IMMIGRATION AND THE ECONOMIC STATUS OF AFRICAN-AMERICAN MEN

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Abstract

The employment rate of black men, and particularly of low-skill black men, fell precipitously from 1960 to 2000. At the same time, the incarceration rate of black men rose markedly. This paper examines the relation between immigration and these trends in black employment and incarceration. Using data drawn from the 1960-2000 U.S. Censuses, we find a strong correlation between immigration, black wages, black employment rates, and black incarceration rates. As immigrants disproportionately increased the supply of workers in a particular skill group, the wage of black workers in that group fell, the employment rate declined, and the incarceration rate rose. Our analysis suggests that a 10-percent immigrant-induced increase in the supply of a particular skill group reduced the black wage by 4.0 percent, lowered the employment rate of black men by 3.5 percentage points, and increased the incarceration rate of blacks by almost a full percentage point.

IMMIGRATION AND THE ECONOMIC STATUS OF AFRICAN-AMERICAN MEN

George J. Borjas, Jeffrey Grogger, and Gordon H. Hanson*

> After a wave of raids by federal immigration agents on Labor Day weekend, a local chicken-processing company called Crider Inc. lost 75% of its mostly Hispanic 900-member work force. The crackdown threatened to cripple the economic anchor of this fading rural town. But for local African-Americans, the dramatic appearance of federal agents presented an unexpected opportunity. Crider suddenly raised pay at the plant. An advertisement in the weekly Forest-Blade newspaper blared "Increased Wages" at Crider, starting at $7 to $9 an hour—more than a dollar above what the company had paid many immigrant workers. (*The Wall Street Journal*, January 17, 2007)

I. Introduction

The employment rate of African-American men—defined as the fraction of weeks worked during a calendar year by the typical black male—fell from 74.9 percent in 1960 to 67.9 percent in 2000.[1] This drop stands in sharp contrast to the slight decline observed among white men during that period, from 87.0 to 85.2 percent. The racial employment gap widened even more for low-skill persons: the employment rate of black high school dropouts fell by 30 percentage points, from 72.1 to 42.1 percent, as compared to an 18 percentage point drop for white high school dropouts, from 82.7 to 64.3 percent.

The decline in labor market participation among black men was accompanied by a rapid increase in the number of black men in correctional institutions. As recently as 1980, only 0.8

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[1] Throughout the paper, the "employment rate" gives the average fraction of weeks worked during the calendar year prior to the Census (i.e., the ratio of weeks worked, including zeros, to 52). The "incarceration rate" gives the fraction of persons who are institutionalized at the time of the Census. The data will be described in greater detail in the next section.

percent of black men (and 1.4 percent of black high school dropouts) were incarcerated. By 2000, 9.6 percent of black men (and 21.2 percent of black high school dropouts) were incarcerated.[2]

A large academic literature examines these trends. One strand of the literature emphasizes the impact of government programs, such as the Social Security disability program or the minimum wage, in driving black men out of the labor market (Bound and Freeman, 1992; Bound, Schauenbaum, and Waidmann, 1995; Parsons, 1980; Stern, 1989; and Welch, 1990). Another focuses on the possibility that the changes in the wage structure, and particularly the decline in the real wage of low-skill workers, may have discouraged low-skill black men from entering the labor market (Juhn, 1992, 2003). Finally, some analysts note that the trend in black incarceration rates was shaped by the crack epidemic of the 1980s and early 1990s. The invention of crack cocaine in the early 1980s represented a technological innovation that greatly increased the profitability of the cocaine trade. As illegal drug markets expanded, crime rose (Grogger and Willis, 2000). Many jurisdictions responded by increasing both drug arrests and the likelihood of imprisonment for convicted arrestees (Boggess and Bound, 1997). Crack and its consequences were concentrated in African-American communities, in part because pre-existing black gangs acted to profit from the expanding drug trade (Fryer et al, 2005).

Although immigration has disproportionately increased the number of low-skill workers in the United States, only a few studies (Altonji and Card, 1991; LaLonde and Topel, 1991) have sought to estimate the effect of immigration on the wages of African Americans, who are disproportionately represented among the low skilled.[3] This paper extends the literature by

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[2] Western and Pettit (2000) show that ignoring the prevalence of incarceration rates provides a very misleading picture of employment trends in the black population.

[3] We are aware of two studies that suggest a potential link between immigration and incarceration. Using cross-section data from the 1980s, Butcher and Piehl (1998) find that metropolitan areas with larger immigrant

examining the relation between immigration and black wages, employment, and incarceration. We use data drawn from the 1960-2000 U.S. Censuses. The data reveal a strong correlation between immigration and black wages, black employment rates, and black incarceration rates. As immigrants disproportionately increased the supply of workers in a particular skill group, we find a reduction in the wage of black workers in that group, a reduction in the employment rate, and a corresponding increase in the incarceration rate.

Our study suggests that a 10-percent immigrant-induced increase in the supply of a particular skill group is associated with a reduction in the black wage of 3.4 percent, a reduction in the black employment rate of 7.9 percentage points, and an increase in the black institutionalization rate of 1.8 percentage points. Among white men, the same ten-percent increase in supply reduces the wage by 4.3 percent, but has much weaker employment and incarceration effects: a 2.9 percentage point reduction in the employment rate and a 0.3 percentage point increase in the incarceration rate. It seems, therefore, that black employment and incarceration rates are more sensitive to immigration than those of whites.

These findings can obviously generate a great deal of controversy in the immigration debate and can be easily misinterpreted. As a result, we are extremely cautious in both the presentation and interpretation of the evidence. Although we have attempted to control for other factors that may account for the large shifts in black employment and incarceration rates over the four-decade period that we examine, it should be obvious that no study can control for all possible factors. It is equally important to emphasize that although the evidence suggests that immigration played a role in generating these trends, much of the decline in employment or

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populations had higher crime rates, but this relationship disappears once they control for the demographic characteristics of the underlying populations. In their replication of the Borjas (2003) study, Raphael and Ronconi (2005) claim that adding incarceration rates as an explanatory variable in a regression of wages on immigrant shares attenuates the wage impact of immigration in national-level data. The Raphael-Ronconi empirical exercise, increase in incarceration in the black population remains unexplained. Put differently, immigration seems to have an effect and this effect seems to be numerically important, but we would have witnessed a sizable decline in black employment and the concurrent increase in black incarceration rates even if there had been no immigration in the past few decades.

II. Data and Descriptive Trends

Our data are drawn from the 1960, 1970, 1980, 1990 and 2000 Integrated Public Use Microdata Samples (IPUMS) of the decennial Censuses. The 1960 file represents a 1 percent sample of the U.S. population, the 1970 file represents a 3 percent sample, and the 1980 through 2000 files represent 5 percent samples. The empirical analysis is restricted to men aged 18 to 64. The Data Appendix describes the construction of the sample extracts and variables used in the study.

We define an immigrant as someone who is either a noncitizen or a naturalized U.S. citizen. All other persons are defined as natives. Similarly, we use information contained in the census race variable to classify persons as "black" or "white." Unless otherwise specified, persons whose race is neither black nor white are excluded from the analysis.

As in Borjas (2003), skill groups are defined in terms of both educational attainment and years of labor market experience. We classify workers into four distinct education groups: (1) high school dropouts (workers who have less than 12 years of completed schooling); (2) high school graduates (workers who have exactly twelve years of schooling); (3) workers who have some college (thirteen to fifteen years of schooling); and (4) college graduates (workers who have at least sixteen years of schooling).

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however, may have the logic backwards: shifts in incarceration rates are likely endogenous and may be partly caused by immigration.

We group workers into a particular years-of-experience cohort by using potential years of experience. We assume that age of entry into the labor market is 17 for high school dropouts, 19 for high school graduates, 21 for persons with some college, and 23 for college graduates, and then calculate years of experience accordingly. The analysis is restricted to persons who have between 1 and 40 years of experience. Workers are aggregated into five-year experience groupings (i.e., 1 to 5 years of experience, 5 to 10 years, and so on) to capture the notion that workers who have roughly similar years of experience are more likely to affect each other's labor market opportunities than workers who differ significantly in their work experience. The resulting data set contain 160 observations (4 education groups, 8 experience groups, and 5 years).

The cell corresponding to educational attainment (*e*), experience level (*x*), and calendar year (*t*) defines a skill group at a point in time for the U.S. labor market. The immigrant supply shock experienced by a particular skill group is given by:

$$(1) \qquad p_{ext} = \frac{M_{ext}}{M_{ext} + N_{ext}},$$

where M_{ext} gives the total number of work hours provided by immigrants in the particular skill group; and N_{ext} gives the corresponding number of work hours provided by native workers.[4] The variable p_{ext} then gives the immigrant share (i.e., the fraction of total supply that is foreign-born).

Figure 1 summarizes some of the (well-known) information regarding trends in the immigrant share, by education and experience group, for the 1960-2000 period. The fraction of the (hours-weighted) workforce that is foreign-born increased most for high school dropouts.

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[4] The counts of immigrants and natives for skill group (*e, x, t*) include persons of all races.

Within any given census year, the immigration-induced increase in supply is largest for workers with lower levels of labor-market experience, due to the preponderance of young adults in the immigrant population. By 2000, 13.8 percent of the male workforce and 40.4 percent of high school dropouts were foreign-born. Among high-school dropouts with 10 to 15 years of experience, 48.0 percent of the workforce was foreign born.

It is useful to begin by illustrating racial differences in national-level trends in employment and incarceration across race, education, and experience groups.[5] The top panel of Figure 2 reports the education-and-experience-specific trends for the black employment rate, while the bottom panel presents the corresponding figure for white men. As noted above, the employment rate is defined as the average fraction of weeks worked during the preceding calendar year (including non-workers). Both figures are drawn to the same scale so that the large racial differences can be grasped easily. Since employment rates and incarceration rates tend to be lower for men in their mid-50s or older, our discussion focuses on the trends for those with up to 30 years of labor market experience.

In 1960, the employment rate hovered around 70 to 75 percent for black workers who were high school dropouts and had between 5 and 30 years of experience. By 2000, the employment rate for these black high school dropouts had fallen to around 40 percent. In contrast, the employment rate of black college graduates with 5 to 30 years of experience hovered around 85 to 90 percent in 1960 and remained in that range by 2000.

The bottom panel of Figure 2 shows the corresponding trend for white men. As with blacks, there has been a decline in employment propensities for the least educated workers, but

[5] To provide some context to the discussion, it is instructive to report the share of black and white workers in each of the education groups. In 2000, 18.8 (8.9) percent of blacks (whites) were high school dropouts, 43.7 (33.2) percent were high school graduates, 26.4 (29.0) percent had some college, and 11.1 (28.9) percent were college graduates.

the decline is modest relative to that seen in the black population. Among the most educated whites, average employment rates remain very high for all but the oldest workers. For college graduates with up to 30 years of experience, the average employment rate was essentially flat during the four-decade period, at around 90 to 95 percent. Among white high school dropouts with 5 to 30 years of experience, however, the average employment rate fell from around 85 percent in 1960 to around 68 percent in 2000, about a 17 percentage point drop. This is a large and important decline, but it is much smaller than the 30 to 35 percentage point drop observed among black high school dropouts with similar levels of work experience.

The rapid disappearance of a large segment of black high school dropouts from the workforce was accompanied by a large increase in the number of black high school dropouts in the institutionalized population. We use information on residence in group quarters available in the decennial censuses to enumerate the number of persons in institutions. These institutions include jails, prisons, and mental hospitals.[6] For young men, the 1980 Census shows that the majority of persons institutionalized are, in fact, incarcerated. Furthermore, the growth in institutionalization in Census data closely tracks the growth in incarceration apparent in Department of Justice data from correctional facilities (Western and Pettit, 2000). For expositional convenience, therefore, we will refer to the fraction of persons institutionalized as the "incarceration rate."

Figure 3 presents the trends in the incarceration rate, by race, education, and experience, over the 1960-2000 period. We again use the same scale in the two graphs so that the very large racial differences can be easily seen. The average incarceration rate among white male high school dropouts with 1 to 30 years of experience increased from around 2 percent in 1960 to between 5 and 10 percent in 2000. For whites with at least a high school diploma, the incarceration rate remained small even by 2000.

In contrast, the incarceration rate for black men increased rapidly beginning after 1980 for all groups except college graduates. Among high school dropouts with 1 to 30 years of experience, for example, the incarceration rate hovered around 5 to 7 percent in 1960. By 2000, however, some of the groups of younger black high school dropouts had astoundingly high incarceration rate. The incarceration rate of black high school dropouts with 5 to 15 years of experience had increased to around 35 percent.

This paper examines if these trends are related to the increases in immigration experienced by the specific skill cohort at a particular point in time. To visually illustrate the nature of this link, Figure 4 presents a scatter diagram relating decadal changes in the immigrant share and decadal changes in employment rates for blacks and whites, after removing decade effects. Figure 5 presents the corresponding scatter diagrams relating decadal changes in the immigrant share and decadal changes in incarceration rates. By removing decade effects, we control for features of the economic environment that are common to all education and experience groups in any given decade, but that might vary over time.

The graphical evidence is striking. Each point in the scatter diagram in Figure 4 represents the change in employment rate for an (e, x, t) cell and the corresponding change in the immigrant share of the workforce for that cell. It is evident that there is a negative correlation between changes in employment propensities and the immigrant share, and that the correlation is stronger for black men. Similarly, Figure 5 shows a corresponding positive correlation between changes in incarceration rates and the immigrant share, with the correlation again being stronger

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[6] Although it is possible to differentiate between correction facilities and mental hospitals before 1990, there is no such information in the 1990 and 2000 census data. To keep the variable definitions constant over time,

for black men. The remainder of this paper examines if these correlations persist after we control for other factors that affected the trends in male employment and incarceration propensities over this time period.

III. Theory

To understand how immigration could reduce employment and increase incarceration among native-born persons, with possibly larger effects among African-Americans, consider a two-sector model of a national labor market. Native labor consists of black and white workers, who are perfectly mobile between a formal sector (i.e., the "market" sector) and a sector dedicated to crime.

In this section, we consider one particularly simple specification of this model. Suppose the market sector employs both natives and immigrants, with all workers being perfect substitutes in terms of their contribution to output.[7] The crime sector employs only native workers. Further, the demand for labor in the crime sector is race-specific, with black and white workers having separate crime production functions. Finally, native labor supply to the market and crime sectors is endogenous, with individuals choosing to increase leisure and participation in crime and to reduce market work as the wage in the market sector falls.

We use this framework to investigate the consequences of an exogenous shift in the supply of immigrant labor. We are interested in determining whether immigration induces some native workers to exit market employment and engage in other activities. The mechanism through which this might occur is straightforward. A positive immigrant supply shift puts downward pressure on the wage in the market sector, causing native workers to substitute out of

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we focus on the number of persons who are in institutions.

market work and into either crime or leisure. One can think of the model as a general-equilibrium extension of Gronau (1977), in which individuals allocate time between work, leisure, and home production. Our framework reinterprets home production as crime (as in Grogger, 1998) and endogenizes the wage.[8]

This specification of the model relies on three key assumptions. First, we assume that all workers (i.e., immigrants, black natives, and white natives) are perfect substitutes in the market sector. Although the assumption of perfect substitutability is not essential for deriving our theoretical results, it greatly simplifies the analysis.[9] More important, we will test for the empirical validity of this assumption in the next section.

We also assume that immigrants do not participate in crime. One could generalize the model to allow immigrant labor to be employed in either the market or crime sector. Our results would hold as long as the elasticity of labor demand in crime is larger for natives than immigrants (or, alternatively, if the elasticity of substitution between native and immigrant labor is higher in market employment than in crime). Since criminal penalties are larger for immigrants than natives (for non-citizens, the penalties for criminal activity are incarceration *and* possible deportation), it seems reasonable to assume immigrants are less likely than natives to substitute into crime in response to a negative wage shock. In fact, relative to observationally equivalent natives, immigrants are much less likely to be incarcerated (Butcher and Piehl, 1998, 2000).

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[7] The race of immigrant workers is left unspecified.

[8] The resulting framework is similar to the specific-factors model of a small open economy (e.g., Feenstra, 2004), extended to allow for the endogenous supply of labor to wage employment.

[9] We also solved a model in which black, white, and immigrant labor are imperfect substitutes in the production of formal-sector output. All of the qualitative results of the simpler model presented here carry through to the more general framework (one minor difference is that imperfect substitutability produces slightly different estimating equations than those reported below in (6a)-(6d)).

Finally, we assume that there are race-specific crime production functions, effectively implying that black and white criminals tend to operate in separate markets.[10] This assumption allows immigration to have different effects by race. Black men, by virtue of being relatively concentrated in inner cities, may have more opportunities to engage in criminal activity. Grogger (1998) finds that black men are more likely to participate in crime than white men even after controlling for alternative labor market options. Further, black-white differences in criminal propensities may have been exacerbated by the advent of crack cocaine. Fryer et al (2005) argue that pre-existing gang organizations, which controlled street corners and other outdoor spaces in many urban areas, gave blacks an advantage in creating and controlling crack distribution networks.[11] Other evidence suggests criminal gangs tend to be organized along racial lines and operate in spatially segmented markets (Venkatesh, 1997; Grogger and Willis, 2000).

Let $L_{fs} = N_{bfs} + N_{wfs} + M_s$, where N_{bfs} denotes employment of native black workers in the (formal) market sector who have skill s; N_{wfs} denotes the corresponding employment of native white workers; and M_s denotes the corresponding number of immigrants. For workers of race i and skill group s, the wage in the market sector is:

(2) $$w_{is} = X_{fs}\left(1-\delta_i\right)(L_{fs})^{\eta_f},$$

where X_{fs} is a labor demand shifter for the market sector; δ_i is a parameter that captures preferences for discrimination on the part of employers, with $\delta_w = 0$ and $1 > \delta_b \geq 0$, so that black

[10] This assumption, like the other two, simplifies the discussion but is not essential for the results. What is essential for immigration to have different effects on black and white workers is that the elasticity of labor demand in crime differs between the two groups.

[11] Responses to a survey of 27 large-city police chiefs administered by Grogger and Willis (2000) indicated that crack was concentrated among blacks or minorities in all but four jurisdictions.

workers may face a lower market wage as a result of discrimination; and $\eta_f < 0$ is the inverse of the labor demand elasticity in the market sector (or the factor price elasticity in this simple framework). Equation (2) assumes that all workers in skill group *s* are perfect substitutes in terms of their contribution to market sector output.

The marginal product of labor in the crime sector for workers in racial group *i* with skill *s* is:

$$(3) \qquad w_{is} = X_{ics}\,(N_{ics})^{\eta_{ic}},$$

where X_{ics} is a demand shifter for criminal activity, N_{ics} gives employment of native workers of race *i* and skill group *s* in crime, and $\eta_{ic} < 0$ is the inverse of the labor demand elasticity in the crime sector.

Black-white wage differences are determined by the extent of discrimination in the market sector, with $w_b/w_w = (1 - \delta_b) \leq 1$.[12] Inter-sectoral labor mobility transmits the discrimination-driven market racial wage gap to the crime sector. The demand shifters, X_{fs} and X_{ics}, embody capital, TFP, and the output price in each sector. Our empirical analysis will allow for changes in sectoral demand shifters by controlling for race-specific changes in the returns to skill over time.

The supply of labor to paid employment (i.e., employment in either the market or crime sectors) is elastic, with the inverse demand for leisure given by:

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[12] For simplicity, we assume that the discrimination coefficient is independent of skills. Relaxing this assumption does not change any of the results of the model.

(4) $$w_{is} = X_{ihs}\,(N_{ihs})^{\eta_{ih}}$$

where X_{ihs} is a leisure demand shift parameter, N_{ihs} gives the number of natives consuming leisure, and $\eta_{ih} < 0$ is the inverse of the demand elasticity for leisure.

Finally, the allocation of native labor to employment in the market sector, employment in the crime sector, and leisure is subject to the constraint:

(5) $$\tilde{N}_{is} = N_{ifs} + N_{ics} + N_{ihs}\,,$$

where $\tilde{N}_{is}$ is the (constant) population of native-born persons of race *i* and skill *s*.

Equations (2)-(5) represent a system of seven equations in seven unknowns. For simplicity, we neglect the impact of immigration on capital accumulation, which would tend to dampen the wage effects of immigration over time.[13] Figure 6 illustrates the equilibrium of the model for black workers of skill group *s*. There is an analogous, and interdependent, set of equilibrium conditions for white workers. The equalization of wages for black workers between the formal sector and the crime sector is shown by the intersection of the two sectoral labor demand schedules at point 1. The allocation of labor to leisure is implicit, since the endogenous leisure allocation defines the value $\tilde{N}_{bs} - N_{bhs} = N_{bcs} + N_{bfs}$, which is total black labor available for employment in either the market or crime sector. Solving for N_{bhs} defines the width of the

[13] We have worked out an extension of the model where the market sector employs both labor and capital in production, with the supply of capital adjusting over time in response to deviations in the return to capital from its long-run rate. As long as complete adjustments in the capital stock are not immediate, the short-run consequences of immigration are qualitatively the same as those implied by the simpler model summarized in this section.

box in Figure 6. The sensitivity of leisure to wages will become apparent once we consider a labor-supply shock due to immigration.

Figure 7 shows the labor-market consequences of an increase in immigrant labor supply. The immediate direct effect is a contraction in the demand for native labor.[14] The contraction in market-sector labor demand puts downward pressure on the market wage, inducing native labor to increase leisure and decrease labor supplied to the market. The increase in leisure implies that labor available for either the market or crime sectors falls from $\tilde{N}_{bs} - N_{bhs}$ to $\tilde{N}_{bs} - N'_{bhs}$, which implies the right vertical axis in Figure 7 shifts to the left, inducing a corresponding leftward shift, from D_c to D'_c, in the demand for labor in the crime sector (whose horizontal position is determined by the position of the right axis).

The net effect of the immigrant labor supply shock is a new equilibrium at point 2, in which the black wage is lower, black employment in the market sector is lower, black employment in the crime sector is higher, and black leisure is higher. Market employment falls because immigrant labor substitutes for black labor; black employment in crime rises because lower market-sector labor demand induces blacks to shift into crime; and black leisure rises because the black wage falls.

The model has similar qualitative predictions for the wage and sectoral distribution of white workers. The model, in addition, yields an important and testable quantitative prediction. Since black and white workers are perfect substitutes in the market sector, the percent impact of immigration on the black and white wage is the same. As long as the discrimination parameter δ_b is invariant to labor-market conditions, immigration changes black and white wages by the same percentage amount, leaving the racial wage differential unchanged. However, racial differences in the demand elasticities of crime and leisure imply that the employment effects of immigration need not have the same magnitude.

Let w_{is}^* be the wage for race group *i* and skill group *s* in the pre-immigration equilibrium, and let N_{fs}^* give the corresponding number of native workers in the market sector at that time $(N_{fs}^* = N_{bfs}^* + N_{wfs}^*)$. We measure the immigrant supply shock by $m_s = M_s / N_{fs}^*$, the immigrant-induced percent increase in labor supply to the market sector. Part 1 of the Mathematical Appendix shows that the race-specific equations relating post-immigration wages, native labor allocations, and the immigrant supply shock are given by:

(6a) $$\ln w_{is} = \ln w_{is}^* + \eta_f \rho \, m_s ,$$

(6b) $$\ln N_{ics} = \ln N_{ics}^* + \frac{\eta_f \rho}{\eta_{ic}} m_s ,$$

(6c) $$\ln N_{ihs} = \ln N_{ihs}^* + \frac{\eta_f \rho}{\eta_{ih}} m_s ,$$

(6d) $$\ln N_{ifs} = \ln N_{ifs}^* - \frac{\eta_f \rho}{\theta_{ifs}} \left(\frac{\theta_{ics}}{\eta_{ic}} + \frac{\theta_{ihs}}{\eta_{ih}} \right) m_s .$$

As shown in the Appendix, the parameter ρ is a positive constant that lies between zero and one and is defined by:

[14] For black workers, part of the inward shift in formal-sector labor demand is mitigated by the exit of white workers from the formal sector. In the graphical analysis, this adjustment is implicit.

$$\rho = \frac{\bar{N}_f}{\bar{N}_f + \frac{\eta_f}{\eta_{bc}}\bar{N}_{bc} + \frac{\eta_f}{\eta_{wc}}\bar{N}_{wc} + \frac{\eta_f}{\eta_{bh}}\bar{N}_{bh} + \frac{\eta_f}{\eta_{wh}}\bar{N}_{wh}}, \tag{7}$$

where $\bar{N}_j$ is the average number of type-*j* natives across skill groups in the pre-immigration equilibrium.

The parameter ρ gives an elasticity-adjusted measure of the market sector participation rate of natives in the pre-immigration equilibrium. Consider, for instance, the special case where the demand elasticities are equal in all activities, so that $\eta_f = \eta_{ic} = \eta_{ih}$. Equation (7) shows that ρ is then exactly equal to the fraction of natives participating in the market sector in the pre-immigration period. It is also worth noting that if the demand for labor in both the crime and leisure sectors is perfectly inelastic (so that the ratios $1/\eta_{ic}$ and $1/\eta_{ih}$ are equal to zero), the parameter ρ is then equal to one. In this extreme case, the relative number of native workers in each of the sectors is effectively fixed.

Equation (6a) implies that more immigration lowers wages ($\eta_f \rho < 0$), with the wage impact being greater the larger the factor price elasticity in the market sector. Two points are worth emphasizing about the wage impact of immigration. First, as noted above, the wage impact is predicted to be the same for black and white workers.[15] Second, the reduced-form regression of the log wage on the immigrant supply shock *m* does *not* identify the factor price elasticity, η_f. Rather, it identifies the product of the factor price elasticity and ρ, the parameter that roughly indicates the sectoral allocation of the native population (up to a linear

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[15] As shown by equation (2), the market sector wage for both black and white workers (who are perfect substitutes in production) is determined by the equilibrium size of the workforce in that sector. As a result, the manner in which native substitution across sectors mitigates the wage consequences of immigration is common to all native workers, regardless of race.

approximation). The parameter ρ equals one when the supply of native labor to the market sector is perfectly inelastic. It is only in this case that the regression coefficient identifies the factor price elasticity. If native labor supply to the market sector is elastic, however, the reduced-form impact of immigration is numerically smaller than the factor price elasticity. The intuition for this result is obvious: Native opportunities to substitute into crime or leisure dampen the impact of immigration on the market wage, relative to the case of inelastic labor supply. Figure 7 illustrates the result. If the demand for leisure were perfectly inelastic the post-immigration equilibrium would be at point 1′, instead of point 2. The fall in the native wage would be larger and the fall in native formal employment would be smaller. We will refer to the product $\eta_f \rho$ as the "reduced-form wage elasticity."

Equation (6b) shows that a larger immigrant supply shock increases the number of natives participating in the crime sector ($\eta_f \rho/\eta_{ic} > 0$), with the impact of immigration being larger the more elastic is the demand for crime labor relative to the demand for formal labor. The immigration-induced change in crime employment for blacks relative to whites depends on the ratio of the elasticities η_{wc}/η_{bc}. Even though the wage impact of immigration is predicted to be the same for blacks and whites, black employment in the crime sector is more responsive to immigration if the elasticity of labor demand in crime is larger for blacks than whites.

Equation (6c) indicates that more immigration is associated with greater native demand for leisure ($\eta_f \rho/\eta_{ih} > 0$), with the impact of immigration on leisure being larger the more elastic is the demand for leisure relative to the demand for formal labor. Similar to participation in crime, the immigration-induced change in leisure for blacks relative to whites depends on the ratio of elasticities η_{wh}/η_{bh}, indicating that black leisure time is more responsive to immigration if the elasticity of demand for leisure is larger for blacks than whites.

Finally, equation (6d) implies that a larger immigrant supply shock is associated with lower native market sector employment, with the impact of immigration being larger the more elastic is the demand for formal labor relative to the demands for crime labor or leisure. The impact of immigration depends on the pre-existing employment shares in the various sectors, where $\theta_{ics} = N^*_{ics} / \tilde{N}_{is}$ (the pre-immigration share of race *i* persons in the crime sector), $\theta_{ihs} = N^*_{ihs} / \tilde{N}_{is}$ (the pre-immigration share of race *i* persons in the leisure sector), and $\theta_{ifs} = N^*_{ifs} / \tilde{N}_{is}$ (the pre-immigration share of race *i* persons in the market sector).[16] If, for expositional convenience, we ignore the skill subscript, equation (6d) implies that the change in market employment for blacks relative to whites is given by the ratio $\theta_{wf}(\theta_{bc}/\eta_{bc} + \theta_{bh}/\eta_{bh})/\theta_{bf}(\theta_{wc}/\eta_{wc} + \theta_{wh}/\eta_{wh})$, which shows that black market employment is more responsive to immigration if the elasticities of demand for crime labor and for leisure are larger for blacks than whites (as long as the market participation rate of whites is at least as high as that of blacks).[17]

This model helps us understand the source of racial differences in the consequences of immigration and motivates why the empirical analysis presented in the subsequent sections allows the impact of immigration on wages, employment, and incarceration rates to differ between black and white men. As we have seen, if the demand for labor in the crime sector is more elastic for blacks than for whites, immigration will have a larger negative impact on black market employment and a larger positive impact on black crime employment.

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[16] Note that none of the coefficients in equations (6a)-(6d) depend on the extent of discrimination in the labor market because the discrimination coefficient is assumed to be constant over time.

[17] The condition for immigration to affect black market sector employment more than for whites depends on initial black-white relative employments in crime and leisure. If the initial black shares of employment in crime and leisure are higher than the white shares ($\theta_{bc} > \theta_{wc}$, $\theta_{bh} > \theta_{wh}$), immigration can then induce relatively larger reductions in market sector employment for blacks even if the black crime and leisure demand functions are less elastic than those for whites.

Our empirical analysis uses data on wage and employment rates for education-experience cohorts by year. Although we do not have data on participation rates in crime, we do have information on incarceration rates for the various groups.[18] These data constraints require that we estimate the reduced-form expressions, as summarized by equations (6a)-(6d), rather than a structural model of sectoral time allocation.

IV. Evidence

The estimating equations implied by the theory built in the assumption of perfect substitution between black and white native workers, as well as perfect substitution between native and immigrant workers. Before proceeding to a discussion of the empirical link between immigration and black economic status, therefore, it is important to test for the validity of these two assumptions using the data set of 160 skill groups (defined by education, experience, and time) introduced in Section II.

Consider a generic two-level nested CES production function (as in Card and Lemieux, 2001), where the first level defines the size of the native-born workforce as a CES aggregate of the number of black (*b*) and white (*w*) workers, and the second level defines output as a function of the (CES-weighted) native-born workforce and immigrants. By equating the wage to the marginal product of labor for each native worker type, it is easy to derive the relative demand function:

$$(8) \qquad \ln\left(w_{bst} / w_{wst}\right) = -\frac{1}{\sigma}\ln\left(N_{bst} / N_{wst}\right) + \frac{1-\sigma}{\sigma}\ln\left(\tau_{bst} / \tau_{wst}\right)$$

[18] The National Longitudinal Survey of Youth (NLSY) has information on participation in criminal activities, but only for a single cross-section.

where σ is the elasticity of substitution between black and white native workers; w_{ist} is the wage of race group i and skill group s at time t; N_{ist} is the total number of manhours provided by the group; and τ_{ist} is a parameter measuring relative efficiency. We proxy the relative efficiency term in equation (8) by introducing vectors of fixed effects indicating education, experience, and time effects, their interactions, and a random error term. The null hypothesis of perfect substitution between black and white native workers states that the coefficient $-1/\sigma$ equals zero.

The first row of the top panel of Table 1 reports the OLS coefficient that examines the extent of substitutability between black and white native labor (i.e., the dependent variable is the log wage ratio between black and white workers and the independent variable is the log ratio of the total number of work hours supplied by black relative to white workers). The results do not provide much support for the hypothesis that black and white workers are imperfect substitutes (within these narrowly defined skill groups). The coefficient is most negative in the specification reported in column 4 of the table (which is the most general specification). In this case, the estimated coefficient is -.045 with a standard error of .027. The implied elasticity of substitution between black and white native workers is 22.2, which for most practical purposes is equivalent to perfect substitution.

One potential problem with the least squares estimates is that the relative size of the black workforce in the right-hand-side of equation (8) may be endogenous. The estimated elasticity of substitution between white and black workers, therefore, may be contaminated by labor supply decisions at both the intensive and extensive margins. We use instrumental variables to correct for the possible endogeneity bias. In particular, we instrument the relative number of manhours worked by blacks with the relative number of men in the particular skill group who are black.[19] Row 2 of the top panel of Table 1 shows that the IV estimates of the elasticity of substitution between black and white workers also provide little evidence that the assumption of perfect substitution between black and white native workers is soundly rejected by the data.

Once we have established that black and white native workers are perfect substitutes, we can then move to the next level of the nested CES system, and derive an analogous relative demand equation that relates the relative log wage of immigrants to the log of the relative supply of immigrants. The bottom panel of Table 1 reports both OLS and IV estimates of the coefficient from regressions that relate the relative wage of immigrant workers to their relative quantity (i.e., the dependent variable is the log wage ratio between immigrant and native workers and the independent variable is the log total hours ratio).[20] Again, there is no evidence to reject the null hypothesis that immigrants and natives are perfect substitutes.[21] For the remainder of the analysis, therefore, we will maintain the assumption that different labor types (within the narrow education-experience categories defined earlier) are perfect substitutes in the formal sector.

Let y_{ext} denote the mean value of a particular labor market outcome for native-born men who have education e, experience x, and are observed at time t. As noted above, we calculated y_{ext} using the sample of natives who are either black or white. The empirical analysis reported in

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[19] More precisely, the instrument is the ratio of the number of black persons in a skill group to the number of white persons in that skill group. The counts of persons in the instrument include both workers and non-workers.

[20] The instrument is given by the ratio of the number of foreign-born persons in a skill group to the total number of native persons in that skill group. The counts of persons in the instrument include both workers and non-workers.

[21] It is worth emphasizing that the literature provides mixed evidence on the extent of substitution between immigrant and native workers, with Jaeger (1996) finding that immigrants and natives are perfect substitutes but Ottaviano and Peri (2006) finding strong complementarities between immigrants and natives. However, Borjas, Grogger, and Hanson (2008) show that the complementarity results in Ottaviano and Peri result from a flaw in the construction of their sample, particularly the classification of currently enrolled junior and senior high school students as high school dropouts. If these students are dropped from the data, the Ottaviano-Peri findings of complementarity vanish.

this section stacks these national-level data across skill groups and calendar years and estimates the following regression model separately by race:

$$(9) \qquad y_{ext} = \theta\, p_{ext} + E + X + T + (E \times T) + (X \times T) + (E \times X) + \varphi_{ext},$$

where E is a vector of fixed effects indicating the group's educational attainment; X is a vector of fixed effects indicating the group's work experience; and T is a vector of fixed effects indicating the time period. The linear fixed effects in equation (9) control for differences in labor market outcomes across schooling groups, experience groups, and over time. The interactions ($E \times T$) and ($X \times T$) control for the possibility that the impact of education and experience changed over time, and the interaction ($E \times X$) controls for the fact that the experience profile for a particular labor market outcome may differ across education groups. The regression specification in (9) implies that the labor market impact of immigration-induced supply shifts is identified using time-variation within education-experience cells. The regressions are weighted by the number of observations used to calculate the dependent variable y_{est}.[22] Finally, the standard errors are clustered by education-experience cells to adjust for possible serial correlation.

We examine the impact of immigration on three distinct outcomes. The alternative dependent variables include: the log weekly earned income, the employment rate, and the incarceration rate.[23] We estimate the employment and incarceration rate regressions using a

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[22] We weigh the cells by the sample size to adjust for differences in precision. An alternative weighting scheme would be to normalize the sum of weights in each cross-section to equal 1, so that each of the censuses "counts" equally in the estimation of the coefficients. This "rescaled" weighting scheme would lead to larger (and more significant) wage effects. For example, the OLS coefficient in the log weekly wage regression reported in Panel A of Table 2 increases from -0.346 (0.137) to -0.617 (0.203) for blacks and from -0.522 (0.254) to -0.672 (0.243) for whites.

[23] Recall that the employment rate is defined by the average fraction of weeks worked during the calendar year prior to the Census (including non-workers).

grouped logit estimator.[24] Let r_{ext} be the relevant employment or incarceration rate for cell (*e, x, t*). The grouped logit estimator is given by the weighted least squares regression of the log odds model:

(9′) $$\ln\left(\frac{r_{ext}}{1-r_{ext}}\right)=\theta^{*}p_{ext}+E+X+T+(E\times T)+(X\times T)+(E\times X)+\varphi_{ext}.$$

To make the results more easily interpretable, we convert the estimated coefficient θ^* (and its standard error) into a marginal impact, which is given by $\theta^{*}\,\overline{r}\,(1-\overline{r})$, where we use the race-specific sample mean of the employment (or incarceration) rate in the calculation.[25]

It is important to emphasize that the incarceration rate is an imperfect measure of participation in crime, as individuals in prison today may have committed crimes several years in the past when different labor market conditions prevailed. To control for lags between shocks to the labor market and changes in the size of the prison population, we report results on incarceration that use either the current share of immigrants in the workforce or the five-year lag of the immigrant share. Together, the contemporaneous immigrant share and the five-year lag bracket the length of the average prison term, which is about two years (Raphael and Stoll, 2005).

Table 2 reports our estimates of the adjustment coefficient θ (or the corresponding marginal impact in the grouped logit regressions). The top panel of the table reports the least

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[24] The use of logit versus a linear probability model leads to relatively similar marginal impacts for the employment rates, but smaller marginal impacts for the incarceration rate. This is not surprising given the clustering of incarceration rates at very small numbers near zero.

squares estimates of the regression model. The first row of the panel reports the results for black men, while the second row reports the results for white men. Consider initially the results when the dependent variable is the mean log weekly earnings of the skill group. The adjustment coefficient θ is -0.346 (with a standard error of 0.137) for blacks, and -0.522 (0.254) for whites. These coefficients are easier to interpret if we convert them into an elasticity that gives the percent change in wages associated with a percent change in labor supply. Let $m_{ext} = M_{ext}/N_{ext}$, or the percentage increase in the labor supply of group (*e*, *x*, *t*) attributable to immigration. We can calculate the reduced-form wage effect (equivalent to the product of parameters $\eta_f \rho_f$ in our theoretical framework) as:

(10) $$\frac{\partial \log w_{ext}}{\partial m_{ext}} = \theta (1 - p_{ext})^2.$$

By 2000, immigration had increased the immigrant share in the total number of hours supplied to the U.S. labor market to 13.8 percent. Equation (10) implies that the reduced-form wage elasticity—evaluated at the mean value of the immigrant supply shift—can be obtained by multiplying θ by approximately 0.74. The reduced-form wage elasticity for weekly earnings is then -0.26 (or -0.346 × 0.74) for blacks, and -0.39 for whites. Put differently, a 10 percent immigrant-induced increase in the number of workers in a particular skill group reduces the wage of that group by 3 to 4 percent.[26] These results closely match the estimated wage impacts

[25] The weighted least squares estimator for grouped logit uses a weight equal to $n_{ext}\hat{r}_{ext}(1-\hat{r}_{ext})$, where the predicted probabilities are calculated from a first-stage unweighted regression on the log odds ratio and n_{ext} is the sample size in the cell.

[26] The regression model in (7) uses the immigrant share, *p*, rather than the (more natural) relative number of immigrants, *m*, as the regressor. The main reason for using *p* as the regressor is that the outcomes used in this paper tend to be nonlinearly related to *m*, and *p* is approximately a linear function of log *m*. Rather than introducing

of immigration across all workers reported in Borjas (2003). Note that the difference in the wage effects between blacks and whites is not statistically significant (t = 0.6), providing additional evidence in favor of the hypothesis that black and white native workers (within narrowly defined skill groups) are perfect substitutes in production.

Columns two through four of Table 2 show the relation between immigration and employment and incarceration rates. There is a strong negative relation between immigration and employment rates and a weaker positive relation between immigration and incarceration rates. A 10 percent increase in supply is predicted to reduce the employment rate of black men by 5.1 percentage points (-0.683×0.74) and that of white men by 1.6 percentage points. Similarly, a 10 percent increase in supply increases the incarceration rate of black men by 1.0 percentage point, but has only a negligible effect on the incarceration rate of white men. Lagged immigration has a roughly similar effect on incarceration rates for blacks and a somewhat larger (and now statistically significant) effect on whites. It seems, therefore, that the impact of immigration at the extensive margin of labor supply is far larger for blacks than for whites.

A potential problem with the least squares estimates is that the immigrant share included in the right-hand-side of equation (9)—that is, the fraction of the total labor supply provided by foreign-born persons—may be endogenous. Although our theoretical model maintained the assumption of inelastic immigrant labor supply, it is important to relax that assumption in the empirical work. We use instrumental variables to correct for the possible endogeneity bias, where the instrument is the immigrant share in the *population*.[27] Panel B of Table 2 reports the estimated IV coefficients. The wage effects from the IV specification are similar to those from

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significant nonlinearity in the regression, we opted for the simpler approach of a generic regression of the outcome on the immigrant share.

[27] In other words, it is the ratio of the number of foreign-born persons in a skill group to the total number of persons in that skill group.

the least squares regression. The least squares coefficient of the immigrant share on the log weekly wage of black men, for example, is -0.346 (0.137), while the corresponding IV coefficient is -0.314 (0.131). In addition, the estimated IV coefficients in the black employment and black incarceration regressions are essentially the same as those obtained in the least squares specification.

Another concern is that the measured effects of immigration may be contaminated by factors that are driving employment and incarceration behavior for blacks and whites within education-experience groups. These factors will not be absorbed by the fixed effects included in the regression and may be correlated with the immigrant supply shifts. The potential existence of these additional factors is not altogether surprising. As we discussed above, the wage structure changed considerably during our sample period and the crack epidemic raised the return to crime during the 1980s and 1990s. It is important, therefore, to examine the impact of various sources of bias on the magnitude of the estimated coefficients.

Consider initially how the crack epidemic may influence our results. In terms of the model from the previous section, the invention of crack raises the marginal product of criminal labor, shifting the curve labeled D_c in Figure 6 to the left. This leftward shift reduces native labor supplied to the formal sector and raises native labor supplied to crime, just like an increase in immigration. Unlike an increase in immigration, however, the increasing productivity of crime should *raise* equilibrium wages. This casts some doubt on the notion that the effects we attribute to immigration are entirely due to crack, since the data reveal that immigration reduced wages. Nevertheless, given the potential importance of crack as an alternative explanation, it is important to account for it explicitly in the regression model.

To do so we make use of the “crack index” developed by Fryer, Heaton, Levitt, and Murphy (2005). This index is a linear combination of several variables related to the crack epidemic, including the share of arrests made for cocaine-related charges, the number of deaths due to cocaine, the number of cocaine busts carried out by the federal Drug Enforcement Administration, and the number of cocaine-related hospital emergency room incidents. Most of these measures can obviously be considered as outcomes of the crack epidemic, whereas the ideal measures for our purposes would be indicators of the extent to which crack raised criminal productivity and exogenous measures of the criminal-justice response to the crack problem. Thus one could argue that the crack index is really an endogenous variable, particularly in the incarceration regressions. This possibility will affect the interpretation of our results.

The Fryer et al crack index varies only by time and race, meaning that its main effects are subsumed by the year fixed effects included in our regression model.[28] To include the index in the model, we interact it with the education-experience fixed effects. In particular, consider the following regression:

$$(11) \qquad y_{ext} = \theta\, p_{ext} + E + X + T + (E \times T) + (X \times T) + (E \times X) + (E \times X)\, C_t + \varphi_{ext},$$

where C_t is the value of the crack index at time t. Note that the specification in (11) essentially introduces a specific type of time variation in the education-experience interaction fixed effects. Obviously, a totally unrestricted type of variation would be impossible since there would then be as many education-experience-time interactions as there are observations. Our specification permits the effects of skill to vary over time in a manner that is related to the spread of crack, while still enabling us to estimate the effects of immigration from changes over time within the same skill groups.

The estimated adjustment coefficients θ (and corresponding marginal impacts in the grouped logit regressions) are reported in panel C of Table 2. It is evident that the estimated marginal effects on black wages, employment, and incarceration rates are not very sensitive to the inclusion of the crack index. A ten-percent immigrant-induced supply shift still lowers the wage of blacks by -3.0 percent, lowers the employment rate by 4.1 percentage points, and increases the incarceration rate by 0.9 percentage points.

There is an important sense, however, in which the Fryer et al crack index does not capture how the crack epidemic affected the sectoral choices of the various skill groups. The index, after all, takes on the same value for black high school dropouts as it does for black college graduates. It is evident that the crack epidemic had a much greater impact on the behavior of low-skill (and younger) persons. To incorporate this notion into our study, we re-estimated the regression models using a "restricted" version of the Fryer et al crack index. In particular, we reset the index to 0 for all persons who have at least a high school diploma or more than 20 years of work experience. The restricted crack index, therefore, effectively assumes that the crack epidemic was mainly a demand shifter for younger, low-skill persons. Panel D of Table 2 reports the regression coefficients obtained from the use of the restricted crack index. Note that this change in the way we account for the crack epidemic barely affects our results.

The use of the crack index helps the regression model partly account for what has been happening at the low-skill end of the U.S. labor market. A large research literature documents that there was also a substantial increase in the rate of return to skills and to labor market experience beginning around 1980 (Katz and Murphy, 1992; Autor, Katz, and Kearny, 2008).

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[28] The index takes on a value of 0 for all cells drawn from the 1960, 1970, or 1980 Censuses. The index for the 1990 cells is the average for 1988 and 1989 of the Fryer et al (2005) race-specific city-level crack index. The index for the 2000 cells is the average of the 1998 and 1999 values.

Our basic regression specification in equation (9) attempts to control for these changes by including interactions of fixed effects between education and time, as well as interactions of fixed effects between experience and time. As noted above, ideally we would control for changes in the wage structure by including a complete set of three-way interactions between the education, experience, and time fixed effects. Such a strategy is not feasible, however, because the full set of interactions would be exactly collinear with the immigrant share variable. As a result, we must balance the collinearity problem against the need to control for changes in the wage structure by allowing year effects to vary by education and experience in a limited way.

We attempt to capture the changes that occurred at both ends of the skill distribution by introducing a set of dummy variables that allow for some interactions between education, experience, and time. In particular, we introduce dummy variables indicating if the (e, x, t) cell refers to a post-1980 observation of high school dropouts, a post-1980 observation of high school graduates or those with some college, or a post-1980 observation of college graduates. Moreover, for each education group we categorize individuals into very young, young, and older experience groups (1-10, 11-20, or 20 plus years of experience). This set of dummy variables is specifically designed to capture changes that occurred at the extremes of the wage distribution (in terms of both education and experience) before and after 1980.

It is worth emphasizing that the inclusion of this restricted set of education-experience-time interactions does not impose any structure on *why* these changes may have occurred. As a result, the interactions capture not only the impact of the widening of the wage structure, but also the impact of the crack epidemic, as well as any kinds of changes in the composition of particular skill groups over time (e.g., the possibility that younger low-skill workers may be selected differently before and after 1980). In particular, one consequence of rising returns to skill may have been to raise the incentive for workers at all ability levels to obtain more education. If, within each skill group, more able workers are those more likely to complete additional schooling, then over time we would observe a decline in the average ability of workers in *all* education groups. Such an outcome would be more likely among the young, since their net benefit to completing more schooling is relatively high. Declining average ability within skill groups could result in all groups, and especially the young, exhibiting declining wages, declining employment rates, and rising incarceration rates, which are the same patterns that would result from an immigration-induced labor supply shock. The included set of additional education-experience-time fixed effects allow for differential effects before and after 1980 among very young, young, and older workers, with low, average, or high levels of education. Hence we can partly control for changes in the average ability of individuals within experience-education group, focusing on those cases where the ability changes are likely to have been most pronounced.

The bottom panel of Table 2 reports the estimated coefficients from the specification with the restricted set of education-experience-time fixed effects. The wage effect of immigration remains negative and significant—even after controlling for these changes in the wage structure at the extremes of the skill distribution. In fact, the adjustment coefficients are very similar for the two race groups: -0.355 (0.171) in the black regression and -0.428 (0.233) in the white regression. A 10-percent immigrant-induced increase in supply, therefore, reduces wages by around 3.0 percent. Second, immigration reduces the employment rate of both black and white workers. These effects are statistically significant, and the marginal impact on blacks is over twice as large as the marginal impact on whites. Finally, immigration has a significant impact on incarceration rates for both blacks and whites, with the marginal impact on blacks being 3.5 times as large as the marginal impact on whites. A 10-percent immigrant induced increase in supply raises black incarceration rates by 1.3 percentage points and that of whites by 0.2 percentage points.[29]

We conclude that our evidence is robust to a number of major specification changes. One remaining question is whether our results would be substantially different if we cast the analysis in terms of state-level labor markets, rather than a national-level labor market. In an earlier draft of this paper (Borjas, Grogger, and Hanson, 2006), we addressed this issue by allowing for natives to move across states in response to immigration from abroad. We show that the correct regression specification in such a model would be one that relates native labor market outcomes to *both* state- and national-level immigration shares on the right hand side of our regression models (where the national-level immigration share captures the "externality" that immigration into one specific region imparts on the wage structure in other regions through native internal migration). The total effect of immigration is then given by the sum of the national- and state-level coefficients. Since the results of our state-level analysis were very similar to those reported in this section, we refer the interested reader to our working paper draft for more detail.

V. Accounting for the Trends in Employment and Incarceration Rates

We now use the regression coefficients reported in Table 2 to determine the extent to which the immigration-induced shift in supply accounts for the decline in black wages and employment and the increase in black incarceration. In particular, we use the coefficients to simulate the impact of the immigrant influx that entered the United States between 1980 and 2000. Suppose the estimated regression coefficient for a particular outcome y (and for a

[29] We experimented with other specifications for the restricted set of education-experience-time interactions (e.g., allowing a dummy variable for each experience cohort for high school dropouts after 1980). This experimentation revealed that the evidence summarized in Table 2 adequately represent the results that can be

particular race group) is $\hat{\theta}$. Equation (10) then implies that the reduced-form impact of an immigrant influx that increases the supply of education group *e* by m_e percent can be approximated by:

$$\text{(12)} \qquad \Delta y_e = \hat{\theta}\,(1-\overline{p})^2\, m_e,$$

where we evaluate the derivative in (10) at the mean value of the immigrant share observed in 2000. We define the immigration-induced labor-supply shock between 1980 and 2000 as:

$$\text{(13)} \qquad m_e = \frac{M_{e,2000} - M_{e,1980}}{0.5(N_{e,1980} + N_{e,2000}) + M_{e,1980}},$$

so that the baseline population used to calculate the percent increase in labor supply averages out the (changing) size of the native workforce during the 1980-2000 period and treats the pre-existing immigrant population as part of the "native" stock.

Table 3 summarizes the results of the simulation. Column 1 of the table reports the actual change in the log weekly wage, employment, and incarceration rates experienced by black men during the 1980-2000 period. The changes in wages, employment and incarceration rates are quite large for black men, and particularly for low-skill black men. Between 1980 and 2000, for example, the real wage of black high school dropouts (a group that made up around 25 percent of the black male population in 1990) fell by 14.2 percent, the employment rate fell by 16.7 percentage points, and the incarceration rate rose by 15.6 percentage points.

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obtained from a large family of alternative specifications—as long as the specifications include interactions that differentiate younger versus older workers separately for the two extreme education groups in the study.

Column 2 summarizes the results of the simulation when we use the adjustment coefficients reported in Panel E of Table 2. Equation (12) predicts that, *other things equal*, the 1980-2000 immigrant influx reduced the wage of black high school dropouts by 5.3 percent, reduced the employment rate by 12.6 percentage points, and increased the incarceration rate by 2.8 percentage points. These predicted impacts are sizable and account for about a third of the observed reduction in wages, for over two-thirds of the observed decline in employment, and for about 15 percent of the rise in incarceration rates. In other words, immigration contributed a numerically significant amount to the trends among low-skill blacks, but much of the decline in black economic status would have been observed even in the absence of the immigrant influx.

Table 3 shows that our approach cannot explain the observed changes in wages or employment for high-skill blacks. In the sample of black college graduates, for example, immigration is predicted to reduce wages by 2.9 percent and employment rates by 6.9 percent, whereas they rose by 13.1 percent and 0.1 percent, respectively. The key assumption of the simulation—that "other things are equal" in the 1980-2000 period—obviously misses an important part of what was driving employment opportunities for high-skill workers at that time. The simulation reported in Table 3 does not account for these shifts.

A second caveat is that the regression results reported above impose the restriction that the various demand elasticities are invariant to skill. We might imagine that the elasticity of labor demand in crime, in particular, varies by education level, with elasticities being larger for the less-educated. If the demand for highly educated labor in crime is in fact relatively inelastic, we would overstate the impact of immigration on their employment and incarceration rates.

Finally, the "all other things equal" assumption ignores the capital adjustments induced by the immigrant supply shock. If the aggregate production function had constant returns to scale, for example, the capital adjustments would eventually return the average wage in the labor market to its pre-immigration equilibrium level. Even in the long run, however, immigration can have distributional effects. In a CES framework, the long-run *relative* wage effects of immigration are obtained by differencing the effects across the different education groups reported in Table 3 (Borjas and Katz, 2007). For example, the 1980-2000 immigrant influx lowered the wage of black high school dropouts relative to that of college graduates by 2.4 percent (or –0.053 – (–0.029)). These relative wage effects suggest that there would also be sizable relative employment effects and relative incarceration effects.

The last two columns of Table 3 report the simulation results for white men. The 1980-2000 immigrant influx reduced the wages of white high school dropouts by about 6.8 percent, or about a third of the observed change. The influx also lowered the employment rate by 4.6 percentage points, or about 40 percent of the decline actually observed in the white sample. Immigration also explains a smaller portion of the increase in incarceration rates observed among low-skill white men.

VI. Summary

It is well known that black employment rates declined substantially over the past few decades, and that this decline was accompanied by a rapid rise in the black incarceration rate. This paper examines a simple yet potentially controversial question: has the resurgence of large-scale immigration in the United States contributed to these trends?

We use data drawn from the microdata files of the 1960-2000 U.S. Censuses to examine the trends in black wages, employment, and incarceration. Our empirical analysis examines the link between immigration and the evolution of these variables over the four-decade period—after adjusting for other factors that could account for shifts in the wage structure as well as shifts in opportunities in the "formal" labor market and in the crime sector. We find a numerically sizable and statistically significant negative correlation between immigration and the wages of black men; a sizable and significant negative correlation between immigration and the employment rate of black men; and a significant positive correlation between immigration and the incarceration rate of black men. It is important to emphasize that we find similar correlations for white men, but the magnitude of these correlations differ in important ways between the two groups. Although the wage effect of immigration is similar for black and white men, the negative employment effect and the positive incarceration effect are larger for blacks.

Our analysis suggests that a 10-percent immigration-induced increase in the supply of a particular skill group reduced the black wage by about 3 percent, lowered the employment rate of black men by about 5 percentage points, and increased the incarceration rate of blacks by one percentage point. Although these effects are statistically and numerically important, much of the decline in employment or increase in incarceration in the black population remains unexplained.

DATA APPENDIX

The data are drawn from the 1960, 1970, 1980, 1990, and 2000 Integrated Public Use Microdata Samples (IPUMS) of the U.S. Census. In the 1960 Census, the data extract forms a 1 percent sample of the population. In the 1970 Census, the data extract forms a 3 percent sample (formed by pooling the state, metropolitan area, and neighborhood samples). Finally, in the 1980, 1990, and 2000, the data extracts form a 5 percent sample. The analysis is restricted to men aged 18-64. A person is classified as an immigrant if he was born abroad and is either a non-citizen or a naturalized citizen; all other persons are classified as natives. We use the information contained in the census race variable to classify persons as "black" (IPUMS variable *raced* = 200) or "white" (*raced* = 100 or 110). Sampling weights are used in all calculations.

Definition of education and experience: We use the IPUMS variables *educrec* to classify workers into four education groups: high school dropouts (*educrec* <= 6), high school graduates (*educrec* = 7), persons with some college (*educrec* = 8), college graduates (*educrec* = 9). We assume that high school dropouts enter the labor market at age 17, high school graduates at age 19, persons with some college at age 21, and college graduates at age 23, and define work experience as the worker's age at the time of the survey minus the assumed age of entry into the labor market. We restrict the analysis to persons who have between 1 and 40 years of experience. Workers are classified into one of 8 experience groups, defined in five-year intervals.

Counts of persons in education-experience groups: The workforce counts are calculated in the sample of men who worked at some point in the calendar year prior to the census and are not enrolled in school. The workforce counts weigh each observation by the number of hours worked in the previous calendar year (so that the total count of immigrants and natives in a particular skill groups effectively represents the total number of hours worked by the group). The population counts are calculated in the sample of men who are not enrolled in school.

Weekly earnings: We use the sample of men who do not reside in group quarters, report positive weeks worked, are not enrolled in school, and report positive earnings. Our measure of earnings is the sum of the IPUMS variables *incwage* and *incbusfm* in 1960, the sum of *incearn*, *incbus*, and *incfarm* in 1970 and 1980, and is defined by *incearn* in 1990-2000. In the 1960, 1970, and 1980 Censuses, the top coded annual salary is multiplied by 1.5. All earnings are deflated to 1990 dollars. In the 1960 and 1970 Censuses, weeks worked in the calendar year prior to the survey are reported as a categorical variable. We imputed weeks worked for each worker as follows: 6.5 weeks for 13 weeks or less, 20 for 14-26 weeks, 33 for 27-39 weeks, 43.5 for 40-47 weeks, 48.5 for 48-49 weeks, and 51 for 50-52 weeks. The average log weekly earnings for a particular education-experience cell is defined as the mean of log weekly earnings over all workers in the relevant population.

Employment and incarceration rates: These variables are calculated in the total sample of men. The employment rate is the sample average of the ratio of weeks worked during the calendar year prior to the Census, including zeros, to 52. A person is institutionalized if the group quarter variable *gq* takes on a value of 3 (indicating the person is in an institution). The 1990 and 2000 Censuses do not provide detailed information on the type of institution where an institutionalized person resides.

MATHEMATICAL APPENDIX

1. Derivation of equations (6a)-(6d)

Equate equations (2) and (3) in the text. This yields:

$$\text{(A1)} \qquad N_{ics} = g\left(\frac{X_{fs}}{X_{ics}}, L_{fs}\right) = \left(\frac{X_{fs}}{X_{ics}}\right)^{1/\eta_{ic}} (L_{fs})^{\eta_f/\eta_{ic}}.$$

We linearize equation (A1) using a multivariate first-order Taylor series approximation. The linearization is conducted at the arbitrary point:

$$\text{(A2)} \qquad \bar{N}_f = \frac{1}{S}\sum_{s=1}^{S} N_{fs}^*, \quad \overline{X_f / X_{ic}} = \frac{1}{S}\sum_{s=1}^{S}(X_{fs}^* / X_{ics}^*),$$

where the asterisks indicate the values of the variables in the pre-immigration equilibrium and S gives the number of skill groups. Recall that L_f gives the sum of natives (N_f) and immigrants (M) in the market sector. In the pre-immigration equilibrium, it must be the case that $L_f = N_f$. The first-order approximation of (A1) can then be written as:

$$\text{(A3)} \qquad \begin{aligned} N_{ics} &= g(\overline{X_f / X_{ic}}, \bar{N}_f) + g_1(\overline{X_f / X_{ic}}, \bar{N}_f)(X_{fs} / X_{ics} - \overline{X_f / X_{ic}}) \\ &+ g_2(\overline{X_f / X_{ic}}, \bar{N}_f)(L_{fs} - \bar{N}_{fs}) \end{aligned}$$

where $g_1 = \partial g / \partial (X_{fs}/X_{ics})$ and $g_2 = \partial g / \partial L_{fs}$. Given (A1) these derivatives are:

$$\text{(A4)} \qquad \begin{aligned} g_1 &= \frac{N_{ics}}{\eta_{ic}}\left(\frac{X_{fs}}{X_{ics}}\right)^{-1}, \\ g_2 &= \frac{\eta_f}{\eta_{ic}}\frac{N_{ics}}{L_{fs}}. \end{aligned}$$

If we evaluate these derivatives at the points defined in (A2), the first-order approximation is:

$$\text{(A5)} \qquad N_{ics} = \alpha_{ics} + \beta_{ic} L_{fs},$$

where:

$$\text{(A6)} \qquad \begin{aligned} \alpha_{ics} &= g(\overline{X_f / X_{ic}}, \bar{N}_f) - g_1(\overline{X_f / X_{ic}}, \bar{N}_f,)\overline{X_f / X_{ic}} + g_1(\overline{X_f / X_{ic}}, \bar{N}_f)(X_{fs} / X_{ics}) \\ &\quad - g_2(\overline{X_f / X_{ic}}, \bar{N}_f)\bar{N}_f \\ \beta_{ic} &= g_2(\overline{X_f / X_{ic}}, \bar{N}_f). \end{aligned}$$

Note that the intercept in (A5) is a function of the skill level s, but the slope β_{ic} is not. By using a similar method, we can derive the analogous linear approximation for N_{ihs}:

(A7) $$N_{ihs} = \alpha_{ihs} + \beta_{ih} L_{fs} \,,$$

In the post-immigration equilibrium, the total number of persons participating in the market sector can be written as:

(A8) $$L_{fs} = (\tilde{N}_{bs} + \tilde{N}_{ws} + M_s) - N_{bcs} - N_{wcs} - N_{bhs} - N_{whs} \,,$$

where $\tilde{N}_{is}$ gives the number of native persons of race i and skill s in the population. By substituting equations (A5) and (A7) into (A8), and solving for L_{fs} we obtain:

(A9) $$L_{fs} = \alpha_s + \rho(\tilde{N}_{bs} + \tilde{N}_{ws} + M_s),$$

where ρ is a constant that is not race-specific and that does not depend on skills. In fact, it follows from the definition of g_2 that:

(A10) $$0 < \rho = \left[\frac{\bar{N}_f}{\bar{N}_f + \frac{\eta_f}{\eta_{bc}} \bar{N}_{bc} + \frac{\eta_f}{\eta_{wc}} \bar{N}_{wc} + \frac{\eta_f}{\eta_{bh}} \bar{N}_{bh} + \frac{\eta_f}{\eta_{wh}} \bar{N}_{wh}} \right] < 1 \,.$$

where $\bar{N}_{ic} = g(\overline{X_f / X_{ic}}, \bar{N}_f)$ and $\bar{N}_{ih}$ is defined analogously. The intercept α_s in equation (A8) does depend on skills and is defined by:

(A11) $$\alpha_s = -(\alpha_{bcs} + \alpha_{wcs} + \alpha_{bhs} + \alpha_{whs})\rho \,.$$

It follows from equation (A9) that $N^*_{fs} = \alpha_s + \rho(\tilde{N}_{bs} + \tilde{N}_{ws})$ in the pre-immigration equilibrium. Substituting this into equation (2) in the text implies that the pre-immigration equilibrium wage is:

(A12) $$\ln w^*_{is} = \ln[X_{fs}(1-\delta_i)] + \eta_f \ln[\alpha_s + \rho(\tilde{N}_{bs} + \tilde{N}_{ws})] \,.$$

Suppose immigrant supply increases from 0 to M_s. The post-immigration wage is:

(A13) $$\begin{aligned} \ln w_{is} &= \ln[X_{fs}(1-\delta_i)] + \eta_f \ln[\alpha_s + \rho(\tilde{N}_{bs} + \tilde{N}_{ws} + M_s)], \\ &= \ln[X_{fs}(1-\delta_i)] + \eta_f \ln[N^*_{fs} + \rho M_s], \\ &= \ln w^*_{is} + \eta_f \rho m_s, \end{aligned}$$

where $m_s = M_s / N^*_{fs}$, which produces the result in equation (6a). Equations (6b) and (6c) in the text follow from combining (A5) with equations (2) and (3). To derive equation (6d), write equation (5) in terms of rates of change as

(A14) $$\ln N_{ifs} - \ln N^*_{ifs} = -\frac{\theta_{ics}}{\theta_{ifs}}(\ln N_{ics} - \ln N^*_{ics}) - \frac{\theta_{ihs}}{\theta_{ifs}}(\ln N_{ihs} - \ln N^*_{ihs}),$$

where $\theta_{ics} = N^*_{ics} / \tilde{N}_{is}$, $\theta_{ihs} = N^*_{ihs} / \tilde{N}_{is}$, and $\theta_{ifs} = N^*_{ifs} / \tilde{N}_{is}$. Equation (6d) follows after substituting equations (6b) and (6c) into (A14).

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Figure 1. The share of immigrants in the workforce

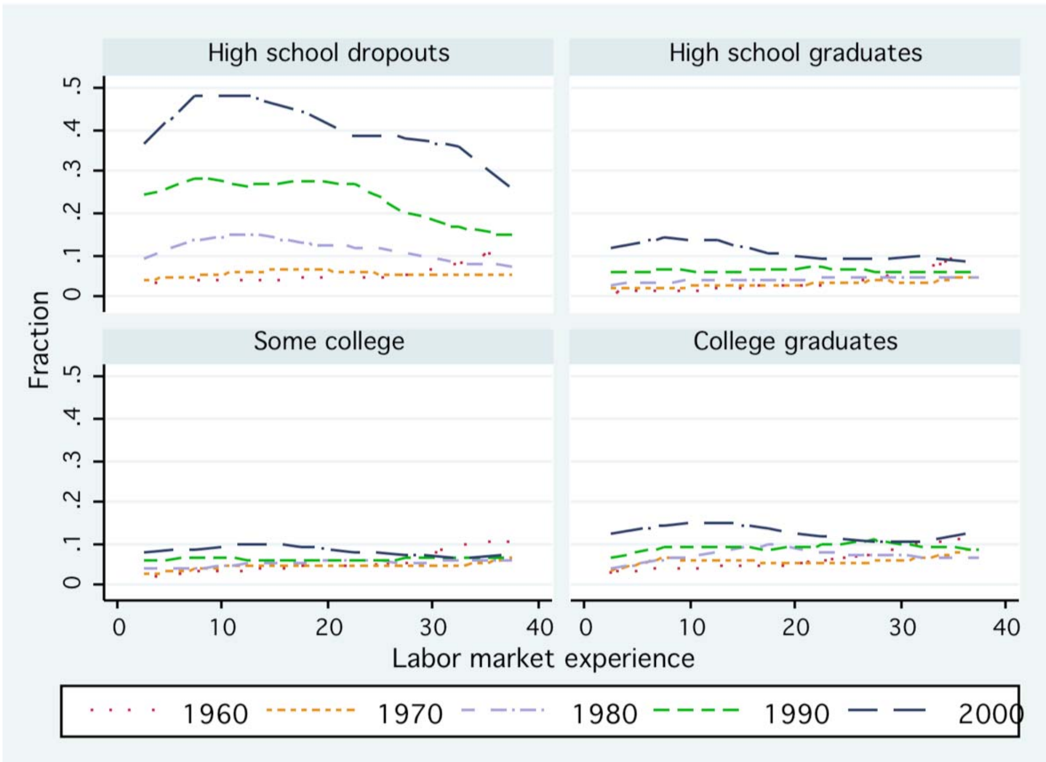
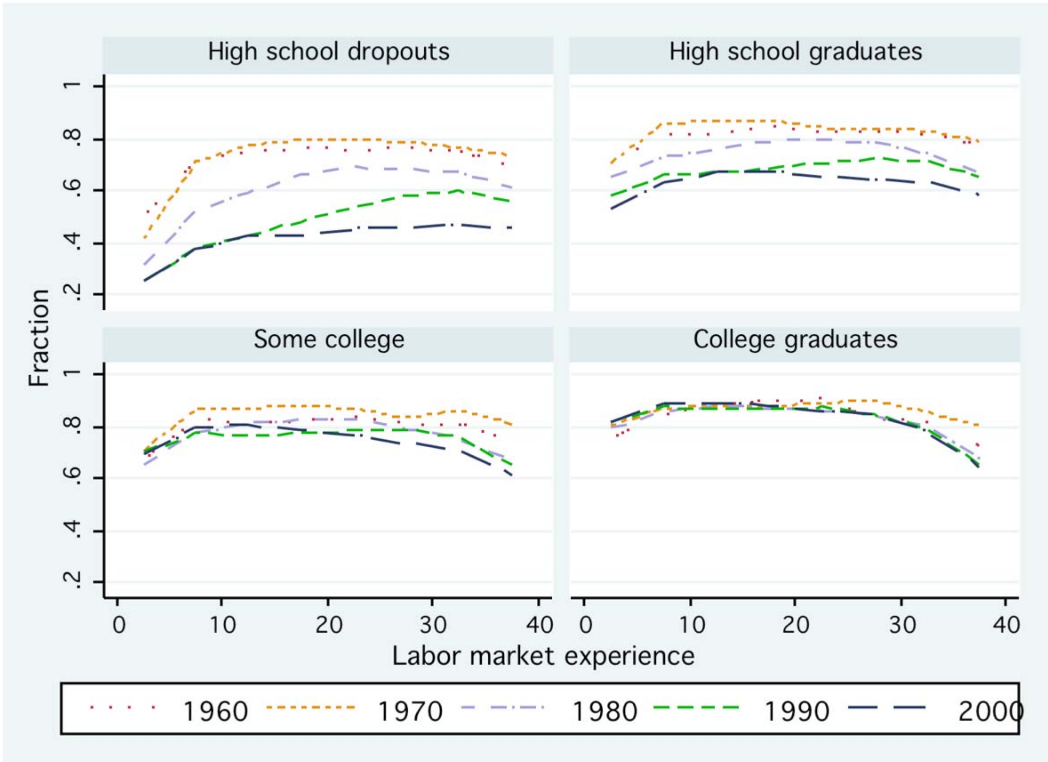


Figure 2. Trends in employment rates, by race, education and experience

A. Blacks



B. Whites

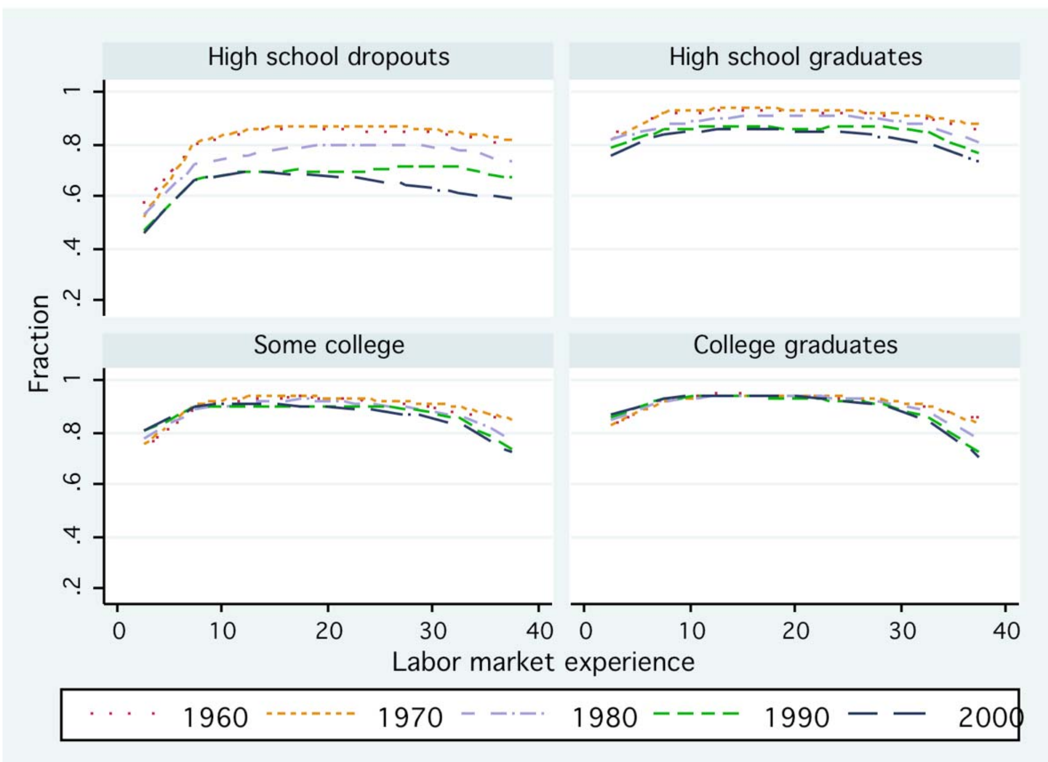
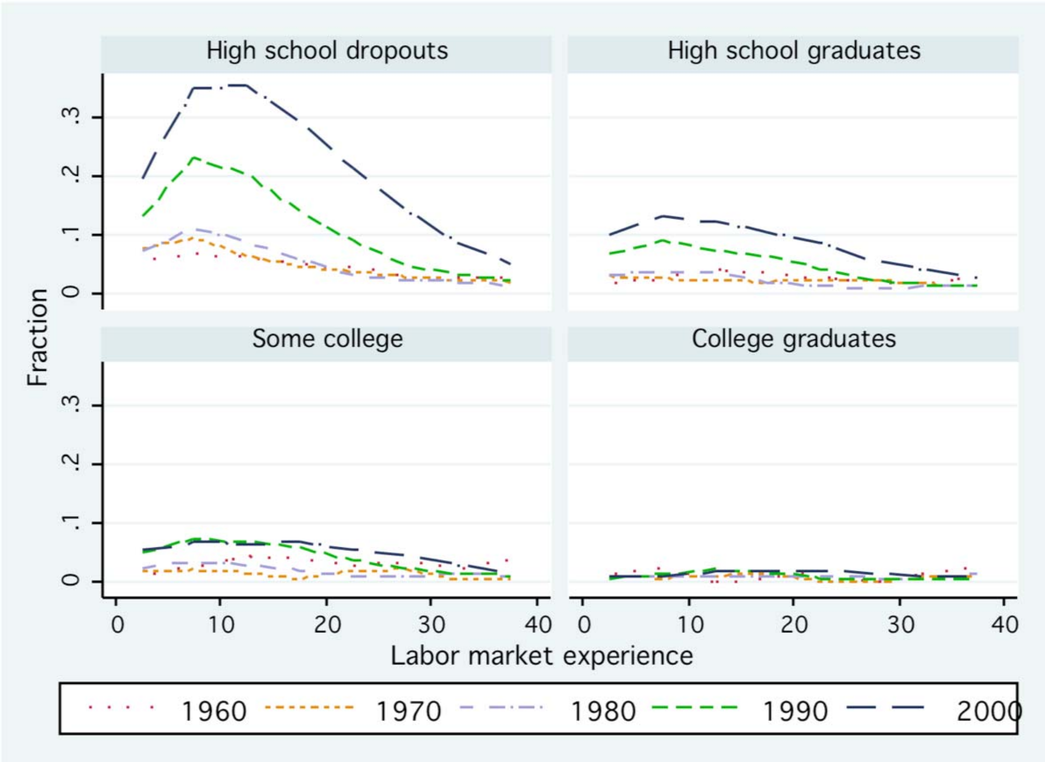


Figure 3. Trends in incarceration rates, by race, education and experience

A. Blacks



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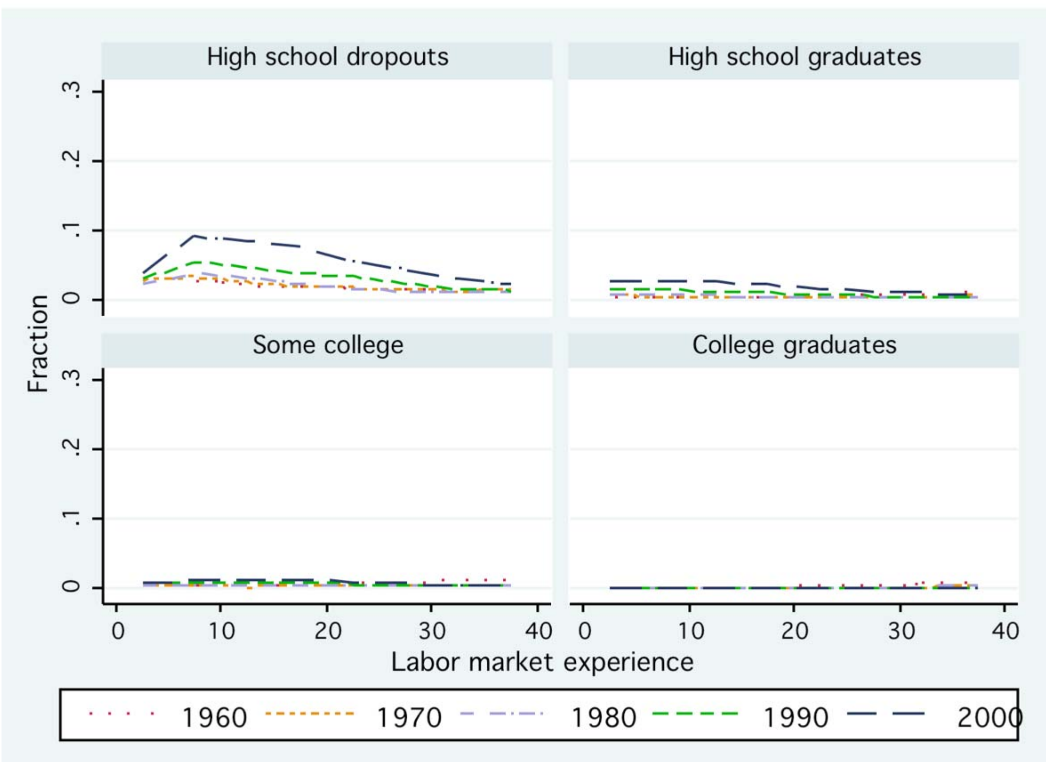
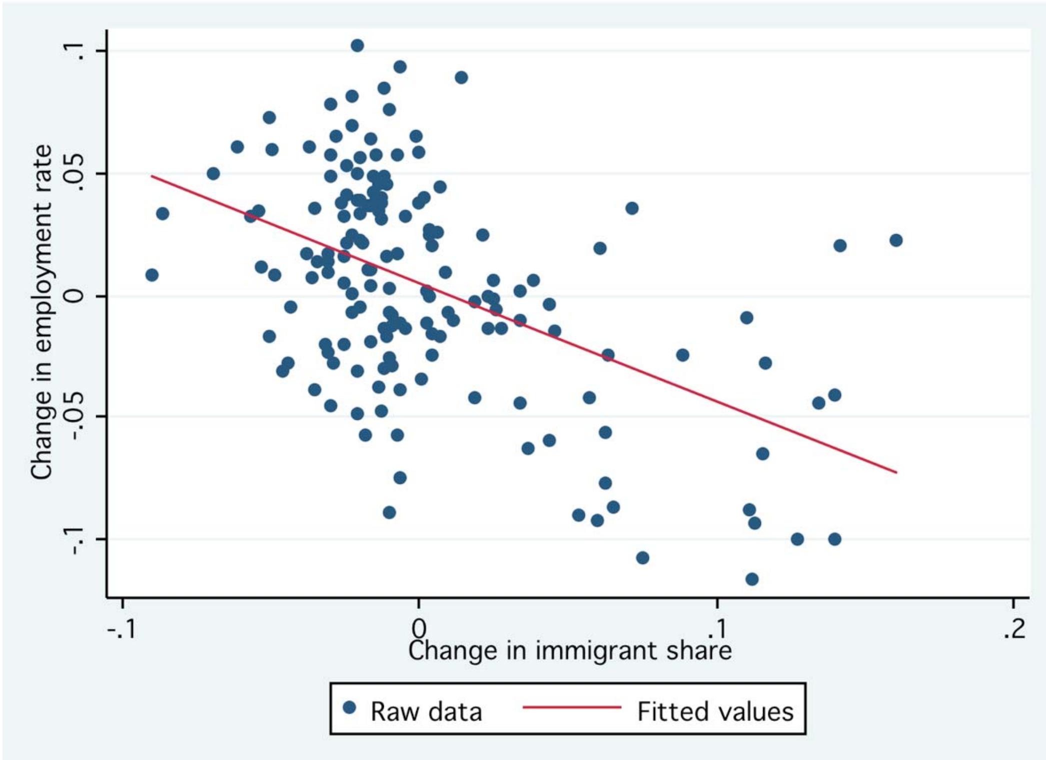


Figure 4. Relation between decadal changes in employment and immigration (removing decade effects)

A. Blacks



B. Whites

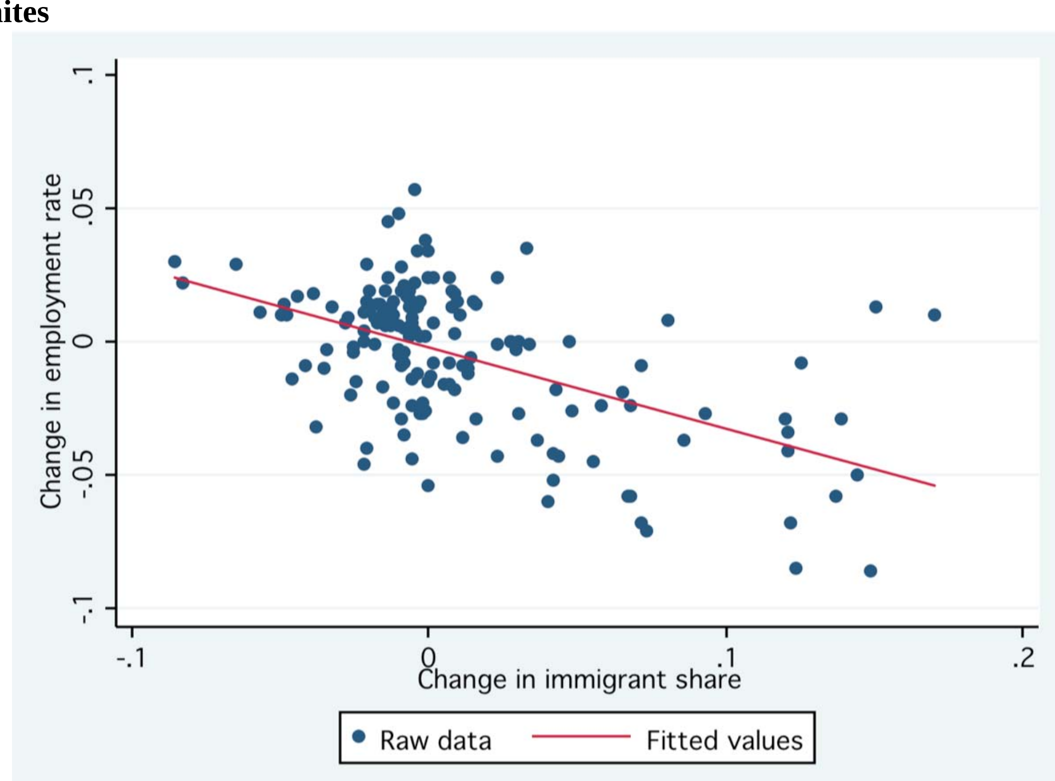
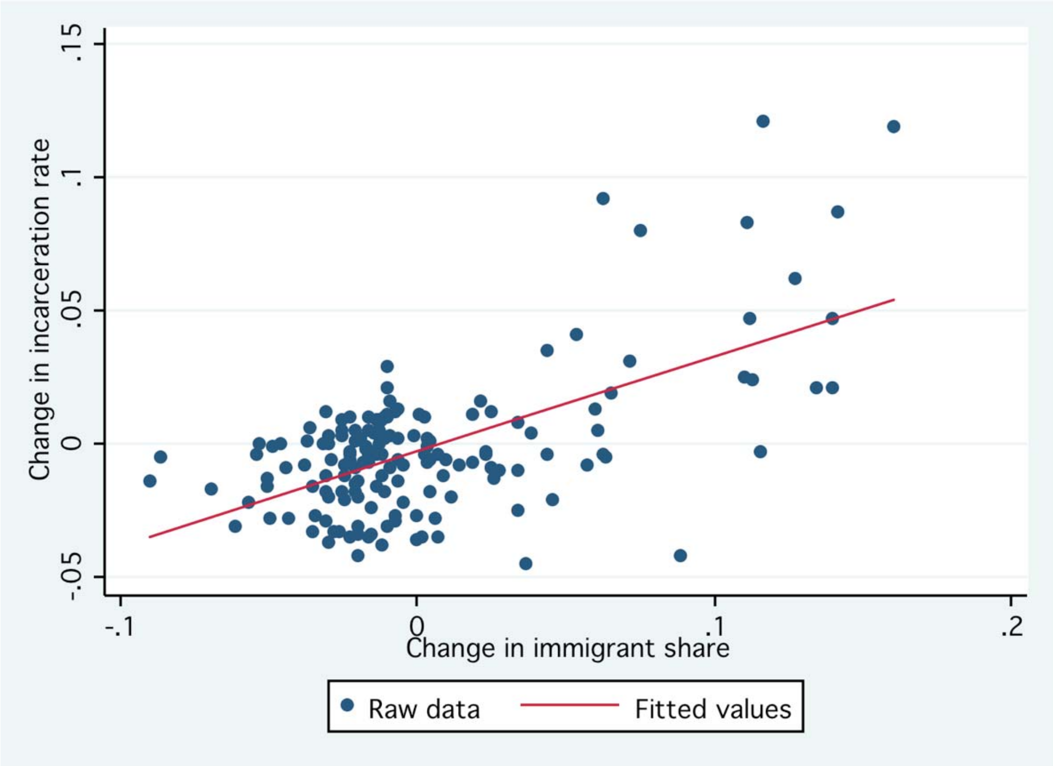


Figure 5. Relation between decadal changes in incarceration and immigration (removing decade effects)

A. Blacks



B. Whites

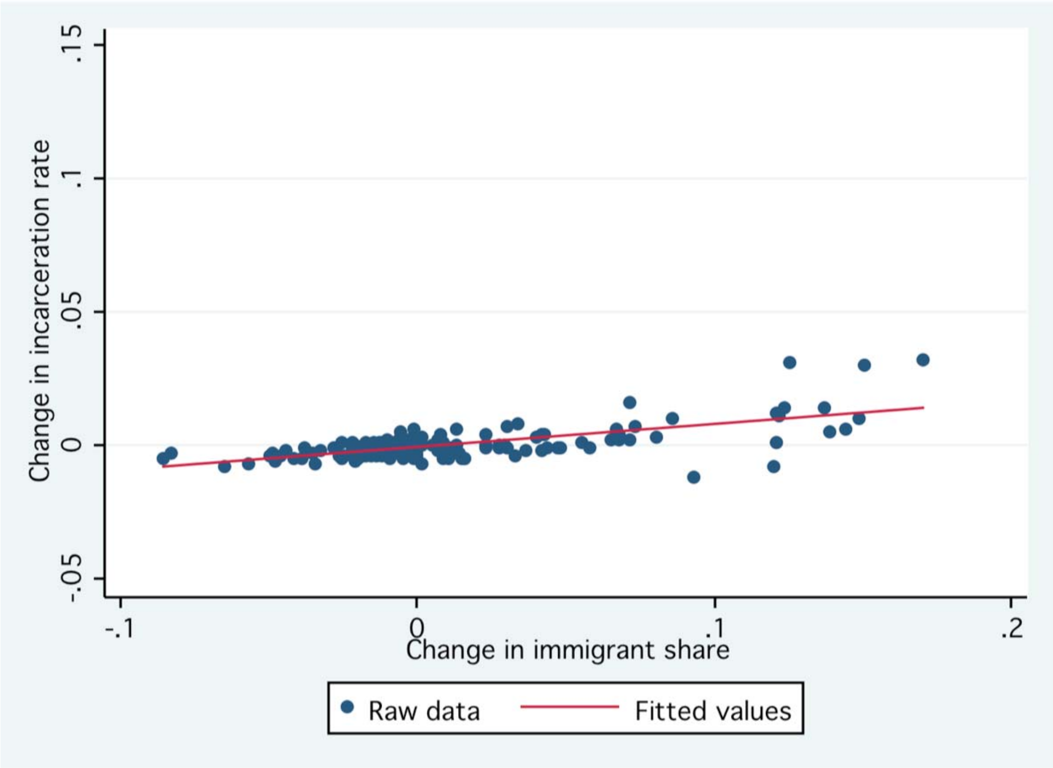
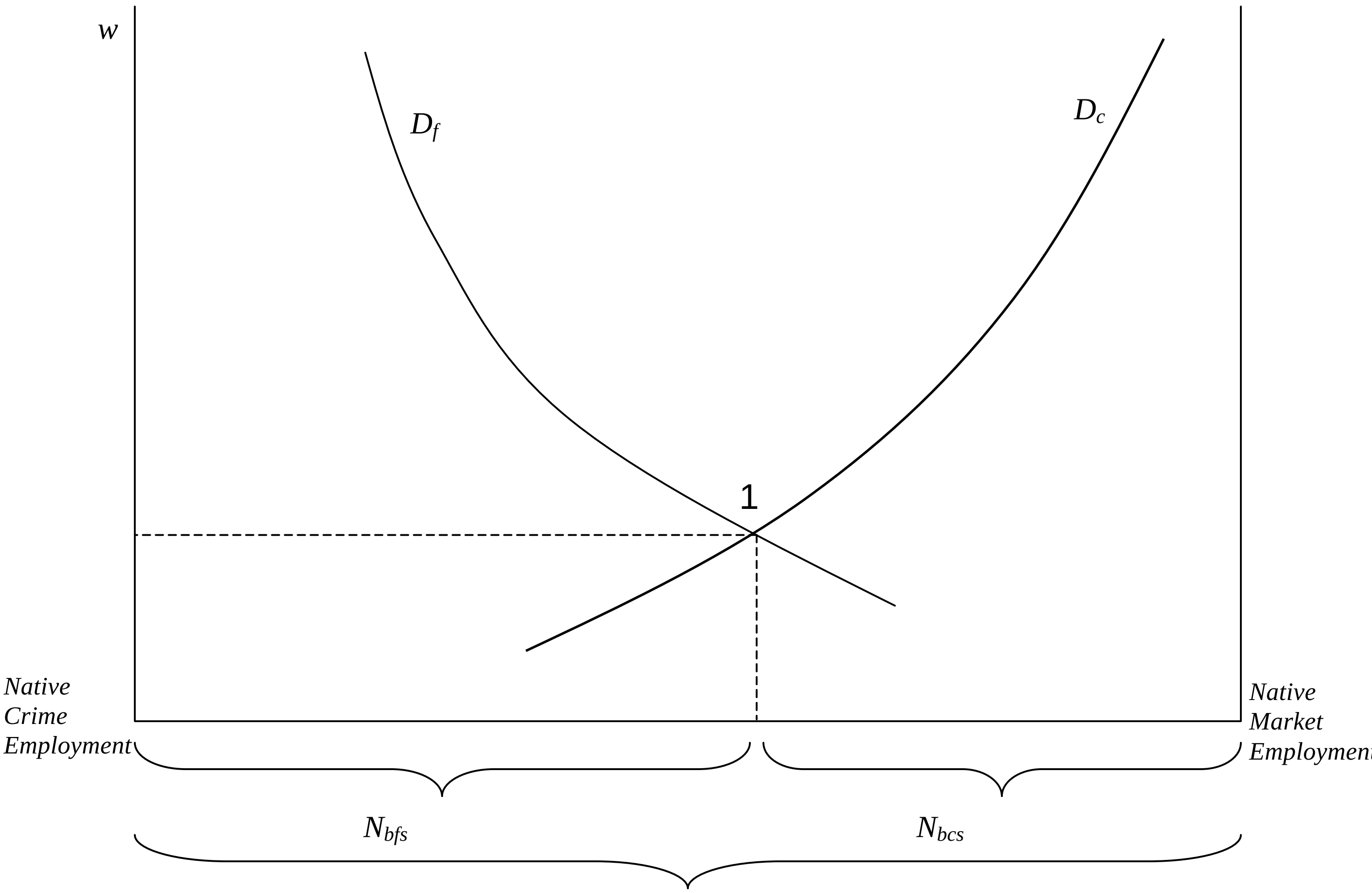


Figure 6. Initial equilibrium: allocation of labor across sectors



$$\tilde{N}_{bs} - N_{bhs} = N_{bcs} + N_{bfs}$$

Figure 7. Impact of immigration on sectoral allocation of labor

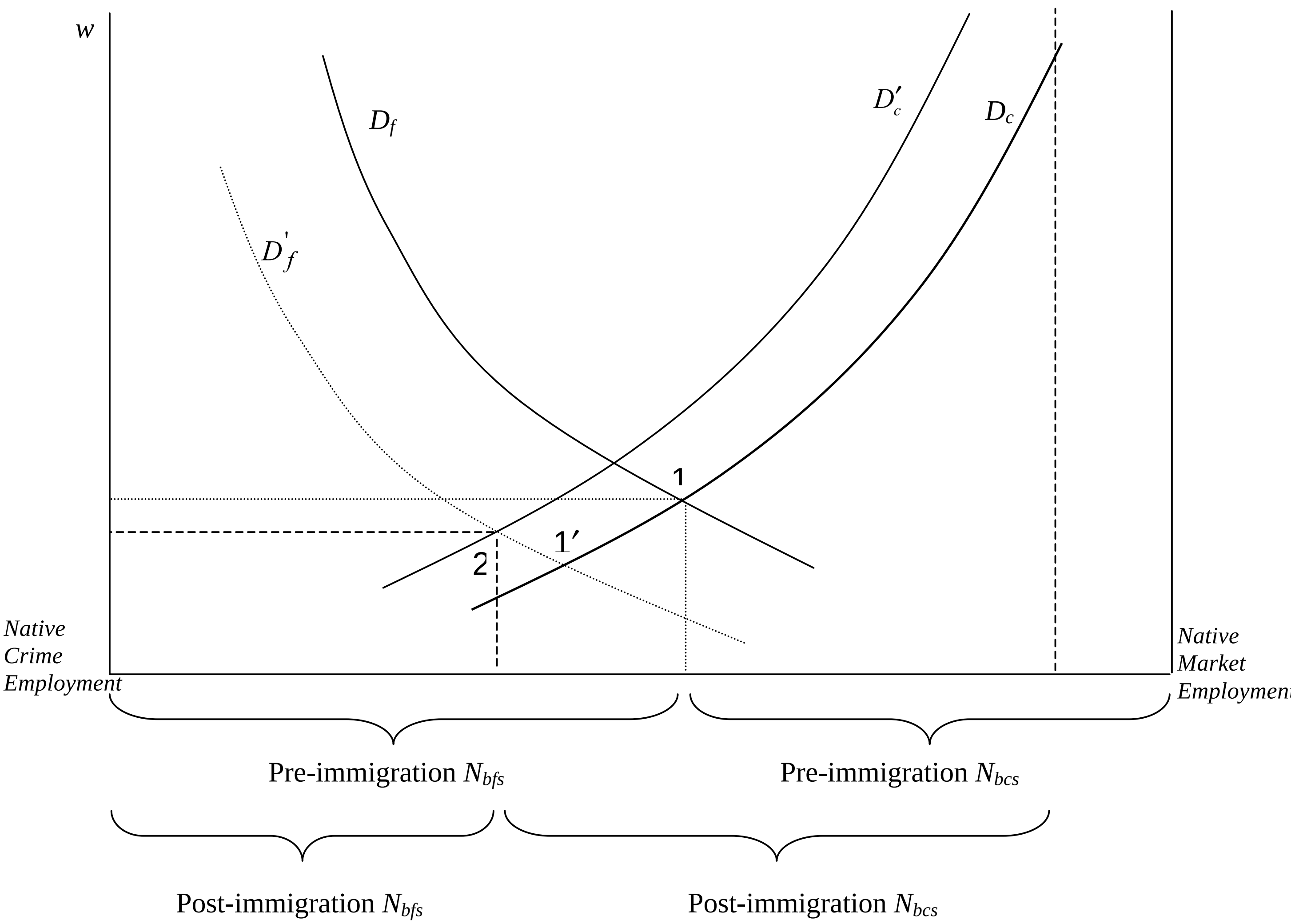


Table 1. Tests for perfect substitution

	Specification			
	1	2	3	4
A. Testing perfect substitution between black and white native workers.				
OLS estimate of $-1/\sigma$	0.013	-0.003	0.033	-0.045
	(0.012)	(0.008)	(0.026)	(0.027)
IV estimate of $-1/\sigma$	0.011	-0.005	0.019	-0.054
	(0.012)	(0.009)	(0.034)	(0.032)
Includes time fixed effects	No	Yes	Yes	Yes
Includes education-experience fixed effects	No	No	Yes	Yes
Interacts education and time fixed effects	No	No	No	Yes
Interacts experience and time fixed effects	No	No	No	Yes
A. Testing perfect substitution between immigrant and native workers				
OLS estimate of $-1/\sigma$	-0.016	0.019	-0.002	0.047
	(0.008)	(0.008)	(0.014)	(0.042)
IV estimate of $-1/\sigma$	-0.019	0.018	-0.002	0.035
	(0.008)	(0.008)	(0.014)	(0.043)
Includes time fixed effects	No	Yes	Yes	Yes
Includes education-experience fixed effects	No	No	Yes	Yes
Interacts education and time fixed effects	No	No	No	Yes
Interacts experience and time fixed effects	No	No	No	Yes

Notes: Standard errors are reported in parentheses and are adjusted for clustering within education-experience cells. All regressions have 160 observations and are weighted by the total number of observations used to calculate the dependent variable. The dependent variable in panel A is the difference between the mean log weekly wage of black and white workers, and the independent variable is the difference between the log of the number of black workers and the log of the number of white workers. The dependent variable in panel B is the difference between the mean log weekly wage of immigrant and native workers, and the independent variable is the difference between the log of the number of immigrant workers and the log of the number of native workers.

Table 2. Estimates of the impact of immigration

	Dependent variable			
	Log weekly earnings	Employment rate	Incarceration rate	Incarceration rate, using lagged immigration
A. Least squares				
Blacks	-0.346	-0.683	0.135	0.086
	(0.137)	(0.183)	(0.078)	(0.012)
Whites	-0.522	-0.222	0.003	0.026
	(0.254)	(0.097)	(0.025)	(0.009)
B. Instrumental variables				
Blacks	-0.314	-0.683	0.135	0.086
	(0.131)	(0.183)	(0.079)	(0.012)
Whites	-0.444	-0.222	0.003	0.026
	(0.243)	(0.097)	(0.025)	(0.009)
C. IV, with crack index				
Blacks	-0.410	-0.557	0.116	0.070
	(0.258)	(0.248)	(0.067)	(0.029)
Whites	-0.332	-0.231	0.039	0.028
	(0.135)	(0.137)	(0.035)	(0.008)
D. IV, with restricted crack index				
Blacks	-0.432	-0.655	0.113	0.066
	(0.166)	(0.176)	(0.052)	(0.025)
Whites	-0.434	-0.272	0.030	0.025
	(0.280)	(0.105)	(0.028)	(0.008)
E. IV, with restricted education-experience-time interactions				
Blacks	-0.335	-0.792	0.176	0.072
	(0.171)	(0.193)	(0.047)	(0.015)
Whites	-0.428	-0.287	0.033	0.020
	(0.233)	(0.074)	(0.015)	(0.005)

Notes: Standard errors are reported in parentheses and are adjusted for clustering within education-experience cells. All regressions have 160 observations and include education, experience, and period fixed effects, and interactions between education and experience fixed effects, education and period fixed effects, and experience and period fixed effects. Regressions on employment and incarceration rates use a grouped logit specification; reported coefficients are marginal effects evaluated at the mean employment and incarceration rates in the particular sample. Instrumental variable regressions instrument the immigrant share in the workforce with the immigrant share in the population. The restricted crack index sets the value of the Fryer et al (2005) crack index to zero if the skill cell has at least a high school education or more than 20 years of experience. The restricted education-experience-time interactions include a vector of fixed effects indicating if the cell refers to a post-1980 observation of high school dropouts with 1-10, 11-20, or more than 20 years of experience; a post-1980 observation of workers with a high school diploma or some college with 1-10, 11-20, or more than 20 years of experience; or a post-1980 observation of college graduates with 1-10, 11-20, or more than 20 years of experience. The lagged immigration variable gives the immigrant share for the particular skill group measured five years prior to the Census.

Table 3. The impact of the 1980-2000 immigrant influx

	Black Men		White Men	
	1. Actual change (1980-2000)	2. Predicted impact	1. Actual change (1980-2000)	2. Predicted impact
Log weekly wage				
High school dropouts	-0.142	-0.053	-0.205	-0.068
High school graduates	-0.076	-0.020	-0.133	-0.025
Some college	0.018	-0.022	-0.034	-0.028
College graduates	0.131	-0.029	0.114	-0.037
Employment rate				
High school dropouts	-0.167	-0.126	-0.110	-0.046
High school graduates	-0.098	-0.046	-0.048	-0.017
Some college	0.001	-0.051	0.000	-0.018
College graduates	0.010	-0.069	-0.003	-0.025
Incarceration rate				
High school dropouts	0.156	0.028	0.032	0.005
High school graduates	0.065	0.010	0.014	0.002
Some college	0.031	0.011	0.005	0.002
College graduates	0.005	0.015	0.001	0.003

Notes: The education-specific immigrant supply shock is defined as the number of immigrants arriving between 1980 and 2000 divided by a baseline population equal to the average size of the native workforce (over 1980-2000) plus the number of immigrants in 1980. The education-specific supply shocks are: 0.214 for high school dropouts; 0.079 for high school graduates; 0.087 for persons with some college; and 0.118 for college graduates. The predicted impacts use the coefficients reported in Panel E of Table 2.