When Work Disappears: Manufacturing Decline and the Falling Marriage-Market Value of Men*

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Abstract

The structure of marriage and child-rearing in U.S. households has undergone two marked shifts in the last three decades: a steep decline in the prevalence of marriage among young adults, and a sharp rise in the fraction of children born to unmarried mothers or living in single-headed households. A potential contributor to both phenomena is the declining labor-market opportunities faced by males, which make them less valuable as marital partners. We exploit large scale, plausibly exogenous labor-demand shocks stemming from rising international manufacturing competition during the period of 1990 to 2014 to test how shifts in the supply of young 'marriageable' males affect marriage, fertility and children's living circumstances. Trade shocks to manufacturing industries have particularly negative impacts on the labor market prospects of men and degrade their marriage-market value along multiple dimensions: diminishing their relative earnings—particularly at the lower segment of the distribution—reducing their physical availability in trade-impacted labor markets, and increasing their participation in risky and damaging behaviors. We document that adverse shocks to the supply of 'marriageable' men reduce the prevalence of marriage and lower fertility but *raise* the fraction of children born to young and unwed mothers and living in poor single-parent households.

Keywords: Marriage Market, Fertility, Household Structure, Single-Parent Families, Trade Flows, Import Competition, Local Labor Markets

JEL Classifications: F16, J12, J13, J21, J23

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"The consequences of high neighborhood joblessness are more devastating than those of high neighborhood poverty. A neighborhood in which people are poor but employed is different from a neighborhood in which people are poor and jobless. Many of today's problems in the inner-city ghettos—crime, family dissolution, welfare, low levels of social organization, and so on—are fundamentally a consequence of the disappearance of work." William Julius Wilson, *When Work Disappears*, 1996, pp. xiii.

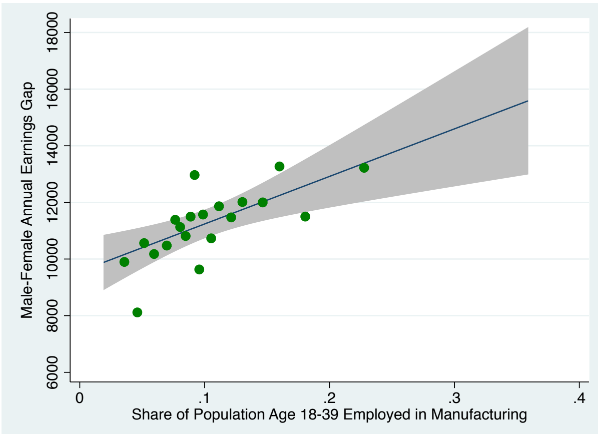
"Wilson's book spoke to me. I wanted to write him a letter and tell him that he had described my home perfectly. That it resonated so personally is odd, however, because he wasn't writing about the hillbilly transplants from Appalachia—he was writing about black people in the inner cities." J.D. Vance, *Hillbilly Elegy: A Memoir of Family and Culture in Crisis*, 2016, p. 144.

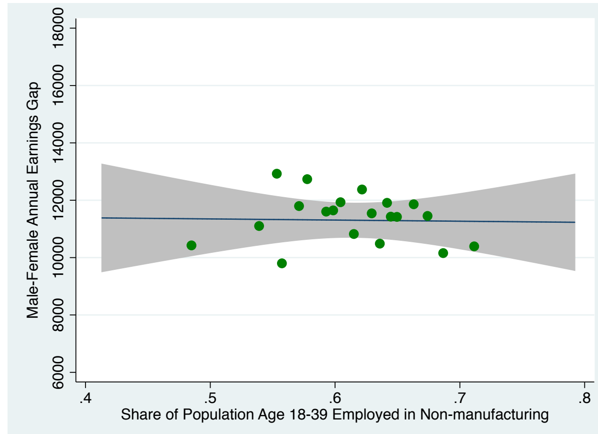
1 Introduction

Marriage and child-rearing in U.S. households has undergone two marked shifts in the last three decades. A first is the steep decline in marriage rates among young adults, particularly for the less-educated. Between 1979 and 2008, the share of U.S. women between the ages of 25 and 39 who were currently married fell by 10 percentage points among the college-educated, by 15 percentage points among those with some college but no degree, and by fully 20 percentage points among women with high-school education or less (Autor and Wasserman, 2013). These declines reflect rising age at first marriage, a decline in lifetime marriage rates and, to a lesser extent, a rise in divorce among less-educated women (Bailey and DiPrete, 2016; Cherlin, 2010; Greenwood, Guner and Vandenbroucke, forthcoming; Heuveline, Timberlake and Furstenberg, 2003). Accompanying the decline in marriage is an increase in the share of children born out of wedlock and living in single-headed households. The fraction of U.S. children born to unmarried mothers more than doubled between 1980 and 2013, rising from 18 to 41 percent (Martin, Hamilton, Osterman, Curtin and Matthews, 2015).

The causes of the decoupling of marriage from child-rearing has drawn decades of research and policy attention. Among the most prominent entries in this debate are William Julius Wilson's pioneering book, *The Truly Disadvantaged* (Wilson, 1987), followed a decade later by *When Work*

Figure 1: Bin-Scatter of the Commuting Zone Level Relationship Between the Manufacturing Employment Share (panel A), the Non-Manufacturing Employment Share (panel B) and the Male-Female Median Annual Earnings Gap: Adults Age 18-39 in 2000





Notes: The regression lines and shaded 95% confidence intervals in each panel are based on bivariate regressions using data from the 2000 Census concorded to 722 commuting zones (CZ) covering the U.S. mainland. Each point of the bin scatter indicates variable averages for subsets of CZs ordered by the x-axis variable that each account for 5% of U.S. population.

Disappears (Wilson, 1996).[1] Although Wilson focuses primarily on outcomes for African-Americans, his work shares a key theme with the larger literature on the rise of single-parent households, which is that the loss of jobs—for men especially—is the root cause of the social anomie found in poor communities. Wilson draws a causal arrow from the secular decline in manufacturing, blue-collar, and non-college employment to the broader social changes occurring in poor neighborhoods.[2]

In this paper, we assess how adverse shocks to the marriage-market value of young adult men, emanating from rising trade pressure on manufacturing employment, affect marriage, fertility, household structure, and children's living circumstances in the United States. While economists and expert commentators have tended to downplay the outsized role assigned to declining manufacturing employment in the U.S. economic debate—what economist Jagdish Bhagwati dubs 'manufacturing fetishism'—simple descriptive statistics support the contention that manufacturing jobs are a fulcrum on which traditional work and family arrangements rest.[3] The lefthand panel of Figure 1

[1]The literature began with the then-controversial report, *The Negro Family: The Case for National Action* (Moynihan, 1965). Elwood and Jencks (2004) and Autor and Wasserman (2013) discuss research on rising single-headship.

[2]Wilson's argument that joblessness is a cause, rather than simply a manifestation, of social decay has precedents in sociology (e.g., Jahoda, Lazarsfeld and Zeisel 1971). But this view has detractors. Focusing on American whites rather than African-Americans, Murray (2012) contends that the expanding social safety net is responsible for the decline in employment and traditional family structures among non-college adults. Putnam (2015) suggests joblessness in poor U.S. communities has cultural and economic causes, which may be self-reinforcing.

[3]See *The Economist* 2011.

illustrates this point with a bin scatter showing the association between manufacturing employment and men's earnings relative to women. Comparing across Commuting Zones (CZs) among young adults ages 18-39 in the year 2000, the male-female annual earnings advantage is substantially larger in CZs where a greater fraction of young adults (both men and women) work in manufacturing. By contrast, the righthand panel reveals that there is no such relationship between *non-manufacturing* employment and the male-female earnings gap. By implication, the male earnings advantage is sharply falling with the share of young adults who are not working (Appendix Figure A1, panel A).[4] Reasoning from the Becker (1973) marriage model and recent variants such as Bertrand, Kamenica and Pan (2015), we would further predict marriage to be less prevalent where the earnings differential between men and women is smaller, as would be the case where fewer adults work in manufacturing.[5] Figure 2 confirms this prediction. In CZs where a larger fraction of young adults (both men and women) are employed in manufacturing, adult women ages 18-39 are substantially more likely to be married (panel A). By contrast, there is a *negative* relationship between the prevalence of marriage and the share of adults working in non-manufacturing (panel B), and this negative relationship also carries over to the share of adults non-employed (Appendix Figure A1, panel B). While these cross-sectional correlations do not admit a causal interpretation, they underscore why manufacturing employment looms large in discussions of traditional gender roles in employment, earnings, and family formation, and they lend credence to the hypothesis that shocks to manufacturing employment may destabilize these roles.

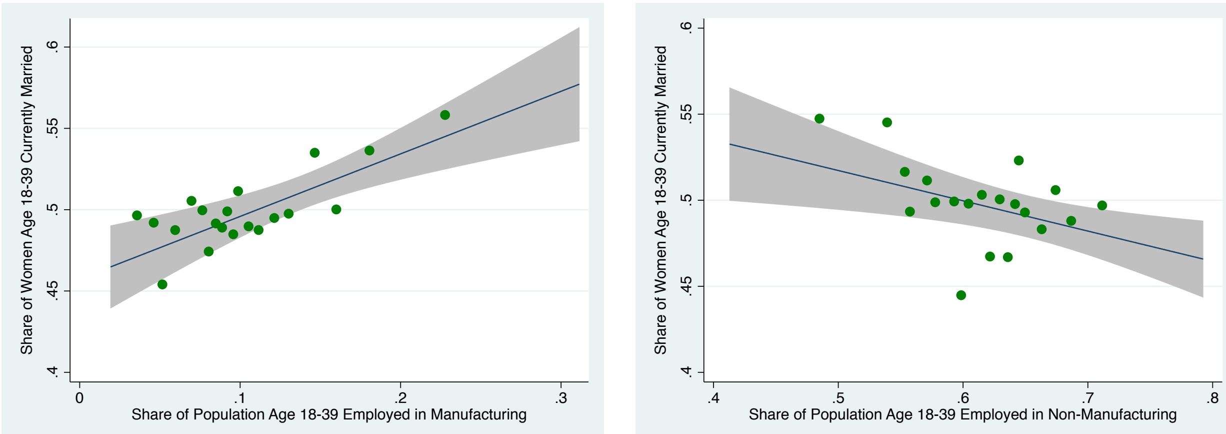
Following Autor, Dorn and Hanson (2013b), we exploit cross-industry and cross-local-labor-market (i.e., commuting zone) variation in import competition stemming from China's rapidly rising productivity and falling barriers to trade to identify market-level labor-demand shocks that are concentrated in the manufacturing sector.[6] In linking local-labor-demand shocks to marriage and fertility outcomes, our work is close in spirit to Black et al. (2003) who document an increasing

[4] Related to these observations, Summers (1986) uses cross-state panels to document that employment growth in the high-wage industries of manufacturing, mining, construction, transportation, and public utilities predicts declines in state unemployment rates while comparable employment growth in low-wage industries is unrelated to unemployment.

[5] Whereas the Becker (1973) marriage model argues that the probability of marriage is increasing in the male-female earnings gap, Bertrand, Kamenica and Pan (2015) additionally posit that men and women care about the earnings ranking within a couple and strongly prefer matches that involve slightly higher male earnings over those that would generate slightly higher female earnings.

[6] Autor, Dorn and Hanson (2013a) and Autor, Dorn and Hanson (2015) show that these trade shocks are an important, but not a unique, reason for local-labor-market declines in U.S. manufacturing employment. Ongoing automation of routine production work is an additional contributing factor.

Figure 2: Bin-Scatter of the Commuting Zone Level Relationship Between the Manufacturing Employment Share (panel A), the Non-Manufacturing Employment Share (panel b) and the Share of Women that Are Currently Married: Adults Age 18-39 in 2000



Notes: See Figure 1.

prevalence of single-headed households in four U.S. states that suffered from a decline of the coal and steel industries, and Kearney and Wilson (2016) who observe rising fertility but no change in marital patterns in U.S. regions that benefitted from the fracking boom during the 2000s. [7] Our study complements the evidence from these episodes of industry-specific booms and busts by assessing whether two decades of contracting U.S. manufacturing employment across a large set of industries and local labor markets has contributed to the rapid, simultaneous decline of traditional household structures. While the decline of manufacturing disproportionately affected males, we exploit gender dissimilarities in industry specialization to separately identify demand shocks that distinctly affect men's and women's employment and earnings.

We apply a large body of harmonized data sources to quantify the link between differential shocks to male and female labor-market opportunities and marriage and fertility outcomes. We first show that shocks to manufacturing labor demand, measured at the commuting-zone level, exert large impacts on men's relative employment and annual wage-and-salary earnings. Although earnings

[7] In other related work, Ananat, Gassman-Pines and Gibson-Davis (2013) find that adverse local economic shocks reduce birthrates and sexual activity among teens—particularly black teens—while increasing the use of contraception and the incidence of abortion. Relatedly, Shenhav (2016) uses gender-specific Bartik shocks and gender differences in occupational choice to predict changes in relative gender earnings in U.S. states, drawing its empirical strategy in part on an earlier version of this paper (Autor et al., 2014a). Shenav's complementary focus is on the economic independence of women rather than the declining marriage-market value of men. Using a strategy similar to Shenhav (2016), Schaller (forthcoming 2016) finds that improvements in men's labor market conditions predict increases in fertility while improvements in women's labor market conditions have the opposite effect. Page et al. (2007)and Lindo et al. (2013) document adverse impacts of parental job loss on children's living circumstances.

losses are visible throughout the earnings distribution, the relative declines in male earnings are largest at the bottom of the distribution. We estimate that a trade shock that increases CZ-level import penetration by one percentage point (a 'unit' shock)—roughly equal to the decadal average trade shock over the 1990s and 2000s—reduces the male-female annual earnings advantage by 3.3 percent at the median and by nearly 9.5 percent at the 25^{th} percentile.[8]

Trade shocks reduce the availability and desirability of potentially marriageable young men along multiple dimensions. The most immediate effect is in populations shifts: a unit rise in Chinese import penetration reduces the ratio of male to female young adults in a CZ by 1.0 percentage points. Where are these men going? Following Case and Deaton (2015) and Pierce and Schott (2016b), we show that trade shocks lead to a differential rise in mortality from drug and alcohol poisoning, liver disease, diabetes, and lung cancer among young men relative to young women. The proportional rise in mortality from these causes is substantial: a one-unit shock raises the relative male death rate from drug and alcohol poisoning by more than 50 percent. But this effect is not nearly large enough to explain the differential decline in the young male population, suggesting that other channels are operative, including migration, homelessness and incarceration. Regarding criminal activity, Deiana (2015), Feler and Senses (2015), and Pierce and Schott (2016b) find significant increases in property- and violent-crime and arrests in trade-exposed CZs during the 1990s and 2000s, which plausibly lead to larger incarceration rates especially for men. The observed rise in the incidence of drug-related deaths and crime in trade-exposed locations suggests that trade shocks contribute to a variety of behaviors that diminish the marriage-market value of males that remain in these locations, including non-lethal substance abuse or illegal activities that do not lead to incarceration.[9]

We next assess marriage-market consequences. Consistent with earlier work (Blau, Kahn and Waldfogel, 2000; Elwood and Jencks, 2004), we find that adverse labor-market shocks reduce the fraction of young women who are currently married. The decline in marriage is not offset by a growth in unmarried cohabitation, as women become less likely to live in couple-headed households regardless of the couple's marital status. More subtly, we find asymmetric marriage-market impacts that depend upon the source of the shock: adverse shocks to labor demand in male-intensive indus-

[8] Autor, Dorn and Hanson (2013a) find that trade shocks reduce CZ-level mean earnings and Chetverikov, Larsen and Palmer (2016) demonstrate that these shocks raise CZ-level earnings inequality, though though they do not study impacts on the gender earnings gap.

[9] Our perspective is akin to Charles and Luoh (2010) and Caucutt, Guner and Rauh (2016), who interpret the rise in male incarceration as an adverse shock to the supply of marriageable men.

tries reduce the prevalence of marriage among young women, whereas analogous shocks to female labor demand significantly *raise* the prevalence of marriage.

Building on these results, we explore outcomes for fertility. Consistent with the general fact that fertility is pro-cyclical, we document that a one-unit import shock lowers births per thousand women of ages 20-39 by 1.9 (a 2% decline). But this decline is not uniform across demographic groups. Fertility among teens and unmarried women falls by proportionately less than fertility among older and married women, so that the share of births to unmarried and—more sizably—teen mothers rises.

Finally, we examine how these changes in children's birth circumstances flow into downstream parental arrangements and child poverty. A one-unit trade shock raises the fraction of children of ages 0-17 living in poverty by 0.6 percentage points (a 3% increase), reduces the fraction living in married households by 0.4 percentage points, and spurs a concomitant rise in the share living in single-parent-headed households. The asymmetric effect of male and female labor demand shocks seen for marriage carries over to household structures. Holding female economic opportunities constant, shocks to male earnings raise the fraction of children living in single-headed households, suggesting that woman are curtailing marriage by more than childbearing. When female earnings fall, however, the share of children in single-parent households *declines* steeply. These shifts in household structure contribute to differential impacts of gender-specific labor demand shocks on child poverty. Adverse shocks to male and female earnings both increase the poverty rate. However, the direct effect of reduced male earnings gets exacerbated as it causes a greater concentration of children in single-parent homes which have an elevated poverty risk; conversely, the direct of effect of lower female earnings is mitigated by the decline in single motherhood. Whereas the male labor-demand shock raises the fraction of children living in poverty, the female labor-demand shock has no significant effect.

Our work contributes to two branches of literature. A first explores how marriage and divorce rates respond to shifts in labor demand or to changes in welfare benefits (Blau, Kahn and Waldfogel, 2000; Elwood and Jencks, 2004).[10] A second, following Wilson and Neckerman (1986) and Wilson (1987), asks whether a shrinking the pool of marriageable low-education men has eroded the incentive

[10]The literature tends to find that better male labor-market opportunities increase marriage rates, whereas better female labor-market opportunities decrease marriage rates. The evidence for a discouragement effect of welfare policies on marriage rates is less certain. Changes in welfare policies are however an unlikely explanation for recent declines in U.S. marriage rates given that the U.S. welfare system has become less generous over the past two decades.

for men to maintain committed relationships, curtailed women's gains from marriage, and strengthened men's bargaining position vis-a-vis casual sex, out-of-wedlock childbirth, and non-custodial parenting (Angrist, 2002; Charles and Luoh, 2010; Edin and Kefalas, 2011; Edin and Nelson, 2013; LeBlanc, 2003; Lundberg, Pollak and Stearns, 2015). Despite a substantial body of evidence, it remains a conceptual and empirical challenge to distinguish cause from effect in the relationship between household structure and labor-market opportunity.[11] Current literature often does not offer tightly identified results delineating whether reductions in the supply of 'marriageable' men are in any meaningful sense responsible for the dramatic changes in marriage and out-of-wedlock fertility observed in the U.S. population. We provide such evidence to the debate.

2 Empirical Approach

2.1 Local labor markets

We approximate local labor markets using the construct of Commuting Zones (CZs) developed by Tolbert and Sizer (1996). Our analysis includes the 722 CZs that cover the entire mainland United States (both metropolitan and rural areas). Commuting zones are particularly suitable for our analysis of local labor markets because they cover both urban and rural areas, and are based primarily on economic geography rather than incidental factors such as minimum population.[12]

2.2 Exposure to international trade

Following Autor, Dorn and Hanson (2013b), we examine changes in exposure to international trade for U.S. CZs associated with the growth in U.S. imports from China. The focus on China is a natural one: rising trade with China is responsible for nearly all of the expansion in U.S. imports from low-income countries since the early 1990s. China's export surge is a consequence of its transition to a market-oriented economy, which has involved rural-to-urban migration of over 250 million workers (Li, Li, Wu and Xiong, 2012), Chinese industries gaining access to long banned foreign technologies,

[11] Bailey and DiPrete (2016) and Greenwood, Guner and Vandenbroucke (forthcoming) review the changing role of U.S. women in the household and the labor market, with the former focusing on educational gender norms and skills development and the latter focusing on technological progress as drivers of these changes. Neither considers the role of the supply of high-quality males in determining women's fertility and marriage decisions.

[12] Parts of our analysis draw on Public Use Microdata from Ruggles, Sobek, Fitch, Goeken, Hall, King and Ronnander (2004) that indicates an individual's place of residence at the level of Public Use Micro Areas (PUMAs). We allocate PUMAs to CZs using the probabilistic algorithm developed in Dorn (2009) and Autor and Dorn (2013).

capital goods, and intermediate inputs (Hsieh and Klenow, 2009), and multinational enterprises being permitted to operate in the country (Naughton, 2007).[13] Compounding the effects of internal reforms on China's trade is the country's accession to the World Trade Organization in 2001, which gives it most-favored nation status among the 157 WTO members (Pierce and Schott, 2016a).

In the empirical analysis, we follow the specification of local trade exposure derived by Autor, Dorn, Hanson and Song (2014b) and Acemoglu, Autor, Dorn, Hanson and Price (2016). Our measure of the local-labor-market shock is the average change in Chinese import penetration in a CZ's industries, weighted by each industry's share in initial CZ employment:

$$\Delta IP^{cu}_{i\tau} = \sum_j \frac{L_{ij90}}{L_{i90}} \Delta IP^{cu}_{j\tau}. \tag{1}$$

In this expression, $\Delta IP^{cu}_{j\tau} = \Delta M^{cu}_{j\tau}/(Y_{j91} + M_{j91} - X_{j91})$ is the growth of Chinese import penetration in the U.S. for industry j over period τ, which in our data include the time intervals 1990 to 2000 and 2000 to 2014. Following Acemoglu, Autor, Dorn, Hanson and Price (2016), it is computed as the growth in U.S. imports from China, $\Delta M^{cu}_{j\tau}$, divided by initial absorption (U.S. industry shipments plus net imports, $Y_{j91} + M_{j91} - X_{j91}$) in the base year 1991, near the start of China's export boom. The fraction L_{ij90}/L_{i90} is the share of industry j in CZ i's total employment, as measured in County Business Patterns data in 1990.

In (1), the difference in ΔIP^{cu}_{it} across commuting zones stems from variation in local industry employment structure in 1990, which arises from differential concentration of employment in manufacturing versus non-manufacturing activities and specialization in import-intensive industries within local manufacturing. Importantly, differences in manufacturing employment shares are not the primary source of variation. In a bivariate regression, the start-of-period manufacturing employment share explains only 35 percent of the variation in ΔIP^{cu}_{it}. In all specifications, we control for the start-of-period manufacturing share within CZs so as to focus on variation in exposure to trade stemming from differences in industry mix within local manufacturing.

The measure $\Delta IP^{cu}_{i\tau}$ captures overall trade exposure experienced by CZs but does not distinguish between employment shocks that differentially affect male and female workers. To add this dimension of variation to $\Delta IP^{cu}_{i\tau}$, we modify (1) to account for the fact that manufacturing industries differ

[13] While China overwhelmingly dominates low-income country exports to the U.S., trade with middle-income nations, such as Mexico, may also matter for U.S. labor-market outcomes. Hakobyan and McLaren (2016) find that NAFTA reduced wage growth for blue-collar workers in exposed industries and locations.

in their male and female employment intensity; hence, trade shocks of a given magnitude will differentially affect male or female employment depending on the set of industries that are exposed. We incorporate this variation by multiplying the CZ-by-industry employment measure in (1) by the initial period female or male share of employment in each industry by CZ (f_{ij90} and $1 - f_{ij90}$), thus apportioning the total CZ-level measure into two additive subcomponents, $\Delta IP^{m,cu}_{i\tau}$ and $\Delta IP^{f,cu}_{i\tau}$:

$$\Delta IP^{m,cu}_{i\tau} = \sum_j \frac{(1 - f_{ij90})\, L_{ij90}}{L_{i90}} \Delta IP^{cu}_{j\tau} \text{ and } \Delta IP^{f,cu}_{i\tau} = \sum_j \frac{f_{ij90} L_{ij90}}{L_{i90}} \Delta IP^{cu}_{j\tau}, \qquad (2)$$

Concretely, consider the hypothetical example of a CZ that houses two import-competing manufacturing industries, leather goods and rubber products, both of which employ the same number of workers and are exposed to industry-specific import shocks equal to 1 percent of initial domestic absorption (thus, $\Delta IP^{cu}_{i\tau} = 1.0$ for this CZ). Imagine that 55 percent of leather goods workers in the CZ are women while 75 percent of rubber products workers in the CZ are men. Equation (2) would apportion these industry by commuting zone trade shocks to males and females according to their local industry employment shares such that $\Delta IP^{m,cu}_{i\tau} = 0.45 \times 1.0 + 0.75 \times 1.0 = 0.6$ and $\Delta IPW^{f}_{uit} = 0.55 \times 1.0 + 0.25 \times 1.0 = 0.4$. In this example, we would assign a larger fraction of a CZ's trade shock to males than to females because males constitute a larger fraction of employment in the CZ's trade-exposed industries. Although the example is hypothetical, the numbers are quite close to the data, as shown in Appendix Table A1. For the period of 1990 - 2000, our data indicate a mean rise of Chinese import penetration of 0.95 percentage points, 59 percent of which accrued to male employment and 40 percent to female employment. In the subsequent 2000 - 2014 period, when Chinese import penetration accelerated, import penetration rose by an additional 1.15 percent per 10 years, with 60 percent of this rise accruing to male employment.

To identify the supply-driven component of Chinese imports, we instrument for growth in Chinese imports to the U.S. using the contemporaneous composition and growth of Chinese imports in eight other developed countries.[14] Specifically, we instrument the measured import-exposure variable ΔIP^{cu}_{it} with a non-U.S. exposure variable ΔIP^{co}_{it} that is constructed using data on industry-level growth of Chinese exports to other high-income markets:

[14] The eight other high-income countries are those that have comparable trade data covering the full sample period: Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland.

$$\Delta IP_{it}^{co} = \sum_{j} \frac{L_{ij80}}{L_{i80}} \Delta IP_{j\tau}^{co}. \tag{3}$$

This expression for non-U.S. exposure to Chinese imports differs from the expression in equation (1) in two respects. In place of computing industry-level import penetration with U.S. imports by industry ($\Delta M_{j\tau}^{cu}$), it uses realized imports from China by other high-income markets ($\Delta M_{j\tau}^{co}$), and it replaces all other variables with lagged values to mitigate any simultaneity bias.[15] As documented by Autor, Dorn and Hanson (2016), all eight comparison countries used for the instrumental variables analysis witnessed import growth from China in at least 343 of the 397 total set of manufacturing industries. Moreover, cross-country, cross-industry patterns of imports are strongly correlated with the U.S., with correlation coefficients ranging from 0.55 (Switzerland) to 0.96 (Australia). That China made comparable gains in penetration by detailed sector across numerous countries in the same time interval suggests that China's falling prices, rising quality, and diminishing trade and tariff costs in these surging sectors are a root cause of its manufacturing export growth.[16]

The exclusion restriction underlying our instrumentation strategy requires that the common component of import growth in the U.S. and in other high income countries derives from factors specific to China, associated with its rapidly evolving productivity and trade costs. Any correlation in product demand shocks across high income countries would represent a threat to our strategy, possibly contaminating both our OLS and IV estimates.[17] To check robustness against correlated demand shocks, Autor, Dorn and Hanson (2013a) develop an alternative estimation strategy based on the gravity model of trade. They regress China exports relative to U.S. exports to a common destination market on fixed effects for each importing country and for each industry. The time difference in residuals from this regression captures the percentage growth in imports from China due to changes in China's productivity and foreign trade costs *vis-a-vis* the U.S. By using China-U.S. relative exports, the gravity approach differences out import demand in the purchasing country,

[15]The start-of-period employment shares L_{ij80}/L_{i80} and the gender shares f_{ij80} are replaced by their 10 year lags, while initial absorption in the expression for industry-level import penetration is replaced by its 3 year lag.

[16]A potential concern about our analysis is that we largely ignore U.S. exports to China, focusing instead on trade flows in the opposite direction. This is because our instrument, by construction, has less predictive power for U.S. exports to China. Nevertheless, to the extent that our instrument is valid, our estimates will correctly identify the direct and indirect effects of increased import competition from China. We note that imports from China are much larger—approximately five times as large—as manufacturing exports from the U.S. to China. To a first approximation, China's economic growth during the 1990s and 2000s generated a substantial shock to the supply of U.S. imports but only a modest change in the demand for U.S. exports.

[17]Note that positive correlation in product demand shocks across high-income economies would make the impact of trade exposure on labor-market outcomes appear smaller than it truly is since these shocks would generate rising imports and rising domestic production simultaneously.

helping to isolate supply and trade-cost driven changes in China's exports. These gravity-based estimation results are quite similar to those from the IV approach that we employ in this paper.[18]

Data on international trade are from the UN Comtrade Database, which gives bilateral imports for six-digit HS products.[19] To concord these data to four-digit SIC industries, we apply the cross-walk in Pierce and Schott (2012), which assigns ten-digit HS products to four-digit SIC industries (at which level each HS product maps into a single SIC industry), and aggregate up to the level of six-digit HS products and four-digit SIC industries (at which level some HS products map into multiple SIC entries). To perform this aggregation, we use data on U.S. import values at the ten-digit HS level, averaged over 1995 to 2005. All dollar amounts are inflated to dollar values in 2015 using the PCE deflator. Data on CZ employment by industry from the County Business Patterns for 1980 and1990 is used to compute employment shares by 4-digit SIC industries in (1) and (3).[20]

3 The Supply of Marriageable Males

We begin by assessing whether trade shocks curtail the supply of marriageable males under age 40, as measured by their employment and absolute and relative earnings, physical availability in trade-impacted labor markets, and participation in risky and damaging behaviors. Across all margins, we find unambiguous evidence that adverse labor-market shocks stemming from trade exposure, whether measured in aggregate or disaggregated by gender, curtail the supply of young men who would likely be judged as good marital prospects.

3.1 Employment effects

The trade shocks that form the basis for our identification strategy are concentrated in manufacturing. We thus set the stage by characterizing the role that manufacturing plays in the employment of young adults. In 1990, 17.4 percent of men and 8.7 percent of women ages 18-39 worked in manufacturing. Focusing only on those currently employed, these shares were 21.1 percent and 12.8 percent respectively—that is, more than one in five young male workers and more than one in eight

[18]See Autor, Dorn and Hanson (2013a) and Autor, Dorn, Hanson and Song (2014b) for further discussion of possible threats to identification using our instrumentation approach, and see Bloom, Draca and Van Reenen (2015) and Pierce and Schott (2016a) for alternative instrumentation strategies for the change in industry import penetration.

[19]See http://comtrade.un.org/db/default.aspx.

[20]Because Census industry categories are somewhat coarser than the SIC codes available in the Country Business Patterns data from which we calculate CZ-by-industry employment, we assign to each SIC industry in a CZ the gender share of the Census industry in the CZ encompassing it when calculating gender-specific employment shocks.

young female workers. These shares fell substantially in the ensuing two decades. By 2014, only 9.6 percent of men and 3.7 percent of women ages 18-39 worked in manufacturing (12.7 and 5.5 percent among those currently employed), corresponding to a fall of 40 percent among young men and more than 55 percent among young women.[21]

While declining manufacturing employment was largely offset by gains in non-manufacturing employment among women, the employment-to-population ratio among young men declined by 7 percentage points among young men. But even absent such an overall employment decline, the sectoral shift away from manufacturing may nonetheless be consequential for marriage and fertility outcomes if manufacturing jobs provide superior hourly earnings or annual hours than non-manufacturing jobs. Descriptive regressions reported in Appendix Table A2 strongly suggest that this is the case. Controlling for an extensive set of covariates, including detailed indicators for age, education, race, nativity, and a complete set of CZ main effects, we estimate that annual earnings of men and women age 18-39 working in manufacturing are approximately 20 to 25 log points higher than annual earnings of demographically comparable adults working in non-manufacturing in the year 2000. Rouugly half of this annual earnings differential is attributable to higher annual hours among man=ufacturing workers, with the remaining half attributable to higher hourly earnings. Although these cross-sectional comparisons may overestimate the causal effect of manufacturing employment on annual earnings despite detailed controls for observable worker characteristics, they are in line with an established literature that documents large industry wage premia in manufacturing (Krueger and Summers, 1988).

We assess the causal effect of trade shocks on employment by fitting models of the form

$$\Delta Y_{sit} = \alpha_t + \beta_1 \Delta IP^{cu}_{it} + \mathbf{X}'_{it}\beta_2 + e_{sit}, \tag{4}$$

where ΔY_{sit} is the decadal change in the manufacturing employment share of the young adult population ages 18 - 39 in commuting zone i among gender group s (males, females, or both) during time interval t, calculated using Census IPUMS samples for 1990 and 2000 (Ruggles, Sobek, Fitch, Goeken, Hall, King and Ronnander, 2004), and pooled American Community Survey samples for 2013 through 2015. Our focus is on employment of young adults because this population is disproportionately engaged in marriage and child-rearing.[22] We estimate (4) separately for the 1990s

[21] These calculations are based on our main Census of populations samples discussed further below.

[22] Our sample is further restricted to individuals who are not residents of institutionalized group quarters such as

and 2000s, and subsequently stack the ten-year equivalent first differences for 1990 to 2000 and 2000 to 2014, while including time dummies for each decade (in α_t). The explanatory variable of interest in this estimate is the change in CZ-level import exposure ΔIP_{it}^{cu} , which in most specifications is instrumented by ΔIP_{it}^{co} as described above. When we turn to gender-specific estimates, we replace ΔIP_{it}^{cu} with $\Delta IP_{it}^{m,cu}$ and $\Delta IP_{it}^{f,cu}$, and use the corresponding gender-specific instruments. The control vector $\mathbf{X}_{it}'$ contains a set of start-of-period CZ-level covariates detailed below.

The first panel of Table 1 presents initial results. As a point of comparison, the first and third column in panel I report OLS estimates of (4) and contain no covariates aside from a constant. Consistent with Autor, Dorn and Hanson (2013a), we find a negative association between rising Chinese import penetration and declining U.S. manufacturing employment in exposed CZs. The highly significant coefficients of -0.65 and -1.29 in columns 1 and 3 indicate that each percentage point rise in import exposure faced by a CZ is associated with a fall of approximately 0.7 points in the share of young adults employed in manufacturing during the 1990s, and a corresponding fall of 1.3 points for the 2000s.

prisons, and who are thus potential participants in the local labor and marriage markets.

Table 1: OLS and 2SLS Estimates of the Relationship between Import Penetration and CZ-Level Manufacturing Employment, 1990-2014 and Pre-Period 1970-1990. Dependent Var: 100 x Change in Share of Population Age 18-39 Employed in Manufacturing (in % pts)

	I. OLS and 2SLS, 1990-2014							
	1990-2000				1990-2014			
	(1)		(2)		(3)		(4)	
Δ Chinese Import Penetration	-0.65	*	-2.12	**	-1.29	**	-1.58	**
	(0.26)		(0.43)		(0.13)		(0.16)	
2SLS First Stage Estimate			0.73	**			0.81	**
			(0.06)				(0.05)	
	II. 2SLS Stacked, 1990-2014							
	Sequential Addition of Control Variables							
	(5)		(6)		(7)		(8)	
Δ Chinese Import Penetration	-1.64	**	-1.05	**	-0.91	**	-1.06	**
	(0.14)		(0.15)		(0.15)		(0.17)	
Census Division Dummies	Yes		Yes		Yes		Yes	
Manufacturing Emp Share$_{-1}$			Yes		Yes		Yes	
Occupational Composition$_{-1}$					Yes		Yes	
Population Composition$_{-1}$							Yes	
2SLS First Stage Estimate	0.83	**	0.68	**	0.65	**	0.64	**
	(0.04)		(0.07)		(0.06)		(0.06)	
	III. Reduced Form OLS, 1970-2014							
	Pre-Periods				Exposure Periods			
	1970-1980		1980-1990		1990-2000		2000-2014	
	(9)		(10)		(11)		(12)	
Δ Predicted Chinese Import Penetration 1990-2014	1.69	**	0.21		-1.09	**	-0.70	**
	(0.36)		(0.33)		(0.30)		(0.10)	

Notes: N=722 in panels I and III, N=1444 (722 commuting zones x 2 time periods) in panel II. All models in panel II comprise a dummy for the 2000-2014 period. Occupational composition controls in columns 7-8 comprise the start-of-period indices of employment in routine occupations and of employment in offshorable occupations as defined in Autor and Dorn (2013). Population controls in column 8 comprise the start-of-period shares of commuting zone population that are Hispanic, black, Asian, other race, foreign born, and college educated, as well as the fraction of women who are employed. The models in panel III regress the outcome on the instrument for growth in Chinese import penetration during the 1990-2014 period and initial Census manufacturing employment shares. Robust standard errors in parentheses are clustered on state. Models are weighted by start of period commuting zone share of national population. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Because observed variation in Chinese import penetration includes both China-based supply shocks—which will tend to reduce competing domestic employment—and domestic demand shocks for specific goods—which will tend to increase both imports and U.S. manufacturing employment simultaneously—we would expect OLS estimates of the relationship between import penetration and domestic employment to be biased towards zero, that is, understating the causal effect of an exogenous increase in import supply on U.S. manufacturing. Columns 2 and 4 of panel I in Table 4, which employ our instrumental variables strategy, confirm this expectation. We find that each percentage-point rise in import penetration causes a decline in U.S. manufacturing employment per working-age population of -2.1 points in the 1990s and -1.6 points in the 2000s. These coefficients are precisely estimated, as are the first-stage coefficients reported at the bottom of each column. The second set of four columns in panel II of Table 1 refine our approach and test robustness. Column 5 performs a stacked first-difference estimate, yielding a point estimate of -1.64. This regression model allows for different time trends across the nine geographic Census Divisions. Columns 6 through 8 cumulatively add further control variables ($\mathbf{X}_{it}$ in equation 4) to account for factors that might independently affect manufacturing employment: the lagged share of CZ employment in manufacturing, absorbing any general shock to manufacturing that leads to a proportional contraction of the sector (column 6); occupational composition controls, accounting for employment in occupations susceptible to automation and offshoring (column 7); and measures of CZ demographics, including race, education, and the fraction of working-age adult women who are employed, which may affect labor supply to manufacturing (column 8).[23] Most of these controls, which we include in all subsequent regression tables, have negligible effects on the magnitude and the precision of the impact estimate. The exception is the lagged manufacturing employment share, which absorbs a secular decline in the size of the sector (possibly including an effect of import competition that is proportional to the size of local manufacturing activity). We estimate in the final column of panel II that a one percentage point rise in import penetration in a CZ causes a -1.06 percentage point change in CZ manufacturing employment as a share of adult population.[24]

[23]Occupational controls in column 7 include, first, the fraction of employment in routine task-intensive occupations, which numerous papers find is a strong predictor of machine-displacement of labor in codifiable clerical, administrative support, production and operative tasks (Autor and Dorn, 2013; Goos, Manning and Salomons, 2014; Michaels, Natraj and Van Reenen, 2014), and second, the mean index of 'offshorability' for occupations in a CZ, where occupations are coded as offshorable if they do not require either direct interpersonal interaction with customers or proximity to a specific work location. Population controls in column 8 comprise the start-of-period shares of CZ population that are Hispanic, black, Asian, other race, foreign born, and college educated, as well as the fraction of women who are employed.

[24]Autor, Dorn and Hanson (2013a) adjust this estimate downward by half to account for the fact that only about 50 percent of the rise in U.S. exposure to Chinese imports during this period can be directly attributed to import supply shocks via the identification strategy described above. We provide rough benchmark numbers here since our objective is to characterize the effect of employment shocks on marriage and household structure, and not to account for aggregate trends in U.S. manufacturing.

How large are these effects? One benchmark is to scale the impact estimate by the interquartile range of rising import exposure across CZs during this time period, equal to 0.66 percentage points per decade (Appendix Table A1). Multiplying the IQR by the column 8 impact estimate of -1.06 implies that rising trade exposure reduced manufacturing employment by 0.7 percentage points more *per decade* in CZs at the 75^{th} percentile of exposure relative to those at the 25^{th} percentile of exposure. As another comparison, the mean per-decade increase in CZ exposure was 1.07 percentage points, implying that the mean CZ lost 1.1 additional percentage points of manufacturing employment per decade relative to a CZ with no exposure. These magnitudes are sizable: only 13.0 percent of young adults age 18-39 were employed in manufacturing in 1990, and this fraction fell by 2.6 percent per decade over the next 24 years (bottom row Table 2).

One possible concern with our empirical analysis is that Chinese import competition may concentrate in local labor markets where manufacturing employment was already differentially declining prior to the 1990s and 2000s, so that the estimates in panels I and II of Table 1 would be counfounded by a pre-trend. This possibility is explored in that table's panel III, which presents OLS reduced form regressions of the decadal change in manufacturing employment on the growth of local labor market exposure to Chinese import competition (averaged over the 1990s and 2000s) not only for the two exposure periods 1990-2000 and 2000-2014, but also for the two decades that preceeded the Chinese export boom, 1970-1980 and 1980-1990. While import competition from 1990 onwards caused a concurrent decline in manufacturing employment (columns 11 and 12), these same import-exposed local labor markets did not experience a differential contraction of manufacturing in the 1980s (column 10), and indeed experienced even a greater expansion of manufacturing during the 1970s. This evidence suggests that the local labor markets with greater exposure to Chinese competition did not fare differentially worse prior to the onset of the trade shock.

The next set of results estimates the consequences of trade shocks on overall employment, unemployment, and non-participation and implements the gender-specific instrumental variables strategy described above. For comparison, column 1 of the upper panel of Table 2 replicates the final estimate from column 8 of Table 1 that includes the full set of covariates that now constitute the baseline specification. The next two columns estimate the impact of trade exposure on manufacturing employment among young men and young women, separately. The point estimates of -0.99 and -1.09 for men and women respectively indicate that the trade shocks seen in this time period had comparable impacts on manufacturing employment rates of both sexes—though the proportional impact for women was larger since roughly twice as large a share of young men as young women was employed in manufacturing at the start of the period (bottom row of panel A in Table 2).

Panel A-II in Table 2 augments these specifications to include male- and female-specific trade exposure measures, each instrumented by contemporaneous changes in China's import penetration to other high income countries during. Despite the relatively high correlation between the gender-specific shock measures ($\rho = 0.80$), there is abundant statistical power for distinguishing their independent effects on labor-market outcomes. The first set of estimates in the lower panel of column 1 indicate that a one-percentage point rise in import penetration of either male or female-dominated industries reduces young adult manufacturing employment by rougly 1 percentage points, as suggested by the by-sex estimates in the upper panel. Columns 2 and 3 demonstrate that the employment effects of sex-specific shocks—constructed by interacting import exposure with gender shares of CZ-by-industry employment—fall almost entirely on their corresponding genders. A one-percentage-point import-penetration shock to male-specific industries reduces employment of young males in manufacturing by 2.6 percentage points ($t = -5.1$) and has a small and statistically insignificant impact on female manufacturing employment.[25] Conversely, a one-percentage-point shock to female-specific industries reduces employment of young women in manufacturing by 2.6 percentage points ($t = -6.7$), while having a modest positive effect on male manufacturing employment.

Panels C and D of 1 detail how these shocks translate to overall employment changes of young men and young women. The overall trade shock causes an employment decline among both sexes, which is offset by greater propensities to be unemployed or non-employed. Among males, the employment decline is entirely due to the sex-specific shock centering on male employment, while female employment falls in response to the female-industry shock (panels C-II and D-II of Table 2). While the qualitative patterns are thus similar across sexes, the quantitative impacts differ. A unit trade shock reduces the male employment-to-population ratio by 1.5 percentage points, while the corresponding number falls by 0.9 percentage points among women.[26] As a result, the gender

[25] We use the terms male-specific and female-specific shocks as a shorthand for the gender-specific trade exposure measure as defined in(2). In reality, industries are not specific to one gender, but the sex composition of manufacturing employment does vary substantially across industries, CZs and industry-CZ pairs.

[26] For women, the overall decline in the employment-to-population is slightly smaller than the decline in manfuacturing employment, while for men, it is somewhat larger. These results are consistent with Autor and Dorn (2013) and Autor, Dorn, Hanson and Song (2014b), who document that adverse shocks to manufacturing employment are offset by sectoral mobility only to a very limited extent, and instead may be exacerbated by additional employment losses outside of the manufacturing sector.

differential in employment rates is reduced by -0.65 percentate points as indicated in panel B of Table 2.

Columns 5 and 6 provide a regression-based decomposition of how shocks to the male-female differential in employment net out across two other domains: unemployment, and non-participation, where each outcome is measured for young adults ages 18 - 39. Due to their common scaling, the net effect of the trade shock must be zero across these three exhaustive and mutually exclusive domains. The column 5 estimate reveals that less than a third of the fall in male relative to female employment accrues to a rise in male relative to female unemployment, while more than two thirds accrue to a rise in male relative to female non-participation (not in the labor force, abbreviated as NILF).[27] Both of the sex-specific components of the trade shock have significant impacts on the gender gap in employment and non-participation. Panel B-II of 2 shows that relative male employment falls by almost than 3.0 percentage points for a unit male shock, and rises by 2.6 percentage point for each unit female shock, while non-participation rates move in the opposite direction.

[27] Aguiar, Bils, Charles and Hurst (2017) document that young men devote an increasing part of their time to video games and recreational computer use, while providing fewer work hours to the labor market.

Table 2: 2SLS Estimates of the Impact of Import Penetration on Employment Status by Gender, 1990-2014. Dependent Var: 100 x Change in Share of Overall/Male/Female Population Age 18-39 Employed in Manufacturing (in % pts); 100 x Change in Share of Male and Female Population and Male-Female Differential in Fraction of Population Age 18-39 that is Employed, Unemployed or Non-Employed (in % pts)

	A. Share Pop Age 18-39 in Manufacturing			B. Male-Female Differential by Employment Status		
	All	Males	Females	Emp	Unemp	NILF
	(1)	(2)	(3)	(4)	(5)	(6)
	I. Overall Trade Shock					
Δ Chinese Import Penetration	-1.06 **	-0.99 **	-1.09 **	-0.65 *	0.19 *	0.46 ~
	(0.17)	(0.17)	(0.20)	(0.26)	(0.09)	(0.24)
	II. Male Industry vs Female Industry Shock					
Δ Chinese Import Penetration × (Male Ind Emp Share)	-1.21 **	-2.59 **	0.20	-2.96 **	0.38	2.58 **
	(0.44)	(0.51)	(0.43)	(0.76)	(0.26)	(0.62)
Δ Chinese Import Penetration × (Female Ind Emp Share)	-0.88 *	0.82 ~	-2.56 **	2.63 **	-0.02	-2.61 **
	(0.35)	(0.46)	(0.38)	(0.58)	(0.26)	(0.55)
Mean Outcome Variable	-2.61	-3.19	-2.06	-2.74	0.03	2.71
Level in 1990	12.98	17.37	8.68	14.64	1.22	-15.87

	C. Employment Status Males Age 18-39			D. Employment Status Females Age 18-39		
	Emp	Unemp	NILF	Emp	Unemp	NILF
	(7)	(8)	(9)	(10)	(11)	(12)
	I. Overall Trade Shock					
Δ Chinese Import Penetration	-1.54 **	0.55 **	0.98 **	-0.88 *	0.36 **	0.53 ~
	(0.29)	(0.15)	(0.20)	(0.35)	(0.11)	(0.31)
	II. Male Industry vs Female Industry Shock					
Δ Chinese Import Penetration × (Male Ind Emp Share)	-3.06 **	1.09 **	1.97 **	0.08	0.71 *	-0.78
	(0.79)	(0.38)	(0.57)	(0.84)	(0.29)	(0.71)
Δ Chinese Import Penetration × (Female Ind Emp Share)	0.19	-0.06	-0.13	-1.97 **	-0.04	2.01 **
	(0.61)	(0.25)	(0.53)	(0.74)	(0.23)	(0.70)
Mean Outcome Variable	-3.00	0.65	2.35	-0.26	0.62	-0.36
Level in 1990	82.33	6.42	11.25	67.69	5.20	27.12

Notes: N=1444 (722 CZ x 2 time periods). All regressions include the full vector of control variables from Table 1. Robust standard errors in parentheses are clustered on state. Models are weighted by start of period commuting zone share of national population. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

3.2 Relative earnings

In this section, we estimate the effect of trade shocks on quantiles of the gender gap in the earnings distribution in local labor markets. Figure 1 in the Introduction provides suggestive evidence for this link by documenting that lower manufacturing employment shares in CZs are correlated with a narrower gender gap in earnings. Whereas the prior tables document that gender-specific trade shocks to manufacturing reduced male relative to female manufacturing employment, our next set of analyses find a similar asymmetry in wage impacts.

For this analysis, we implement the Chetverikov, Larsen and Palmer (2016) approach for performing instrumental-variable estimates of the distributional effects of group-level treatments. Let $y_{it_0}(u)$ equal the unconditional male-female annual earnings gap (in real 2015 US$) in commuting zone i in year t_0 at quantile u among CZ residents ages 25-39.[28] Let $\Delta y_{it}(u)$ equal the change in this gap between time periods t_0 and t_1, corresponding to either $1990-2000$ or $2000-2014$. Our estimating equation takes the form

$$Q_{\Delta y_{it}|\alpha_t, \Delta IP_{it}^{cu}, \mathbf{X}_{it}, \varepsilon_{it}}(u) = \alpha_t(u) + \Delta IP_{it}^{cu}\beta_1(u) + \mathbf{X}_{it}'\beta_2(u) + \varepsilon_{it}(u), \tag{5}$$

where $Q_{\Delta y_{it}|, x_{it}, \varepsilon_{it}}(u)$ is the u^{th} conditional quantile of Δy_{it} given $(\alpha_t, \Delta IP_{it}^{cu}, \mathbf{X}_{it}, \varepsilon_{it})$, α_t is an intercept, ΔIP_{it}^{cu} is the China-Shock measure (instrumented as above), $\mathbf{X}_{it}$ is the vector of observable group-level covariates used in our prior models, $\beta_1(u)$ and $\beta_2(u)$ are conformable coefficient vectors, $\varepsilon_{it} = \{\varepsilon_{it}(u), u \in \mathcal{U}\}$ is a set of unobservable group-level random scalar shifters, and $\mathcal{U}$ is a set of quantile indices of interest. The object of interest for this estimation is $\beta_1(u)$, equal to the causal effect of a trade shock on the conditional quantiles of Δy_j.[29]

The first panel of Table 3 presents estimates of the effect of trade shocks on the CZ-level male-female earnings gap for the 25^{th}, 50^{th}, and 75^{th} percentiles of the distribution. Following our analysis above, we estimate (5) in stacked first differences. Summary statistics for CZ-level wage quantiles in the bottom rows of Table 3 reveal that, within-CZs, male earnings substantially exceed female earnings at all three quantiles considered, with the size of the gap rising steeply with the quantile

[28] Because unemployment and labor-force exit are important margins of response to trade shocks—as our results above demonstrate—the earnings measure includes all CZ residents ages 25-39, including those with zero earnings.

[29] The approach developed by Chetverikov, Larsen and Palmer (2016) provides a two-step procedure for estimating the effects of both person-level (step 1) and group-level (step 2) covariates on the distribution of the outcome variable. In our application, the outcome of interest is the CZ-level distribution of the unconditional male-female earnings gap. Hence, our estimating equation does not include person-level covariates, and our estimation corresponds to step 2 of the Chetverikov, Larsen and Palmer (2016) procedure.

index. In 1990, this gap was \$6,925, \$13,376, and \$17,489 at the 25^{th}, 50^{th}, and 75^{th} quantiles, respectively. The table also shows that these male-favorable gaps shrank at each quantile between 1990-2014, falling by \$1,894, \$2,125 and \$2,491 *per decade* at the 25^{th}, 50^{th}, and 75^{th} quantiles respectively. Given that the absolute size of the contraction was relatively uniform across quantiles, the proportionate reduction in the male-favorable earnings gap was larger at lower quantiles. Specifically, the average male-female earnings gap at the 25^{th} percentile fell by 65% from 1990 to 2014, versus 39% at the median and 35% at the 75^{th} percentile.

In the first row of estimates in Table 3, we find that trade shocks differentially reduced male relative to female earnings. The column 1 estimate indicates that a one-point rise in a CZ's import penetration reduces male relative to female earnings by a sizable \$672 at the 25^{th} percentile ($t = -3.5$) relative to a base of \$6,926 in 1990. A trade shock of comparable size also significantly reduced male relative to female earnings at the median and 75^{th} percentiles by similar dollar amounts (\$445 and \$847, respectively), which however are proportionately much smaller relative to the baseline gaps at these percentiles of the distribution.

To provide a more readily interpretable scaling for these relative earnings effects, columns 4 through 6 of Table 3 express the male-female earnings differential as a percentage of start-of-period male earnings at the indicated percentile. This scaling adjusts for the fact that male earnings exceed those of females at all quantiles and moreover that the gap is larger at higher quantiles; fluctuations in this scaled gap are sizable only if they are large relative to the base earnings of the higher-earning comparison group. As shown at the bottom of the table, the male-female earnings advantage at the 25^{th} percentile fell by 22.6 percentage points per decade relative to initial male earnings between 1990 and 2014. By comparison, the fall in the male-female differential at the median and 75^{th} percentiles was only about one-quarter as large. When we reestimate the impact of trade shocks on male-female earnings rescaled by baseline male earnings in columns 4 through 6 of the table, we find that a one-unit CZ-level import shock reduces the male-female earnings differential at the 25^{th} percentile by 5.3 percentage points. By comparison, impacts on higher quantiles are considerably smaller in magnitude.

Table 3: 2SLS Estimates of the Impact of Import Penetration on Gender Differentials in Annual Earnings, 1990-2014. Dependent Var: Change in the Male-Female Annual Earnings by Percentile of CZ Earnings Distribution (in 2015 US$ or as Percentage of Initial Male Earnings by Percentile), Population Age 18-39

	A. Male-Female Earnings Diff ($)						B. M-F Diff in % of Male Earnings					
	P25		Median		P75		P25		Median		P75	
	(1)		(2)		(3)		(4)		(5)		(6)	
						I. Overall Trade Shock						
Δ Chinese Import Penetration	-672	**	-445	*	-847	*	-5.29	**	-1.01		-1.46	*
	(193)		(191)		(334)		(1.40)		(0.71)		(0.64)	
						II. Male Industry vs Female Industry Shock						
Δ Chinese Import Penetration × (Male Ind Emp Share)	-2,216	**	-2,945	**	-3,685	**	-14.16	**	-9.62	**	-7.26	**
	(516)		(593)		(1081)		(4.46)		(1.77)		(2.03)	
Δ Chinese Import Penetration × (Female Ind Emp Share)	1,086	*	2,400	**	2,384	**	4.81		8.79	**	5.14	**
	(529)		(630)		(814)		(4.88)		(2.01)		(1.57)	
Mean Outcome Variable	-1,894		-2,126		-2,491		-22.59		-7.58		-5.11	
Level of Male Earnings 1990	8,145		26,433		46,533		n/a		n/a		n/a	
Level of Female Earnings 1990	1,104		12,834		28,753							

Notes: N=1444 (722 CZ x 2 time periods). The dependent variables measure the change in the differential between the 25th, 50th and 75th percentile of the male earnings distribution in a CZ and the corresponding percentile of the female earnings distribution. The earnings measure is annual wage and salary income, and earnings distributions include individuals with zero earnings. In columns 1-3, the outcome is measured in 2015 US$, while in columns 4-6, the earnings differential is expressed as a percentage of start-of-period male earnings at the indicated percentile, which in case of the P25 is winsorized at the 5th percentile of the CZ distribution. All regressions include the full set of control variables from Table 1. Regressions are weighted by start-of-period population and standard errors are clustered on state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

The lower panel of Table 3 indicates that trade shocks centered on male employment have a larger effect on the gender earnings gap than shocks to economic opportunities for females: a one-unit rise in import penetration in male-intensive manufacturing reduces male relative to female earnings at the 25^{th} percentile by \$2,216 ($t = -4.3$), whereas a one-unit rise in penetration in female-intensive manufacturing raises male relative to female earnings by a more modest $\$1,086$ ($t = -2.1$). At higher quantiles, however, the relative magnitude of these effects is more comparable (columns 2 and 3), implying that trade shocks differentially reduce the earnings of low-earnings males relative to low-earnings (or non-working) females. The righthand panel of the table replicates this pattern when we scale gaps by initial male earnings: a one-unit import penetration shock to male-intensive employment reduces the male-female 25^{th} percentile earnings gap by 14.2% relative to initial male earnings, whereas a one-unit import shock to female-intensive employment increases the gender earnings gap only by one-third of that amount. At higher quantiles, male- and female-specific shocks have nearly symmetric and smaller impacts on the male-female wage gap expressed in terms of initial male earnings.

This result—that trade shocks differentially reduce the prospects of low-earnings men relative to low-earnings women—will prove central to our subsequent analysis of how these earnings shocks catalyze changes in marriage, fertility, and children's living conditions.

3.3 Satisfaction of gender earnings norms

Our conceptual model predicts that a reduction in the supply of 'high-quality' males will reduce marriage and fertility. One metric of partner quality, suggested by the work of Bertrand, Kamenica and Pan (2015), is whether or not that partner satisfies the so-called gender identity norm in which the male partner in a marriage earns more than the female partner. It is thus of interest to ask whether import shocks reduce the fraction of young males who meet this expectation.

Table 4: 2SLS Estimates of the Impact of Import Penetration on Relative Wages of Potential Spouses (Panel A) and CZ-Level Population Gender Ratio (Panel B), 1990-2014. Dependent Var: 100 x Change in Probability that a Women Earns More than a Randomly Matched Potential Male Partner in the CZ (Panel A) and Male/Female Gender Ratio Among Adults Ages 18-39 (Panel B)

	A. Δ Pr[Woman's Earnings > Earnings of Potential Male Partner], Women Ages 22-43		B. Δ 100 x CZ Male/Female Ratio, Adults Ages 18-39	
	(1)	(2)	(1)	(2)
Δ Chinese Import Penetration	0.23 (0.16)		-1.02 * (0.46)	
Δ Chinese Import Penetration × (Male Ind Emp Share)		1.99 ** (0.44)		-2.52 * (1.09)
Δ Chinese Import Penetration × (Female Ind Emp Share)		-1.76 ** (0.53)		0.69 (0.60)
Mean Outcome Variable	2.11		0.42	
Level in 1990	27.3		98.6	

Notes: N=1444 (722 CZ x 2 time periods). Panel A: For women age 22-41, a potential marriage partner is defined as a man age 24-43 with the same CZ of residence, the same race/ethnicity (non-hispanic white, black, or hispanic), and the same education level (college or non-college). Panel B: Sample comprises all CZ residents ages 18-39 who are not in insitutionalized group quarters. All regressions include the full set of control variables from Table 1 and are weighted by start-of-period population. Standard errors are clustered on state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

For each woman of age 22-41 in a CZ, we calculate the fraction of male potential marriage partners whose earnings fall below hers. Using the patterns of homophily and gender age-gaps employed by Bertrand, Kamenica and Pan (2015), we define potential marriage partners for women of ages 22-41 as men of ages 24-43 in the same CZ of the same race/ethnicity (non-Hispanic white, non-Hispanic black, or Hispanic) and possessing the same broad level of educational attainment (non-college, college).[30] Summary statistics reported in the bottom rows of panel A of Table 4 show that among young women in 1990, 27 percent of the men in their geographic and demographic marriage set (CZ, age, race/ethnicity, education) earned less than they did, and this fraction rose by approximately two percentage points per decade until 2014. Regression estimates for the impact of trade shocks on this outcome measure are consistent with the notion that rising trade exposure reduces the availability of males meeting the gender identity norm, although the estimate is imprecise and not statistically significant. The point estimate of 0.23 ($t = 1.4$) in column 1 of Panel A implies that a one-unit import shock reduces by about a quarter of a percentage point the fraction of demographically suitable young men who meet the gender identity norm for their female counterparts in the same CZ. When we enrich the model in column 2 to include gender-specific trade shocks, we find that a unit shock to male-intensive manufacturing reduces male suitability by 2.0 percentage points ($t = 4.5$) while a unit shock to female-intensive manufacturing raises male suitability by 1.8 percentage points ($t = -3.3$). Thus, the male-specific component of the trade shocks decreases the supply of males meeting the gender-identity norm, while female-specific shocks have an offsetting effect.

3.4 Physical presence, mortality, and risky behaviors

In panel B of Table 4, we consider a more tangible measure of the availability of suitable marital partners: the ratio of young men to young women among non-institutionalized CZ residents age 18-39. Trade shocks may affect this ratio if they induce differential migration, military enlistment, incarceration, homelessness, or mortality by gender. The regression estimates in panel B show that trade shocks reduce the ratio of male to female young adult CZ residents. This ratio averaged 98.6 young men per 100 young women in 1990, and it increased by 0.47 points per decade between 1990 and 2014. The column 1 estimate finds that a unit import-penetration shock reduces the gender ratio by 1.0 points ($t = -2.2$), a unit male-specific shock reduces the gender ratio by 2.5 points

[30]This calculation does not condition on current marital status as the set of suitable marriage partners will include many extant marriages. For the present analysis of gender earnings norms, we follow Bertrand, Kamenica and Pan (2015) in omitting individuals who are neither white, black nor Hispanic.

($t = -2.3$), while a unit female-specific shock has an insignificant positive effect ($\beta = 0.69, t = 1.1$).

What is the proximate source of the trade-induced drop in the gender ratio in trade-impacted CZs? One potential source is military enlistment. The U.S. military is a large employer of young adults, and less-educated young men in particular.[31] We use administrative data from the Office of Economic and Manpower Analysis at West Point (OEMA) to assess the impact of trade shocks on applications to and enlistments in the U.S. Army. The Army is the largest branch of the U.S. Armed Forces, accounting for 40 percent of all active duty enlisted personnel in 2010 (Office of the Deputy Under Secretary of Defense, 2011, Table 2.02). The OEMA data indicates applicants' county of residence, which we concord to commuting zones.[32] Focusing on cumulative applications and enlistments among adults ages 17-40 by sex, age, marital status and education, we show in Appendix Table A3 that a unit shock to CZ-level import penetration spurs an additional 0.18 percentage points of young men to apply for Army service and an additional 0.07 percentage points of young men to enlist. These effects are concentrated among young (17-25), unmarried, and non-college males. We also detect qualitatively similar effects among young women, but the magnitudes are one-quarter to one-third as large: an additional 0.07 percentage points of young women apply, an additional 0.02 percentage points enlist. These estimates confirm that military enlistment contributes to the fall in the gender ratio in trade-impacted locations. But the magnitude of this contribution comprises no more than one-eight of the 1.02 percentage point drop in the virtual gender ratio per unit trade shock reported in Table 4.[33]

A second potential cause of the falling gender ratio is differential mortality. Case and Deaton (2015) document a sharp rise in deaths among non-elderly white non-Hispanics Americans, stemming largely from drug and alcohol poisoning. If the mortality spike was larger among men in trade-impacted locations, it might contribute to the effect of trade exposure on the gender ratio. We assess

[31] In 2010, 86 percent of the 1.18 million enlisted military personnel were men (Office of the Deputy Under Secretary of Defense, 2011, Table 2.02).

[32] These confidential data were used under agreement with the U.S. Army OEMA (we do not have data for other service branches). The regression models in Appendix Table A3 apply an identical vector of start-of-period control variables used in previous tables. The population dominators are drawn from the Census of Populations for the year 2000 and are adjusted for the length of the outcome window (i.e., the number of adults who are ever between the ages of 18-39 during the years 2000 and 2011 exceeds the number in that age range in the year 2000). We additionally control for cumulative applications or enlistments for the years 1998-2000 to account for pre-existing cross-CZ differences in military participation.

[33] Applying the point estimates in Appendix Table A3 and multiplying the effect sizes by 2.5 to account for the fact that the Army accounts for only 40 percent of active duty enlisted personnel implies a reduction in the gender ratio from 98.6 to 98.47 ($98.6 - 100 \times [(98.6 - 2.5 \times 0.07) / (100 - 2.5 \times 0.02)] = 0.125$), which is only about 12 percent of the observed 1.02 point effect.

this possibility using U.S. Vital Statistics mortality files containing person-level death certificates of all U.S. residents, used under agreement with the U.S. Center for Disease Control.[34] The dependent variable for the mortality analysis is the male-female gap in deaths, overall and by cause, per 100K adults ages 20-39. Our analysis here is similar to recent work by Pierce and Schott (2016b), who establish a link between county-level trade exposure and rising mortality due to suicide, accidental poisoning, and alcohol-related liver disease. In addition to exploiting a different source of variation in trade exposure from Pierce and Schott (2016b), there are two differences that distinguish our analysis: first, we analyze mortality outcomes for young adults ages 20-39, rather than studying mortality across all age groups; and second, we analyze CZs rather than counties as the treatment unit for the analysis. Both choices—age groups and geographic units—are guided by our focus on understanding the interplay between labor markets and marriage markets.[35]

Table 5 present results.[36] The first row of estimates finds that shocks to import penetration increase the male-female mortality gap among young adults. The point estimate of 5.36 in column 1 of the upper panel implies that a percentage point rise in import penetration predicts an increase in differential annual male-female mortality of five deaths per 100K adults.[37] This effect is marginally statistically insignificant ($t = 1.8$) but numerically small relative to the baseline mortality rates of 213 and 79 per 100K respectively among men and women in this age bracket, as reported in the bottom of the table. The subsequent six columns decompose this overall mortality estimate into deaths by cause using the classification scheme developed in Case and Deaton (2015, Figure 2). Column 2 shows that trade shocks significantly raise differential male mortality from drug and alcohol poisoning (D&A). The point estimate of 2.6 is large relative to start-of-period D&A mortality rates, which were equal to 6.4 among males and 1.9 among females. As Case and Deaton

[34]Our vital statistics data covers the years up to 2010, and we thus study changes over the pooled periods 1990-2000 and 2000-2010.

[35]As a source of county-level variation in trade exposure, Pierce and Schott (2016b) focus on the rise in Chinese import competition induced by the U.S. Congressional vote in 2000 to provide Permanent Normal Trade Relations to China. This is distinct from our trade exposure measure, which spans a longer time interval and exploits temporal variation in China's revealed comparative advantage in other high income countries as an instrument for Chinese import penetration in the U.S. The mortality measured used in Pierce and Schott is the age-adjusted death rate by county, which incorporates deaths at all age levels. Results by age bracket reported in Figure 6 of Pierce and Schott demonstrate that the county-level mortality response to trade exposure stems almost entirely from increased mortality among adults ages 35-59, most of whom are excluded from our analytic population of young adults.

[36]Our regression models include the same vector of start-of-period control variables used in previous tables, and additionally control for the start-of-period level of the mortality outcome variable in all models to allow for serial correlation in CZ-level mortality rates.

[37]We obtain similar results when using age-adjusted mortality rates, likely because there is only a minimal mortality age gradient among adults under age 40.

(2015) document, D&A mortality rose by epidemic proportions among working-age adults in this time period, and this is also seen in the summary statistics in the bottom of the table. Columns 3 through 6 document corresponding trade-induced increases in differential male mortality from liver disease (which is often alcohol-related), diabetes, lung cancer, and suicide. We do not find an impact of trade shocks on deaths from these causes among young adults. The final column of the table combines all other causes of death beyond those emphasized by Case and Deaton (2015). The trade shock insignificantly increases the gender gap in mortality due to such other causes, which including infectuous diseases, neoplasms and accidents, and account for more than 4 out of every five deaths among young adults.

The lower panel of Table A4 re-estimates the same models using gender-specific trade shocks. As anticipated, male-specific employment shocks raise the male-female differential mortality rate while female-specific employment shocks reduce this differential. Nevertheless, male-specific shocks generally have larger effects, which explains why the pooled trade shock in the upper panel differentially raises overall male deaths and deaths from D&A poisoning. Appendix Table A4, which presents analogous estimates performed separately by sex, documents that the gender gap in D&A deaths widens because the trade shock raises male D&A mortality while having no discernible impact on female mortality.

Table 5: 2SLS Estimates of the Impact of Import Penetration on CZ-Level Gender Differences in Death Rates, 1990-2010. Dependent Var: Difference between Male and Female Deaths per 100k Adults Ages 20-39 by Cause of Death

	Male-Female Death Rate Differential										
	Total		Drug/ Alc Poison		Liver Disease	Diabetes	Lung Cancer	Suicide	All Other		
	(1)		(2)		(3)	(4)	(5)	(6)	(7)		
	I. Overall Trade Shock										
Δ Chinese Import Penetration	5.36	~	2.55	**	0.06	-0.11	-0.22	-0.07	3.05		-
	(3.06)		(0.92)		(0.23)	(0.25)	(0.18)	(0.90)	(2.59)		
	II. Male Industry vs Female Industry Shock										
Δ Chinese Import Penetration	29.62	*	8.27	*	-0.60	-0.69	-0.96	1.45	20.68	~	
× (Male Ind Emp Share)	(12.01)		(3.36)		(0.77)	(0.83)	(0.75)	(2.28)	(10.73)		
Δ Chinese Import Penetration	-19.91	*	-3.43		0.75	0.50	0.56	-1.64	-15.29	*	
× (Female Ind Emp Share)	(8.37)		(2.32)		(0.67)	(0.68)	(0.50)	(1.94)	(6.90)		
Mean Outcome Variable	-27.97		4.02		-0.69	-0.05	-0.23	-1.10	-29.92		
Male Death Rate in 1990	213.43		6.39		4.11	1.95	1.55	25.12	174.31		
Female Death Rate in 1990	78.89		1.92		1.91	1.44	1.00	5.57	67.05		

Notes: N=1444 (722 CZ x 2 time periods). All regressions include the full set of control variables from Table 1 and the start-of-period value of the outcome variable. Regressions are weighted by start-of-period population and standard errors are clustered on state. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

These estimates yield two conclusions. First, male deaths from drug and alcohol abuse are differentially affected by trade shocks. Second, the quantitative magnitude of these effects is too small to explain a noteworthy part of the trade-induced fall in the male/female population ratio documented in the prior table.[38] However, only a small minority of adults who engage in adverse health behaviors experience fatal consequences such as drug overdose; most others survive the most extreme consequences but are likely to be less attractive as marital partners as a result. Thus, the rise in fatal overdoses in trade-exposed locations may imply a diminution in the marriage-market value of a far larger set of males.

This leaves three other likely channels through which trade shocks reduce the ratio of young males to young females in CZs: migration, incarceration and homelessness. Since the Census and ACS data used in our analysis do not provide an individual-level panel, they do not allow a quantification

[38] A one-percentage-point increase in import penetration reduces male population by 5.4 for every $100K$ women. Cumulating over 10 years, this implies about 54 missing men per $100K$. Table 4 finds that a one-percentage-point increase in import penetration reduces the overall male population by 1.02 per 100 women ($1,020$ per $100K$), which is 20 times as large as the mortality effect.

of the gender-specific flows of individuals between commuting zones, into or out of prison, and into or out of the coverage of the Census data (some of the homeless population may not be covered in the data). Using confidential data from the Social Security Administration, Autor, Dorn, Hanson and Song (2014b) found little evidence for large migration responses to local trade shocks, but their analysis did not investigate gender-specific mobility.[39] Homelessness and imprisonment are both plausible channels that could remove males from the pool of available spouses that is observed in the Census data. At the end of our sample period in 2007, men accounted for an estimated 64% of homeless adults, and for a whopping 93% of all prisoners. The second channel, incarceration, is challenging to measure with available data, but indirect evidence suggests that it is likely to be an important contributor.[40] Deiana (2015), Feler and Senses (2015), and Pierce and Schott (2016b) all document statistically significant increases in property-crime and violent-crime incidents and arrests in trade-exposed CZs during the 1990s and 2000s. At the end of our sample period in 2007, 93% of U.S. federal and state prisoners were male and 64% of these males were between the ages of 20 and 39, and younger still at the start of their sentences (West and Sabol, 2008, Table 1 and Appendix Table 7). It is thus plausible that rising crime and arrest rates in trade-impacted local labor markets ultimately yield higher incarceration rates and declining (local) population shares of young males.[41] The third channel, homelessness, again disproportionately affects males who account for 64% of homeless adults according to survey evidence from 2007 (U.S. Conference of Mayors, 2007, Exhibits 2.3 and 2.4). While substance abuse is responsible for the increased incidence in drug-related male deaths that we report in Table 5, it is also also one of the most frequently reported causes for homelessness.[42] We interpret our results on drug and alcohol poisonings as one manifestation of a broader set of risky behaviors and adverse outcomes. This observation again suggests that, beyond removing potentially marriageable men from the labor market, trade shocks diminish the set of marriageable young males among those *remaining* in the trade-impacted labor market.

[39] For an analysis of labor-demand shocks on migration between versus within CZs, see Monte, Redding and Rossi-Hansberg (2015).

[40] Our analysis of Census and ACS data omits residents of institutional group quarters, including prisoners. The Census enumerates inmates at the locations where they are imprisoned but not at their home residences, and consequently, we cannot observe population flows from trade-impacted CZs to correctional facilities.

[41] Combined federal and state prisoner population statistics are from Carson (2014), Tables 2 and 3.

[42] The U.S. Conference of Mayors (2007, Exhibit 2.2) asked city officials to select the three most common causes for homelessness among single persons out of a list of twelve possible causes. Substance abuse (61%) was the second most frequent answer after after mental illness (65%), and was cited more often than direct economic causes of homelessness such as lack of affordable housing (43%) or poverty (39%).

4 The Labor Market and the Marriage Market

The constellation of evidence so far demonstrates that trade shocks shocks differentially reduce males' marriage-market value along multiple margins: relative employment, relative earnings, physical presence in trade-impacted labor markets, and possibly participation in risky, abusive, and illegal activities. We next test how trade shocks have shifted patterns of marriage, fertility and household structures.

4.1 The prevalence of marriage

Both theory and the qualitative evidence discussed in the Introduction suggest that a fall in the marriage-market stature of young men relative to young women in trade-impacted locations will reduce marriage formation and perhaps spur divorce. Conversely, a reduction in the labor-market opportunities available to women will increase marriage. While our repeated cross-section data do not allow us to observe flows into and out of marital statuses, we can assess the impact of trade shocks on the contemporaneous marital status of young adults.[43] Table 6 presents results.

Panel A of the table estimates whether trade shocks affect the fraction of young women ages 18 through 39 who are in each of three marital states: never married, currently married, or, alternatively, widowed, divorced or separated.[44] The first row finds that the aggregate trade shock (i.e., not distinguishing between gender components) reduce the prevalence of marriage: a one-unit trade shock predicts a decline in the the fraction of young women who are currently married by about 1.0 percentage points on a base of 53% (column 3). As predicted by Becker's marriage market model, shocks to male and female-intensive employment have opposite signed effects on marital outcomes (lower panel): adverse shocks to male-intensive employment increase the fraction of young women who were never married and reduce the fraction currently married or formerly married (widowed, divorced or separated); adverse shocks to female-intensive employment reduce the fraction of young women who have never married while raising the fraction who are divorced, widowed or separated.

[43] The collection of flow data on marriages and divorces by the CDC's National Vital Statistics System was suspended in 1996.

[44] These categories are exhaustive and mutually exclusive. If a woman is currently married, we cannot determine if she was previously widowed, divorced or separated.

Table 6: 2SLS Estimates of the Impact of Import Penetration on Marital Status of Young Women, 1990-2014. Dependent Var: Change in Percentage of Women that Are Currently Married, Widowed/Divorced/Separated, or Never Married; Change in Percentage of Mothers that Are Currently Married

	A. Women's Marital Status			B. Women's Household Structure			C. Pct of Mothers Currently Married
	Married	Widowed Divorced Separated	Never Married	Living w/ Spouse	Living w/ Partner	Other Househ Structure	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				I. Overall Trade Shock			
Δ Chinese Penetration	-0.95 **	-0.21 *	1.16 **	-0.81 **	-0.22 ~	1.03 **	-0.52 ~
	(0.30)	(0.11)	(0.33)	(0.27)	(0.12)	(0.30)	(0.31)
				II. Male Industry vs Female Industry Shock			
Δ Chinese Penetration × (Male Share)	-3.57 **	-0.66 **	4.23 **	-3.21 **	0.04	3.17 **	-3.28 **
	(0.62)	(0.22)	(0.64)	(0.55)	(0.28)	(0.60)	(0.73)
Δ Chinese Penetration × (Female Share)	2.03 **	0.29	-2.32 **	1.93 **	-0.52 **	-1.41 **	2.62 **
	(0.55)	(0.19)	(0.58)	(0.54)	(0.20)	(0.52)	(0.85)
Mean Outcome Var	-6.92	-1.62	8.55	-7.57	1.65	5.93	-6.56
Level in 1990	53.05	12.11	34.84	50.30	5.25	44.45	76.02

Notes: N=1444 (722 CZ x 2 time periods). Column 4 refers to households in which the women is the household head and has a spouse who is living in the household, or where the women is the spouse of the household head. Column 5 refers to households in which the women is the household head and has cohabiting partner in the household, or where the women is the cohabiting partner of the household head. Column 6 comprises all other household structures. All regressions include the full set of control variables from Table 1. Regressions are weighted by start-of-period CZ population and standard errors are clusterd by state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Panel B of Table 6 shifts the focus to women's living arrangements. This allows us to explore whether the decline in marriage is to a substantial part compensated by an increase in cohabitation of unmarried couples. We distinguish between three different living arrangements: (i) women living with their spouse, with one of the spouses being head of the household, (ii) women living with an unmarried partner, with one of the partners being head of the household, and (iii) all other living arragements, such as one-person households or households that are headed by a parent, other relative, or friend of a woman.[45] The descriptive statistics at the bottom of the table indicate a rapid change in young women's household structures over the sample period. The fraction of young women living with a spouse declined by a third from 50.3 percent in 1990 to 32.1 percent in 2014, a decline of 7.6 percentage points. The share of women living with an unmarried partner

[45] The Census and ACS data defines as household head the person who owns or rents the dwelling. The data indicates every household member's relationship to the household head, but not the relationship between the non-head household members.

instead nearly doubled, from 5.3 percent in 1990 to 9.2 percent in 2014. However, the regression estimates indicate that the trade shock did not contribute to the growth of cohabiting: a one-unit shock reduced the fraction of young women in married couple households by 0.8 percentage points, and caused an additional decline of 0.2 percentage points in unmarried couple households. The lower two rows of the panel B, which disaggregate the trade shock into gender-intensity components, again show strikingly countervailing effects. A one-unit shock to male-intensive employment raised by 3.2 percentage points the fraction of young women in married couple households while raising the fraction of women outside of couple households by the same amount. Conversely, a one-unit shock to female-intensive employment significantly reduced the fraction of women in cohabiting and non-couple households, while raising the fraction in married couple households by 1.9 percentage points.

The final column of Table 6 assesses the change in marital status for the subset of women age 18 to 39 who live with own children in the household. We find a modest negative impact of the overall trade shock on the probability that young mothers are currently married, and again a notable contrast between the male component of the trade shock, which reduces marriage among mothers, and the female component of the shock, which increases marriage. The lack of flow data on marriage dissolution does not allow us to determine whether the trade shock increases the risk of divorce among young mothers. However, in the following section on fertility and birth circumstances, we investigate the change in the percentage of mothers that are unwed at the time their children are born.

4.2 Fertility and birth circumstances

The U.S. birthrate has not fallen by nearly as much over the last three decades as has marriage, meaning that a rising fraction of births occurs outside of marriage. Both the decline in marriage rates and the rise in non-marital births have been concentrated among low-education adults, a demographic group that has seen its wages and employment fall simultaneously, and hence a natural hypothesis is that declining labor-market opportunities contribute to both falling marriage rates and rising out-of-wedlock fertility (Autor and Wasserman, 2013). Our evidence above is qualitatively consistent with the first tenet of this hypothesis: adverse shocks to labor demand reduce marriage. We now turn to the second: fertility and non-marital births. For this analysis, we use birth-certificate data from the U.S. Vital Statistics to study the impact of trade-induced employment shocks overall and by gender on overall birth rates as well as births to teens and unmarried mothers. We again use a variant of the stacked first-difference estimating equation in (4), where here our dependent variables are decadal changes in Vital Statistics birth measures for the periods 1990 - 2000 and 2000 - 2010.

The first column of Table 7 considers the impact of trade shocks on births per thousand adult women ages 15-39.[46] The estimate in the upper panel finds that a one-unit trade shock reduces total fertility by approximately 2.3 births per thousand women per year, which is roughly a 3-percent decline relative to the 1990 baseline level reported in the bottom row of the table. When we disaggregate labor-demand shocks into gender-specific components, we detect, consistent with earlier results, countervailing effects of male and female labor-market conditions on outcomes. A one-unit shock concentrated on male employment is estimated to reduce births by 7.9 per thousand women (t = 4.2) while an equivalent shock concentrated on female employment is found to raise births by 3.5 per thousand women ($t = 3.0$) .

The following three columns of the table investigate changes in the composition of births. A one-unit trade shock moderately increases the fraction of teen births among all birth to mothers age 15-39 by 0.36 percentage points ($t = 2.0$), corresponding approximately to a 3 percent increase relative to the 1990 baseline level. The share of birth to unmarried mothers, which include many of the teen moms, also increased in response to the economic shock, but this effect is measured with less statistical precision (column 3).

Recent literature hypothesizes that the high rate of teen fertility in the United States relative to other rich countries is partly attributable to the dearth of opportunities facing young non-college women (Kearney and Levine, 2012). The evidence in the lower panel of Table 7 indicates that a trade shock centered on female employment opportunities may indeed rise fertility more broadly, while having a modest and imprecisely estimated effect on the share of teen moms.

The final column of Table 7 investigates an additional proxy of mother's family circumstances at the time of birth. We measure whether the Vital Statistics records of a birth contain either the father's age or race. In 1990, 14 percent of birth records lack such basic information on the father. This absence of father information occurs almost exclusively in case of out-of-wedlock births, and

[46]The count of live births is calculated using Vital Statistics. County population denominators by age group are drawn from Census Survey Tabulation files. All variables are aggregated to the Commuting Zone level.

may correlate with a lack of paternity acknowledgement. The descriptive statistics at the bottom of the table indicate that the fraction of birth certificates without father information declined slightly over the sample period despite a growing fraction of out-of-wedlock births, which may be related to states' increased efforts of obtaining paternity acknowledgement (Almond and Rossin-Slater, 2013). Such changes in recording practices could confound our analysis of this outcome if they were correlated with the spatial distribution of the trade shock. Having this caveat in mind, our estimates for this outcome yield the same sign pattern as for as for the share of birth to teen and unmarried mothers, thus reinforcing the overall impression that the fraction of births into more challenging family circumstances increased in trade-exposed locations.

Table 7: 2SLS Estimates of the Impact of Import Penetration on Births to Young Women, 1990-2010. Dependent Var: 100 x Change in Birth Rate, Share of Births to Teen Mothers, Share of Births to Unmarried Mothers, Share of Births with no Information about Fathers on Birth Certificate (in %pts).

	Births per 1,000 Women Age 15-39		Share of Births to: Teenage Mothers		Unmarried Mothers		No Father on Birth Certificate	
	(1)		(2)		(3)		(4)	
			I. Overall Trade Shock					
Δ Chinese Import Penetration	-2.30	**	0.36	*	0.44		1.43	*
	(0.64)		(0.18)		(0.34)		(0.59)	
			II. Male Industry vs Female Industry Shock					
Δ Chinese Import Penetration	-7.91	**	1.20	~	2.55	~	2.89	
× (Male Ind Emp Share)	(1.86)		(0.64)		(1.42)		(2.18)	
Δ Chinese Import Penetration	3.51	**	-0.52		-1.92	~	-0.45	
× (Female Ind Emp Share)	(1.17)		(0.38)		(1.02)		(1.32)	
Mean Outcome Variable	-4.05		-1.70		6.94		-0.45	
Level in 1990	82.3		12.8		27.7		14.4	

Notes: N=1444 (722 CZ x 2 time periods). Regressions weighted by start-of-period CZ population. Standard errors clusterd on state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

4.3 Children's household living circumstances

We finally trace these marriage and fertility effects downstream to their consequences for the household circumstances of children ages 0 to 17. We measure household circumstances along two dimensions: household poverty status and household headship. Table 8 presents estimates for the period 1990 through 2014 using the stacked first-difference specification. The column 1 estimate suggests a sizable impact of labor market shocks on the incidence of household poverty among U.S. children. We find that a one-unit trade trade shock raises the fraction of children under age 18 living below the poverty line by approximately 0.6 percentage points ($t = 2.3$), which is roughly a 3% increase in the prevalence of childhood poverty relative to the base poverty rate of 18.0% of U.S. children in 1990. Columns 2 through 5 show that an aggregate one-unit trade shock reduces the fraction of children in married-couple households by 0.35 percentage points, most of which is accounted for by a rise in the fraction living in single-headed households.

Since poverty is far more prevalent among single-headed than married households (bottom row of Table 8), one might surmise that the effect of trade shocks on the prevalence of childhood poverty (column 1) stems to a large extent from the a rise in poor single-headed households. Appendix Table A5 shows that this is indeed the case. When we decompose the impact of the shock on the fraction of children living in poor and non-poor households for each of the five household types above, we find that half of the rise in poverty is due to a greater likelihood that children liven in poor single-headed homes. Conversely, the only household setting whose probability is significantly reduced by the main trade shock is non-poor married households.

Table 8: 2SLS Estimates of the Impact of Import Penetration on Socio-Economic Outcomes of Children Age 0 - 17, 1990-2014. Dependent Var: 100 x Change Share of Age Group in Indicated Condition (in %pts)

	Household Income < Poverty Line		Household Headed by: Married Parents		Parent and Unmarried Patner		Parent without Partner		Grand-Parents	Any Other	
	(1)		(2)		(3)		(4)		(5)	(6)	
					I. Overall Trade Shock						
Δ Chinese Import Penetration	0.61	*	-0.35	~	-0.11		0.30	**	0.19	-0.03	
	(0.26)		(0.19)		(0.07)		(0.11)		(0.13)	(0.07)	
					II. Male Industry vs Female Industry Shock						
Δ Chinese Import Penetration × (Male Ind Share)	2.13	**	-1.85	**	0.28		1.43	**	0.43	-0.29	
	(0.70)		(0.50)		(0.23)		(0.32)		(0.33)	(0.18)	
Δ Chinese Import Penetration × (Female Ind Share)	-1.12		1.36	*	-0.55	*	-0.98	*	-0.09	0.27	~
	(0.82)		(0.55)		(0.25)		(0.42)		(0.23)	(0.15)	
Mean Outcome Variable	1.65		-4.69		1.62		1.79		1.01	0.27	
Mean Level in 1990	17.99		71.39		2.82		16.82		5.43	3.53	
Poverty Rate (%) in 1990	n/a		8.1%		42.3%		47.4%		25.1%	34.4%	

Notes: N=1444 (722 CZ x 2 time periods). The Census records every household member's relationship to the household head, who is the person that owns or rents the household's dwelling. All regressions include the full set of control variables from Table 1. Regressions are weighted by start-of-period CZ population and standard errors are clusterd by state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

A second pattern conveyed by the lower panel of Table 8 is that the entirety of the impact of adverse economic shocks on child poverty arises from shocks to male employment, while female-specific labor-market shocks have an insignificant negative effect on the incidence of child poverty. Appendix Table A5 helps to interpret this finding. The male-specific shock greatly reduces the fraction of children in non-poor married households (by 2.4 percentage points) while raising primarily the fraction of children in poor single-parent homes (by 1.1 percentage points), as well as the fractions of children in poor married and unmarried couple homes (by 0.5 and 0.3 percentage points, respectfully). This finding suggests that an adverse shock to *male* earnings increases child poverty not only through its direct effect on male earnings, but also by raising the prevalence of households structures such as single-parent homes which face an elevated poverty risk more generally. The complementary finding that adverse shocks to *female* employment have no significant impact on the prevalence of childhood poverty results from the combination of lower female earnings which raise poverty, and a shift towards married households, which reduces it. The estimates in Appendix Table A5 lend weight to this interpretation: adverse shocks to female earnings opportunity reduce the fraction of children living in poor single-parent households and poor cohabiting households while raising the proportion of children in non-poor married households.

As noted in the Introduction, the decline in marriage and rise in single-headedness in the U.S. is pervasive across major race and ethnic groups, which raises the question of whether similar labor-market and marriage-market relationships are present for all groups. A detailed breakdown of our results by race and ethnicity is complicated by the fact that many local labor markets comprise only very small headcounts of minorities like blacks and Hispanics. However, Appendix Table A6 confirms that key results of our analysis, such as the rise in unmarried women, single-headed households and child poverty are also present when we focus the analysis exclusively on non-Hispanic whites, thus suggesting that our results are not primarily driven by minorities.

5 Conclusions

The multiple complementary analyses in this paper provide an integrated narrative for the impact of labor-market shocks on fertility and household structure. Adverse shocks to local employment opportunities stemming from rising international competition from China in manufactured goods yield a fall in men's relative employment and earnings; an increase in the rate of male mortality from drug and alcohol abuse; a reduction in the net availability of marriage-age males in affected labor markets; a reduction in the fraction of young adults entering marriage; a fall in fertility accompanied by a weak rise in the fraction of births to teen and unmarried mothers; and a sharp jump in the fraction of children living in impoverished and single-headed households.

Two cardinal results help to weave these many empirical strands together. A first is that trade shocks faced by the U.S. manufacturing sector—which employs a disproportionate share of male workers—reduce the economic stature of men relative to women. Consistent with this pattern, shocks to male-intensive manufacturing industries are particularly destabilizing to marriage-markets and reduce marriage rates. A second broad result, is that gender-specific shocks to labor-market outcomes have strikingly non-parallel impacts on fertility patterns. Male-specific shocks reduce overall fertility, but reduce it by *less* among teens and unmarried mothers than among older and married mothers, thereby increasing the fraction of children born out of wedlock and living in poverty. Conversely, female-specific shocks raise overall fertility but reduce the share of births to teens and unmarried mothers, thus contributing to a reduction in the fraction of children living in single-headed households.

We conclude that the declining employment and earnings opportunities faced by young (i.e., under 40) U.S. males are a plausible contributor to the changing structure of marriage and childbirth in the United States. Although our analysis does not imply that surging import competition from China over the last two decades has been the sole or primary driver of these trends, it highlights that broader declines in the labor market for U.S. males have likely made a substantial contribution—assuming of course that these declines exert qualitatively similar causal effects on marriages, births and household structures as do trade shocks concentrated on manufacturing employment.

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Appendix Figures and Tables

Figure A1: Bin-Scatter of the Commuting Zone Level Relationship in the Year 2000 Between the Share of Adults Age 18-39 who are Not Currently Employed and the Male-Female Mean Annual Earnings Gap (panel A) and the Share of Women Age 18-39 that Are Currently Married (panel b)

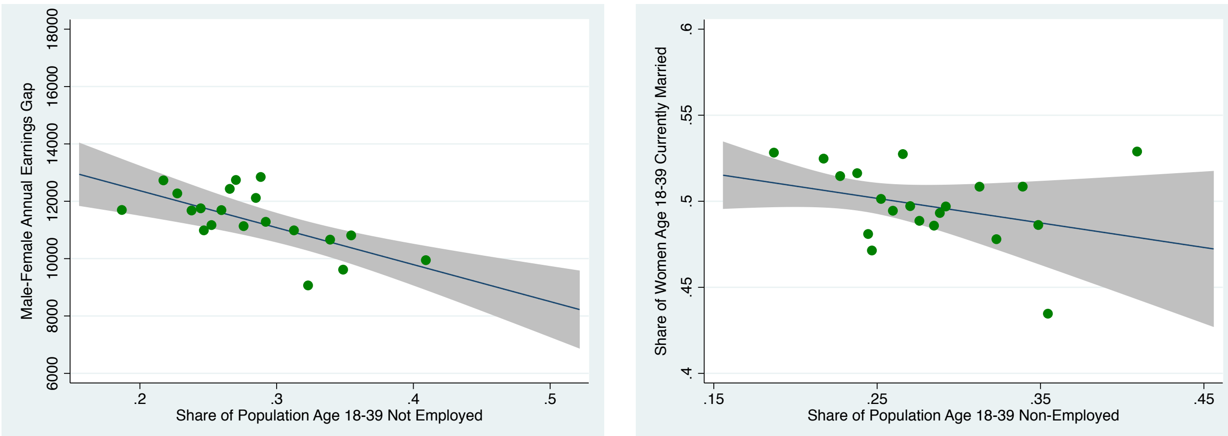

Table A1: Mean and Percentiles of Decadal Growth in Import Penetration, Overall and Sex-Specific Measures

	Δ Chinese Import Penetration		
	1990-2014	1990-2000	2000-2014
	(1)	(2)	(3)
		I. Overall Shock	
Mean	1.07	0.95	1.15
	(0.71)	(0.61)	(0.77)
P25	0.64	0.54	0.73
P50	0.92	0.89	1.01
P75	1.30	1.22	1.30
P75-P25	0.66	0.68	0.57
		II. Male Industry Shock	
Mean	0.63	0.56	0.69
	(0.40)	(0.33)	(0.43)
P25	0.38	0.35	0.40
P50	0.58	0.53	0.62
P75	0.80	0.73	0.80
P75-P25	0.42	0.38	0.41
		III. Female Industry Shock	
Mean	0.43	0.39	0.46
	(0.35)	(0.31)	(0.38)
P25	0.23	0.21	0.25
P50	0.35	0.34	0.37
P75	0.50	0.48	0.52
P75-P25	0.27	0.27	0.27

Notes: N=1444 (722 commuting zones x 2 time periods) in column 1, N=722 in columns 2 and 3. Observations are weighted by start of period commuting zone share of national population.

Table A2: Regression Adjusted Differences in Earnings and Hours between Manufacturing and Non-Manufacturing Workers. Dependent Variables: Level and Log of Annual Wage and Salary Income, Log Annual Work Hours and Log Earnings per Hour for Workers Age 18-39 in the 2000 Census

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	I. Annual Wage and Salary Income (US$)				II. Log Annual Wage and Salary Income			
Male x Employed in Manufacturing	1518 ~	4077 **	3906 **	4959 **	0.13 **	0.17 **	0.17 **	0.19 **
	(827)	(587)	(603)	(408)	(0.01)	(0.01)	(0.01)	(0.01)
Female x Employed in Manufacturing	2503 **	5306 **	5333 **	5920 **	0.15 **	0.25 **	0.25 **	0.26 **
	(464)	(341)	(348)	(261)	(0.02)	(0.01)	(0.01)	(0.01)
Mean (S.D.) Outcome Var Males / Females	40871 (43768)				10.18 (1.07)			
	28099 (29465)				9.77 (1.14)			
	III. Log Annual Work Hours				IV. Log Hourly Wage			
Male x Employed in Manufacturing	0.08 **	0.08 **	0.08 **	0.08 **	0.05 **	0.09 **	0.09 **	0.11 **
	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)
Female x Employed in Manufacturing	0.14 **	0.17 **	0.17 **	0.17 **	0.00	0.08 **	0.08 **	0.09 **
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Mean (S.D.) Outcome Var Males / Females	7.39 (0.72)				2.67 (0.70)			
	7.14 (0.85)				2.51 (0.69)			
Age x Gender	yes	yes	yes	yes	yes	yes	yes	yes
Education x Gender		yes	yes	yes		yes	yes	yes
Race/Nativity x Gender			yes	yes			yes	yes
CZone Fixed Effects				yes				yes

Notes: N=243,071 (130,181 male and 112,890 female workers). The analysis is based on the 5% IPUMS sample of the 2000 Census, and includes all individuals age 18-39 who report a positive wage and salary income and who are not classified as self-employed, unpaid family members, or residents of institutional group quarters. The fraction of workers employed in the manufacturing sector is 15.2% for males and 11.2% for females. Annual hours are the product of weeks worked and usual works hours per week. Top-coded incomes are replaced by the average income for top-coded observations within a state. The full control vector in column 4 includes a gender dummy interacted with 22 indicators for age in years, 9 indicators for eduction levels, 3 indicators for race and ethnicity, and an indicator for foreign-born individuals. It also includes 721 indicators for Commuting Zones (CZs). Regressions are weighted by the product of the Census person weight and the weighting factor that attributes individuals from Census PUMAs to CZs. Standard errors are clusterd by state. ~ $p \le 0.10$, * $p \le 0.05$, ** $p \le 0.01$.

Table A3: 2SLS Estimates of the Impact of Import Penetration on Cumulative Applications and Enlistments in the U.S. Army, 2001 - 2011. Dependent Variable: Number of Applicants or Enlistments 2001-2011 by CZ Denominated by Relevant Population Group

	Age 17-40		Age 17-25		Age 26-40		Un-Married		Married		Non-College		Some College	
	I. Men: Cumulative Applications per Population, 2001-2011													
Δ Chinese Penetration	0.179	*	0.266	*	0.042	~	0.312	*	0.045	~	0.304	*	0.033	
	(0.076)		(0.114)		(0.022)		(0.149)		(0.026)		(0.145)		(0.020)	
Mean of Outcome	3.25		5.15		0.62		5.87		0.75		6.08		0.49	
	II.Men: Cumulative Enlistments per Population, 2001-2011													
Δ Chinese Penetration	0.069	**	0.094	*	0.019	**	0.096	~	0.022	*	0.108	~	0.007	
	(0.027)		(0.043)		(0.007)		(0.053)		(0.010)		(0.057)		(0.007)	
Mean of Outcome	1.28		2.05		0.23		2.33		0.29		2.41		0.18	
	III. Women: Cumulative Applications per Population, 2001-2011													
Δ Chinese Penetration	0.068	*	0.099	*	0.011	~	0.114	~	0.027	**	0.111	~	0.014	~
	(0.030)		(0.048)		(0.007)		(0.059)		(0.008)		(0.067)		(0.007)	
Mean of Outcome	0.96		1.60		0.15		1.71		0.23		2.06		0.14	
	IV. Women: Cumulative Enlistments per Population, 2001-2011													
Δ Chinese Penetration	0.022	*	0.033	*	0.006	*	0.035	*	0.010	*	0.047	*	0.003	
	(0.010)		(0.016)		(0.003)		(0.017)		(0.004)		(0.023)		(0.003)	
Mean of Outcome	0.30		0.51		0.05		0.54		0.08		0.66		0.05	

Notes: n=722 CZs in each cell. Source Army application and enlistment records aggregated to the CZ by year level, used under agreement with the U.S. Army Office of Economic and Manpower Analysis. All models control for: eight Census region trends; start-of-period shares of commuting zone population that are Hispanic, black, Asian, other race, foreign born, and college educated; and the fraction of adults that are employed in manufacturing, the fraction of women who are employed, and the start-of-period indices of employment in routine occupations and of employment in offshorable occupations as defined in Autor and Dorn (2013). Population denominators by CZ for application and enlistment counts are calculated using the Census 2000 IPUMS. Robust standard errors in parentheses are clustered on state. Models are weighted by start of period commuting zone share of national population. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Table A4: 2SLS Estimates of the Impact of Import Penetration on CZ-Level Death Rates by Sex, 1990-2014. Dependent Var: Male and Female Deaths per 100k Adults Ages 20-39 by Cause of Death

	Total (1)		Drug/ Alc Poison (2)		Liver Disease (3)	Diabetes (4)	Lung Cancer (5)		Suicide (6)		All Other (7)
	I. Male Death Rates										
	Overall Trade Shock										
Δ Chinese Penetration	3.86		2.25	~	-0.07	-0.18	-0.20	~	0.25		1.79
	(2.80)		(1.28)		(0.21)	(0.22)	(0.11)		(0.79)		(2.22)
	Male Industry vs Female Industry Shock										
Δ Chinese Penetration	22.87	**	6.27		-1.01	-0.43	-0.52		0.42		14.47
× (Male Ind Share)	(8.39)		(4.10)		(0.86)	(0.66)	(0.46)		(1.83)		(8.02)
Δ Chinese Penetration	-15.95	*	-1.94		0.90	0.09	0.13		0.07		-11.40
× (Female Emp Share)	(6.99)		(2.74)		(0.66)	(0.50)	(0.34)		(1.94)		(6.04)
Mean of Outcome	-32.38		8.34		-1.07	-0.02	-0.43		-1.10		-38.08
Level in 1990	213.43		6.39		4.11	1.95	1.55		25.12		174.31
	II. Female Death Rates										
	Overall Trade Shock										
Δ Chinese Penetration	-0.86		0.07		-0.18	-0.07	0.02		0.36		-1.02
	(2.31)		(0.67)		(0.17)	(0.15)	(0.13)		(0.41)		(1.91)
	Male Industry vs Female Industry Shock										
Δ Chinese Penetration	-8.48		-0.65		-0.62	0.25	0.42		-1.15		-7.94
× (Male Ind Share)	(10.24)		(2.28)		(0.79)	(0.55)	(0.53)		(1.32)		(9.01)
Δ Chinese Penetration	7.07		0.82		0.27	-0.41	-0.39		1.93	~	6.15
× (Female Emp Share)	(6.87)		(1.45)		(0.58)	(0.48)	(0.44)		(1.03)		(5.96)
Mean of Outcome	-4.41		4.32		-0.38	0.02	-0.20		-0.01		-8.16
Level in 1990	78.89		1.92		1.91	1.44	1.00		5.57		67.05

Notes: N=1444 (722 CZ x 2 time periods). All regressions include the full set of control variables from Table 1 and the start of-period value of the outcome variable. Regressions are weighted by start-of-period population and standard errors are clustered on state. ~ p ≤ 0.10, * p ≤ 0.05, ** p ≤ 0.01.

Table A5: Decomposition of the Impact of Import Penetration on the Prevalence of Childhood Poverty, Overall and by Household Structure, 1990-2014 (2SLS Estimates). Dependent Var: 100 x Change Share of Children Living in Each Household Type and Poverty Status

	All HH's		Household Headed by: Married Parents		Parent and Unmarried Patner		Parent without Partner		Grand-Parents		Any Other
	(1)		(2)		(3)		(4)		(5)		(6)
	A. Percent in Household Type w/ Income < Poverty Line										
	A1. Overall Trade Shock										
Δ Chinese Penetration	0.61	*	0.19		-0.03		0.29	**	0.12	*	0.04
	(0.26)		(0.17)		(0.04)		(0.10)		(0.05)		(0.04)
	A2. Male Industry vs Female Industry Shock										
Δ Chinese Penetration	2.13	**	0.53		0.28	~	1.12	**	0.15		0.06
× (Male Ind Share)	(0.70)		(0.37)		(0.16)		(0.37)		(0.14)		(0.11)
Δ Chinese Penetration	-1.12		-0.19		-0.37	*	-0.65	*	0.08		0.02
× (Female Ind Share)	(0.82)		(0.55)		(0.18)		(0.31)		(0.13)		(0.11)
	B. Percent in Household Type w/ Income >= Poverty Line										
	B1. Overall Trade Shock										
Δ Chinese Penetration	-0.61	*	-0.54	**	-0.08		0.01		0.07		-0.07
	(0.26)		(0.21)		(0.06)		(0.11)		(0.09)		(0.05)
	B2. Male Industry vs Female Industry Shock										
Δ Chinese Penetration	-2.13	**	-2.38	**	0.01		0.31		0.28		-0.35
× (Male Ind Share)	(0.70)		(0.55)		(0.14)		(0.33)		(0.23)		(0.12)
Δ Chinese Penetration	1.12		1.55	*	-0.18		-0.33		-0.17		0.25
× (Female Ind Share)	(0.82)		(0.70)		(0.13)		(0.33)		(0.15)		(0.10)

Notes: N=1444 (722 CZ x 2 time periods). The Census records every household member's relationship to the household hea who is the person that owns or rents the household's dwelling. All regressions include the full set of control variables from Table 1. Regressions are weighted by start-of-period CZ population and standard errors are clusterd by state. ~ p ≤ 0.10, * p 0.05, ** p ≤ 0.01.

Table A6: 2SLS Estimates of the Impact of Import Penetration on Employment, Earnings, Marital Status and Children's Environment among non-Hispanic Whites, 1990-2014. Dependent Var: 100 x Change in Marital Status among Women Age 18-39 and Living Environment Children Age <18 (in %pts)

	A. Marital Status Women 18-39		B. Children's Environment Age <18	
	% Nev. Marr. Women	% Single Mothers	% Child Poverty	% Child w/ Single Parent
	(1)	(2)	(3)	(4)
	Ia. Full Population: Overall Trade Shock			
Δ Chinese Penetration	1.16 **	0.52 ~	0.61 *	0.30 **
	(0.33)	(0.31)	(0.26)	(0.11)
	Ib. Full Population: Male Industry vs Female Shock			
Δ Chinese Penetration × (Male Ind Share)	4.23 **	3.28 **	2.13 **	1.43 **
	(0.64)	(0.73)	(0.70)	(0.32)
Δ Chinese Penetration × (Female Ind Share)	-2.32 **	-2.62 **	-1.12	-0.98 *
	(0.58)	(0.85)	(0.82)	(0.42)
	IIa. Non-Hispanic Whites: Overall Trade Shock			
Δ Chinese Penetration	0.73	0.96 ~	0.59 **	0.40 **
	(0.51)	(0.58)	(0.21)	(0.11)
	Ib. Non-Hispanic Whites: Male Industry vs Female Shock			
Δ Chinese Penetration × (Male Ind Share)	4.91 **	5.05 **	1.35 *	1.17 **
	(1.12)	(1.29)	(0.60)	(0.28)
Δ Chinese Penetration × (Female Ind Share)	-4.45 **	-4.12 **	-0.29	-0.49 ~
	(1.58)	(1.42)	(0.57)	(0.30)

Notes: N=1444 (722 CZ x 2 time periods). Panel I reiterats results from previous tables for the full population, while panel II analyzes outcomes for the non-Hispanic white poulation only. Standard errors are clustered on state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.