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When Work Disappears: Manufacturing Decline and the Falling Marriage Market Value of Young Men†

By David Autor, David Dorn, and Gordon Hanson*

We exploit the gender-specific components of large-scale labor demand shocks stemming from rising international manufacturing competition to test how shifts in the relative economic stature of young men versus young women affected marriage, fertility, and children's living circumstances during 1990–2014. On average, trade shocks differentially reduce employment and earnings of young adult males. Consistent with Becker's model of household specialization, shocks to males' relative earnings reduce marriage and fertility. Consistent with prominent sociological accounts, these shocks heighten male idleness and premature mortality, and raise the share of mothers who are unwed and the share of children living in below-poverty, single-headed households. (*JEL* F16, J12, J13, J16, J23, J31, L60)

The consequences of high neighborhood joblessness are more devastating than those of high neighborhood poverty … Many of today's problems in the inner-city ghetto neighborhoods—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, p. 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 a Family and Culture in Crisis* (2016, p. 144)

*Autor: Department of Economics, Massachusetts Institute of Technology, Building E52, Room 438, 77 Massachusetts Avenue, Cambridge, MA 02142, and NBER (email: dautor@mit.edu); Dorn: Department of Economics, University of Zürich, Schoenberggasse 1, CH-8001, Zürich, Switzerland, and CEPR (email: david.dorn@econ.uzh.ch); Hanson: University of California, San Diego, 9500 Gilman Drive, La Jolla, CA 92093-0519, and NBER (email: gohanson@ucsd.edu). Pete Klenow was the coeditor for this article. This paper previously circulated under the title "The Labor Market and the Marriage Market" (first circulating draft May 12, 2014). Autor, Dorn, and Hanson acknowledge funding from the Russell Sage Foundation (RSF Project #85-12-07). Dorn acknowledges funding from the Spanish Ministry of Science and Innovation (grants CSD2006-00016 and ECO2010-16726) and the Swiss National Science Foundation (grant BSSGI0-155804). Autor and Hanson acknowledge funding from the National Science Foundation (grant SES-1227334). We thank Andrew Cherlin, Janet Currie, Marianne Page, Ann Huff Stevens, Kathleen Vohs, Jane Waldfogel, two anonymous referees, and numerous seminar and conference participants for valuable suggestions. We are grateful to Juliette Fournier, Julius Lüttge, Ante Malenica, Timothy Simmons, Oscar Suen, Juliette Thibaud, and Melanie Wasserman for expert research assistance.
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An influential body of work associated with sociologist William Julius Wilson (Wilson 1987, 1996; Wilson and Neckerman 1986) hypothesizes that the decline of US blue-collar employment has diminished the pool of economically secure young adult men, thereby reducing women's gains from marriage, eroding traditional parental roles, and imperiling children.[1] Wilson's narrative is a close relative of the classic Becker (1973) framework in which the economic gains to marriage arise in part from spousal earnings differences, which spur household specialization.[2] Reflecting the difficulty of distinguishing cause from effect in the correlations between labor market opportunity and family structure, the literature has, with important exceptions, faced a challenge in testing these hypotheses.[3] We surmount this challenge by assessing how adverse labor market shocks for young adults, emanating from rising trade pressure on US manufacturing, affect marriage, fertility, and children's living circumstances. Following Autor, Dorn, and Hanson (2013b) and Acemoglu et al. (2016), we exploit cross-industry and cross-local labor market variation in import competition stemming from China's market reforms to trade to identify labor-demand shocks that are concentrated on manufacturing.

In linking local labor demand shocks to marriage and fertility, our work is close in spirit to Black, McKinnish, and Sanders (2003) who document an increasing prevalence of single-headed households in four US states that suffered a decline in their coal and steel industries, and Kearney and Wilson (2017) who observe rising fertility but no change in marital patterns in US regions that benefited from the 2000s fracking boom. Our study complements the evidence from these episodes of industry-specific booms and busts by assessing whether two decades of contracting US manufacturing employment across a large set of industries and localities has affected marital and family outcomes. Distinct from much prior work, we exploit gender dissimilarities in industry specialization to identify demand shocks that distinctly affect men's and women's employment and earnings. Our strategy thus complements work by Shenhav (2016), who uses gender-specific Bartik shocks and gender differences in occupational choice to predict changes in relative gender earnings in US states.[4] We provide three main results.

First, shocks to manufacturing labor demand, measured at the commuting-zone (CZ) level, exert large negative impacts on men's relative employment and earnings. Although losses are visible throughout the earnings distribution, the relative declines in male earnings are largest at the bottom of the distribution.[5]

Second, these shocks curtail the availability and desirability of potentially marriageable young men along multiple dimensions: reducing the share of men among young adults in a CZ and increasing the prevalence of idleness—the state of being neither employed nor in school—among young men who remain. Underscoring the

[1] See also Jahoda, Lazarsfeld, and Zeisel (1971); Murray (2012); Bailey and DiPrete (2016); and Greenwood, Guner, and Vandenbroucke (2017).

[2] Whereas Becker focuses on *relative* economic stature, Wilson's argument further implies that holding gender differentials constant, an *absolute* fall in male economic stature reduces the value of marriage.

[3] Exceptions include Angrist (2002) and Charles and Luoh (2010).

[4] Shenhav (2016) focuses on the economic independence of women rather than the declining marriage-market value of men, drawing in part on the empirical strategy in an earlier version of this paper (Autor, Dorn, and Hanson 2014).

[5] 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.

acuity of economic distress, we find, related to Case and Deaton (2015, 2017) and Pierce and Schott (2016b), that these forces induce a differential and economically large rise in male mortality from drug and alcohol poisoning, HIV/AIDS, and homicide.

Finally, we link manufacturing decline to marriage, fertility, and children's household circumstances. Much literature shows that adverse labor market shocks reduce the fraction of young women who are currently married. The Becker (1973) model makes a stronger prediction: a fall in the *relative* economic stature of men diminishes the gains from household specialization and therefore reduces the prevalence of marriage, while a decline in women's economic opportunities has the opposite effect. We confirm the Becker prediction using the gender-specific components of manufacturing decline, and further show that these shocks raise the fraction of mothers who are unwed, the fraction of children in single-headed households, and the fraction of children living in poverty.

These results complement work by Schaller (2016) showing that improvements in mens labor market conditions predict increases in fertility while improvements in womens labor market conditions do the opposite.[6] They also support Wilson's observation that manufacturing contractions shrink the pool of economically secure young adult men and erode traditional household arrangements. Because trade-induced manufacturing shocks generate both an *absolute* fall in the employment and earnings of young adult men and a fall in these outcomes *relative* to women, our empirical setting does not allow us to cleanly distinguish between the Becker hypothesis—focusing on *relative* economic stature—and Wilson's thesis, focusing on men's *absolute* economic stature.

Alongside providing support for the argument that contracting blue-collar employment catalyzes changes in gender roles and household structures, our analysis indicates that Wilson's conclusions apply to a far broader group of adults than the urban poor African Americans on whom he focused, and that the magnitude of these effects is sizable relative to observed declines in male employment rates, female fertility, and prevalence of marriage among US young adults.

I. Empirical Approach

We examine changes in exposure to international trade for US CZs associated with the growth in US imports from China. Rising trade with China is responsible for nearly all of the expansion in US imports from low-income countries since the early 1990s (Pierce and Schott 2016a). Our empirical strategy builds on Autor, Dorn, and Hanson (2013a) and Acemoglu et al. (2016). We approximate local labor markets using the construct of CZs developed by Tolbert and Sizer (1996), and include the 722 CZs that cover the US mainland.

[6] Work by Ananat, Gassman-Pines, and Gibson-Davis (2013) shows that adverse local economic shocks reduce teen birthrates and sexual activity, while raising contraceptive use and abortion, while Page, Stevens, and Lindo (2009) and Lindo, Schaller, and Hansen (2013) document adverse impacts of parental job loss on children's living circumstances.

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:

$$(1) \qquad \Delta IP_{i\tau}^{cu} = \sum_j \frac{L_{ij90}}{L_{i90}} \Delta IP_{j\tau}^{cu}.$$

Here, $\Delta IP_{j\tau}^{cu} = \Delta M_{j\tau}^{cu}/(Y_{j91} + M_{j91} - X_{j91})$ is the growth of Chinese import penetration in the US for industry j over period τ, which in our data include the time intervals 1990 to 2000 and 2000 to 2014. It is computed as the growth in US imports from China, $\Delta M_{j\tau}^{cu}$, divided by initial absorption (US industry shipments plus net imports, $Y_{j91} + M_{j91} - X_{j91}$) in the base year 1991, near the start of China's export boom. The fraction L_{ij90}/L_{i90} is the share of industry j in CZ i's total employment, as measured in County Business Patterns data in 1990. Differences in $\Delta IP_{i\tau}^{cu}$ across CZs stem from variation in local industry employment structure in 1990, which arises from differential concentration of employment in manufacturing versus non-manufacturing activities and specialization in import-intensive industries within local manufacturing. 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 $\Delta IP_{i\tau}^{cu}$ 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, we modify (1) to exploit the fact that manufacturing industries differ in their male and female employment intensity—so that trade shocks of a given magnitude will differentially affect male or female employment depending on the set of industries that are exposed. We incorporate this variation by multiplying the CZ-by-industry employment measure in (1) by the initial period female or male share of employment in each industry by CZ (f_{ij90} and $1 - f_{ij90}$), thus apportioning the total CZ-level measure into two additive subcomponents, $\Delta IP_{i\tau}^{m,cu}$ and $\Delta IP_{i\tau}^{f,cu}$:

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

As shown in online Appendix Table A1, Chinese import penetration rose by 0.95 percentage points between 1990–2000, with an additional 1.15 percent rise per decade over 2000–2014. Sixty percent of this rise accrued to male employment.[7]

We identify the supply-driven component of Chinese imports by instrumenting for growth in Chinese imports to the United States using the contemporaneous composition and growth of Chinese imports in eight other developed countries.[8] Our

[7] We construct (1) using trade data from UN Comtrade that we harmonize to 4-digit SIC industries, and data on CZ employment by industry from the County Business Patterns. In (2), we further use census IPUMS data to compute gender shares within industries and CZs, and assign to each SIC industry in a CZ the gender share of the census industry in the CZ encompassing it. Most outcome variables are based on census IPUMS samples for 1990 and 2000 (Ruggles et al. 2004) and pooled American Community Survey samples for 2013 through 2015. We allocate PUMAs to CZs using the algorithm in Dorn (2009) and Autor and Dorn (2013).

[8] The eight comparison countries—determined by the availability of comparable trade data for the full sample period—are Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland.

instrument for the measured import-exposure variable ΔIP_{it}^{cu} is a non-US exposure variable ΔIP_{it}^{co} that is constructed using data on industry-level growth of Chinese exports to other high-income markets:

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

This expression differs from (1) by using realized imports from China by other high-income markets ($\Delta M_{j\tau}^{co}$) in place of China-US import penetration ($\Delta M_{j\tau}^{cu}$); and by replacing other variables with lagged values to mitigate simultaneity bias.[9] The exclusion restriction underlying our instrumentation strategy is that the common component of import growth in the United States and in other high income countries derives from factors specific to China, associated with its rapidly evolving productivity and trade costs. Autor, Dorn, and Hanson (2013a) provide tests against correlated demand shocks and develop an alternative estimation strategy based on the gravity model of trade.

II. Main Results

A. *Employment and Earnings*

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

$$\Delta Y_{si\tau} = \alpha_t + \beta_1 \Delta IP_{i\tau}^{cu} + \mathbf{X}_{it}'\beta_2 + e_{si\tau}, \tag{4}$$

where ΔY_{sit} is the decadal change in the manufacturing employment share of the young adult population ages 18–39 in CZ i among gender group s (males, females, or both) during time interval τ. Our focus is on employment of young adults because this population is disproportionately engaged in marriage and child-rearing.[10] We estimate (4) by stacking ten-year equivalent first differences for 1990 to 2000 and 2000 to 2014, while including dummies for each decade (α_t). The explanatory variable of interest is the change in CZ-level import exposure $\Delta IP_{i\tau}^{cu}$, instrumented by $\Delta IP_{i\tau}^{co}$ as in (3).[11] The control vector $\mathbf{X}_{it}'$ contains start-of-period CZ-level covariates, including: time trends for US census divisions; the lagged share of CZ employment in manufacturing, absorbing general shocks to the sector; controls for employment

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

[10] Our sample is individuals who are not residents of institutionalized group quarters such as prisons. Here, we pool individuals of all races and ethnicities; we show in online Appendix Table A5 that our results also hold separately for non-Hispanic whites, for whom we have sufficient statistical power to conduct a separate CZ-level analysis.

[11] When performing gender-specific estimates, we replace $\Delta IP_{i\tau}^{cu}$ with $\Delta IP_{i\tau}^{m,cu}$ and $\Delta IP_{i\tau}^{f,cu}$, and use the corresponding gender-specific instruments. In online Appendix Table A2, we report OLS and 2SLS estimates separately by time period, with corresponding first stages for 2SLS models. A falsification test in this table shows that the negative effect of the China trade shock on employment is not present in the 1980s (i.e., prior to its occurrence) and has the opposite sign in the 1970s, when labor-intensive US manufacturing was still expanding in some US regions.

Table 1—Estimated Impact of Manufacturing Trade Shock on Manufacturing Employment by Gender and Gender Differential in Employment Status, Earnings, and Idleness, 1990–2014: 2SLS Estimates

	Panel A. Manufacturing employment as a share of population, age 18–39			*Panel B. Male-female differential by employment status, age 18–39*		
	M + F (1)	Males (2)	Females (3)	Emp (1)	Unemp (2)	NILF (3)
I. Overall trade shock						
Δ Import penetration	−1.06	−0.99	−1.09	−0.65	0.19	0.46
	(0.17)	(0.17)	(0.20)	(0.26)	(0.09)	(0.24)
II. Male industry versus female industry shock						
Δ Import penetration	−1.21	−2.59	0.20	−3.13	0.38	2.75
× (male ind emp share)	(0.44)	(0.51)	(0.43)	(0.78)	(0.26)	(0.62)
Δ Import penetration	−0.88	0.82	−2.56	2.17	−0.02	−2.15
× (female ind emp share)	(0.35)	(0.46)	(0.38)	(0.65)	(0.26)	(0.64)
Mean outcome variable	−2.61	−3.19	−2.06	−2.74	0.03	2.71
Level in 1990	12.98	17.37	8.68	14.64	1.22	−15.87
	Panel C. Male-female differential in annual earnings ($), age 18–39			*Panel D. Male-female differential in idleness, age 18–25*		
	P25 (1)	Median (2)	P75 (3)	Emp (1)	No emp in school (2)	No emp no school (3)
I. Overall trade shock						
Δ Import penetration	−672	−445	−847	−0.64	−0.02	0.66
	(193)	(191)	(334)	(0.34)	(0.26)	(0.20)
II. Male industry versus female industry shock						
Δ Import penetration	−2,216	−2,945	−3,685	−3.16	0.56	2.60
× (male ind emp share)	(516)	(593)	(1,081)	(1.03)	(0.73)	(0.60)
Δ Import penetration	1,086	2,400	2,384	2.24	−0.68	−1.55
× (female ind emp share)	(529)	(630)	(814)	(0.92)	(0.74)	(0.56)
Mean outcome variable level	−1,894	−2,126	−2,491	−2.83	−0.25	3.08
Level in 1990	6,926	13,376	17,489	7.70	0.87	−8.56

Notes: Observations = 1,444 (722 CZ × 2 time periods). Dependent variables: change in percentage of population age 18–39 that is employed in manufacturing, change in gender differentials in employment status (in percent pts); change in gender differential in annual earnings (in dollars); change in gender differential in percentage of young adults age 18–25 that are employed, not employed but in school, or neither employed nor in school. Panel C analyzes the change over time in the difference between a percentile of the unconditional male earnings distribution in a commuting zone and the corresponding percentile in the unconditional female earnings distribution. All models include a dummy for the 2000–2014 period, occupational composition controls (start-of-period indices of employment in routine occupations and of employment in offshorable occupations as defined in Autor and Dorn 2013), start-of-period shares of commuting zone population that is Hispanic, black, Asian, other race, foreign born, and college educated, as well as the fraction of women who are employed. Models are weighted by the product of period length and commuting zone share of start-of-period US mainland population. Robust standard errors in parentheses are clustered on state.

in occupations susceptible to automation and offshoring (see Autor and Dorn 2013 and Goos, Manning, and Salomons 2014); and CZ demographics (race, education, and the fraction of working-age adult women who are employed).

Panel A of Table 1 estimates the impact of rising trade exposure on the locus of proximate impact: manufacturing employment. In 1990, 17.4 percent of young men and 8.7 percent of young women ages 18–39 were employed in manufacturing (bottom row of panel A).[12] Estimates in panel A, columns 1–3, find that rising import competition reduces manufacturing employment among both sexes. A one unit trade shock—roughly equal to the average decade-level CZ-level rise in trade exposure over the 1990–2014 period—depresses the share of young adults employed in manufacturing by 1.06 percentage points ($t = -6.3$) with similar effects on young men ($\hat{\beta} = -0.99$, $t = -5.8$) and young women ($\hat{\beta} = -1.09$, $t = -5.5$). These estimates imply large declines in manufacturing employment among young adults over our 24-year window: a drop of 2.5 points (20 percent) among men and of 2.7 points (31 percent) among women.

We introduce the gender-specific trade shocks in panel A-II. Despite the high correlation between these by-gender measures ($\rho = 0.80$), there is abundant power for distinguishing their independent effects. Panel A-II, column 1, finds that a unit rise in import penetration of either male or female-dominated industries reduces manufacturing employment by 1 point. Panel A-II, columns 2 and 3, demonstrate that the employment effects of sex-specific shocks fall mostly on their corresponding genders. A unit trade shock to male-specific industries reduces male manufacturing employment by 2.6 points ($t = -5.1$) and has a small and statistically insignificant impact on female employment; a unit trade shock to female-specific industries reduces female manufacturing employment by 2.6 points ($t = -6.7$), while having a modest positive effect on male employment. Our ability to cleanly differentiate male- and female-specific shocks is crucial to the analysis that follows.

We consider the broader impacts on overall employment in Table 1, panel B. Here and below, we report the causal effects of trade shocks on the male-female *gap* in outcomes rather than on their levels so as to measure impacts on relative economic stature. Although trade shocks have similar impacts on male and female manufacturing employment, the estimate in panel B-I, column 1, shows that these shocks significantly depress the male relative to female employment-to-population rate—a unit trade shock reduces male relative employment by 0.65 points ($t = -2.5$).[13] While one might have predicted this differential effect based upon men's overrepresentation in manufacturing, the results in panel A—showing that trade shocks reduce male and female manufacturing employment in lockstep—underscore that differential manufacturing exposure is *not* the explanation. Rather, these estimates indicate that trade shocks differentially reduce male employment in non-manufacturing. Panel A of online Appendix Table A3 indicates that the overall employment loss of males due to a unit trade shock is larger than the decline in manufacturing employment seen in panel A of Table 1 (-1.5 versus -1.0 points) while the overall employment decline for women is slightly smaller than in manufacturing (-0.9 versus -1.1 points).[14]

We next quantify the impact of gender-specific trade shocks on the distribution of annual wage and salary income. For this analysis, we implement the Chetverikov, Larsen, and Palmer (2016) technique for performing instrumental-variable estimates of the distributional effects of group-level treatments. Panel C of Table 1 shows

[12] The denominator for this calculation is the noninstitutionalized adult population ages 18–39. Among *employed* adults in this demographic group, these fractions were 21.1 and 12.9 percent, respectively.

[13] Here, "relative" means the difference in levels rather than the ratio of levels.

[14] Acemoglu et al. (2016) find substantial employment losses in industries that sell their outputs to import-competing manufacturing, including mining whose workforce is strongly male-dominated.

estimates of the effect of trade shocks on the CZ-level male-female earnings gap for the twenty-fifth, fiftieth, and seventy-fifth percentiles of the distribution. Within CZs, male earnings substantially exceed female earnings at all quantiles, with the size of the gap rising steeply with the quantile index. In 1990, this gap was \$6,926, \$13,376, and \$17,489 at the twenty-fifth, fiftieth, and seventy-fifth quantiles, respectively (bottom rows of panel C). Between 1990 and 2014, these gaps compressed by \$1,894, \$2,126, and \$2,491 *per decade* at the twenty-fifth, fiftieth, and seventy-fifth quantiles. Reinforcing the panel B findings for the gender gap in employment, the first row of estimates in panel C demonstrates that trade shocks differentially curtail male earnings. A one-unit trade shock reduces male relative to female earnings by \$672 at the twenty-fifth percentile (panel C, column 1, $t = -3.5$), by \$445 at the median (panel C, column 2, $t = -2.3$), and by \$847 at the seventy-fifth percentile (panel C, column 3, $t = -2.5$).[15]

Since the male-female earnings gap is smaller at lower wage quantiles, the relative impact of trade shocks on the male-female wage gap is largest among low earners, as seen in Figure 1. Panel A details that trade-induced earnings losses are larger for males than females at every quantile from the fifteenth to ninety-fifth percentile.[16] Panel B reports the impact of a unit trade shock on the male-female annual earnings *gap* expressed as a percentage of baseline male earnings in 1990 at the corresponding percentile. Trade shocks modestly compress the male-female annual earnings gap in the upper half of the annual earnings distribution. The effect is more dramatic below: the male-female wage compression is 2 points at the median, 4 points at the thirty-fifth percentile, and 6 points at the twentieth percentile.[17]

The gender-specific estimates in online Appendix Table A3 confirm that trade shocks reduce employment and earnings of both genders; that employment and absolute earnings losses are larger for males than for females; and that proportional earnings losses for both sexes are larger at low than high percentiles. These findings support Wilson's observation that manufacturing contractions shrink the pool of economically secure young adult men, generating both an *absolute* fall in the employment and earnings of young adult men and a fall in these outcomes *relative* to women.

B. *Gender Gaps in Idleness, Absence, and Mortality*

The heart of the Wilson hypothesis is that adverse shocks to blue-collar employment catalyze a broader deterioration in adult social function. We test for such consequences with three nonmarket measures: idleness, absence, and mortality. Idleness is the state of being neither employed nor in school; we focus on the ages 18–25, which cover the transition between school and work.[18] In panel D of Table 1, we estimate a variant of (4) where the dependent variable is the male-female gap

[15] Panel C-II of Table 1 shows that shocks centered on male employment have a larger effect on the gender earnings gap than do shocks centered on female employment.

[16] Annual earnings for both genders are zero below the tenth percentile. Above the ninety-fifth percentile, earnings are largely censored and then imputed by the Census Bureau.

[17] These reductions are relative to the baseline male earnings *level* not the baseline gap, indicating large changes. We truncate the estimates at the twentieth percentile because the low values of the denominator below this point make estimates uninformative.

[18] Aguiar et al. (2017) document that young men devote more time to video games and recreational computer use, while working fewer hours.

FIGURE 1. IMPACT OF MANUFACTURING TRADE SHOCK ON EARNINGS OF MALES AND FEMALES AGE 18–39, 1990–2014

Notes: Panel A measures the impact of a unit trade shock on the unconditional distribution of annual earnings (in \$2015) separately for males and females. Each dot indicates a coefficient estimate from a separate IV quantile regression with group-level treatment (Chetverikov, Larsen, and Palmer 2016) that controls for the covariates indicated in Table 1, and shaded areas indicate a 95 percent confidence interval. Panel B reports the effect of a unit trade shock on the difference in the male-female annual earnings gap expressed as a percentage of male earnings in 1990 at the indicated percentile.

in three main (mutually exclusive) activity statutes: currently employed (panel D, column 1), not employed but enrolled in school (panel D, column 2), and neither employed nor enrolled in school (panel D, column 3), which we refer to as idleness. Panel D, column 1, shows that a unit trade shock lowers the fraction of young men employed by 0.64 percentage points relative to women of the same age range ($t = -2.5$). This is nearly identical to the effect found for the broader set of adults

Table 2—Estimated Impact of Manufacturing Trade Shock on Male Share of Adult Residents and Gender Differentials in Death Rates 1990–2015: 2SLS Estimates

	Panel A. Percentage of male residents		*Panel B. Male-female death rate differential ages 20–39*						
	18–39 (1)	18–25 (2)	Total (3)	D&A Poison (4)	HIV (5)	Homicide (6)	Suicide (7)	Accident (8)	All other (9)
I. Overall trade shock									
Δ Import penetration	−0.25	−0.28	64.4	19.5	21.6	14.0	−2.4	4.0	7.7
	(0.11)	(0.16)	(22.3)	(6.7)	(8.6)	(8.5)	(4.3)	(8.4)	(5.5)
II. Male versus female industry shock									
Δ Import penetration	−0.62	−0.76	189.7	60.3	66.5	103.0	−9.8	−40.2	9.9
× (male ind emp share)	(0.27)	(0.43)	(60.0)	(23.6)	(20.3)	(27.5)	(11.1)	(27.2)	(16.3)
Δ Import penetration	0.18	0.26	−77.1	−26.6	−29.1	−86.4	6.0	53.9	5.1
× (female ind emp share)	(0.15)	(0.29)	(49.0)	(18.9)	(15.5)	(29.2)	(10.9)	(27.1)	(16.7)
III. Summary stats: cumulative mortality 1990–2015 (decadal averages)									
Male-female gap			936.0	93.8	110.3	154.4	168.9	262.5	146.10
Males			1,644.6	153.3	146.6	198.6	218.6	378.6	548.92
Females			708.7	59.5	36.3	44.3	49.7	116.1	402.82

Notes: Observations = 1,444 (722 CZ × 2 time periods). Dependent variables: Change in percentage of male share of residents; cumulative male-female difference in death rates per 100,000 population by cause of death. The percentage of male residents is measured for the period 1990–2014 among all individuals who do not reside in institutionalized group quarters. Male share of CZ residents in 1990 was 49.6 percent among ages 18–39 and 50.2 percent among ages 18–25. Weighted mean changes in these variables were 0.11 and 0.19, respectively. Cumulative decadal mortality rates cover the period 1990–2015. All regressions include the full set of control variables from Table 1, and regressions in panel B control for a ten-year lag of the male-female differential in total mortality. Regressions are weighted by the product of period length and CZ population share, and standard errors are clustered on state.

ages 18–39 considered in panel B, column 1. Panel D, column 3, finds that the entire differential rise in nonparticipation among young males is due to increased idleness (0.66 points, $t = -2.5$), with little effect on school enrollment (panel D, column 2). This pattern is reinforced when focusing on the gender-specific components of trade exposure (panel D-II): shocks to male-intensive manufacturing generate a larger differential increase in male idleness (2.6 points, $t = 4.3$) than do shocks to female-intensive manufacturing (-1.6 points, $t = -2.8$).[19] Panel C of online Appendix Table A3 reports these impacts separately by gender. The differential effect of manufacturing shocks on the male-female idleness gap stem entirely from increases in male idleness. By contrast, reductions in female employment accrue almost entirely to increases in female school enrollment.

Table 2 presents complementary evidence on absence and mortality. Columns 1 and 2 indicate that trade shocks significantly reduce the supply of noninstitutionalized young men in a local labor market. A one unit trade shock reduces the fraction of males among adults ages 18–39 and 18–25 by about 0.25 points, with the effect on the broader age group more precisely estimated ($t = -2.2$) than the effect on the narrower age group ($t = -1.8$). The lower row of estimates (panel A-II) demonstrates that these (modest) shifts in the relative availability of young men

[19] In panel D, unemployed adults are categorized as either students or as idle. If we define idleness as the state of being neither employed, unemployed, nor in school, we continue to find a significant differential impact of trade shocks on male idleness.

stem from shocks to male-intensive employment: a unit shock to male-intensive manufacturing reduces the male fraction of young adults by two-thirds to three-quarters of a percentage point. A unit shock to female-intensive manufacturing generates a countervailing effect, but it is only one-third as large and is statistically insignificant.

The reduced supply of young adult males in trade-impacted CZs may reflect gender differences in migration, incarceration, homelessness, or mortality.[20] We focus on mortality, which is well-measured, has an unambiguous interpretation, and has attracted attention following Case and Deaton (2015, 2017). Using US Vital Statistics files enumerating person-level death certificates for all US residents, Table 2 reports the impact of trade shocks on the gender gap in cumulative mortality per decade—overall and by cause—per 100,000 adults ages 20–39.[21] Our analysis is related to Pierce and Schott (2016b), who link county-level trade exposure to rising mortality due to accidental poisoning and suicide in the working-age population. Guided by our focus on the interaction between labor markets and marriage markets, our analysis examines mortality among young adults ages 20–39 and differential effects on males versus females.

Shocks to import penetration significantly increase the male-female mortality gap among young adults. The point estimate in Table 2, panel B, column 3, indicates that a unit trade shock induces an additional 64.4 male relative to female deaths per 100,000 adults (of each gender) per decade. Given an average differential mortality rate of 936 per 100,000 adults per decade over 1990–2015, this increment is large. Subsequent columns decompose the overall mortality effect into by-cause categories. Case and Deaton (2015, 2017) show that drug and alcohol (D&A) related mortality rose by epidemic proportions among working-age adults in this time period. The bottom of panel B, column 4, indicates that D&A deaths accounted for 10 percent of all young adult male deaths between 1990–2015, while the upper row of the column demonstrates that the male-female gap in D&A deaths surged in trade-impacted CZs. The point estimate of 19.5 ($t = 2.9$) accounts for 30 percent of the total contribution of trade shocks to differential male mortality.

Panel B, columns 5–9 test for corresponding trade shock-related increases in differential male mortality from HIV/AIDS (often related to IV drug use), homicide, suicide, accidents, and all other causes.[22] A one unit trade shock causes a differential increase in male mortality due to HIV that is strongly significant ($t = 2.5$) and to homicides that is marginally significant ($t = 1.7$), where the former impact (21.6)

[20] Autor, Dorn, and Hanson (2013b) and Autor et al. (2014) do not find robust evidence for increased outmigration from import-competing CZ, but do not test for gender-specific migration patterns. Deiana (2015); Feler and Senses (2017); and Pierce and Schott (2016b) document statistically significant increases in crime incidents and arrests in such CZs during the 1990s and 2000s. Because incarceration and homelessness are disproportionately prevalent among males (West and Sabol 2008, Table 1 and online Appendix Table 7; US Conference of Mayors 2007, Exhibits 2.3 and 2.4), a rise in either may reduce the number of males enumerated in the noninstitutional population.

[21] These data, used under agreement with the US Center for Disease Control, cover deaths occurring in 1990 through 2015. The corresponding birth data (used below) extend through 2016. The denominator for death rates is the CZ-level population reported by the Census Bureau, which is available for the age bracket 20–39. The dependent variable is normalized to correspond to a 10-year cumulative value. Our regressions include the start-of-period control variables used in previous tables and the ten-year lag of the male-female differential in total mortality, such that we capture how trade shocks induce deviations in male-female mortality from long-run CZ-specific trends.

[22] The first row of section III in Table 2 shows that combined with D&A poisoning, the first four of these causes account for 84 percent of the male-female mortality gap over 1990–2015.

is slightly larger than the D&A effect and the latter impact (14.0) is moderately smaller. In panel B, columns 7–9, trade shocks have small and insignificant effects on differential male mortality related to suicide, accidents, and all other causes.

In net, the differential increase in male mortality can account for 14 percent of the fall in the fraction of males among young adults in trade-impacted CZs (panel A, column 1, of Table 2).[23] While only a small minority of adults who engage in risky behaviors experience fatal consequences, the remainder may be less attractive marital partners due to substance abuse and exposure to HIV and violent crime, suggesting that marriage-market values may fall for a broader set of young males.[24]

C. *Fertility, Marriage, and Children's Living Circumstances*

We test finally for impacts of trade shocks on fertility, marriage, and children's circumstances. Panel A of Table 3 presents the impact of trade exposure on marital status among women ages 18–39, whom we classify as currently married, currently widowed, divorced or separated, or never-married.[25] Trade shocks deter marriage formation: a one-unit trade shock predicts a 0.95 percentage-point decline in the fraction of young women who are currently married (panel A, column 1, $t = -3.1$), a further 0.21 point decline (panel A, column 2, $t = -2.0$) in the fraction of women who are previously married, and a corresponding rise of 1.2 points in the fraction of women never married (panel A, column 3, $t = 3.5$). Shocks to male and female-intensive employment have opposing and precisely estimated effects on marriage formation (panel A-II, columns 1–3): a one unit shock to male-intensive employment reduces the fraction of young adult women ever married by 4.2 points ($t = 6.6$, a 12 percent rise on a 1990 base of 34.8 percent) and the fraction currently married by 3.6 points ($t = -5.8$); a unit shock to female-intensive employment has a countervailing impact on marital status that is about two-thirds as large as the impact of a shock to male-intensive employment.

We find corresponding results for fertility, measured as births per 1,000 women ages 20–39. Trade shocks significantly deter fertility, with a one-unit shock reducing births by 1.5 per 1,000 women (panel B, column 4, $t = -4.2$). While it is tempting to interpret this pattern as indicative of the procyclicality of fertility, the lower rows of estimates (panel B-II) show otherwise. Shocks to male-intensive employment diminish fertility ($\hat{\beta} = -4.7$, $t = -5.4$) while shocks to female-intensive employment raise it ($\hat{\beta} = 2.0$, $t = 2.3$).

These results support Becker (1973), in which the gains to household formation are increasing in gender-based specialization. Here, shocks that diminish earnings capacity for the high-earning spouse (typically male) reduce these gains, deterring marriage and fertility—and vice versa for shocks that diminish earnings

[23] A unit trade shock reduces the male fraction of population by 0.25 per 100 adults among those ages 18–39 over the course of a decade (Table 2), implying an effect of 500 per 100,000 men. A unit trade shock raises excess male versus female mortality by 64 for every 100,000 adults of each sex among those ages 20–39 over the course of a decade (Table 2). Adjusting for the wider age range of the population versus mortality bracket (22 versus 20 years), this number rises to 71 per 100,000 adults. Thus, excess mortality can account for a share of $71/500 = 0.14$ of the decline in the male share of the young adult population in trade-impacted CZs.

[24] See also Charles and Luoh (2010) and Caucutt, Guner, and Rauh (2016).

[25] If a woman is currently married, we cannot determine if she was previously widowed, divorced, or separated.

TABLE 3—ESTIMATED IMPACT OF MANUFACTURING TRADE SHOCK ON MARRIAGE, FERTILITY, MATERNAL STATUS, CHILDHOOD POVERTY, AND HOUSEHOLD STRUCTURES OF ADULT WOMEN AND DEPENDENT CHILDREN, 1990–2014: 2SLS ESTIMATES

	Panel A. Women's marital status			*Panel B. Fertility and maternity*			*Panel C. Percent of children in HH < poverty line*
	Married (1)	Widowed divorced separated (2)	Never married (3)	Births per 1,000 women (4)	Percent of women w/ children (5)	Percent mothers unmarried (6)	(7)
I. Overall trade shock							
Δ Import penetration	−0.95	−0.21	1.16	−1.54	−0.66	0.52	0.61
	(0.30)	(0.11)	(0.33)	(0.37)	(0.23)	(0.31)	(0.26)
II. Male versus female industry shock							
Δ Import penetration	−3.57	−0.66	4.23	−4.65	−1.79	3.28	2.13
× (male share)	(0.62)	(0.22)	(0.64)	(0.84)	(0.63)	(0.73)	(0.70)
Δ Import penetration	2.03	0.29	−2.32	2.01	0.62	−2.62	−1.12
× (female share)	(0.55)	(0.19)	(0.58)	(0.87)	(0.52)	(0.85)	(0.82)
Mean outcome variable	−6.92	−1.62	8.55	−1.45	−3.53	6.56	1.65
Level in 1990	53.05	12.11	34.84	86.87	53.24	23.98	17.99

	Panel D. Women's household type			*Panel E. Children's household type*			
	Living w/ spouse (1)	Living w/ partner (2)	Other HH structure (3)	Married couple (4)	Parent + unmarried partner (5)	Single parent, no partner (6)	Grandparent or other (7)
I. Overall trade shock							
Δ Import penetration	−0.81	−0.22	1.03	−0.35	−0.11	0.30	0.15
	(0.27)	(0.12)	(0.30)	(0.19)	(0.07)	(0.11)	(0.16)
II. Male versus female industry shock							
Δ Import penetration	−3.21	0.04	3.17	−1.85	0.28	1.43	0.14
× (male share)	(0.55)	(0.28)	(0.60)	(0.50)	(0.23)	(0.32)	(0.42)
Δ Import penetration	1.93	−0.52	−1.41	1.36	−0.55	−0.98	0.17
× (female share)	(0.54)	(0.20)	(0.52)	(0.55)	(0.25)	(0.42)	(0.29)
Mean outcome variable	−7.57	1.65	5.93	−4.69	1.62	1.79	1.28
Level in 1990	50.30	5.25	44.45	71.39	2.82	16.82	8.96
Poverty rate 1990	n/a	n/a	n/a	8.7%	42.3%	47.4%	28.8%

Notes: Observations = 1,444 (722 CZ × 2 time periods). Dependent variables: changes in women's marital status, births per 1,000 women, fraction of women with children, and fraction of mothers unmarried; fraction of children living in poverty; and household type of women and children. Outcomes in panels A, B, and D consider adult women ages 18–39 while those in panels C and E consider children ages 0–17. Fertility in panel B, column 4 is measured through 2016 while all other outcomes are measured through 2014. Dependent variables are: fraction of women with any biological, adopted, or stepchildren in the household (panel B, column 5); fraction currently married among women with children in the household (panel B, column 6); the fraction of children in households below the official census poverty line (panel C, column 7). Panel D, columns 1 and 2, refer to households where either (1) the woman is the spouse or partner of the household head or (2) she is the household head and has a spouse or partner who is living in the household. Panel D, column 3, comprises all other household structures. Dependent variables in panel E, columns 4–7 are the fraction of children in each household type: household head is a married parent of the child (panel E, column 4); household head is a parent with cohabiting partner (panel E, column 5); household head is a single parent (panel E, column 6); or household head is a grandparent, other relative, or non-related caregiver (panel E, column 7). All regressions include the full set of control variables from Table 1, are weighted by the product of period length and CZ population share, and standard errors are clustered on state.

and employment for the low-earnings spouse. This reasoning, and the corroborating evidence in Table 3, helps explains why shocks to manufacturing employment are so damaging to adult social function: by differentially impairing male earnings capacity, such shocks reduce the attractiveness of marriage, fertility, and joint child-rearing.

Despite its strong predictions for marriage and fertility, the Becker (1973) framework is silent on the implications of shifts in relative economic status on children's living circumstances since this framework does not consider non-marital fertility. As an empirical matter, if a fall in males' relative economic stature deters fertility by at least as much as it deters marriage, more children will live in two-parent, married, and non-poor households.[26] Conversely, if motherhood is less elastic than marriage to shocks to relative economic stature, then children's household circumstances will move in the opposite direction. This latter possibility is implicit in the Wilson hypothesis.

Panel B, columns 5 and 6, and panel C, column 7 of Table 3 affirm Wilson's prediction. Panel B, column 5, shows that a unit trade shock reduces by 0.66 points the fraction of adult women ages 18–39 with children in the household ($t = -2.9$). Because this effect is only half as large as the increase in the fraction of women ages 18–39 who are never-married (panel A, column 3), the shock raises the share of mothers who are unmarried (panel B, column 6, $\hat{\beta} = 0.52, t = 1.70$), while the share of children living in poverty also increases (panel C, column 7, $\hat{\beta} = 0.61, t = 2.3$). Disaggregating the trade shock into its gender-specific components (panel B-II, columns 5 and 6, and panel C-II, column 7), trade shocks to male employment *reduce* the fraction of women with children (by 1.8 points) while *raising* the share of mothers who are unmarried by 3.3 points ($t = 4.5$) and the share of children living in poverty by 2.1 points ($t = 3.0$); shocks to female employment raise the prevalence of motherhood, reduce the fraction of mothers who are unmarried, and reduce the fraction of children living in poverty.

Panels D and E consider women's and children's living circumstances. Consistent with the panel A findings for marriage, a unit trade shock reduces the fraction of women living with a married partner by 0.81 points and the fraction cohabiting with an unmarried partner by additional 0.22 points. The declining marriage rate is thus not compensated by a rising propensity of young unmarried women to live with a partner.

Panel E documents how these countervailing effects on fertility, marriage, and single motherhood net out for children's circumstances. In panel E, column 4, the fraction of children living in married two-parent households falls by 0.35 points per unit trade shock ($t = -1.7$), while the fraction living in single-parent, non-cohabiting households rises by 0.30 points (panel E, column 6, $t = 2.8$). Echoing our findings for marriage and fertility, in panel E-II these adverse effects on children run entirely through shocks to male employment, which raise the share of children living in single-headed, non-cohabiting couples. Adverse shocks to female employment have *protective* effects for children, significantly raising the share of children in married households, reducing the share in non-married cohabiting and single-headed households, and weakly reducing the fraction of children living in poverty.

III. Conclusions

Our analysis confirms Wilson's (1996) hypothesis that contracting blue-collar employment catalyzes changes in marriage, fertility, household structures, and

[26] As the bottom row of Table 3 panel E shows, the fraction of children living in poverty in 1990 was 8.7 percent in married households, 42.3 percent in unmarried cohabiting households, and 47.4 in single-headed non-cohabiting households.

children's living circumstances. Contractions in the supply of economically secure young adult men stemming from rising trade pressure spur a surge in male idleness and premature mortality, a decline in marriage and fertility, an increase in the fraction of mothers who are unmarried and who are heads of single, non-cohabiting households, and a growth in the fraction of children raised in poverty. Whereas Wilson was writing about the African-American urban poor, our analysis encompasses all US young adults. Online Appendix Table A5 shows that our findings for employment, earnings, women's marital status, and childhood poverty are also confirmed when we focus on non-Hispanic whites.[27]

The implied magnitudes of our estimated impacts are quantitatively important. Scaling the observed rise in China import penetration between 1990 and 2014 by the estimates above, we would infer that rising trade pressure reduced the employment to population rate of young adult males by 3.9 percentage points as compared to an observed decline of 7.2 points; reduced the prevalence of marriage among young adult women by 2.4 percentage points as compared to an observed decline of 16.7 points; and increased the fraction of children living in poor households by 1.6 percentage points as compared to an observed increase of 4.0 points.

A key question unanswered by our work is whether reversing these adverse currents in blue-collar employment would undo their effects on marriage, fertility, and childhood poverty, or whether—as in Kearney and Wilson (2017)—some of these consequences would persist even where opportunities for blue-collar men improve.

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[27] Our census and ACS-based samples lack sufficient statistical power to analyze race *differences* in the consequences of trade shocks at the CZ level.

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