TRADING PLACES: MOBILITY RESPONSES OF NATIVE- AND FOREIGN-BORN ADULTS TO THE CHINA TRADE SHOCK

DAVID AUTOR, DAVID DORN, AND GORDON HANSON*

Previous research finds that the greater geographic mobility of foreign-born workers compared to native-born workers facilitates labor market adjustment to shifting regional economic conditions. The authors examine immigration's role in enabling US commuting zones to respond to manufacturing job loss caused by import competition from China. Although foreign-born population headcounts fell by a larger proportion than those of the native-born in trade-exposed regions, the contribution of immigration to labor market adjustment in the study period was small. Because most US immigrants arrived in the country after manufacturing regions were already mature, few took jobs in industries that later saw import surges. The foreign-born population share in regions with high trade exposure was only three-fifths that of regions with low trade exposure. Immigration may do more to aid adjustment to cyclical shocks, in which job loss follows recent hiring booms, than to aid adjustment to secular decline, in which hiring booms occurred longer ago.

Empirical analysis of migration provides abundant evidence that foreign-born and native-born workers differ in how they make location choices within national borders. As adults, the native-born tend to settle close to where they lived as children (Sprung-Keyser, Hendren, and Porter 2022),

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which may contribute to why supplies of less-educated native-born labor are largely unresponsive to adverse changes in local labor demand (Topel 1986; Bound and Holzer 2000; Notowidigdo 2020; Zabek 2024). For the foreign-born, mobility is often built into their working lives. Mexican immigrants long traveled back and forth across the US–Mexico border, working in US agriculture and construction during warmer months and returning to Mexico for the winter season (Durand, Massey, and Zenteno 2001; Woodruff and Zenteno 2007). Since 1990, the expansion of US temporary work visas has directed new immigrants into US regions in which employment growth happens to be strong in their arrival year (Clemens and Lewis 2022). For other migrants, mobility may serve to maximize short-run nominal income, so as to support family members back home or to attain savings objectives (Dustmann and Görlach 2016; Albert and Monras 2020). Whatever the source of differential mobility patterns, if the foreign-born are indeed more footloose than the native-born, then, as Borjas (2001) hypothesized, immigration may "grease the wheels of the labor market" by easing adjustment to shocks.

The literature documents instances in which immigration has helped smooth labor market adjustment. The original analysis in Borjas (2001) found that over the 1960 to 1980 period, the supply of newly arrived immigrants was larger in US states with higher initial earnings. During the same period, wage convergence across US states was more rapid among skill groups that had a larger immigrant presence. More broadly, immigration helped accommodate changes in the US economy that after 1960 induced the population to shift to the South and West and from cities to suburbs (Borjas et al. 1997; Amior 2024).[1] Immigration also appears to aid in adjustment to cyclical fluctuations. During the Great Recession, the collapse of the US housing market caused sudden job loss in regions that had been caught up in the subprime mortgage lending boom (Mian and Sufi 2014). Cadena and Kovak (2016) showed that over the 2006 to 2010 period, whereas net migration of less-educated native-born men was unresponsive to regional changes in labor demand, less-educated foreign-born men were substantially responsive to the same shocks.[2]

This article examines the role of immigration in adjustment to another well-studied labor market shock, the decline in manufacturing due to global import competition. Increased Chinese manufacturing exports during the 1990s and early 2000s caused widespread job loss in many countries (Autor, Dorn, and Hanson 2016; Redding 2020; Dorn and Levell 2021). US commuting zones (CZs) that were exposed to the China trade shock had larger reductions in manufacturing employment, earnings, and employment–

[1]Related work considers how immigration affects regional innovation and productivity (e.g., Hunt and Gauthier-Loiselle 2010; Kerr and Lincoln 2010; Peri 2012, 2016; Stuen, Mobarak, and Maskus 2012; and Burchardi et al. 2020).

[2]The findings in Monras (2020) suggest that immigration contributes to labor-market adjustment more through the in-migration of labor than through the out-migration of labor.

population ratios, while also suffering deteriorating outcomes across a wide range of other indicators (Autor, Dorn, and Hanson 2013, 2019; Autor, Dorn, Hanson, and Song 2014; Pierce and Schott 2020). Although trade-exposed regions did see larger net declines in working-age populations, these were small in the aggregate and concentrated among the young (Greenland and Lopresti 2016; Greenland, Lopresti, and McHenry 2019; Faber, Sarto, and Tabellini 2022). Extending the analysis out to 2019, Autor, Dorn, and Hanson (2021) found that long-run differences in population growth across regions related to trade exposure were also muted, as well as suggestive evidence that headcounts of the foreign-born may have been more responsive than those of the native-born.[3] The literature still lacks a full accounting of how immigration intersects with adjustment to trade shocks, and in particular whether immigration may have eased labor-market pressures on the less-educated workers who appear to have been most exposed to import competition. We ask whether trade-exposed regions that had larger initial foreign-born populations had larger net out-migrations of labor, if those adjustments varied across individuals according to their educational attainment, and if the magnitude of those adjustments were sufficient to offset the impact of trade shocks.

As a persistent contractionary shift in labor demand, increased import competition from China represents a type of shock that the literature on immigration and labor market adjustment has yet to consider. In Borjas et al. (1997) and Borjas (2001), the shifts in motion were ones that increased the desirability of the Sunbelt. After 1960, the availability of automobiles and air conditioning, the construction of interstate highways, and growth-friendly regulations increased population flows into Southern and Western cities (Baum-Snow 2007; Glaeser and Tobio 2008; Arkolakis, Costinot, and Rodríguez-Clare 2012; Mangum and Coate 2019). The mobility of the foreign-born, combined with rising immigration nationally, may have helped growing regions achieve steady-state size more rapidly. The analysis in Cadena and Kovak (2016) considered the role of immigration in adjustment to a negative shock—the Great Recession—but one that was ostensibly cyclical in nature. Because the early 2000s housing boom pulled workers into construction jobs in growing cities, adjustment to the ensuing housing bust may have been aided by the exodus of those recent arrivals. Indeed, Cadena and Kovak (2016) found that recession-induced reductions in supplies of foreign-born workers occurred in part by workers returning to their origin countries. We examine a case in which there is pressure for labor to leave regions that had been doing neither particularly well nor particularly poorly prior to the shock.

Further motivating our analysis is the unfortunate frequency with which large, persistent, negative, and localized labor demand shocks tend to

[3]See appendix figure A6 of Autor et al. (2021). Their results differentiate population adjustment to trade exposure by nativity but not by educational attainment.

occur. Import competition from China is one of several factors that have contributed to regional manufacturing job loss in recent decades (Charles, Hurst, and Schwartz 2019). Another is the automation of manufacturing production fueled by the adoption of industrial robots (Acemoglu and Restrepo 2020). Outside of manufacturing, the precipitous decline of coal mining after 1980 has triggered long-lasting and geographically concentrated employment declines (Black, McKinnish, and Sanders 2005; Autor et al. 2021; Hanson 2022; Krause 2022). These episodes highlight the value of understanding the characteristics that make regions resilient to negative shocks, among which may be having larger supplies of foreign-born workers.

Empirical Setting: The Geography of Immigration and Trade

We begin by comparing exposure to import competition from China with the allocation of foreign-born workers across US regions, whereby we use commuting zones as our concept of local labor markets (Tolbert and Sizer 1996; Dorn 2009). Exposure to the China trade shock in the 2000s was greater in regions that previously had attracted relatively few foreign-born workers. Data for employment and population are from the 5% samples of 1990 and 2000 U.S. Census and the combined annual 1% samples of the American Communities Survey for 2006–2008 (which we use for 2007), 2009–2011 (which we use for 2010), 2011–2013 (which we use for 2012), and 2017–2019 (which we use for 2018), sourced from IPUMS USA (Ruggles et al. 2022).[4] Trade data are from UN Comtrade, and industry shipments data are from the NBER Manufacturing Database.

Import Competition from China

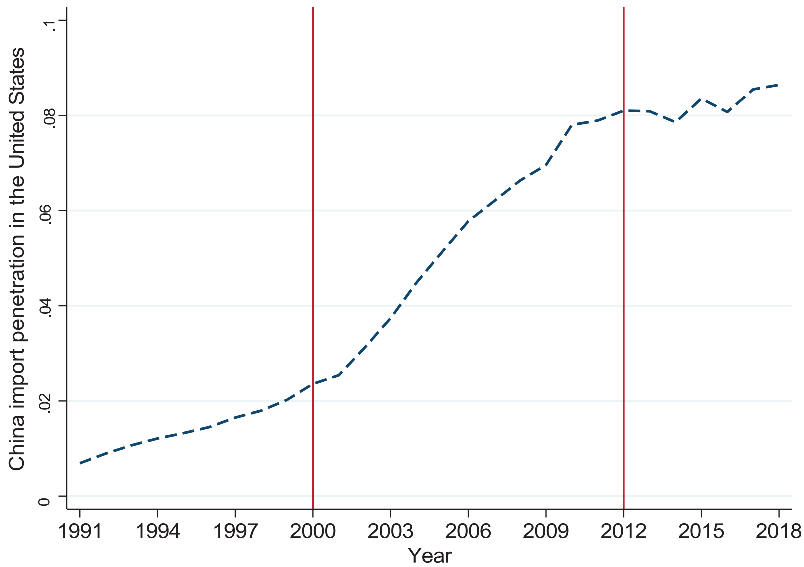
In recent decades, China's manufacturing exports have boomed. Key to this growth were reforms in China (Naughton 2007) that reallocated resources from state-owned enterprises to the private sector (Song, Storesletten, and Zilibotti 2011; Khandelwal, Schott, and Wei 2013; Hsieh and Song 2015), allowed labor to move from farms to cities (Brandt, Tombe, and Zhu 2013), and reduced barriers to foreign trade and investment (Feenstra and Hanson 2005; Yu 2010; Bai, Krishna, and Ma 2017; Brandt and Morrow 2017).

We define the growth of import penetration by China in US industry j and over period τ as,

$$\Delta IP^{cu}_{j\tau} = \frac{\Delta M^{cu}_{j\tau}}{Yjt + Mjt - Xjt} \tag{1}$$

[4]We allocate IPUMS data for Public Use Micro Areas (PUMAs) to commuting zones (CZs) using the crosswalks of Autor and Dorn (2013) and Autor et al. (2019).

Figure 1. US Manufacturing Imports from China



Notes: Import penetration is the ratio of US imports of manufactured goods to US domestic absorption (defined as manufacturing gross output plus imports minus exports). Values exclude oil and gas industries. Data are from UN Comtrade (for imports and exports) and the St. Louis Federal Reserve Bank (for gross output).

where the numerator ($\Delta M_{j\tau}^{cu}$) is the increase of annual US industry imports from China during τ, and the denominator is US industry domestic absorption (industry shipments, Y_{jt}, plus imports, M_{jt}, minus exports, X_{jt}) in a base year t. Autor et al. (2021) highlighted the gradual initiation of China's export boom in the early 1990s, the dramatic acceleration of China's export growth following its accession to the World Trade Organization (WTO) in 2001, and the plateauing of China's export expansion after 2012 (Lardy 2019; Brandt et al. 2020). These phases are evident in Figure 1, which plots the value in Equation (1) averaged across US manufacturing industries. The share of China in US domestic manufacturing absorption rose modestly from 0.7% in 1991 to 2.0% in 2000, then jumped to 8.1% in 2012 during the peak period of the China trade shock, and stabilized over the ensuing decade. We measure changes in industry trade exposure to China over the primary shock period of 2000 to 2012. In extended results, we examine the entire 1992 to 2012 period.

Regional Exposure to Import Competition

We first examine exposure to import competition from China across the 722 commuting zones in the continental United States. As in Acemoglu et al. (2016) and Autor et al. (2021), our measure of trade exposure is the sum of changes in Chinese import penetration across manufacturing industries, by industry shares in initial CZ employment:

$$\Delta IP^{cu}_{i\tau} = 100 \times \sum_j s_{ijt} \Delta IP^{cu}_{j\tau}. \tag{2}$$

Here, $\Delta IP^{cu}_{j\tau}$ is the growth of Chinese import penetration for US industry j over time interval τ (2000 to 2012 in our baseline), t is the initial period (2000 in our baseline), and $s_{ijt} \equiv L_{ijt}/L_{it}$ is the share of industry j in CZ i's total employment (including non-manufacturing) in the initial year. Differences in $\Delta IP^{cu}_{i\tau}$ across CZs stem from variation in local industry employment in the initial year, which arises from differential specialization in manufacturing, and in import-intensive industries specifically. The trade shock in Equation (1) is taken from Autor et al. (2021) for the 2000–2012 time period. The decadalized mean of Equation (2) is 0.89 percentage points, with values of 1.2 percentage points at the 75th percentile and 0.5 percentage points at the 25th percentile (see Online Appendix Table A1; hereafter, numbering for Online Appendix material is prefaced with an "A").[5]

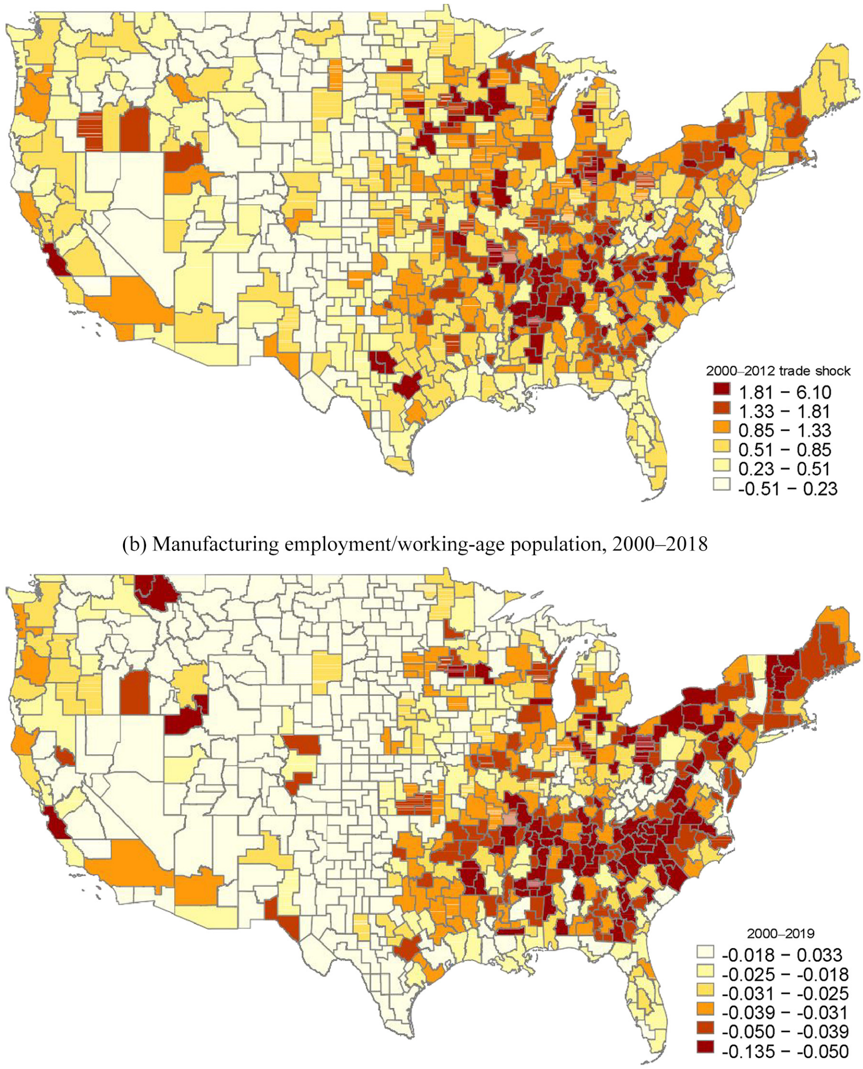
In Figure 2(a), we map the China trade shock in Equation (2) across commuting zones from 2000 to 2012. The most impacted CZs, shown in darker shades as those in the top two deciles of increased import penetration, are concentrated in the eastern half of the United States, and especially in the Southeast and the Midwest. These CZs are where US manufacturing relocated as it moved out of major cities in the Northeast and northern Midwest in the middle of the 20th century (Eriksson, Russ, Shambaugh, and Xu 2019). As US manufacturing matured over the past century, the locus of innovation shifted from industry to services. The rise of advertising, finance, insurance, other business services, and later information technology, pushed manufacturing out of Northern cities and into smaller towns, some of which were located nearby in the Midwestern hinterland and others of which were located in the South and Southeast. Most of this relocation occurred between 1920 and 1980 (Kim and Margo 2004), and therefore was largely complete before large-scale immigration of less-educated workers from Latin America and the Caribbean was in full swing (Hanson, Orrenius, and Zavodny 2022). Although the Latin American immigration wave was triggered by changes in US immigration policy in the 1960s, it did not accelerate until the region's economic crises of the 1980s and 1990s.

Figure 2(b) reports the evolution of manufacturing employment as a share of the working-age population between 2000 and 2018. While the national manufacturing employment rate declined by 2.5 percentage points

[5]This decadalized mean is smaller than the corresponding change in Figure 1 because the expression in Equation (2) takes the value in Equation (1) and weights by industry shares in total CZ employment, including non-manufacturing.

Figure 2. Regional Exposure to Import Competition from China

(a) Import penetration, 2000–2012



(b) Manufacturing employment/working-age population, 2000–2018

Notes: Data are from UN Comtrade (for imports and exports), the NBER-CES Manufacturing Industry Database (for industry shipments), and the 2000 Census and 2017–2019 American Community Survey (ACS) samples (for employment and population).

over this period (see Table A2), the map reveals considerable spatial variation in these changes, with deep employment contractions in parts of the South, Midwest, and Northeast, and modest employment expansions in the Great Plains and some Southern and Western coastal areas. A visual comparison of Figures 2(a) and 2(b) reveals a strong correlation: Many of the CZs that lost manufacturing employment overall (seen in panel (b)) were also more exposed to the China trade shock (seen in panel (a)). This visual evidence is supported by causal analysis of the negative impacts of Chinese import competition on manufacturing employment across US local labor markets (see, e.g., Autor et al. 2013, 2016; Redding 2020).

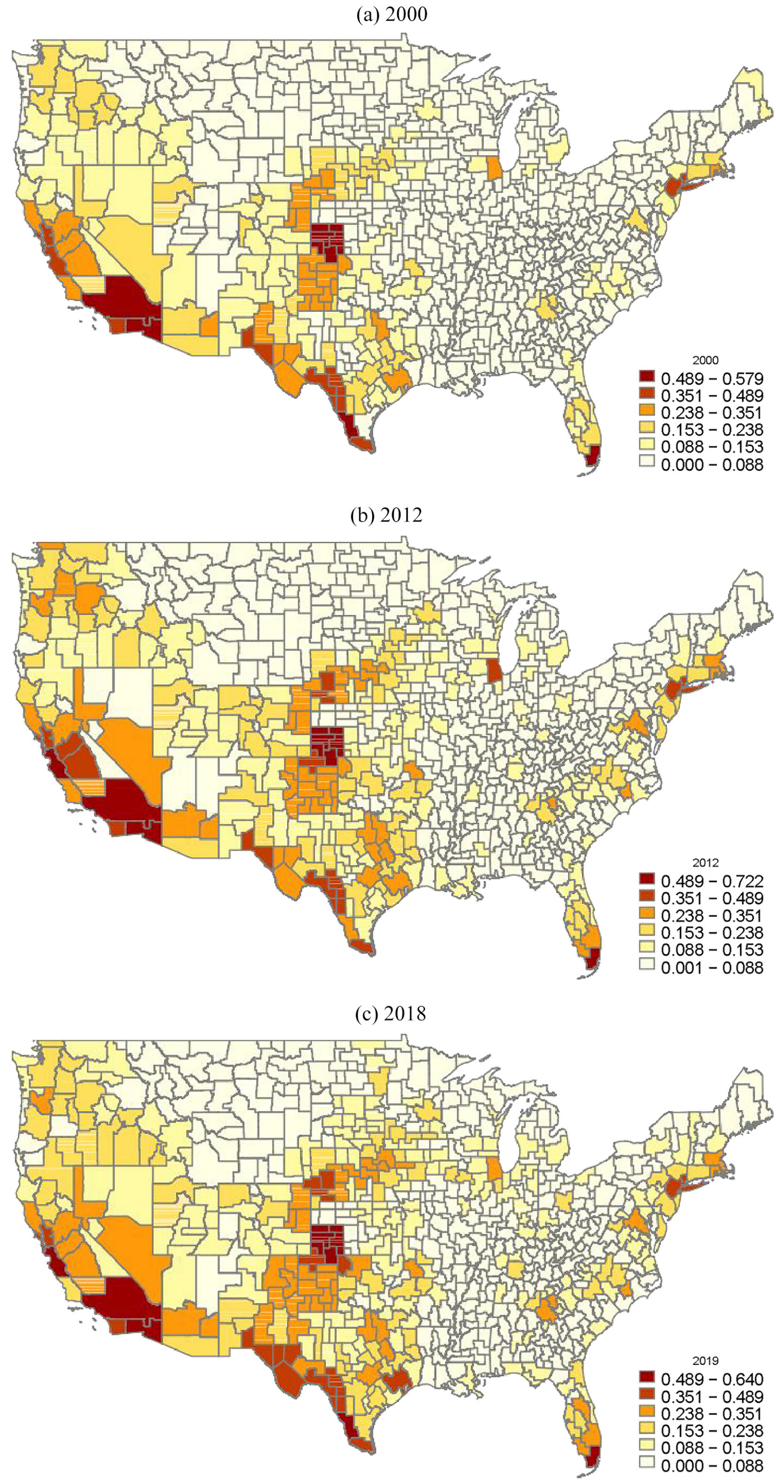
Regional Exposure to Immigration

We next examine the presence of foreign-born workers in manufacturing across US commuting zones, before and after the intensification of the China trade shock in 2001. Figure 3 displays the foreign-born share of total manufacturing employment in 2000, 2012, and 2018, while Figure A1 does so limiting workers to those with a high school education or less. We observe an apparent disconnect between the location of foreign-born manufacturing workers, shown in Figure 3(a), and the geographic dispersion of the China trade shock, shown in Figure 2(a): The regions that were most exposed to Chinese import competition after 2000 had few foreign-born manufacturing workers as of 2000. For CZs at the 75th percentile of exposure to the trade shock, the foreign-born share of the working-age population in 2000 was 8.5%, as compared to 14.1% for CZs at the 25th percentile of trade exposure (see Table A3). This difference potentially limited the role of immigration in easing adjustment to trade-induced manufacturing job loss.[6]

To understand the origins of this disconnect, consider that in 2000 foreign-born manufacturing employment was concentrated on the West Coast, the Southwest, South Florida, and a handful of large cities. These locations were gateway regions for immigration from Latin America and the Caribbean. In 2000, 51.7% of the working-age foreign-born—and 69.7% of the working-age foreign-born with a high school education or less—were from these origin regions (Hanson et al. 2022). Just as previous generations of immigrants had tended to settle in enclaves comprising their country people (Abramitzky and Boustan 2017), so too did arrivals from the Western Hemisphere. Immigrants from Mexico located in states near the US–Mexico border, immigrants from Cuba and elsewhere in the Caribbean concentrated around Miami, and immigrants from South America clustered in the New York City area. As the Latin American immigration wave continued, immigrant clusters emerged in regions with strong job growth for less-educated workers, including farming communities in central California and the inland Northwest; the meatpacking belt of Colorado, Kansas, and Nebraska; and growing larger cities, such as Atlanta, Charlotte, Dallas-Fort Worth, Denver, Houston, and Washington, D.C. (Durand et al. 2001; Champlin and Hake 2006; Card and Lewis 2007). The geographic pattern of foreign-born manufacturing employment in Figure 3(a) mirrors these settlement patterns. Comparing Figure 3(a) to 3(c), we see that immigrant

[6]In Appendix Figure A2, we plot the 2000–2012 China trade shock, mapped in Figure 2(a), against the 2000 foreign-born share of manufacturing employment, mapped in Figures 3(a) and A1. We report zero correlation between the two series.

Figure 3. Share of Foreign-Born Workers in Manufacturing Employment



Notes: Data are from the 2000 Census and the 2011–2013 and 2017–2019 combined one-year American Community Survey (ACS) samples. Employment is of those 18 to 64 years of age.

presence in manufacturing expanded around existing immigrant clusters over the 2000 to 2012 period and then showed little change after 2012, during which time US immigration slowed sharply (Hanson, Liu, and McIntosh 2017).

When the China trade shock began to intensify after the year 2000, immigrant workers were modestly *overrepresented* in manufacturing. In 2000, foreign-born workers accounted for 15.2% of manufacturing employment, as compared to 13.3% of total employment; among workers with a high school education or less, these shares were 18.0% and 16.8%, respectively (see Table A3). Yet, because foreign-born manufacturing workers were concentrated around existing immigrant population centers, the foreign-born were *underrepresented* in the regions exposed to the China trade shock, a fact that foreshadows the empirical results we present later.

Empirical Specification

Our empirical specification builds on Autor et al. (2013), Autor et al. (2021), and much other previous work. We identify the causal impact of import competition from China on population headcounts for the native-born and foreign-born across US commuting zones. Changes in headcounts are indicative of net migration and therefore of labor supply responses to changes in economic conditions.

Baseline Specification

To quantify the impact of the China trade shock on labor supply, we estimate first-difference models for time differences of varying lengths. Our regressions have the form,

$$\Delta Y^{g}_{it+h} = \alpha_t + \beta_{1h}\Delta IP^{cu}_{i\tau} + \mathbf{X}'_{it}\beta_{2h} + \varepsilon_{it+h} \quad (3)$$

where ΔY^{g}_{it+h} is the change in log headcounts for group g in CZ i between the initial year t and later year $t+h$. Our baseline specifications consider outcomes over three time periods: 2000 to 2007, which as seen in Figure 1 is the period of the most rapid increase in import penetration from China, overlapping with the period of analysis in Autor et al. (2013); 2000 to 2012, which spans the period during which the China trade shock reached its full expression; and 2000 to 2018, which extends the time period up to just before the COVID-19 pandemic and the ensuing economic disruptions, and overlaps with the analysis in Autor et al. (2021) (see Table A4). Our definition of the trade shock, $\Delta IP^{cu}_{i\tau}$, is for the period 2000 to 2012, for which the first year is one year prior to China's WTO entry and the final year postdates both the plateauing of the trade shock in 2010 and the volatility in global trade that followed the 2008 to 2010 global financial crisis.

The impact of import competition on CZ population headcounts summarizes the net effect of trade shocks on the pool of both potential workers and non-working residents. Because our interest is in the impact of trade shocks on labor supply, we focus on individuals of working age, defined as those 18 to 64 years old. Native-born and foreign-born workers may differ in their migration responses to labor demand shocks, owing to the potentially stronger attachment of the former to their existing place of residence, which may arise from localized family connections, friend networks, or other bonds and which those born abroad may be less likely to possess. Labor supply responses to labor demand shocks may also differ by worker age and educational attainment. Younger workers and more-educated workers, for instance, appear to be relatively mobile geographically (Bound and Holzer 2000). We therefore examine the responsiveness of population headcounts to greater import exposure separately for workers with a high school education or less and with some college education or more, and for workers ages 18 to 39 and ages 40 to 64.

In Equation (3), the control vector $\mathbf{X}_{it}^{'}$ contains time trends for U.S. Census divisions and start-of-period CZ-level covariates: the manufacturing share of employment, which allows us to focus on variation in trade exposure arising from CZs' differential within-manufacturing industry mix; specialization in occupations according to their routine-task intensity and offshorability (Autor and Dorn 2013), thus accounting for exposure to automation and non-China-specific globalization; the fractions of foreign-born, non-whites, and the college educated in the population, and the fraction of working-age women who are employed, which absorb variation in outcomes related to labor-force composition; and the population shares of residents ages 0 to 17, 18 to 39, and 40 to 64, which control for variation in migration incentives across age groups (see Table A1). We weight regressions by the CZ population in the initial year and cluster standard errors by state.

The analysis is complicated by strong secular trends in population growth across US regions, which began well before the China trade shock (Blanchard and Katz 1992). Greenland et al. (2019) suggested that results on the impact of trade shocks on population headcounts are sensitive to controlling for such trends.[7] Accordingly, we include the log change in CZ population over 1970 to 1990 as a control to absorb historical factors driving population growth.

An additional potential complication is spatial correlation in exposure to import competition (e.g., Adao, Arkolakis, and Esposito 2019). Borusyak, Dix-Carneiro, and Kovak (2022) studied how the analysis of internal migration is affected by regions with stronger bilateral migration links having common labor demand shocks. If, for instance, regions more exposed to the China trade shock were geographically clustered (Eriksson et al. 2019)

[7]Much of the analysis of the China trade shock focuses on outcomes expressed as ratios—for example, the employment–population ratio, earnings per worker, income per capita. Taking ratios effectively differences out secular trends in regional employment or population growth, making impacts of trade exposure on these outcomes immune to the inclusion of controls for lagged population growth (see Autor et al. 2021).

and if bilateral migration flows are decreasing in bilateral distance (Bertoli and Moraga 2013), then in Equation (3) we may need to account for import competition in surrounding commuting zones. In extended results, we expand our baseline regression accordingly.

Causal Identification

US imports may change because of shocks to US product demand and foreign product supply, and the former may be correlated with the disturbance term, ε_{it+h}. To identify the foreign-supply-driven component of US imports from China, we follow Autor et al. (2013) and Acemoglu et al. (2016) in instrumenting US import exposure, $\Delta IP^{cu}_{i\tau}$, using non-US China exposure, $\Delta IP^{co}_{i\tau}$, which we measure as the industry-level growth of Chinese exports to eight other high-income countries:

$$\Delta IP^{co}_{i\tau} = \sum_j s_{ijt-10} \Delta IP^{co}_{j\tau} \tag{4}$$

where $\Delta IP^{co}_{j\tau} = \frac{\Delta M^{co}_{j\tau}}{Y_{jt-3} + M_{jt-3} - X_{jt-3}}$. This expression differs from Equation (2) by using imports by other high-income markets ($\Delta M^{co}_{j\tau}$) in place of US imports ($\Delta M^{cu}_{j\tau}$), the 3-year lag of industry absorption ($Y_{jt-3} + M_{jt-3} - X_{jt-3}$) in place of its year t value, and the 10-year lag of CZ industry employment shares, $s_{ijt-10} \equiv L_{ijt-10}/L_{it-10}$, in place of year t values (see Table A1).

Analyses of the China trade shock have used $\Delta IP^{co}_{i\tau}$ as a shift-share instrument in local labor market regressions (e.g., Autor et al. 2013). Recent literature formalizes the basis for identification and inference in such shift-share settings. Borusyak, Hull, and Jaravel (2022) treated identification as based on exogeneity of the shifts—that is, the industry-level changes in import penetration, and Adao, Kolesár, and Morales (2019) presented a related method for estimating standard errors. Conversely, Goldsmith-Pinkham, Sorkin, and Swift (2020) studied a setting in which industry shifts (import penetration) are taken as given while initial industry employment shares are assumed to be exogenous. Applying the framework in Borusyak, Hull et al. (2022), for the instrument, $\Delta IP^{co}_{i\tau}$, to be orthogonal to the residual, ε_{it+h}, in Equation (4), the following condition must hold: $\mathbb{E}\left[\sum_j s_j \Delta IP^{co}_{j\tau} \bar{\varepsilon}_j\right] = 0$, where s_j is the national employment share of industry j and $\bar{\varepsilon}_j = \sum_i s_{ijt-10} \varepsilon_{it+h} / \sum_i s_{ijt-10}$ is the exposure-weighted average of unobserved shocks for industry j. This orthogonality condition is satisfied if either the large-sample covariance between the industry-level instrument $\Delta IP^{co}_{i\tau}$ and unobserved shocks $\bar{\varepsilon}_j$ is zero (exogeneity of the shifts), or if the employment shares s_{ijt-10} are exogenous and uncorrelated with these shocks (exogeneity of the shares). The substantial industry-level variation in the timing and intensity of the China trade shock documented by Autor et al. (2021) suggests that our approach is more consistent with assuming shift exogeneity than share exogeneity. To check for orthogonality, Borusyak, Hull et al. (2022) recommended regressing current shocks on past outcomes, which are likely correlated with current residuals. Autor et al. (2013), Acemoglu et al. (2016), and Borusyak, Hull et al. (2022) performed such validation exercises for CZs and industries and failed to reject orthogonality in the large majority of instances.

Empirical Analysis

We use Equation (3) to estimate how the 2000–2012 trade shock affected CZs from 2000 to 2018 in terms of population headcounts for the working-age population (in total or by education subgroup), population headcounts broken down by nativity, and population headcounts broken down by age. We extend the analysis by separating CZs by the initial size of their foreign-born population share, and accounting for changes in the attractiveness of alternative domestic migration locations.

Baseline Results

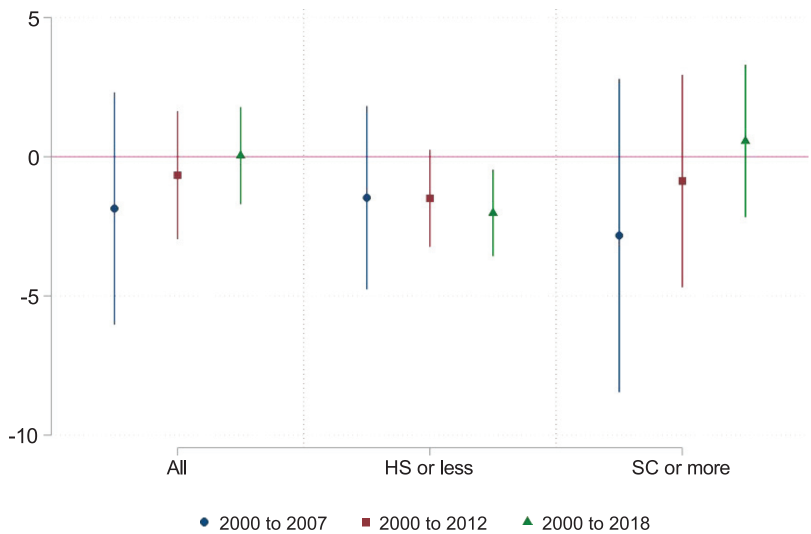
Population Headcounts by Educational Attainment

We begin with the impact of the China trade shock on population headcounts for all those of working age and this population separated by level of education. Figure 4 displays two-stage least-squares (2SLS) point estimates (with vertical bars showing 95% confidence intervals) for the 2000–2012 trade shock, as defined in Equation (1) and instrumented by Equation (2). The top figure presents results for the regression in Equation (3) with a sparse set of controls that include the initial manufacturing employment share and lagged population growth only; the bottom figure presents results with full controls included. Within each figure, we show results for log changes in headcounts for three time periods (2000–2007, 2000–2012, 2000–2018) and three education groups (all, high school and less, some college and more).

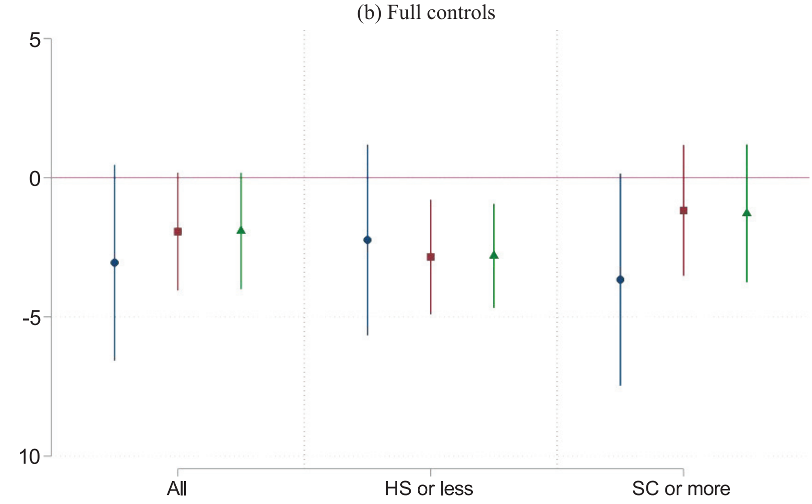
Consider, first, results for the full working-age population, shown in the left-hand panel of Figures 4(a) and 4(b). With either sparse or full controls, the impact of trade exposure on population headcounts is negative but imprecisely estimated, consistent with Autor et al. (2013). Because existing research has shown that CZs exposed to greater import competition from China had larger reductions both in manufacturing employment and in total employment, we might expect a negative impact of greater import competition on local population, as workers migrated out of regions subject to adverse changes in labor demand. Yet, we see weak evidence of such shifts when looking across CZs for all workers. Although precision improves somewhat when we move from regressions with sparse controls in Figure 4(a) to full controls in Figure 4(b), the trade-shock coefficient for the full working-age sample is statistically insignificant in both specifications in each of the three time periods.

Figure 4. Trade Shock Impact on Population Headcounts Ages 18 to 64 by Education, 2000–2018

(a) Sparse controls



(b) Full controls



Notes: Panels (a) and (b) report two-stage least-squares (2SLS) coefficient estimates for β_{1h} in Equation (3) and 95% confidence intervals for these estimates (shown using vertical bars). The dependent variable is the change in the log population over the indicated time period and for the indicated group (all those ages 18 to 64, those with a high school [HS] education or less, those with some college [SC] education or more). The trade shock is the decadalized 2000–2012 change in commuting zone (CZ) import exposure, as defined in Equation (2) and instrumented by Equation (4). Sparse controls (panel (a)) are initial manufacturing employment shares and log population growth over 1970 to 1990; full controls (panel (b)) include initial-period CZ employment composition (shares of employment in manufacturing, routine-task-intensive occupations, and offshorable occupations, as well as the employment share among women), initial-period CZ demographic conditions (shares of the college-educated, the foreign-born, non-whites, and those ages 0–17, 18–39, and 40–64 in the population), Census division dummies, and the change in log population over 1970 to 1990. Regressions are weighted by the CZ working-age population in 2000; standard errors are clustered by state. See Table A5 for tabulated results.

Next, consider results for the working-age population with a high school education and less, shown in the middle portion of Figures 4(a) and 4(b). Because manufacturing is intensive in the employment of less-educated workers, the high-school-and-less group was highly exposed to the China trade shock (Autor et al. 2013). We might therefore expect their net migration responses to be larger than for more-educated workers. Alternatively, previous research has shown that less-educated workers are less geographically mobile in response to adverse labor demand shocks when compared to more-educated workers (see, e.g., Bound and Holzer 2000; Notowidigdo 2020), which could indicate that migration responsiveness to the China trade shock would be weaker for the less-educated. Focusing on results with full controls in Figure 4(b), the impact coefficient is –2.24 (t value = –1.27) for the period 2000 to 2007, which is the end year of analysis in Autor et al. (2013); reaches –2.85 (t value = –2.71) for the 2000 to 2012 period, by which point the China trade shock had reached is maximum intensity; and remains close to this value at –2.81 (t value = –2.93) for the full 2000 to 2018 period. The negative and imprecise results for 2000 to 2007 are consistent with Autor et al. (2013), although the specifications in Figure 4 include lagged population growth as a control whereas the earlier work did not. The negative and statistically significant results for the later time periods are broadly consistent with the analysis in Autor et al. (2021), who examined trade-induced changes in the total population but not in populations broken by education, nativity, and age, as we do here. With sufficient time, workers do begin on net to leave trade-exposed regions, although it takes a full decade for these results to materialize.

To interpret the magnitude of the point estimates, compare CZs at the 25th and 75th percentiles of trade exposure. Over 2000 to 2018, the latter would be predicted to have a decadalized reduction in its high-school-and-less working-age population that is 1.86 (= –2.81 × [1.17 – 0.51]) percentage points larger than the former. This compares to the 25th–75th percentile differential change in log population headcounts of –11.85 (= –9.31 – 2.54) percentage points for the working-age population with no college education over the same period (see Table A4). The observed change in population headcounts for the less-educated over the first two decades of this century dwarfs that predicted by differential exposure to trade shocks, suggesting that any trade-shock-induced net migration was small in the aggregate, an issue we will examine further.

When we turn our attention to the some-college-and-more population, shown in the third panels of Figures 4(a) and 4(b), the impacts of trade exposure on log headcounts are smaller than for the high-school-and-less population and are less precisely estimated. This is initial evidence that the greater trade exposure of the less-educated may have dominated the stronger migration responsiveness of the more-educated, when it comes to population impacts of the China trade shock. These results become clearer when we next disaggregate workers by education *and* nativity.

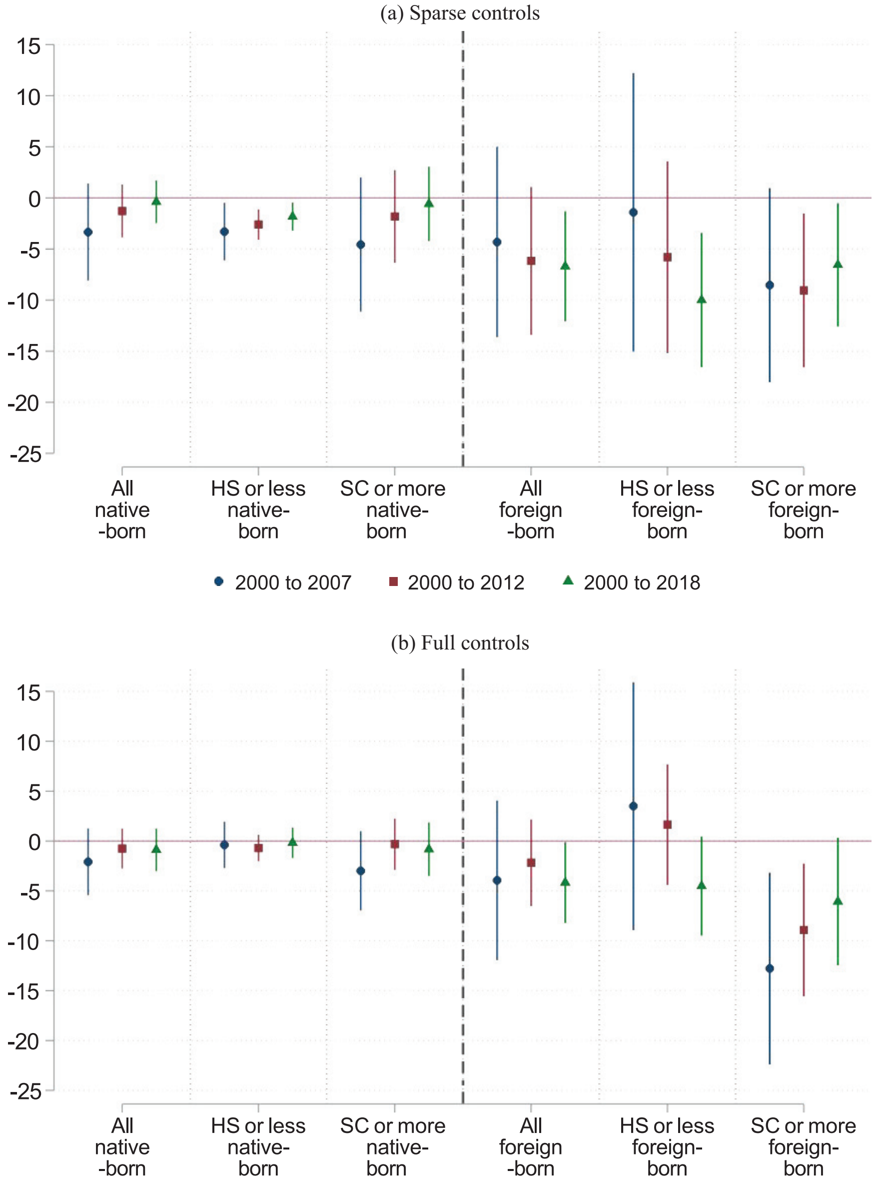
Population Headcounts by Nativity

Figure 5 disaggregates results into effects on foreign- and native-born adults. This figure retains the structure of Figure 4, reporting results in Figure 5(a) with sparse controls and in Figure 5(b) with full controls, with three panels in each figure for three education levels (all workers, high school or less, some college or more) with results for three time periods, 2000–2007, 2000–2012, 2000–2018.

Focusing first on native-born adults in the specification with full controls in Figure 5b, we find small and imprecisely estimated impacts of trade exposure on native-born population headcounts across education groups and time periods. Over the 2000 to 2018 period, impact coefficients are –0.88 (t value = –0.80) for all native-born, –0.18 (t value = –0.23) for native-born with high school or less, and –0.83 (t value = –0.60) for native-born with some college or more. Consistent with previous work, the more-educated have a stronger mobility response to adverse labor demand shocks than do the less-educated, though the difference is not statistically significant. When comparing CZs at the 25th and 75th percentiles of trade exposure, the latter would be predicted to have a decadalized reduction in its native-born working-age population that is just 0.58 (= –0.88 × [1.17 – 0.51]) percentage points larger than the former, which compares to the 25th–75th percentile difference in CZ population changes for this group of –10.31 (= 1.12 – 11.43) percentage points (see Table A4). Overall, we see little impact of shocks to import competition on population headcounts for the native-born. Our finding of weak net migration responses of the native-born is similar to that reported by Cadena and Kovak (2016) for the Great Recession.

The impacts of trade exposure on population headcounts are quantitatively larger and statistically more precise for the foreign-born, as shown in the right-hand trio of panels in Figure 5. Over the 2000 to 2018 period, the impact coefficient is –4.16 (t value = –2.01) for all foreign-born workers. This value is 4.7 (= 4.16/0.88) times that of the corresponding impact coefficient for the native-born. When comparing CZs at the 25th and 75th percentiles of trade exposure, over the 2000 to 2018 period the latter would be predicted to have a decadalized reduction in its foreign-born working-age population that is 2.8 (= –4.16 × [1.17 – 0.51]) percentage points larger than the former. This compares to the 25th–75th percentile differential change in log population headcounts of the working-age foreign-born of –23.3 (= 12.2 – 35.5) percentage points over the same period. Turning to education subgroups, for those with high school or less, the impact coefficient is –4.51 (t value = –1.76), and for those with some college or more, the impact coefficient is –6.06 (t value = –1.86), each of which is marginally statistically significant. For the foreign-born, as for the native-born, the more-educated appear to be more responsive to adverse labor demand shocks than the less-educated (although, as in the earlier results, this difference is not statistically significant).

Figure 5. Trade Shock Impact on Population Headcounts by Nativity, 2000–2018



Notes: Panels (a) and (b) report two-stage least-squares (2SLS) coefficient estimates for β1h in Equation (3) and 95% confidence intervals for these estimates (shown using vertical bars). The dependent variable is the change in the log population over the indicated time period and for the indicated group (all those ages 18 to 64, those with a high school [HS] education or less, those with some college [SC] education or more, either for native-born or foreign-born). The trade shock is the decadalized 2000–2012 change in commuting zone (CZ) import exposure, as defined in Equation (2) and instrumented by Equation (4). Sparse controls (panel (a)) are initial manufacturing employment shares and log population growth over 1970 to 1990; full controls (panel (b)) include initial-period CZ employment composition (shares of employment in manufacturing, routine-task-intensive occupations, and offshorable occupations, as well as the employment share among women), initial-period CZ demographic conditions (shares of the college-educated, the foreign-born, non-whites, and those ages 0–17, 18–39, and 40–64 in the population), Census division dummies, and the change in log population over 1970 to 1990. Regressions are weighted by the CZ working-age population in 2000; standard errors are clustered by state. See Tables A6 and A7 for tabulated results.

Figure A4 further divides the sample by age, reporting regressions for each nativity and education group separately for those ages 18 to 39 and those ages 40 to 64. Previous work suggests that younger workers have stronger migration responses than do older workers, especially among the more-educated (e.g., Greenland et al. 2019). Similar to the results in Figure 5, we see larger responsiveness in population headcounts for the foreign-born when compared to the native-born across all education-by-age subgroups. As expected, these differences are most pronounced among workers 18 to 39 years of age with some college education or more.

Interpreting the Results

Although the foreign-born have stronger net migration responses to trade shocks than the native-born, they were underrepresented in the most trade-exposed US regions and hence played a minor role in spatial labor-market adjustments to trade-induced manufacturing job loss.

To characterize the contribution of the foreign-born to aggregate labor-supply responses to the China trade shock, we again compare CZs at the 25th and 75th percentiles of exposure to imports from China to calculate the implied change in the aggregate labor supplies of these two CZs based on the initial presence of foreign-born workers in each. Given that the foreign-born were 8.5% of the working-age population in 2000 for CZs at the 75th percentile of trade exposure, we can use the impact-coefficient estimate for all foreign-born over 2000 to 2018 in Figure 5(b) to derive a trade-induced decadalized decrease in total working-age population of 0.41 ($= -4.16 \times 1.17 \times 0.085$) percentage points. When we perform a similar calculation for CZs at the 25th percentile of trade exposure, for which the foreign-born were 14.1% of those of working age in 2000, we arrive at a trade-induced decadalized decrease in total potential workers of 0.30 ($= -4.16 \times 0.51 \times 0.141$) percentage points. Because of the initial spatial allocation of foreign-born adults away from traditional manufacturing centers, they contributed only an extra 0.11 percentage-point reduction in potential labor supply in more-trade-exposed relative to less-trade-exposed local labor markets.

To put this quantity in perspective, Autor et al. (2021) estimated that for the 2000 to 2018 period, the impact coefficient for the China trade shock on the log total employment–population ratio was –0.78 (t value = –2.90) percentage points, using an empirical specification very similar to that used here. The implied differential reduction in the log employment–population ratio between more- and less-trade-exposed CZs would have been 0.52 ($= -0.78 \times [1.17 - 0.51]$) percentage points.

More-trade-exposed CZs would have effectively needed to shed an extra half percentage point of the working-age population (while retaining the same number of jobs) to have maintained parity in their employment–population ratios with less-trade-exposed CZs. Of this notional gap, net changes in the foreign-born population would have contributed just 17.5% (0.11/[0.52 + 0.11]) of the needed adjustment, assuming (somewhat optimistically) that the departure of foreign-born adults would reduce population without reducing total jobs.

Extended Results

Separating Commuting Zones by Initial Foreign-Born Population

Literature on the location decisions of international migrants highlights the constructive role that migrant enclaves play in providing access to housing, a sense of community, and job networks for newly arrived co-ethnics (e.g., Borjas 1995; Munshi 2003). Yet, how enclaves relate to the geographic mobility of immigrants is unclear. On the one hand, enclaves offer security and support to immigrants, which may make them wary about leaving gateway locations. This logic underlies the motivation for the Altonji and Card (1991) instrument for immigration-related local labor supply shocks. On the other hand, because recently arrived immigrants may not yet have formed strong personal ties to their new communities, they may be relatively responsive to economic shocks that change the attractiveness of alternative locations. To examine the relative strength of these two forces on immigrant geographic mobility, Figure 6 presents regression results that separate commuting zones according to whether their share of the foreign-born in the local working-age population was above (panel (a)) or below (panel (b)) the national median in 2000.[8]

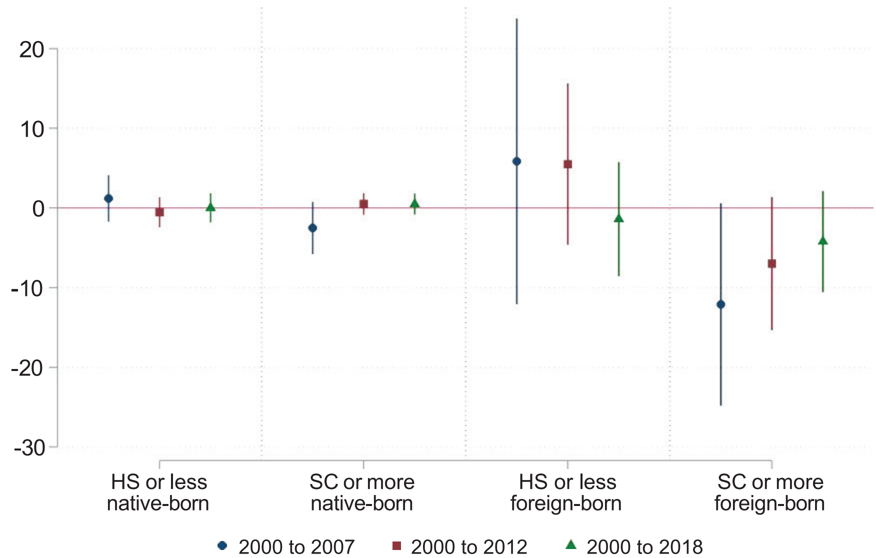
Examining results for the foreign-born, shown in the right two panels of Figure 6, we see much larger trade-induced adjustments in net populations for CZs with larger initial foreign-born populations. For the 2000 to 2018 period, the impact coefficient for foreign-born workers with a high school education and less of –10.71 (t value = –3.04) in Figure 6(b) compares to that of –1.40 (t value = –0.38) in Figure 6(a). Turning to foreign-born workers with some college education and more for the 2000 to 2018 period, the impact coefficient of –8.10 (t value = –1.96) in Figure 6(b) compares to that of –4.21 (t value = –1.30) in Figure 6(a). For the foreign-born, impacts are quantitatively larger and more precisely estimated in CZs with larger initial foreign-born populations. Among the native-born, differences in impact coefficients are much less pronounced across CZs. For the native-born with some college education or more, impact coefficients are near zero in both sets of CZs; for the native-born with high school education or less, coefficients range from precisely estimated zeros in Figure 6(a) to small, positive, and noisily estimated values in Figure 6(b).

On balance, it does not appear that immigrant enclaves are more sticky when it comes to how foreign-born workers respond to negative local labor

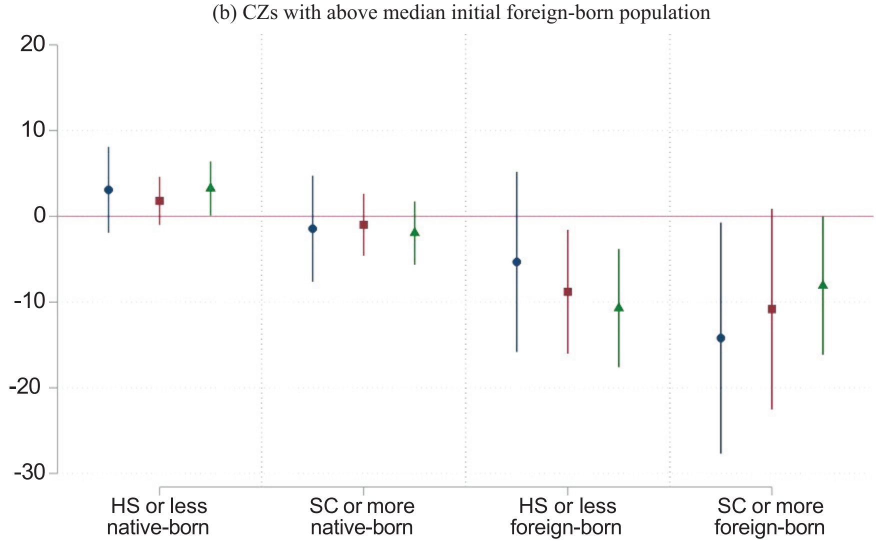
[8]Small sample sizes for immigrants from specific origin countries in many smaller commuting zones prevent us from imposing sample splits based on foreign-born population shares by origin country.

Figure 6. Trade Shock Impacts by Initial Foreign-Born Population, 2000–2018

(a) CZs with below median initial foreign-born population



(b) CZs with above median initial foreign-born population



Notes: Panels (a) and (b) report two-stage least-squares (2SLS) coefficient estimates for β_{1h} in Equation (3) and 95% confidence intervals for these estimates. The dependent variable is the change in the log population over the indicated time period and for the indicated group (those with a high school [HS] education or less, those with some college [SC] education or more, either for the native-born or the foreign-born); the trade shock is the decadalized 1991–2012 change in commuting zone (CZ) import exposure, as defined in Equation (2) and instrumented by Equation (4). In panel (a), the sample is commuting zones with a below median share of the foreign-born in the working-age population in 2000; in panel (b), the sample is commuting zones with an above median share of the foreign-born in the working-age population in 2000. Control variables include initial-period CZ employment composition (shares of employment in manufacturing, routine-task-intensive occupations, and offshorable occupations, as well as the employment share among women), initial-period CZ demographic conditions (shares of the college-educated, the foreign-born, non-whites, and those ages 0–17, 18–39, and 40–64 in the population), Census division dummies, and the change in log population over 1970 to 1990. Regressions are weighted by the CZ working-age population in 2000; standard errors are clustered by state. See Tables A8 and A9 for complete results.

demand shocks. Indeed, the out-migration of immigrants in response to trade exposure appears to be stronger in regions with larger initial concentrations of immigrants. Although the connection between immigrant enclaves and immigrant geographic mobility merits more attention, our results appear to fail to support the rationale underlying the Altonji and Card (1991) instrumentation strategy.

Accounting for the Attractiveness of Alternative Destinations

In recent work, Borusyak, Dix-Carneiro et al. (2022) evaluated the literature on migration responses to local labor demand shocks. Standard spatial economic models (Adao, Arkolakis et al. 2019; Redding 2020) imply that labor supply responses to a localized shock will reflect not just economic conditions in a given location but also those in alternative destinations that residents consider to be in their choice set. Failure to account for exposure to shocks in other regions may lead to biased coefficient estimates in specifications similar to ours. In the context of the China trade shock, we may estimate a low responsiveness of population headcounts to trade exposure, not because the elasticity of migration with respect to local economic conditions is low but because the alternative destinations for residents in exposed local labor markets become unattractive simultaneously (e.g., because they are exposed to similarly negative labor demand shocks).

Inspired by the analysis in Borusyak, Dix-Carneiro et al. (2022), we add the following control variable to Equation (3):

$$\Delta IP^{co}_{-i\tau} = \sum_{k \neq i} \gamma_{ik} \Delta IP^{co}_{k\tau} \tag{5}$$

where $\Delta IP^{co}_{k\tau}$ is the China trade shock facing CZ k and γ_{ik} is the importance of CZ k as a migration location for residents of CZ i.[9] The quantity γ_{ik} should capture the strength of migration flows between CZs i and k. We take two approaches to measuring this value. First, we assume that the attractiveness of other locations is driven entirely by geographic distance, as in gravity models of trade and migration (e.g., Bertoli and Moraga 2013); we also assume that the importance of distance is the same for native-born and foreign-born workers. In this case, γ_{ik} is the bilateral distance between i and k. Second, we focus specifically on the migration propensities of foreign-born workers. Because their mobility appears to be substantially larger than the mobility of the native-born, and because they may evaluate other locations based on the presence of immigrant enclaves in those locations, we alternatively specify γ_{ik} as the Euclidean distance between population shares for all non-US national-origin groups as of 2000, for CZs i

[9]This expression is motivated by equation (17) in Borusyak, Dix-Carneiro et al. (2022). In their general formulation, they differentiated among CZs according to potential sources of migrants to CZ i and potential migrant destinations for residents of CZ i. We implicitly assume that these sets are identical for each CZ.

and *k*. This second approach implicitly assumes that foreign-born workers in a given CZ evaluate other CZs based entirely on the presence of foreign-born residents in those locations.

Estimation results when adding the value in Equation (5) to the specification in Equation (3), and instrumenting for this value using the analogous version of the variable in Equation (4), appear in Figures A5 and A6. When we add the control in Equation (5), we obtain nearly identical impact coefficients on the direct China trade shock, no matter whether we specify bilateral migration connections as depending on geographic distance (Figure A5) or on initial similarity of foreign-born populations (Figure A6). As for the control itself, coefficient estimates are positive but imprecisely estimated in Figure A5, which weights trade exposure in other CZs by geographic distance; coefficient estimates remain positive and become precisely estimated in Figure A6, which weights trade exposure in other CZs by population similarity. These results suggest that adverse trade shocks to likely destination locations reduce the propensity for out-migration from the origin location, which is consistent with the logic of Borusyak, Dix-Carneiro et al. (2022). Overall, accounting for how bilateral regional migration propensities correlate with regional exposure to import competition leaves our core results on how own-CZ trade shocks affect population headcounts unchanged.

Concluding Discussion

The United States has undergone major changes in regional labor demand and supply over the past four decades. The supply of foreign-born workers, and particularly of less-educated migrants from Latin America and the Caribbean, increased sharply after 1980, while adverse labor demand shifts hit regions that had been specialized in traditional manufacturing industries. As it turns out, the first shock appears to have contributed only modestly to adjustments to the second shock. This experience stands in contrast to that for the Great Recession, during which the greater migration elasticity of the foreign-born appeared to play a larger role in regional adjustment to the crash in the US housing market and the severe localized disruptions that ensued.

US local labor markets that were more exposed to the China trade shock had substantially larger net reductions in the population of foreign-born workers but not in the population of native-born workers. The greater sensitivity of foreign-born workers relative to native-born workers to negative labor demand shocks is consistent with the findings in Cadena and Kovak (2016).[10] Despite this differential sensitivity, immigration appears to have had a limited role in aggregate labor-market adjustment to the China trade

[10]Our results also mirror those in Albert and Monras (2022), who found that inflows of less-educated immigrants are highly responsive to positive local labor demand shocks, and Monras (2020), who found that outflows of less-educated immigrants respond quickly to the arrival of new immigrants.

shock. Simply put, at the time of the surge in import competition from China, foreign-born workers were in the wrong locations to contribute much to regional changes in labor supply. The differential change in labor supply associated with the presence of foreign-born workers between labor markets with high versus low trade exposure was effectively zero.

The minor role played by immigration in adjustment to the China trade shock underscores a fundamental albeit straightforward lesson for the spatial equilibration of labor markets. It is insufficient that the foreign-born are relatively willing to move into places with more rapid job growth. For their mobility to buffer adverse shocks, they must initially be present in places subject to negative shocks. The regions exposed to import competition from China in the 1990s and 2000s were specialized in mature manufacturing industries (Eriksson et al. 2019). After 1950, these industries had left larger, more expensive Northern cities for smaller, less expensive locations in the Midwest and Southeast. The relocation of manufacturing was largely complete by 1980, at which point the US immigration wave of less-educated labor from Latin America and the Caribbean was still building momentum. Most of the post-1980 immigrant arrivals from the Western Hemisphere followed earlier cohorts from their home countries by settling in US states on the Mexican border, South Florida, and a handful of large cities. Few were attracted to traditional manufacturing regions. These regions were not growing relative to the nation as a whole and lacked the established immigrant enclaves that tend to attract new arrivals from abroad (Borjas 1995; Munshi 2003).[11]

Although immigration may grease the wheels of the labor market, it may do so inadvertently. This interpretation is consistent with Amior (2024), who found that immigration supplied labor to rapidly expanding US regions after 1960 in part because these places already had immigrant enclaves. Whether arriving immigrants tend to concentrate in regions with strong current job growth or existing communities populated by their country people, immigration may be better suited to ease adjustment to cyclical shocks—in which today's regionalized job growth may be followed by tomorrow's regionalized job loss—than to long-run structural shocks, in which regions with sagging labor demand may be decades past the period they attracted footloose labor.

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[11]Looking ahead, efforts to decarbonize the US economy may divert resources away from fossil-fuel industries, which may trigger a future round of spatially focused job loss (Hanson 2022; Popp, Vona, Marin, and Chen 2022).

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