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 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. As predicted by a simple model of marital decision-making under uncertainty, 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 in poor single-parent households. The falling marriage-market value of young men appears to be a quantitatively important contributor to the rising rate of out-of-wedlock childbearing and single-headed childrearing in the United States.

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.

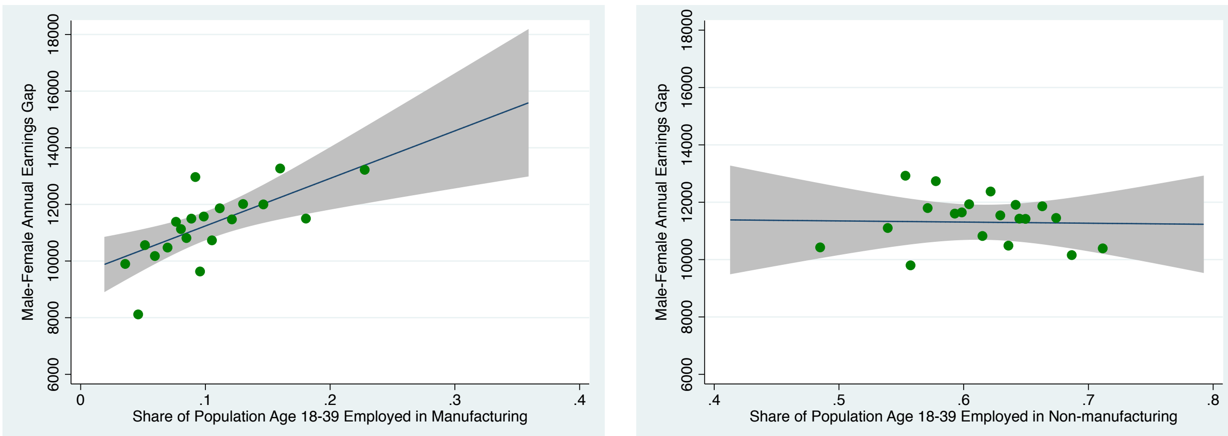
"Wilson's book spoke to me. I wanted to write him a letter and tell him that he had described my home perfectly. That it resonated so personally is odd, however, because he wasn't writing about the hillbilly transplants from Appalachia—he was writing about black people in the inner cities." J.D. Vance,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

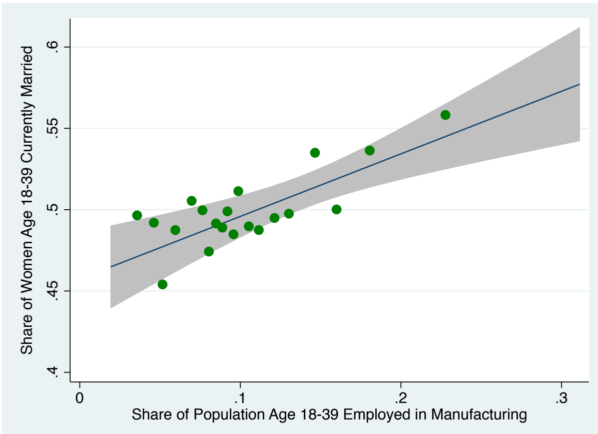
Following Autor, Dorn and Hanson (2013b), we exploit cross-industry and cross-local-labor-market (i.e., commuting zone) variation in import competition stemming from China's rapidly rising productivity and falling barriers to trade to identify market-level labor-demand shocks that are concentrated in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

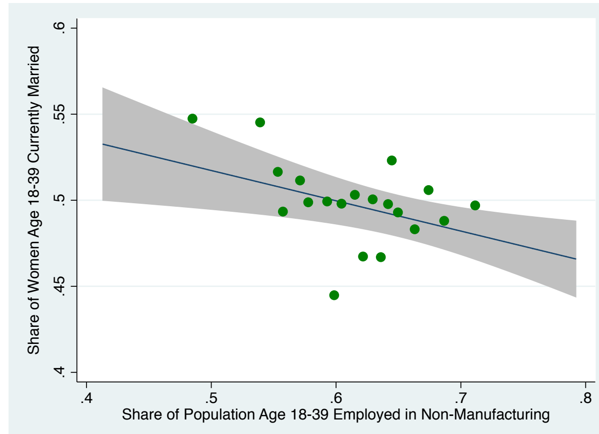
[4] Related to these observations, Summers (1986) uses cross-state panels to document that employment growth in the high-wage industries of manufacturing, mining, construction, transportation, and public utilities predicts declines in state unemployment rates while comparable employment growth in low-wage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

Kearney and Wilson (2016), Schaller (forthcoming 2016), and Shenhav (2016).[7] We exploit gender dissimilarities in industry specialization to identify demand shocks that distinctly affect men's and women's employment and earnings. Our use of trade shocks as a source of variation further allows us to assess whether two decades of contracting U.S. manufacturing employment in labor-intensive sectors, stemming in substantial part from rising international competition from China, has contributed to the rapid, simultaneous decline of traditional household structures.

We offer a simple conceptual model of marital decision-making under uncertainty based on Kane and Staiger (1996) that guides the interpretation of the analysis. Less-skilled unmarried women have a preference for becoming married mothers but face uncertainty about the quality of their male partners, which is not fully revealed until after conception of a child. Women who strongly prefer marriage to single-parenthood (to whom we will refer as having *traditional* preferences) will curtail *both* fertility and marriage when high-quality men become scarce rather than risk becoming pregnant by and then marrying a low-quality man. Women who are willing to exercise the option

[7] Ananat, Gassman-Pines and Gibson-Davis (2013) find that adverse 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. Kearney and Wilson (2016) find that positive shocks to male earnings do not increase marriage, but do raise fertility and reduce the non-marital birth share.

of single-motherhood in the event of a bad draw of a male partner (those with *non-traditional* preferences) will curtail marriage but not (necessarily) fertility when high-quality males become scarce. Because of the asymmetric fertility responses of women who do and do not value the option of single-motherhood, this framework predicts that marriage is more elastic than fertility to the supply of high-quality males. A fall in the supply of high-quality males reduces fertility while *raising* the fraction of children born out-of-wedlock and living with unmarried mothers.

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 annual wage-and-salary earnings. Although earnings 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 2.2 percent at the median and by nearly 17 percent at the 25^{th} percentile. It also increases the share of young men in a local labor market who earn less than women of the same age, race and education.[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 1.7 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 more than doubles the relative male death rate from drug and alcohol poisoning. 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

[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.

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, with especially large effects on the youngest adult women, ages 18-25. More subtly, we find asymmetric marriage-market impacts that depend upon the source of the shock: adverse shocks to labor demand in male-intensive industries reduce the prevalence of marriage among young women, whereas analogous shocks to female labor demand significantly *raise* the prevalence of marriage. These asymmetric responses are predicted by our conceptual model: negative shocks concentrated on males reduce their marriage-market value, thereby discouraging both fertility and marriage; negative shocks concentrated on women raise women's disutility of single-motherhood, discouraging fertility but *encouraging* 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 3.3 (a 4% 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. Recent literature hypothesizes that the high U.S. teen birth rate is in substantial part due to the dearth of economic opportunity facing young non-college women (Kearney and Levine, 2012). Consistent with this hypothesis, we find that adverse shocks to female labor demand increase the fertility rate among teens, although the fraction of teen and unmarried births *declines* by as much as adverse shocks to male labor demand increase these shares. This asymmetry is consistent with our simple option-value model of fertility: holding women's economic opportunities constant, a decline in male earnings spurs some women to curtail both motherhood and marriage while spurring others to exercise the option of single-headedness (curtailing marriage but not fertility), thus raising teen and out-of-wedlock birth shares; conversely, holding men's economic opportunities constant, a decline in female earnings raises the relative attractiveness of male partners, which encourages

[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.

fertility and marriage while single motherhood becomes a less attractive option.

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 2.2 percentage points (a 12% increase), reduces the fraction living in married households by 0.4 percentage points, and spurs a concomitant rise in the share living in single- and grandparent-headed households. The asymmetric effect of male and female labor demand shocks seen for marriage and fertility 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 (i.e., exercising the option of single parenthood). 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 male labor-demand shocks raise the fraction of children living in poverty, female labor-demand shock have no 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 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 does

[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.

[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

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 Conceptual Underpinnings

We consider a setting where unmarried women have a preference to become married mothers but face uncertainty about the availability of high-quality men who may serve as marital partners.[12] A substantial literature documents that the marriage decision tends to follow the fertility decision: upon becoming pregnant, a woman may choose to marry the child's father but absent pregnancy would not elect marriage.[13] We impose this setting on decision-making by assuming that women choose to remain childless, to have a child and marry, or to become a single mother. Removing the option of marriage without children narrows the generality of the model but is not restrictive empirically since nearly 90% of women ages 18-39 are either mothers, or unmarried without children.[14]

Suppose that at the time of considering motherhood, women are uncertain whether the potential father is a high-quality parent, as father quality is not revealed until after pregnancy has occurred. Women who choose pregnancy face two options at the time that partner quality is revealed: those who find that their partners are high-quality will elect marriage; those who find that their partners are low-quality will choose either to marry their low-quality partners or to raise their children out-of-wedlock, whichever has greater utility. In this setting, an inward shift in the supply of high-quality men unambiguously reduces marriage and fertility. Simultaneously, this supply shift may *increase* the fraction of births that are out of wedlock and hence the share of children raised in single-headed households. The intuition, formalized below, is as follows: for women who are committed to raising

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] Our conceptual framework is adapted from Kane and Staiger (1996), who analyze the interaction between abortion restrictions, fertility, and out-of-wedlock births. In related work, Akerlof et al. (1996) consider how women's forward-looking decisions about sexual intercourse and marital pre-commitment are shaped by their outside options, which may include abortion, in the event of pregnancy. We abstract from the option of abortion in the model and in our subsequent empirical analysis (it is not observed in our data).

[13] Edin and Tach (2012) calculate that among women currently over the age of 24 in 2006 through 2008, 53% were mothers by the age of 24, and 65% of those mothers were unmarried at the time of their first birth. Seventy-six percent of first births in 2007 were to mothers under the age of 30, and 46% were to women under the age of 25 (Martin et al., 2010, Table 3).

[14] The share of women age 18-39 who are married but have no children declined from 12.5% to 10.2% from 1990 to 2007 in Census/ACS data. This demographic status is less prevalent than any other combination of marital and motherhood status—married with children (decline from 40.6% to 31.9%), unmarried and childless (increase from 34.3% to 42.9%), and unmarried with children (increase from 12.7% to 15.0%).

children in wedlock—that is, who have high disutility of non-marital childrearing whom we refer to as having *traditional* preferences—a decline in the availability of high-quality males deters both pregnancy and marriage. For women who have a comparatively low psychic cost of non-marital childrearing (*non-traditional* preferences), a shrinking pool of high-quality men deters fertility by less than it deters marriage, shifting the composition of fertility towards non-marital births.

Formally, let potential male partners be either high or low-quality, where quality denotes ability to provide economic and emotional inputs for parenting that are valued by the mother. At a given point in time t, a fraction P_{ij} of the potential partners for a woman i in commuting zone j are high-quality while $1 - P_{ij}$ are of low-quality. The expectation of P_{ij} is common knowledge, but a woman cannot verify the quality of an individual male partner until she conceives a child with that partner. We normalize the utility of not having a child at zero and the utility of having a child with a high-quality father at one, and we assume that women maximize expected utility. The utility for woman i of marrying a low-quality father is $-M_i$, while her utility of raising a child out of wedlock is $-S_i$, with $M_i, S_i > 0$ for all i. The variables P_{ij}, M_i and S_i can all very according to a women i's individual tastes or demographic characteristics.

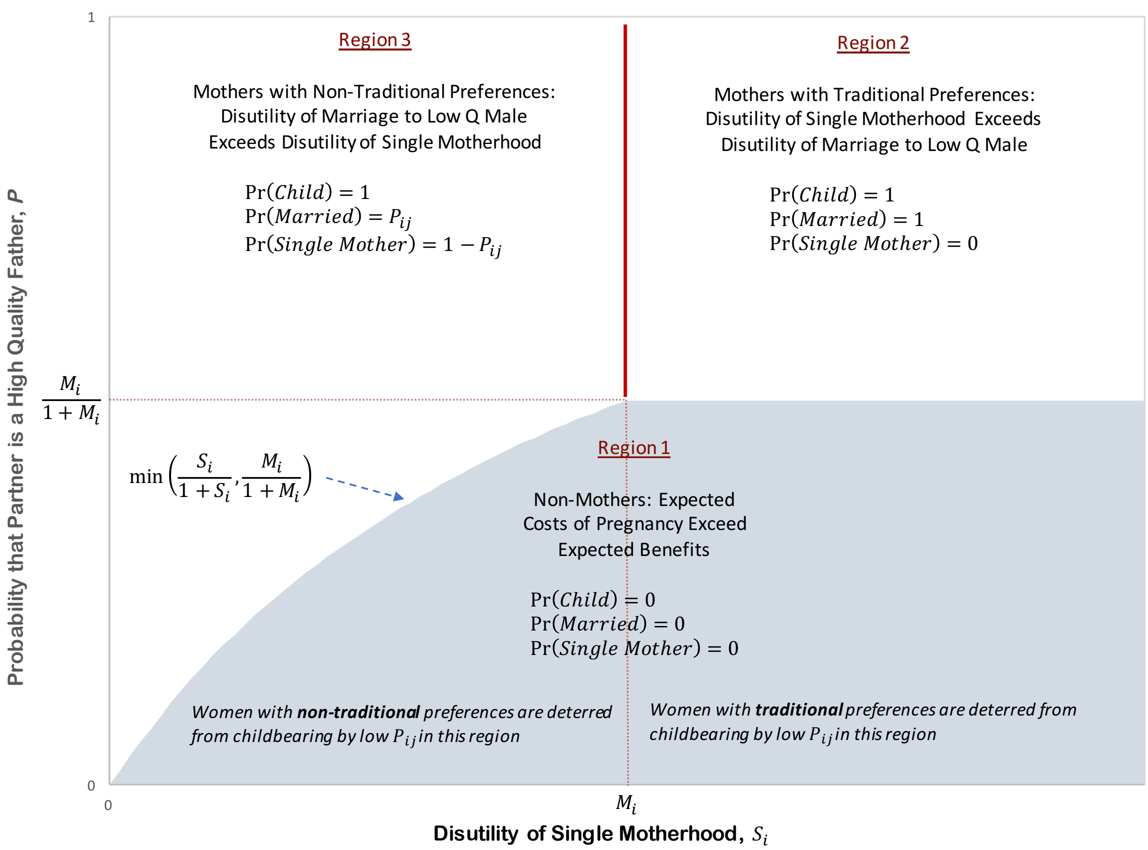
Labor-market conditions in commuting zone j will affect women's choices by changing the fraction P_{ij} of men that are perceived to be attractive partners. A negative shock to labor demand for males reduces the proportion of high-quality male partners P_{ij} in the local labor market, while a negative shock to female labor demand increases P_{ij}, as a set of males with a given income level looks more attractive to women whose own earnings are low. The general premise that higher relative earnings of males make them more attractive partners is consistent with a long literature going back to Becker (1973). The more specific and simplifying assumption that male quality depends (in part) on male *relative* earnings relates to recent work by Bertrand, Kamenica and Pan (2015) who argue that potential partners have a strong preference for matches in which the man's earnings exceed the woman's, eschewing matches that combine a higher-earning women with a lower-earning male. In this setting, a woman will choose to conceive a child if

$$\frac{P_{ij}}{1 - P_{ij}} > \min\left[M_i, S_i\right],$$

that is, if the odds that her partner is revealed to be high-quality after conception are sufficiently high to overcome the risk that she will have to marry a low-quality father (if $M_i \leq S_i$) or become a single mother (if $M_i > S_i$) .

Figure (3) illustrates the operation of this simple model. The $y-$axis of the figure corresponds to P_{ij}, the probability that a male partner will be revealed to be of high quality following conception. The $x-$axis corresponds to the disutility of single-motherhood, S_i. For concreteness, we depict the preferences of women who have a given level of disutility M_i of marrying a low-quality father.

Figure 3: Women's Choices Over Pregnancy, Marriage and Single-Motherhood as a Function of Expected Male Quality (P_{ij}) and Disutility of Single Motherhood (S_i) for a Given Level of Disutility of Marrying a Low-Quality Male (M_i)



Region 1 in the figure depicts the area in which $P_{ij} < \min\left(\frac{S_i}{1-S_i}, \frac{M_i}{1-M_i}\right)$. In this region, women will choose against motherhood because the probability that a father proves to be high-quality are too small to overcome the downside risk of either marrying a low-quality father ($-M_i$) or raising a child as a single mother ($-S_i$).[15] Region 2 captures women for whom the disutility of single-motherhood exceeds the disutility of marrying a low-quality father ($-S_i < -M_i$) *and* for whom

[15]Note that the condition for fertility, $\frac{P_{ij}}{1-P_{ij}} > \min(S_i, M_i)$, can be rewritten as $P_{ij} > \min\left(\frac{S_i}{1-S_i}, \frac{M_i}{1-M_i}\right)$.

the benefits of pregnancy exceed the downside risk. Women with these *traditional* preferences will choose motherhood *and* will marry the child's father whether or not he is ultimately revealed to be of high or low quality. Region 3 depicts the set of women for whom single-motherhood is preferable to marrying a low-quality father ($-M_i < -S_i$) *and* for whom the benefits of pregnancy exceed the downside risk of raising a child out of wedlock. Women with these *non-traditional* preferences will choose motherhood but will marry the child's father only if he is revealed to be high-quality.

Consider the effects on fertility and marriage of a shock in commuting zone j that worsens the labor-market opportunities for males, thus reducing the local supply of high-quality fathers (a drop in P_{ij}). Among women with traditional preferences (right half of the figure), outcomes are unchanged with the decline in P_{ij} for those women who remain in either Region 1 or Region 2. For some women, however, P_{ij} falls below $M_i/(1 - M_i)$ and pushes them from Region 2 into Region 1, where they abstain from both motherhood and marriage. Thus, for women with traditional preferences, the probability of motherhood and marriage falls in lockstep.

Women with non-traditional preferences (left half of the figure) face a more complex choice set since single-motherhood provides option value should the father of the child prove to be low-quality. A first group is already initially in Region 1 and continues to forgo motherhood and marriage as the quality of potential spouses deteriorates. A second group moves from Region 3 to Region 1 as P_{ij} falls below $S_i/(1 - S_i)$, and thus abstains from both fertility and marriage. Finally, a third group of women with non-traditional preferences remains in Region 3 as P_{ij} continues to be larger than $S_i/(1 - S_i)$. These women will not adjust fertility, but as they prefer single-motherhood to marrying a low-quality father, their marriage rate declines as the supply of high-quality males falls.

In combination, the adjustments among women with traditional and non-traditional preferences sum to an unambiguously negative impact of a deterioration in male partner quality on both fertility (by shifting women from Regions 2 and 3 to Region 1) and marriage (by shifting women into Region 1 and reducing marriage rates within Region 3). Under mild additional assumptions, marriage rates will fall more than fertility, thus increasing the fraction of out-of-wedlock births. Intuitively, the shift of women into Region 1 has modest implications for the rate of single motherhood among the remaining mothers, as it reduces both the number of traditional mothers who are always married and the number of marginal non-traditional mothers who are mostly single. The main impact of the economic shock on single motherhood thus operates via the declining likelihood of marriage among mothers with non-traditional preferences (those who remain in Region 3).[16]

Consider next the comparative statics for a shock in commuting zone j that reduces labor-market opportunities for women, raises the relative earnings of males, and thus increases the likelihood that males will be perceived as being of high quality (an increase in P_{ij}). The shift of women from Region 1 to Regions 2 and 3 increases both fertility and the number of marriages. In addition, women in Region 3 become more likely to choose marriage over single motherhood, and contrary to the impact of a shock to male labor demand, the share of out-of-wedlock births is thus likely to decline. Thus, opposite to the case of adverse shocks to relative male earnings, adverse shocks to relative female earnings are likely to increase transitions into marriage by more than transitions into motherhood.

Despite its simplicity, this framework encapsulates a broadly applicable insight: women must make forward-looking, irreversible fertility decisions in a setting where the consequences of pregnancy—father quality in particular—are uncertain until at least the time that the child is conceived. Facing the possibility of obtaining a low-quality partner, women's outside options play a critical role in determining their willingness to risk childbearing. This simple framework offers a number of predictions that we subsequently test and confirm:

1. Adverse shocks to male earnings capacity: (a) reduce overall fertility and the prevalence of marriage, and (b) *reduce marriages by more than births* (because some women with non-traditional preferences chose to become single mothers rather than marrying low-quality males), thereby increasing the share of children born out-of-wedlock and raised in single-headed households;

2. Adverse shocks to women's earnings capacity: (a) increase overall fertility and marriage rates, and (b) *increase marriages more than births,* thus decreasing the share of children born out-

[16] More formally, denote by m the probability that a women has traditional preferences, by b the fraction of traditional women that choose motherhood both before or after a given shock $\Delta P_{ij} < 0$, and by β the fraction of traditional women that would become mothers only absent the shock, while a and α correspondingly denote the fractions of non-traditional women who would choose motherhood regardless of the shock, or only absent the shock. Finally, p and π are the average pre-shock marriage rates for the two groups of non-traditional women who become mothers either regardless of the shock or only absent the shock. The shock $\Delta P_{ij} < 0$ increases the share of out-of-wedlock births if it causes a greater relative decline in married mothers than in single mothers, $(-m\beta - (1-m)\alpha) / (m(b+\beta) + (1-m)(a+\alpha)) < (-a(1-\pi) - a\Delta P_{ij}) (a(1-p) + \alpha(1-\pi))$. A sufficient but not necessary condition for this inequality to hold is $(1 + m(b/a - 1)) / (1 + m(\beta/\alpha - 1)) \leq (1-p) / (1-\pi)$, which implies that the combined effect of women moving from Regions 2 and 3 to Region 1 in Figure (3) will weakly increase the rate of out-of-wedlock birth among the remaining mothers. Since the probability of single motherhood is smaller for inframarginal than for marginal non-traditional women, $(1-p) / (1-\pi) < 1$, this condition requires that the relative shift of traditional women into non-motherhood is larger than the corresponding shift for non-traditional mothers, $\beta/b > \alpha/a$. If this condition is not fulfilled, then the decline in fertility induced by the shock will depress the share of single mothers, but this effect still trades off against the rise in single motherhood due to the reduced marriage rate within Region 3, thus allowing the share of single motherhood to rise in the aggregate.

of-wedlock and raised in single-headed households.

This model, of course, ignores many salient considerations for fertility and marriage, including the option for women to seek abortions, and the reality that adults marry for reasons other than childrearing. Nevertheless, it captures a subtle mechanism by which shocks to male and female earnings produce distinct effects on fertility, marriage, and the prevalence of single-motherhood.

3 Data and Measurement

3.1 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.[17]

3.2 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, capital goods, and intermediate inputs (Hsieh and Klenow, 2009), and multinational enterprises being permitted to operate in the country (Naughton, 2007).[18] 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).

[17] 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).

[18] 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 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:

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

In this expression, $\Delta IP^{cu}_{j\tau} = \Delta M^{cu}_{j\tau}/(Y_{j0} + M_{j0} - X_{j0})$ 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 2007. 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, $Y_{j0} + M_{j0} - X_{j0}$) in the base year 1991, near the start of China's export boom. The fraction L_{ijt}/L_{it} is the share of industry j in CZ i's total employment, as measured in County Business Patterns data at the start of each period.

In (1), the difference in ΔIP^{cu}_{it} across commuting zones stems from variation in local industry employment structure at the start of period t, 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 less than 40 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 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_{ijt} and $1 - f_{ijt}$), thus apportioning the total CZ-level measure into two additive subcomponents, $\Delta IP^{m,cu}_{i\tau}$ and $\Delta IP^{f,cu}_{i\tau}$:

$$\Delta IP_{i\tau}^{m,cu} = \sum_j \frac{(1-f_{ijt})\,L_{ijt}}{L_{it}} \Delta IP_{j\tau}^{cu} \text{ and } \Delta IP_{i\tau}^{f,cu} = \sum_j \frac{f_{ijt}L_{ijt}}{L_{it}} \Delta IP_{j\tau}^{cu}, \tag{2}$$

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, $\Delta IP_{i\tau}^{cu} = 1.0$ 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 $\Delta IP_{i\tau}^{m,cu} = 0.45 \times 1.0 + 0.75 \times 1.0 = 0.6$ and $\Delta IPW_{uit}^{f} = 0.55 \times 1.0 + 0.25 \times 1.0 = 0.4$. 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 0.94 percentage points, 60 percent of which accrued to male employment and 40 percent to female employment. In the subsequent 2000 - 2007 period, when Chinese import penetration accelerated sharply, import penetration rose by an additional 1.33 percent, with 65 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.[19] Specifically, we instrument the measured import-exposure variable ΔIP_{it}^{cu} with a non-U.S. exposure variable ΔIP_{it}^{co} that is constructed using data on industry-level growth of Chinese exports to other high-income markets:

$$\Delta IP_{it}^{co} = \sum_j \frac{L_{ijt-10}}{L_{uit-10}} \Delta IP_{j\tau}^{co}. \tag{3}$$

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.[20] As documented by

[19]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.

[20]The start-of-period employment shares L_{ijt}/L_{it} and the gender shares f_{ijt} 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.

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.[21]

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.[22] 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, helping to isolate supply and trade-cost driven changes in China's exports. These gravity-based estimation results are quite similar to the IV approach that we employ in this paper.[23]

Data on international trade are from the UN Comtrade Database, which gives bilateral imports for six-digit HS products.[24] 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

[21]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 for the simple reason that 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.

[22]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.

[23]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.

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

(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 2007 using the PCE deflator. Data on CZ employment by industry from the County Business Patterns for 1990 and 2000 is used to compute employment shares by 4-digit SIC industries in (1) and (3).[25]

4 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.

4.1 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 21.8 percent and 12.9 percent respectively—that is, more than one in five young male workers and more than one in eight young female workers. These shares fell substantially in the ensuing two decades. By 2007, only 10.9 percent of men and 4.6 percent of women ages 18-39 worked in manufacturing (14.1 and 6.8 percent among those currently employed), corresponding to a fall of more than 35 percent among young men and more than 45 percent among young women.[26]

Although declining manufacturing employment was largely offset by gains in non-manufacturing employment—in net, employment-to-population fell by 2.0 percentage points among men and by

[25] 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.

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

roughly zero among women—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 log points higher than annual earnings of demographically comparable adults working in non-manufacturing in the year 2000. Approximately 60 percent of this annual earnings differential is attributable to higher annual hours among manufacturing workers, with the remaining 40 percent attributable to higher hourly earnings.[27] 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). The wage premia for male and female workers are similar in percentage terms, but larger for males in absolute value due to the overall higher wage level for males. The sectoral shift out of manufacturing may thus have eroded earnings especially for males, not only because men make up a disproportionate share of the sector's workforce, but also because the average male manufacturing job has a larger dollar wage premium than the average female manufacturing job.

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 2006 through 2008. Our focus is on employment of young adults because this population is disproportionately engaged in marriage and child-rearing.[28] We estimate (4) separately for the 1990s

[27]In untabulated results, we find that among men age 18-39 employed in manufacturing in 2000, 35% were in the top quartile of unconditional male annual wage and salary earnings in their commuting zone, another 33% were in the second quartile, and 32% were in the bottom *two* quartiles. The distribution of earnings in manufacturing was similarly skewed towards upper quartiles for women.

[28]Our sample is further restricted to individuals who are not residents of institutionalized group quarters such as prisons, and who are thus potential participants in the local labor and marriage markets.

and 2000s, and subsequently stack the ten-year equivalent first differences for 1990 to 2000 and 2000 to 2007, 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.

Table 1: OLS and 2SLS Estimates of the Relationship between Import Penetration and CZ-Level Manufacturing Employment, 1990-2007. Dependent Var: 100 x Change in Share of Population Age 18-39 Employed in Manufacturing (in % pts)

	OLS 1990-'00		OLS 2000-'07		OLS 1990-'07		2SLS 1990-'00		2SLS 2000-'07	
	(1)		(2)		(3)		(4)		(5)	
Δ Chinese Import Penetration	-0.65	*	-1.85	**	-1.44	**	-2.14	**	-2.54	**
	(0.27)		(0.14)		(0.17)		(0.43)		(0.18)	
2SLS First Stage Estimate	n/a		n/a		n/a		0.73	**	0.86	**
							(0.06)		(0.06)	
R^2							0.33		0.62	
	2SLS: 1990-'07									
	(6)		(7)		(8)		(9)		(10)	
Δ Chinese Import Penetration	-2.44	**	-2.64	**	-2.33	**	-2.32	**	-2.52	**
	(0.20)		(0.35)		(0.34)		(0.36)		(0.40)	
Manufacturing Emp Share$_{-1}$			Yes		Yes		Yes		Yes	
Census Division Dummies					Yes		Yes		Yes	
Occupational Composition$_{-1}$							Yes		Yes	
Population Composition$_{-1}$									Yes	
2SLS First Stage Estimate	0.82	**	0.60	**	0.62	**	0.60	**	0.59	**
	(0.05)		(0.05)		(0.05)		(0.05)		(0.06)	
R^2	0.55		0.60		0.61		0.63		0.63	

Notes: N=722 in columns 1-2 and 4-5, N=1444 (722 commuting zones x 2 time periods) in columns 3 and 6-10. All stacked first differences regressions in column 3 and 6-10 include a dummy for the 2000-2007 period. Occupational composition controls in columns 9-10 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 10 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. Robust standard errors in parentheses are clustered on state. Models are weighted by start of period commuting zone share of national population. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

The first panel of Table 1 presents initial results. As a point of comparison, the first two columns 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 -1.85 in columns 1 and 2 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 non-elderly adults employed in manufacturing during the 1990s, and a corresponding fall of 1.9 points for the 2000s. Column 3 stacks these two first differences, yielding an OLS point estimate of -1.44.

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 4 and 5 of 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 -2.5 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 five columns in Table 1 refine our approach and test robustness. Column 6 performs a stacked first-difference estimate, yielding a point estimate of -2.44. Columns 7 through 10 cumulatively add a rich vector of controls ($\mathbf{X}_{it}$ 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 7); Census division dummies, allowing for regional employment trends (column 8); occupational composition controls, accounting for employment in occupations susceptible to automation and offshoring (column 9); 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 10).[29] These controls, which we include in all subsequent regressions, have negligible effects

[29]Occupational controls in column 9 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 10 comprise the start-of-period shares of CZ population that

on the magnitude and the precision of the impact estimate. We estimate in the final column that a one percentage point rise in import penetration in a CZ causes a -2.52 percentage point change in CZ manufacturing employment as a share of adult population.[30]

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 0.74 percentage points per decade (Appendix Table A1). Multiplying the IQR by the column 10 impact estimate of -2.52 implies that rising trade exposure reduced manufacturing employment by 1.9 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 1.13 percentage points, implying that the mean CZ lost 3.35 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 3.1 percent per decade over the next seventeen years (bottom row Table 2).

The next set of results estimates the consequences of trade shocks on various employment and non-employment outcomes by sex—manufacturing employment, non-manufacturing 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 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 -2.60 and -2.40 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 Table 2).

The lower panel of Table 2 augments these specifications to include male- and female-specific trade exposure measures, each instrumented by contemporaneous changes in China's import pene-

are Hispanic, black, Asian, other race, foreign born, and college educated, as well as the fraction of women who are employed.

[30] 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.

tration to other high income countries during. Despite the relatively high correlation between the gender-specific shock measures ($\rho = 0.82$), there is abundant statistical power for distinguishing their independent effects on labor-market outcomes. The first set of estimates in the lower panel indicate that a one-percentage point rise in import penetration of either male or female-dominated industries reduces young adult manufacturing employment by approximately 2.5 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 5.0 percentage points ($t = -4.2$) and has a small and statistically insignificant impact on female manufacturing employment.[31] Conversely, a one-percentage-point shock to female-specific industries reduces employment of young women in manufacturing by 5.9 percentage points ($t = -5.1$), while having no measurable effect on male manufacturing employment. The fourth column in Table 2 combines these sex-specific outcomes by using as the dependent variable the male-female difference in manufacturing employment. It shows that a one-percentage-point rise in male-specific import penetration reduces the male-female manufacturing employment differential by 5.1 percentage points while a corresponding shock to female-specific import penetration raises this differential by 6.9 percentage points.[32]

Columns 5 through 8 provide a regression-based decomposition of how shocks to male-female differential in manufacturing employment net out across three other domains: non-manufacturing employment, 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 four exhaustive and mutually exclusive domains. The first row of estimates in the lower panel reveals that fully half of the fall in male relative to female manufacturing employment induced by the male shock (2.6 of 5.1 percentage points) accrues to a *rise* in male relative to female

[31] 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.

[32] Note that the column 4 coefficient for the effect of the male-specific shock on the male-female difference in manufacturing employment is equal by construction to the difference between the male-specific shock coefficients in columns 2 and 3 for male and female manufacturing employment separately ($-5.03 - 0.02 = -5.05$), while a parallel equality holds for the female-specific shock coefficient in column 4 ($0.94 - (-5.02) = 6.86$). Since the overall trade shock in the upper panel of the table reduces manufacturing employment almost evenly among males and females, it has only a modest impact on the male-female manufacturing employment differential ($-2.60 - (-2.40) = -0.19$, using rounded numbers).

non-manufacturing employment, while another 45% (2.3 of 5.1 percentage points) accrues to a rise in male relative to female non-participation (not in the labor force, abbreviated as NILF), with the remaining 5% accruing to a rise in the male-female unemployment gap. Thus, each one-point trade-induced fall in male relative to female manufacturing employment yields a half point net reduction in male relative to female labor-force participation. The second row of estimates documents parallel findings for female-specific trade shocks: 40% of the reduction in female relative to male employment in manufacturing accrues to rising female relative to male employment in non-manufacturing, while 55% accrues to non-participation, with only 5% to unemployment. 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 only partially offset by sectoral mobility, leading to large net reductions in employment. What these results add is a well-identified gender-specific dimension to employment shocks, which is crucial for the analysis that follows.

Table 2: 2SLS Estimates of the Impact of Import Penetration on Employment Status by Gender, 1990-2007. Dependent Var: 100 x Change in Share of Overall/Male/Female Population Age 18-39 Employed in Manufacturing (in % pts); 100 x Change in 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	Mfg	Non-Mfg	Unemp	NILF
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				I. Overall Trade Shock			
Δ Chinese Import Penetration	-2.52 **	-2.60 **	-2.40 **	-0.19	0.41	0.04	-0.26
	(0.40)	(0.47)	(0.36)	(0.29)	(0.34)	(0.17)	(0.34)
				II. Male Industry vs Female Industry Shock			
Δ Chinese Import Penetration × (Male Ind Emp Share)	-2.51 **	-5.03 **	0.02	-5.05 **	2.61 *	0.19	2.26 *
	(0.87)	(1.20)	(0.74)	(0.98)	(1.09)	(0.44)	(0.97)
Δ Chinese Import Penetration × (Female Ind Emp Share)	-2.54 *	0.94	-5.92 **	6.86 **	-2.77 *	-0.18	-3.91 **
	(1.10)	(1.39)	(1.16)	(1.37)	(1.31)	(0.58)	(1.32)
Mean Outcome Variable	-3.13	-3.86	-2.48	-1.38	-0.03	-0.06	1.46
Level in 1990	12.98	17.37	8.68	8.69	3.59	1.22	-13.50

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.

4.2 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. The conceptual model in Section 2 emphasizes the central role of the gender earnings gap in determining marriage market and fertility outcomes. Shocks that compress the gender earnings differential in a local market reduce the attractiveness of men as potential spouses, thus reducing fertility and especially marriage rates. The trade-induced decline in manufacturing employment may contribute to a reduction on the gender earnings gap since male manufacturing jobs earn higher industry wage premia than female manufacturing jobs. Figure 1 in the Introduction provides suggestive evidence for this link by documenting that lower manufacturing employment shares in CZs are correlated with the narrower gender gap in earnings. Whereas the prior tables document that gender-specific trade shocks to manufacturing only modestly reduce male relative to female manufacturing employment, our next set of analyses find a larger 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 2007 US\$) in commuting zone i in year t_0 at quantile u among CZ residents ages 25-39.[33] 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 $2000-2007$. Our estimating equation takes the form

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

where $Q_{\Delta y_{it}|,x_{it},\varepsilon_{it}}(u)$ is the u^{th} conditional quantile of Δy_{it} given $(\alpha_t, \Delta IP^{cu}_{it}, \mathbf{X}_{it}, \varepsilon_{it})$, α_t is an intercept, ΔIP^{cu}_{it} 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.[34]

[33] 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.

[34] 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.

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 index. In 1990, this gap was $\$6,247$, $\$12,065$, and $\$15,775$ 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 1990-2007, falling by $\$1,169$, $\$1,119$ and $\$1,696$ *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 25% from 1990 to 2007, versus 15% at the median and 18% at the 75^{th} percentile.

In the first row of estimates in Table 3, we find that trade shocks differentially reduce male relative to female earnings, and moreover, that this earnings reduction is substantially larger at low earnings quantiles. 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 $\$1,325$ at the 25^{th} percentile ($t = -5.9$) relative to a base of $\$7,226$. A trade shock of comparable size also significantly reduces male relative to female earnings at the median and 75^{th} percentiles, but the earnings impact is roughly only half as large ($613 and $695, respectively) and is proportionately much smaller relative to the baseline gap.

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 13.6 percentage points per decade relative to initial male earnings between 1990 and 2007. By comparison, the fall in the male-female differential at the median and 75^{th} percentiles was only about one-third 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

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.

find that a one-unit CZ-level import shock reduces the male-female earnings differential at the 25^{th} percentile by 17.4 percentage points ($t = -5.2$). By comparison, impacts on higher quantiles, while statistically significant, are only 10 to 15 percent as large in magnitude.

Table 3: 2SLS Estimates of the Impact of Import Penetration on Gender Differentials in Annual Earnings, 1990-2007. Dependent Var: Change in the Male-Female Annual Earnings by Percentile of CZ Earnings Distribution (in 2007 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	-1,325.3	**	-612.5	**	-694.6	**	-17.40	**	-2.22	*	-1.63	**
	(226.2)		(238.4)		(234.8)		(3.32)		(1.04)		(0.54)	
					II. Male Industry vs Female Industry Shock							
Δ Chinese Import Penetration × (Male Ind Emp Share)	-2,360.2	**	-2,860.3	**	-3,341.2	**	-26.62	**	-11.97	**	-8.61	**
	(668.5)		(807.0)		(976.1)		(8.33)		(3.20)		(2.08)	
Δ Chinese Import Penetration × (Female Ind Emp Share)	176.0		2,648.3	*	3,144.8	*	-4.03		11.91	**	8.51	**
	(940.3)		(1,072.0)		(1,298.9)		(12.85)		(4.45)		(2.89)	
Mean Outcome Variable	-1,169		-1,119		-1,696		-13.61		-4.31		-3.87	
Level of Male Earnings 1990	7,226		23,452		41,285		n/a		n/a		n/a	
Level of Female Earnings 1990	979		11,387		25,510							

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 2007 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. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

One hypothesis for why trade shocks have asymmetric effects on male versus female earnings is that manufacturing workers are disproportionately male, such that a rise in CZ import penetration differentially exposes men to earnings losses. This explanation seems unlikely, however, since we found above that these shocks do not cause a strong differential reduction in male relative to female manufacturing or non-manufacturing employment (Table 2, row 1, columns 4 - 8). The lower panel of Table 3 confirms that the underlying gender-bias of trade shocks is not the explanation. Gender-specific trade shocks of identical magnitude have substantially different impacts on male versus female earnings: a one-unit rise in import penetration in male-intensive manufacturing reduces male relative to female earnings at the 25^{th} percentile by \$2,360 ($t = -3.53$), whereas a one-unit rise in penetration in female-intensive manufacturing raises male relative to female earnings by a small and statistically insignificant \$176. At higher quantiles, however, these effects are 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 one-quarter (26.6%) relative to initial male earnings, whereas a one-unit import shock to female-intensive employment has no effect on this gap. At higher quantiles, male- and female-specific shocks have symmetric (and much smaller) impacts on the male-female wage gap expressed in terms of initial male earnings.

These patterns indicate that a simpler explanation for the asymmetry between the relative-employment and relative-wage effects of trade shocks is likely to be more apt. As shown in the the bottom row of Table 3, annual earnings at the 25^{th} percentile of the female earnings distribution in 1990 were extremely low—reflecting the fact that many women at this quantile are either non-participants or work a small number of hours annually, leaving little room for earnings to fall further. Average male earnings at this quantile were by comparison seven times as large (\$7,226 versus \$979), suggesting that a substantial fraction of males at this quantile in 1990 were gainfully employed. Although trade shocks reduce male and female employment equally, displaced females must primarily be drawn from higher quantiles of the female wage distribution (where female employment is concentrated), whereas displaced males are likely drawn from deeper down in the male-specific distribution. 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.

4.3 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), 1990-2007 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	0.42 *		-1.65 **	
	(0.17)		(0.50)	
Δ Chinese Import Penetration × (Male Ind Emp Share)		1.93 **		-2.87 **
		(0.60)		(0.90)
Δ Chinese Import Penetration × (Female Ind Emp Share)		-1.78 ~		0.13
		(0.96)		(1.35)
Mean Outcome Variable	1.88		1.70	
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 ≤ 0.10, * p ≤ 0.05, ** p ≤

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).[35] 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 over the next two decades. Regression estimates for the impact of trade shocks on this outcome measure confirm that rising trade exposure reduces the availability of males meeting the gender identity norm. The point estimate of 0.42 (t = 2.4) in column 1 of Panel A implies that a one-unit import shock reduces by about half a percentage point the fraction of

[35] 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.

demographically suitable young men who meet the gender identity norm for their female counterparts in the same CZ. Recalling that in aggregate, trade shocks have comparable impacts on male and female *employment*, this result again underscores the asymmetric effects of trade on male versus female *earnings*. 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 1.9 percentage points ($t = 3.2$) while a unit shock to female-intensive manufacturing raises male suitability by 1.8 percentage points ($t = -1.9$). Thus, both aggregate trade shocks and male-specific trade shocks decrease the supply of males meeting the gender-identity norm, while female-specific shocks have a partially offsetting effect.

4.4 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 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 unambiguously 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 1.7 points per decade between 1990 and 2007. The column 1 estimate finds that a unit import-penetration shock reduces the gender ratio by 1.7 points ($t = -3.3$), a unit male-specific shock reduces the gender ratio by 2.9 points ($t = -3.2$), and a unit female-specific shock has no detectable effect on the gender ratio ($\beta = 0.13, t = 0.1$). The fact that the *aggregate* trade shock differentially reduces the presence of young men relative to young women again underscores that trade-induced employment losses of comparable magnitude for men and women have distinct economic consequences for the two sexes.

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.[36] 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

[36] 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).

Deputy Under Secretary of Defense, 2011, Table 2.02). The OEMA data indicates applicants' county of residence, which we concord to commuting zones.[37] 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 10 percent of the1.65 percentage point drop in the virtual gender ratio per unit trade shock reported in Table 4.[38]

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 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. 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

[37]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.

[38]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 8 percent of the observed 1.65 point effect.

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.[39]

Table 5 present results.[40] The first row of estimates finds that shocks to import penetration have an insignificant positive effect on the male-female mortality gap among young adults. The point estimate of 4.27 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 four deaths per $100K$ adults. This effect is statistically insignificant ($t = 1.2$) and 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). Though these six categories must sum to the column 1 total, the precision of the estimates is much higher in finer subcategories. Column 2 shows that trade shocks significantly raise differential male mortality from drug and alcohol poisoning (D&A). The point estimate of 3.2 is large relative to start-of-period D&A mortality, which was equal to 6.4 among males and 1.9 among females. As Case and Deaton (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 5 document corresponding trade-induced increases in differential male mortality from liver disease (which is often alcohol-related), diabetes, and lung cancer.[41] Distinct from Pierce and Schott (2016b), we find no impact of trade shocks on suicides.

The lower panel of Table A4 re-estimates these 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 gener-

[39]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.

[40]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.

[41]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.

ally have larger effects, which explains why the pooled trade shock in the upper panel differentially raises male deaths from D&A poisoning, liver disease, diabetes and lung cancer. Appendix Table A4, which presents analogous estimates performed separately by sex, documents sharp male-female differences in mortality impacts of trade shocks. Contrary to men, the only mortality margin along which we detect a significant effect for women is suicide. In all other cases, the trade-induced rise on the male-female mortality differential is driven almost entirely by male responses.

Table 5: 2SLS Estimates of the Impact of Import Penetration on CZ-Level Gender Differences in Death Rates, 1990-2007. 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	4.27		3.16	*	0.72	*	0.62	~	0.40	*	0.01	0.08
	(3.54)		(1.35)		(0.32)		(0.32)		(0.17)		(0.99)	(3.08)
					II. Male Industry vs Female Industry Shock							
Δ Chinese Import Penetration × (Male Ind Emp Share)	18.84	~	10.11	**	1.68	~	1.38		0.30		1.12	3.13
	(11.28)		(2.80)		(0.91)		(0.91)		(1.00)		(3.22)	(9.89)
Δ Chinese Import Penetration × (Female Ind Emp Share)	-16.79		-6.93	~	-0.68		-0.49		0.54		-1.59	-4.34
	(17.46)		(3.61)		(1.33)		(1.69)		(1.53)		(5.17)	(17.94)
Mean Outcome Variable	-21.93		5.54		-0.73		0.20		-0.25		-1.28	-25.40
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. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

These estimates yield two conclusions. First, male deaths from adverse health behaviors are differentially affected by trade shocks: male-female differential death rates from drug and alcohol overdoses (D&A), liver damage, poor diet, and cigarette smoking rise in trade-exposed CZs. Second, the quantitative magnitude of these effects is too small to explain the bulk of the trade-induced fall in the male/female population ratio documented in the prior table.[42] However, only a small minority of

[42] A one-percentage-point increase in import penetration reduces male population by 4 for every $100K$ women. Cumulating over 10 years, this implies about 40 missing men per $100K$. Table 4 finds that a one-percentage-point increase in import penetration reduces the overall male population by 1.6 per 100 women ($1,600$ per $100K$), which is 40 times as large as the mortality effect.

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, it does not allow a quantification 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.[43] 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.[44] 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.[45] 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

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

[44] 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.

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

homelessness.[46] 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.

5 The Labor Market and the Marriage Market

Though trade shocks have comparable effects on young men's and women's aggregate employment rates, the constellation of evidence so far demonstrates that these shocks differentially reduce males' marriage-market value along multiple margins: relative earnings, physical presence in trade-impacted labor markets, and possibly participation in risky, abusive, and illegal activities. Our simple conceptual model above makes a number of specific predictions about how shocks to the marriage-market value of men will affect marriage, fertility, and the composition of births and family living arrangements. We test these predictions in this section, beginning with the prevalence of marriage.

5.1 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.[47] Table 6 presents results.

Panel A of the table estimates a series of linear probability models that test 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.[48] The first row finds that aggregate trade shocks (i.e., not distinguishing between gender components) reduce the

[46] 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%).

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

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

prevalence of marriage: a one-unit aggregate trade shock modestly reduces the fraction of young women who are currently married by about 0.7 percentage points on a base of 53% (column 3). As predicted by the option-value 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; 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.

Table 6: 2SLS Estimates of the Impact of Import Penetration on Marital Status of Young Women, 1990-2007. 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. Marital Status (% pts): Women Ages 18-39			B. Marital Status (% pts): Women Age 18-25			C. Pct of Mothers Currently Married	
	Never Married	Widowed Divorced Separated	Married	Never Married	Widowed Divorced Separated	Married	Age 18-39	Age 18-25
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				I. Overall Trade Shock				
Δ Chinese Penetration	0.44	0.28 ~	-0.72 *	1.03 ~	0.19 ~	-1.22 *	-0.76 *	-1.01
	(0.30)	(0.15)	(0.34)	(0.53)	(0.11)	(0.50)	(0.31)	(0.77)
				II. Male Industry vs Female Industry Shock				
Δ Chinese Penetration × (Male Share)	1.77 ~	-0.22	-1.55 ~	3.90 **	-0.06	-3.84 **	-1.56 ~	-3.57 ~
	(0.91)	(0.41)	(0.88)	(1.25)	(0.43)	(1.18)	(0.81)	(2.04)
Δ Chinese Penetration × (Female Share)	-1.50	1.01 ~	0.49	-3.12 ~	0.55	2.57	0.39	2.70
	(1.35)	(0.58)	(1.39)	(1.60)	(0.59)	(1.66)	(1.32)	(3.53)
Mean Outcome Var	8.62	-1.49	-7.14	9.00	-1.32	-7.69	-5.14	9.59
Level in 1990	34.84	12.11	53.05	67.30	4.96	27.74	76.02	61.23

Notes: N=1444 (722 CZ x 2 time periods). Columns 3 and 6 refer to the percentage of women in the indicated age group who report to be married but not separated. 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.

Panel B of Table 6 tightens the focus to a narrower age bracket of young women, those ages 18-25, who are more likely to be new entrants to the marriage market. As shown in the bottom row of the table, women in this age group are only half as likely to be currently married, and twice as likely to be never married, as are those in the broader 18-39 age bracket. Trade shocks have larger impacts on marital status among women ages 18-25. A one-unit shock reduces the fraction of young women currently married by 1.2 percentage points, raises the fraction never married by 1.0 percentage point, and increases the fraction who are widowed, divorced or separated already at a young age by 0.2 percentage points. All point estimates are significant with $p < 0.10$. The lower two rows of the panel B, which disaggregate the trade shock into gender-intensity components, reveal even larger effects on marriage formation. A one-unit shock to male-intensive employment raises by 3.9 percentage points the fraction of young women ages 18-25 who have not entered into marriage. Conversely, a one-unit shock to female-intensive employment reduces the fraction of never-married young women by 3.1 percentage points.

Our conceptual model makes a further prediction regarding the interaction between labor and marriage markets: adverse shocks to the supply of marriageable males should *raise* the fraction of mothers who are unmarried while adverse shocks to women's outside (i.e., non-marital) economic opportunities should *lower* this share. Panel C confirms this prediction. Shocks to male-intensive employment significantly boost the fraction of mothers who are single, particularly among the youngest women (ages 18-25), while shocks to female-intensive employment diminish the share of mothers who are not married, though this latter estimate is imprecise. We again find that aggregate manufacturing shocks—not distinguishing by gender—generate the same sign pattern as shocks to male-intensive industries, though with smaller magnitude. Consistent with the panoply of evidence above, trade shocks appear detrimental to the marriage-market value of men.

5.2 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 (also reflected in our conceptual model) 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 - 2007.

The first column of Table 7 considers the impact of trade shocks on births per thousand adult women ages 20-39.[49] The estimate in the upper panel finds that a one-unit trade shock reduces total fertility by approximately 3.3 births per thousand women, which is roughly a 4-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 8.2 per thousand women ($t = -3.9$) while an equivalent shock concentrated on female employment is found to raise births by 3.7 per thousand women, though this latter effect is not precisely estimated.

At the start of our sample in 1990, the U.S. teen birth rate was nearly 75 percent as high as the corresponding adult (ages 20-39) rate, but it fell rapidly between 1990 and 2007. Based on the column 1 estimate, one might speculate that trade shocks contributed to the decline in teen births. The column 2 estimate, which considers births to women ages 15-19, provides only limited support for this mechanism. A one-unit trade shock reduces teen births by 1.2 per thousand women ($t = -1.7$), an effect that is only one-third as large as the corresponding impact for adult women (and one half as large in proportional terms). The smaller proportional fall in teen relative to adult births suggests that the share of births accruing to teens rises in trade-impacted labor markets. The column 3 estimate confirms this implication: a one-unit shock raises the teen birth share by 0.6 percent, which is a five-percent increase relative to baseline. 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 4).[50]

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). In line with this argument, the evidence in the lower panel of Table 7 shows that the female-specific shock has a sizable but imprecisely estimated positive impact on teen fertility. Since adult fertility also expands, the female shock however lowers the fraction of

[49]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.

[50]In terms of how these results relate to our conceptual model, it is natural to think of teen births as occurring out of wedlock. A fall in the marriage-market value of men would thus be expected to reduce births among teens by less than among adult women, for whom marriage rates are much higher. By this reasoning a reduction of in-wedlock births should primarily reduce non-teen births, consistent with our findings.

teen births in overall births, and produces an even greater decline in the proportion of out-of-wedlock births. Consistent with our simple model, these consequences of the shock to female labor-market opportunities provide a mirror image to the impacts of the male-specific shock, which reduces overall fertility while raising the share of out-of-wedlock births.

Table 7: 2SLS Estimates of the Impact of Import Penetration on Births to Young Women, 1990-2007. Dependent Var: 100 x Change in Birth Rate, Share of Births to Teen Mothers, Share of Births to Unmarried Mothers (in %pts).

Table 7. Imports from China and Birth Outcomes, 1990-2007: 2SLS Estimates.
Dep Var: 100 x Change in Birth Rate or Share of Births with Indicated Conditions (in %pts).

	Births per 1,000 Women				Share of Births to			
	Adults Age 20-39		Teens Age 15 - 19		Teenage Mothers		Unmarried Mothers	
	(1)		(2)		(3)		(4)	
	I. Overall Trade Shock							
Δ Chinese Import Penetration	-3.30	**	-1.22	~	0.63	**	0.76	~
	(0.51)		(0.72)		(0.17)		(0.40)	
	II. Male Industry vs Female Industry Shock							
Δ Chinese Import Penetration × (Male Ind Emp Share)	-8.16	**	-5.30	**	1.92	**	3.70	**
	(2.10)		(1.98)		(0.53)		(0.95)	
Δ Chinese Import Penetration × (Female Ind Emp Share)	3.74		4.71		-1.25	~	-3.51	*
	(3.00)		(3.05)		(0.69)		(1.45)	
Mean Outcome Variable	3.86		-11.08		-1.44		7.74	
Level in 1990	86.9		60.0		12.8		27.7	

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

5.3 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 2007 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 2.2 percentage points ($t = 5.1$), which is roughly a 12% increase in the prevalence of childhood poverty relative to the base poverty rate of 18.5% of U.S. children in 1990. Columns 2 through 5 show that an aggregate one-unit trade shock reduces the fraction of children in two-parent households by 0.44 percentage points, about half of which is accounted for by a rise in the fraction living in single-headed households, and the other half due to a rise in children living with grandparents.

Since poverty is far more prevalent among single-headed than married households (bottom row of Table 8), one might surmise that the substantial effect of trade shocks on the prevalence of childhood poverty (column 1) stems primarily from the rise in single-headedness. Surprisingly, this is not 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 four household types above in Appendix Table A5, we find that most of the rise in poverty is *within* household type. For example, a one-unit overall trade shock reduces the fraction of children living in non-poor two-parent households by 1.39 points but raises the fraction living in poor two-parent households by 0.99 points, which sums to the modest decline of children in two-parent households by 0.44 percentage points that is reported in Table 8. Similarly, a one-unit trade shock strongly reduces the fraction of children living in non-poor single-headed households by 0.65 points, while increasing the fraction living in poor single-headed households by 0.89 points. Thus, the trade-induced rise in childhood poverty is not primarily accounted for by rising single-headedness but rather by pervasive income losses across all household types.

Table 8: 2SLS Estimates of the Impact of Import Penetration on Socio-Economic Outcomes of Children Age 0 - 17, 1990-2007. Dependent Var: 100 x Change Share of Age Group in Indicated Condition (in %pts)

	Income < Poverty Line		Parent Head, Spouse Present		Parent Head, Spouse Absent		Grand-Parent Headed		Any Other Person Headed
	(1)		(2)		(3)		(4)		(5)
					I. Overall Trade Shock				
Δ Chinese Import Penetration	2.17	**	-0.44	~	0.22		0.24	~	-0.01
	(0.42)		(0.25)		(0.23)		(0.13)		(0.13)
					II. Male Industry vs Female Industry Shock				
Δ Chinese Import Penetration × (Male Ind Share)	3.99	**	-1.00		1.98	**	-0.77	~	-0.20
	(0.85)		(0.62)		(0.59)		(0.40)		(0.30)
Δ Chinese Import Penetration × (Female Ind Share)	-0.48		0.38		-2.34	*	1.70	*	0.26
	(1.28)		(1.01)		(1.06)		(0.67)		(0.42)
Mean Outcome Variable	0.51		-4.98		3.91		0.56		0.50
Mean Level in 1990	0.17		71.43		19.59		5.43		3.55
Poverty Rate (%) in 1990	n/a		8.13		45.26		23.99		34.73

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 almost the entirety of the impact of adverse economic shocks on child poverty arises from shocks to male employment; we find no statistically or economically significant effect of female-specific labor-market shocks 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 two-parent households (by 3.3 percentage points) while raising the fractions of children in poor two-parent and poor single-parent homes (by 2.3 and 1.8 percentage points, respectfully). In conjunction with our earlier results on fertility, and reasoning through the lens of our simple model, this finding suggests that adverse shocks to *male* earnings disadvantage women and children along several margins: raising the fraction of children born to teen and unmarried mothers, and expanding the share of married mothers who are married to low-quality (i.e., low earnings) men.[51] The complementary finding that adverse shocks to *female* employment have no measurable impact on the prevalence of childhood poverty is also consistent with our model and the findings above wherein the direct negative effect of lower female earnings

[51] Recall that all else equal, the share of non-marital births falls as the fraction of high-quality men P rises.

on household income trades off against a favorable compositional shift as the rate of out-of-wedlock birth falls, and children are less likely to be in single-parent homes that by far have the highest poverty risk among all household types. Though imprecise, the estimates in Appendix Table A5 lend weight to this interpretation: adverse shocks to female earnings opportunity reduce the fraction of children living in non-poor single-parent households by much more than the fraction living in poor single-parent homes, suggesting that the female-specific shock *increases* child poverty conditional on being in a single-parent home. However, the overall decline in the fraction of children living in single parent homes is to a large extent balanced by an increasing fraction living in non-poor two-parent or grandparent-headed homes, an effect that contributes to a *decrease* in child poverty.

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. Appendix Table A6 explores this question and documents that the causal chain running from the labor market to the marriage market to the prevalence of non-marital births, single-headedness, and childhood poverty appears to hold consistently across all three major demographic groups in our data—white Non-Hispanics, blacks, and Hispanics—though, given often much smaller sample sizes, precision is far lower when the data are split into these subpopulations.

6 Economic Magnitudes: Some Simple Calculations

Our results document a causal chain running from adverse shocks to male earnings capacity to declining marriage prevalence among young adults, a rising fraction of births accruing to teenage and unmarried mothers, and an increasing share of children living in single-headed and impoverished households. To assess the economic magnitudes of these effects, we consider some simple benchmarks, focusing on six outcome variables that feature in the analysis: the prevalence of marriage among women age 18-25 and 18-39; the share of U.S. births to unmarried mothers and to teenagers age 15-19; and the fraction of children age 0-17 living in below-poverty households and in parent-headed, spouse-present (i.e., married) households. Many of these outcomes changed substantially in our analytic interval, as summarized in panel A of Table 9. The fraction of young women ages 18-25 and 18-39 currently married fell by 12 and 11 percentage points respectively, the fraction of births out-of-wedlock rose by 13 percentage points, the fraction of children living in two-parent married families dropped by 8.3 percentage points, and the share of births to teenagers fell by 2.3 percentage points (an 18 percent drop). The fraction of children living in poor households was largely unchanged, however, declining in the 1990s and then rising in the 2000s.

Panel B of Table 9 provides a first benchmark, quantifying the effect of the 'China Shock' on marriage outcomes and children's birth and living circumstances between 1990 and 2007. This panel reports the estimated reduced-form impact of the 'China Shock´ on each of the six outcome measures, scaled by the instrumented change in trade exposure between 1990-2007, as well as the counterfactual change in each outcome variable when setting the change in China trade exposure to to zero.[52] In presenting these numbers, we stress that this paper's intellectual target is to assess the causal effect of secular changes in the labor-market stature of men and women on marriage, fertility, and children's living circumstances; we harness the China trade shock primarily for causal variation rather than for accounting purposes. Nevertheless, its impacts are directly quantifiable from our regression analysis, and we report them here.

The panel B estimates indicate that the China trade shock modestly but meaningfully altered patterns of marriage, fertility, and family structure: reducing marriage prevalence among young women by 0.75 to 1.25 percentage points; raising the share of teen and non-marital births by approximately two-thirds of a percentage point each, decreasing the fraction of children living in married two-parent households by roughly a half percentage point, and substantially hiking—by 2.2 percentage points, a 13 percent increase—the fraction of children living in poverty. The inference that we draw, consistent with our priors, is that rising China trade is a contributor to, but *not* the primary driver of the broader demographic shifts on which we focus.

[52]Each estimate is the product of three terms: the reduced-form coefficient for the impact of a one-unit China trade shock on the relevant outcome, taken from earlier tables; the observed rise in aggregate China trade exposure during 1990-2007, reported in Appendix Table A1; and a scaling coefficient of 0.55, corresponding to the R-squared of the first-stage regression of the observed China exposure measure by on the instrument (column 6 of Table 1)—that is, the portion of observed China import penetration that we can attribute to the exogenous supply-shock component.

Table 9: Implied Impacts of (1) the 'China Shock´ and (2) the Decline in Male-Female Relative Earnings on the Prevalence of Marriage, the Share of Births to Teens and Unmarried Mothers, the Fraction of Children Living in Poverty and in Two-Parent Households

Table 9. Implied Impacts of (1) the `China Shock´ and (2) the Decline in Male-Female Relative Earnings on the Prevalence of Marriage, the Share of Births to Teens and Unmarried Mothers, the Fraction of Children Living in Poverty and in Two-Parent Households

	% Women 18-25 Married (1)	% Women 18-39 Married (2)	% Births to Teens (3)	% Births Out of Wedlock (4)	% Children 0-17 in Poor HHs (5)	% Children 0-17 in Two-Parent HHs (6)
			A. Summary Statistics			
Δ 90-07: Actual	-11.86	-11.02	-2.33	12.58	0.25	-8.25
1990 Level	27.74	53.05	12.77	27.71	17.99	71.43
			B. Implied Impact of `China Shock´			
Δ 90-07: Implied impact	-1.26	-0.74	0.64	0.78	2.23	-0.45
Δ 90-07: Counterfactual	-10.60	-10.28	-2.97	11.80	-1.98	-7.80
			C. Using Observed ΔMale-Female P50 Annual Earnings Gap as Explanatory			
Δ 90-07: Implied impact	-3.63	-2.13	1.86	2.25	6.44	-1.30
Δ 90-07: Counterfactual	-8.22	-8.89	-4.18	10.33	-6.19	-6.96

This table reports levels and observed and counterfactual changes for six outcome variables analyzed in the main analysis. Panel A reports the 1990 level and the 1990-2007 change in each outcome. Panel B reports the reduced form impact of the `China Shock´ on each outcome and the counterfactual change in that outcome while setting the China shock to zero. These calculations correspond to the product of (1) the reduced form coefficient for the relevant outcome, (2) the observed China shock value reported in Appendix Table 1 and (3) a scaling coefficient of 0.55, corresponding to the R-squared of the first stage regression of the observed China exposure measure on the predicted exposure instrument (as explained in the text). Panel C reports counterfactual calculations that treat the change in the male-female P50 annual earnings gap as the hypothetical forcing variable through which the China Shock affects other outcomes. This gap, measured in real 2007 dollars, fell by $1,820 between 1990 and 2007. The estimates in Table 3 imply that the exogenous component of the China trade shock caused a decline of $631 in the male-female P50 gap. Interpreting the reduced form estimates in panel B as corresponding to the causal effect of a $631 fall in the male-female P50 gap, we can rescale the panel B impact estimates by 1,820/631=2.88 to estimate the implied effect of the overall decline in the male-female P50. This calculation is reported in panel C.

We next use these reduced-form causal relationships to generate a more ambitious set of impact estimates. This further step requires us to take a stand on which intermediating variable is the forcing channel through which trade-shocks move other demographic margins. A logical candidate for a forcing variable is the male-female annual earnings gap, analyzed in Table 5. This metric incorporates changes along both earnings and hours margins, and is arguably our best omnibus measure of how trade shocks affect the relative economic stature of young men and women. We focus on the $P50$ gap since it is representative. The $P50$ gap, measured in real 2007 dollars, fell by $\$1,820$ between 1990 and 2007, a 15-percent decline on its 1990 base of $\$12,065$. The estimates in Table 5 imply that the exogenous component of the China trade shock explains slightly more than one-third of this decline ($631).[53] If we interpret the reduced-form estimates in panel B as measuring the causal effect of a $631 fall in the male-female $P50$ gap, we can rescale the panel B impact estimates by 1,820/631 = 2.88 to infer the implied effect of the overall decline in the male-female $P50$. These simple calculations, reported in panel C, imply that the drop of $\$1,820$ in the male-female $P50$ gap between 1990 and 2007 reduced the prevalence of marriage by 2.1 to 3.6 percentage points (equal to 20 to 30 percent of the observed change, similar to the estimate obtained by Shenhav (2016) using a distinct methodology), raised the teen and out-of-wedlock birth shares by 1.9 and 2.3 percentage points, increased the fraction of children living in poverty by an impressive 6.4 percentage points, and lowered the share of children in two-parent married households by 1.3 percentage points. These are substantial effects. But they should of course be interpreted with care. Our calculations imply that, ceteris paribus, the falling gender wage gap should have raised the teen birth share between 1990 and 2007. In reality, this share fell sharply in this time interval, driven in part by the influence of social media (see Kearney and Levine, 2015). Although the evolving labor market has altered marriage, fertility, and children's living circumstances during this time period, it is not the only consequential force at play.

7 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 both male and female employment; a reduction in men's relative earnings, particularly at the lower tail of the earnings distribution; an increase in the rate of male mortality from risky and unhealthful behaviors; 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 rise in the fraction of births to teen and unmarried mothers; and a sharp jump in the fraction of children living in impoverished and, to a lesser degree, single-headed households.

Two cardinal results help to weave these many empirical strands together. A first is that trade

[53] Specifically, a one-unit China trade shock is estimated to compress this gap by $613. The combined trade shock for 1990-2007, adjusted by the scaling factor of 0.55 is equal to 1.03, implying a causal effect of $\$631.4 = \$613 \times 1.03$.

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. A second broad result, predicted by our model and strongly affirmed by the data, is that gender-specific shocks to labor-market outcomes have strikingly non-parallel impacts on marriage-market outcomes. 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 have more modest effects on overall fertility but reduce the share of births to teens and unmarried mothers, thus raising in-wedlock births and reducing the fraction of children living in single-headed households. These patterns are consistent with our model in which a decline in the quality of male partners makes single motherhood a more attractive option to young mothers, while a decline in female earnings potential increases marriage rates conditional on fertility. Netting over the effects of secularly falling male earnings and improving women's labor-market conditions during recent decades, our model predicts a reduction in both fertility and marriage, a rise in the fraction of children born out of wedlock, and an increase in the prevalence of children living in single-headed and poor households. These patterns are evident in the aggregate data and, moreover, hold as causal relationships within local labor markets when we isolate plausibly exogenous shocks to earnings opportunities overall and by gender.

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)

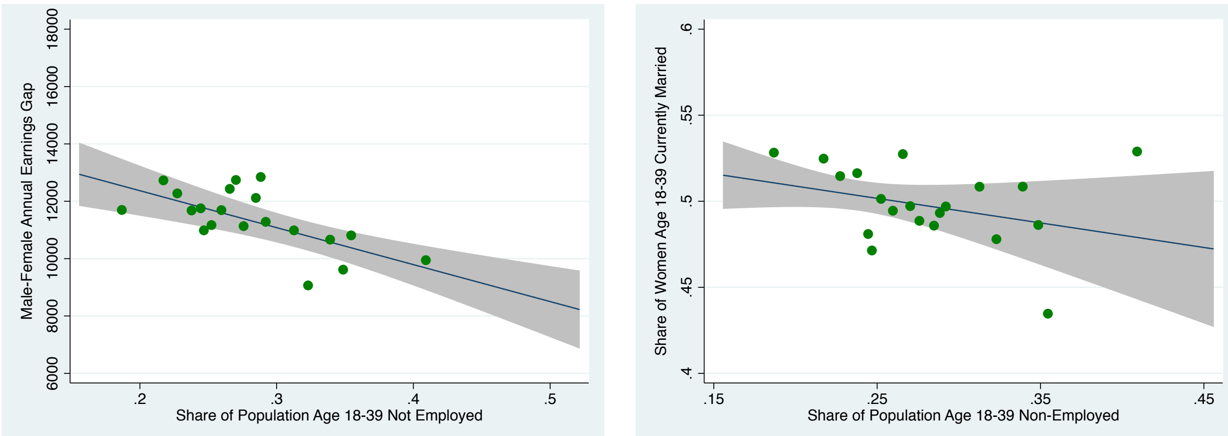

Table A1: Mean and Percentiles of Decadal Growth in Import Penetration, Overall and Sex-Specific Measures

	1990-2007 (1)	1990-2000 (2)	2000-2007 (3)
		I. Overall Shock	
Mean	1.13	0.94	1.33
	(0.75)	(0.61)	(0.83)
P25	0.68	0.54	0.83
P50	0.95	0.88	1.14
P75	1.43	1.22	1.59
P75-P25	0.74	0.68	0.76
		II. Male Industry Shock	
Mean	0.71	0.56	0.86
	(0.47)	(0.33)	(0.53)
P25	0.43	0.35	0.54
P50	0.60	0.53	0.77
P75	0.90	0.73	1.09
P75-P25	0.47	0.38	0.56
		III. Female Industry Shock	
Mean	0.42	0.39	0.46
	(0.32)	(0.31)	(0.33)
P25	0.25	0.21	0.27
P50	0.37	0.34	0.39
P75	0.52	0.48	0.54
P75-P25	0.27	0.27	0.28

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)	(1)	(2)	(3)	(4)
	I. Annual Wage and Salary Income (US$)				II. Log Annual Wage and Salary Income			
Male x Employed in Manufacturing	1334 **	1890 **	1879 **	2153 **	0.18 **	0.19 **	0.19 **	0.20 **
	(291)	(277)	(265)	(268)	(0.02)	(0.02)	(0.01)	(0.01)
Female x Employed in Manufacturing	508	1464 **	1483	1764 **	0.13 **	0.20 **	0.20	0.21 **
	(319)	(273)	(268)	(280)	(0.02)	(0.02)	(0.02)	(0.02)
Mean (S.D.) Outcome Var Males / Females	20,781 (23,171) 15,518 (19,795)				9.12 (1.28) 8.79 (1.29)			
	III. Log Annual Work Hours				IV. Log Hourly Wage			
Male x Employed in Manufacturing	0.13 **	0.13 **	0.13 **	0.12 **	0.05 **	0.06 **	0.06 **	0.08 **
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Female x Employed in Manufacturing	0.12 **	0.14 **	0.14	0.13 **	0.02	0.06 **	0.06	0.07 **
	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Mean (S.D.) Outcome Var Males / Females	6.79 (0.99) 6.60 (1.03)				2.34 (0.76) 2.19 (0.77)			
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. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$

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, 1990-2007. 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	3.67		4.28	*	0.64	*	0.31		0.33	**	0.72		-0.24
	(3.47)		(1.87)		(0.32)		(0.27)		(0.12)		(0.87)		(2.45)
	Male Industry vs Female Industry Shock												
Δ Chinese Penetration × (Male Ind Share)	16.61		11.33	**	1.12		1.83	**	0.28		-0.73		1.98
	(12.85)		(3.97)		(0.70)		(0.65)		(0.47)		(2.33)		(11.53)
Δ Chinese Penetration × (Female Emp Share)	-15.06		-5.96		-0.05		-1.88	~	0.41		2.81		-3.44
	(15.28)		(4.99)		(1.04)		(1.11)		(0.69)		(3.74)		(16.55)
Mean of Outcome	-25.71		10.27		0.19		-1.26		-0.53		-1.42		-32.95
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	2.26		1.05		-0.12		-0.30		-0.07		0.81	*	1.69
	(2.51)		(0.78)		(0.20)		(0.22)		(0.15)		(0.37)		(2.05)
	Male Industry vs Female Industry Shock												
Δ Chinese Penetration × (Male Ind Share)	-3.97		0.46		-0.55		0.49		-0.05		-2.04		-1.86
	(7.72)		(2.23)		(0.64)		(0.63)		(0.69)		(1.43)		(6.23)
Δ Chinese Penetration × (Female Emp Share)	11.31		1.89		0.51		-1.43		-0.11		4.94	*	6.85
	(11.12)		(3.55)		(1.12)		(1.01)		(1.12)		(2.09)		(9.32)
Mean of Outcome	-3.79		4.73		-0.01		-0.53		-0.28		-0.14		-7.55
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, 1990-2007 (2SLS Estimates). Dependent Var: 100 x Change Share of Children Living in Each Household Type and Poverty Status

	All HH's (1)		Parent Head, Spouse Present (2)		Parent Head, Spouse Absent (3)		Grand-Parent Headed (4)		Any Other Person Headed (5)	
	A. Percent in Household Type w/ Income < Poverty Line									
	A1. Overall Trade Shock									
Δ Chinese Penetration	2.17	**	0.99	**	0.89	**	0.19	**	0.10	~
	(0.42)		(0.23)		(0.22)		(0.06)		(0.05)	
	A2. Male Industry vs Female Industry Shock									
Δ Chinese Penetration × (Male Ind Share)	3.99	**	2.33	**	1.76	**	-0.07		-0.03	
	(0.85)		(0.52)		(0.50)		(0.19)		(0.16)	
Δ Chinese Penetration × (Female Ind Share)	-0.48		-0.97		-0.36		0.56		0.30	
	(1.28)		(0.81)		(0.75)		(0.35)		(0.26)	
	B. Percent in Household Type w/ Income >= Poverty Line									
	B1. Overall Trade Shock									
Δ Chinese Penetration	-2.17	**	-1.39	**	-0.65	**	0.06		-0.19	*
	(0.42)		(0.33)		(0.19)		(0.11)		(0.09)	
	B2. Male Industry vs Female Industry Shock									
Δ Chinese Penetration × (Male Ind Share)	-3.99	**	-3.25	**	0.28		-0.70	*	-0.32	
	(0.85)		(0.54)		(0.58)		(0.33)		(0.20)	
Δ Chinese Penetration × (Female Ind Share)	0.48		1.32		-2.00	*	1.15	*	-0.01	
	(1.28)		(1.09)		(0.87)		(0.50)		(0.22)	

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. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Table A6: 2SLS Estimates of the Impact of Import Penetration on Employment Status by Race and Sex, 1990-2007. Dependent Var: 100 x Change in Male-Female Differential in Fraction of Population Age 18-39 that is Employed, Unemployed or Non-Employed (in % pts).

	A. Labor Market		B. Marriage Market		C. Fertility		D. Children's Environment	
	M-F Gap Emp Mfg	M-F Gap Median Earn	M/F Gender Ratio	% Female Never Marr	Births/1000 Women	Unmarried Women	% Child Poverty	% Child w/ Single Parent
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				Ia. Non-Hispanic Whites: Overall Trade Shock				
Δ Chinese Penetration	-0.30	-1,100	-1.44 *	1.18 **	-1.09	0.54 ~	1.76 **	0.35
	(0.32)	(688)	(0.60)	(0.44)	(1.17)	(0.29)	(0.32)	(0.22)
				Ib. Non-Hispanic Whites: Male Industry vs Female Industry Shock				
Δ Chinese Penetration	-5.61 **	-4,138 **	-2.38 ~	2.87 *	-6.33 *	3.23 **	3.12 **	0.93 *
(Male Ind Emp, Whites)	(0.82)	(1,284)	(1.24)	(1.22)	(2.58)	(0.86)	(0.65)	(0.43)
Δ Chinese Penetration	7.87 **	3,726 **	0.00	-1.41	6.97	-3.59 **	-0.33	-0.55
(Female Ind Emp, Whites)	(1.19)	(1,189)	(1.54)	(1.83)	(5.24)	(1.34)	(1.01)	(0.67)
				IIa. Blacks: Overall Trade Shock				
Δ Chinese Penetration	0.30	304	-2.59	1.45	-1.57	-0.10	4.14 **	0.93
	(0.59)	(769)	(1.63)	(0.89)	(1.08)	(0.72)	(1.43)	(0.94)
				IIb. Blacks: Male Industry vs Female Industry Shock				
Δ Chinese Penetration	-4.76	2,528	-0.03	-0.19	-6.24	-4.18	6.50	7.73 ~
(Male Ind Emp, Blacks)	(3.10)	(2,154)	(6.02)	(4.35)	(6.26)	(3.40)	(4.77)	(4.36)
Δ Chinese Penetration	6.92 ~	-2,774	-5.93	3.58	4.54	5.25	1.05	-7.95
(Female Ind Emp, Blacks)	(4.06)	(3,204)	(8.59)	(5.24)	(8.69)	(5.64)	(5.40)	(5.40)
				IIIa. Hispanics: Overall Trade Shock				
Δ Chinese Penetration	-0.29	-898	-4.99 *	1.83	-4.92	0.64	2.64 **	-0.22
	(0.74)	(848)	(2.46)	(1.20)	(3.94)	(1.37)	(0.83)	(1.16)
				IIIb: Hispanics: Male Industry vs Female Industry Shock				
Δ Chinese Penetration	-4.34	-685	2.38	1.13	-11.57	4.53	-4.44	-0.23
(Male Ind Emp, Hispanics)	(3.94)	(1,516)	(11.22)	(3.99)	(9.28)	(3.60)	(4.30)	(0.30)
Δ Chinese Penetration	4.53	-1,251	-13.78	2.66	3.00	-3.99	11.08 ~	-0.01
(Female Ind Emp,	(4.66)	(3,686)	(11.43)	(5.43)	(13.85)	(4.74)	(6.00)	(0.02)

Notes: N=1444 (722 CZ x 2 time periods). Panel I analyzes outcomes within the non-Hispanic white population, panel II within the non-Hispanic black population and panel III within the Hispanic population. Gender-specific import shocks are constructed using the gender composition in the indicated race/ethnicity group within each industry-CZ cell. The outcomes are computed for adults age 18-39 (columns 1-3), age 18-25 (column 4) and age 20-39 (columns 5-6), as well as for children age 0-17 (columns 7-8). All regressions include the full set of control variables from Table 1. Regressions are weighted by a CZ's share in national start-of-period population of 18-39 year olds within the respective race and ethnicity group (non-Hispanic whites in panel I, non-Hispanic blacks in panel II, Hispanics in panel III). Standard errors are clustered on state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.