notes: Selection criteria maintained from Table 2. Destinations not selected by anyone from a source during one census year are omitted from the entire sample. Analysis limited to source-destination pairs with valid observations for all covariates. Pairs are weighted by the square root of the number of source observations nationwide.

Table 7: IV Coefficient on Change in Female Employment - By Source Region

	β	se	First Stage F-stat
Canada	0.07	0.08	8.93
Mexico	-0.24	0.14	8.56
Central America	-0.17	0.08	17.24
Caribbean	-0.11	0.08	13.96
South America	0.00	0.06	26.26
Europe	0.00	0.04	33.05
East Asia	-0.09	0.07	31.62
Southeast Asia	-0.07	0.06	16.45
Southwest Asia	0.01	0.07	20.71
Middle East	-0.10	0.08	17.98
Africa	0.03	0.06	27.21

Note: The specification and sample selection criteria are identical to column (6) of Table 6.

Table 8: Reduced Form Estimates of Change in Selection Probability by Arrival Year

	All New Female Immigrants	Arrived 1990-1994 Same MSA	Arrived 1995-2000	Arrived 1990-1994 Moved MSAs
	(1)	(2)	(3)	(4)
Welfare Participation Rate - Low Education Women	-0.0624***	-0.0550***	-0.0652***	-0.0673***
	(0.0128)	(0.0142)	(0.0158)	(0.0150)
Change in average employment growth rate	0.0492*	0.0290	0.0329	0.00985
	(0.0250)	(0.0243)	(0.0296)	(0.0363)
Immigrant Share of MSA Population, 1990	-0.0193***	0.00184	-0.0258***	-0.0501***
	(0.00330)	(0.00424)	(0.00427)	(0.00581)
Change in Average Increase in Log(Real Minimum Wage)	0.205	-0.369	0.664*	0.393
	(0.347)	(0.457)	(0.394)	(0.491)
Constant	1.013***	0.631***	0.993***	1.590***
	(0.177)	(0.229)	(0.206)	(0.237)
Observations	157	157	157	156
R-squared	0.314	0.086	0.309	0.476

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Source: Author's Calculations from 1980-2000 PUMS. Selection criteria maintained from Table 2. Early arrivers entered the US prior to 1995, and have not moved since 1995. Late arrivers entered the US after 1995. Movers arrived prior to 1995, but have since moved to a different MSA than where they were living in 1995.

Table 9: Percentage Change in Share of Male Immigrants Selecting a City

	OLS	IV	IV	IV	IV	IV	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Change in Native Female Employment Rate	-0.0359*	-0.203***	-0.224***	-0.132***	-0.114***	-0.0775***	-0.0661**
	(0.0198)	(0.0589)	(0.0588)	(0.0312)	(0.0321)	(0.0269)	(0.0262)
Change in decade average employment growth rate			0.287***	0.111***	0.0809**	0.0374	-0.00243
			(0.0624)	(0.0363)	(0.0397)	(0.0379)	(0.0403)
Immigrant Share of MSA Population, 1990				-0.0538***	-0.0493***	-0.0391***	-0.0339***
				(0.00536)	(0.00526)	(0.00578)	(0.00635)
Change in Average Increase in Log(Real Minimum Wage)					1.109**	0.866*	-0.685
					(0.496)	(0.445)	(0.812)
Change in log(Share of Immigrants - College Degree)						0.530***	0.568***
						(0.158)	(0.151)
Maximum cash benefit, family of three							-0.0563
							(0.0421)
Change in Percent of Low-Skilled Female Immigrants Receiving Welfare							0.0138
							(0.0162)
MSA in state that restored Food Stamp benefits							-0.328***
							(0.102)
MSA in state that restored TANF benefits							0.120
							(0.164)
MSA in state that restored Medicaid benefits							-0.486***
							(0.146)
MSA in state that restored SSI benefits							0.380***
							(0.144)
Constant	0.718***	1.279***	1.508***	1.538***	1.103***	0.849***	1.670***
	(0.0993)	(0.198)	(0.221)	(0.139)	(0.248)	(0.212)	(0.408)
Number of Cities	156	156	156	156	156	156	156
First Stage Coefficient		0.442	0.413	0.487	0.474	0.446	0.436
First Stage F-stat		25.51	27.44	50.05	41.98	30.67	31.86

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Source: Authors Calculations from the 1990 and 2000 PUMS. Selection criteria maintained from Table 2. The excluded instrument is the female welfare participation rate in 1990. Immigrant shares are calculated using person-level weights. The regressions are unweighted.

Appendix - For Online Publication

A-1 Implications for Immigration Flows and Native Wages

This appendix provides a more complete treatment of the extent to which the earnings-sensitive migration in this study implies that a cross-geography approach will tend to underestimate the effect of immigration on wages.

A-1.1 A Conceptual Framework

Consider the following wage regression:

$$\Delta w_k = \beta m_k + \Delta\epsilon_k \tag{9}$$

Here k indexes geographic labor markets, Δw_k represents changes in local wages (typically measured in logs) and m_k measures the proportionate change in labor supply due to new immigrant inflows, i.e. new immigrants divided by the size of the labor force($\frac{\Delta M}{M+N}$). β is the true elasticity of demand for low-skilled labor, and $\Delta\epsilon_k$ captures all other components of the wage determination process. The typical concern is that $Cov(m, \Delta\epsilon) > 0$ and $\beta < 0$, and thus that the estimated demand elasticity is less negative than the true elasticity. To fix ideas, consider the coefficient that would result (in expectation) from a bivariate estimate of this regression:

$$\hat{\beta} = \frac{Cov(m, \Delta w)}{Var(m)} \tag{10}$$

$$\hat{\beta} = \beta + \frac{Cov(m, \Delta\epsilon)}{Var(m)} \tag{11}$$

How then can one think about using the locational response information I have identified in this paper to determine whether this bias might be large relative to β? This paper suggests that new immigrant inflows are a function of local expected earnings, i.e. $m_k = f_k(\Delta\epsilon_k)$, and thus the data generating process can be written as:

$$\Delta w = \beta f_k(\Delta\epsilon_k) + \Delta\epsilon_k \tag{12}$$

Note that I have allowed for the possibility that the function relating wage shocks to immigrant inflows may be different across locations. Importantly, this allows for the possibility that immigrant inflows to location k are a function of factors other than the wage shock experienced in location k. The results in this paper clearly imply that $f'(\cdot) > 0$, but without any additional structure, it is not possible generally to determine $Cov(f(\Delta\epsilon), \Delta\epsilon)$, which is what is required to determine the bias in a regression of observed wage changes on observed immigrant inflows.

A-1.2 Additional Structure Including My Results

I have identified the change in immigrant entry due to one component of $\Delta\epsilon$. Letting n_k denote the reform-induced native entrants as a share of the local labor force, we can rewrite the error term as

$$\Delta\epsilon_k = \beta n_k + \beta\nu_k \tag{13}$$

with β the same labor demand elasticity, n the exogenous entry of natives due to welfare reform (measured in percentage terms analogously to immigrant entry), and $\beta\nu_k$ containing all other unobserved determinants of changes in local wages such as local changes in demand.[43] Note that by writing the wage effect of native entry as βn, I have assumed that immigrants and natives are perfect substitutes and provide the same number of efficiency units of labor. I have also written any other shocks in "worker entry equivalents," i.e. I have simply rescaled any other shocks so that the effect on local wages of a one-unit change in those other unobservables is equivalent to the effect of a one percent native entry shock.

What then, do the estimates in this paper allow to be calculated? One can know for each location, $f_k(\beta n_k + \beta\nu_k)$ and $f_k(\beta\nu_k)$ as well as n_k. In words, I can use my estimation results to predict the immigrant inflows that would have occurred in each city both with and and without the welfare-reform induced native entry. Without imposing some additional structure on $f_k(\cdot)$, this still does not provide any insight into the generic covariance between immigrant inflows and the error term in the wage regression. In order to make progress, one must assume a form for $f_k(\cdot)$ such that the results provided by the empirical model are informative. One way to do so is to make the following assumption:

$$f_k(\Delta\epsilon) = m_k^0 + \gamma\Delta\epsilon \tag{14}$$

In words, the immigrant inflows into city k are the level of inflows that would occur if the unobserved wage shock were zero plus a (constant across cities) elasticity of supply of additional immigrants times the value of the wage shock. This is a particularly useful way of parameterizing $f_k(\cdot)$ because it implies the following relationship among the predicted and counterfactual inflows:

$$f_k(\beta n_k + \beta\nu_k) - f_k(\beta\nu_k) = \beta\gamma n_k \tag{15}$$

I can therefore estimate $\phi \equiv \beta\gamma$ from a regression of the additional immigrant inflows implied by the main results in my paper on the size of the welfare reform induced native supply shock implied by the first stage. In fact, this is precisely what Figure 7 provides. The resulting coefficient is roughly -0.5, i.e. each reform-induced native entrant leads to 0.5 fewer immigrant entrants. In order to see how this number informs a bias calculation, it is useful to use the previous assumption of the form of $f(\cdot)$ to rewrite the observed immigrant inflows in Equation (11) as $m_k = m_k^0 + \phi n_k + \phi\nu_k$. Doing so implies that the expected value

[43] For brevity, I have suppressed the intercept term in (12). Thus n_k and ν_k should be considered deviations from their respective means.

of $\hat{\beta}$ can be rewritten as:

$$\hat{\beta} = \beta + \frac{Cov(m^0 + \phi(n+\nu), \beta(n+\nu))}{Var(m)} \tag{16}$$

$$= \beta \left(1 + \frac{Cov(m^0, (n+\nu)) + \phi Var(n+\nu)}{Var(m^0) + 2Cov(m^0, \phi(n+\nu)) + \phi^2 Var(n+\nu)}\right). \tag{17}$$

As the above equation demonstrates, even with some fairly restrictive assumptions, the bias formula does not depend in a straightforward way on the responsiveness of immigrant inflows to wage shocks (ϕ). Instead the bias can be decomposed into two pieces. First, there is the "coincidental" portion - the part due to the fact that, even if immigrant inflows were not actually *affected* by wage shocks ($\phi = 0$), immigrant inflows could nevertheless be correlated with the error term. The second portion of the bias is due to the fact that immigrant inflows respond to changes in wages. Importantly, without knowing $Cov(m^0, n+\nu)$, it is not possible even to sign the total bias, although we can say confidently that earnings-sensitive inflows move the bias in the opposite direction of the sign of β because $\phi < 0$. Given this difficulty in making general statements, it is therefore instructive to consider some special cases.

A-1.2.1 Case 1: No Unobserved Wage Shocks other than Welfare Reform

One could consider the case in which the supply shocks induced by welfare reform are the only unobserved factor affecting wages that is correlated with the error term, i.e. $Var(\nu) = 0$. In this case, equation (17) simplifies to:

$$= \beta \left(1 + \frac{Cov(m^0, n) + \phi Var(n)}{Var(m^0) + 2Cov(m^0, n) + \phi^2 Var(n)}\right) \tag{18}$$

$$= \beta \left(1 + \frac{Cov(\hat{m}, n)}{Var(\hat{m})}\right) \tag{19}$$

with $\hat{m}$ the immigrant inflow predicted by my model. I can estimate this regression coefficient (i.e. regress $\hat{m}$ on n), and it implies that $\hat{\beta} \approx 0.9\beta$, i.e. that a wage regression would underestimate the true elasticity by roughly ten percent. It is unlikely, however, that the welfare reform shock was the only shock happening over the 1990s, and thus it is not clear how informative such an estimate is to the general question.

A-1.2.2 Case 2: No "Coincidental" Correlation between Inflows and Wage Shocks

As an alternative, suppose that $Cov(m^0, n+\nu) = 0$, i.e. the only reason that inflows are correlated with wage shocks is due to earnings-sensitive location choices among immigrants. In this case, equation (17) simplifies to:

$$= \beta \left(1 + \frac{\phi Var(n + \nu)}{Var(m^0) + \phi^2 Var(n + \nu)} \right) \tag{20}$$

In order to understand this expression, it is useful to define $\rho \equiv \frac{\phi^2 Var(n+\nu)}{Var(m^0)+\phi^2 Var(n+\nu)}$, or the share of the variance in immigrant inflows that results from endogenous entry. Then, the above equation can be written as:

$$= \beta \left(1 + \frac{\rho}{\phi} \right) \tag{21}$$

Here, there are several useful observations. First, conditional on the degree of endogeneity (ϕ), a wage regression will less biased when more of the geographic variation in inflows is exogenous (i.e. as $Var(m^0)$ increases). Second, if $\rho < -\phi$, $\hat{\beta}$ will be of the correct sign, although smaller in magnitude. Finally, if $\rho > -\phi$, it is possible that $\hat{\beta}$ will be sufficiently biased as to reverse the sign.

A-1.3 Conclusion

This appendix provides a useful framework for considering how endogenous immigrant inflows are likely to bias a wage equation. Importantly, the degree to which immigrant inflows respond to unobserved changes in wages is only one of a number of necessary estimates that determine how much bias is likely to result from earnings-sensitive entry. Nevertheless, the results in this paper suggest a strong immigrant response to an exogenous wage shock, which reinforces the idea that a cross-geography approach likely requires an instrumental variables approach in order to provide a causal estimate. Given that nearly all studies rely on a single instrument (the "supply push" or "previous migrants" instrument), future research should examine this instrument with greater scrutiny.