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

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Abstract

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, raise the prevalence of idleness, and elevate premature mortality among young males. Consistent with Becker's model of household specialization, shocks to male relative stature reduce marriage and fertility. Consistent with sociological accounts, these shocks raise the share of mothers who are unwed and share of children living in below-poverty, single-headed households.

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

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

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1 Introduction

An influential body of work associated with sociologist William Julius Wilson (1986; 1987; 1996) hypothesizes that the decline of U.S. 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 economically imperiling children.[1] Reflecting the difficulty of distinguishing cause from effect in the observed correlations between labor-market opportunities and marriage, fertility, and household structures, the literature has, with important exceptions, faced a challenge in testing the Wilson hypothesis.[2] We surmount this challenge by assessing how adverse shocks to the labor-market opportunities of young adults, emanating from rising trade pressure on U.S. manufacturing, affect marriage, fertility, household structure, and children's living circumstances. Following Autor et al. (2013b) and Acemoglu et al. (2016), we exploit cross-industry and cross-local-labor-market (i.e., commuting-zone) 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 et al. (2003) who document an increasing prevalence of single-headed households in four U.S. 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 U.S. 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 U.S. manufacturing employment across a large set of industries and localities has reduced the prevalence of marriage and increased the fraction of children living in poor and single-headed households. 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.[3] We provide three main results.

First, shocks to manufacturing labor demand, measured at the commuting-zone (CZ) level,

[1] See also Jahoda et al. (1971), Murray (2012), Bailey and DiPrete (2016), and Greenwood et al. (2017) for alternative treatments.

[2] Exceptions include Angrist 2002; Charles and Luoh 2010.

[3] Ananat et al. (2013) find that adverse local economic shocks reduce teen birthrates and sexual activity, while raising contraceptive use and abortion. 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 similar strategy, Schaller (2016) finds that improvements in men's labor market conditions predict increases in fertility while improvements in women's labor market conditions have the opposite effect. Page et al. (2007) and Lindo et al. (2013) document adverse impacts of parental job loss on children's living circumstances.

exert large negative impacts on men's relative employment and 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.[4]

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 acuity of economic distress, we find, related to Case and Deaton (2015; 2017) and Pierce and Schott (2016b), that these forces induce a differential rise in male mortality from drug and alcohol poisoning that explains a non-negligible share of the decline in the male-female gender ratio in trade-exposed CZs.

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 canonical Becker (1973) marriage 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 and fertility.[5] 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, non-cohabiting households, and the fraction of children living in poverty.

Alongside providing robust, representative support for Wilson's argument that contracting blue-collar employment catalyzes broad changes in gender roles and household structures, our analysis also provides an economic rationale for these findings: by differentially impairing male earnings capacity, shocks to manufacturing erode the gains from household specialization and thus deter joint investments in marriage, fertility, and child-rearing.

2 Empirical Approach

We examine changes in exposure to international trade for U.S. CZs associated with the growth in U.S. imports from China. Rising trade with China is responsible for nearly all of the expansion in U.S. imports from low-income countries since the early 1990s (Naughton, 2007; Hsieh and Klenow,

[4] 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.

[5] This idea is also tested by Schaller (2016) and Shenhav (2016), as discussed in footnote 3.

2009; Li et al., 2012; Pierce and Schott, 2016a). Our empirical strategy builds on Autor et al. (2013a) and Acemoglu et al. (2016). We approximate local labor markets using the construct of CZs developed by Tolbert and Sizer (1996). Commuting zones are suited to analysis of local labor markets because they encompass both urban and rural areas, and are based primarily on economic geography rather than incidental factors such as minimum population. Our analysis includes the 722 CZs that cover the U.S. mainland

Our measure of the local-labor-market shock is the average change in Chinese import penetration in a CZ's industries, weighted by each industry's share in initial CZ employment:

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

Here, $\Delta IP^{cu}_{j\tau} = \Delta M^{cu}_{j\tau}/(Y_{j91} + M_{j91} - X_{j91})$ is the growth of Chinese import penetration in the U.S. for industry j over period τ, which in our data include the time intervals 1990 to 2000 and 2000 to 2014. It is computed as the growth in U.S. imports from China, $\Delta M^{cu}_{j\tau}$, divided by initial absorption (U.S. industry shipments plus net imports, $Y_{j91} + M_{j91} - X_{j91}$) in the base year 1991, near the start of China's export boom. The fraction L_{ij90}/L_{i90} is the share of industry j in CZ i's total employment, as measured in County Business Patterns data in 1990. Differences in $\Delta IP^{cu}_{i\tau}$ 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 $\Delta IP^{cu}_{i\tau}$ 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, 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[6], $\Delta IP^{m,cu}_{i\tau}$ and $\Delta IP^{f,cu}_{i\tau}$:

[6] We use 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

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

As shown in 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.

We identify the supply-driven component of Chinese imports by instrumenting for growth in Chinese imports to the U.S. using the contemporaneous composition and growth of Chinese imports in eight other developed countries.[7] Our instrument for the measured import-exposure variable ΔIP^{cu}_{it} is a non-U.S. exposure variable ΔIP^{co}_{it} that is constructed using data on industry-level growth of Chinese exports to other high-income markets:

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

This expression differs from (1) by (a) using realized imports from China by other high-income markets ($\Delta M^{co}_{j\tau}$) in place of China-U.S. import penetration ($\Delta M^{cu}_{j\tau}$); and (b) replacing other variables with lagged values to mitigate any simultaneity bias.[8] The exclusion restriction underlying our instrumentation strategy is 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. Autor et al. (2013a) provide a large number of specific tests against correlated demand shocks and develop an alternative estimation strategy based on the gravity model of trade.

Dorn (2013). Data on CZ employment by industry from the County Business Patterns for 1980 and 1990 are used to compute employment shares by 4-digit SIC industries in (1) and (3). Because Census industries are coarser than the SIC codes in the Country Business Patterns 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.

[7] Trade data are from the UN Comtrade Database, which gives bilateral imports for 6-digit HS products. We harmonize Comtrade data with U.S. industry data as in Autor et al. (2014b). The eight comparison countries—determined by the availability of HS trade data for the full sample period-–are Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland. See Bloom et al. (2015) and Pierce and Schott (2016a) for alternative instrumentation strategies.

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

3 Main results

3.1 Employment and earnings

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

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

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.[9] 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 $\Delta IP^{cu}_{i\tau}$, instrumented by $\Delta IP^{co}_{i\tau}$ as in (3).[10] The control vector $\mathbf{X}'_{it}$ contains start-of-period CZ-level covariates, including: time trends for U.S. Census Divisions; the lagged share of CZ employment in manufacturing, absorbing general shocks to the sector; controls for employment in occupations susceptible to automation and offshoring (see Autor and Dorn, 2013 and Goos et al., 2014); and CZ demographics (race, education, and the fraction of working-age adult women who are employed).

The first panel 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).[11] Estimates in columns A1 through A3 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 ($\hat{\beta} = -1.06$, $t = -5.5$).

[9] Our sample is restricted to individuals who are not residents of institutionalized group quarters such as prisons, and who are thus potential participants in local labor and marriage markets. While the analysis pools individuals of all races and ethnicities, our results also hold separately for non-Hispanics whites, for whom we have sufficient statistical power to conduct a CZ-level analysis.

[10] When performing gender-specific estimates, we replace $\Delta IP^{cu}_{i\tau}$ with $\Delta IP^{m,cu}_{i\tau}$ and $\Delta IP^{f,cu}_{i\tau}$, and use the corresponding gender-specific instruments. In 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 U.S. manufacturing was still expanding in some U.S. regions.

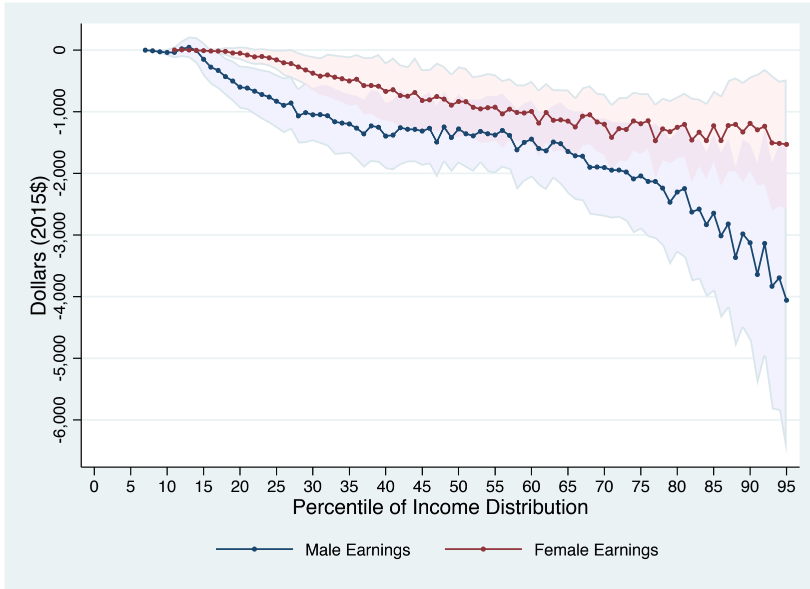
[11] The denominator for this calculation is the non-institutionalized adult population ages 18-39. Among employed adults in this demographic group, these fractions were 21.1 and 12.9 percent respectively.

Table 1: Impact of Manufacturing Trade Shock on Manufacturing Employment by Gender and Gender Differential in Employment Status, Earnings, and Idleness, 1990-2014: 2SLS Estimates. Dependent Variables: Changes in Percentage of Population Age 18-39 Employed in Manufacturing, Changes in Gender Differentials in Employment Status (in % pts); Change in Gender Differential in Annual Earnings (in $); Change in Gender Differential in Percentage of Young Adults Age 18-25 that is Employed, Not Employed but in School, or Neither Employed nor in School

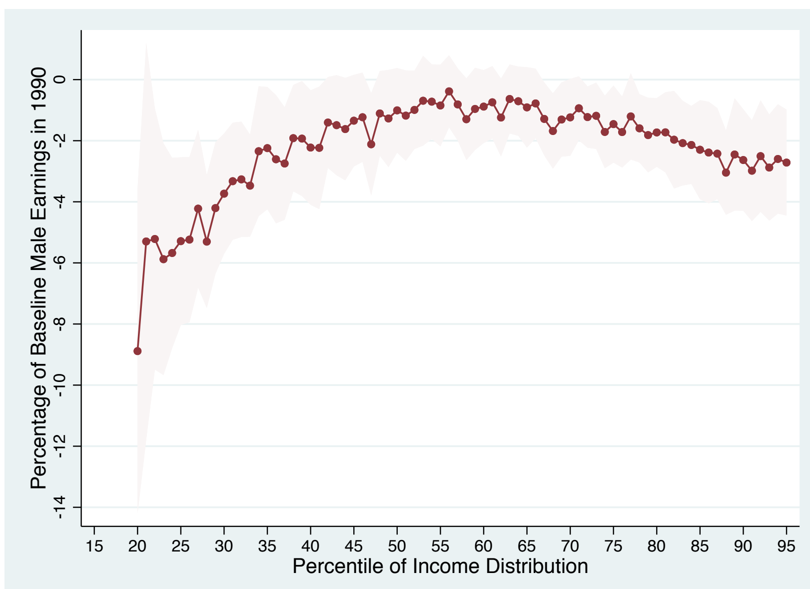
	A. Manufacturing Employment as a Share of Population, Age 18-39			*B. Male-Female Differential by Employment Status Age 18-39*		
	M+F	Males	Females	Emp	Unemp	NILF
	(1)	(2)	(3)	(1)	(2)	(3)
			I. Overall Trade Shock			
Δ Chinese Import Penetration	-1.06 **	-0.99 **	-1.09 **	-0.65 *	0.19 *	0.46 ~
	(0.17)	(0.17)	(0.20)	(0.26)	(0.09)	(0.24)
			II. Male Industry vs Female Industry Shock			
Δ Chinese Import Penetration	-1.21 **	-2.59 **	0.20	-2.96 **	0.38	2.58 **
× (Male Ind Emp Share)	(0.44)	(0.51)	(0.43)	(0.76)	(0.26)	(0.62)
Δ Chinese Import Penetration	-0.88 *	0.82 ~	-2.56 **	2.63 **	-0.02	-2.61 **
× (Female Ind Emp Share)	(0.35)	(0.46)	(0.38)	(0.58)	(0.26)	(0.55)
Mean Outcome Variable	-2.61	-3.19	-2.06	-2.74	0.03	2.71
Level in 1990	12.98	17.37	8.68	14.64	1.22	-15.87
	C. Male-Female Differential in Annual Earnings ($) Age 18-39			*D. M-F Diff in Idleness, Age 18-25*		
					No Emp	No Emp
	P25	Median	P75	Emp	In School	No School
	(1)	(2)	(3)	(1)	(2)	(3)
			I. Overall Trade Shock			
Δ Chinese Import Penetration	-672 **	-445 *	-847 *	-0.64 ~	-0.02	0.66 **
	(193)	(191)	(334)	(0.34)	(0.26)	(0.20)
			II. Male Industry vs Female Industry Shock			
Δ Chinese Import Penetration	-2,216 **	-2,945 **	-3,685 **	-3.16 **	0.56	2.60 **
× (Male Ind Emp Share)	(516)	(593)	(1081)	(1.03)	(0.73)	(0.60)
Δ Chinese 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	-1,894	-2,126	-2,491	-2.83	-0.25	3.08
Level in 1990	7,041	13,599	17,780	7.70	0.87	-8.56

Notes: N=1444 (722 CZ x 2 time periods). 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. Robust standard errors in parentheses are clustered on state. Models are weighted by start of period commuting zone share of national population. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Figure 1: Impact of Manufacturing Trade Shock on Earnings of Males and Females Age 18-39, 1990-2014



A. Impact on Male and Female Annual Earnings by Percentile, 1990 - 2014



B. Impact on Male-Female Annual Earnings Gap 1990-2014 as a Percentage of 1990 Male Earnings

Panel A reports 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% confidence interval. Panel B reports the effect of a unit trade shock on difference in the male-female annual earnings gap expressed as a percentage of male earnings in 1990 at the indicated percentile.

These estimates imply large cumulative declines in manufacturing employment among young adults over our 24-year sample window: a drop of 2.5 percentage points (20 percent) among men and of 2.7 percentage points (31 percent) among women. While these results are consistent with Autor et al. (2013a), they cover a longer time interval and differentiate employment impacts by gender.

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 on labor-market outcomes.[12] Column A1-II finds that a unit rise in import penetration of either male or female-dominated industries reduces manufacturing employment by 1 percentage point. Columns A2-II and A3-II 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. Conversely, 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. Because our objective is to assess how shocks to the relative economic stature of men and women affect fertility, marriage, and household structure, the ability to cleanly differentiate gender-specific shocks is crucial to what follows.

While the proximate locus of the trade shock is manufacturing, it may have broader consequences. We consider impacts on overall employment in panel B of Table 1. 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 column B1-I shows that these shocks significantly depress the male relative to female employment-to-population rate—a unit trade shock reduces the male relative to female employment-to-population ratio among young adults 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

[12] We use the terms male-specific and female-specific shocks as a shorthand for the gender-specific trade exposure measure as defined in (2).

[13] For brevity, we use the term 'relative' to mean the difference in levels rather than the ratio of levels.

reduce male employment in non-manufacturing.

Panel C of Table 1 provides a fuller picture of the gender-non-neutrality of trade shocks by quantifying their effect on the gender gap on the distribution of annual wage and salary income. For this analysis, we implement the Chetverikov et al. (2016) technique for performing instrumental-variable estimates of the distributional effects of group-level treatments. Let $y_{it_0}(u)$ equal the unconditional male-female annual earnings gap (in real 2015 US\$) in CZ i in year t_0 at quantile u among residents ages 18-39.[14] Let $\Delta y_{i\tau}(u)$ equal the change in this gap between time periods t_0 and t_1, corresponding to either $1990-2000$ or $2000-2014$.

Our estimating equation takes the form

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

where $Q_{\Delta y_{i\tau}|,x_{it},\varepsilon_{it}}(u)$ is the u^{th} conditional quantile of Δy_{it} given $(\alpha_t, \Delta IP^{cu}_{i\tau}, \mathbf{X}_{it}, \varepsilon_{it})$, where α_t is an intercept, $\Delta IP^{cu}_{i\tau}$ is the China-Shock measure (instrumented as above), $\mathbf{X}_{it}$ is the 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 $\Delta y_{i\tau}$. As above, we estimate (5) in stacked first differences.[15]

Panel C of Table 1 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. 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 \$7,041, \$13,599, and \$17,780 at the 25^{th}, 50^{th}, and 75^{th} quantiles, respectively (bottom rows of panel C). Between 1990 and 2014, these gaps compressed by \$1,894, \$2,125 and \$2,491 *per decade* at the 25^{th}, 50^{th}, and 75^{th} quantiles respectively. 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 25^{th} percentile (column C1, $t = -3.5$), by \$445 at

[14] 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 18-39, including those with zero earnings.

[15] Chetverikov et al. (2016) provide 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. Because the outcome of interest is the CZ-level distribution of the unconditional male-female earnings gap, our estimating equation excludes person-level covariates and thus corresponds to step 2 of the Chetverikov et al. (2016) procedure.

the median (column C2, $t = -2.3$), and by \$847 at the 75^{th} percentile (column C3, $t = -2.5$).[16]

Since the male-female earnings gap is smaller at lower wage quantiles, these results suggest that the relative impact of trade shocks on the male-female wage gap is largest among low-earners. Figure 1 confirms this intuition. The first panel details that trade-induced earnings losses are larger for males than females at every quantile from the 15^{th} to 95^{th} percentile.[17] The second panel 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 and reaches 4 points at $p35$ and 6 points at $p20$.[18]

These results support Wilson's hypothesis that manufacturing contractions shrink the pool of economically secure young adult men—though notably, differential male employment losses accrue *outside* of manufacturing.

3.2 Gender gaps in idleness, absence, and mortality

The heart of Wilson's thesis is that shocks to blue-collar labor demand reshape adult social function. We test these consequences with three non-market measures: idleness, absence, and mortality. Idleness is the state of being neither employed nor in school; we focus on those ages 18-25 since many of this age are transitioning between school and work.[19] In panel D of Table 1, we estimate a variant of (4) where the dependent variable is the male-female gap in three mutually exclusive outcomes: currently employed (D1), not employed but enrolled in school (D2), and neither employed nor enrolled in school (D3). Column D1 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 ages 18-39 considered in column B1. Column D3 finds that the entire differential rise in non-participation among young males is due

[16]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.

[17]Annual earnings for both genders are zero below the 10^{th} percentile. Above the 95^{th} percentile, earnings are largely censored and then imputed by the Census Bureau.

[18]Recall that these reductions are relative to the baseline male earnings *level* not the baseline gap, so these are large changes. We truncate the estimates at $p20$ because the low values of the denominator below this point make for uninformative estimates.

[19]Aguiar, Bils, Charles and Hurst (2017) document that young men devote more time to video games and recreational computer use, while working fewer hours.

to increased idleness (0.66 points, $t = -2.5$), with little effect on school enrollment (column D2). 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, $t = -2.8$).[20]

Table 2: Impact of Manufacturing Trade Shock on Male Share of Adult Residents and Cumulative Gender Differential in Mortality, 1990-2015: 2SLS Estimates. Dependent Variables: Change in Percentage of Male Share of Residents, Adults 18-39 and 18-25; Difference in Male-Female Cumulative Deaths per 100k Population by Cause of Death, Adults Age 18-39

	A. Percentage of Male Residents		*B. Male-Female Death Rate Differential Ages 20-39*						
	18-39	18-25	Total	D&A Poison	Liver Disease	Diabetes	Lung Cancer	Suicide	All Other
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
			I. Overall Trade Shock						
Δ Chinese Import Penetration	-0.25 *	-0.28 ~	74.32 *	22.80 **	-0.26	-1.48	0.94	-5.00	43.81 ~
	(0.11)	(0.16)	(29.34)	(7.42)	(1.22)	(1.03)	(0.62)	(4.75)	(24.89)
			II. Male vs Female Industry Shock						
Δ Chinese Import Penetration	-0.62 *	-0.76 ~	168.40 ~	70.22 **	-2.06	1.91	-1.70	-22.99 *	82.62
× (Male Ind Emp Share)	(0.27)	(0.43)	(85.95)	(23.03)	(3.18)	(3.36)	(1.50)	(11.72)	(78.68)
Δ Chinese Import Penetration	0.18	0.26	-27.07	-30.03	1.75	-5.26	3.87 *	14.90	2.07
× (Female Ind Emp Share)	(0.15)	(0.29)	(75.58)	(20.79)	(3.39)	(3.79)	(1.83)	(10.71)	(69.28)
			III. Summary Stats: Cumulative Mortality 1990-2015 (Decadal Averages)						
Males			2,113.3	203.5	31.8	25.3	12.0	282.4	1,558.2
Females			914.8	79.6	16.4	18.1	9.9	64.5	726.3
Male-Female Gap			1,198.5	123.9	15.4	7.3	2.1	217.9	831.9

Notes: N=1444 (722 CZ x 2 time periods). 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% among 18-39 and 50.2% among ages 18-25. Weighted mean changes in these variables were 0.11 and 0.19 respectively. Mortality rates cover the period 1990-2015. 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 2 presents complementary evidence on absence and mortality. The first two columns test whether trade shocks reduce the supply of non-institutionalized young men in a local labor market. Columns 1 and 2 confirm that they do. A one unit trade shock reduces the fraction of males among adults ages 18-39 and 18-25 by 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 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.

[20] 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. The corresponding point estimates and standard errors for column D3 under this definition are: 0.34 (0.15), 2.13 (0.41), and -1.69 (0.43) for rows 1, 2, and 3 respectively.

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.[21] We focus on mortality, because it is well-measured, has an unambiguous interpretation, and has attracted attention following Case and Deaton (2015; 2017). Using U.S. Vital Statistics files enumerating person-level death certificates for all U.S. residents, Table 2 reports the impact of trade shocks on the gender gap in cumulative mortality per decade—overall and by cause—per 100K adults ages 20-39.[22] Our analysis is related to Pierce and Schott (2016b), who link county-level trade exposure to rising mortality due to suicide, accidental poisoning, and liver disease. Distinct from prior work, we analyze mortality among young adults ages 20-39, consider differential effects on males versus females, and employ CZs rather than counties as the unit of analysis, choices guided by our focus on how labor and marriage markets interact.

Shocks to import penetration increase the male-female mortality gap significantly among young adults. The point estimate in column B3 of the upper panel of Table 2 indicates that a unit trade shock induces an additional 74.3 male relative to female deaths per 100K adults (of each gender) per decade. Given an average differential mortality rate of 1200 per 100K adults per decade over 1990 - 2015, this increment is large. Subsequent columns decompose the overall mortality effect into by-cause categories using the scheme in Case and Deaton (2015, Figure 2). 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 column B4 indicates that D&A deaths accounted for 10 percent of all young adult male deaths between 1990 - 2015, while the point estimate in the upper row of the column demonstrates that male D&A deaths surged in trade-impacted CZs. The point estimate of 22.8 ($t = 3.1$) accounts for one-third of the total contribution of trade shocks to differential mortality.

[21] See also Monte et al. (2015). Deiana (2015), Feler and Senses (2015), and Pierce and Schott (2016b), who document statistically significant increases in crime incidents and arrests in trade-exposed CZs during the 1990s and 2000s. Because incarceration and homelessness are disproportionately prevalent among males (West and Sabol 2008, Table 1 and Appendix Table 7; U.S. Conference of Mayors, 2007, Exhibits 2.3 and 2.4), a rise in either may reduce the number of males enumerated in the non-institutional population.

[22] These vital statistics data, used under agreement with the U.S. Center for Disease Control, cover deaths occurring in 1985, 1989, and 1990 through 2015. The corresponding birth data (used below) extend through the year 2016. The dependent variable is normalized to correspond to a 10-year cumulative value. Our regression models include the same vector of start-of-period control variables used in previous tables and account for serial correlation in CZ-level mortality rates by additionally controlling for lagged cumulative decadal mortality.

Columns B5 through B8 test for corresponding trade shock-related increases in differential male mortality from liver disease (often alcohol-related), diabetes, lung cancer, and suicide. No effect is significant. The final column (B9) combines all other causes of death beyond those emphasized by Case and Deaton (2015), including infectious diseases, neoplasms and accidents, which account for three of every four young-adult deaths. The point estimate of 43.9 indicates that trade shocks contribute to an overall increase in the gender mortality gap among all remaining causes, with this estimate marginally significant ($t = 1.8$).[23]

Overall, the differential increase in male mortality can account for 16 percent of the fall in the fraction of males among young adults in trade-impacted CZs (column A1 of Table 2).[24] While only a small minority of adults who engage in adverse health behaviors experience fatal consequences, the remainder may be less viable as marital partners due to substance abuse. The differential rise in fatal overdoses among young adult males in trade-exposed locations may imply a fall in the marriage-market value of a far larger set of males.[25]

3.3 Fertility, marriage, and children's living circumstances.

We test finally for impacts 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.[26] Trade shocks deter marriage formation: a one-unit trade shock predicts a 0.95 percentage-point decline in the the fraction of young women who are currently married (column A1, $t = -3.1$), a further 0.21 point decline (column A2, $t = -2.0$) in the fraction of women who are previously married, and a corresponding rise of 1.2 points in the fraction of women never married (column A3, $t = 3.5$). Shocks to male and female-intensive employment have opposing and precisely estimated effects on marriage formation (columns A1-II through A3-II): a one unit shock to male-intensive employment reduces

[23] Models using gender-specific trade shocks (panel B-II) find that essentially all of the gender-differential in mortality effects stems from shocks to male-intensive employment. Similarly, models for by-gender mortality find that the mortality response to trade shocks stems almost entirely from male deaths (Appendix Table A3).

[24] 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 100K men. A unit trade shock raises excess male versus female mortality by 74 for every 100K 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 81 per 100K adults. Thus, excess mortality can account for percent ($81/500 = 0.16$) of the decline in the male share of the young adult population in trade-impacted CZs.

[25] See also Charles and Luoh (2010) and Caucutt et al. (2016).

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

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 ($t = -5.8$). Adverse shocks to female-intensive employment have the opposite effects, and these effects are two-thirds as large as the impacts of shocks to male employment.

We find corresponding results for fertility, measured as births per $1,000$ women ages 18-39. Trade shocks significantly deter fertility, with a one-unit shock shock reducing births by 2.0 per $1,000$ women (column B4, $t = -3.6$). 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 ($\hat{\beta} = -6.1, t = -4.8$) while shocks to female-intensive employment raise it ($\hat{\beta} = 2.5, t = 2.0$).

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 capacity 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 strong predictions for marriage and fertility, the implications of Becker (1973) for children's living circumstances are ambiguous. 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.[27] 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.

[27] The bottom row of Table 3 panel E shows that in 1990, the fraction of children living in poverty was 8.1 percent in married households, 42.3 percent in unmarried cohabiting households, and 47.4 in single-headed non-cohabiting households.

Table 3: Impact of Manufacturing Trade Shock on Marriage, Fertility, Maternal Status, Childhood Poverty, and Household Structures of Adult Women and Dependent Children, 1990-2000/2014: 2SLS Estimates. Dependent Variables: Women's Marital Status (Ages 18-39); Births per 1,000 Women Ages 18-39; Fraction of Women Ages 18-39 with Children, and Fraction of Mothers Ages 18-39 Unmarried; Fraction of Children Living in Poverty; and Household Type of Women Ages 18-39 and Children Ages 0-17

	A. Women's Marital Status			*B. Fertility and Maternity*			*C. % of Children in HH < Poverty Line*
	Married	Widowed Divorced Separated	Never Married	Births per 1,000 Women	% of Women w/ Children	% Mothers Unmarried	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				I. Overall Trade Shock			
Δ Chinese Penetration	-0.95 **	-0.21 *	1.16 **	-2.02 **	-0.66 **	0.52 ~	0.61 *
	(0.30)	(0.11)	(0.33)	(0.57)	(0.23)	(0.31)	(0.26)
				II. Male vs Female Industry Shock			
Δ Chinese Penetration × (Male Share)	-3.57 **	-0.66 **	4.23 **	-6.10 **	-1.79 **	3.28 **	2.13 **
	(0.62)	(0.22)	(0.64)	(1.28)	(0.63)	(0.73)	(0.70)
Δ Chinese Penetration × (Female Share)	2.03 **	0.29	-2.32 **	2.51 *	0.62	-2.62 **	-1.12
	(0.55)	(0.19)	(0.58)	(1.25)	(0.52)	(0.85)	(0.82)
Mean Outcome Var	-6.92	-1.62	8.55	-1.90	-3.53	6.56	1.65
Level in 1990	53.05	12.11	34.84	86.87	53.24	23.98	17.99

	D. Women's Household Type			*E. Children's Household Type*			
	Living w/ Spouse	Living w/ Partner	Other HH Structure	Married Couple	Parent + Unmarried Partner	Single Parent, No Partner	Grand-parent or Other
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				I. Overall Trade Shock			
Δ Chinese 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 vs Female Industry Shock			
Δ Chinese Penetration × (Male Share)	-3.21 **	0.04	3.17 **	-1.85 **	0.28	1.43 **	0.14
	(0.55)	(0.28)	(0.60)	(0.50)	(0.23)	(0.32)	(0.42)
Δ Chinese Penetration × (Female Share)	1.93 **	-0.52 **	-1.41 **	1.36 *	-0.55 *	-0.98 *	0.17
	(0.54)	(0.20)	(0.52)	(0.55)	(0.25)	(0.42)	(0.29)
Mean Outcome Var	-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	8.1%	42.3%	47.4%	28.8%

Notes: N=1444 (722 CZ x 2 time periods). 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 column B4 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 (B5); fraction currently married among women with children in the household (B6); the fraction of children in households below the official Census poverty line (C7). Columns D1 and D2 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. Column D3 comprises all other household structures. Dependent variables in columns E4-E7 are the fraction of children in each household type: household head is a married parent of the child (E4); household head is a parent with cohabiting partner (E5); household head is a single parent (E6); or household head is a grandparent, other relative, or non-related caregiver (E7). All regressions include the full set of control variables from Table 1 and are weighted by start-of-period CZ population. Standard errors are clusterd by state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Columns B5, B6, and C7 of Table 3 affirm Wilson's prediction. Column B5 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 (column A3), the shock raises the share of mothers who are unmarried (column B6, $\hat{\beta} = 0.52, t = 1.70$) and the share of children living in poverty (column C7, $\hat{\beta} = 0.61, t = 2.3$). Disaggregating the trade shock into its gender-specific components (columns B5-II, B6-II, and C7-II) underscores their distinct effects on children's living circumstances. 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 complete the picture of household adjustment by considering 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 percentage points and the fraction cohabiting with an unmarried partner by additional 0.22 points. The reduction in cohabitation is slightly smaller than the corresponding reduction in marriage, indicating that trade shocks reduce the prevalence of married cohabiting couples, while have little effect on the household status of non-married adults.

Panel E documents how these countervailing effects on fertility, marriage, and single motherhood net out for children's circumstances. In column E4, 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 (column E6, $t = 2.8$). 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 on children's circumstances: 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.

4 Conclusions

Our analysis confirms William Julius Wilson's hypothesis that contracting blue-collar employment catalyzes changes in marriage, fertility, household structures, and 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.

Two economic mechanisms explain these adverse outcomes. First, trade-induced manufacturing decline reduces the economic stature of men relative to women. Second, consistent with Becker's model of specialization in marriage, relative declines in male versus female stature have countervailing impacts on fertility, marriage, and cohabitation, reflecting their opposing effects on gains from specialization.

A key remaining question is whether reversing trends in blue-collar employment would undo effects on marriage, fertility, and childhood poverty, or whether—as in Kearney and Wilson (2017)—these consequences may persist even where opportunities for blue-collar men improve.

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5 Appendix Tables

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

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

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

Table A2: Impact of Manufacturing Trade Shock on Employment, 1970-2014: OLS and 2SLS Estimates. Dependent Variable: Change in Percentage of Population Age 18-39 Employed in Manufacturing

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

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

Table A3: Impact of Manufacturing Trade Shock on Cumulative Mortality by Gender 1990-2015: 2SLS Estimates. Dependent Variable: Male and Female Cumulative Mortality per 100k Population Ages 20-39 by Cause of Death

	Total		D&A Poison		Liver Disease	Diabetes	Lung Cancer	Suicide		All Other	
	(1)		(2)		(3)	(4)	(5)	(6)		(7)	
					I. Male Mortality						
					Overall Trade Shock						
Δ Chinese Penetration	103.70	**	29.62	**	-0.82	-1.25	0.65	-2.67		57.51	~
	(40.48)		(11.27)		(1.31)	(0.84)	(0.59)	(5.69)		(33.08)	
					Male Industry vs Female Industry Shock						
Δ Chinese Penetration	225.50	*	91.77	**	-2.83	-0.81	0.40	-27.54	~	97.20	
× (Male Ind Share)	(110.90)		(31.48)		(3.91)	(2.83)	(1.37)	(14.74)		(95.99)	
Δ Chinese Penetration	-27.14		-39.60		1.42	-1.73	0.93	24.86	*	15.07	
× (Female Emp Share)	(114.60)		(30.71)		(3.06)	(2.73)	(1.99)	(12.24)		(101.10)	
					II. Female Mortality						
					Overall Trade Shock						
Δ Chinese Penetration	21.83		7.06		0.26	0.14	-0.29	0.52		8.51	
	(16.61)		(4.72)		(1.08)	(0.57)	(0.37)	(1.46)		(13.28)	
					Male Industry vs Female Industry Shock						
Δ Chinese Penetration	33.04		21.34	~	1.05	-2.67	1.97	-10.67	*	0.32	
× (Male Ind Share)	(50.98)		(11.40)		(2.38)	(1.82)	(1.44)	(4.71)		(40.26)	
Δ Chinese Penetration	9.63		-8.83		-0.61	3.28	-2.80	12.97	**	17.41	
× (Female Emp Share)	(56.11)		(13.16)		(2.50)	(2.38)	(1.77)	(5.04)		(46.56)	

Notes: N=1444 (722 CZ x 2 time periods). All regressions include the full set of control variables from Table 1 and the lagged 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.