DISCUSSION PAPER SERIES

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ABSTRACT

Immigration, Working Conditions, and Compensating Differentials

The large inflow of less-educated immigrants that the United States has received in recent decades can worsen or improve U.S. natives' labor market opportunities. Although there is a general consensus that low-skilled immigrants tend to hold "worse" jobs than U.S. natives, the impact of immigration on U.S. natives' working conditions has received little attention. This study examines how immigration affected U.S. natives' occupational exposure to workplace hazards and the return to such exposure over 1990 to 2018. The results indicate that immigration causes less-educated U.S. natives' exposure to workplace hazards to fall, and instrumental variables results show a larger impact among women than among men. The compensating differential paid for hazard exposure appears to fall as well, but not after accounting for immigration-induced changes in the returns to occupational skills.

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The United States experienced sustained high levels of immigration in recent decades, with the foreign-born population share almost tripling between 1970 and 2018. This tremendous influx of immigrants created both challenges and opportunities for U.S.-born workers. Research reaches disparate conclusions on whether immigration caused earnings and employment to fall among U.S. natives (e.g., Borjas, Grogger, and Hanson, 2012; Ottaviano and Peri, 2012), but several studies agree that immigration had a perhaps-unexpected upside: immigrant inflows pushed U.S. natives into better jobs. Peri and Sparber (2009, 2011a) show that, as a result of increased immigration, U.S. natives moved into higher-paying jobs that utilize communication skills more intensively, and Dillender and McInerney (2020) show that Mexican immigration caused U.S. natives to move into safer jobs. Bond, Giuntella, and Lonsky (2020) find that immigration resulted in U.S. natives moving from shift work into daytime jobs.

This study adds to the literature an examination of how immigration affected less-educated U.S.-born workers' exposure to workplace hazards and the wage premium – if any – workers receive for such exposure. Previous research shows that immigrants work in occupations with more exposure to hazardous conditions and higher injury and fatality rates than similarly educated U.S. natives (Orrenius and Zavodny, 2009; Zavodny, 2015), and some evidence suggests that immigrants earn smaller compensating differentials than U.S. natives for occupational risks (Hersch and Viscusi, 2010; Hall and Greenman, 2015).[1] Only one study, Dillender and McInerney (2020), examines whether immigration affects U.S. natives' exposure to occupational hazards, but it is limited to men only. No previous study has examined whether immigration affects the compensating differential that workers receive for exposure to

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[1] Dávila, Mora, and González (2011) find that Hispanic immigrant men with limited English proficiency earn larger compensating differentials than other men for occupational risk. However, they note that those differentials could be due to differences in risk within occupations, which cannot be observed in the data.

occupational hazards.[2] We present a comprehensive examination of how immigration affected

exposure to occupational hazards among less-educated U.S.-born men and women and the

compensating differential that workers received for exposure to occupational hazards during

1990 to 2018.

Understanding how immigration affects natives' exposure to workplace hazards and the

compensating differential for such exposure is important for several reasons. The workplace is a

major source of exposure to hazardous conditions for working-age adults. In 2016, about 1.5

million U.S. workers had a work-related injury or illness, and an average of 14 workers died per

day as a result of a workplace injury.[3] Less-educated and minority workers tend to be more

exposed to workplace hazards. The measures of workplace hazards we examine include exposure

to disease or infection, which has become of paramount importance since the emergence of

COVID-19 and outbreaks at meatpacking plants and poultry processors, among other

workplaces.

U.S. natives benefit if immigration results in natives moving into less-hazardous jobs.

Further, if immigrants are healthier than U.S. natives, workplace injury and illness rates may fall

if immigrants are concentrated in riskier occupations. Dillender and McInerney (2020) document

that such changes led to a drop in workers' compensation claims, which reduced costs for

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[2] Bond, Giuntella, and Lonsky (2020) find that immigration reduced U.S. natives' compensating differential for shift work. Several studies examine how immigration affected European natives' exposure to occupational risks. Alacevich and Nicodemo (2019) find that immigration led to reductions in natives' injury rates in Italy and report suggestive, although not statistically significant, evidence that changes in the physical intensity of natives' occupations underlies the observed effect. Bellés-Obrero, Bassols, and Castello (2020) find that increased immigration caused the number of workplace accidents among natives to fall in Spain as natives shifted to communications-intensive occupations from manual occupations. Giuntella and Mazzonna (2015) find that immigration improved natives' health outcomes in Germany as a result of improvements in natives' working conditions. Giuntella et al. (2019) find immigration caused UK natives to shift into jobs with lower physical burdens and injury risks.

[3] See https://www.bls.gov/spotlight/2019/25-years-of-worker-injury-illness-and-fatality-case-data/home.htm (accessed July 3, 2020).

employers and insurers. However, movement into less-hazardous jobs may involve a tradeoff: workers may earn less if they forego a risk premium.

Movement of U.S. natives into jobs that do not pay a risk premium may underlie wage declines among less-educated natives that some researchers attribute to immigration (e.g., Borjas, 2003). If so, U.S. natives may be better off than their wages would otherwise suggest. However, U.S.-born workers who remain in relatively hazardous jobs may be worse off if immigration causes compensating differentials to fall. Our study is the first to examine whether this occurs.

Combining O*NET data on occupational conditions with U.S. Census Bureau data on occupations held by U.S. natives who have at most finished high school and the immigrant share in U.S. states between 1990 and 2018, we find surprisingly small differences between less-educated U.S.- and foreign-born men's exposure to workplace hazards. Differences are larger among women. For both sexes, the immigrant-native gap appears to have risen over time, coinciding with a rise in the immigrant share. Whether this is because immigration caused U.S.-born men and women to move into less-hazardous occupations is one of our two areas of focus here. The other is how immigration affected the return to working in a hazardous occupation.

We exploit several changes in immigration patterns that occurred between 1990 and 2018. Most notably, the origins and destinations of immigrants shifted. The share of less-educated working-age immigrants from Mexico rose through the mid-2000s and then fell, the share from Asia fell steadily, and the share from Central America rose. Immigrants became more geographically dispersed across the United States, with fewer living in most of the traditional destination states, particularly California. We use changes in origins and destinations to create measures of changes in immigrant shares within states over time that are plausibly exogenous to

labor market conditions that might affect U.S. natives' occupational choices and compensating differentials. Results using these instrumental variables are generally consistent with the ordinary least squares results, suggesting little endogeneity bias.

The analysis includes separate results for men and women since they tend to work in different occupations and therefore have different patterns of exposure to workplace hazards. Men work in occupations with considerably more hazard exposure than women, on average, and the immigrant-native gap is larger among women than men. This suggests that effects of immigration may differ by sex. We indeed find this to be the case: during our sample period, immigration caused U.S.-born women's exposure to workplace hazards to fall 1.5 times as much as men's exposure. U.S.-born men and women alike shifted into communication-intensive occupations as immigration increased, but the instrumental variables results indicate that the returns to communication skills rose only among women as immigrant inflows increased. For men, the return to manual skills fell as immigrant inflow rose. Simple bivariate regressions indicate a decline in the compensating differential paid for hazard exposure due to immigration, but this effect disappears when controlling for changes in the returns to manual and communication skills.

The next section discusses why immigration might affect U.S. natives' exposure to occupational hazards and the compensating differential for such exposure. We then explain the data used to examine the effect of immigration on natives' job characteristics, the construction of the instrumental variables used to control for potential endogeneity, and the regression models. We then present the results.

Background

Immigration may affect native-born workers' exposure to workplace hazards by changing the types of jobs that natives hold. First, natives might shift to jobs that utilize their comparative advantage in order to reduce labor market competition with immigrants. If so, natives would move to jobs that emphasize skills that are relatively scarce among immigrants and away from jobs that emphasize skills that are relatively abundant among immigrants. Consistent with this, less-educated U.S. natives tend to move into communication-intensive jobs and out of manual labor-intensive jobs in response to immigration (Peri and Sparber, 2009). Manual labor intensity does not perfectly overlap with exposure to workplace hazards, but the two are strongly correlated, particularly for occupations typically held by less-educated workers.[4] Construction jobs exemplify this, with high manual labor intensity, relatively high exposure to high places, hazardous conditions, and hazardous equipment, and the highest injury and fatality rates among major occupational groups. Natives might also increase their human capital or move to a different geographic area in order to reduce labor market competition with immigrants (e.g., Hunt, 2017; Borjas, 2006), and these shifts could affect natives' exposure to workplace hazards by changing the types of jobs they hold.[5] Dillender and McInerney (2020) confirm that U.S.-born men work in less-hazardous occupations as Mexican immigration increases.

In addition to supply-side shifts, immigration may affect native-born workers' hazard exposure by changing the demand side of the labor market. Employers may change the types of jobs they offer in response to immigration.[6] If immigrants disproportionately have manual skills,

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[4] Appendix Figure 1 shows manual skill intensity and hazard exposure for occupations dominated by less-educated workers. Most of the dots are on or near the 45-degree line that indicates having the same percentile score for the two measures.

[5] However, other research concludes that there is little evidence that U.S. natives move in response to immigrant inflows, e.g., Card (2001, 2009) and Peri and Sparber (2011b).

[6] Employers also may change conditions within jobs in response to immigration and thereby affect natives' exposure to workplace hazards. Since the data we use hold constant conditions within jobs, we do not focus on this channel.

employers may increase the share of jobs that are manual labor intensive in areas that experience more immigration. Employers also might substitute labor for capital or slow their adoption of manual labor-saving technology in response to low-skilled immigration inflows. For example, employers in Miami used production technologies that were more unskilled labor intensive and were slower than employers in other cities to adopt computers after the Mariel boatlift dramatically increased Miami's low-skilled labor force in 1980 (Lewis, 2004). If employers change the occupational distribution of jobs in response to immigrant inflows, natives may see their own occupational distribution shift toward jobs that emphasize skills that are relatively abundant among immigrants. Since less-educated immigrants tend to be relatively young and healthy, have relatively little formal education, and lack English fluency, these jobs are likely to emphasize manual skills over communication skills and involve more exposure to workplace hazards. The theoretical impact of immigration on natives' occupational risks is thus ambiguous—supply-side factors predict a reduction, while demand-side responses to immigration can predict an increase in occupational risks.

The standard model of efficient labor markets predicts that riskier occupations pay higher wages. Most workers are willing to hold in a risky job only in exchange for a compensating differential, but workers differ in their willingness to trade off wages and occupational risks. Employers also differ in the compensating differential they are willing to pay workers to take on risk, based on their costs of reducing risk. In equilibrium, workers and employers sort into wage-risk combinations that involve workers with low willingness to incur risk holding low-risk jobs and employers with low costs of reducing risk offering low-risk jobs, and the converse. The magnitude (and direction) of the risk premium depends on the distributions of workers' preferences and employers' costs.

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Immigration can affect the risk premium by changing the distribution of workers'
preferences or employers' costs. Immigrants may have smaller wage-risk tradeoffs than U.S.
natives since they tend to be healthier than U.S natives. In addition, limited labor market
opportunities or different frames of reference may make immigrants more willing than native-
born workers to accept job risks (Dávila, Mora and González, 2011; Orrenius and Zavodny,
2013). Immigration then would cause the risk premium to fall. For employers, immigration may
mean higher costs of reducing risk. Costs of workplace safety training may be higher for workers
who are not fluent in English or not familiar with U.S. workplace norms, for example (Hersch
and Viscusi, 2010). If immigration increases employers' costs of reducing risk, the risk premium
falls (Hersch and Viscusi, 2010). Either way, areas experiencing larger immigrant inflows should
see compensating differentials for occupational risks fall.

Whether there actually is a compensating differential for occupational risks is unclear,
however. Most studies agree that there is a wage premium for jobs with higher fatality risks, but
there is mixed evidence of a wage premium for jobs with higher injury and illness rates or other
measures of poor working conditions (Viscusi, 1993; Viscusi and Aldy, 2003). Indeed, some
studies report wage penalties for worse jobs. Self-selection, unobservable characteristics of
workers and jobs, and incomplete information among workers about job conditions may underlie
researchers' failure to find evidence of higher wages in worse jobs. If a wage penalty, not a
premium, is observed for hazardous occupations, that penalty is likely to increase as a result of
immigration.

Immigration is also likely to cause a decrease in returns to skills that become relatively
more abundant as a result of immigration. Since less-educated immigrants typically have a
comparative advantage in manual skills, not communication skills, returns to manual skills are

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likely to fall and returns to communication skills to rise when less-skilled immigration increases. Previous research indicates that less-educated immigrants are more concentrated in manual labor-intensive occupations than U.S. natives. Peri and Sparber (2009) show that less-educated U.S. natives shift into communication-intensive occupations in response to immigration and estimate the elasticity of substitution between manual and communication-intensive work. We differ from their approach by instead assessing how immigration affects the rate of return paid to these skills and also examining hazard exposure.

We next explain the data on occupation-level hazard exposure and skill intensity and the methods we use to examine whether immigration affects U.S. natives along these dimensions as well as their returns to hazard exposure, manual skills, and communication skills.

Data and Methods

Our analysis requires measuring the extent to which workers' occupations expose them to hazardous conditions and utilize manual and communication skills. To do this, we turn to the U.S. Department of Labor's Occupational Information Network (O*NET), a program that surveys workers and experts about a wide variety of characteristics of occupations and workers in those occupations. Economists from a diverse set of subfields have used O*NET data to study labor issues such as immigration, wage disparities, offshoring, technological development, and COVID-19.[7] Our analysis uses version 24.2 of the O*NET Work Context and Worker Abilities surveys and most closely follows the methodologies of Peri and Sparber (2009) and Dillender and McInerney (2020).

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[7] See, for example, Goos et al. (2009), Chiswick and Miller (2010), Oldenski (2014), Fan et al. (2017), De Arcangelis and Mariani (2010), and Avdiu and Nayyar (2020).

O*NET respondents answer questions using a five-point Likert Scale. The Work Context component of the survey asks how frequently workers face particular job conditions, for example, "How often does your current job require that you be exposed to radiation?" with 1 representing "Never" and 5 representing "Every day." The Worker Abilities component of the survey asks about the importance of a skill to job performance, for example, "How important is oral comprehension to the performance of your current job?" with 1 representing "Not Important" and 5 representing "Extremely Important." We use the O*NET data for the questions summarized in Table 1.

O*NET provides mean Likert values for every occupation in the Standard Occupation Classification (SOC) system. We use a crosswalk to merge O*NET scores with 2010 American Community Survey (ACS) data from IPUMS (Ruggles et al., 2020) to create raw average O*NET values for each of 439 ACS occupations (IPUMS variable *occ2010*). We create composite measures for hazard exposure, manual skills, and communication skills using the mean values for the variables listed in Table 1. Following Peri and Sparber (2009), we convert those raw means into percentile scores for each occupation that measure the share of the U.S. labor force in 2010 that was less exposed to a given work condition or used less of a given skill. We then merge these occupation-specific measures to workers in the 1990, 2000, 2010, and 2018 Census/ACS.[8]

Table 1 shows the raw average and percentile average values for each O*NET question we use and our composite values. While the weighted median value of each percentile equals 0.50 by construction, weighted averages may differ. Averages among subsets of workers differ

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[8] Unless otherwise indicated, throughout the analysis we use IPUMS data on labor force participants ages 18-67 who are not enrolled in school, have at most completed high school, are not living in group quarters, and report an occupation that can be merged with the O*NET data.

even more. Figure 1 displays average hazard exposure percentile values for men (left panel) and women (right panel) by educational attainment and nativity in 2010.[9] Not surprisingly, male, less-educated, and foreign-born workers tend to be in more-hazardous occupations than female, more-educated, and U.S-born workers. We focus the remainder of the analysis on workers with at most a high school diploma. We do so because this group is particularly exposed to hazardous work conditions and because immigration has been disproportionately concentrated among less-educated workers.

Figure 2 displays the trends in hazard exposure for less-educated workers over time. Large male/female differences become more apparent. More interesting, however, is that while there is a pronounced and growing gap in hazard exposure between native- and foreign-born women, the corresponding difference among men is small and switches over time, with U.S.-born men slightly more exposed to hazardous conditions in 1990 but less exposed in 2010 and 2018. This motivates us to examine how hazard exposure among U.S. natives responds to immigration separately for men and women.

Since our measures of hazard exposure are occupation-specific and do not vary across time, sex, educational attainment, nativity, or other dimensions, differences in average exposure across groups emerge only due to differences in the distribution of workers across occupations. To get a better sense of sex differences, Table 2 provides examples of male-majority (top panel) and female-majority (bottom panel) occupations characterized by high levels of hazard exposure. The table also provides the female and immigrant shares of less-educated workers in these occupations in 2010. Hazardous jobs disproportionately held by native-born men include

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[9] The Appendix includes similar graphs for manual and communication skill intensity (Appendix Figures 2 and 3, respectively). The Appendix also shows trends over time in average manual and communication intensity for less-educated workers (Appendix Figures 4 and 5).

firefighters and police officers. Foreign-born men are more heavily represented among construction workers, janitors, and meat processors. Native-born women, in contrast, are more likely to work as dental assistants, flight attendants, and hairstylists, while foreign-born women disproportionately work as launderers, housekeepers, food servers, and filling machine operators.

The previous tables and figures illustrate that the demographic composition of workers varies considerably across occupations. Our occupation-level measure of hazard exposure therefore is correlated with workers' demographic characteristics. The same is true for our measures of manual and communication skills. This leads us to create a second set of measures for hazard exposure and skills that controls for demographic characteristics. To do so, we begin with the percentile values described above after merging them with individuals in the Census and ACS samples. We then run individual-level regressions of log-percentile values on educational attainment (high school dropout versus high school degree), age, race/ethnicity, employment status, sex, and nativity, as applicable.[10] The exponential of these residuals represents an individual's occupational hazard exposure, manual skill intensity, or communication skill intensity after accounting for demographic characteristics.

Regression Models

Our main focus in this study is how immigrant inflows to regional labor markets (i.e., states) affect U.S. natives' hazard exposure and the wage premium for workplace exposure to hazards. Thus, state-by-year cells represent our unit of analysis, and we collapse our measures of natives' occupational hazard exposure and skills to the state-by-year level. Again, we use two sets of

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[10] We use different samples and regressions depending upon the group of interest. For example, if we examine only native-born men, we remove women and foreign-born workers and adjust the regression variables accordingly. In all cases, log-residual values have a mean of zero by construction.

measures: average percentile scores and average residuals from regressions that control for observable characteristics, as described above.

Our regressions examining the effect of immigration on U.S. natives' exposure to occupational hazards are motivated by the simple model

$$ln\left(Haz_{s,t}\right) = \alpha + \beta\ ImmigShare_{s,t} + \delta\ Unemp\ rate_{s,t} + FE_s + FE_t + Trend_{s,t} + \varepsilon_{st}, \quad (1)$$

where *Haz* is the average hazard exposure among U.S.-born workers in state *s* and year *t*. We focus on the variable *ImmigShare*, which measures the share of workers who are foreign born. The coefficient of interest, β, therefore represents the percentage change in natives' hazard exposure from a 1 percentage point increase in a state's immigrant share. The regression includes the state's unemployment rate to control for business cycle effects on workplace hazard exposure and state and year fixed effects to capture unobservable state- or time-specific factors. The state-specific linear time trends capture smooth trends that vary across states.

We estimate a first-differenced version of equation 1 to absorb omitted factors specific to states in a particular year that might be correlated with our explanatory and outcome variables. This model is akin to models commonly used to examine the effect of changes in the immigrant share on wages:

$$\Delta\ ln\left(Haz_{s,t}\right) = \alpha + \beta\ \Delta ImmigShare_{s,t} + \delta\ \Delta Unemp\ rate_{s,t} + FE_s + FE_t + \Delta\varepsilon_{st}, \quad (2)$$

where first differencing effectively transforms the state-specific linear trends into state fixed effects. We continue to include year fixed effects in the model, but their interpretation is slightly different. We follow Card and Peri (2016) by measuring the change in the immigrant share as

$$\Delta ImmigShare_{s,t} = \frac{F_{s,t} - F_{s,t-1}}{F_{s,t-1} + N_{s,t-1}}, \tag{3}$$

where F is foreign-born workers and N is native-born workers. This variable is closely related to the change in the share of workers who are foreign born but has the advantage of being driven entirely by changes in the foreign-born workforce and not by movement of native-born workers across state lines or into/out of the labor force. The mean (and standard deviation) of this variable for the whole sample, men, and women is 0.036 (0.047), 0.040 (0.055), and 0.031 (0.041), respectively.

We estimate equation 2 for changes in U.S. natives' average occupational manual and communication skill intensity as well. As discussed above, we expect that increases in the immigrant share decrease U.S. natives' occupational hazard exposure and manual skill intensity (i.e., estimates of β are negative) and increase their occupational communication skill intensity (positive estimates of β).

Instrumental Variables

A common concern in studies of the labor market effects of immigration is that immigrant shares are related to shocks to the labor market outcomes under examination. Such endogeneity results in biased estimates of the effects of immigration on labor marker outcomes. The endogeneity concern for our study is the possibility that immigrants are attracted to states where demand for

occupations that involve hazards (or skills) is changing in a consistent manner, either positively

or negatively. To control for such endogeneity, we rely upon variants of the shift-share

instrumental variable to establish a causal link between immigration and changes in natives'

exposure to occupational hazards. The shift-share instrument was pioneered by Bartik (1991) and

popularized in the immigration literature by Card (2001). It is both widely used and controversial

within the immigration literature. Jaeger et al. (2018, p. 5) argue, "It is difficult to overstate the

importance of this instrument for research on the impact of immigration. Few literatures rely so

heavily on a single instrument or variants thereof."

The basic intuition of the shift-share instrument is that new immigrants are attracted to

the same areas that attracted previous immigrants. By combining immigrants' base-year

settlement patterns at the state level with subsequent growth rates in the immigrant workforce at

the national level, one can predict the size of a state's immigrant workforce in later years.

Distinguishing between immigrants by region of origin typically leads to better predictions. If

$F_{b,s,t}$ is the foreign-born workforce from birthplace b in state s at time t, then the predicted

foreign-born workforce, $\widehat{F_{b,s,t}}$, is

$$\widehat{F_{b,s,t}} = \left(\frac{F_{b,s,0}}{F_{b,0}}\right) \times F_{b,t}. \tag{4}$$

The first term in equation 4 represents the share of immigrants from birthplace b in state s during

some base year ($t = 0$). The second term represents the total number of immigrants from that

birthplace in the United States in year t. Under the assumption that – from the perspective of a

state in year t – base-year settlement patterns and national-level totals are exogenous, then the

sum of $\widehat{F_{b,s,t}}$ across all B birthplaces, $\widehat{F_{s,t}}$, can serve as an instrument for the observed foreign-born workforce in a state and year:

$$\widehat{F_{s,t}} = \sum_{b=1}^{B} \left(\frac{F_{b,s,0}}{F_{b,0}}\right) \times F_{b,t}. \tag{5}$$

We create two versions of the shift-share instrument: a basic version and a geographic version.

Basic Shift Share

First, we limit our sample to birthplaces that are recorded in all Census and ACS years in our study. Observed immigrant shares at the state level do not rely upon birthplace identification, so we make no such sample restriction to observed values. The instrument, however, requires base-year populations (at the state level) and subsequent total populations (at the national level), so we use a balanced panel to ensure a clear calculation of $\widehat{F_{b,s,t}}$ for all countries in our resulting instrument.[11]

Second, we predict the native-born workforce (N) in much the same way as in equation 5. One important difference, however, is that we replace a state's base-year workforce with the number of U.S. natives who were born in their state of residence, $\widetilde{N_{s,0}}$. National workforces in both the base year and year t, however, continue to include all U.S. natives regardless of their state of residence or birth. The predicted native-born workforce in a state is therefore:

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[11] We also combine some birthplaces, such as combining England, Wales, Scotland, Northern Ireland, the Falklands, and Gibraltar. The instrument is constructed using 112 countries based on 307 unique birthplaces (none of them U.S. states or territories) during the period covered by our dataset.

$$\widehat{N_{s,t}} = \left(\frac{\widetilde{N_{s,0}}}{N_0}\right) \times N_t. \tag{6}$$

Our explanatory variable measures the change in the foreign-born labor force normalized by the initial total labor force, $\Delta ImmigShare_{s,t} = \frac{F_{s,t} - F_{s,t-1}}{F_{s,t-1} + N_{s,t-1}}$. We therefore construct the final instrument in much the same way:

$$\Delta IV_{s,t} = \frac{\widehat{F_{s,t}} - \widehat{F_{s,t-1}}}{\widehat{F_{s,t-1}} + \widehat{N_{s,t-1}}}. \tag{7}$$

We calculate separate instruments for our subpopulations of interest (e.g., less-educated men, less-educated women, etc.). We use 1980 as the base year in all cases.

Geographic Shift Share

In principle, the basic shift-share methodology provides an exogenous prediction of observed foreign-born shares (or changes in shares) since it forecasts state workforces if original migration patterns and national population sizes are the sole determinants of a state's later workforce. It predicts state labor forces if natives and immigrants do not move across state lines in search of better economic opportunities. In practice, however, economists such as Adão et al. (2019) and Jaeger et al. (2018) point to serious weaknesses of this strategy. One important limitation is that if the factors that determined initial settlement patterns persist and influence subsequent settlement patterns as well, the instrument will not be exogenous. Similarly, factors might simultaneously influence the national-level growth of foreign-born workforces and state-level labor market outcomes, thereby biasing coefficients. As a step toward overcoming these

limitations, we adopt a modified shift-share methodology that exploits exogenous, mostly

geographic, variation across destination states and origin countries.

We begin by separating the two factors in equation 4, recognizing that the first

component, $\left(\frac{F_{b,s,0}}{F_{b,0}}\right)$, represents a state's share of immigrants from country b in the base year and

the second component, $F_{b,t}$, is simply the total number of immigrants from b in the United States

in year t.

Again choosing 1980 as the base year, we form a dataset composed of B^*S observations

and estimate the following regression:

$$\frac{F_{b,s,1980}}{F_{b,1980}} = \alpha + \delta_1 ASINH\left(F_{b,s,1920}\right) + \sum_{b=1}^{B} \delta_2^b \ln\left(Distance_{b,s}\right) + \delta_3 CAN_{b,s} + \delta_4 MEX_{b,s}$$

$$+FE_b + FE_s + \mu, \tag{8}$$

where $ASINH\left(F_{b,s,1920}\right)$ is the inverse hyperbolic sine[12] of the stock of foreign workers from b

in s in 1920, $Distance_{b,s}$ is the distance between the capitals of b and s, and CAN and MEX are

indicators for when b and s jointly identify a state border with Canada or Mexico.[13] The

regression includes fixed effects for origin countries and destination states. Predicted values from

the regression then represent a state's expected share of immigrants from a particular country in

the base year (1980) according to geographic features and the 1920 stock of immigrants.

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[12] The change in the inverse hyperbolic sine, like the change in the natural log, approximates a percentage change. Unlike natural logs, however, it is defined at zero.
[13] Distance data is from CEPII (http://www.cepii.fr/anglaisgraph/bdd/distances.htm).

Persistence in economic shocks after that date should not affect the exogeneity of the instrument.[14]

We follow a similar procedure to create the second term of equation 4, $F_{b,t}$, the total number of immigrants from b in the United States at year t. Here we create a cross-country panel dataset and estimate the regression

$$ln(F_{b,t}) = \alpha + \delta_1 \, ln(GDP_{b,t}) + \delta_2 \, ln(POP_{b,t}) + FE_b + FE_t + \mu, \tag{9}$$

where *GDP* and *POP* represent real gross domestic product and total population of origin country b in year t.[15] The intuition behind the regression model is that higher GDP in the origin should lead to smaller numbers of immigrants in the United States, while a higher population in the origin should lead to larger numbers of immigrants. The model is based on gravity models commonly used to predict immigration stocks and flows.

Finally, we multiply the predicted state shares from equation 8 by exponentials of the predicted log-workforces from equation 9 to create estimates of $\widehat{F_{b,s,t}}$ driven by exogenous forces alone. We then sum these values across countries and substitute the resulting measure of $\widehat{F_{s,t}}$ into equation 7, along with predicting the native-born workforce based on birthplace and national size, to create our geographic shift-share instrumental variable.

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[14] We omit observations in which $\frac{F_{b,s,1980}}{F_{b,1980}} = 0$. We similarly replace negative predicted values with 0.

[15] GDP and population data are from the World Bank's World Development Indicators database.

Hazard Compensation

As discussed above, if immigrants have different wage-risk tradeoffs than U.S. natives or if immigration increases employers' costs of reducing risk, an increase in immigrant labor supply could cause the compensating differential for hazardous work to decline. Estimating the effect of immigration on this premium (or penalty) requires first measuring it, but it is not directly observable. To construct a proxy for the compensating differential, we follow a methodology closely related to Peri and Sparber (2009) that is similar to the construction of our residual hazard exposure and skills measures described above.

First, we construct a measure of the average log-wage paid to workers in a given occupation (*o*), state (*s*), and year (*t*) after accounting for demographic characteristics. To do so, we use data on less-educated workers and estimate 204 separate state-by-year (51 states and 4 years) individual-level log-wage regressions.[16] Each regression controls for educational attainment (high school dropout versus high school degree), age, race/ethnicity, sex, and nativity. Crucially, the model also includes occupation fixed effects. The estimated coefficients on the occupation variables (when added to the constant term) represent the average residual log-wage paid to workers in a given occupation in a state and year after controlling for observable characteristics. We call this variable $ln(\widetilde{w_{o,s,t}})$.

We are interested in understanding whether changes in a state's foreign-born labor force share affect workers' compensating differential for hazardous conditions. Suppose we begin with the following model of residual log-wages paid to occupation *o* in state *s* and year *t*:

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[16] Only people who report a positive wage are included in the regressions.

$$ln\left(\widetilde{w_{o,s,t}}\right) = \alpha + \beta\, FB_{s,t} + \gamma\, ln(Haz_o) + \eta\, FB_{s,t} \times ln(Haz_o) + \rho\, ln(Haz_o) \times Trend_t +$$

$$\varepsilon_{o,s,t}. \tag{10}$$

This model explicitly recognizes that wages are related to the state's foreign-born share ($FB_{s,t}$), an occupation's hazard exposure (Haz_o), the interaction between the two, and a linear trend in the compensation paid for hazard exposure. As discussed above, economic theory predicts a wage premium for occupational hazards, but estimates of γ may not be positive because of unobservable heterogeneity that our residual approach cannot capture.

The error term, $\varepsilon_{o,s,t}$, may include a time-invariant component that is correlated with the regressors. This would bias the estimated coefficients, so we instead write the model in first differences over time:

$$\Delta ln\left(\widetilde{w_{o,s,t}}\right) = \beta\, \Delta FB_{s,t} + \eta\, \Delta FB_{s,t} \times ln(Haz_o) + \rho\, ln(Haz_o) + \Delta\varepsilon_{o,s,t}. \tag{11}$$

The first-differences regression sacrifices our ability to estimate the return to hazard exposure, γ, directly, but it still includes the parameter of interest, η, which measures how a 1 percentage point change in a state's immigration share affects the compensating differential for occupational hazards above and beyond its trend, ρ.

A potential concern about equation 11 is that the error term in the first-differences regression may still include occupational trends that are correlated with hazard exposure or state trends that are correlated with the growth in the foreign-born share. We can better account for these possibilities by including fixed effects that fully absorb $\Delta FB_{s,t}$ and $ln(Haz_o)$ but not their interaction:

$$\Delta ln\left(\widetilde{w_{o,s,t}}\right) = \eta \, \Delta FB_{s,t} \times ln(Haz_o) + FE_{s,t} + FE_o + \Delta\varepsilon_{o,s,t}. \qquad (12)$$

We use the regression model in equation 12 to estimate η. One small deviation from this approach, however, is that our measure of ΔFB is, as above in equation 3, measured as the change in the foreign-born labor force normalized by the initial total labor force rather than the change in the foreign-born share. As before, we instrument for the change in the immigrant share because of concerns about potential endogeneity, and we estimate equation 12 with occupation-level measures of manual and communication skills as well as with hazard exposure.

Results

We first estimate the relationship between changes in the immigrant share within states over time and changes in U.S. natives' hazard exposure at work. Those results, together with results on the corresponding relationships for manual and communication skill intensity, then motivate our examination of changes in the returns to hazard exposure and skill intensity.

Hazard Exposure and Skill Intensity

We begin by examining the effect of immigration on the characteristics of the occupations held by U.S.-born workers. As Table 3 shows, immigration reduces less-educated U.S. natives' occupational hazard exposure. In the OLS regressions, there is a significant negative coefficient overall and among men but not among women. The IV results indicate larger declines in occupational hazard exposure than the OLS results, and all of the results for hazard exposure are

significant when using either instrumental variable.[17] The IV results suggest that a 1 percentage point increase in the foreign-born labor force relative to the initial total labor force reduces less-educated U.S.-born men's hazard exposure by about 0.2 percent, and women's by 0.2 to 0.4 percent. Alternatively, the average shock to a state's immigrant share in our sample led to about a 0.8 percent decline in native men's hazard exposure and a 0.6 percent to 1.2 percent decline for native women.

The results for manual skills are similar: immigration reduces manual skill intensity among U.S. natives. The estimated effects are bigger for women than for men, and the IV results are again bigger than the OLS results, albeit not significantly so for women. As immigration pushes U.S. natives out of occupations that are manual intensive, they move into occupations that are communication intensive. As the third row of Table 3 shows, increases in the immigrant share boost communication skill intensity for both men and women in all but one of the specifications estimated.

The finding that increases in the immigrant share caused less-educated U.S. natives to shift from manual-intensive jobs to communication-intensive ones over 1990 to 2018 is consistent with Peri and Sparber's (2009) findings for 1970 to 2000. The post-2000 period encompassed a shift in less-educated immigration, with smaller inflows from Mexico and more from Central America. Nonetheless, the effects were similar among U.S. natives.

The results in Table 3 could arise from changes in labor market composition. If, for example, immigrants displace natives whose demographic characteristics are correlated with

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[17] The increase in the point estimates (in absolute value) could be due to classical measurement error in the immigration variable that is smaller in the instrumental variables. Alternatively, the local average treatment effects identified by the IV regressions may be bigger than the average treatment effects in the OLS regressions. Weak instruments seem unlikely given the F-statistics for the first-stage regressions reported in the bottom row of the table. We cannot test the validity of the exclusion restriction, but there is little reason to believe the geographic shift share measure is invalid.

high hazard exposure, high manual skill intensity, and low communication skill intensity, then an increase in immigration will change the composition of the native workforce in such a way that the remaining sample has hazard, manual, and communication values consistent with the results in Table 3. It is not possible to measure O*NET values for individuals without an occupation reported in the Census/ACS data, nor do we think regressions with state-level controls for average demographics would provide an adequate solution for addressing this limitation. There is fairly little variation in changes in average demographics, such as sex and age, of the U.S.-born workforce during the time period we examine.

Instead, we construct an alternative set of residual O*NET values as described above. This uses individual- and occupation-level variation to remove demographic correlates with skill and job characteristics. Table 4 displays results from regressions using these alternative measures. We find qualitatively similar results that are somewhat larger in magnitude than those reported in Table 3. This implies that immigration caused a decrease in the measurable portion of U.S. natives' manual skill intensity and hazard exposure that is not explained by demographic characteristics such as education, sex, age, and race. Similarly, immigration caused natives' communication skill intensity to rise even when accounting for their demographic characteristics. Thus, the main results presented in Table 3 are driven by U.S. natives' movement across occupations in response to immigration, not by labor market entry and exit that is systematically related to their demographic characteristics.

Unfortunately, our methodology does not permit identification of the occupations U.S. natives are moving to or from – a limitation shared by previous studies that rely on spatial variation to identify how natives' skill levels respond to immigration. ACS and Census data do not provide information on individuals' occupations over time. While the Current Population

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Survey does do so, that sample is much smaller and using it to look at data by state and occupation would lead to considerable measurement error when looking at data defined for the intersection of states and occupations. That sample would also create selection bias since we could look only at workers who do not move housing units.

Compensating Differentials and Returns to Skills

As the immigrant share increases, the return to working in an occupation with more exposure to hazardous conditions appears to fall. Column 1 of Table 5 reports OLS results from estimating equation 12, and columns 1 and 5 of Table 6 report the corresponding IV results. Most of the results indicate that a 1 percentage point increase in the immigrant share reduces an occupation's return to hazard exposure by about 0.1 to 0.2 percent, on average. Alternatively, the coefficient in column 5 of Table 6 (-0.169) implies that the average immigration shock to a state during this period (0.036) caused the wage elasticity of hazard exposure to decline by 0.006 points relative to its trend – a small but statistically significant effect. The estimates indicate larger effects among men than among women, and the estimated coefficient on the interaction term is not significantly different from zero for women when using the geographic shift-share IV.

The results for manual skill intensity are again similar to the results for hazard exposure. As column 2 of Table 5 shows for the OLS regressions and columns 2 and 6 of Table 6 show for the IV regressions, the return to manual skills falls as the immigrant share rises. The effect is again larger for men than for women and statistically insignificant for women when using the geographic shift-share IV. The results for communication skills are the opposite, with the return to communication skills rising as the immigrant share rises. The estimated coefficients for that interaction effect do not differ between men and women.

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Finally, we include all three occupation characteristics interacted with the change in the immigrant share in the wage change regressions to examine how immigration affects the return to hazard exposure when controlling for its impact on the returns to manual and communication skills requirements. The results, shown in column 4 of table 5 and columns 4 and 8 of Table 6, no longer indicate that changes in the immigrant share affect the return to hazard exposure. Overall and for women, only the interaction term with communication skills is statistically significant. For men, the interaction term with manual skills is significant in all three models, and the interaction term with communication skills is significant as well in the OLS regression.

Taken together, the results indicate that men in jobs with high levels of hazard exposure earn less as the immigrant share increases, with the effect occurring because immigration reduces the return to manual skills and hazardous jobs are typically manual-intensive jobs. Men with hazardous jobs that are not manual intensive do not see their wages fall as immigration increases. The estimates suggest that, among men, immigrants are more concentrated in manual-intensive jobs than in hazardous jobs compared with U.S. natives. Indeed, this is the case, as confirmed by male immigrants' and U.S. natives' average hazard exposure and manual intensity in Figure 1 and Appendix Figure 2. However, the difference narrowed slightly over 1990 to 2018, as suggested by the trends in Figure 3 and Appendix Figure 4. That said, there was little change over time in U.S.-born men's hazard exposure relative to their manual skill intensity. This is not surprising since the two measures are highly correlated, as noted above. In short, men who hold hazardous, manual-intensive jobs see the return to those jobs fall as the immigrant share rises.

For women, the increase in the return to communication skills as immigration increases is not surprising. Communication-intensive jobs are U.S. natives' comparative advantage, and less-educated U.S.-born women are more likely to work in communication-intensive jobs than less-

educated men (Appendix Figures 3 and 5). Jobs that emphasize communication skills tend to have relatively low hazard exposure, although there are exceptions. U.S.-born women thus appear to benefit from immigrant inflows: they move into jobs that emphasize communication skills, and the return to those skills rises.

Conclusion

In recent decades, the United States experienced large inflows of immigrants, many of them with relatively low levels of education. The surge in less-educated immigrants generated considerable social and political controversy. Much of this controversy centered on concerns that less-educated immigrants negatively affected labor market opportunities for less-educated American workers. Research supporting such concerns is mixed. This study adds to a growing literature that shows that increased immigration resulted in less-skilled U.S. natives moving into "better" jobs. We show that over the period 1990 to 2018, less-educated U.S. natives moved out of occupations that involved relatively high exposure to hazards or intensive manual labor and into occupations with greater emphasis on communication skills. The movement away from hazardous occupations was more pronounced among women than men, a new and perhaps surprising finding.

We further contribute to the literature new findings on the effect of immigration on the returns to job characteristics. Increases in immigration appear to have reduced the compensating differential for hazard exposure. However, that negative effect occurred because increases in immigration reduced the return to manual-intensive jobs, which also tend to expose workers to hazardous conditions. Meanwhile, the return to communication-intensive jobs rose for women as a consequence of immigration. This benefitted U.S.-born women who moved into those jobs as

the immigrant share rose. The results thus indicate that large-scale increases in immigration are not detrimental to native-born workers, on average, even at the lower end of the education distribution.

Research that follows individuals over time and examines how their jobs change in response to immigration would shed additional light on immigration's impacts. Individual-specific measures of job characteristics, rather than occupation-level measures as used here, are particularly needed to more fully understand how immigration affects working conditions and the returns to those conditions. Regardless of future immigration patterns, understanding how immigration – or the lack thereof – affects natives' working conditions and returns to those conditions are important issues.

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Table 1 O*NET scores for hazard exposure, manual skills and communication skills

Variable	Raw Mean	Percentile Mean
Hazard exposure composite	*1.72*	*0.52*
(Std. dev.)	*(0.53)*	*(0.29)*
Exposed to radiation	1.16	0.54
Exposed to disease or infections	1.84	0.48
Exposed to high places	1.53	0.52
Exposed to hazardous conditions	1.80	0.52
Exposed to hazardous equipment	1.93	0.53
Exposed to minor burns, cuts, bites, or stings	2.05	0.51
Manual skills composite	*1.93*	*0.51*
(Std. dev.)	*(0.56)*	*(0.30)*
Arm-hand steadiness	2.44	0.52
Manual dexterity	2.39	0.52
Finger dexterity	2.65	0.53
Control precision	2.27	0.52
Multilimb coordination	2.24	0.52
Response orientation	1.72	0.54
Rate control	1.70	0.54
Reaction time	1.83	0.53
Wrist-finger speed	1.73	0.53
Speed of limb movement	1.60	0.54
Static strength	2.06	0.52
Explosive strength	1.34	0.53
Dynamic strength	1.82	0.51
Trunk strength	2.35	0.51
Stamina	1.93	0.52
Extent flexibility	1.99	0.52
Dynamic flexibility	1.13	0.63
Gross body coordination	1.84	0.52
Gross body equilibrium	1.71	0.53
Communication skills composite	*3.51*	*0.51*
(Std. dev.)	*(0.47)*	*(0.29)*
Oral comprehension	3.72	0.53
Written comprehension	3.44	0.53
Oral expression	3.67	0.51
Written expression	3.22	0.52

Note: Shown are average scores from O*NET data across 439 occupations.

Table 2 Examples of occupations with high levels of hazard exposure

Occupation	Hazard score	% female	% immigrant
Male-majority occupations:			
Firefighters	1.00	0.03	0.03
Automotive service technicians and mechanics	0.96	0.01	0.21
Police officers and detectives	0.95	0.14	0.04
Construction laborers	0.93	0.02	0.40
Carpenters	0.92	0.01	0.30
Janitors and building cleaners	0.84	0.34	0.32
Couriers and messengers	0.82	0.16	0.20
Farmers, ranchers, and other agricultural managers	0.79	0.13	0.09
Butchers and other meat, poultry, and fish processing workers	0.75	0.28	0.42
Female-majority occupations:			
Dental assistants	0.97	0.97	0.18
Flight attendants and transportation workers and attendants	0.97	0.83	0.14
Nonfarm animal caretakers	0.87	0.67	0.15
Packaging and filling machine operators and tenders	0.86	0.55	0.44
First-Line supervisors of food preparation and serving workers	0.78	0.65	0.18
Laundry and dry-cleaning workers	0.69	0.65	0.44
Food servers, nonrestaurant	0.66	0.73	0.28
Hairdressers, hairstylists, and cosmetologists	0.66	0.92	0.20
Maids and housekeeping cleaners	0.63	0.89	0.52

Note: Shown are O*NET percentile scores for the occupations indicated and the share of less-educated workers who are female or foreign-born in those occupations in the 2010 ACS.

Table 3 Estimated effect of immigration on U.S. natives' job characteristics

	OLS			Shift-share IV			Geographic shift-share IV		
	Men & women	Men	Women	Men & women	Men	Women	Men & women	Men	Women
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Hazard exposure	-0.105***	-0.092***	-0.125	-0.166***	-0.195***	-0.203**	-0.211***	-0.202***	-0.429***
	(0.037)	(0.034)	(0.118)	(0.034)	(0.033)	(0.096)	(0.045)	(0.071)	(0.104)
Manual skills	-0.145***	-0.106***	-0.217*	-0.199***	-0.211***	-0.289***	-0.199***	-0.216***	-0.301***
	(0.032)	(0.031)	(0.113)	(0.034)	(0.034)	(0.093)	(0.056)	(0.074)	(0.106)
Communication skills	0.233***	0.243***	0.217***	0.270***	0.391***	0.233***	0.269***	0.419***	0.194
	(0.046)	(0.060)	(0.069)	(0.044)	(0.063)	(0.058)	(0.068)	(0.099)	(0.133)
First-stage F-statistic	--	--	--	38.94	29.27	43.96	39.69	38.19	17.72

*** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$

Note: Shown are estimated coefficients on the change in the immigrant share as in equation 2. Each estimate is from a separate regression. All regressions have 153 observations and also include the change in the unemployment rate and state and year fixed effects.

Table 4 Estimated effect of immigration on U.S. natives' job characteristics, residual measures

	OLS			Shift-share IV			Geographic shift-share IV		
	Men & women (1)	Men (2)	Women (3)	Men & women (4)	Men (5)	Women (6)	Men & women (7)	Men (8)	Women (9)
Hazard exposure	-0.187*** (0.060)	-0.144*** (0.049)	-0.210 (0.164)	-0.288*** (0.056)	-0.303*** (0.044)	-0.312** (0.149)	-0.365*** (0.068)	-0.352*** (0.089)	-0.556*** (0.132)
Manual skills	-0.268*** (0.068)	-0.178*** (0.049)	-0.354** (0.165)	-0.390*** (0.065)	-0.366*** (0.050)	-0.484*** (0.143)	-0.371*** (0.106)	-0.384*** (0.116)	-0.465*** (0.151)
Communication skills	0.404*** (0.070)	0.360*** (0.066)	0.393*** (0.141)	0.480*** (0.068)	0.535*** (0.074)	0.470*** (0.114)	0.476*** (0.124)	0.570*** (0.144)	0.420** (0.206)
First-stage F-statistic	--	--	--	38.94	29.27	43.96	39.69	38.19	17.72

*** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$

Note: Shown are estimated coefficients on the change in the immigrant share as in equation 2. The dependent variable is based on residuals after controlling for education, age, race/ethnicity, employment status, and sex, as applicable. Each estimate is from a separate regression. All regressions have 153 observations and also include the change in the unemployment rate and state and year fixed effects.

Table 5 Estimated effect of immigration on compensating differentials and returns to skills, OLS

	(1)	(2)	(3)	(4)
Men & women				
Hazard exposure × Change in immigrant share	-0.149***			-0.020
	(0.035)			(0.062)
Manual skills × Change in immigrant share		-0.190***		-0.060
		(0.025)		(0.055)
Comm. skills × Change in immigrant share			0.173***	0.143***
			(0.024)	(0.028)
Men				
Hazard exposure × Change in immigrant share	-0.139***			0.069
	(0.042)			(0.069)
Manual skills × Change in immigrant share		-0.219***		-0.166**
		(0.032)		(0.064)
Comm. skills × Change in immigrant share			0.157***	0.122***
			(0.035)	(0.044)
Women				
Hazard exposure × Change in immigrant share	-0.091*			-0.029
	(0.052)			(0.063)
Manual skills × Change in immigrant share		-0.103**		0.047
		(0.050)		(0.074)
Comm. skills × Change in immigrant share			0.168***	0.177***
			(0.057)	(0.063)

*** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$

Note: Shown are estimated coefficients on the interaction of the occupation characteristic and the change in the immigrant share as in equation 12. The dependent variable is the change in the residual log-wage in an occupation, state, and year. All regressions also include state-year and occupation fixed effects.

Table 6 Estimated effect of immigration on compensating differentials and returns to skills, IV

	Shift-share IV				Geographic shift-share IV			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Men & women								
Hazard exposure × Change in immigrant share	-0.168*** (0.030)			-0.076 (0.055)	-0.169*** (0.045)			-0.051 (0.053)
Manual skills × Change in immigrant share		-0.189*** (0.028)		-0.035 (0.054)		-0.210*** (0.046)		-0.089 (0.058)
Comm. skills × Change in immigrant share			0.153*** (0.023)	0.117*** (0.026)			0.150*** (0.037)	0.100** (0.034)
Men								
Hazard exposure × Change in immigrant share	-0.151*** (0.028)			0.054 (0.064)	-0.203*** (0.041)			0.035 (0.065)
Manual skills × Change in immigrant share		-0.229*** (0.022)		-0.201*** (0.072)		-0.293*** (0.049)		-0.278*** (0.091)
Comm. skills × Change in immigrant share			0.129*** (0.034)	0.079 (0.048)			0.123*** (0.029)	0.047 (0.033)
Women								
Hazard exposure × Change in immigrant share	-0.076** (0.038)			-0.043 (0.048)	-0.029 (0.052)			0.011 (0.067)
Manual skills × Change in immigrant share		-0.076* (0.043)		0.068 (0.063)		-0.045 (0.059)		0.050 (0.078)
Comm. skills × Change in immigrant share			0.138*** (0.047)	0.151*** (0.051)			0.119** (0.061)	0.144** (0.061)

*** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$

Note: Shown are estimated coefficients on the interaction of the occupation characteristic and the change in the immigrant share as in equation 12. The change in the immigrant share is instrumented as indicated. The dependent variable is the change in the residual log-wage in an occupation, state, and year. All regressions also include state-year and occupation fixed effects.

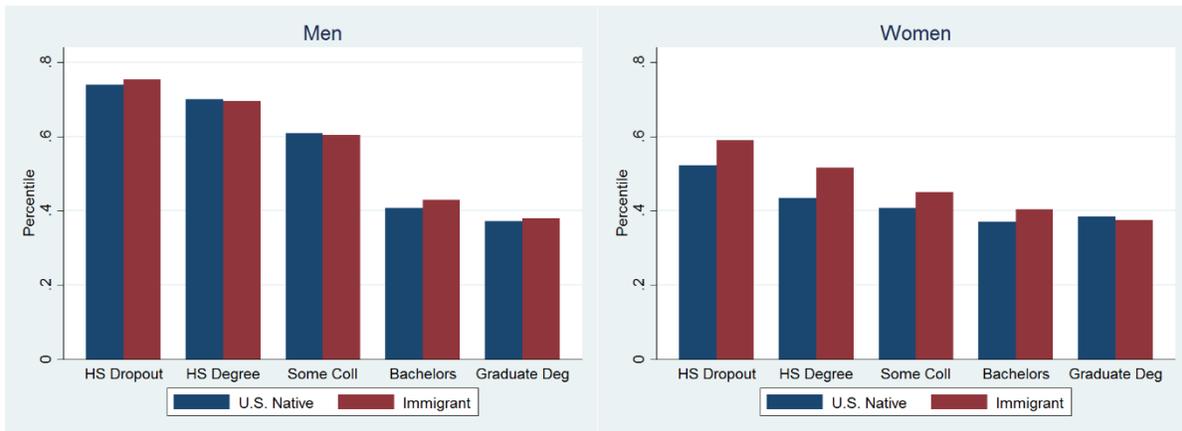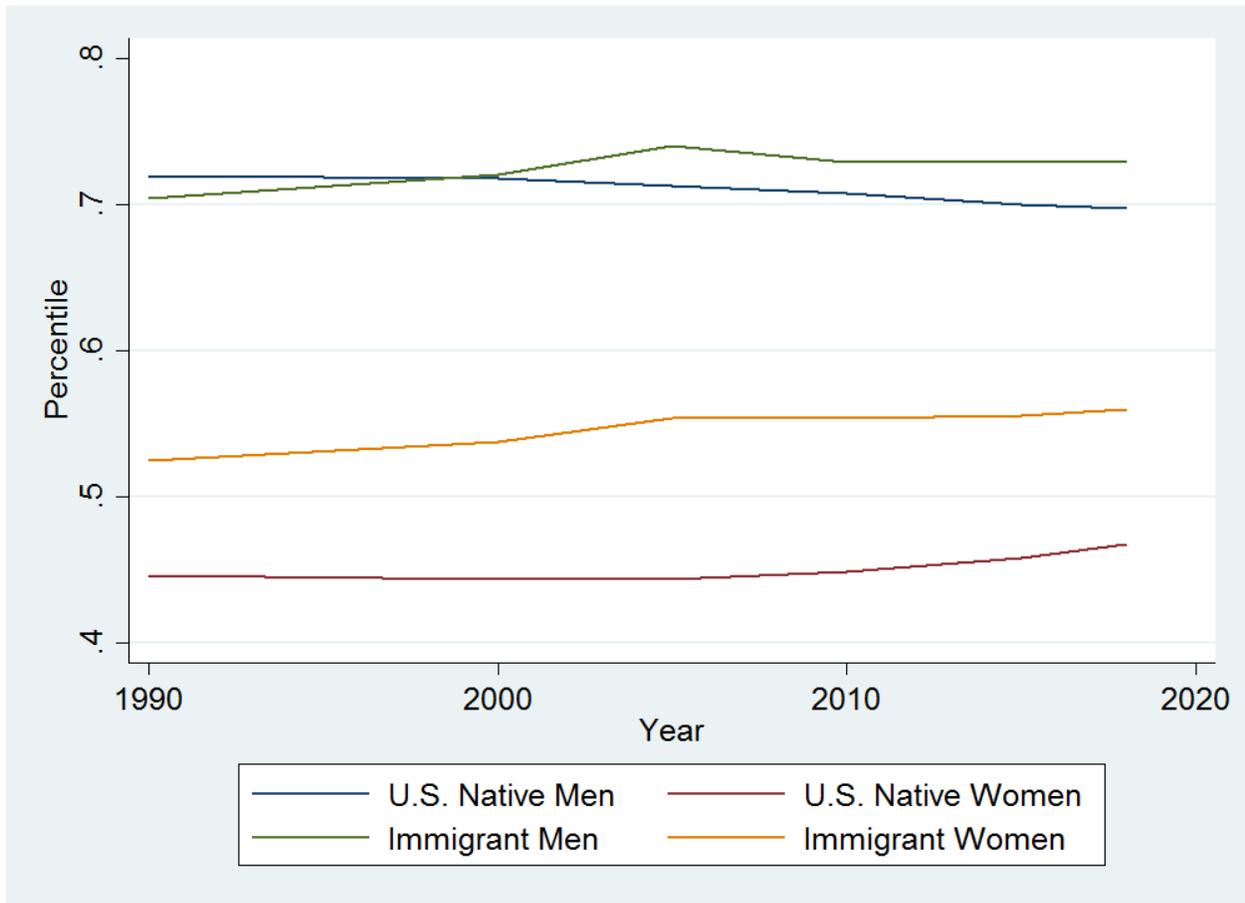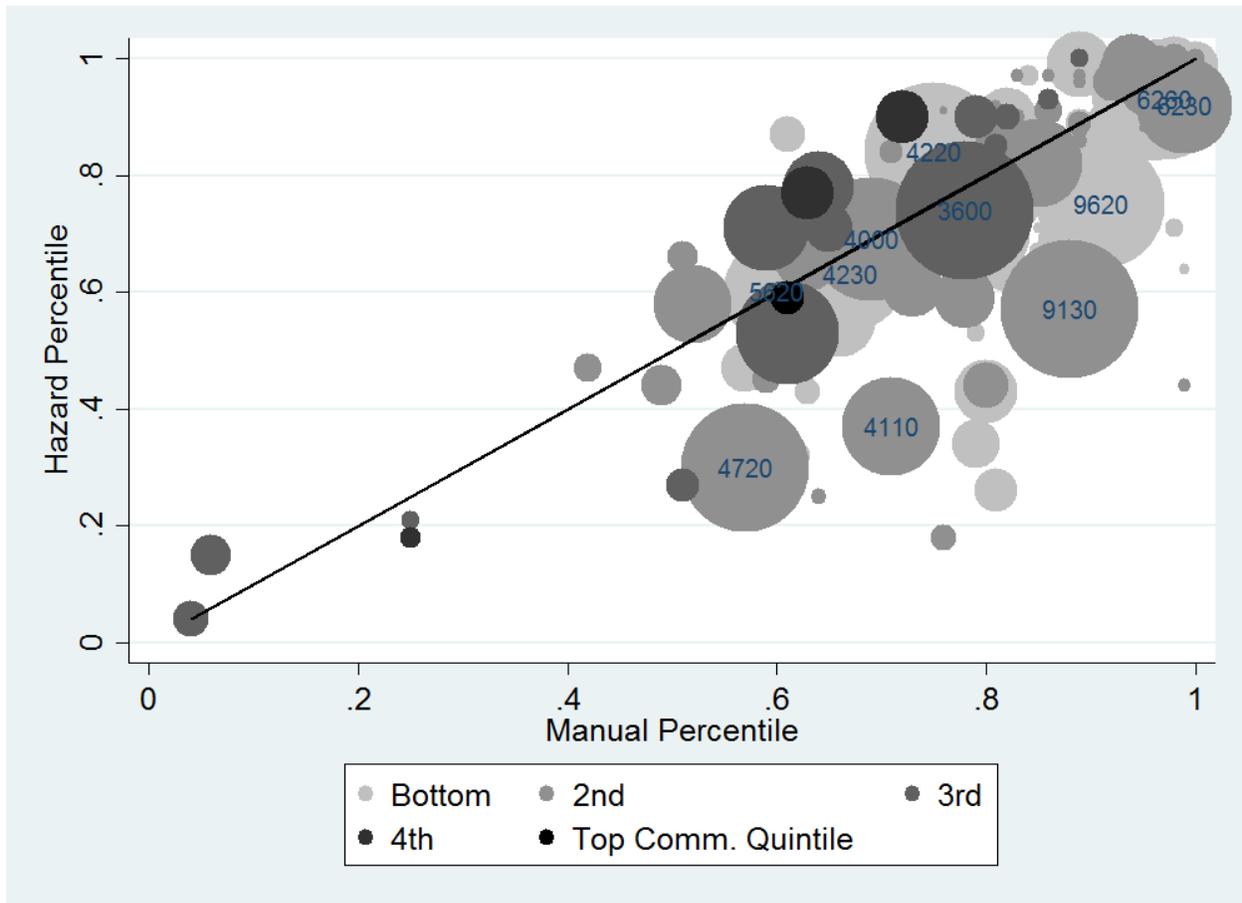**Figure 1 Exposure to hazardous work conditions, by education, sex and nativity, 2010**

Figure 2 Exposure to hazardous work conditions, by sex and nativity, 1990–2018

Appendix Figure 1 Manual skills and hazard exposure, by communication skills quintile


Note: Each dot represents an occupation in which at least half of workers have at most a high school diploma. Bigger dots have more workers. Occupations on the 45-degree line are have the same percentile for manual skill and hazard exposure (such as 4000=Chefs and Cooks). Occupations below the line are relatively higher in manual skill intensity than hazard exposure (4720=Cashiers; 4110=Wait staff; 9130=Truck Drivers; 9620=Laborers). Those above the line are disproportionately risker (4220=Janitors). Darker dots have higher communication intensity (3600=Nursing Aides).
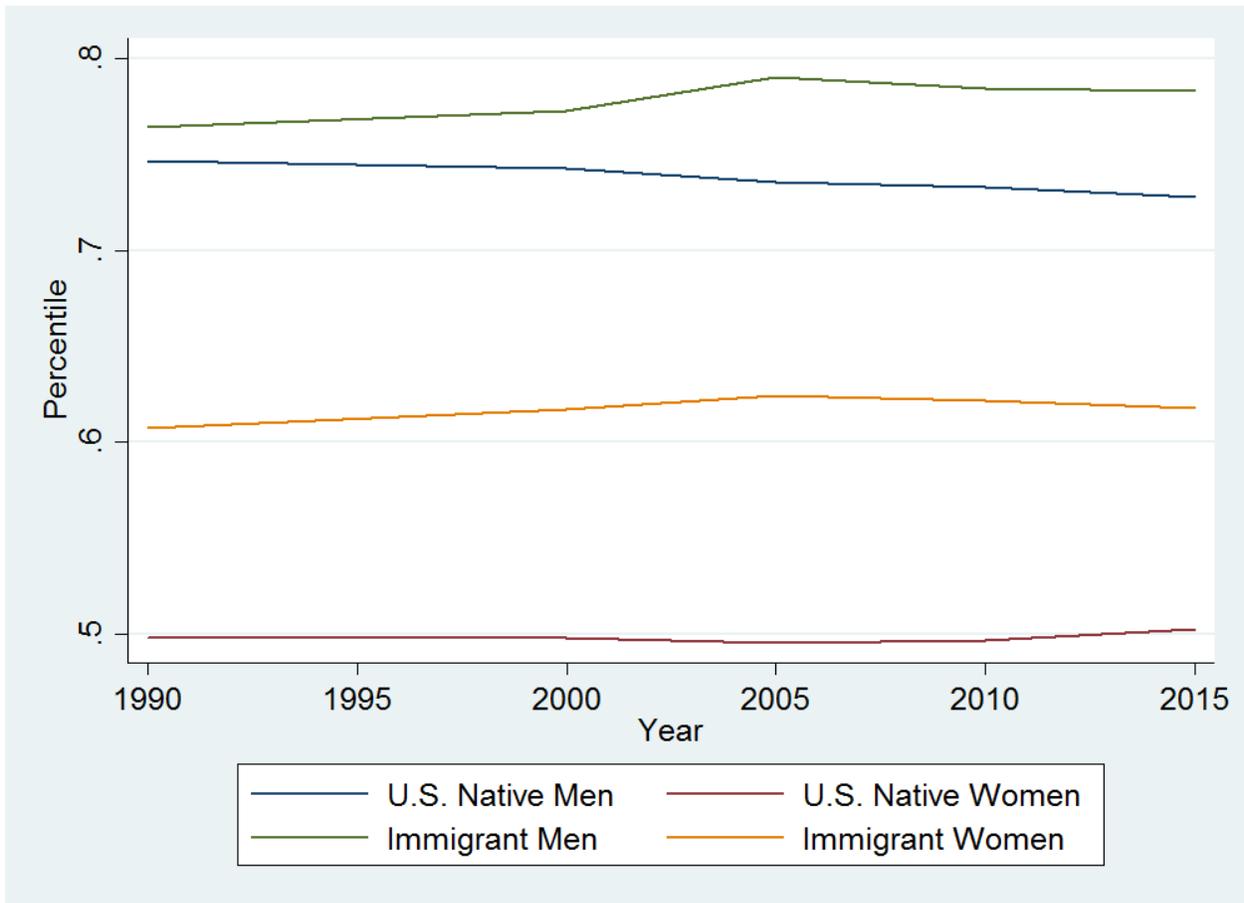
Appendix Figure 2 Manual skills, by education, sex, and nativity, 2010
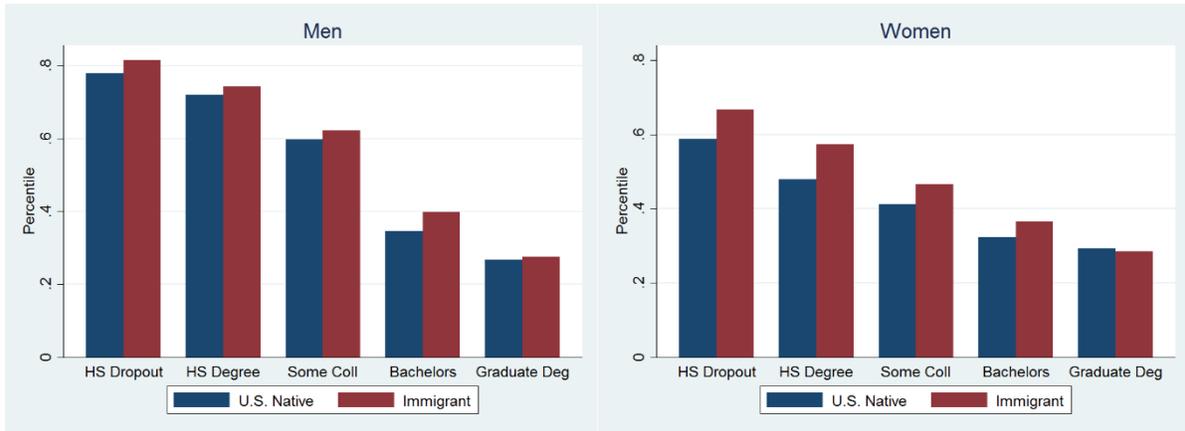
Appendix Figure 3 Communication skills, by education, sex, and nativity, 2010
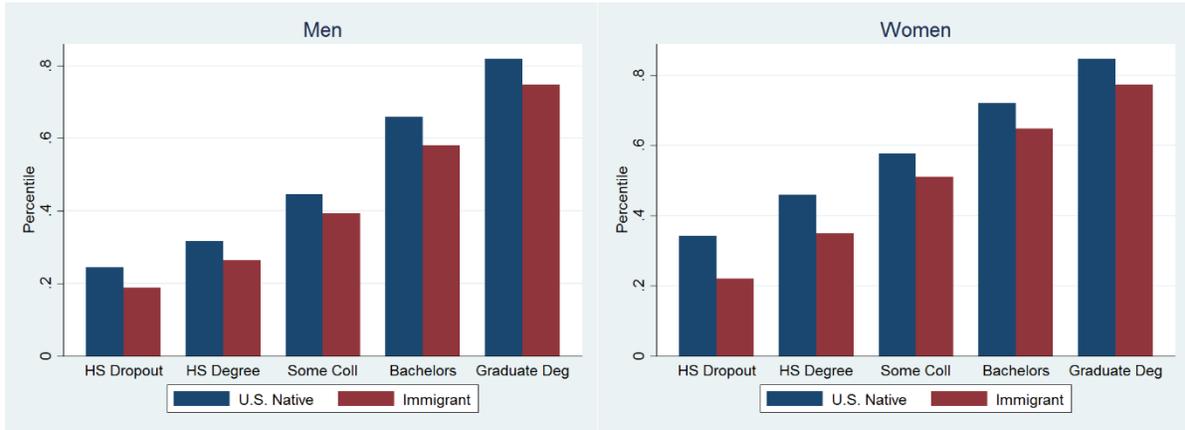
Appendix Figure 4 Manual skills, by sex and nativity, 1990–2018

Appendix Figure 5 Communication skills, by sex and nativity, 1990–2018