Empowered Young Women: Trade Liberalization and Women's Family Decisions in China *

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Abstract

Do trade-induced labor market opportunities affect women's marriage and fertility decisions? Exploiting regional variation in the exposure to the U.S. granting of Permanent Normal Trade Relations (PNTR), we find that more exposed Chinese prefectures experience a relative increase in the fraction of unmarried young women. This relative increase is due to young women delaying their first marriage and more married women choosing to divorce. The share of young women with children, as a result of changed marriage decisions, also experiences a relative decline in more exposed areas. We show that these shifts in family decisions coincide with a trade-induced increase in female workforce participation and reallocation of women relative to men to the service sector, where wages are higher.

JEL Classification: F16, J12, J13, J16

Keywords: Trade Liberalization; Gender Inequality; Family Decisions, Chinese Economy

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1 Introduction

How do family decisions respond to globalization? Autor et al. (2019) finds that import competition from China reduces marriage and fertility in the U.S. What is the other side of this coin? Does export expansion necessarily increase marriage and fertility in China? More generally, does international trade necessarily reduce marriage and fertility in one country and increase them in another, thus producing a global redistribution of marriages and gender power, just as it does for manufacturing jobs?[1] This paper aims to answer this question, and surprisingly finds the answer is no.

In this paper, we document the link between Chinese young women's marriage and fertility outcomes and a large, plausibly exogenous shock to local labor markets driven by a change in U.S. trade policy: the granting of permanent normal trade relationship (PNTR) to China in 2000. This trade shock differentially exposes Chinese regions to increased export opportunities via their initial industry structure. While Chinese exports to the U.S. have already been subject to low normal trade relations (NTR) tariffs prior to 2000, these tariff rates are subject to annual renewal, and the possibility of tariff increases due to China's loss of NTR status has created significant uncertainty for the U.S.-China trade prior to the policy change (Pierce and Schott, 2016).

Exploiting regional variations in PNTR exposure stemming from the initial differences in industry structure, we find that Chinese prefectures more exposed to PNTR experience a relative decline in the fraction of married young women,[2] primarily among women with low levels of education or rural hukou. Coefficient estimates suggest that the effects are sizable: a prefecture at the 90th percentile in terms of the PNTR exposure experiences a 2.7 percentage point larger decline in the fraction of married young women compared to a prefecture at the 10th percentile. A simple back-of-the-envelope calculation indicates that PNTR shock accounts for 37% of the observed decline in the share of married young women in China between 2000 and 2010. Noting that women aged 20 to 39 may not have finalized their decisions to marry, hence we interpret these findings as evidence that PNTR has reduced or *delayed* women's marriage decisions in China.

Our findings are robust to the inclusions of other contemporaneous shocks and various start-of-period prefecture characteristics, and there is no pre-existing trend in more exposed

[1] Autor et al. (2013), Pierce and Schott (2016) find that U.S. imports from China and the U.S. granting of PNTR to China reduce U.S. manufacturing jobs, while Erten and Leight (2021), Ouyang and Yuan (2019) find a corresponding manufacturing job gain in China.

[2] The young women (men) refer to those aged 20-39 throughout the paper.

prefectures. A decomposition reveals that the decline in the fraction of married young women is jointly contributed by the rise in the fraction of those who are divorced and those who have never been married, especially the latter. As expected, the fraction of young married men also declines in prefectures with greater exposures. However, when gender-specific trade shocks are introduced, only female-specific trade shocks matter for both young women and men, demonstrating that the effect of PNTR on marriage is through affecting young women in China.

The decline in the fraction of married young women naturally has an impact on regional fertility outcomes. We find that PNTR also reduces the share of young mothers by a similar magnitude as the share of married young women. This synchronous pattern may stem from China's hukou system, which makes it difficult for children of unmarried women to access local welfare services.[3] Specifically, a prefecture at the 90th percentile in terms of the PNTR exposure experiences a 2.1 percentage point decrease in the share of young mothers compared to a prefecture at the 10th percentile. Again, because women aged 20-39 are still able to have children, we interpret this result as evidence that PNTR reduces or *delays* childbearing. Although statistically insignificant, we find that PNTR increases the number of children per young woman with children, possibly due to the fact that PNTR also increases the income of local households. In other words, PNTR has opposite effects on the extensive margin (fraction of young mothers) and the intensive margin (number of children per young mother). This results in the impact of PNTR on the number of children per young woman at a region statistically indistinguishable from zero.

Why do trade shocks contribute to the decline or delay in marriage and fertility decisions? We lay out five hypotheses based on the trade-induced major economic or demographic changes that occurred at the time. The first hypothesis, "economic stature," assumes that trade liberalization may affect marriage and fertility outcomes by changing the economic status of young women relative to men. The second hypothesis, "the role of education," posits that trade liberalization may affect women's family decisions by influencing their school enrollment or educational attainment. The third hypothesis is that trade may reduce marriage rates by creating a more gender-segregated labor market. The fourth hypothesis is that trade may result in a spatial Balassa-Samuelson effect, thus raise the cost of marriage and childbearing more in trade-exposed regions. Finally, China's trade liberalization has led to a massive reallocation of labor across regions. Young people's marriage and fertility rates could also be affected if migration creates compositional imbalances in the spatial dimension

[3]This institutional setup also explains why unmarried mothers are almost nonexistent in China.

(e.g., gender or local hukou), or if migrants are less likely to marry due to economic and institutional constraints.

After providing direct and indirect evidence for or against each hypothesis, we find that the economic stature hypothesis appears to be most consistent with data patterns. Young women have been positively, and differentially affected in two major ways. Firstly, in prefectures that are more exposed to PNTR, there is a greater increase in women workforce participation, which is mostly coinciding with the decrease in the share of women who do not work due to family obligations. Secondly, while the granting of the PNTR results in a reallocation of young men from the service sector to manufacturing, the opposite is true for women: the more exposed prefectures see only a very small increase in female manufacturing employment, while more young women move to the higher paying service sector. We find no evidence to support the other hypotheses.

Our paper contributes to the large literature on the economics of marriage and fertility.[4] In linking globalization to marriage and fertility outcomes, Autor et al. (2019) finds import competition from China reduces marriage and fertility in the U.S. by differentially reducing employment and earnings of young adult males. Keller and Utar (2022) finds the same shock contributes to gender disparity in Denmark by influencing women's fertility decisions near the end of their biological clock. In a developing country's context, our work is close in spirit with Jensen (2012), who finds that young women in villages that are randomly getting jobs in processing outsourcing are significantly less likely to get married or have children during this period in rural India, and Heath and Mobarak (2015) and Kis-Katos et al. (2018), who find that young women delay their marriage and childbirth decisions in response to job opportunities created by positive trade shocks in Bangladeshi and Indonesia, respectively. Consistent with these studies, our results are primarily driven by women from rural areas with low levels of education, and we find that the observed decrease in marriage is consistent with a relative improvement in women's job opportunities. Among the other potential channels, however, we do not find evidence that education (Heath and Mobarak, 2015; Jensen, 2012) or sectoral gender segregation (Kis-Katos et al., 2018) plays a role.[5] These differences may be due to the fact that China has a very different institutional and

[4]Chiappori (2020) and Doepke et al. (2022) provide reviews of the relevant literature.

[5]In particular, Jensen (2012) finds that young women in villages that are randomly getting jobs in processing outsourcing are significantly less likely to get married or have children during this period, choosing instead to enter the labor market or obtain more schooling or post-school training. Similarly, Heath and Mobarak (2015) finds that in villages closer to export processing boom in ready-made garment industries, girls exposed to the garment sector delay marriage and childbirth, which stems from young girls becoming more likely to be enrolled in school after garment jobs (which reward literacy and numeracy) arrive, and older girls becoming more likely to be employed outside the home in garment-proximate villages.

economic setting. For example, the legal age of marriage for Chinese women is 20 years old, at which age the vast majority of women have finished high school and pre-high school education. Whereas in, for example, Jensen (2012), the median age of marriage for Indian women is 17, which means that on average women in their sample are much younger. The nine years of compulsory education in China, which has been in place since 1986, means that most women are educated enough to become blue collar workers. Manufacturing jobs in China are also concentrated in urban rather than rural areas, so the labor market is less segregated across sectors. Finally, our analysis also identifies a novel general equilibrium effect, namely women are differentially more employed in the service sector. In this regard, our finding also connects to Olivetti and Petrongolo (2014), Ngai and Petrongolo (2017), and Ngai et al. (2020), which highlight the role of production in services in explaining the gender gap.

Existing research suggests that trade-induced improvements in male labor market conditions lead to higher fertility, while the opposite is true for improvements in female labor market conditions (Schultz, 1985; Schaller, 2016; Giuntella et al., 2022; Do et al., 2016; Li, 2021).[6] In contrast, we find a positive, although imprecisely estimated, response in the number of births among young women with children. This may be due to the fact that the one-child policy has artificially depressed China's fertility rate at a very low level. The trade-induced income improvements alleviate economic pressures caused by the cost of raising children and the additional cost of having more than one child associated with the one-child policy, thereby encouraging some families to have more children.

This paper finds that trade liberalization is most likely to affect the marriage and fertility outcomes of young Chinese women by changing their relative economic status. This result also complements existing research on the differential impact of trade on male and female workers. In the context of other countries, Aguayo-Tellez et al. (2014) finds that tariff reductions in the aftermath of the NAFTA agreement expand sectors that are initially female intensive and increase the hiring of women in skilled blue-collar occupations in Mexico. Similarly, Juhn et al. (2013) and Juhn et al. (2014) find the same shock improves the relative wage and employment of female blue-collar workers. The findings from other countries are

[6] Schultz (1985) finds that improved women's wages relative to men due to world commodity price changes contribute to the decline of fertility in Sweden. Giuntella et al. (2022) finds that in Germany, workers in industries that benefit from increased exports have better employment prospects and higher fertility, and these results are driven by low-educated, married men and full-time workers. Schaller (2016) finds that improvements in men's labor market conditions are associated with fertility increases, while improvements in women's labor market conditions have negative effects in the U.S. Cross-country studies also show that countries with a comparative advantage in female labor-intensive sectors tend to have lower fertility (Do et al., 2016; Li, 2021) and more equal gender norms (Li, 2021).

nevertheless mixed. Connolly (2022) shows that import demand from China generates larger employment gains for female workers in Brazil. Gaddis and Pieters (2017) finds that reduced trade protection due to Brazil's trade liberalization narrows the gender gap in employment in levels but not in proportion terms. In the case of Peru, Mansour et al. (2022) finds that Chinese competition has a negative and persistent impact on female workers only and thus widens the gender gap in employment, the opposite of Connolly (2022) and Gaddis and Pieters (2017). The paper closest to our spirit in this line of research is Molina and Tanaka (2020), where the authors find that women living near exporting factories in Myanmar are significantly more likely to be working, have lower tolerance of domestic violence, and are less likely to be victims of domestic violence.

In the international trade literature, there is a large body of work that uses the granting of PNTR as an exogeneous trade shock and examines its impacts on various economic outcomes. This policy shift is found to have caused declines in U.S. employment (Pierce and Schott, 2016), consumer prices (Handley and Limão, 2017; Amiti et al., 2020), and domestic investment (Pierce and Schott, 2018), as well as altered patterns of voter turnout (Che et al., 2022). Pierce and Schott (2020) also finds that U.S. areas that are more exposed to PNTR shocks see a relative increase in fatal drug overdoses, particularly among whites. On the Chinese side, regions that are more exposed to PNTR shocks are characterized by increased exports (Pierce and Schott, 2016; Amiti et al., 2020) and improved export participation of firms (Zhou and Zhang, 2021), increased foreign direct investment, a shrinking agricultural sector, an expanding secondary sector, and increases in total GDP and GDP per capita (Erten and Leight, 2021). The more exposed regions also see greater migration inflows (Facchini et al., 2019), more innovative activity (Liu and Ma, 2020), and upgrading of skills (Zhou and Zhang, 2021) and product quality (Feng et al., 2017). None of these papers, however, examine non-labor market outcomes nor the gender-specific effects of the trade shock.

Finally, our paper contributes to the literature on how trade liberalization affects non-labor market adjustments in China. The unprecedented growth of Chinese exports following China's accession to the WTO has generated a large literature documenting significant trade-induced labor market responses in China, ranging from employment and internal migration (Zi, 2020), regional adjustments in wages (Dai et al., 2020), sectoral worker reallocation (Ouyang and Yuan, 2019) to legislative changes (Tian, 2019). In a smaller set of analyses focusing on non-labor market outcomes, Fan et al. (2020) finds a negative link between input trade liberalization and worker health. Crozet et al. (2018) finds that increased export

opportunities have a positive impact on the perceived life satisfaction of Chinese adults. Tang and Zhang (2021) finds that foreign affiliates from countries with a more gender-equal culture tend to employ proportionally more women and appoint more female managers in China. Dai et al. (2021) finds that in regions adversely affected by Chinese tariff cuts, households respond to trade-induced wage declines by increasing labor force participation of women and the elderly, reducing savings, and opting to live with parents more frequently. Li (2018) and Lin and Long (2020) document a wide range of educational adjustments across Chinese regions in response to export expansions. As part of our discussion of potential mechanisms, we find that the PNTR shock increases Chinese young women's labor force participation, but has no significant effect on their educational attainment, thus complementing the work of Dai et al. (2020) and Li (2018). We add to this literature by examining how trade liberalization affects marriage and fertility decisions, and by uncovering the possible causes of the trade-induced delay and decline in young women marriage.

The rest of the paper is organized as follows. Section 2 describes the data and the empirical strategy, Section 3 presents the main results and robustness checks, and Section 4 discusses mechanisms. Section 5 concludes.

2 Data and Empirical Strategy

2.1 Marriage and Fertility Rates across Prefectures

We compute the number of individuals in 337 time-consistent Chinese prefectures by demographic categories using 1‰ random-sampled microdata from the 2000 and 2010 Chinese Population Census. Observable demographics include age, gender, marital status, fertility history (female only), hukou status, education, migration, employment, and prefecture of residence. We use these information to calculate variables of interest, such as the decadal change in the fraction of young women (aged 20-39) who are married in each prefecture. For some analyses, we complement the data using variables computed from the 1% random-sampled microdata of the 1990 census.[7] Summary statistics of the variables are presented in Table A1.

The primary focus of the paper is on the marriage and childbearing decisions of young women. In 2000, the median age of first marriage for women aged 20 to 39 is 22. Divorce rates are low (0.81%), but there is substantial regional variation (s.d. = 0.79%). Between

[7]Note that China's administrative divisions undergo considerable changes between 1990 and 2000, so the corresponding time-consistent geo-boundary in this case may have more measurement errors.

2000 and 2010, the fraction of the young women who are unmarried rises from 15.72% to 27.09%, almost entirely driven by the increases in never-married women and the incidences of divorce.[8] The fraction of young women without children also increases significantly from 20.63% to 33.00% during the same period.

Among those who ever married, the median age at the first marriage is deterred slightly to 23. Alongside these facts, the share of young women in school has increased by about 4.46 percentage points, from 1.11% to 5.57%. Among the rest, women's paid labor force participation has decreased slightly, from 84.54% to 80.17%. Much of this decline in overall labor force participation may be the result of overpopulation in the agricultural sector in the initial year (Lewis, 1954). Between 2000 and 2010, the employment of women in the manufacturing and service sectors increase by a total of 19.61 percentage points.[9]

Figure A1 presents the decadal change in the married young women shares across Chinese regions. The decline is relatively concentrated in the central and coastal regions, with the notable exception of the Hakka regions (eastern Guangdong, southern Guangxi, and western Fujian), which imposes a more patriarchal family for historical and cultural reasons and emphasis on female marriage and childbearing. Across prefectures, the unweighted decadal decline in the married young women shares averages 0.09 and has a standard deviation of 0.06 (Table A1, row 1). The fraction of young women without children increases by an average of 10 percentage points (Table A1, row 8); its regional variation is highly correlated with that of the share of young women never married, as shown in Figure A3.

2.2 Measuring Exposure to PNTR

Our measure of PNTR exposure is based on two sets of tariff rates in the U.S. tariff schedule. The first set, known as the NTR tariff, applies to goods imported from World Trade Organization (WTO) members. The second, known as the non-NTR tariff, is established by the Smoot-Hawley Tariff Act of 1930 and is typically larger than the corresponding NTR rate. Prior to China's accession to the WTO, the President of the United States grants China a temporary access to NTR rate, but subject to annual congressional approval.

As discussed in detail in Pierce and Schott (2016), the possibility of tariff increases due to the withdrawal of China's NTR status has created significant uncertainty for U.S.-

[8] The increase in the fraction of never-married women and the increased incidences of divorce together account for 99.69% of the increase in the share of the young women unmarried between 2000 and 2010.

[9] We find a similar pattern in male employment. Between 2000 and 2010, the employment of men in the manufacturing and service sectors increase by a total of 18.32 percentage points.

China trade.[10] U.S. companies doing business in China cite "uncertainty surrounding the annual renewal of China's most-favored-nation trade status as the single most important issue affecting U.S. trade relations to China," according to the U.S. General Accounting Office (1994, p. 3). This uncertainty is removed with the U.S. granting of PNTR status to China in 2000, which becomes effective after China's accession to the WTO in 2001.[11] This trade policy change has created new export opportunities for Chinese firms and can explain more than a third of the increase in Chinese exports to the U.S. from 2000 to 2005, the largest export market for Chinese goods during the period (Handley and Limão, 2017).

Following Pierce and Schott (2016), we measure the trade uncertainty as the gap between non-NTR and NTR rates, i.e. the rise in U.S. tariffs on Chinese goods that would have occurred in the event of a failed renewal of NTR status:

$$NTRGap_j = NonNTRRate_j - NTRRate_j.$$

We refer to this difference as the NTR gap, and compute it for each 4-digit Chinese manufacturing (CIC4) industry j using the *ad valorem* equivalent NTR and non-NTR rates provided by Feenstra et al. (2002) and the crosswalk by Brandt et al. (2017).[12]

We then compute regional exposure to PNTR as the initial-employment-share weighted average NTR gap across the 4-digit manufacturing industries active in a prefecture c:

$$NTRGap_c = \sum_j \frac{L_{cj99}}{L_{c99}} \times NTRGap_j, \tag{1}$$

where c indexes prefectures, j indexes industries, and L presents employment. The fraction L_{cj99}/L_{c99} presents the share of industry j in total manufacturing employment in prefecture c in 1999, which is computed using the Chinese Annual Survey of Industrial Firms (ASIF). Across prefectures, the unweighted NTR gap averages 0.25 and has a standard deviation of 0.07 (Table A1). As illustrated in Figure A2, China's coastal regions are among the biggest beneficiaries of trade policy changes but there are also a reasonable number of prefectures

[10] Following the Chinese government's 1989 crackdown on the Tiananmen Square protests, the U.S. House of Representatives passes resolutions to terminate China's NTR status in 1990, 1991, and 1992, although the Senate does not.

[11] In October 2000, the U.S. grants China PNTR status - permanent Most Favored Nation (MFN) status in the U.S. terminology. This status becomes effective on January 1, 2002, following China's accession to the WTO in December 2001.

[12] Specifically, we use the *ad valorem* equivalent NTR and non-NTR rates provided by Feenstra et al. (2002) at the HS8 level for 1999, the year before PNTR is granted to China. We first aggregate these two rates to the HS6 level and then calculate the differences, then aggregate to the CIC4 level using a simple average. There are 423 CIC4 industries in total.

in Middle and Western regions with large PNTR exposures.

We also explore trade shocks that differentially affect female and male workers. To include this gender dimension of variation, we follow Autor et al. (2019) and modify (1) by multiplying the initial period female or male share of employment in each industry by prefecture, apportioning the prefecture-level overall exposure into two additive subcomponents, $NTRGap_c^f$ and $NTRGap_c^m$:

$$\begin{aligned} NTRGap_c^f &= \sum_j \frac{f_{cj} L_{cj99}}{L_{c99}} \times NTRGap_j, \\ NTRGap_c^m &= \sum_j \frac{(1-f_{cj}) L_{cj99}}{L_{c99}} \times NTRGap_j, \end{aligned} \tag{2}$$

where f_{cj} is the initial female employment share in industry j, prefecture c. Because the ASIF data does not include information on employment by gender, we use the 2000 Census to calculate f_{cj}. The census reports employment at the 3-digit industry level, so f_{cj} reflects the share of female employment in the 3-digit CIC industry to which j belongs.

2.3 Other Variables

We control for several changes in Chinese policy: regional exposure to Chinese import tariff cuts and input trade liberalization associated with the country's accession to the WTO, as in Zi (2020). We also control for start-of-period values of several prefecture demographic and development attributes: manufacturing employment share and rural population share, as controls for the potential effect of structure change and urbanization; fraction of female employed (aged 20-39) and minority population share, to account for the possibility that social or cultural norms may lead to differential changes in family decisions across prefectures; the fraction of population with college or higher education, to help identify region more exposed to the introduction of labor-saving technical change; a dummy indicating whether a prefecture has export processing zones, which accounts for potential changes in FDI or processing trade policy. We also use income information by gender and sector from the 2005 Census, wage by industry from the 1999 ASIF data, and housing and education information from regional and city statistical year books when discussing potential mechanisms. These variables' summary statistics are presented in Appendix A1.

2.4 Identification Strategy

Our baseline specification examines whether prefectures with higher NTR gaps experience differential changes in family decisions after the change in the U.S. trade policy,

$$\Delta y_c = \alpha + \beta NTRGap_c + \gamma \mathbf{X}_c + \mu_p + \varepsilon_c \tag{3}$$

where Δy_c is the decadal change of a variable of interest for prefecture c, such as the change in the fraction of young women (aged 20-39) who are currently married.[13] The main explanatory variable, $NTRGap_c$, is prefecture c's exposure to PNTR as we have described in equation (1). The term $\mathbf{X}_c$ represents a set of additional controls, and μ_p denotes the province-fixed effects. In the main specification, $\mathbf{X}_c$ includes contemporaneous Chinese tariff reductions[14] and a set of start-of-period conditions that may lead to differential changes in local marriage market across prefectures: manufacturing employment share, fraction of female employed (aged 20-39), rural population share, minority population share, the fraction of population with college or higher education, and a dummy indicating whether a prefecture has export processing zones.

When we are interested in distinguishing the effects of female industry versus male industry shocks, we access the effects of PNTR by fitting a model of the form

$$\Delta y_c = \alpha + \beta_1 NTRGap_c^f + \beta_2 NTRGap_c^m + \gamma \mathbf{X}_c + \mu_p + \varepsilon_c, \tag{4}$$

where $NTRGap_c^f$ and $NTRGap_c^m$ are the gender-specific trade shocks that we have described in equation (2).

3 Empirical Results

3.1 Marriage

Starting from this section, we analyze the effect of PNTR on young women's marriage and fertility outcomes. Table 1 presents the results of regressing the change in the fraction

[13] In China, the legal age of marriage for women is 20 years old.

[14] Specifically, China engages in phased tariff reductions after its WTO accession. In 2000, China's simple average applied tariff is 17%, with the standard deviation across the six-digit Harmonized System (HS6) products being 12%. By the end of 2005, the average tariff level is reduced to 6%, and the standard deviation almost halved. The average tariff level stabilizes after 2005. We construct regional output and input tariff cuts across Chinese prefectures between 1999 and 2007, in a similar way as NTR gaps. Similar measures have been used in Lu and Yu (2015), Brandt et al. (2017), and Zi (2020).

of young married women (20-39 years) on regional PNTR exposure. All regressions are weighted by the beginning-of-period population of the same age cohort. Column (1) of Table 1 presents the OLS results only controlling for initial manufacturing employment shares. Column (2) additionally controls the contemporaneous trade liberalization upon China's WTO accession, and column (3) includes the full set of baseline controls. Column (4) is the preferred specification, in which we further add province fixed effects to control for province-specific trends. In all cases, the point estimates of $NTRGap_c$ are negative and statistically significant at the 1% level. Including province fixed effects slightly reduces the estimated coefficient, but the magnitude is similar overall. The estimate of -0.17 ($s.e. = 0.07$) in column (4) of Table 1 implies that a unit increase in PNTR exposure is associated with an approximately 0.17 relative decrease in the fraction of married young women. The difference between PNTR exposure in regions at the 90th and 10th percentiles is 0.16. Therefore, evaluated using the estimate in column (4), a prefecture at the 90th percentile experiences about 2.7 percentage point larger decline in the fraction of married young women compared to a prefecture at the 10th percentile. The mean of $NTRGap_c$ is 0.25, indicating that the trade policy change on average is associated with a 4.3 percentage point decline in the share of married young women. Therefore, PNTR explains approximately 37% of the observed aggregate trend in the share of married young women between 2000-2010.[15]

As expected, prefectures with smaller shares of manufacturing employment and greater shares of educated populations in 2000 experience relatively larger declines in the fraction of married young women. This regional change is also negatively correlated with the fraction of minority populations at the initial period (column (3) of Table 1). Yet, this relationship becomes insignificant once we control for province fixed effects. On the other hand, the reduction in output (input) tariffs seems to have a negligible (and statistically insignificant) impact on the change in the regional female marriage share.

Marital Status by Type and Gender.— Having presented the negatived relationship between $NTRGap_c$ and married young women shares, we next explore the impact of $NTRGap_c$ on people's marital status by type. In particular, we split the reason of being single to two, never married or divorced. The results are reported in Table 2-I, columns (2) and (3). We find that high PNTR prefectures experience relatively larger increases in the fraction of women who have never married and in the fraction of divorces, which are statically significant at the 5% level. Because of the relatively low divorce rate in China (0.81% in 2000),

[15] $0.17 \times 0.25 \approx 0.0425$. The aggregate decline in the share of married young women is 11.37%, and $4.25/11.37 = 0.37$.

about 97.50% of the change in the fraction of married young women in response to $NTRGap_c$ can be attributed to the change in the share of never-married young women. Men in the same age group also experience a relative decline in the share of marriages in high PNTR prefectures, as shown in column (4) of Table 2-I. However, this is entirely accounted for by the increased share of never-married young men; the estimated coefficient on the share of divorced men, on the other hand, is negative and statistically insignificant (columns (5)-(6), Table 2-I).

Gender Specific Trade Shocks.— Table 2-II introduces gender-specific trade shocks using the specification (2). The correlation between these by-gender shocks is moderate ($\rho = 0.4$), leaving sufficient power to distinguish their independent effects. As presented in columns (1)-(6) of Table 2-II, the estimation results all point to the unique importance of female-specific trade shocks. A unit increase in female-specific trade shocks decreases married young women shares by 0.31 ($s.e. = 0.10$), increases the share of women never married (divorced) by 0.27 (0.02) ($s.e. = 0.11, 0.01$, respectively). Female-specific trade shocks also reduce the share of married young men by a slightly smaller magnitude ($\beta = 0.24, s.e. = 0.10$), almost fully accounted for by the increases in never-married young men. On the other hand, trade shocks to men-specific industries affect neither young women nor men's marital status: the point estimates of $NTRGap_c^m$ are close to zero and statistically insignificant in all cases. These results suggest that the granting of PNTR affects marriage outcomes in China mainly by affecting women's marriage decisions – this will be important for interpreting our baseline results and discussing mechanisms in the later section.

Pre-trends.— There could be a possibility that prefectures more exposed to PNTR have different trends in their marriage market. For example, more exposed prefectures somehow have started a declining trend in the fraction of married young women long before China's WTO accession. To address this concern, we carry out a placebo test in Table 2-III, where the dependent variables are replaced by the decadal change in different outcomes between 1990-2000. If any differential trend that starts before 2000 in the local marriage markets confounds our results, the PNTR exposure should be correlated with it. However, all estimates are statistically insignificant and close to zero, mitigating such a concern.

Results by Age Cohorts.— Figure 1-(a) visualizes the effect of PNTR on the marital status by types for young women in different age cohorts. We expect that PNTR will have a greater effect on young people's first marriages because older individuals are more likely to be married when the shock occurs. Similarly, we anticipate that PNTR will have a greater

effect on the divorce rate among older age cohorts. Consistent with the expectation, we find that the younger the cohort, the stronger the positive effect of trade shocks on the fraction of never-married women, and the impact is statistically significant for the 20-24 age cohorts. On the other hand, the positive effect of PNTR on the fraction of divorced young women increases with the increases in age cohorts, and the impact is most significant for women aged 35-39. As a result, the effect of $NTRGap_c$ on the fraction of *married* young women is negative and statistically significant for the age cohorts 20-24 and 35-39, caused by distinct reasons. Note that because women in these age groups may not have completed their decision to marry, we interpret these results as evidence that PNTR reduces or *delays* women's marriage in China.

3.2 Fertility

In this section, we analyze the effect of PNTR on the fertility outcomes of young women. The Chinese population census only surveys the fertility status of women who (ever) get married. This has two implications. First, never-married mothers are excluded from the survey, even though they should represent a small portion of the population in the context of China.[16] Second, unlike the marital status, we cannot observe fertility outcomes on the male side. With these limitations in mind, we repeat the exercises in Section 3.1 but examine fertility outcomes instead. The results are reported in Table 3.

Table 3-I presents the results of regressing the changes in three outcome variables on regional PNTR exposure: the number of children per young woman (20-39 years), the fraction of young women with children, and the number of children per young woman with children. All regressions are weighted by the beginning-of-period population of the same age cohort. Not surprisingly, we find a tight link between marriage and fertility decisions: a unit increase in PNTR exposure is associated with an approximately 0.13 relative decrease in the fraction of young women with children ($s.e. = 0.07$).[17] However, conditional on having children, a one-unit increase in PNTR exposure *increases* the number of children for those with children by 0.15 ($s.e. = 0.15$), though the effect is imprecisely estimated.[18]The two effects work in

[16]Specifically, a new birth of an unmarried woman can hardly acquire an official hukou, which affects children's schooling and other welfare services. Beyond institutional constraints, fertility choice is tightly associated with marriage decisions in China for cultural reasons as well.

[17]Recall that a unit increase in PNTR exposure is associated with an approximately 0.17 relative decrease in the fraction of married young women.

[18]The positive response in the number of births among those with children to PNTR can due the high cost of raising children and the additional cost of having more than one child associated with the one-child policy. The one-child policy began to be relaxed in late 2011.

opposite directions and lead to a statistically insignificant decrease in the number of children per woman in more PNTR exposed prefectures ($\beta = -0.06$, $s.e. = 0.19$).

Table 3-II introduces the gender-specific trade shocks. A unit increase in female-specific trade shocks decreases the share of women with children by 0.26 ($s.e. = 0.11$), which is statistically significant at the 5% level. Female-specific trade shocks are also negatively associated with the number of children per woman and positively associated with the number of children per woman with children, although both are not statistically significant. The male-specific shocks, on the other hand, do not have statistically significant effects in all cases. These results are consistent with the response of marriage to PNTR and again point to the unique importance of female-specific trade shocks.

Table 3-III reports the placebo test in which the dependent variables are replaced by the decadal change in the corresponding outcomes between 1990-2000. Again, all estimates are statistically insignificant and close to zero, mitigating the concern of the presence of differential pre-trends.

Finally, Figure 1-(b) visualizes the effect of PNTR on women's fertility outcomes in different age cohorts. As expected, the older the cohort, the weaker the negative effect of PNTR on the number of children per woman and the share of young women with children. This effect is only significant for the 20-24 age group, possibly because a greater share of women in the older age group has already completed their reproductive decisions by the time PNTR occurs. Again, since women may not have completed their fertility decisions at the age of 20-39, we interpret these results as evidence that PNTR has reduced the share of young women having children or *delayed* women's fertility decisions.

3.3 Heterogeneous Effects on Young Women of Different Types

We further explore the heterogeneous effects of PNTR on different types of young women's marriage and fertility decisions. A natural conjecture is that more educated women in urban areas have a greater say in their personal matters; thus, their decisions may be more responsive to PNTR. However, on the other hand, PNTR may have a greater impact on the opportunity cost of marriage and childbearing for rural or low-educated women through various channels (which we discuss in the next section) and thus affect their family choices more.

Specifically, we consider two comparisons. The first is between rural (agriculture hukou) and urban (non-agriculture hukou) young women. The second is between low-educated (below high school education) and high-educated (high school education or above) young

women. Rural or low-educated women presumably should have lower incomes and economic stature than urban and high-educated women; they account for 73.4% and 79.8% of total young women in the 2000 Census, respectively.

Table 4 provides the corresponding results on marriage decisions. Columns (1) and (7) of Table 4-I show that prefectures with higher PNTR exposure experience greater declines in marriage shares for both hukou types. Nevertheless, the effect is more pronounced and only statistically significant for women with agriculture hukou. Further examination reveals that PNTR affects rural females primarily by increasing the share of never-married women, while urban females by increasing the share of divorces (columns (5) and (9), respectively). Since divorced women represent only a small fraction (0.81% in 2000) of the total young female population, the impact of PNTR on overall marriage shares is expected to be greater and more accurately estimated for rural women compared to that for urban women – consistent with the estimates reported in columns (1) and (7). On the other hand, although PNTR also increases the fraction of the divorced (never married) population among rural (non-rural) women, the effect is not statistically significant, as reported in column (3) (column (11)).

Table 4-II compares young women with low- and high-educational attainment. The baseline results are driven almost exclusively by the responses of young women with less than high school education. As shown in columns (1)-(6), PNTR significantly reduces the fraction of married women by increasing the fraction of divorced and never-married women in the low-education group, with point estimates being similar to the baseline case. In contrast, the marriage decisions of high-educated women do not appear to be correlated with PNTR shocks (Columns (7)-(12)).

To alleviate the concern that rural and low-education groups may largely overlap, we control for the 2000-2010 changes in the share of high-educated (rural) young women when regressing by household (education) type. Adding the additional control does not change our results in any meaningful manner, as presented in even columns of Table 4-I, II.

Table A2 provides results on fertility decisions. The PNTR shock reduces the fraction of young mothers in the rural group more, which is consistent with the marriage results. The remaining outcomes are imprecisely estimated but qualitatively in line with our baseline findings: a prefecture's PNTR exposure is positively associated with the regional decline in fraction of women with children and the increase in the number of children per woman with children, regardless of women types.

3.4 Shift-share Designs

Recent research on identification and inference in shift-share designs indicates that consistency can be attained if either the shares or the shifts are exogenous (Adão et al., 2019; Borusyak et al., 2022; Goldsmith-Pinkham et al., 2020). In particular, Goldsmith-Pinkham et al. (2020) formalizes an approach for shift-share identification and consistency based on the exogeneity of the shares, imposing no explicit assumption of shock exogeneity. Borusyak et al. (2022) proposes a different framework based on the quasi-random assignment of shocks while allowing the shares to be endogeneous.

In our setting, the validity of our shift-share design relies on the exogeneity of the shifts. The U.S. sets the NTR and non-NTR tariffs against all its trading partners and thus is unlikely to be affected by the demographic conditions of one particular country-China. More importantly, seventy-nine percent of the variation in the NTR gap across industries arises from variation in non-NTR rates (Pierce and Schott (2016)), set 50 years before the normalization of U.S.-China relations and 70 years before China's WTO accession. Therefore, we apply the equivalence results of Borusyak et al. (2022) and transform our prefecture-level specification into a shift-level specification, re-estimating the effect of PNTR on young women's marital and fertility decisions across industries.

We first justify the quasi-experimental view of our research design. Panel (a) of Table A3 summarizes the distribution of industry shocks (g_j) and the industry-level weights from the equivalence result (s_j). The shock, measured by $NTRGap_j$, is well distributed with a mean of 0.27, a standard deviation of 0.16, and an interquartile range of 0.25. The effective sample size of our equivalent regression (the inverse HHI of the s_j) is also relatively high: 76.31 across CIC4 industries and 54.98 when aggregated to the 3-digit Chinese manufacturing industry (CIC3) level. The largest shock only weights 6% in the CIC4 industry and 7% in the CIC3 industry, suggesting a considerable variation across industries. The intra-class-correlations (ICCs) of shocks within the CIC3 and CIC2 industries are 0.15 and 0.05, respectively. These reveal moderate clustering of shock residuals at the CIC3 level but less evidence of clustering of shocks at a higher CIC2 level, so it will be sufficient to cluster standard errors at the level of CIC3 groups. The inverse HHI estimates in column (5) of Panel (a) indicate that there is still an adequate, effective sample size at this level of shock clustering.

Panel (b) of Table A3 reports industry-level and region-level tests of the balance of shocks. The industry-level covariate is log wages in 1999, with which we find no statistically significant correlation for NTR Gap. The region-level covariates are the beginning-of-period baseline controls: fraction of female employed (aged 20-39), the share of population with

agriculture hukou, the share of the minority population, the share of the population with college or higher education, number of dummies in processing zones; and pre-trend (1990-2000) changes in the fraction of married young women. Throughout the analysis, we control for the share of manufacturing employment. We again find no statistically significant relationship between these variables and the shocks across time, except for the share of the minority population. Areas that benefit from larger NTR Gaps tend to have smaller minority populations, which may lead to different regional marriage dynamics. We include the baseline controls in the shift-level regressions below.

Having assessed the plausibility of the quasi-experimental nature of the shift-share design, we next revisit the effect of PNTR on young women's marriage and fertility decisions. Panel (c) of Table A3 reports the shift-share coefficients and valid exposure-robust standard errors[19] obtained from equivalent industry-level regressions. Based on the ICCs results reported above, we cluster standard errors at the CIC3 level. Overall, the estimated coefficients are economically and statistically significant, similar in magnitude to the baseline estimates.

4 Understanding the Mechanism

Having established the link between PNTR and regional change in marriage and fertility outcomes among young women, we now turn to understand the mechanism. Based on the main development of the Chinese economy in the period, we lay out five hypotheses. The first hypothesis is the role of economic stature: trade liberalization may shift the relative employment opportunities of young women versus men, affecting young women's marriage decisions. A second hypothesis is the role of education. Trade liberalization tends to change the returns to education through various channels (Heath and Mobarak (2015), Blanchard and Olney (2017), Li (2018), Lin and Long (2020)), in which case young women may choose to continue their education and hence postpone marriage decisions. A third hypothesis is that trade can reduce marriage incidence by creating a more gender-segregated labor market. For example, trade liberalization may lead to more young women working in female-intensive industries, increasing cost of searching potential partners and thus deterring their marriage (Goni (2022)). The fourth hypothesis is that trade liberalization can also create a spatial Balassa-Samuelson effect. In this case, locations with greater PNTR exposures tend to

[19]The method by Borusyak et al. (2022) estimates shift-share coefficients at the level of shocks and yields asymptotically valid standard errors, thus providing an alternative solution to the invalid conventional standard errors in shift-share regression (because observations with similar exposure shares may have correlated residuals) raised and solved by Adão et al. (2019).

have higher population densities, urban population shares, and higher relative prices of nontraded goods (Fajgelbaum and Redding, 2022). This may raise the cost of marriage and childbearing by placing a strain on housing and educational resources. Finally, China's trade liberalization has led to massive labor reallocation across regions (e.g., Facchini et al. (2019); Zhou and Zhang (2021); and Zi (2020)). If young female workers tend to move to a different location than male workers due to regional differences in industry specialization, trade may create a gender imbalance in the spatial dimension, thus affecting young people's marriage opportunities.

We present multiple evidence for or against each hypothesis and find that the empirical patterns appear most consistent with the economic stature hypotheses. So we first present evidence for this hypothesis and then discuss the rest.

4.1 Economic Stature

The importance of economic status in influencing family decisions can be traced to Becker (1973), who argues that an increase in the relative economic status of women reduces the gains from family specialization and therefore reduces the prevalence of marriage. In contrast, a rise in men's economic opportunities has the opposite effect. Because it is the shifts in women's *relative* economic status that matters, we modify the baseline specification to examine the differential effects of PNTR on young men and young women:

$$\Delta y_{cg} = \alpha + \beta_1 NTRGap_c \times Female_g + \beta_2 NTRGap_c + \beta_3 Female_g + \gamma \mathbf{X}_c + \mu_p + \varepsilon_c \quad (5)$$

where Δy_{cg} is the decadal change of the outcome variable y for gender g in city c and $Female_g$ represents a female dummy. With this specification, the effect of $NTRGap_c$ on young men is captured by β_2, while β_1 captures the differential effect of $NTRGap_c$ on young women relative to that on young men.

First, we study the differential impact of PNTR on the share of young women who are unemployed or not in the labor force compared to that of young men. Panel (a) of Table 5 presents the results. Columns (1)-(3) examine each of the three most common destinations for young people who leave the labor force,[20] while column (4) looks at the unemployment response. The estimates in columns (1)-(3) indicate that PNTR significantly decreases the fraction of women out of the labor force relative to that of men, and almost

[20]School/training, family obligations, and health limitations account for 84.06% and 79.07% of young people exiting the labor force in 2000 and 2010, respectively.

exclusively through a reduction in the fraction of women who leave the labor force due to family issues. As indicated in column (2), a unit increase in PNTR exposure reduces the ten-year change in the proportion of young women who do not work due to family obligations by 0.15 ($s.e = 0.04$) relative to that of men, with the latter being barely affected by PNTR shocks. Regardless of gender, PNTR has a negligible effect on the proportion of individuals out of the labor force due to education or health limitations, as reported in columns (1) and (3). Finally, the estimates in column (4) indicate that while PNTR does not affect the fraction of unemployed young men, a unit increase in PNTR reduces the fraction of unemployed young women by 0.03 ($s.e = 0.02$) relative to that of young men. Overall, estimates in Panel (a) of Table 5 indicate that PNTR enhances young women's labor market participation rate by reducing the proportion of housewives and decreasing the fraction of unemployed young women relative to that of young men.

We next investigate the impact of PNTR on the employment of young women versus young men in Panel (b) of Table 5. Column (1) considers the broad impacts on overall employment, and the rest columns consider employment by sectors. The estimates in column (1) show that PNTR significantly increases female employment relative to males - a unit increase in PNTR exposure increases the ten-year change in young female employment share by 0.14 ($s.e = 0.05$) over that of males. Estimates in columns (2)-(4) of the Panel (b) suggest that PNTR leads to a different redistribution of employment between industries for young women and men. For the directly affected sector, manufacturing, a unit increase in PNTR exposure leads to an increase of the share of manufacturing employment for young men by 0.34 ($s.e = 0.06$), but the share of employment for young women increases by 0.28 ($s.e = 0.07$) *less* compared to men. On the other hand, the share of service employment for young men decreases by 0.18 ($s.e = 0.07$). However, the share of services employment of young females experiences a relative and absolute increase: the interaction term is estimated at 0.37 ($s.e = 0.08$). The point estimates for the agricultural sector are one order of magnitude smaller and statistically insignificant. In sum, these estimates suggest that PNTR leads to a reallocation of young males from services to manufacturing, but there is no evidence of net employment growth. On the other hand, PNTR has significantly improved female labor participation, particularly in the services sector.

Overall, the results presented in Table 5 indicate that PNTR not only differentially improves female employment, but also causes relatively more young women to move to the services sector, where wages are higher. These findings are consistent with our conjecture that PNTR leads to a relative increase in women's economic status, which ultimately reduces or

delays their marriage and fertility. Table A4 further demonstrates that the estimates of Table 5 remain robust when focusing on married individuals: the relative improvement in young women's employment persists after marriage, indicating that the rise in females' relative economic status is not transitory. Consistent with our baseline findings, Table 6 indicates that the differential employment effects are predominantly driven by the population with low educational attainment and those with a rural hukou.

Figure 2 illustrates the monthly income information by gender and sector in 2005, notwithstanding the absence of location- and gender-specific salary data for the years 2000-2010.[21] As expected, agricultural workers have the lowest wages, while men workers across all industries earn more than their female counterparts. However, two additional stylized facts emerge. First, service sector employees have a higher average wage than manufacturing employees. Second, the gender wage gap is less in the service compared to the manufacturing sector. This indicates that the expansion of female employment in the service sector, driven by PNTR, is likely to narrow the average wage difference between young women and young men rather than widening it.

In conclusion, although we cannot directly evaluate how PNTR affects the gender pay gap, the observed differential employment adjustments to PNTR and suggestive evidence on wages across sector and gender indicate that PNTR improves the economic standing of young women in China relative to young men.

4.2 Women Education

In addition to economic statures, PNTR may also affect young women's family choices by affecting their educational decisions. First, positive trade shocks can lead to higher local incomes, which may enable more families to pay for their children's higher education. The increased share of young women in school may naturally cause women to delay marriage, as students are less likely to marry before graduation. Second, trade shocks can affect the return on education, thereby incentivizing more women to pursue higher education, which in turn alters their opportunity cost of marriage and childbirth (Becker, 1991). We examine these two possibilities in Table A5.

Column (1) of Table A5 shows that PNTR has a negligible and statistically insignificant effect on the decadal change in the fraction of young women currently in school. Since 99.46% women in China complete school education before the age of 24, the result may be subject to attenuation bias because the age range of 20-39 may be too wide. Therefore, column (2)

[21] Only the 2005 Chinese Population Census reports each respondent's monthly income.

repeats the regression of column (1), but focuses on women aged 20-24. The point estimate increases slightly but remains statistically insignificant. In conclusion, we find no evidence that PNTR is associated with regional changes in the share of young women in school.

Columns (3)-(6) of the Table A5 examine whether PNTR has incentivized more females to pursue higher education. Since China has a nine-year compulsory education system,[22] we examine the change in the fraction of young women with more than a junior high school education. In particular, we separate high school education from college or higher, and we examine both the full sample of young women (20-39 years old) as well as women aged 20-24. We do not find any evidence that greater PNTR exposure is associated with an increase in the fraction of women with higher education; point estimates even have the wrong sign for changes in the fraction of women with a college or higher degree for the 20-24 age group.

Thus, education access, either by affecting the incidence of young women in school or their overall educational attainment, cannot explain our main findings.

4.3 Gender Segregation in Local Labor Market

Our third hypothesis is that trade can reduce marriage incidence by creating a more gender-segregated labor market. For example, trade liberalization may lead to more young women working in female-intensive industries, reducing their chances of meeting potential partners. To probe whether there is a greater increase in gender concentration in high PNTR regions, we construct an industry gender-concentration index (GCI) inspired by the Herfindahl-Hirschman Index (HHI):

$$GCI_{ct} = \sum_j w_{jc} \times \left(f_{jct}^2 + (1 - f_{jct})^2 \right),$$

where w_{jc} is the start-of-period employment share of industry j in prefecture c, and f_{jct} is the share of young female over total young employment of industry j in prefecture c in year t. Intuitively, if the young people in an industry are gender-balanced, $f_{jct} = 0.5$ and $f_{jct}^2 + (1 - f_{jct})^2 = 0.5$. More imbalanced the gender gets, greater the value of $f_{jct}^2 + (1 - f_{jct})^2$.

We then test whether Chinese prefectures that are more exposed to PNTR experience a relative increase in their GCI index. Table A6 presents the results. Note that we only observe prefecture gender-specific employment by each CIC 3-digit industry, therefore the GCI index is computed from that level. As suggested in column (1) of Table A6, we find no effect of regional PNTR exposure on changes in gender concentration in industries. Column (2) finds

[22] That is, the education at the junior high school level or below is compulsory for young people.

that gender-specific trade shocks also do not significantly change the degree of the local labor market's gender segmentation. Columns (3)-(4) instead use aggregate sectors (agriculture, manufacturing, service and others) to compute the GCI index and find the opposite sign of the point estimates than expected. However, this should not come as a surprise. Female employment in 2000 is largely concentrated in agriculture,[23] as shown in Section 4.1, while the PNTR shock leads to more female working in manufacturing and services sectors. Therefore, when gender concentration is measured in very aggregated sectors, PNTR are expected to result in more, not less, gender-balanced industrial structures.

4.4 Marriage and Childbearing Costs

Trade liberalization can also create a spatial Balassa-Samuelson effect: locations with greater export opportunities may have higher population densities, urban population shares, and higher relative prices of nontraded goods (Fajgelbaum and Redding, 2022). This means that areas with high PNTR exposure may face, among other things, more rapid increases in housing prices and more severe shortages of educational resources, both of which can have a significant impact on the marriage and fertility decisions of young people in China. We explore these possibilities in Table A7.

In Table A7, column (1), we compute the decadal change in commercial housing prices (per square meter) between 2000 and 2010 and correlate it with the regional NTR gap.[24] Interestingly, we find that prefectures with higher PNTR exposure experience a relative decrease (or smaller increase) in housing prices, which may be due to an increase in housing supply in high PNTR regions. In column (3), we find that prefectures with higher exposure to PNTR also experience a relative increase in the teacher-student ratio.[25] While the point estimate is statistically insignificant, the overall findings contradict our hypothesis that prefectures with higher PNTR exposures will experience greater housing price increases or more severe teacher shortages.

Finally, as seen in columns (2), (4), and (5) of Table A7, the point estimates of PNTR on share of young women married continue to be robust and quantitatively comparable to the baseline estimate when controlling for changes in housing prices and teacher-student ratios. Thus, it does not appear that changes in the potential costs of marriage and childbearing

[23] In 2000, the fraction of employment in agriculture is 62.27% for young female and 53.86% for young male.

[24] The data comes from China Statistical Yearbook for Regional Economy; the housing price is computed as total sales of commercial houses divided by the total floor spaces sold (square meter).

[25] We use the teacher-student ratio for elementary school because it is the stage of education closest to marriage. The data are from the China City Statistics Yearbook.

are a main reason in young women's decisions to delay or avoid marriage.

4.5 Internal Migration

Finally, trade shocks often lead to substantial labor force adjustment across firms, sectors, and regions. Existing studies suggest that PNTR induces considerable migration across regions in China (e.g., Facchini et al. (2019); Zhou and Zhang (2021)). Migration can affect the local marriage outcome in various ways. First, migrant workers tend to marry later,[26] so trade may induce a compositional effect on regional marriage outcomes via affecting migration. Secondly, if young female and male workers migrate to different destinations due to regional differences in industry specialization, trade may create gender imbalances in the spatial dimension, affecting young people's marriage opportunities. Finally, young immigrants may prefer to marry the locals to enjoy the various social benefits offered by the local hukou system in China. The more immigrants there, the smaller the fraction of marriageable locals, making it harder to find a match.[27] In conclusion, both the quantity and gender composition of migrants may affect the local marriage market. We explore these three possibilities in Table A8.

Table A8, Panel (a) examines the possibility of the compositional effect. Column (1) of Panel (a) confirms, consistent with the findings of Facchini et al. (2019), that more exposed prefectures indeed experience a relative increase in the share of young migrants.[28] Columns (2)-(4) of Table A8-(a) repeat the baseline regressions in columns (1)-(3) of Table 1, but also control for the decadal change in the proportion of young migrants. Prefectures with greater inflow of migrants do experience a more significant decline in the fraction of married young women and an increase in the fraction of never-married young women. However, the point estimates of PNTR continue to be robust and quantitatively comparable to the baseline. Thus, local marriage outcomes do not appear to be driven by the compositional change. This may not be surprising, because despite the size of internal migration, they are still small compared to the total population of each region.

Table A8, Panel (b) considers the possibility of the trade-induced imbalances in the local marriage market. Column (1) of Table A8-(b) demonstrates that more exposed prefectures

[26] In 2000, the share of immigrants getting married is 15.34 percentage points lower than that of locals.

[27] Related, Lichter et al. (2015) finds that white immigrants are far more likely than other groups to marry U.S. born natives, while black immigrants are much less likely to marry black natives or out-marry with other groups.

[28] A migrant is an individual who resides in a prefecture but with a hukou outside the province or a hukou under the pending status. This is a consistent measure in the 2000 and 2010 Census questionnaires. Other measures comparing the residence locality five years ago and the current locality generate similar results.

have more young male migrants than young female migrants. However, as indicated in column (2), this effect is rather modest and does not translate into a significant local gender disparity, mitigating the concern of trade-induced gender disparities in the spatial dimension. Columns (3) and (4) further demonstrate that there is no evidence that the PNTR exposure is associated with an increase in the ratio of young female locals to total young males, nor the ratio of young male locals to total young females. That is, even assuming that all migrants prefer to marry locals, there is no evidence that the incidence of marriage among young individuals is affected.

Overall, the results presented in Table A8 suggest that the influx of migrants is unlikely to be the primary explanation for our baseline findings.

5 Conclusion

By examining the granting of Permanent Normal Trade Relations (PNTR) by the U.S. to Chinese exports, we study the impact of trade-induced labor market opportunities on young Chinese women's marriage and fertility decisions. We find that Chinese prefectures with a greater PNTR exposure experience a relative increase in the fraction of unmarried young women. This relative increase is due to young women reducing or delaying their first marriages and more married women choosing to divorce. As a result of changed marriage decisions, the number of children per woman and especially the share of women with children also experience a relative decline in more exposed areas. We find that PNTR can approximately explain 37% of the observed aggregate decline in marriage among young women between 2000-2010. We show that these shifts in family decisions coincide with a trade-induced reallocation of women relative to men to the service sector, where wages are higher. Trade-induced changes in educational choices or migration patterns, marriage costs, or gender segregation in the labor market that may be caused by PNTR, on the other hand, do not appear to explain the decline in female marriage and fertility decisions in China.

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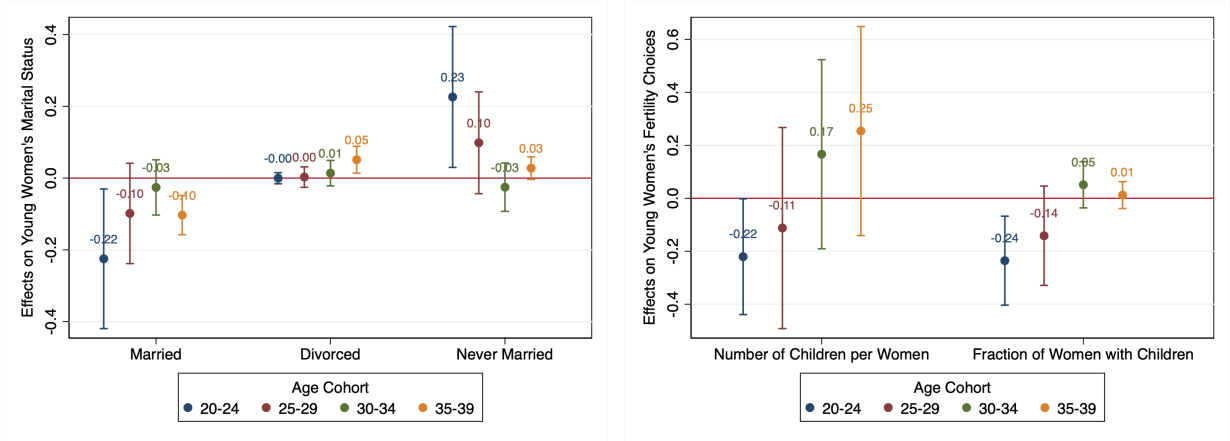
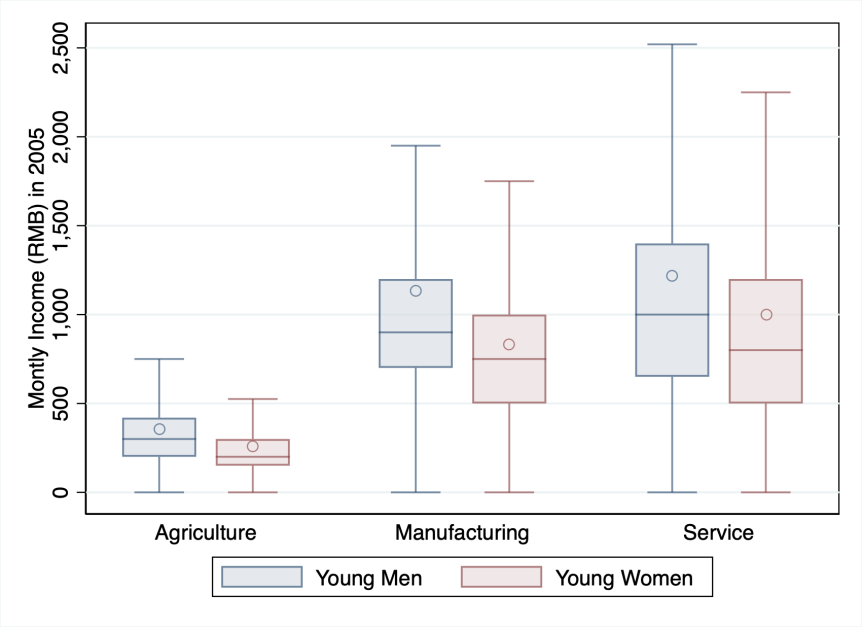


(a) Marital Status (b) Fertility Choices

Notes: (a): We replicate Table 2-I, columns (1)-(3) for young women in four different age cohorts: 20-24, 25-29, 30-34, 35-39 and plot the estimates and 10% confidence intervals. (b): We replicate Table 3-I, columns (1)-(2) for young women in four different age cohorts: 20-24, 25-29, 30-34, 35-39 and plot the estimates and 10% confidence intervals.

Figure 1: The Effect of PNTR on the Family Decisions of Young Women by Age Groups



Notes: This figure visualizes the box plots of individual wages by sector and gender in China. We use individual monthly income information from the 2005 Chinese Population Census and plot the summary statistics by gender and sector. The circle, central line, upper and lower hinges of the box are the mean, median, the 75th and the 25th percentile of individual monthly income. The upper and lower adjacent lines are upper and lower adjacent values, as defined by Tukey (1977).

Figure 2: Monthly Income by Gender and Sector in 2005

Table 1: Effects of PNTR on Fraction of Young Women Married

	Main			
	(1)	(2)	(3)	(4)
NTR Gap	-0.31***	-0.35***	-0.26***	-0.17***
	(0.07)	(0.09)	(0.07)	(0.07)
Fraction of employment in manufacturing	0.25***	0.25***	0.34***	0.33***
	(0.06)	(0.07)	(0.04)	(0.05)
Output tariff changes		0.16	-0.20	-0.04
		(0.24)	(0.20)	(0.18)
Input tariff cuts		-0.85	0.65	0.74
		(0.62)	(0.60)	(0.60)
Fraction of young women employed			0.01	0.05
			(0.04)	(0.05)
Fraction of population with agricultural hukou			0.06	0.05
			(0.05)	(0.05)
Fraction of minority population			0.07***	0.02
			(0.01)	(0.02)
Fraction of population with college or above education			-0.74***	-0.87***
			(0.22)	(0.22)
Processing zone dummy			0.03	0.06
			(0.03)	(0.04)
Province fixed effects				Yes
Observations	337	337	337	337
R-squared	0.18	0.18	0.43	0.57

Notes: The dependent variable is the 10-year change in fraction of young women (aged 20-39) currently married. The sample contains 333 prefectures and four directly controlled municipalities. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table 2: Effects of PNTR on Young People's Marital Status

	Young Women			Young Men		
	Married	Divorced	Never Married	Married	Divorced	Never Married
	(1)	(2)	(3)	(4)	(5)	(6)
I. Overall trade shocks						
NTR Gap	-0.17***	0.02**	0.14**	-0.14**	0.00	0.15**
	(0.07)	(0.01)	(0.07)	(0.06)	(0.01)	(0.07)
R-squared	0.57	0.16	0.59	0.52	0.18	0.51
II. Gender-specific trade shocks						
NTR Gap × (female share)	-0.31***	0.02*	0.27***	-0.24**	0.01	0.24**
	(0.10)	(0.01)	(0.11)	(0.10)	(0.02)	(0.10)
NTR Gap × (male share)	-0.02	0.01	-0.01	-0.02	-0.01	0.03
	(0.10)	(0.02)	(0.10)	(0.10)	(0.02)	(0.10)
R-squared	0.58	0.16	0.60	0.52	0.19	0.51
III. Pre-trend: 1990-2000						
NTR Gap	0.04	0.01	-0.03	0.00	0.01	-0.00
	(0.05)	(0.01)	(0.05)	(0.08)	(0.01)	(0.08)
R-squared	0.54	0.54	0.54	0.36	0.50	0.36
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	337	337	337	337	337	337

Notes: The dependent variable is the 10-year change in marital status of young people (aged 20-39). Columns 1-3 are fraction of young women currently married, divorced, never married. Columns 4-6 are fraction of young men currently married, divorced, never married. The sample contains 333 prefectures and four directly controlled municipalities. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table 3: Effects of PNTR on Young Women's Fertility Decisions

	Number of Children per Woman	Fraction of Women with Children	Number of Children per Woman with Children
	(1)	(2)	(3)
I. Overall trade shocks			
NTR Gap	-0.06	-0.13*	0.15
	(0.19)	(0.07)	(0.15)
R-squared	0.53	0.60	0.64
II. Gender-specific trade shocks			
NTR Gap × (female share)	-0.21	-0.26**	0.19
	(0.32)	(0.11)	(0.27)
NTR Gap × (male share)	0.11	0.01	0.10
	(0.29)	(0.11)	(0.26)
R-squared	0.53	0.60	0.64
III. Pre-trend: 1990-2000			
NTR Gap	-0.12	-0.01	0.00
	(0.28)	(0.05)	(0.33)
R-squared	0.36	0.62	0.48
Baseline controls	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes
Observations	337	337	337

Notes: The dependent variable is the 10-year change in fertility decisions of young women (aged 20-39). The sample contains 333 prefectures and four directly controlled municipalities. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Heterogeneous Effects of PNTR on Young Women's Marital Status

	Married		Divorced		Never Married		Married		Divorced		Never Married	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
I. By Hukou	Agricultural						Non-agricultural					
NTR Gap	-0.15**	-0.14**	0.01	0.01	0.14*	0.12*	-0.10	-0.10	0.05*	0.05*	0.05	0.05
	(0.07)	(0.07)	(0.01)	(0.01)	(0.07)	(0.07)	(0.12)	(0.11)	(0.03)	(0.03)	(0.12)	(0.11)
Fraction of high-educated young women		-0.38**		-0.00		0.37***		-0.17***		0.00		0.17***
		(0.07)		(0.01)		(0.07)		(0.05)		(0.01)		(0.05)
Observations	337	337	337	337	337	337	337	337	337	337	337	337
R-squared	0.58	0.62	0.14	0.14	0.59	0.63	0.42	0.44	0.19	0.19	0.44	0.46
II. By Edu	Low						High					
NTR Gap	-0.15**	-0.15**	0.02*	0.02*	0.12*	0.12*	-0.02	-0.01	0.02	0.02	-0.01	-0.02
	(0.06)	(0.06)	(0.01)	(0.01)	(0.06)	(0.06)	(0.13)	(0.13)	(0.03)	(0.03)	(0.13)	(0.13)
Fraction of rural young women		0.00		-0.02		0.02		-0.12*		0.00		0.12*
		(0.08)		(0.01)		(0.07)		(0.07)		(0.01)		(0.07)
Observations	337	337	337	337	337	337	336	336	336	336	336	336
R-squared	0.54	0.54	0.15	0.16	0.54	0.54	0.45	0.45	0.17	0.17	0.45	0.46
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The dependent variable is the 10-year change in marital status of young women (aged 20-39). Columns (1), (2), (7) and (8) are fraction of young women currently married. Columns (3), (4), (9) and (10) are fraction of young women divorced. Columns (5), (6), (11) and (12) are fraction of young women never married. The sample contains 333 prefectures and four directly controlled municipalities. Columns (7)-(12) in Panel II contain only 332 prefectures and four directly controlled municipalities as the Golog Tibetan Autonomous Prefecture has no sample of high educated women in 2000. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table 5: Effects of PNTR on Gender Gaps in (Non-) Labor Markets

(a) Fraction of Young People in Non-Labor Markets

	School or Training	Family Obligations	Health Limitations	Unemployment
	(1)	(2)	(3)	(4)
NTR Gap × Female	0.01	-0.15***	-0.00	-0.03**
	(0.03)	(0.04)	(0.00)	(0.02)
NTR Gap	0.02	0.01	-0.00	-0.01
	(0.03)	(0.04)	(0.01)	(0.02)
Female	-0.00	0.07***	0.00	0.00
	(0.01)	(0.01)	(0.00)	(0.00)
Baseline controls	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes
Observations	674	674	674	674
R-squared	0.57	0.40	0.13	0.64

Notes: The dependent variable is the 10-year change in fraction on young people (aged 20-39) in non-labor markets. Columns (1)-(4) are fraction of young women (men) who are out of the labor force due to school or training, family obligations, health limitations, and who are job seekers, respectively. The sample contains 333 prefectures and four directly controlled municipalities × 2 gender groups (women or men). Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

(b) Fraction of Young People in Labor Markets

	Employed	Agriculture	Manufacturing	Service
	(1)	(2)	(3)	(4)
NTR Gap × Female	0.14***	0.03	-0.28***	0.37***
	(0.05)	(0.09)	(0.07)	(0.08)
NTR Gap	0.03	-0.13	0.34***	-0.18**
	(0.05)	(0.08)	(0.06)	(0.07)
Female	-0.06***	-0.02	0.06***	-0.10***
	(0.01)	(0.02)	(0.02)	(0.02)
Baseline controls	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes
Observations	674	674	674	674
R-squared	0.54	0.64	0.57	0.40

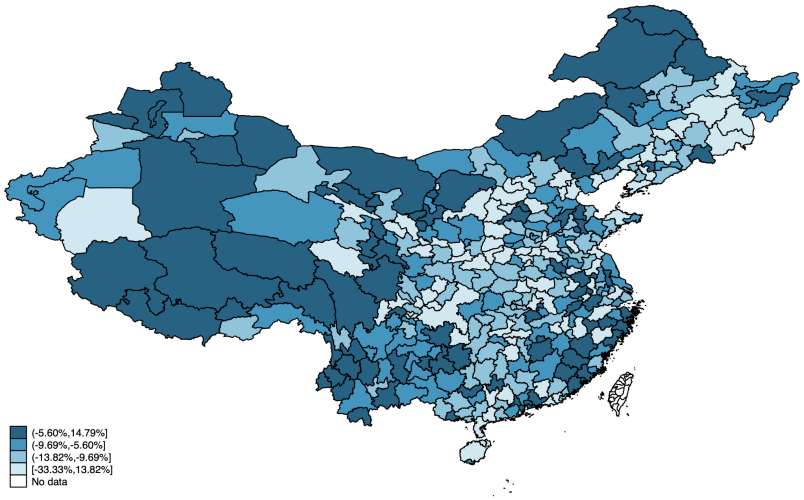
Notes: The dependent variable is the 10-year change in fraction on young people (aged 20-39) in labor markets. Columns (1)-(4) are fraction of young women (men) employed, employed in agricultural sector, employed in manufacturing sector, employed in service sector. The sample contains 333 prefectures and four directly controlled municipalities × 2 gender groups (women or men). Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Heterogenous Effects of PNTR on Gender Gaps in Labor Markets

	Employed	Agriculture	Manufacturing	Service	Employed	Agriculture	Manufacturing	Service
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
I. By Hukou	Agricultural				Non-agricultural			
NTR Gap × Female	0.12**	0.03	-0.32***	0.40***	0.15*	0.03	-0.09	0.17
	(0.05)	(0.12)	(0.09)	(0.08)	(0.08)	(0.05)	(0.08)	(0.12)
NTR Gap	0.00	-0.15	0.36***	-0.22**	0.10	-0.02	0.21**	-0.07
	(0.06)	(0.11)	(0.06)	(0.08)	(0.08)	(0.05)	(0.08)	(0.11)
Female	-0.06***	-0.02	0.08***	-0.11***	-0.05**	-0.01	-0.01	-0.02
	(0.01)	(0.03)	(0.02)	(0.02)	(0.02)	(0.01)	(0.02)	(0.03)
Observations	674	674	674	674	674	674	674	674
R-squared	0.48	0.57	0.53	0.39	0.31	0.32	0.32	0.25
II. By Edu	Low				High			
NTR Gap × Female	0.13**	0.03	-0.27***	0.34***	0.03	-0.14**	-0.02	0.15
	(0.06)	(0.10)	(0.09)	(0.08)	(0.09)	(0.07)	(0.08)	(0.14)
NTR Gap	0.06	-0.14	0.32***	-0.13	0.06	0.05	0.22***	-0.16
	(0.06)	(0.10)	(0.07)	(0.08)	(0.08)	(0.08)	(0.08)	(0.11)
Female	-0.07***	-0.02	0.07***	-0.11***	-0.03	0.08***	-0.03	-0.06
	(0.02)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.04)
Observations	674	674	674	674	673	673	673	673
R-squared	0.52	0.59	0.57	0.48	0.32	0.30	0.38	0.18
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

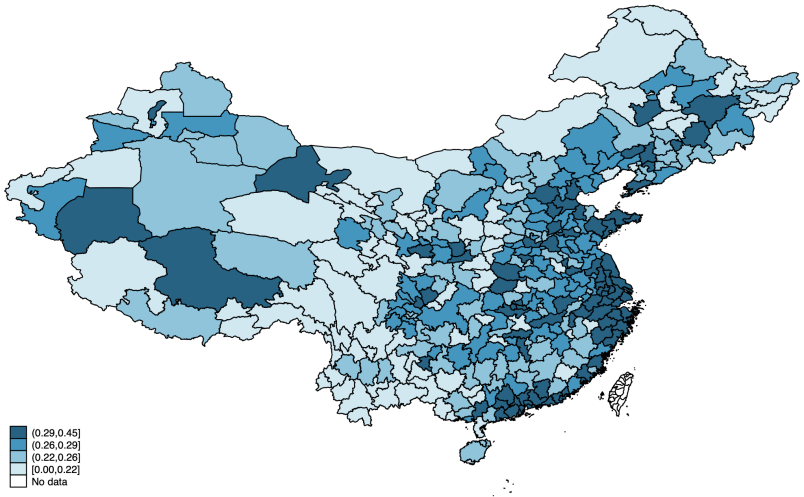
Notes: The dependent variable is the 10-year change in fraction on young people (aged 20-39) in labor markets. In Panel I, Columns (1)-(4) are fraction of young women or men (with agricultural hukou) employed, employed in agricultural sector, employed in manufacturing sector and employed in service sector; Columns (5)-(8) are fraction of young women or men (with non-agricultural hukou) employed, employed in agricultural sector, employed in manufacturing sector and employed in service sector. In Panel II, Columns (1)-(4) are fraction of young women or men (below high school education) employed, employed in agricultural sector, employed in manufacturing sector and employed in service sector; Columns (5)-(8) are fraction of young women or men (with high school education or above) employed, employed in agricultural sector, employed in manufacturing sector and employed in service sector.The sample contains 333 prefectures and four directly controlled municipalities × 2 gender groups (women or men). Columns 4-8 of Panel II have one less as the Golog Tibetan Autonomous Prefecture has no sample of high educated women in 2000. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Appendix



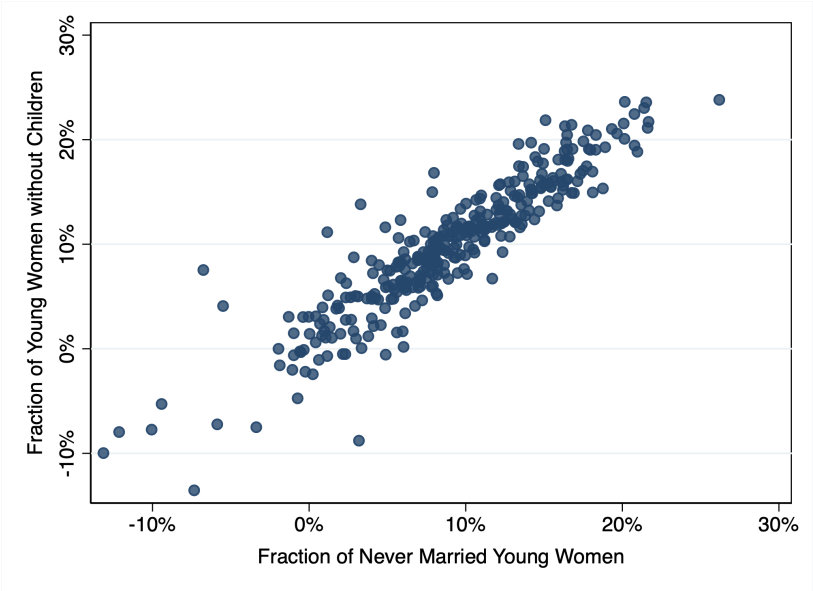
Notes: This figure presents the regional variation in decadal change in the fraction of young women aged 20-39 who are married. The darker prefectures experience greater increase, or less decrease in the share of married young women.

Figure A1: Regional Change in Share of Married Young Women, 2000-2010



Notes: This figure presents the regional variation in their exposure to trade shocks due to the U.S. granting of PNTR to China. The calculation of the regional PNTR exposure is given by equation (1), with darker prefectures experiencing greater trade shocks.

Figure A2: Regional exposure to PNTR



Notes: This figure presents the decadal change in the share of never married young women against that of young women without children across 337 Chinese prefectures between years 2000 and 2010.

Figure A3: Regional Change in Share of Young Women Never Married vs. without Children

Table A1: Descriptive Statistics

Variable	Mean	Std. Dev.	Min.	Max.	N
I. Dependent variables (10-Year Δ)					
(a) Year 2000-2010:					
Fraction of young women married	-0.09	0.06	-0.33	0.15	337
Fraction of young women divorced	0.00	0.01	-0.05	0.13	337
Fraction of young women never married	0.09	0.06	-0.13	0.26	337
Fraction of young men married	-0.09	0.06	-0.29	0.17	337
Fraction of young men divorced	0.00	0.01	-0.06	0.06	337
Fraction of young men never married	0.08	0.06	-0.16	0.28	337
Number of birth per young women	-0.27	0.17	-1.55	0.16	337
Fraction of young women with children	-0.10	0.06	-0.24	0.14	337
Number of birth per young women with children	-0.16	0.17	-1.5	0.23	337
Fraction of young women as students	0.04	0.03	-0.02	0.17	337
Fraction of young women as housekeepers	0.02	0.06	-0.20	0.17	337
Fraction of young women disabled or aged	0.00	0.01	-0.06	0.03	337
Fraction of young women as job seekers	-0.02	0.02	-0.19	0.05	337
Fraction of young men as students	0.04	0.70	-0.04	0.16	337
Fraction of young men as housekeepers	0.00	0.01	-0.03	0.08	337
Fraction of young men disabled or aged	0.00	0.01	-0.04	0.03	337
Fraction of young men as job seekers	-0.01	0.03	-0.17	0.05	337
Fraction of young women employed	-0.07	0.08	-0.30	0.24	337
Fraction of young women employed in agriculture	-0.15	0.11	-0.54	0.19	337
Fraction of young women employed in manufacturing	0.00	0.05	-0.16	0.21	337
Fraction of young women employed in service	0.08	0.07	-0.15	0.39	337
Fraction of young men employed	-0.05	0.05	-0.24	0.12	337
Fraction of young men employed in agriculture	-0.14	0.10	-0.46	0.22	337
Fraction of young men employed in manufacturing	0.01	0.06	-0.12	0.25	337
Fraction of young men employed in service	0.08	0.08	-0.14	0.42	337
Fraction of young women (agricultural hukou) married	-0.10	0.07	-0.35	0.16	337
Fraction of young women (agricultural hukou) divorced	0.00	0.01	-0.05	0.13	337
Fraction of young women (agricultural hukou) never married	0.09	0.07	-0.15	0.27	337
Fraction of young women (non-agricultural hukou) married	-0.07	0.10	-0.67	0.19	337
Fraction of young women (non-agricultural hukou) divorced	0.00	0.02	-0.1	0.11	337
Fraction of young women (non-agricultural hukou) never married	0.07	0.10	-0.19	0.67	337
Fraction of young women (high education) married	-0.15	0.13	-1	0.47	336
Fraction of young women (high education) divorced	-0.00	0.02	-0.13	0.11	336
Fraction of young women (high education) never married	0.15	0.13	-0.47	1	336
Fraction of young women (low education) married	-0.05	0.06	-0.35	0.15	337
Fraction of young women (low education) divorced	0.01	0.02	-0.04	0.14	337
Fraction of young women (low education) never married	0.04	0.05	-0.15	0.26	337

Table A1: Descriptive Statistics (Continued)

Variable	Mean	Std. Dev.	Min.	Max.	N
Fraction of young women (agricultural hukou) employed	-0.07	0.08	-0.31	0.21	337
Fraction of young women (agricultural hukou) employed in agriculture	-0.19	0.13	-0.68	0.26	337
Fraction of young women (agricultural hukou) employed in manufacturing	0.03	0.06	-0.15	0.35	337
Fraction of young women (agricultural hukou) employed in service	0.09	0.07	-0.06	0.37	337
Fraction of young women (non-agricultural hukou) employed	-0.04	0.11	-0.40	0.38	337
Fraction of young women (non-agricultural hukou) employed in agriculture	0.01	0.08	-0.28	0.50	337
Fraction of young women (non-agricultural hukou) employed in manufacturing	-0.07	0.07	-0.27	0.20	337
Fraction of young women (non-agricultural hukou) employed in service	0.03	0.13	-0.83	0.44	337
Fraction of young women (high education) employed	-0.11	0.10	-0.36	0.24	337
Fraction of young women (high education) employed in agriculture	-0.01	0.08	-0.19	0.50	337
Fraction of young women (high education) employed in manufacturing	-0.04	0.07	-0.25	0.20	337
Fraction of young women (high education) employed in service	-0.06	0.13	-0.67	0.46	337
Fraction of young women (low education) employed	-0.04	0.08	-0.48	-0.24	337
Fraction of young women (low education) employed in agriculture	-0.14	0.12	-0.60	0.29	337
Fraction of young women (low education) employed in manufacturing	0.02	0.06	-0.20	0.32	337
Fraction of young women (low education) employed in service	0.07	0.06	-0.12	0.34	337
Housing price	0.78	0.27	-0.57	1.53	293
Teacher-student ratio	0.01	0.03	-0.44	0.05	260
(b) Year 1990-2000:					
Fraction of young women married	0.01	0.05	-0.22	0.18	337
Fraction of young women divorced	0.01	0.01	0	0.12	337
Fraction of young women never married	-0.01	0.05	-0.22	0.22	337
Fraction of young men married	0.02	0.06	-0.14	0.25	337
Fraction of young men divorced	0.01	0.01	0	0.08	337
Fraction of young men never married	-0.02	0.06	-0.24	0.14	337
Number of birth per young women	-0.28	0.24	-1.22	1.19	337
Fraction of young women with children	0.03	0.06	-0.26	0.32	337
Number of birth per young women with children	-0.43	0.26	-1.67	0.33	337
II. Explanatory variables					
NTR Gap	0.25	0.07	0	0.45	337
NTR Gap × (female share)	0.12	0.05	0	0.27	337
NTR Gap × (male share)	0.14	0.03	0	0.21	337
III. Control variables					
Fraction of employment in manufacturing	0.11	0.10	0	0.75	337
Output tariff change	-0.08	0.03	-0.24	0	337
Input tariff cuts	-0.05	0.01	-0.09	0	337
Fraction of young women employed	0.82	0.11	0.45	0.97	337
Fraction of population with agricultural hukou	0.73	0.15	0.15	0.94	337
Fraction of minority population	0.16	0.26	0	0.99	337
Fraction of population with college or above education	0.03	0.03	0	0.17	337
Processing zone dummy	0.02	0.13	0	1	337

Notes: This table provides descriptive statistics for the variables used in the main empirical analyses.

Table A2: Heterogeneous Effects of PNTR on Young Women's Fertility Decisions

	Number of Children per Woman	Fraction of Women with Children	Number of Children per Woman with Children	Number of Children per Woman	Fraction of Women with Children	Number of Children per Woman with Children
	(1)	(2)	(3)	(4)	(5)	(6)
I. By Hukou		Agricultural			Non-agricultural	
NTR Gap	0.01	-0.14*	0.24	-0.09	-0.04	0.01
	(0.22)	(0.08)	(0.17)	(0.22)	(0.13)	(0.16)
Observations	337	337	337	337	337	337
R-squared	0.57	0.58	0.66	0.34	0.46	0.35
II. By Edu		Low			High	
NTR Gap	-0.05	-0.11	0.17	-0.15	-0.13	0.20
	(0.20)	(0.08)	(0.16)	(0.13)	(0.11)	(0.16)
Observations	337	337	337	337	337	334
R-squared	0.72	0.82	0.66	0.73	0.73	0.36
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The dependent variable is the 10-year change in fertility decisions of young women (aged 20-39). The full sample contains 333 prefectures and four directly controlled municipalities. Column 6 in Panel II contain only 330 prefectures and four directly controlled municipalities as three prefectures have no sample of young women with children. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table A3: Equivalent Shift-level Estimates

(a) Shock summary statistics

Mean	SD	Interquartile Range	Effective sample		Largest s_j weights		No. of		ICCs within	
			CIC4	CIC3	CIC4	CIC3	CIC4	CIC3	CIC3	CIC2
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
0.27	0.16	0.25	76.31	54.98	0.06	0.07	423	161	0.15	0.05

Notes: This table reports the summary statistics of shocks. Columns (1)-(7) summarize the distribution of shock g_j across industries. Shocks are measured by the NTR gap. All statistics are weighted by the average industry exposure shares s_j. Columns (8)-(9) report the number of industries at CIC4 and CIC3 level, repsectively. Columns (10)-(11) reports the intra-class-correlation coefficients (ICCs) of shocks within CIC3 group and CIC2 group, respectively.

(b) Shock balance tests

Balance variables	Coef.	SE
	(1)	(2)
Industry level log wage, 1999	0.00	(0.03)
City level fraction of young women employed, 2000	0.02	(0.01)
City level fraction of population with agricultural hukou, 2000	0.02	(0.02)
City level fraction of minority population, 2000	-0.09***	(0.03)
City level fraction of population with college or above education 2000	-0.00	(0.00)
City level processing zone dummy, 2000	0.00	(0.01)
City level fraction of young women currently married, 1990-2000	-0.00	(0.00)

Notes: The first row of this table reports coefficients from regressions of the industry-level covariates on the NTR Gap, weighted by average industry exposure shares, with standard errors (provided in parentheses) clustered at the CIC3 level. The remaining rows of this table report coefficients from regression of the city-level covariates and pre-trends on the shift-share, controlling for manufacturing employment share in 2000. CIC3-clustered exposure-robust standard errors are reported and obtained from equivalent industry-level regressions.

(c) Shift-share estimates from an equivalent shift-level regression

	Married	Divorced	Never Married	Number of Children per Woman	Fraction of Women with Children	Number of Children per Woman with Children
	(1)	(2)	(3)	(4)	(5)	(6)
Coefficient	-0.23***	0.02***	0.20**	-0.13	-0.20**	0.18
	(0.09)	(0.01)	(0.08)	(0.21)	(0.08)	(0.20)
Baseline Controls	Yes	Yes	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
No. of regions	335	335	335	335	335	335
No. of industries	423	423	423	423	423	423

Notes: This table reports the shift-share coefficients from regression of the change in regional young women's marital and fertility decisions on the city's exposure to PNTR. Exposure-robust standard errors (provided in parentheses) are obtained from equivalent industry-level IV regression, allowing for clustering of shocks at the CIC3 level. The sample includes 335 prefecture-level cities (two prefectures have no sample of manufacturing firms in ASIF data thus are not included in transformation to equivalent industry-level regression) and 423 industries.

Table A4: Effects of PNTR on Gender Gaps in Labor Markets After Marriage

	Employed	Agriculture	Manufacturing	Service
	(1)	(2)	(3)	(4)
NTR Gap × Female	0.19***	-0.02	-0.23***	0.41***
	(0.05)	(0.09)	(0.08)	(0.08)
NTR Gap	0.02	-0.07	0.29***	-0.18**
	(0.06)	(0.09)	(0.06)	(0.08)
Female	-0.10***	-0.01	0.05***	-0.13***
	(0.01)	(0.02)	(0.02)	(0.02)
Baseline controls	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes
Observations	674	674	674	674
R-squared	0.55	0.62	0.59	0.49

Notes: The dependent variable is the 10-year change in the fraction of ever married young people (aged 20-39) in labor markets. Columns (1)-(4) are fraction of young women (men) employed, employed in agricultural sector, employed in manufacturing sector, employed in service sector. The sample contains 333 prefectures and four directly controlled municipalities × 2 gender groups (women or men). Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table A5: Effects of PNTR on Young Women's School Enrollment

	Aged 20-39	Aged 20-24	Aged 20-39	Aged 20-24	Aged 20-39	Aged 20-24
	(1)	(2)	(3)	(4)	(5)	(6)
	At School		High School Educated		College-educated or Above	
NTR Gap	0.02	0.08	0.05	0.10	0.02	-0.01
	(0.03)	(0.09)	(0.04)	(0.09)	(0.04)	(0.10)
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	337	337	337	337	337	337
R-squared	0.64	0.47	0.69	0.47	0.81	0.55

Notes: The dependent variable is the 10-year change in school enrollment of young women. Columns (1)-(2) are fraction of young women currently at school. Columns (3)-(4) are fraction of young women with high school education. Columns (5)-(6) are fraction of young women with college education or above. The sample contains 333 prefectures and four directly controlled municipalities. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table A6: Effects of PNTR on Labor Markets' Segmentation

	Changes in Gender Concentration Index (GCI), 2000-2010			
	CIC 3-digit Industry		CIC Aggregate Sector	
	(1)	(2)	(3)	(4)
NTR Gap	-0.00		-0.03**	
	(0.03)		(0.01)	
NTR Gap × (female share)		-0.05		-0.05*
		(0.05)		(0.03)
NTR Gap × (male share)		0.07		-0.00
		(0.06)		(0.02)
Baseline controls	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes
Observations	337	337	337	337
R-squared	0.52	0.53	0.32	0.32

Notes: The dependent variable is the 10-year change in local labor markets' segmentation. Columns (1)-(2) use the 3-digit industry (CIC3) to construct the segmentation index. Columns (3)-(4) use the sector (agriculture, manufacturing, service and others) to construct the segmentation index. The sample contains 333 prefectures and four directly controlled municipalities. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table A7: PNTR and the Costs of Marriage and Childbearing

	Housing Price	Young Women Married	Teacher-student Ratio	Young Women Married	Young Women Married
	(1)	(2)	(3)	(4)	(5)
NTR Gap	-0.76**	-0.17**	0.02	-0.16**	-0.15**
	(0.33)	(0.07)	(0.05)	(0.09)	(0.08)
Housing price		-0.01			-0.01
		(0.02)			(0.02)
Teacher-student ratio				0.16***	0.16***
				(0.04)	(0.04)
Baseline controls	Yes	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	291	291	260	260	255
R-squared	0.52	0.60	0.25	0.61	0.62

Notes: The dependent variables are the 10-year change in housing price (local CPI deflated), fraction of young women married, teacher-student ratio (elementary school), and fraction of young women married. The full sample contains 333 prefectures and four directly controlled municipalities. The missed samples are due to lack of data. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

Table A8: Migration as a Mechanism

(a) Inflow of Migrants

	Migrants	Married	Divorced	Never Married
	(1)	(2)	(3)	(4)
NTR Gap	0.07*	-0.15**	0.02***	0.12*
	(0.04)	(0.06)	(0.01)	(0.06)
Migrants		-0.29***	-0.00	0.31***
		(0.11)	(0.01)	(0.11)
Baseline controls	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes
Observations	337	337	337	337
R-squared	0.82	0.59	0.16	0.61

Notes: The dependent variable in column (1) is the 10-year change in fraction of young people (aged 20-39) who are migrants. The dependent variables in column (2)-(4) are the 10-year change in fraction of young women (aged 20-39) currently married, divorced or widowed, never married. The sample contains 333 prefectures and four directly controlled municipalities. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.

(b) Gender Imbalances

	Sex Ratio of Migrants	Sex Ratio	Female Locals to Male Locals and Male Migrants	Male Locals to Female Locals and Female Migrants
	(1)	(2)	(3)	(4)
NTR Gap	3.11*	0.03	-0.15	0.00
	(1.65)	(0.10)	(0.11)	(0.09)
Baseline controls	Yes	Yes	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes
Observations	320	337	337	337
R-squared	0.10	0.47	0.61	0.49

Notes: The dependent variable in column (1) is the 10-year change in the sex ratio of young migrants (aged 20-39). The dependent variable in column (2) is the 10-year change in the sex ratio of young people (aged 20-39). The dependent variable in column (3) is the 10-year change in the ratio of young women locals (aged 20-39) to total young men (aged 20-39). The dependent variable in column (4) is the 10-year change in the ratio of young men locals (aged 20-39) to total young women (aged 20-39). The sample in column (1) contains 316 prefectures and four directly controlled municipalities as 17 prefectures have no sample of young female migrant in year 2000 or 2010. The sample in columns 2-4 contains 3333 prefectures and four directly controlled municipalities. Robust standard errors are provided in parentheses. Models are weighted by the beginning-of-period prefecture young population (aged 20-39). *** p<0.01, ** p<0.05, * p<0.1.