

The Labor Market and the Marriage Market: How Adverse Employment Shocks Affect Marriage, Fertility, and Children's Living Circumstances*

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Abstract

The structure of marriage and child-rearing in U.S. households has undergone two marked shifts in the last three decades: a steep decline in the prevalence of marriage among young adults, and a sharp rise in the fraction of children born to unmarried mothers or living in single-headed households, the latter of which is concentrated among non-college and minority households and thus particularly affects lower-SES children. A potential contributor to both phenomena is the declining labor market opportunities faced by non-college and minority males, which make these males less valuable as marital partners. We explore the impact of the labor market on the marriage market by exploiting large scale, plausibly exogenous trade-induced shocks to local manufacturing employment, stemming from rising import competition from China. We trace out how these shocks impact marriage, divorce, childbearing, and the prevalence of children growing up in poor and single-parent households. We find that trade shocks between 1990 and 2010 have had quite modest impacts on household structure in aggregate. When we disaggregate these shocks into components affecting male versus female employment, however, we find impacts that are both economically and statistically significant. Import shocks concentrated on male employment reduce marriage rates and fertility, raise the fraction of births due to teen mothers, and, most significantly, increase the fraction of children living either in poverty or in single-headed households. On net, our findings do not suggest that rising import competition from China has been an important contributor to changing marital behavior in this time interval, since these shocks have not been particularly biased against males. But our analysis strongly supports the hypothesis that changes in labor demand that reduce male employment opportunities—and in particular, the sharp decline in labor market conditions facing non-college U.S. males over the last three decades— may be a quantitatively important contributor to the rise in the share of U.S. children living in poor and in single-headed households.

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Introduction

The structure of marriage and child-rearing in U.S. households has undergone two marked shifts in the last three decades. A first is the steep decline in the prevalence of marriage among young adults, and among less educated adults in particular. Between 1979 and 2008, the fraction of U.S. women between the ages of 25 and 39 who were currently married fell by roughly 10 percentage points among college-educated women, by 15 percentage points among women with some college but no degree, and by fully 20 percentage points among women with high school or lower education.¹ These declines reflect a combination of a higher age at first marriage, a decline in lifetime marriage rates and, to a lesser extent, a rise in divorce rates among less educated women (Cherlin, 2010; Heuveline, Timberlake and Furstenberg, 2003). Accompanying the fall in marriage rates, a second change in household structures is a sharp rise in the fraction of children born to unmarried mothers or living in single-headed households. Between 1980 and 2009, the fraction of all U.S. births accounted for by unmarried mothers more than doubled from 18 percent in 1980 to 41 percent. The rise in non-marital births was mirrored by an increase in the fraction of children living without biological fathers—and, as with falling marriage rates, this increase is concentrated among less-educated adults.² For example, the fraction of white non-Hispanic children under age 18 born to mothers with less than a high school degree living in mother-headed households rose from 17 to 42 percent between 1980 and 2010. The comparable increase was from 12 to 29 percent among children born to mothers with exactly a high school diploma, and from 11 to 18 percent among children born to mothers with post-high school education.³ Thus, U.S. women of all education levels have chosen to postpone marriage, yet less-educated women have chosen *not* to postpone childbearing—at least, not nearly to the same degree.

The causes of the gradual decoupling of marriage from child-rearing has drawn decades of sustained research and policy attention, exemplified by William Julius Wilson’s pioneering 1987 book,

¹All statistics in this paragraph are drawn from Autor and Wasserman (2013), who in employ data from U.S. Census of Population and American Community Survey samples, as well as published reports from the National Vital Statistics Data System.

²A concern raised by these demographic trends, though not the subject of this paper, is that children born into low-education and minority households experience relatively few years in which a stable father is present in the household, as their mothers repeatedly change adult partner relationships through cohabitation and remarriage (Cherlin, 2005). Children reared in single-headed households also fare relatively poorly on many behavioral and educational metrics including disciplinary suspensions, involvement in risky behaviors (e.g., smoking and early childbearing), and high school dropout (Bertrand and Pan, 2013; McLanahan, Tach and Schneider, 2013; Autor et al., 2014).

³Among Blacks, the comparable figures for children of high school dropout, exactly high school diploma, and post high school educated mothers were: 56 to 80 percent; 44 to 70 percent; and 41 to 56 percent. Among non-black Hispanics, the increase in mother-headedness was relatively evenly distributed across education groups: from 24 to 34 percent, 18 to 35 percent, and 20 to 31 percent, among children of high school dropout, exactly high school educated, and post high school educated mothers.

The Truly Disadvantaged (Wilson, 1987), and prior to that, the controversial 1965 U.S. Department of Labor report, *The Negro Family: The Case for National Action* (Moynihan, 1965).⁴ The traditional economic model of marriage implies that falling male earnings, rising female earnings, and increasing public support for unmarried mothers will generally reduce marriage rates and increase the prevalence of single-headed households (Becker, 1973, 1974). A sizable literature explores these ideas by testing whether and to what degree marriage and divorce rates respond to shifts in labor market demand or to changes in welfare benefits. It generally finds that better male labor market opportunities tend to increase marriage rates, while better female labor market opportunities tend to decrease marriage rates. The evidence for a discouragement effect of welfare policies on marriage rates is much less certain, however (Blau, Kahn and Waldfogel, 2000; Elwood and Jencks, 2004).⁵

A central question animating this literature is why declining marriage and rising single-parenthood have occurred disproportionately among less educated and minority adults. Following the lead of Wilson and Neckerman (1986) and Wilson (1987), scholars have pursued an additional implication of Becker’s model of marriage markets, which is that declining labor market opportunities, high rates of imprisonment, and a low ratio of available men to women will shift the marriage market in a direction that favors men. Specifically, by shrinking the pool of marriageable low education and minority men, the combined effects of mass incarceration and slackening labor markets may erode the incentive for men to maintain committed relationships, curtail women’s gains from marriage, and strengthen men’s bargaining position vis-a-vis casual sex, out-of-wedlock childbirth, and non-custodial parenting. Consistent with this hypothesis, Angrist (2002) and Charles and Luoh (2010) find that shifts in the ratio of available men to women—stemming from immigration (Angrist, 2002) or incarceration (Charles and Luoh, 2010)—shift the apparent distribution of gains from marriage in the direction predicted by theory. In settings where males are relatively abundant (i.e., a high ratio of males to females), marriage rates rise and the fraction of married women who participate in the workforce falls (Angrist, 2002). Conversely, where males are relatively scarce, women are more likely to work, more likely to have children out of wedlock, less likely to marry, and more likely to marry less educated males conditional on marriage (Charles and Luoh, 2010).⁶

The current paper contributes to this established literature by offering a direct and, arguably, unusually powerful test of the causal effect of labor market opportunities on the operation of mar-

⁴Elwood and Jencks (2004) provide a detailed overview of research on causes of rising single-headedness. See also Murray (2012) for a recent perspective.

⁵Changes in welfare policies are also an unlikely explanation for falling U.S. marriage rates given that the U.S. welfare system has become substantially less generous over the past two decades.

⁶Rich descriptive studies of household formation and child-rearing among poor, low education urban men and women in the U.S. provide abundant ethnographic evidence that appears consistent with this line of reasoning (LeBlanc, 2003; Edin and Kefalas, 2011; Edin and Nelson, 2013)

riage markets. Following Autor, Dorn and Hanson (2013*b*), we identify shocks to local employment opportunities using cross-industry and cross local labor market variation in import competition stemming from China’s rapidly rising productivity and falling barriers to trade. In exploiting localized employment shocks, our paper is closest in spirit to Ananat, Gassman-Pines and Gibson-Davis (2013), who use administrative data on county-level business closings and layoffs from the state of North Carolina to assess the impact of adverse employment shocks on teen birth rates.⁷ A specific innovation of our paper is to exploit differences in industry specialization by gender to identify trade shocks that differentially affect men’s and women’s employment and potential earnings. This allows us to test whether, in accordance with theory, shocks to male and female employment opportunities produce asymmetric affects on marriage rates and single-headedness. By pairing employment data from the U.S. Census and American Community Survey with Vital Statistics birth records data by year and county, we are further able to trace these shocks through not only to household formation but also to their effects on fertility and non-marital births. An additional contribution, which may be of independent interest, is that the research design allows us to assess specifically whether the recent two decades of contracting U.S. manufacturing employment in labor-intensive sectors, stemming in substantial part from rising international competition from China, has measurably contributed to the rapid, simultaneous decline of traditional household structures among less educated adults.⁸

Our principle results are as follows. Central to our identification strategy, we first demonstrate that trade-induced competitive shocks to manufacturing industries differentially affect employment of men and women according to whether the affected sectors are relatively male-intensive or female-intensive. Consistent with existing literature, we find that trade-induced declines in male employment decrease the fraction of young women ages 20 through 44 who are currently married or ever married. We find little robust evidence, however, that female-specific labor market shocks affect marriage rates—though, as theory would predict, these estimates are always opposite in sign to those of male-specific shocks (and also appear to significantly reduce divorces). Notably, in specifications that do not disaggregate labor market shocks into their sex-specific components, the estimated causal effect of employment shocks on marital status are only one-fourth to one-half as large as our main estimates, consistent with the proposition that male and female labor market opportunities have countervailing marriage market impacts.

Paralleling the asymmetric effect of labor market shocks on marriage markets, we document

⁷Ananat, Gassman-Pines and Gibson-Davis (2013) find that adverse shocks reduce birthrates and sexual activity among teens—particularly black teens—while increasing the use of contraception and abortion.

⁸Wilson (1996, p29) cites the erosion of U.S. manufacturing as the most significant economic factor affecting low-skilled African American workers, though he was not focused on import competition from China specifically.

that male and female-specific employment shocks have countervailing impacts on fertility. Adverse trade shocks reduce births to women ages 15 through 44 in net, but female-specific adverse shocks increase birth rates while male-specific adverse shocks depress birth rates alongside marriage rates.⁹ Adverse shocks to male employment, however, appear to contribute disproportionately to changes in child-rearing conditions. Specifically, they contribute to a rise in the fraction of births due to teenage mothers and to a rise in the probability that children under age 18 live in single-headed or below-poverty households.¹⁰

In net, our analysis strongly supports the hypothesis that changes in labor demand that reduce male employment opportunities—and in particular, those stemming from trade-induced contractions in manufacturing employment—contribute to declining marriage rates and rising single-headedness, with accompanying increases in child poverty.

1 Data and Measurement

1.1 Local labor markets

Our analysis requires a time-consistent definition of regional economies in the U.S. We approximate local labor markets using the construct of Commuting Zones (CZs) developed by Tolbert and Sizer (1996), who analyzed county-level commuting data from the 1990 Census data to create 741 clusters of counties that are characterized by strong commuting ties within CZs, and weak commuting ties across CZs. Our analysis includes the 722 CZs that cover the entire mainland United States (both metropolitan and rural areas). Commuting zones are particularly suitable for our analysis of local labor markets because they cover both urban and rural areas, are based primarily on economic geography rather than incidental factors such as minimum population.¹¹

1.2 Exposure to international trade

Following Autor, Dorn and Hanson (2013*b*), we examine changes in exposure to international trade for U.S. CZs associated with the growth in U.S. imports from China. The focus on China is a natural

⁹Ananat, Gassman-Pines and Gibson-Davis (2013) find that adverse shocks reduce birthrates and sexual activity among teens. Though not directly comparable, our results are potentially consistent with their findings in that they imply that adverse (male) employment shocks reduce both teen and adult fertility but reduce adult fertility by relatively more (thus increasing the teen fraction of births).

¹⁰The finding that adverse employment shocks predict an increase in the relative rate of teen births is consistent with Kearney and Levine (2012), who hypothesize that poor economic prospects of less educated U.S. women, stemming in part from high levels of U.S. inequality, contribute to the high U.S. teen birth rate.

¹¹Parts of our analysis draw on Public Use Microdata from Ruggles et al. (2004) that indicates an individual's place of residence at the level of Public Use Micro Areas (PUMAs). We allocate PUMAs to CZs using the probabilistic algorithm developed in Autor and Dorn (2013).

one: rising trade with China is responsible for nearly all of the expansion in U.S. imports from low-income countries since the early 1990s. China’s export surge is a consequence of its transition to a market-oriented economy, which has involved rural-to-urban migration of over 150 million workers (Chen, Jin and Yue, 2010), Chinese industries gaining access to long banned foreign technologies, capital goods, and intermediate inputs (Hsieh and Klenow, 2009), and multinational enterprises being permitted to operate in the country (Naughton, 2007).¹² Compounding the positive effects of internal reforms on China’s trade is the country’s accession to the World Trade Organization in 2001, which gives it most-favored nation status among the 157 WTO members (Pierce and Schott, 2012).

How can examining the impacts of trade exposure at the level of Commuting Zones be justified in terms of trade theory? Because trade shocks play out in general equilibrium, one needs empirically to map many industry-specific shocks into a small number of aggregate outcomes. For national labor markets at annual frequencies, one is left with few observations and many confounding factors. By taking regional economies as the unit of analysis, we circumvent the degrees-of-freedom problem endemic to estimating the labor-market consequences of trade. Our approach is valid for identifying the labor-market consequences of trade insofar as (i) CZs differ in their pattern of industry specialization, and (ii) frictions in labor markets allow regional differences in wages, unemployment, and labor-force non-participation to persist over the medium run. Autor, Dorn and Hanson (2013b) find strong evidence that greater exposure to trade with China affects local labor market outcomes across CZs.

Following the empirical specification derived by Autor, Dorn and Hanson (2013b), our main measure of local labor market exposure to import competition is the change in Chinese import exposure per worker in a region, where imports are apportioned to each region according to its share of national industry employment:

$$\Delta IPW_{uit} = \sum_j \frac{L_{ijt}}{L_{ujt}} \frac{\Delta M_{ucjt}}{L_{it}}. \quad (1)$$

¹²While China overwhelmingly dominates low-income country exports to the U.S., trade with middle-income nations, such as Mexico, may also matter for U.S. labor-market outcomes. The North American Free Trade Agreement (1994), for instance, lowered U.S. barriers to imports to a country in which U.S. firms already had extensive supply networks. Finding exogenous sources of variation in Mexico’s export growth, however, is tricky. Whereas China has had dramatic productivity growth in manufacturing—making internal supply shocks an important source of its export growth—Mexico has not (Hsieh and Klenow, 2012). The expansion of U.S. trade with Mexico is thus primarily driven by changes in U.S. bilateral trade policy which could be influenced by economic conditions in U.S. industries. Arguably, such simultaneity concerns are less an issue with regards to U.S. trade with China because of China’s phenomenal productivity surge, which has been due in large part to how far inside the global technology frontier the country remained at the end of the Maoist era. In recent work, McLaren and Hakobyan (2010) do not detect substantial effects of NAFTA on local U.S. labor markets, though they do find effects on wage growth nationally in exposed industries.

In this expression, L_{it} is the start of period employment (year t) in region i and ΔM_{ucjt} is the observed change in U.S. imports from China in industry j between the start and end of the period.

In equation (1), the difference in ΔIPW_{uit} across local labor markets stems entirely from variation in local industry employment structure at the start of period t . This variation arises from two sources: differential concentration of employment in manufacturing versus non-manufacturing activities and specialization in import-intensive industries within local manufacturing. Differences in manufacturing employment shares are not the primary source of variation, however; in a bivariate regression, the start-of-period manufacturing employment share explains less than 25% of the variation in ΔIPW_{uit} . In our main specifications, we control for the start-of-period manufacturing share within CZs so as to focus on variation in exposure to Chinese imports stemming from differences in industry mix within local manufacturing sectors.

The measure ΔIPW_{uit} measures overall trade exposure experience by CZs but does not distinguish between employment shocks that may differentially affect male versus female workers. To add this dimension of variation to ΔIPW_{uit} , we modify equation (1) to take account of the fact that manufacturing industries differ in their male and female employment intensity; hence, trade shocks of 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 male or female share of employment in each industry nationally (\bar{m}_{jt} and $1 - \bar{m}_{jt}$), thus apportioning the total CZ-level measure into two additive subcomponents, ΔIPW_{uit}^m and ΔIPW_{uit}^f :

$$\Delta IPW_{uit}^m = \sum_j \frac{\bar{m}_{jt} L_{ijt}}{L_{ujt}} \frac{\Delta M_{ucjt}}{L_{it}} \quad \text{and} \quad \Delta IPW_{uit}^f = \sum_j \frac{(1 - \bar{m}_{jt}) L_{ijt}}{L_{ujt}} \frac{\Delta M_{ucjt}}{L_{it}}, \quad (2)$$

Concretely, consider the hypothetical example of a CZ that houses two import-competing manufacturing industries, leather goods and rubber products, both of which employ the same number of workers and are exposed to industry-specific import shocks of \$500 per working age adult (thus, $\Delta IPW_{uit} = \$1,000$ for this CZ). Imagine that 55% of leather goods workers in U.S. employment are women while 75% of rubber products workers in U.S. employment are men. Equation (2) would apportion these the industry by commuting zone trade shocks to males and females according to their national industry employment shares such that $\Delta IPW_{uit}^m = 0.45 \times 500 + 0.75 \times 500 = \600 and $\Delta IPW_{uit}^f = 0.55 \times 500 + 0.25 \times 500 = \400 . In this example, we would assign a larger fraction of a CZ's trade shock to males than to females because males constitute a larger fraction of national employment in the CZ's trade-exposed industries. While the example is hypothetical, the numbers are quite close to the data. For the period of 1990 - 2000, our data indicate a mean rise of Chinese

trade exposure per U.S. worker of \$1,140, 55% of which accrued to male employment and 45% to female employment. In the subsequent 2000 - 2007 period, when Chinese import penetration accelerated sharply, trade exposure by an additional \$1,840 per worker, and 59% of this rise accrued to male employment.

A concern for our subsequent estimation is that realized U.S. imports from China in (1) may be correlated with industry import demand shocks. In this case, OLS estimates of the relationship between increased imports from China and changes in U.S. manufacturing employment may understate the true impact, as both U.S. employment and imports may be positively correlated with unobserved shocks to U.S. product demand. To identify the causal effect of rising Chinese import exposure on U.S. manufacturing employment and other local labor-market outcomes, we employ an instrumental variables strategy that accounts for the potential endogeneity of U.S. trade exposure. We exploit the fact that during our sample period, much of the growth in Chinese imports stems from the rising competitiveness of Chinese manufacturers (a supply shock from the U.S. producer perspective) and China’s lowering of trade barriers, dismantling of the constraints associated with central planning, and accession to the WTO. This approach requires that import demand shocks in high-income countries are not the primary cause of China’s export surge.

To identify the supply-driven component of Chinese imports, we instrument for growth in Chinese imports to the U.S. using the contemporaneous composition and growth of Chinese imports in eight other developed countries.¹³ Specifically, we instrument the measured import exposure variable ΔIPW_{uit} with a *non*-U.S. exposure variable ΔIPW_{oit} that is constructed using data on contemporaneous industry-level growth of Chinese exports to other high-income markets:

$$\Delta IPW_{oit} = \sum_j \frac{L_{ijt-1}}{L_{ujt-1}} \cdot \frac{\Delta M_{ocjt}}{L_{it-1}}. \quad (3)$$

This expression for *non*-U.S. exposure to Chinese imports differs from the expression in equation (1) in two respects. First, in place of realized U.S. imports by industry (ΔM_{ucjt}), it uses realized imports from China to other high-income markets (ΔM_{ocjt}). Second, in place of start-of-period employment levels by industry and region, this expression uses employment levels from the *prior* decade. We use 10-year-lagged employment levels because, to the degree that contemporaneous employment by region is affected by anticipated China trade, the use of lagged employment to apportion predicted Chinese imports to regions will mitigate this simultaneity bias.¹⁴ As instruments for our sex-specific

¹³The eight other high-income countries are those that have comparable trade data covering the full sample period: Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland. Our identification strategy is related to that used by Bloom, Draca and Van Reenen (2011), who consider the relationship between imports from China and innovation in Europe.

¹⁴By David H Autor, David Dorn and Gordon H Hanson (2013a) provide an extensive discussion of possible threats

labor demand measures, ΔIPW_{uit}^m and ΔIPW_{uit}^f , we construct analogues of (3) that multiply the employment measures in (1) by the corresponding lagged industry-specific gender shares nationally.

2 Results

2.1 The impact of trade shocks on employment, overall and by gender

The first panel of Table 1 presents initial results for the impact of trade shocks on employment. We fit models of the form

$$\Delta Y_{ijt} = \alpha_t + \beta_1 \Delta IPW_{uit} + \mathbf{X}'_{jt} \beta_2 + e_{ijt}, \quad (4)$$

where ΔY_{ijt} is the decadal change in the manufacturing employment share of the working-age population (ages 16 - 64) in commuting zone i among gender group j (males, females, or both) during time interval t , calculated using Census IPUMS samples for 1980, 1990 and 2000 (Ruggles et al., 2004), and pooled American Community Survey samples for 2006 through 2008. We estimate this model by stacking ten-year equivalent first differences for the two decades of our sample, 1990 to 2000 and 2000 to 2007, while including separate time dummies for each decade (in α_t).¹⁵ The variable of interest in this estimate is the change in CZ-level import exposure ΔIPW_{uit} , which is instrumented by ΔIPW_{oit} as described above. When we turn to gender-specific estimates, we replace ΔIPW_{uit} with ΔIPW_{uit}^m and ΔIPW_{uit}^f , and use the corresponding gender-specific instruments.¹⁶

To account for labor force and demographic composition factors that might independently affect manufacturing employment, the vector \mathbf{X}_{it} contains a rich set of controls for CZs' population and demographic structure at the start of each decade, including: the college-educated population share, the foreign-born population share, the female employment to population rate, the fraction of employment in manufacturing industries, and a set of eight Census geographic division dummies that allow for differential regional trends in manufacturing employment. We also include two variables from Autor and Dorn 2013 that measure CZ-level exposure to computerization and offshoring

to the validity of this approach, as well as a large set of robustness tests and complementary identification exercises.

¹⁵This stacked first difference estimating equation is analogous to a three-period fixed effects model in which the dependent variable is the manufacturing employment share (by CZ, year and gender) and the covariate vector includes Commuting Zone and year fixed effects. The first difference and fixed effects estimators are asymptotically equivalent as the number of periods grows. Estimating (4) as a fixed-effects regression assumes that the errors are serially uncorrelated, while the first-difference specification we use here is more efficient if the errors are a random walk (?). Since we apply Newey-West standard errors clustered on US state in all models, our estimates should be robust to either error structure

¹⁶Following Autor, Dorn and Hanson (2013b), we construct ΔIPW_{uit} by concurring UN Comtrade import data to 397 4-digit manufacturing industries and combining this trade data with employment counts from the County Business Patterns. The gender-specific shocks additionally use information on the male employment shares by industry, which are based on U.S. Social Security data and taken from unpublished results of Autor, Dorn, Hanson and Song (2014).

of job tasks. The first is the fraction of employment in routine task-intensive occupations, which numerous papers find is a strong predictor of machine-displacement of labor in codifiable clerical, administrative support, production and operative tasks (Autor and Dorn, 2013; Goos, Manning and Salomons, 2014; Michaels, Natraj and Reenen, 2014). The second is the mean index of ‘offshorability’ occupations in a CZ, where occupations are coded as offshorable if they do not require either direct interpersonal interaction with customers or proximity to a specific work location. Autor, Dorn and Hanson (2013a) discuss each of these variables in detail and find that none substantially affects the estimated relationship between (instrumented) CZ-level trade shocks and manufacturing employment. We conservatively include these variables in all specifications.

The column 1 point estimate of -0.596 on the import exposure measure indicates that a CZ that experiences a \$1,000 increase in per-worker import exposure is predicted to experience a fall of approximately 0.6 percentage points in the share of working-age adults employed in manufacturing. The \$1,000 increase is a natural benchmark as it is approximately the difference between the change in imports per worker in a Commuting Zone at the 75th percentile of trade exposure and one at the 25th percentile of trade exposure, when exposure measures are averaged over the full sample period. To gauge the magnitude of this effect, note that 12.7 percent of the working-age population was employed in manufacturing at the start of the sample in 1990, and that this share fell to 8.5 percent by 2007, a one-third reduction. In the same time period, mean trade exposure per U.S. worker rose by \$2,980, implying that the direct effect of trade exposure can explain roughly 1.8 percentage points (40 percent) of the 4.2 percentage point drop in the share of U.S. adults employed in manufacturing in this period.¹⁷ The next two columns perform estimates separately for male and female employment in manufacturing. The impact of point estimate -0.63 for males is slightly larger than the corresponding female estimate of -0.56 , partly reflecting the fact that roughly two-thirds of U.S. manufacturing workers are male throughout our sample period.¹⁸

The lower panel replaces the ΔIPW_{uit} measure with separate (instrumented) gender-specific exposure measures. Although these gender-specific measures are quite highly correlated across CZs, we detect distinct effects of each on manufacturing employment by sex.¹⁹ A \$1,000 rise in import

¹⁷Autor, Dorn and Hanson (2013a) adjust this estimate downward by half to account for the fact that only about 50 percent of the rise in U.S. exposure to Chinese imports during this period can be directly attributed to import supply shocks via the identification strategy described above. We provide rough benchmark numbers here since our objective is to characterize the effect of employment shocks on marriage and household structure, and not to account aggregate trends in U.S. manufacturing.

¹⁸This pattern could also indicate that male-intensive manufacturing industries are relatively more trade-exposed, but our data suggest that if anything, that the ratio of female to male exposure to Chinese import shocks exceeds the ratio of female to male employment in U.S. manufacturing. This pattern is consistent with the observation that Chinese import penetration is greatest in labor-intensive U.S. manufacturing sectors such as textiles, footwear, and apparel, which are relatively intensive in female production labor.

¹⁹The cross-CZ correlation is 0.90 for 1990 - 2000 and 0.87 for 2000 - 2007. These high correlations reflect the

exposure per male worker is estimated to reduce the fraction of working-age males employed in manufacturing by 1.3 percentage points ($t = 4.0$) but has essentially no impact on female manufacturing employment. Conversely, an equivalent rise in female import exposure reduces the fraction of working-age females in manufacturing by 1.4 percentage points ($t = 2.7$) while weakly increasing male employment in manufacturing (with an insignificant point estimate of 0.3). These estimates are of central importance for the analysis that follows because they indicate that we are able to identify economically meaningful and statistically distinct exogenous shocks to female and male employment. This finding provides the basis for exploring how changes in labor market conditions facing males versus females affect household formation and children’s living conditions.

2.2 The labor market and the marriage market

We now turn our attention to the impact of trade-induced labor market shocks on household structure, focusing initially on marriage. For this analysis, we estimate a variant of equation (4) using one of three dependent variables measuring the marital status of female CZ residents ages 20 through 44: the fraction never married; the fraction currently married; and the fraction widowed or divorced.²⁰ Due to data limitations, we focus our analysis of marriage outcomes to the time period 2000 - 2010, thus estimating a single decadal first-different for each CZ.²¹ Canonical theory makes no clear prediction for how aggregate labor market shocks will affect the operation of marriage markets since these shocks may simultaneously make men less marriageable and reduce women’s capacity to remain financially independent. The results in the upper panel of Table 2 are consistent with this ambiguity. We estimate that a \$1,000 per-worker rise in import exposure increases the fraction of women never married by 0.13 percentage points, and reduces the fraction who are currently married or divorced or widowed by 0.08 and 0.05 percentage points respectively. These three point estimates are all small in magnitude, and only the first is marginally significant.

Theory however makes the clear prediction that holding female labor market opportunities constant, a decline in male earnings potential should reduce marriages, a prediction that is born out

fact that CZs that are relatively manufacturing employment-intensive have large fractions of both males and females working in manufacturing.

²⁰We focus on women in this age bracket because they are most likely to change marital status (from single to married or from married to divorced).

²¹Because the trade shock measure only extends to 2007, we multiply the 2000/07 change by 10/7 to place it on a decadal footing. In the course of the analysis, we concluded that the IPUMS and American Community Survey data do not offer sufficient precision capture to the impacts of trade shocks on non-labor market outcome variables at the CZ-level, such as marriage and divorce. We thus make use of the Census and ACS Short Form Survey Tabulation Files (STF) of the U.S. Census, which are based on either a 15 percent or 100 percent extract of the underlying Census data (depending upon on the survey question). We currently only have STF data for 2000 and 2010, but will add in 1990 data shortly.

by the data. We estimate that a \$1,000 per worker rise in trade exposure concentrated on male employment increases the fraction of women ages 20-44 who have never been married by 0.25 percentage points, reduces the fraction currently married by 0.34 percentage points, and increases the fraction widowed or divorced by 0.1 percentage points. These estimated effects are not highly precise ($p < 0.10$ for the never-married and currently-married coefficients), but they are two to five times as large as the corresponding estimates that fail to account for the gender-specific components of trade shocks. Also consistent with theory, we find that adverse shocks to female labor demand have countervailing effects on marital statuses across all columns: modestly reducing the fraction of women never married, and substantially increasing the fraction currently married while reducing the fraction widowed or divorced. Though precision is also limited for these estimates, the panel 2 estimates paint a clear picture in which declines in male labor demand reduce marriage rates by reducing entry into marriage and modestly raising divorce, while, conversely declines in female labor demand raise marriage rates primarily by reducing divorce rather than by catalyzing additional marriages.

It is tempting to interpret the Table 2 estimates by scaling them relative to the Table 1 findings for the impact of trade shocks on manufacturing employment by gender. We caution against this scaling, however, since it would implicitly assume that the sole channel by which trade shocks affect marriage markets is through their impact on displaced manufacturing workers. In reality, trade-shocks that impact manufacturing employment affect the economic and, presumably, the marital prospects of CZ residents through multiple channels, as documented by Autor, Dorn and Hanson (2013a). These channels include reduced employment in local *non*-manufacturing industries, stemming from Keynesian demand spillovers; increases in the fraction of CZ residents applying for and receiving Social Security Disability Insurance benefits; and growth in public transfer benefits payments under the Unemployment Insurance, Trade Adjustment Assistance, Medicare and Medicaid, TANF, and SNAP programs, as well as through early Social Security retirement. These findings underscore that the set of CZ residents impacted by rising trade exposure at local manufacturing establishments is far larger than the set of workers initially employed in those establishments.

A more straightforward (and plausible) way to interpret these estimates is to scale them by the size of the trade shocks observed in our data. Using the figures reported above for the gender-specific changes in trade exposure that occurred between 2000 and 2007, our point estimates imply that rising import competition over this period reduced the fraction of women ages 20-44 currently married by 0.21 percentage points, reflecting the net impact of two countervailing effects of trade shocks on marriage formation and dissolution: a 0.35 percentage point rise in the fraction of

women never married and a 0.14 percentage point fall in the fraction of women currently widowed or divorced.²² These impacts are quantitatively modest relative to the net fall between 2000 and 2010 of -8.2 percentage points in the fraction of women ages 20 through 44 currently married. They nevertheless demonstrate that labor market shocks do impinge upon marriage market outcomes and lend credence to the hypothesis that the general decline in employment and earnings opportunities among U.S. males in this time interval may have contributed to the marked contemporaneous fall in marriage rates.

2.3 Impacts of labor market shocks on births

The U.S. birthrate has not fallen by nearly as much as has marriage during the last three decades, implying that a rising fraction of births occurs outside of marriage.²³ Since the decline in marriage rates and rise in non-marital births have been concentrated among low-education adults—a demographic group that has seen its wages and employment fall simultaneously—a natural hypothesis is that declining labor market opportunities contribute to both falling marriage rates and rising out-of-wedlock fertility. Our evidence above is qualitatively consistent with the first tenet of this hypothesis: adverse shocks to male employment reduce marriage. We now turn to the second. For this analysis, we use birth certificate data from the U.S. Vital Statistics to study the impact of trade-induced employment shocks overall and by gender on overall birth rates as well as births to teens and unmarried mothers. We again use a variant of the stacked first-difference estimating equation in (4), where here our dependent variables are decadal changes in Vital Statistics birth measures for the periods 1990 - 2000 and 2000 - 2010.²⁴

The first column of Table 3 considers the impact of labor demand shocks on the birthrate of women ages 15-44, defined as the number of live births per thousand women. These are calculated by denominating Vital Statistics live birth counts to women ages 15-44 by Census Short Form tabulations of county populations in the relevant age brackets. The estimate in the upper panel finds that an import exposure shock of \$1,000 per worker reduces total fertility by approximately 0.6 births per thousand women, which is roughly a 1 percent decline relative to the average rate over

²²Consistent with the logic of footnote 17, one should multiply these already small estimated impacts by approximately 0.5 to account for the fact that only about half of the observed increase in U.S. imports from China can be explained by import growth in other high-income countries.

²³While data from the National Vital Statistics System are not consistently available for as long a panel as desired, the fraction of all births due to unmarried women rose from 31.0 to 40.8 between 1993 and 2010, and this rise was evident in all racial and ethnic subgroups, except for Asians and Pacific Islanders. In the same time interval, the overall number of annual births per thousand U.S. women ages 15-44 held roughly constant at 67 (source http://www.cdc.gov/nchs/data_access/vitalstats/VitalStats_Births.htm, accessed 6/30/2014).

²⁴We rescale the 2000-07 trade shock measure to extrapolate a ten year change as in Table 2.

our sample of approximately 67 births per thousand women. When we disaggregate labor demand shocks into gender-specific components, we detect, consistent with earlier results, countervailing effects of male and female labor market conditions on outcomes. A \$1,000 per worker import shock concentrated on male employment is estimated to reduce births by 1.7 per thousand women ($t = 2.9$) while an equivalent shock concentrated on female employment is found to raise births by 1.1 per thousand women ($t = 1.9$). Since birth *rate* measures are potentially less precise than birth *count* measures, we provide an alternative estimate of the impact of labor market shocks on fertility by using as our dependent variable the percentage change in the count of births to women ages 15-44 by CZ.²⁵ As shown in column 2, we again find a modest negative effect of employment shocks on births (though in this case, the point estimate is not significant). When we subdivide labor market shocks by gender, however, we find substantially larger and more precise fertility impacts: male shocks of \$1,000 magnitude are estimated to reduce fertility by 4.7 percent ($t = 3.1$) while female shocks of the same order raise fertility by 4.6 percent ($t = 2.8$).

These sizable effects present an intriguing puzzle. While it is logical that declines in male earnings and employment opportunities would reduce fertility, it is less obvious why similar declines in female labor market opportunities would contribute to a rise in births. One possibility suggested by Kearney and Levine (2012) is that out of wedlock births—and teen births in particular—are partly driven by a lack of opportunity; young women who perceive few viable earnings or career opportunities see little benefit to delaying parenting and hence are more likely to give births as teens. We explore this idea further in columns 3 through 6 by regressing the percentage change in the number of births to married mothers, unmarried mother non-teen births (mothers 20-plus), and teen births (ages 15 to 19) on the trade shock measures.²⁶ These estimates also find countervailing effects of male and female labor market shocks on fertility among these subgroups: adverse shocks to male employment reduce births across all categories; adverse shocks to female employment raise births across all categories. However, the upper panel suggests that the falloff in births in response to the combined male and female shock is greatest for married and non-teen births, somewhat smaller for unmarried births, and actually opposite-signed (though insignificant) for teen births.

One implication of these results is that adverse economic conditions may tend to raise the *share* of births accruing to teens and unmarried parents since these categories of births appear relatively

²⁵These count measures occasionally contains zeros, which rules out a logarithmic specification. We apply the technique developed by Davis, Haltiwanger and Schuh (1996) for approximating percentage changes in variables where a start or end period value may equal zero: $\text{Pct Change} \equiv 2(n_2 - n_1) / (n_2 + n_1)$, where n_1 and n_2 are, respectively, the start-of-period and end-of-period count of births in each CZ. This measure has a range from -2 to $+2$ and is robust to either n_1 or n_2 taking a value of zero (though not both).

²⁶Note that most teen births are out-of-wedlock births, but that only a minority of out-of-wedlock births are teen births.

inelastic to economic shocks. This implication is confirmed by the final pair of columns in Table 3, which present results for the impact of economic shocks on unmarried and teen births as a *share* of all births. We estimate that trade shocks raise both the fraction of children born to unmarried mothers and the fraction of children born to teenage mothers, though only the latter point estimate is statistically significant. The magnitudes of these effects are modest, however. The point estimates imply that a \$1,000 per worker import shock raises the fraction of births due to teens by 0.13 percentage points, and 0.35 percentage points if the shock is concentrated solely on males. Approximately 11 percent of U.S. births in our sample are due to mothers under ages 20, and consistent with other published tabulations (Kearney and Levine, 2012), this fraction fell substantially in the last two decades (from 12.6 percent in 1990 to 11.5 percent in 2000 to 9.2 percent in 2010).

These results provide a nuanced set of inferences about the relationship between economic shocks and out-of-wedlock childbearing. Although the *relative* rates of non-marital and teen births rise when plausibly exogenous economic shocks reduce employment opportunities, this primarily reflects the fact that these categories of birth fall *less rapidly* than do in-wedlock and teen births in the face of comparable shocks. Thus, our results do not suggest that young and unmarried women increase childbearing when labor market conditions deteriorate. But since these groups are less responsive to these shocks, they account for a relatively larger fraction of all births during adverse conditions. As noted above, teen births declined rapidly during our sample frame (by almost 25 percent) even as non-marital births rose steeply (by almost 20 percent). Thus, the responses we document do not contribute to an explanation for the widely-remarked fall in the U.S. teen birth rate. But they may help us to better understand the rising incidence of non-marital childbearing among non-teens.

2.4 Impacts of labor market shocks on the household circumstances of children

We finally turn to the impact of labor market shocks on the household circumstances of children ages 0 to 17. We measure household circumstances along two dimensions: presence of parents (two, one or none), and household poverty status. Table 4 presents estimates for the period 1990 through 2007 using the stacked first-difference specification. The column 1 estimate suggests a sizable impact of labor market shocks on the incidence of household poverty among U.S. children. We find that a \$1,000K per worker import shock robustly raises the fraction of children under age 18 living below the poverty line by approximately 0.7 percentage points ($t = 5.0$), which is roughly a four percent increase relative to the base poverty rate of U.S. children in 1990. Columns 2 through 4 suggest that some—but by no means all—of this poverty effect stems from changes in household structure.

A \$1,000 import shock is estimated to reduce the fraction of children in two-parent households by 0.18 percentage points, about two-thirds of which is accounted for by a rise in the fraction living in single-headed households, with the remainder due to a rise in children living apart from their parents (i.e., with relatives or under institutional care). These estimated impacts should be viewed as merely suggestive, however, since they are for the most part statistically insignificant.

A striking pattern conveyed by the lower panel of Table 4 is that almost the entirety of the impact of adverse economic shocks on the household structure of children arises from shocks to male employment. Indeed, we find no significant effect of female labor market shocks on the incidence of household poverty, single-headedness, or non-parental living arrangements among children under 18. By contrast, shocks to male employment substantially raise childhood poverty rates, reduce the fraction of children living in two-parent households, and raise the fraction of children living in single-parent households. For all three outcome measures—poverty, dual-headedness and single-headedness—the impact of male-specific shocks on children’s circumstances are more than twice as large as the upper panel estimates that do not distinguish employment shocks by gender. This result lends credence to the hypothesis that deteriorating employment conditions faced by males contribute to an increase in the share of U.S. children living in poor and single-headed households. Focusing exclusively on the influence of shocks to male employment opportunities stemming from trade, we calculate using the point estimates in the lower panel of Table 4 that these shocks raised the share of U.S. children living in poverty by 2.4 percentage points (13%) and increased the share living in single-parent households by 0.8 percentage points (3.3%) between 1990 and 2007.

3 Conclusions

The four sets of analyses in this paper provide a coherent narrative for the impact of labor market shocks on fertility and household structure. Adverse shocks to local employment opportunities stemming from rising competition in manufactured goods from China lower employment, reduce entry into marriage, increase the fraction of births accruing to teen and unmarried mothers, and raise the fraction of children living in poverty and single-headed households. Many of these results are expected, and their economic magnitudes are modest. The central finding of the paper, however, is that gender-specific shocks to labor market outcomes have strikingly non-parallel impacts on marriage, births, and children’s household circumstances. While female-specific trade shocks reduce female manufacturing employment by roughly the same amount as male-specific trade shocks reduce male manufacturing employment, the similarity of their impacts ends there.

Male-specific shocks sharply reduce the fraction of women entering marriage and slightly increase the fraction divorcing. While these shocks reduce fertility overall, they reduce fertility by less among teens and unmarried women, thus increasing the fraction of children that are born to teens and, to a lesser extent, to unmarried mothers. Perhaps most striking is the outsized impact of shocks to male employment on the living circumstances of children. Since male employment shocks reduce both marriages and fertility, their net impact on the household structure experienced by children is potentially ambiguous. As an empirical matter, however, the marriage-reducing effects of male-specific labor market shocks appears to dominate the dampening effect on fertility. We estimate that a \$1,000 per-worker male-specific employment shock raises the fraction of children living in single-headed households by half a percentage point (a 2% rise), with essentially all of this rise coming from a reduction in two-parent households rather than a rise in non-parent households, and increases the fraction of children living in poverty by 1.4 percentage points. Since the induced rise in poverty is roughly three times as large as the rise in single-headedness, we infer that a substantial share of the poverty effect stems from higher rates of poverty among married households. By contrast, while adverse shocks to female manufacturing employment generally increase marriage rates and raise fertility—opposite to shocks to male employment—these impacts do not significantly affect the fraction of children living in poverty or in single-headed households.

In summary, a crucial takeaway of our analysis is that the declining labor market faced by U.S. males is a plausible contributor to the changing structure of marriage and childbirth in the United States. While our analysis does *not* suggest that surging import competition from China over the last two decades has been a major contributor to these trends, it highlights that broader declines in the labor market for U.S. males may have made a substantial contribution—assuming of course that these declines exert qualitatively similar causal effects on marriages, births and household structures as do trade shocks concentrated on male employment. Future work on this paper will attempt to quantify these linkages.

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Table 1. Imports from China and Employment by Sector, 1990-2007: 2SLS Estimates.

Dependent Var: 100 x Change in Share of Working Age Population Employed in Indicated Sector (in %pts).

| | Share Working Age Pop in Manufacturing | | | | | |
|---|--|-----|-------------------|-----|-------------------|-----|
| | All | | Males | | Females | |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| <u>I. Overall Trade Shock</u> | | | | | | |
| (Δ Imports from China to US)/Worker | -0.596 (0.099) | ** | -0.625 (0.123) | ** | -0.555 (0.088) | ** |
| <u>II. Male Industry vs Female Industry Shock</u> | | | | | | |
| (Δ Imports from China to US)/Worker (Male Industries) | -0.626 (0.269) | * | -1.275 (0.320) | ** | 0.021 (0.278) | |
| (Δ Imports from China to US)/Worker (Female Industries) | -0.552 (0.370) | | 0.332 (0.338) | | -1.400 (0.521) | ** |

Notes: N=1444 (722 CZ x 2 time periods). All regressions include the full vector of control variables from China Syndrome. Regressions are weighted by start-of-period population and standard errors are clustered on state. $\sim p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Table 2. Imports from China and Marital Status, 2000-2010: 2SLS Estimates.
 Dependent Var: 100 x Change in Share of Women Age 20-44 with Indicated Marital Status
 (in %pts).

| | Share of Women Age 20-44 | | |
|---|---|---------------------|--------------------|
| | Never Married | Married | Divorced/Widowed |
| | (1) | (2) | (3) |
| | <u>I. Overall Trade Shock</u> | | |
| (Δ Imports from China to US)/Worker | 0.132 (0.078) | ~ -0.078 (0.070) | -0.054 (0.046) |
| | <u>II. Male Industry vs Female Industry Shock</u> | | |
| (Δ Imports from China to US)/Worker (Male Industries) | 0.252 (0.151) | ~ -0.343 (0.184) | ~ 0.091 (0.093) |
| (Δ Imports from China to US)/Worker (Female Industries) | -0.045 (0.200) | 0.312 (0.263) | -0.266 (0.150) |

Notes: N=722 CZ. All regressions include the full vector of control variables from China Syndrome, and control for start-of-period population shares in 9 age groups and 5 race/ethnicity groups. Regressions are weighted by start-of-period CZ population and standard errors clustered on state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Table 3. Imports from China and Birth Outcomes, 1990-2010: 2SLS Estimates.
 Dependent Var: 100 x Change in Birth Rate or Share of Births with Indicated Conditions (in %pts).

| | Births to Women 15-44 | | Percentage Point Change in # Births | | | | Share of Births to | | |
|---|-----------------------|------------------------|-------------------------------------|----------------------|----------------------|----------------------|--------------------|-------------------|----|
| | Change in Birth Rate | Pct Change in # Births | Married Mothers | Unmarried Mothers | Non-Teen Mothers | Teenage Mothers | Unmarried Mothers | Teenage Mothers | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | |
| <u>I. Overall Trade Shock</u> | | | | | | | | | |
| (Δ Imports from China to US)/Worker | -0.586 (0.217) | ** -0.971 (0.602) | -1.247 (0.809) | -0.722 (0.434) | ~ -1.117 (0.636) | ~ 0.372 (0.416) | 0.154 (0.164) | 0.126 (0.056) | * |
| <u>II. Male Industry vs Female Industry Shock</u> | | | | | | | | | |
| (Δ Imports from China to US)/Worker (Male Industries) | -1.734 (0.589) | ** -4.747 (1.541) | ** -5.259 (1.834) | ** -5.161 (1.877) | ** -5.155 (1.571) | ** -1.082 (1.223) | 0.254 (0.327) | 0.348 (0.110) | ** |
| (Δ Imports from China to US)/Worker (Female Industries) | 1.101 (0.579) | ~ 4.580 (1.611) | ** 4.652 (1.696) | ** 5.804 (2.589) | * 4.820 (1.647) | ** 2.508 (1.723) | 0.006 (0.464) | -0.200 (0.139) | |

Notes: N=1444 (722 CZ x 2 time periods). Regressions weighted by start-of-period CZ population. Birth rate in column (1) is equal to the number of live births per 1,000 women ages 15-44. Percentage point changes in columns 2 - 6 are calculated using the Davis-Haltiwanger-Schuh (1996) method: percentage change = two times the end-of-period minus start-of-period level divided by the sum of the end-of-period level and the start-of-period level. Standard errors clustered on state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.

Table 4. Imports from China and Socio-Economic Outcomes for Children and Young Adults, 1990-2007:
2SLS Estimates.

Dependent Var: 100 x Change Share of Age Group in Indicated Condition (in %pts).

| | Share Children Age 0-17 | | | | | |
|--|---------------------------|----------------------------|-------------------------------|---------------------------|-------------------|---------------------|
| | in HH with Inc<Poverty | in Two-Parent Household | in Single-Parent Household | in No-Parent Household | | |
| | (1) | (2) | (3) | (4) | | |
| <u>I. Overall Trade Shock</u> | | | | | | |
| (Δ Imports from China to US)/Worker | 0.683 (0.136) | ** | -0.179 (0.113) | 0.111 (0.106) | 0.069 (0.038) | ~ |
| <u>II. Male Industry vs Female Industry Shock</u> | | | | | | |
| (Δ Imports from China to US)/Worker (Male Industries) | 1.393 (0.416) | ** | -0.456 (0.247) | ~ | 0.466 (0.234) | * -0.009 (0.126) |
| (Δ Imports from China to US)/Worker (Female Industries) | -0.360 (0.471) | | 0.228 (0.316) | | -0.411 (0.276) | 0.183 (0.175) |

Notes: N=1444 (722 CZ x 2 time periods). All regressions include the full vector of control variables from China Syndrome, and control for start-of-period population shares in 9 age and 5 race/ethnicity groups. Regressions are weighted by start-of-period CZ population and standard errors are clustered by state. ~ $p \leq 0.10$, * $p \leq 0.05$, ** $p \leq 0.01$.