THE IMPACT OF IMMIGRATION ON THE LABOR MARKET FOR NATIVE-BORN WORKERS:

INCORPORATING THE DYNAMICS OF INTERNAL MIGRATION

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The dramatic increase in immigration over the past two decades has led to a growing concern about its impact on the labor market outcomes of native-born workers. There now exists an expanding body of literature that attempts to address these two questions. Do immigrants depress the earnings of native-born workers by competing with them for jobs? Do immigrants reduce the employment levels of natives?

However, one crucial and well recognized flaw remains: namely, the failure in existing empirical studies to capture the dynamics of internal migration of natives in response to increased immigration [Card, 1990; Butcher and Card, 1991; Borjas, 1994; Borjas, Freeman, and Katz, 1996]. This paper utilizes a unique approach developed in the demography literature to obtain better estimates of the impact of immigration on natives' wages and employment by directly controlling for internal migration flows.

THEORETICAL AND EMPIRICAL SETTING

Many empirical studies have attempted to capture the effects of immigration on wages and employment. The most common method is a regional, or area, approach. This approach typically compares the effects of immigration on native wages across regions by introducing a variable which controls for immigration as a percentage of the total labor force. The basic empirical test, thus, consists of estimating a wage (or employment) equation which regresses the log of weekly wages for native-born wage and salary workers on a set of human capital variables, and then also includes a variable which captures the proportion of immigrants in a given geographical area.

This approach has two serious problems. One problem is the possible endogeneity of immigrant concentrations, since it is expected that immigrants will be attracted to those areas that offer the greatest rewards and lowest costs to migration. Estimates of the impact of immigration, therefore, may be biased since the ratio of immigrants is expected to be positively correlated with uncaptured influences on wages in the error term. This problem can be mitigated by isolating areas to which immigration is largely exogenous [Card, 1990], or by eliminating the endogeneity problem with the use of an instrumental variables procedure [Altonji and Card, 1991; Borjas, 1987; DeFreitas, 1991; Schoeni, 1997].

The second problem is that results from cross-sectional studies can be biased because they fail to account for the migratory response of natives to increased immigration. These models implicitly assume that immigrants enter and compete in closed local economies. The typical responses of native-born workers to increased immigration, namely migration and movement out of the labor force, are ignored. Since these models fail to measure shifts in native-born labor supply, the coefficients on the immigration variable are not necessarily indicative of the "true" impact of immigration.

The response of natives to lower wages caused by immigration should be to reduce labor supply, thereby increasing wages and employment for the remaining native-born workers. As immigrants begin to compete with native workers for jobs, wages and employment (for natives) will fall. The expected result is that some native workers will migrate away from those labor markets causing equilibrium wages to rise. Since immigration and native net internal migration should be inversely related, excluding controls for net migration will generate immigration coefficients that are biased down. Having failed to control for shifts in native-born labor supply, it is not surprising that the estimated impact of immigration in the existing models is small and/or insignificant.[1] Indeed, Borjas, Freeman, and Katz [1996] warn that the relatively small effects of immigration on the labor market outcomes of native workers may be due to the "diluting effect of native migration across regions and failure to take adequate account of other regional labor-market conditions."

Some evidence suggests that the migration of native workers responds to immigration. Filer [1992] finds that in-migration of natives is lower in areas with higher concentrations of immigrants. Similarly, Frey [1995] claims that increased immigration into metropolitan areas is associated with larger out-migration of natives. On the other hand, White and Imai [1994] find that areas with high concentrations of immigrants exhibit both lower rates of native in-migration and out-migration. In addition, Butcher and Card [1991] conclude that native in-migration to cities during the 1980s was positively correlated with immigration, except for New York, Los Angeles, and Miami. So, while there seems to be no consensus on how native migration responds to immigration, there is no doubt that internal migration flows may cause variations in native wages and employment across metropolitan areas. The next obvious step, therefore, is to obtain estimates of net-migration in these areas and to control for it in the wage and employment equations.

REVISING THE BASIC EMPIRICAL FRAMEWORK

Estimating Net Migration

The method used to measure net migration in this study will be the intercensal cohort-component (or census survival) method used by demographers.[2] The information required to employ this technique is available in the same public use microdata samples from the federal censuses that will be used to estimate the wage equations. The first step is to calculate the survival ratio for the entire male native-born population at each specific age. This is accomplished by dividing the sample population of a given age in a census year by the sample population of those aged ten years younger in the previous census year. This can be estimated from the public use samples using the following formula:

$$SR_x = POP_{1990,\, x+10}/POP_{1980,\, x} \tag{1}$$

where the x subscript represents age, POP is the size of the sample population, and SR is the survival ratio.[3]

Once the survival ratio has been calculated, estimates of the native male population that one would expect to find of a given age in a Metropolitan Statistical Area (MSA) at the end of the decade can be produced, assuming that there has been no migration. The number of individuals of a given age in a MSA are multiplied by their corresponding survival ratio to obtain the expected number of individuals in a MSA ten years later. Then, the actual number of individuals in a given MSA of residence is subtracted from this expected number to obtain an estimate of net migration over the decade [Carter and Sutch, 1996]. This process is accomplished by using the following formulas:

$$E(POP)_{1990,\, x+10} = (SR_x)(POP_{1980,\, x}) \tag{2}$$

$$NM_{1990,\, x+10} = POP_{1990,\, x+10} - E(POP_{1990,\, x+10}) \tag{3}$$

where the x subscript represents age, $E(POP)$ is the expected size of the population, SR is the survival ratio, POP is the actual size of the population, and NM is net migration.

A possible criticism to this approach is that survival ratios will not be accurate because censuses are subject to substantial underenumeration and age heaping. For example, individuals commonly and systematically report their young children to be older than their actual age, which leads to underenumeration of children in the 0 to 5 year old cohort. Age heaping, on the other hand, results when many individuals report their age within the 20 to 29 year old cohort when they are in fact slightly younger or older. These phenomena will lead to calculations of survival ratios that either overstate or understate true mortality. This is likely to be a valid criticism, but it does not imply that the estimates of net internal migration will be flawed. In fact, to the extent that underenumeration and age heaping are uniform across the MSAs under study, the census survival ratios will automatically correct for this bias [Sutch, 1975].

The Wage and Employment Equation

Once an estimate of the net migration of native-born individuals is obtained, this can be added to the wage and employment equations to correct for possible omitted variable bias. The estimated wage equation, following the specification of LaLonde and Topel [1991] and DeFreitas [1991], will have the form:

$$\log w_{i,s} = \alpha_0 + \alpha_1\,(\mathbf{EDUC})_i + \alpha_2\,(EXP)_i + \alpha_3(EXP)_i^2 + \alpha_3(\mathbf{REG})_i + \alpha_5\,(MARRIED)_i + \alpha_6\,(HEALTH)_i + \alpha_7\,(\mathbf{OCC})_i + \alpha_8\,(\mathbf{IND})_i + \alpha_9\,(I/POP)_s + \alpha_{10}\,(NM/POP)_s + \epsilon_{i,s} \tag{4}$$

where the i and s subscripts represent individual and MSA, respectively, $\log w$ is the natural logarithm of the weekly wage (or number of weeks worked), **EDUC** is a vector of categorical variables representing education groups (e.g., high school graduate, some college, etc.), *EXP* represents years of estimated post-school work experience (age − years of schooling − 6), EXP^2 is *EXP* squared, **REG** is a vector of categorical variables representing region of residence (e.g., south, west, etc.), *MARRIED* is a marital status dummy, *HEALTH* is a health status dummy, **OCC** is a vector of categorical variables representing occupation (e.g., operative, laborer, etc.), **IND** is a vector of categorical variables representing industry (e.g., manufacturing, construction, etc.), I/POP is the ratio of immigrants to the total working age population, and NM/POP is the ratio of net migration of natives to the total working age population.[4]

The same process can be used to estimate the impact of immigration on employment. The number of weeks employed is simply substituted for the natural log of wages in the equations. This is an important part of the analysis, especially if labor supply is relatively elastic. Also, in some low-skilled jobs the minimum wage may constrain downward flexibility in wages. If this is the case, the impact of immigration may be more accurately captured by employment fluctuations.

According to Blanchard and Katz [1992], regional labor markets may take as long as 10 years to reequilibrate through internal migration in response to an economic shock. In other words, immigration during the 1980s may affect migration patterns through 1990. Consequently, a specific form of this general equation will attempt to capture the impact of varying immigrant concentrations on wages and employment in 1990 while controlling for net migration from 1980 to 1990.

Instrumenting the Concentration of Immigrants and Net Migration?

The problem with the specification of equation (4) is that it treats both immigration and net migration of natives as an exogenous process. In other words, it assumes that shifts in the labor supply curve occur exogenously while the labor demand curve remains fixed. However, it is likely that foreign-born and/or native-born migrants choose to locate (or re-locate) in booming labor markets and, as a result, the I/POP and NM/POP variables are endogenous in the wage and employment equations. In order to correct for possible endogeneity bias the concentration of immigrants and the ratio of native net migration across MSAs are instrumented. Both variables are instrumented with the fraction of immigrants in 1980, its square, the standard deviation of the weekly wage in 1980, and the average yearly income from public assistance in 1980.

Bound, Jaeger, and Baker [1995] provide a word of caution in regards to estimation with instrumental variables. According to these authors, instrumental variable estimation can result in coefficients that are largely inconsistent when the relationship between the instruments and the endogenous explanatory variable is weak. Furthermore, Bound, Jaeger, and Baker [1995] suggest that the partial R-squared and F-statistic on the excluded instruments in the first-stage regression should be reported and used to evaluate the quality of the instrumental variable estimates. These are reported for each group of regressions in the corresponding tables. In addition to

these statistics, a Generalized Method of Moments Specification Test was used to measure the relationship between the instruments and the errors in the structural equation and to ensure the validity of the instruments.

A Hausman Specification Test was performed in order to determine whether, or not, the immigration and net migration coefficients are significantly different when they are instrumented [Griffiths, Hill, and Judge, 1993, 475-76]. The null hypothesis that the coefficients are the same cannot be rejected in the case of the net migration coefficients in the wage regression of less-educated blacks, and the employment regressions of less-educated blacks and higher-educated blacks and Hispanics.[5] Therefore, the net internal migration variable is treated as endogenous in all, but the above mentioned cases. The concentration of immigrants, on the other hand, is treated as endogenous in only the wage regression for higher educated Hispanics and the employment regression for less educated whites, since those are the only regressions that result in significantly different coefficients when the variable is instrumented.[6]

The Data

The data set used in this analysis is the Integrated Public Use Microdata Series (IPUMS). The IPUMS consists of a series of compatible-format individual-level representative samples from the U.S. Census [Ruggles and Sobek, 1995]. This study utilizes the 1980 "B" sample and 1990 "1%" sample. The sample density for both of these is 1/100.

The samples for this analysis include males aged 16-64 who were living in MSAs identified on both the 1980 and 1990 Public Use Samples.[7] The variables constructed are the proportion of immigrants to the total working age population and the ratio of net migration of natives for each MSA. In addition, weekly wages are defined as annual earnings divided by the number of weeks worked in 1989. The regression samples were restricted to the civilian, non-student, native-born wage and salary workers who were working for pay during 1989 and reported all the necessary personal and employment information.

EMPIRICAL RESULTS

In general, the results are consistent with previous work in this area. That is, increases in immigration are often times associated with an improvement in labor market outcomes, and negative results, when present, tend to be relatively small. Specifically, the positive effects of immigration on wages for some groups and the negative effects on employment are both consistent with the existing literature (see Table 1).

The Effect of Immigration on Wages

Table 2 presents a summary of the immigration and internal migration coefficients for male natives without a high school diploma. Table 3 shows results for those with at least a high school diploma. For each education group, two models are esti-

TABLE 1
Review of Previous Studies on Labor Market Impact of Immigration

Wage/Earnings Results

Study	All	Native Group Whites	Blacks	Hispanics
Schoeni [1997]	———	-.20 to .46	-.31 to .63	-.54 to .29
Borjas, Freeman, and Katz [1996]	-.02 to .29	———	———	———
Altonji and Card [1991]	.02 to .47	-.26 to -.18	-.15 to .74	———
LaLonde and Topel [1991]	———	———	-.06 to -.02	-.02 to -.01
Borjas [1990]	———	-.01	-.02	———
Bean, Lowell, and Taylor [1988]	———	———	-.003 to .06	-.005 to .05
Grossman [1982]	-.03 to -.02	———	———	———

Employment Results

Study	All	Native Group Whites	Blacks	Hispanics
Schoeni [1997]	———	-.09 to .10	-.29 to .23	-.14 to .04
Simon, Moore, and Sullivan [1993]	.001	———	———	———
Altonji and Card [1991]	-.19 to -.09	-.31 to -.22	-.15 to -.05	———
Winegarden and Khor [1991]	———	.01	-.003	———
Muller and Espenshade [1985]	———	———	-.01	———
Grossman [1982]	-.10 to -.04	———	———	———

These figures correspond to the cross-sectional results. The specification of the wage/earnings and employment regressions vary across studies. Simon, Moore, and Sullivan [1993], Winegarden and Khor [1991], and Muller and Espenshade [1985] use the unemployment rate as the dependent variable.

mated. All models regress the log of weekly wages on a set of human capital variables, regional dummies, controls for occupation and industry, and a variable controlling for the proportion of immigrants. A variable to control for the net migration of natives is added to the second model. These models are then run separately for native whites, blacks, and Hispanics.

The results suggest that the estimates of the impact of immigration on native wages contain omitted variable bias when controls for net migration are not included. A Hausman Specification Test was used to test the hypothesis that the immigration coefficients are significantly different when controls for net migration are added. The results revealed that the coefficients are significantly different in all the wage regressions, except for less-educated blacks.

Among those with lower levels of education, the immigration coefficient is positive for whites and blacks, and negative for Hispanics. However, when net migration controls are added, the coefficient remains positive only for blacks. A 10 percent increase in the share of immigrants is associated with an increase in weekly wages for blacks of 4.2 percent without the net migration control and 4.6 percent with the net migration control. For whites, however, a positive coefficient on the immigration variable becomes negative when the net migration control is added. Without controlling for net migration, a 10 percent increase in the share of immigrants is associated

TABLE 2
Estimated Effects of Immigration on Male Native-born Wages, 16-64
(Less than High School Education)

Dependent Variable: Log Weekly Wages	Explanatory Variables: I/POP	NM/POP
Whites:[c]		
(1)	.1226	———
	(1.59)	
(2)	-.0121	-.7485[a]
	(0.14)	(3.64)
Blacks:[d]		
(1)	.4199[a]	———
	(2.64)	
(2)	.4559[a]	.1970
	(2.86)	(1.90)
Hispanics:[e]		
(1)	-.0130	———
	(0.16)	
(2)	-.0989	-.4223[a]
	(1.13)	(3.44)

T-ratios in parentheses. All equations estimated with the natural log of 1989 weekly wages as the dependent variable. Samples restricted to non-student civilian wage and salary workers reporting positive labor earnings and weeks worked in 1989. The standard deviation of the weekly wage in 1980, the average yearly income from public assistance in 1980, the fraction of immigrants in 1980, and its square are used to predict the concentration of immigrants in 1990 and the rate of net migration. A Generalized Method of Moments Specification Test was used to ensure the validity of the instruments.

a. Significant at the 0.01 level.
b. Significant at the 0.05 level.
c. The sample size is 16,640. A Hausman Test revealed that *I/POP* should be treated as exogenous and *NM/POP* should be instrumented. The partial R^2 for the internal migration regression is .0890; the partial F-statistic is 469.90.
d. The sample size is 3,931. A Hausman Test revealed that both *I/POP* and *NM/POP* should be treated as exogenous.
e. The sample size is 10,854. A Hausman Test revealed that *I/POP* should be treated as exogenous and *NM/POP* should be instrumented. The partial R^2 for the internal migration regression is .1649; the partial F-statistic is 332.93.

with an increase in weekly wages of 1.2 percent, but with the net migration control, the same increase in immigration is associated with a 0.1 percent decrease in weekly wages. Similarly, for Hispanics, a 10 percent increase in the share of immigrants is correlated with a 0.1 percent decrease in weekly wages without controlling for net migration, but once this control is added an identical increase in immigration is associated with a 1.0 percent decrease in weekly wages.[8]

For those with at least a high school education, increased immigration tends to be positively associated with weekly wages. But this positive correlation becomes weaker

TABLE 3
Estimated Effects of Immigration on Male Native-born Wages, 16-64 (High School Graduates +)

Dependent Variable: Log Weekly Wages	Explanatory Variables: I/POP	NM/POP
Whites:[c]		
(1)	.1208[a]	———
	(4.36)	
(2)	.0600[b]	-.6382[a]
	(2.07)	(9.46)
Blacks:[d]		
(1)	.1235	———
	(1.38)	
(2)	.0584	-.6756[a]
	(0.64)	(3.64)
Hispanics[e]:		
(1)	.0217	———
	(0.26)	
(2)	-.3083[a]	-.9793[a]
	(2.69)	(4.66)

T-ratios in parentheses. All equations estimated with the natural log of 1989 weekly wages as the dependent variable. Samples restricted to non-student civilian wage and salary workers reporting positive labor earnings and weeks worked in 1989. The standard deviation of the weekly wage in 1980, the average yearly income from public assistance in 1980, the fraction of immigrants in 1980, and its square are used to predict the concentration of immigrants in 1990 and the rate of net migration. A Generalized Method of Moments Specification Test was used to ensure the validity of the instruments.

a. Significant at the 0.01 level.
b. Significant at the 0.05 level.
c. The sample size is 99,866. A Hausman Test revealed that I/POP should be treated as exogenous and NM/POP should be instrumented. The partial R^2 for the internal migration regression is .0901; the partial F-statistic is 2,658.40.
d. The sample size is 9,497. A Hausman Test revealed that I/POP should be treated as exogenous and NM/POP should be instrumented. The partial R^2 for the internal migration regression is .0739; the partial F-statistic is 221.03.
e. The sample size is 9,386. A Hausman Test revealed that I/POP and NM/POP should be instrumented. The partial R^2 for the immigration and internal migration regressions are .9354 and .0878, respectively; the partial F-statistics are 25,085.80 and 207.23, respectively.

when net migration controls are added, and for Hispanics, a positive (but insignificant) coefficient on the immigration variable becomes negative (and significant) when the net migration control is added. The results imply that a 10 percent increase in the fraction of immigrants is associated with a 1.2 percent increase in weekly wages for whites and blacks when the net migration control is excluded, but this falls to 0.6 percent when the net migration control is added. For Hispanics, a 10 percent increase in the share of immigrants is associated with a 0.2 percent increase in weekly

TABLE 4
Estimated Effects of Immigration on Male Native-born Employment, 16-64 (Less than High School Education)

Dependent Variable: Log Weekly Wages	Explanatory Variables: I/POP	NM/POP
Whites:[c]		
(1)	-.1889[a]	———
	(3.25)	
(2)	-.3176[a]	-.5727[a]
	(4.23)	(3.01)
Blacks:[d]		
(1)	-.2325	———
	(1.67)	
(2)	-.2411	-.0423
	(1.72)	(0.53)
Hispanics:[e]		
(1)	-.0377	———
	(0.54)	
(2)	-.0900	-.2567[a]
	(1.24)	(2.55)

T-ratios in parentheses. All equations estimated with the natural log of 1989 weekly wages as the dependent variable. Samples restricted to non-student civilian wage and salary workers reporting positive labor earnings and weeks worked in 1989. The standard deviation of the weekly wage in 1980, the average yearly income from public assistance in 1980, the fraction of immigrants in 1980, and its square are used to predict the concentration of immigrants in 1990 and the rate of net migration. A Generalized Method of Moments Specification Test was used to ensure the validity of the instruments.

a. Significant at the 0.01 level.
b. Significant at the 0.05 level.
c. The sample size is 16,640. A Hausman Test revealed that *I/POP* and *NM/POP* should be instrumented. The partial R^2 for the immigration and internal migration regressions are .9394 and .0890, respectively; the partial F-statistics are 37,599.69 and 496.90, respectively.
d. The sample size is 3,931. A Hausman Test revealed that *I/POP* and *NM/POP* should be treated as exogenous.
e. The sample size is 10,854. A Hausman Test revealed that *I/POP* should be treated as exogenous and *NM/POP* should be instrumented. The partial R^2 for the internal migration regression is .1649; the partial F-statistic is 332.93.

wages prior to the addition of the net migration control, but an identical increase in immigration is associated with a 3.1 percent decrease in weekly wages for this group after the addition of the net migration control.

In addition to the immigration coefficients responding as expected to the inclusion of controls for net migration, the net migration coefficients are significant and appear with the anticipated negative sign (except for less-educated blacks). Among

TABLE 5
Estimated Effects of Immigration on Male Native-born Employment, 16-64 (High School Graduates +)

Dependent Variable: Log Weekly Wages	Explanatory Variables: I/POP	NM/POP
Whites:[c]		
(1)	-.0845[a] (5.29)	———
(2)	-.0955[a] (5.76)	-.1161[a] (3.05)
Blacks:[d]		
(1)	-.1100 (1.61)	———
(2)	-.1084 (1.58)	.0170 (0.39)
Hispanics:[e]		
(1)	.0190 (0.36)	———
(2)	.0272 (0.49)	.0321 (0.88)

T-ratios in parentheses. All equations estimated with the natural log of 1989 weekly wages as the dependent variable. Samples restricted to non-student civilian wage and salary workers reporting positive labor earnings and weeks worked in 1989. The standard deviation of the weekly wage in 1980, the average yearly income from public assistance in 1980, the fraction of immigrants in 1980, and its square are used to predict the concentration of immigrants in 1990 and the rate of net migration. A Generalized Method of Moments Specification Test was used to ensure the validity of the instruments.

a. Significant at the 0.01 level.
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c. The sample size is 99,866. A Hausman Test revealed that I/POP can be treated as exogenous and NM/POP should be instrumented. The partial R^2 for the internal migration regression is .0901; the partial F-statistic is 2,658.40.
d. The sample size is 9,497. A Hausman Test revealed that I/POP and NM/POP should be treated as exogenous.
e. The sample size is 9,386. A Hausman Test revealed that I/POP and NM/POP should be treated as exogenous.

those with lower levels of education, the results imply that a 10 percent increase in the rate of native net migration is associated with a decrease in weekly wages of 7.5 percent for whites and 4.2 percent for Hispanics, and an increase in weekly wages of 2.0 percent for blacks. For those with at least a high school education, a 10 percent increase in the rate of net migration is correlated with a decrease in weekly wages of 6.4 percent for whites, 6.8 percent for blacks, and 9.8 percent for Hispanics.

The Effect of Immigration on Employment

A set of similar regressions is also used to estimate the effect of immigration on employment. In these models, the log of weeks worked substitutes for the log of weekly wages as the dependent variable. Tables 4 and 5 present a summary of the immigration and internal migration coefficients for less-educated and higher-educated natives, respectively. These regressions are also run separately for whites, blacks, and Hispanics.

A Hausman Specification Test is also used with the employment regressions to test the hypothesis that the immigration coefficients are statistically different when net migration controls are added. The results indicate that only the immigration coefficients for less-educated whites are significantly different with the net migration control added to the regression.

With the exception of higher-educated Hispanics, the immigration coefficient is negative for all racial and educational groups. However, it is only significant in the regression for whites. Although the immigration coefficients are insignificant for other racial groups, the negative coefficients become larger (in absolute terms) when the net migration control is added. Nevertheless, including the net migration variable does not change the significance of any of the immigration coefficients in the employment regressions. Therefore, a significant negative relationship between immigration and employment remains only for whites. For whites with lower levels of education, a 10 percent increase in the ratio of immigrants is associated with a decrease in weeks worked of 1.9 percent without controlling for net migration and 3.2 percent with the control for net migration. Similarly, for whites with at least a high school education, a 10 percent increase in the share of immigrants is associated with a 0.8 percent decrease in weeks worked without internal migration controls and a 1.0 percent decrease with internal migration controls.

The net migration control also seems to have less explanatory power in the employment regressions. The net migration coefficients are significant only in the regressions for whites and less-educated Hispanics. A 10 percent increase in net migration is associated with a decrease in weeks worked of 5.7 percent for less-educated whites, 2.6 percent for less-educated Hispanics, and 1.2 percent for higher-educated whites.

Summary

The results indicate that the wage regressions are much more responsive than the employment regressions to the inclusion of the net migration variable. Once the control for internal migration is included, many of the immigration coefficients in the wage regressions change significantly, but only the coefficient for less-educated whites changes significantly in the employment regressions. The positive effect of immigration on the wages of higher-educated whites and blacks, for example, remains, but the magnitude is 50 percent smaller. Also, the negative correlation between immigration and weeks worked is 70 percent larger for less-educated whites when the net migration control is added. And perhaps the most evident impact of controlling for net migration is seen in the wage regression for higher-educated Hispanics where a positive but insignificant association between immigration and wages becomes negative and significant.

The net migration control does not only influence the immigration coefficient, but also frequently appears with the expected sign and significance. The net migration coefficient is negative and significant in all the wage regressions, except for less-educated blacks. In the employment regressions, the net migration coefficient is not consistently significant, but appears with the expected negative sign, except for the regressions of higher-educated blacks and Hispanics. Also as expected, the negative effects on wages and employment associated with native internal migration are larger than the negative effects associated with increases in immigration.

CONCLUSIONS

This paper has investigated the impact of immigration on wages and employment of natives. While previous work has examined these issues using a variety of econometric techniques (e.g., instrumental variables, first-differencing, etc.) they have all contained omitted variable bias. The principle source of this bias is addressed in this paper by directly controlling for the net migration of natives in the estimates. First, the census survival method was used to obtain estimates of male net migration for all compatible MSAs in the 1980 and 1990 censuses. Then, these estimates are applied to the wage and employment regressions to mitigate omitted variable bias and generate better estimates of the effect of immigration on labor market outcomes.

Two noteworthy results are consistent with the common neoclassical view of labor substitution. First, the larger negative effects of immigration when controls for net migration are added suggests that the arguments made by Borjas [1994] and Borjas, Freeman, and Katz [1996] are valid. It seems, therefore, that previous work that has ignored changes in native-born labor supply across regional labor markets has underestimated the effect of immigration on wages and employment. Second, the negative effects of native net migration are relatively larger than the effects of immigration. This is perhaps due to the fact that native-born migrants are more substitutable for non-migrant natives than immigrants.

Some results, however, are not easily explained. For example, the ambiguous effects of immigration on less-educated workers does not seem to be consistent with evidence of larger low-skilled immigration flows [Borjas, 1994]. This is especially true for less-educated blacks and Hispanics who are presumed to be the most substitutable for recent immigrants. Furthermore, there are also negative effects on employment of higher-educated white natives and wages of higher-educated Hispanic natives. These groups are often presumed to be complements to low-skilled labor.

This study has addressed a major short-coming of previous work on the impact of immigration on the labor market, but other issues must be examined before robust conclusions can be drawn. The significance of the net migration controls suggests that other forms of labor market migration on the part of natives may also be of vital importance in eliminating omitted variable bias. In this paper, net migration measures physical movements from one MSA to another, but labor supply shifts can also result if natives change their labor force status or move from one sector (e.g., the wage/salary sector) to another (e.g., the self-employed sector) without any change in residence. Furthermore, the existence of some negative (as well as ambiguous) effects of immigration on both less-educated and higher-educated natives may imply, as Borjas [1994] suggests, that controlling for the skill distribution of immigrants across labor markets will provide more precise estimates of the impact of immigration. Or, if immigrants are assimilating over time into the U.S. labor market, perhaps controlling for the timing of immigration, as LaLonde and Topel [1991] point out, is of critical importance. These issues must be addressed in future research in order to accurately determine which native groups are affected by immigration and, more specifically, by what type of immigration.

NOTES

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1. See Borjas [1994] and Friedberg and Hunt [1995] for a review of this literature.
2. This procedure was first introduced by Hamilton [1934] to measure rural-urban migration in Tennessee. It has since become a technique adopted by demographers and a general description of the method can be found in Shryock, Siegel, and Associates [1976].
3. This formula treats migration abroad as equivalent to death and survival rates are assumed to be independent of occupational status and region of residence [Carter and Sutch, 1996].
4. In order to calculate potential experience, the education categories must be converted to years of education. Individuals who reported educational attainment of less than 1st grade are assigned zero years of schooling, 1st to 4th grade are assumed to have 2.5 years of schooling, and those with 5th to 8th grade attainment are assumed to have 6.5 years of schooling [Schoeni, 1997]. Those with some college, but no degree, an associate degree, a bachelor's degree, a master's degree, and a professional/doctorate degree are assumed to have 13, 14, 16, 18, and 20 years of schooling, respectively.
5. This may be the result of internal native migrants being primarily white workers that do not base their migration decisions on the wages and employment levels of black workers in the receiving areas.
6. This may be due to greater concern on the part of immigrants to locate in areas where a large existing concentration of immigrants already exists [Bartel, 1989; Zavodny, 1997] and where they are more likely to build ethnic and socioeconomic ties.
7. The geographic regions of metropolitan areas were matched across the 1980 and 1990 samples using a procedure similar to that adopted by Bound, Jaeger, Loeb, and Turner [1997] in their working paper "Coding Geographic Areas Across Census Years: Creating Consistent Definitions of Metropolitan Areas."
8. The coefficients are not statistically significant for whites and Hispanics.

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