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Trade liberalization and gender gaps in local labor market outcomes: Dimensions of adjustment in the United States

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We provide empirical results that trade liberalization with China reduced gender gaps in local U.S. labor markets. In MSAs with higher exposure to trade liberalization, the simple wage gender gap decreased, while the residual wage gap increased, indicating important selection effects in labor force participation decisions. The reduction in the gender labor force participation gap was driven by higher entry of women, in particular more educated women, and exit of the less educated men. This results in intrahousehold adjustments in work dynamics, with women entering the labor force to offset the lost income of male partners who left the labor force. We show that trade liberalization increased female workers' unemployment rate and reliance on part-time jobs.

1. Introduction

This paper considers the impact of trade liberalization on gender inequality in local labor markets. We examine the effect of the U.S. granting China Permanent Normal Trade Relations (PNTR) on the gender wage and labor-force-participation gaps in the U.S. We show that the liberalization of U.S. trade relationship with China *reduced* both gaps in local labor markets. These changes occurred because women entered the labor force, while men left the labor force as the exposure to China receiving PNTR status affected manufacturing industries more which tend to employ relatively more men. Our results indicate that women entered the labor force in part to offset the reduction in family income that occurred as their male partners lost jobs and left the labor market.

We use the Pierce and Schott (2016, 2020) approach to measure a local labor market's exposure to trade liberalization. Pierce and Schott (2016) show that granting permanent normal trade relations to China in 2001 caused a sharp decline in U.S. manufacturing employment in the 2000s. They argue that Chinese exporters faced significant risks of increased tariffs before 2001 since China's Normal Trade Relations (NTR) status, guaranteeing low tariffs when exporting to the U.S., required annual renewals by the Congress. Pierce and Schott (2016) define PNTR exposure as the difference between the high non-NTR tariffs and the much lower NTR tariffs, which averages to 33 percentage points in 1999. The conferral of PNTR status to China eliminated tariff uncertainty and brought the decline in U.S. employment by encouraging Chinese exporters to scale up

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and U.S. firms to do more offshoring/outsourcing. In a follow-up paper, Pierce and Schott (2020) calculate the exposure level to PNTR for each U.S. county and show that a county's higher exposure to PNTR is associated with increases in mortality from stress-related causes (e.g., suicides), specifically among white males.

Our investigation focuses on Metropolitan Statistical Areas (MSAs), which are defined along the lines of local labor markets with sufficiently high population density at their core. We find that a Metropolitan Statistical Area (MSA) with higher PNTR exposure experienced a decrease in gender gaps in wage and labor force participation rates (LFPR) after trade liberalization with China. We examine both the simple and residual wage gender gaps. The simple gender wage gap decreased after China received PNTR status, and this finding is unaffected once we control for selection. However, we also show that there were some selection effects in the wake of China receiving PNTR status. The residual wage gap actually increased, showing that, once controlling for observable worker characteristics such as education levels, men earned higher wages than women. The two results are reconcilable as long as this shock made the male labor market relatively more competitive by generating different selections for the two genders as less educated men left the labor force or more educated women entered the labor force in greater numbers. We find evidence of both of these effects. MSAs with higher PNTR exposure experienced larger increases in female relative labor force participation rates, and this increase was largely driven by the entry of more educated women into the labor force, where the more educated women are those with at least some college-level education. We also find that labor force participation rates decreased for men, somewhat more strongly for the less educated men.

The greater participation in the labor force by women was a response to how China receiving PNTR status affected households. We show that MSAs with higher exposure did not experience change in the share of households with both spouses working, but there was a significant decrease in the share of households with only the husband (or male partner) working and an increase in households with just the wife (or female partner) working. In addition, the share of family income accounted for by the female partner has increased in MSAs with greater exposure to China's PNTR status. The greater labor force participation rates of women also resulted in higher unemployment for women, while males exiting the labor force did not prevent higher unemployment among them which was also accompanied by a reduction in employment among men. Furthermore, in MSAs with higher PNTR exposure, female workers spent more time in part-time employment. The increase in part-time employment is a consequence of individuals not being able to find full-time employment, rather than wanting to work only part-time.

Other studies have found significant effects of trade liberalization on labor market outcomes and gender inequality using different methodologies. Autor et al. (2013) explain the decline in U.S. manufacturing employment with Chinese import penetration. Their estimation strategy, to instrument the growth of Chinese exports to the U.S. by the growth of Chinese exports to other high-income countries, was adopted by a number of follow-up papers. Following their identification strategy, Brussevich (2018) shows that U.S. commuting zones with higher import penetration show greater reduction in the gender wage gap and that wage and welfare gains from trade are higher for female workers since the import competition shock in manufacturing sector disproportionally affected the labor market outcomes of the two gender groups. Greenland et al. (2019) found that commuting zones with greater exposure to China gaining PNTR status experienced reduced population growth, particularly for men. Benguria and Ederington (2017) show that increased competition from China lowered the gender wage gap in Brazil, which is driven by the underperformance of male workers.

Other trade liberalization episodes have also been found to have reduced gender wage gaps. Aguayo-Tellez et al. (2013) show that tariff reductions, accompanied by the North American Free Trade Agreement (NAFTA), increased the demand for female labor and raised their relative wage in Mexico. Juhn et al. (2014) find that NAFTA raised the relative wage and employment of female workers, especially in blue-collar tasks in Mexico. They explain that higher competition encouraged firms to modernize their technology, thus reducing their dependence on physical ability. Ederington et al. (2009) report that tariff reductions in Chile, as a result of Chile's entry into GATT/WTO, raised the number of female workers (relative to male ones) in blue-collar jobs. Similar to Brussevich (2018), Black and Brainerd (2004) show that import penetration is associated with greater reductions in the gender wage gap in the U.S.

Some researchers have shown that gender inequality can increase with trade liberalization. Sauré and Zoabi (2014) report that the formation of NAFTA widened the gender gaps in the U.S. labor market. Since female intensive sectors tend to be capital intensive, trade liberalization between capital-rich and capital-poor countries may raise the gender gap in the capital-rich country (i.e., the U.S.) by reallocating male workers into capital intensive sectors. Bøler et al. (2018) report higher wage gaps for exporting firms compared to non-exporters in Norway. They claim that exporters may require a greater commitment from employees. Hence, female workers, who tend to have less flexible schedules, receive lower relative wages.

Gender inequality and trade liberalization have been an important issue investigated in the literature. This paper shares similar insights and findings with previous papers in part. Our paper shares similar insights with Brussevich (2018), Benguria and Ederington (2017), and Sauré and Zoabi (2014), in that trade liberalization affects gender inequality in the labor market through reallocation of male labor from the male-intensive sector. However, our paper makes an important contribution since we explain how the various channels are connected by examining both the simple and residual wage gaps. We explore gender gaps both in wage and labor force participation whereas most existing papers focus on wage inequality. We also propose possible mechanisms through multiple channels including education level and marital status, the former of which help explain the selection effects induced in local labor markets by China receiving PNTR status. We also provide additional insights into the effects of trade liberalization in developed economies and show that with respect to the effect on gender gaps, trade liberalization has similar effects in developed and developing countries. The latter conclusion is based

Table 1
CPS sample composition across time.

	1990			2007			2013		
	Female	Male	Total	Female	Male	Total	Female	Male	Total
By sector									
Manufacturing	5.3%	9.9%	15.2%	3.4%	7.0%	10.4%	2.8%	6.4%	9.2%
Services	29.7%	24.0%	53.7%	34.0%	26.4%	60.4%	34.3%	27.1%	61.3%
Other	17.2%	14.0%	31.2%	14.8%	14.4%	29.2%	15.2%	14.2%	29.4%
By education									
No college experience	31.0%	26.1%	57.1%	20.1%	20.5%	40.6%	17.9%	19.1%	37.0%
Has college experience	21.2%	21.7%	42.9%	32.1%	27.3%	59.4%	34.4%	28.7%	63.0%
Total	52.2%	47.9%	100.0%	52.2%	47.8%	100.0%	52.3%	47.8%	100.0%

on the similarity between our results for the U.S. and Benguria and Ederington's (2017) results for Brazil. We show that trade liberalization has affected the overall quality of female jobs by increasing part-time work at the expense of full-time employment.

Our paper is related to Charles et al. (2019), who show that the manufacturing decline in a local labor market in the 2000s had negative effects on local employment rates, hours worked, and wages. They find larger negative effects employment of men and for the less educated workers. Their findings are consistent with ours in that the negative effects from China gaining permanent normal trade relations are larger for male workers and the less educated group.

Lastly, our paper is related to a set of papers which examine intrahousehold or intrafamily adjustments to trade liberalization and offers new and interesting findings. Autor et al. (2015) find that local labor markets more exposed to Chinese imports experience a decrease in employment, especially in manufacturing and among non-college workers. Autor et al. (2019) examine the effects of increased competition from China on young adults in the U.S. finding an increase in male idleness and premature mortality as well as reductions in marriage and fertility, and an increase in the fraction of single mothers who are heads of households, as well as an increase in the number of children living in poverty. Keller and Utar (2019) show that increased competition from Chinese imports in Denmark resulted in a shift towards the family, with an increase in parental leave, fertility, marriage, and a reduction in divorce rates.[1] We find opposite results the married women in the U.S. entering the labor force and to some extent shifting away from the family. Hakobyan and McLaren (2017) show that the impact of trade liberalization on female wages depends on marital status. They observe that married low-skilled women experienced larger reductions in wage growth with respect to NAFTA tariff reductions than other demographic groups since high-skilled women drop out of the labor market. Our paper indicates that the shock due to China gaining PNTR status was different than NAFTA as the former induced more educated and higher-skilled women to enter the labor force. Thus, our results stand in contrast to both Keller and Utar (2019) and Hakobyan and McLaren (2017) as we examine the reduction in gender gaps by exploring monetary incentives of married females who had to compensate for lost family income from the negative income shock. The shock created by China receiving PNTR status is well recognized to have brought important changes to the U.S. society. Our paper provides a good understanding of this important event from the perspective of female workers.

2. Data

2.1. Labor market outcomes

We measure labor market outcomes using Current Population Survey's (CPS) Annual Social and Economic Supplement (ASEC), a nationally representative household data with detailed information of each household member's earnings, work hours, gender, and race, among other indicators. We restrict our sample to individuals who are older than 25 and younger than 64 and are either in or out of the labor force for reasons not related to being on active military duty or having a disability.[2] Our sample in 2000 includes 49,700 individuals who resided in 272 MSAs with trade liberalization exposure data. Table 1 shows the demographic characteristics of our sample in three years, 1990, 2007, and 2013. Women account for 52% of our sample in every year. Male workers' dependency on manufacturing in the labor market is almost twice as large as female workers' before and after the conferral of permanent trade relations to China.[3] Artuç et al. (2010) show that high switching costs across sectors slow down the readjustment of an economy in response to a trade shock. Hence, the

[1] Using our data, we examined the effect on divorce rates and childbirth rate, but find no significant effects. These results are available on request.

[2] For variables that require only those who are working, we further restrict the sample to include workers with a stronger attachment to the labor market – worked for more than 20 weeks in the previous year, and more than 35 h per week in the previous year (both inclusive), following Maasoumi and Wang (2019).

[3] We can observe a worker's industry from industry codes provided by CPS. We classify a worker's industry as services if she or he is in transportation, communications, public utilities, wholesale trade, retail trade, finance, insurance, real estate, business and repair services, personal services, entertainment and recreation services, or professional and related services. If the worker does not belong to either manufacturing or services, we assign her or him "other" sector.

Table 2
Summary statistics of CPS variables (means).

	1990			2007			2013		
	Female	Male	All	Female	Male	All	Female	Male	All
Wage	9.52	12.66	10.79	17.96	22.62	21.41	20.47	25.65	23.16
	(1.69)	(2.29)	(2.61)	(3.62)	(5.50)	(13.35)	(4.35)	(5.80)	(6.55)
Work hours	1705.6	2113.6	1920.8	1818.6	2132.9	1994.0	1794.5	2081.1	1937.9
	(139.8)	(123.4)	(113.2)	(127.0)	(107.8)	(91.3)	(140.1)	(146.3)	(99.7)
Labor force	70.18%	89.92%	79.72%	77.67%	91.64%	84.41%	77.37%	89.76%	83.33%
Participation rate	(0.10)	(0.05)	(0.07)	(0.08)	(0.05)	(0.52)	(0.09)	(0.06)	(0.06)

Standard deviation in parentheses.

heavier dependency on manufacturing sector on the part of male workers may have aggravated the impact on the male labor market. Table 1 displays the familiar patterns by now: the share of manufacturing in the U.S. economy has steadily declined, from 15.2% of individuals in 1990 to 9.2% in 2013. This decrease has occurred at the expense of services which have grown from 53.7% of individuals to 61.3%.

To understand the differential impact by educational attainment, we divide all individuals in our sample into two levels of education, those without any college education and those with at least some, even if they do not have a college degree.[4] About 43% of our sample had some college education in 1990 while the education levels of two gender groups are not significantly different from each other. By the end of our sample the share of individuals without college education has declined to 37%, with women comprising a larger share of individuals with some education by 2013.

Our variables of interest, summarized in Table 2, include the average hourly wage, w_{it}, the number of total hours worked, h_{it}, and labor force participation status, l_{it}, of an individual i in year t. The number of total work hours, h_{it}, is calculated by multiplying "Usual work hours worked per week last year" and "Weeks worked last year," and reflects the total number of hours the individual spent working in the previous year. The hourly wage, w_{it}, is calculated as "Wage and salary income" divided by the number of total hours worked, h_{it}. We directly observe an individual's labor participation status, l_{it}, from the variable "Labor force status" and consider individual i is in the labor force if she or he worked, was looking for a job, or was temporarily absent/laid-off during the reference period.

In 1990, the average hourly wage for the whole population was \$10.79, with the average female hourly wage of \$9.52, equivalent to 75% of the average male hourly wage. In 2007, the female worker's average wage increased to about 79% of the male worker's and to 80% in 2013.[5] We observe similar patterns for other variables. In 1990 female workers' average work hours were equivalent to 81% of male average work hours. It increased to 85% by 2007 and 86% in 2013. In 1990 the labor force participation rate (LFPR) of female workers was about 20 percentage points lower than that of their male counterparts. In 2007, the gap decreased to about 15 percentage points and to 13% by 2013. Our data reveal the well-known patterns: while male workers tend to outperform female workers, gender gaps have been on the decline.

We are interested in regional differences in changes in gender gaps. For this purpose, we calculate weighted averages of the above variables at MSA levels:[6]

$$w_{mt}^S = \frac{1}{\hat{N}_{mt}^S}\sum_{i\in S\cap m} w_{it},\ \ l_{mt}^S = \frac{1}{N_{m,t}^S}\sum_{i\in S\cap m} l_{it},\ \ h_{mt}^S = \frac{1}{\hat{N}_{mt}^S}\sum_{i\in S\cap m} h_{it}$$

where m indicates an MSA, S refers to demographic groups, $N_{m,t}^S$ is the number of individuals of group S who resided in MSA m in year t, and $\hat{N}_{mt}^S$ is the number of individuals with a positive w_{it}. For example, if we let F be the set of females, then l_{mt}^F refers to labor force participation rate of females in MSA m and year t. Unlike Pierce and Schott (2020) who conduct their analysis using county-level data, we perform our analysis using MSA-level data. As we are interested in labor market outcomes, we conducted our analysis at the MSA level since they are largely defined by boundaries of local labor markets.[7]

2.2. NTR gap

Our measure of exposure of an MSA to trade liberalization follows Pierce and Schott (2020). Their measure is based on the difference between two tariff rates in the U.S. tariff schedule that could be assessed on imports from China. Imports from a country which does not have normal trade relations with the U.S. are assessed tariff rates established by the Smoot–Hawley Tariff Act of 1930. These rates are significantly higher than the normal trade relation tariffs rates, which are assessed

[4] Our definition for education group follows Maasoumi and Wang (2019), who split their samples into four classes: below high school education, high school degree, some college experience, and above college degree. We aggregate their classifications into two groups, at least some college education and less than college education, due to the insufficient number of individuals in an MSA if we apply the same classification as Maasoumi and Wang (2019).

[5] The median female wage increased from 75% of the male wage in 1990 to 80% in 2010.

[6] When we calculate MSA-level variables, we use ASEC ASECWT weights as CPS suggests.

[7] Using Commuting Zones (CZs), which cover the entire country, may be a better option, but not a feasible one since CPS's county identifier needed to calculate CZ-level labor market outcomes is only available since 1996.

Table 3
MSAs with the highest and lowest NTR Gaps.

Rank	Metropolitan statistical area	NTR Gap
1	Hickory-Morganton, NC	0.235
2	Burlington, NC	0.212
3	Danville, VA	0.198
4	Elkhart-Goshen, IN	0.170
5	Rocky Mount, NC	0.165
316	Bismarck, ND	0.038
317	Billings, MT	0.038
318	Great Falls, MT	0.036
319	Las Vegas, NV	0.034
320	Farmington, NM	0.031

on imports from countries that are members of the World Trade Organization (WTO). China was first granted temporary NTR status in 1980 with a provision that its status be reaffirmed on an annual basis. The uncertainty associated with renewal was a function of various crises in U.S.–China relations during the 1990s. China was finally granted permanent normal trade relations with the U.S. in October 2000 as a prelude to its entry into the WTO in December 2001.

We follow Pierce and Schott (2020)'s methodology to measure a local labor market's exposure to trade liberalization. We start with their industry-level measure, $NTR\ Gap_j$, defined as the difference between non-NTR rates and NTR rates in a six-digit NAICS sector j:

$$NTR\ Gap_j = non\text{–}NTR\ tariff_j - NTR\ tariff_j.$$

$NTR\ Gap_j$ refers to the potential tariff increase on Chinese imports and captures the uncertainty faced by Chinese exporters in industry j.

Using $NTR\ Gap_j$, Pierce and Schott (2020) calculate a county's exposure to PNTR. We follow the same steps and calculate the exposure to PNTR for metropolitan areas:

$$NTR\ Gap_m = \sum_j \frac{L_{jm}^{1990}}{L_m^{1990}} NTR\ Gap_j$$

where L_{jm}^{1990} refers to the number of employees in sector j in MSA m in the year 1990 and L_m^{1990} refers to the total number of workers in MSA m. The information about employment weights, L_{jm}^{1990} and L_m^{1990}, are from the County Business Patterns (CBP), an annual dataset with information on employment and payroll by sector and county. Higher $NTR\ Gap_m$ indicates a higher exposure of MSA m to trade liberalization with China. $NTR\ Gap_m$ has a mean 0.145 and standard deviation 0.05. Table 3 lists the MSAs with the highest and lowest $NTR\ Gap_m$, while Fig. 1 illustrates the geographical distribution of $NTR\ Gap_m$. MSAs in the mid-Atlantic and Midwest regions are more exposed to trade liberalization.

3. Estimation

3.1. DID identification strategy

Our estimation strategy follows Pierce and Schott (2020). Our baseline difference-in-differences (DID) specification examines whether MSAs more exposed to PNTR (first difference) experience differential changes in gender gaps in wage and labor force participation rates after the change in U.S. trade policy versus before (second difference),

$$LHS_{mt} = \theta \cdot Post\ PNTR_t \times NTR\ Gap_m + \beta \boldsymbol{X}_{mt} + \gamma \cdot PostPNTR_t \times \boldsymbol{Z_m} + \delta_m + \delta_t + \varepsilon_{mt}. \tag{1}$$

The dependent variable LHS_{mt}, a local labor market outcome such as the gender wage or labor-force-participation-rate gap, is defined in year t for an MSA m. The first term on the right-hand side is the DID term of interest, an interaction of a post-PNTR (i.e., $t > 2000$) indicator with the (time-invariant) MSA-level NTR Gap. $\boldsymbol{X}_{mt}$ represents the (time-varying) overall U.S. import tariff rates associated with the industries active in the MSA as well as exposure to the elimination of the Multi-Fibre Agreement quotas which took place in 2002 and 2005. $\boldsymbol{Z}_m$ represents the initial-period MSA attributes, 1990 median household income, 1990 share of population without any college education, 1990 share of population that are veterans, and its exposure to changes in Chinese imports tariffs. δ_m and δ_t refer to MSA and year fixed effects. We cluster standard errors at MSA levels. The sample period is 1990 to 2013 as in Pierce and Schott (2020).

3.2. Estimates of the gender wage gap

We begin our analysis by focusing on the gender wage gap. We find that MSAs with higher NTR Gap and higher exposure to PNTR have lower labor market gender gaps after the conferral of PNTR in 2001. We use specification (1) with female-male wage ratio, w_{mt}^F / w_{mt}^M, the gender wage gap, as the dependent variable and estimate the DID point estimate of interest, θ. The

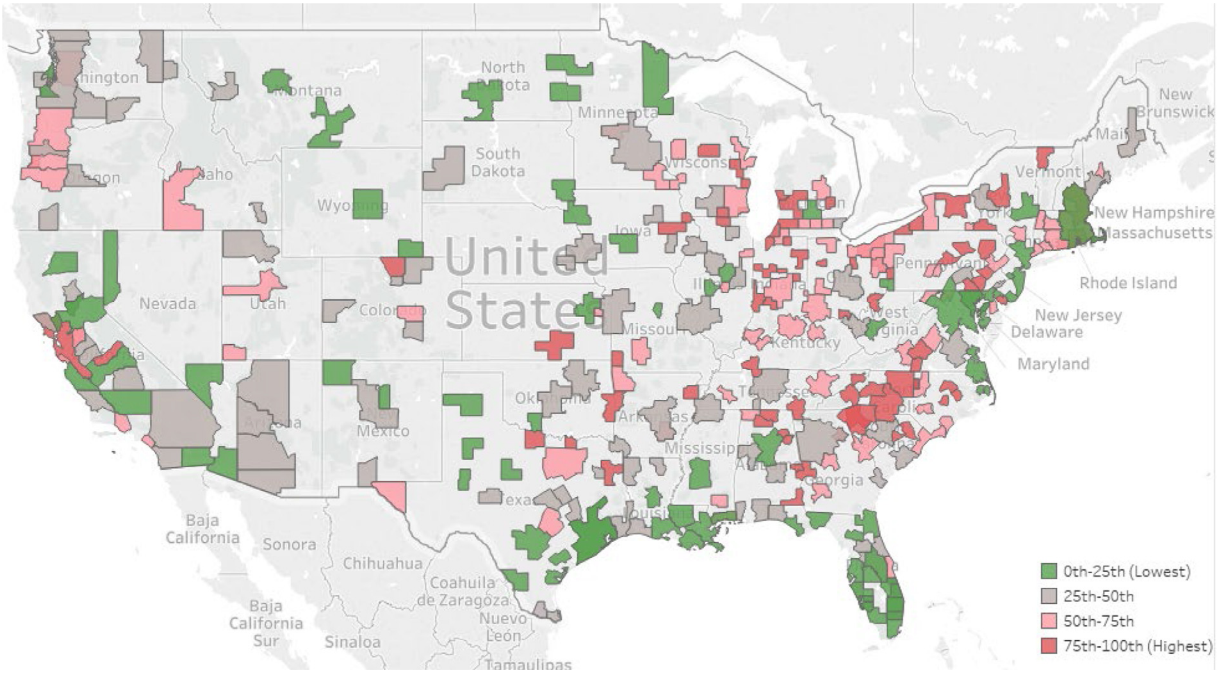

Fig. 1. Geographic distribution of $NTR\ Gap_m$.

Table 4
Gender wage gap and wages.

	All data (benchmark sample)				Machado sample	
	w^F_{mt}/w^M_{mt}		$\log(w^F_{mt})$	$\log(w^M_{mt})$	w^F_{mt}/w^M_{mt}	
Post * NTR Gap	0.418*	0.678**	0.597*	−0.206	−0.705	1.447*
	(0.241)	(0.338)	(0.343)	(0.413)	(0.948)	(0.828)
NTR rate		1.632*	1.166	−0.567		7.261*
		(0.854)	(0.749)	(0.780)		(3.875)
MFA rate		1567	−628.0	−2391		9929
		(3202)	(3180)	(4623)		(13,200)
Post * Chinese tariff		−0.0371	0.834	0.942		2.965
		(0.605)	(0.720)	(0.841)		(3.202)
Post * No College		−0.000324	−0.167***	−0.169***		0.211
		(0.0638)	(0.0595)	(0.0612)		(0.274)
Post * Veteran		0.235*	0.235*	−0.00420		0.890
		(0.141)	(0.120)	(0.107)		(0.794)
Post * Median HHI		0.0291	0.0639***	0.0318		0.198
		(0.0221)	(0.0209)	(0.0252)		(0.121)
Observations	5429	5356	5356	5357	5372	5302
R^2	0.124	0.129	0.781	0.741	0.067	0.071

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

first and second columns of Table 4 report the results for the female-male wage ratio. The first column reports coefficient estimates for a specification containing just the DID term of interest and fixed effects. The second column adds controls for policy changes $\boldsymbol{X}_{mt}$ and demographic variables $\boldsymbol{Z}_m$.[8] The DID point estimates of interest are positive ($\theta > 0$) and statistically significant at conventional levels across all columns. Our empirical results suggest that higher import competition from China is associated with a lower gender wage gap in U.S. labor market outcomes, as in Brussevich (2018).

Our main question is to understand which changes in local labor market outcomes drive changes in gender gaps. As a first step, we examine changes in wages of women and men. We estimate specification (1) by using the log value of the wage, $\log(w^S_{mt})$, for each gender $S \in \{F, M\}$ as dependent variables and collecting results in columns 3 and 4 of Table 4. Our results suggest that female wages increased, while that of men may have decreased, though the latter coefficient is not estimated

[8] Given the number of tables and results in the paper, in certain tables we only report the estimates from the specification with additional controls out of concerns for space. All other results are available on request.

Table 5
Sectoral differences in the wage gender gap and wages.

	w^F_{mt}/w^M_{mt}			Wages					
	Manufacturing	Services	Other	Manufacturing		Services		Other	
				Female	Male	Female	Male	Female	Male
Post * NTR Gap	−1.670**	0.891**	−0.847	−1.111	−0.0863	0.541	−0.358	1.107	0.761
	(0.836)	(0.444)	(1.020)	(0.814)	(0.646)	(0.401)	(0.460)	(0.875)	(0.517)
NTR rate	−3.198*	3.028**	0.803	−2.657	−1.483	1.009	−1.749	2.042	1.433
	(1.785)	(1.196)	(2.025)	(1.769)	(1.308)	(0.818)	(1.085)	(1.733)	(1.202)
MFA rate	11,610	5562	−25,908**	9438	−4745	−2283	−8433	−3426	15,760**
	(8335)	(4847)	(12,471)	(8926)	(7078)	(3977)	(6363)	(8353)	(6311)
Post * ΔChinese tariff	−0.559	−0.0815	0.919	0.611	1.018	0.679	0.899	2.083	0.979
	(1.470)	(0.868)	(2.089)	(1.302)	(1.156)	(0.794)	(1.019)	(1.441)	(1.079)
Post * No College	−0.0855	0.0231	−0.0610	−0.201	−0.189*	−0.190***	−0.212**	0.0599	0.0678
	(0.232)	(0.0818)	(0.171)	(0.159)	(0.108)	(0.0649)	(0.0829)	(0.121)	(0.0962)
Post * Veteran	0.223	0.288*	−0.130	−0.181	−0.342*	0.286**	0.0825	−0.284	−0.217
	(0.403)	(0.174)	(0.374)	(0.310)	(0.199)	(0.136)	(0.146)	(0.239)	(0.197)
Post * Median HHI	−0.0150	0.0353	−0.0438	0.0834	0.0237	0.0684***	0.0262	0.00343	0.0561
	(0.0753)	(0.0319)	(0.0661)	(0.0571)	(0.0448)	(0.0244)	(0.0344)	(0.0484)	(0.0360)
Observations	4613	5353	4865	4658	5192	5355	5355	4865	5301
R^2	0.108	0.106	0.094	0.482	0.506	0.734	0.631	0.431	0.459

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

with much statistical precision. The reduction in the gender wage gap is likely a consequence of both an increase in female wages and a decrease in male wages, but our data preclude us from identifying those changes in wages precisely. To put it differently, the increase in female wages and the decrease in male wages on their own may be statistically insignificant, but taken together their changes in opposite directions combine to result in a statistically meaningful reduction in the gender wage gap.

Before examining selection and the residual wage gap, we examine whether there are sectoral differences in the behavior of the wage gap and wages themselves. To that end, we take advantage of the information on the sector in which employed individuals were working and classify them as either manufacturing, services, or other in Table 1. We then estimate for each of the three sectors the wage gap as well as female and male wages and collect results in Table 5. [9]

The reduction in the wage gap in the wake of China attaining permanent normal trade relations seems to be largely driven by the services sector which experienced a large and significant reduction in the gender gap, while the manufacturing sector experienced a widening gender wage gap. Unfortunately, similar to our wage regressions pooling across all sectors, our results for changes in wages in each sector are not precise enough, preventing us from drawing strong conclusions. Changes for male and female workers alone are not precisely estimated, but taken together they indicate a reduction in the gender gap in the services sector and an increase in the gender gap in the manufacturing sectors. Our estimates suggest that while both male and female wages decline in manufacturing, female wages decline more. That wages in manufacturing declined is not a surprise given the Pierce and Schott (2016) finding of a reduction in manufacturing employment.

In order to consider the role of selection and entry/exit dynamics in our wage outcomes, we estimate specification (1) with samples restricted as Machado (2017) who proposed an estimator for the wage gap that allows for arbitrary and unobserved heterogeneity in selection. Using the National Longitudinal Survey of Youth 1979, she found that the "always employed" female subpopulation has similar characteristics as the male subpopulation in terms of labor market experience and cognitive tests. Following her methodology of restricting CPS samples, we restrict our sample to individuals who worked 50 weeks per year and 35 hours per week in the previous year to perform an apple-to-apple comparison and re-estimate coefficients regarding the gender wage gaps in Table 4, collecting results in columns 5 and 6 of Table 4. The estimated DID coefficient for the sparse specification is no longer significant, while the estimated coefficient for the specification with additional controls is marginally statistically significant and larger in magnitude, suggesting that even after controlling for sample selection issues, the gender wage gap has decreased in the wake of China receiving permanent normal trade relations status. The reduction in the precision of our estimates could be due to the selection effect, in that some of the reduction is due to changes in characteristics of individuals in the labor market, not by the changes in performance of "always employed" workers. We will return to this point later.

We next examine the residual wage gap between males and females. The simple wage gap analyzed above may mask the effect of having different occupations, education, tenure, etc., on the wage gap. Following Lemieux (2006), we define residual wage as the residual term from a regression of individual worker's wage on his or her observed characteristics. To obtain the residual wage we regress wages using individual level data on the following variables: age, sex, marriage status, veteran status, occupation, level of education, industry of employment, and MSA where the individual resides. We then use

[9] To conserve space for the remainder of the paper we present results only from the specification which includes additional explanatory variables. Complete results are available on request from authors.

Table 6
Residual wage gender gap.

Post * NTR Gap * Male	0.185***
	(0.064)
Post * NTR Gap	−0.098
	(0.087)
Post * Male	−0.014***
	(0.005)
NTR rate	−0.001
	(0.145)
MFA rate	−15.37
	(989.8)
Post * Chinese tariff	−0.000
	(0.006)
Post * No College	0.000
	(0.014)
Post * Veteran	−0.000
	(0.031)
Post * Median HHI	−0.001
	(0.137)
Observations	1151,394
R^2	0.000

Standard errors clustered on MSA × Sex × Year in parentheses, MSA, year, and sex fixed effects.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

the obtained residuals to estimate a variant of the specification (1) at the individual level. The dependent variable is the residual wage of a worker and along with the usual control variables, we include a dummy variable identifying males in the post-PNTR period, as well as an interaction of the post-PNTR dummy, the NTR gap, and the male dummy, effectively estimating a triple difference specification.[10] Our results are collected in Table 6.

Our results indicate that the residual gender wage gap has increased. Taken together, our results indicate that while the relative female wage has increased, the residual female wage has decreased. The implication is that the relative female wage may have increased due to the China PNTR shock generating selection effects, which would be consistent with how controlling for selection effects affected our simple wage gap regression results. It is possible that the relative female wage has increased because more educated women entered the labor force, or similarly, that more educate men left the labor force. The residual wage regression results, which control for the level of education, indicate that the quality of the male labor force has increased, as long as we assume that wages are correlated with labor productivity. Thus, the shock created by China being granted PNTR status may have had a more negative effect on the male labor market, by increasing import competition in sectors where male participation is relatively higher such as manufacturing, inducing more women to enter the labor force to compensate for the reduction in family income due to men leaving the labor force. However, there was selection effect in the female subpopulation with more educated women more likely to enter the labor force. To better understand these effects, we now turn our attention to labor force participation effects.

3.3. Estimates of the gender labor force participation gap

We begin our investigation of the gender gap in labor force participation rates (LFPR) by estimating specification (1) using the LFPR gap as the dependent variable and collecting our result in the first two columns of Table 7. The last two columns estimate specification (1) separately for female and male labor force participation rates. Our estimates for changes in labor force participation rates are much more precise. The gender gap in labor force participation rates has declined in the more exposed MSAs. This reduction is driven by both a statistically significant decline in male and a statistically significant increase in female labor force participation rates as reported by results in Table 7. Such changes are potentially indicative of female workers replacing male workers in the labor force or may be a consequence of a structural change in available jobs skewed in favor of women. While the latter change is beyond the scope of our paper, later in the paper we examine whether this shock precipitated women replacing men in the labor force within households.

In the previous subsection we noted that results from the residual wage regression could be explained by the negative shock of China receiving PNTR status may have induced not only greater labor force participation on the part of women, but a particular pattern in selection, namely that it was the more educated women who tended to enter the labor force. We now

[10] Note that as with all out regressions we estimate MSA and year fixed effects as well as a sex fixed effect. The addition of the sex fixed effect is only so that this specification is econometrically a true difference-in-differences-in-differences specification. Since the wage regression we estimate to obtain residual wages includes sex as an explanatory variable, it usually would not be used in the residual wage regression. Not surprisingly, its inclusion does not affect our estimates in a material way. It is possible to group occupations and we examined one such approach grouping 'Managerial and Professional Specialty' and 'Technical, Sales, and Administrative Support' occupations. Similar to sectoral regressions, we found no significant results. Out of concern for space, these results are not reported and are available on request.

Table 7
Gender gap and labor force participation rates.

	l^F_{mt}/l^M_{mt}		l^F_{mt}	l^M_{mt}
Post * NTR Gap	0.424***	0.671***	0.395**	−0.206**
	(0.139)	(0.200)	(0.166)	(0.0973)
NTR rate		1.020***	0.751***	−0.154
		(0.364)	(0.263)	(0.212)
MFA rate		−2552	1498	4206**
		(2124)	(1700)	(1806)
Post * Chinese tariff		0.156	0.338	0.233
		(0.326)	(0.289)	(0.171)
Post * No College		−0.000118	0.0532	0.0345*
		(0.0361)	(0.0372)	(0.0189)
Post * Veteran		−0.0184	−0.0381	−0.0624
		(0.0747)	(0.0616)	(0.0413)
Post * Median HHI		−0.0152	0.0311*	0.0323***
		(0.0146)	(0.0159)	(0.00800)
Observations	5400	5400	5401	5400
R^2	0.308	0.310	0.433	0.273

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 8
Labor force participation gap and labor force participation by sex and education.

	l^F_{mt}/l^M_{mt}		l^F_{mt}		l^M_{mt}	
	Less educated	More educated	Less educated	More educated	Less educated	More educated
Post * NTR Gap	0.497	0.701**	0.162	0.388**	−0.346*	−0.204
	(0.379)	(0.279)	(0.233)	(0.193)	(0.182)	(0.130)
NTR rate	0.045	1.122**	0.097	0.867**	−0.377	−0.196
	(0.954)	(0.459)	(0.567)	(0.362)	(0.413)	(0.285)
MFA rate	−6183*	−3466	1477	−798.8	6311***	2022
	(3737)	(3254)	(2350)	(2193)	(2153)	(2007)
Post * ΔChinese tariff	0.370	0.125	0.312	0.239	0.115	0.196
	(0.639)	(0.450)	(0.427)	(0.358)	(0.360)	(0.254)
Post * No College	0.038	0.010	0.084*	0.078*	−0.001	0.085***
	(0.063)	(0.046)	(0.046)	(0.042)	(0.030)	(0.031)
Post * Veteran	−0.004	0.091	−0.106	0.034	−0.087	−0.004
	(0.129)	(0.107)	(0.089)	(0.072)	(0.061)	(0.067)
Post * Median HHI	0.026	−0.042**	0.035*	0.029	0.005	0.064***
	(0.025)	(0.019)	(0.019)	(0.018)	(0.012)	(0.012)
Observations	5394	5398	5397	5401	5396	5399
R^2	0.174	0.143	0.326	0.213	0.215	0.180

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

examine whether trade liberalization had a different effect on individuals with different levels of education. We estimate specification (1) with labor force participation rates of different education groups and report the estimated coefficients in Table 8. [11] We separate our sample by education into less and more educated, where the less educated are those with no college education and more educated are those individuals with at least some college education. The reduction in the gender gap in labor force participation rates identified in pooled results in Table 7 is largely driven by the increased participation of the more educated women, while the estimate for less educated women is also positive but imprecisely estimated. Labor force participation rates of all men decreases, while only that of less educated men is precisely estimated.

One interpretation of these changes is that the negative income shock experienced by families may have induced some of the more educated women who chose to be stay-at-home parents to re-enter the labor force. This is related to Hakobyan and McLaren's (2017) finding that married high-skilled females drop out from labor market with respect to worsened labor market conditions due to NAFTA tariff reductions. Our results indicate that the effects of NAFTA tariff reductions and granting of PNTR status to China have at least some different effects and highlight that high-skilled females are the demographic group whose labor supply is relatively elastic with respect to negative income shocks.

With respect to the observed sectoral results, as we discussed in Section 2.1 and seen from Table 1, the service sector tends to employ more women. Hence, the more educated women tend to find jobs in the service sector. By the same logic, many of the less educated men seem to lose their jobs in the manufacturing sector. We conjecture that the less educated

[11] As defined earlier, individuals with less education are those with no college experience while those with more education are those with at least some college education.

Table 9
Intrahousehold work dynamics.

	Working spouses			Share of female income in household income
	Both	Husband only	Wife only	
Post * NTR Gap	0.0455	−0.171*	0.148***	0.228*
	(0.205)	(0.102)	(0.0503)	(0.138)
NTR rate	0.772*	−0.386*	0.000165	0.433*
	(0.421)	(0.205)	(0.141)	(0.256)
MFA rate	3467	−340.1	−1393	−2683*
	(2601)	(1290)	(1030)	(1430)
Post * Chinese tariff	0.0679	−0.241	0.207**	0.165
	(0.356)	(0.177)	(0.0983)	(0.202)
Post * No College	0.00503	−0.0202	0.0177*	0.0155
	(0.0334)	(0.0157)	(0.0102)	(0.0217)
Post * Veteran	−0.115*	−0.00393	0.0616***	0.00584
	(0.0665)	(0.0302)	(0.0195)	(0.0355)
Post * Median HHI	0.0509***	−0.0169***	−0.00854**	−0.00353
	(0.0122)	(0.00630)	(0.00408)	(0.00755)
Observations	5401	5401	5401	0.323***
R^2	0.351	0.316	0.223	(0.0510)

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects.
*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

women in the manufacturing sector are less affected than the less educated men since their roles are different. Pierce and Schott (2016) show that U.S. firms in sectors more exposed to trade liberalization introduced more labor-saving technologies. Less educated men are more affected since their roles may depend more on physical abilities. This is consistent with Juhn et al. (2014) who show that NAFTA tariff reductions raised the relative wage of female workers in blue-collar tasks in Mexico. They explain that higher competition encouraged Mexican firms to modernize their technology and thus reduced their dependence on physical ability of workers.

Changes we observe in how labor force participation rates are affected across different levels of education are consistent with our conjecture based on residual wage regression results. The increase in the observed relative female wage, coupled with higher residual wages for men, are due to the effect of China being granted PNTR status being stronger for the male labor market and inducing more women to join the labor force. A selection effect operated in the female labor market which favored the entry of more educated women. In the next section we investigate whether women entered the labor force in order to substitute the lost income from their partners leaving the labor force, voluntarily or involuntarily.

3.3. Intra-household adjustments

China receiving PNTR status resulted in more women joining the labor force, while men became discouraged and left the labor force or left the more exposed MSAs, as shown by Greenland et al. (2019). An interesting question we can ask is how intra-household employment dynamics were affected. If men become discouraged and do not move from their current MSA, the increase in female labor force participation may be indicative of women taking on a larger role of earning an income in households with married or cohabiting couples. To that end, we examine whether there are changes in the fraction of households with both spouses working, or just one spouse, either husband or wife working, estimating specification (1) with the respective ratios as dependent variables. We restrict the sample to only those households with married or cohabiting couples. Our results are collected in Table 9.

From the first column, we can see that there does not seem to be a significant change in the fraction of households with both spouses working. At the same time, there is a significant reduction in the fraction of households with only the husband working and a significant increase in the fraction of households with only the wife working. These results should not necessarily be surprising. The reduction in labor force participation on the part of men as they become discouraged creates the need within households for women to play an increasingly important role in income generation. As a result, women enter the labor force in growing numbers to make up for the shortfall created by men leaving the labor force.

To further examine whether such intra-household adjustments took place we examine changes in the share of income earned by women in married or cohabiting households. We estimate specification (1) with all control variables and use the share of female income in household income as the dependent variable. Results are shown in the last column of Table 9. As we can see from the table, in MSAs that were more exposed to the effects of China receiving PNTR status, the female share of household income did increase.

Our results contrast with Keller and Utar (2019). They show that lower labor market opportunity due to the growing Chinese import competition raised gender inequality in Denmark by inducing female workers to stay away from the labor market (more parental leave, higher fertility, more marriage, and fewer divorces). The different outcomes may be explained by Denmark's better social safety net or the more flexible labor market in the U.S. If the latter is the main reason, the change in female labor force participation may not depend on marital status. Table 10 reports the impact of PNTR with

Table 10
Labor force participation rate by marital status.

	Single households		Married or cohabiting household	
	Male LFPR	Female LFPR	Male LFPR	Female LFPR
Post * NTR Gap	−0.335*	0.207	−0.152	0.458**
	(0.183)	(0.237)	(0.107)	(0.222)
NTR rate	−0.696	−0.110	0.0933	1.276***
	(0.434)	(0.456)	(0.260)	(0.399)
MFA rate	4969*	2655	2525	1014
	(2668)	(2921)	(1890)	(2309)
Post * ΔChinese tariff	0.172	−0.0119	0.224	0.476
	(0.348)	(0.397)	(0.201)	(0.408)
Post * No College	0.0430	−0.00268	0.0350*	0.0804*
	(0.0377)	(0.0380)	(0.0197)	(0.0444)
Post * Veteran	−0.154**	−0.0245	−0.0409	−0.0432
	(0.0780)	(0.0766)	(0.0466)	(0.0752)
Post * Median HHI	0.0231*	0.0250*	0.0321***	0.0437**
	(0.0131)	(0.0150)	(0.00854)	(0.0191)
Observations	5386	5395	5400	5401
R^2	0.179	0.234	0.215	0.359

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 11
Unemployment and employment rates.

	Unemployment rate		Employment rate	
	Female	Male	Female	Male
Post * NTR Gap	0.231***	0.168*	0.197	−0.347**
	(0.0742)	(0.0997)	(0.180)	(0.144)
NTR rate	−0.134	−0.210	0.780***	0.0665
	(0.262)	(0.260)	(0.291)	(0.300)
MFA rate	−1536	−1798	2360	5527***
	(1034)	(1143)	(1601)	(1903)
Post * ΔChinese tariff	0.143	0.0432	0.222	0.162
	(0.145)	(0.189)	(0.312)	(0.249)
Post * No College	0.00373	0.0159	0.0475	0.0184
	(0.0157)	(0.0162)	(0.0353)	(0.0233)
Post * Veteran	−0.0208	0.0430	−0.0172	−0.0948**
	(0.0299)	(0.0317)	(0.0570)	(0.0453)
Post * Median HHI	−0.0266***	−0.0171**	0.0504***	0.0455***
	(0.00706)	(0.00711)	(0.0149)	(0.0100)
Observations	5400	5400	5401	5400
R^2	0.259	0.320	0.458	0.368

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

China on labor force participation rate changes in our sample for males and females with different marital status. This table shows that labor force participation rate of females in married or cohabiting households significantly increased while those in single households were not significantly affected, supporting the claim that women entered the labor force to replace the lost income from their male partners' leaving the labor force. While the effect on married male labor force participation rates is not statistically significant, it is estimated with a negative sign, which is in line with the substitution hypothesis.

3.4. Other changes in local labor markets

Our approach precludes us from drawing welfare implications due to both modeling and data limitations. However, we can take advantage of additional data to better understand the driving forces behind the increased labor force participation of women and decreased participation of men. Labor force participation numbers reflect both individuals who are working and those who are not working but are actively seeking employment. Thus, increases in labor force participation could come from the ranks of the officially unemployed: those actively seeking employment. The other possibility is that even if the increase in labor force participation comes from more individuals working, it is not clear whether the increase comes from individuals holding full- or part-time jobs. Granting China PNTR status may have resulted in a redistribution of jobs from full- to part-time employment, which would be indicative of structural changes in local labor markets which were more exposed to the shock.

We begin by examining how unemployment and employment for both female and male workers have changed because of China being granted permanent normal trade relations with the U.S. Our results, collected in Table 11, show that unemployment rates for both female and male workers rose. The last two columns of Table 11 indicate that the employment rate

Table 12
Total hours worked and part-time work.

	Total hours worked (log value)		Weeks in part-time work (log value)	
	Female	Male	Female	Male
Post * NTR Gap	−0.188	−0.409***	2.597*	2.440
	(0.172)	(0.148)	(1.333)	(1.747)
NTR rate	0.470	−0.286	2.076	0.607
	(0.393)	(0.296)	(1.848)	(3.624)
MFA rate	806.9	3017*	−2653	−26,460
	(1630)	(1739)	(17,361)	(21,831)
Post * ΔChinese tariff	0.0438	0.00605	0.993	−0.368
	(0.310)	(0.248)	(2.037)	(3.046)
Post * No College	0.0320	0.00845	−0.191	0.0892
	(0.0272)	(0.0225)	(0.188)	(0.318)
Post * Veteran	−0.0336	0.0675	−0.532	0.700
	(0.0477)	(0.0499)	(0.393)	(0.632)
Post * Median HHI	0.00687	0.0207**	−0.0247	0.0697
	(0.0107)	(0.00914)	(0.0719)	(0.107)
Observations	5358	5357	5348	5249
R^2	0.299	0.296	0.275	0.175

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

of men decreased in MSAs with greater exposure while the female employment rate was not significantly affected. Taken together, these changes imply that much of the increase in the female labor force participation comes from more women joining the labor force and searching for jobs. Since our data do not allow us to track individuals across time, we cannot decompose net changes in labor force participation into its separate constituents. In other words, understanding how many women who join the labor force due to China receiving PNTR status are unable to find jobs, but keep looking for one, is beyond our scope. For men, the conclusion is clear cut. Male unemployment rate increases, while their employment rate decreases. Both changes have likely resulted in some men becoming discouraged and dropping out of the labor force or moving out of the MSA, the latter being consistent with the findings of Greenland et al. (2019).

To examine the kinds of jobs individuals in our sample hold, we take advantage of two additional variables, total hours worked and weeks spent in part-time employment. The former reflects the number of hours spent working in both full- and part-time jobs in a calendar year, while the latter reflects the number of weeks individuals spent in part-time employment, defined as working less than 35 hours a week.

We estimate specification (1) with total hours worked and weeks in part-time employment as dependent variables collecting our results in Table 12. While our results indicate that both men and women spend less time working after China is granted PNTR, only for men is the coefficient estimated precisely. Similarly, while our results indicate both men and women spent more time in part-time employment, only the coefficient for women is estimated precisely. Taken together, these results indicate that both men and women experience a change in the fundamental nature of jobs they held, facing limited opportunities for full-time work and greater reliance on part-time work.

To better understand whether the nature of jobs held changed because individuals started preferring part- to full-time employment, we take advantage of our data asking respondents to identify the reasons behind their working a part-time job focusing on two of them: whether they reported being unable to find a full-time job and whether they reported preferring a part-time job. We then estimate specification (1) with the proportion of individuals who could not find full-time employment and then the proportion who wanted part-time employment,[12] separately for women and men, collecting our results in Table 13. During our sample, there is no significant change in the proportion of either women or men who want part-time work. However, granting China PNTR status is associated with a large and statistically significant increase in the proportion of both women and men who state that they were unable to find full-time employment and settled for part-time work instead. Thus, the change in the nature of jobs occurred due to labor market conditions and increased exposure to import competition from China, rather than due to changes in preferences.

4. Robustness

We conduct several robustness exercises. The first one replicates the two exogeneity robustness checks of Pierce and Schott (2016). The use of NTR gap to identify the effect of granting China PNTR status relies on the exogeneity of the NTR gap. As Pierce and Schott (2016) argue, most of the variation in NTR gap is due to the variation in non-NTR tariff rates which were set in 1930 and any increase in NTR to protect industries would result in a smaller NTR gap. They perform two checks of exogeneity of NTR gap. One was to instrument the baseline DID term, $Post\ PNTR_t \times NTR\ Gap_m$, with an interaction of the post-PNTR indicator and the Smoot–Hawley-based non-NTR tariffs rates, $Post\ PNTR_t \times non - NTR\ Tariff_j$. The second one was to re-estimate the baseline specification using the NTR gap observed in 1990, ten years prior to PNTR implementation.

[12] The denominator for both measures is the number of individuals in each gender group who held a part-time job in the previous year.

Table 13
Reasons for part-time work.

	Cannot find a full-time job		Wanted a part-time job	
	Female	Male	Female	Male
Post * NTR Gap	0.573**	0.537*	−0.663	0.270
	(0.251)	(0.310)	(0.464)	(0.389)
NTR rate	−0.328	0.135	0.479	0.168
	(0.449)	(0.655)	(0.985)	(0.833)
MFA rate	602.4	−3472	4242	10,070
	(3375)	(3294)	(5308)	(6567)
Post * ΔChinese tariff	0.319	−0.576	−1.895**	0.463
	(0.424)	(0.551)	(0.768)	(0.647)
Post * No College	0.065	−0.005	−0.026	0.031
	(0.041)	(0.051)	(0.071)	(0.074)
Post * Veteran	−0.129	0.047	0.390***	−0.072
	(0.080)	(0.110)	(0.118)	(0.124)
Post * Median HHI	−0.020	0.001	0.025	−0.004
	(0.017)	(0.022)	(0.024)	(0.026)
Observations	5348	5249	5348	5249
R^2	0.177	0.125	0.235	0.147

Standard errors clustered on MSAs in parentheses, MSA and year fixed effects. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 14
Robustness checks.

	Benchmark	2000 NTR Gap	NNTR IV	Additional variables
w^F_{mt}/w^M_{mt}	0.678**	0.595*	0.596**	2.192*
$\log(w^F_{mt})$	0.597*	0.582	0.579**	1.192*
$\log(w^M_{mt})$	−0.206	−0.115	−0.117	−0.145
l^F_{mt}/l^M_{mt}	0.671***	0.617***	0.629***	0.673***
l^F_{mt}	0.395**	0.352**	0.356***	0.404**
l^M_{mt}	−0.206**	−0.207**	−0.214**	−0.180*

Standard errors clustered on MSAs in parentheses, MSA and year, fixed effects. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

The second robustness exercise adds several additional explanatory variables to our benchmark regressions. We interact the share of foreign-born individuals with the Post PNTR indicator since a larger immigrant population may be associated with higher labor force participation among higher-skilled workers (Peri, 2014). We control for the MSA's share of employment in routine occupations given that areas exposed to import competition are also more susceptible to worker displacement by automation. We add the share of employment in manufacturing in an MSA interacted with a Post PNTR indicator as manufacturing jobs are at greatest risk according to Pierce and Schott (2016). Finally, we add the average offshorability index to control for the exposure to offshorability of jobs for an MSA (Autor and Dorn, 2013).

In the interest of space, we present only the estimates of the DID term, $Post\ PNTR_t \times NTR\ Gap_m$, in regressions involving the wage and labor-force-participation gaps, as well as for gender-specific wage and labor force participation rates in Table 14. [13] Across all three robustness checks, our results are qualitatively similar. Robustness results for wage related regressions are the least precisely estimated. In each regression, the wage gap decreases in the most exposed MSAs, while the male wage decreases and the female wage increases.

In all three robustness checks the gender labor-force-participation gap decreases in the most exposed MSAs, with similar magnitudes across all regressions and is precisely estimated. Female labor force participation in the more exposed MSAs is also precisely estimated to increase by a similar magnitude across all regressions. Male labor force participation is decreasing by a similar magnitude in all regressions and is precisely estimated. Thus, our robustness results confirm our main conclusions: the gender wage gap and the gender labor-force-participation gap both decrease in the wake of China receiving PNTR status, with the labor force participation results more precisely estimated.

The last robustness check we perform is to examine to what extent our results are affected by the Great Recession which started in 2008. To that end we add a new variable to our regression specification given by Eq. (1) which is the interaction of the $Post\ PNTR_t \times NTR\ Gap_m$ variable and a dummy variable identifying the 2008 through 2013 period. In Table 15 we only present the estimated coefficients for the $Post\ PNTR_t \times NTR\ Gap_m$ term and the new triple interaction term to identify what effect the Great Recession had for the four key regressions, simple wage gap in the full and Machado sample, residual wage gap, and labor force participation gap.[14] As can be seen, the simple wage gender gap decreased prior to the Great Recession

[13] Complete results are available on request.

[14] The residual wage regression is the exception as the dependent variable of interest here is the male wage in the post-2001 period interacted with NTR Gap, the triple difference term. Therefore, the additional independent variable is the interaction between that variable and a dummy identifying the 2008–2013 period.

Table 15
Wage and labor force participation rate gap and the Great Recession.

	Full sample	Machado sample	Residual wage†	LFPR
Post * PNTR Gap	0.671*	0.154	0.189***	0.480**
	(0.379)	(0.933)	(0.066)	(0.192)
Post * PNTR Gap * (2008–2013 Dummy)	0.114	3.111**	−0.007	0.445**
	(0.394)	(1.402)	(0.032)	(0.225)

†The independent variable reported for the residual wage regression are $Post\ PNTR_t \times NTR\ Gap_m \times$ Male Dummy and $Post\ PNTR_t \times NTR\ Gap_m \times$ Male Dummy $\times$ (2008 – –2013 Dummy).
Standard errors clustered on MSAs in parentheses, MSA and year fixed effects.
*** $p<0.01$, ** $p<0.05$, * $p<0.1$.

which had no statistically significant effect on it. The results for the Machado sample indicate that there is no change in the wage gap prior to the Great Recession, but the Great Recession had a large effect on the wage gap decreasing it in the post-2008 period. The residual wage regression indicates that the residual wage gap increased prior to the Great Recession, which had no additional effect on the residual wage gap. The results for the simple wage gap in the Machado sample taken together with the residual wage regression results indicate that there was important selection effect in the post-PNTR period and prior to the Great Recession as discussed earlier in the paper. Finally, in terms of the labor force participation gender gap, the gap decreased prior to the Great Recession, which decreased it further by roughly a similar amount.

5. Conclusion

China being granted PNTR status caused several important changes in local labor markets. Gender wage and labor-force-participation gaps both declined. We show that the simple gender wage gap decreased, but that there were also important selection effects in the wake of the PNTR change. Our residual wage regression shows that the male residual wage increased which is indicative of increased quality of the male labor force. Our results further show that less educated men left the labor force in greater numbers and that more educated women entered the labor force in greater numbers, both of which help explain the residual wage gap result. As a result of both of these forces, the gender gap in labor force participation decreased in the most exposed MSAs due to more women joining the labor force, while men became discouraged and left the labor force. These changes were accompanied by increased unemployment among women and men, while the employment rate of women did not change and that of men decreased. Concurrently, both women and men spent less overall time working. While the reduction in time spent working for women is imprecisely estimated, women did spend more time in part-time employment indicating a reduction in full-time work. Our results indicate both women and men spent more time in part-time work because they were unable to find full-time employment, rather than because they preferred part-time employment.

This episode induced changes in intrahousehold work dynamics. While there were no changes in the number of households where both spouses worked, in the more exposed MSAs there was a decrease in the households with just the husband working and an increase in the households with only the wife working. In addition, the share of family income accounted for by the female partner increased in the more exposed MSAs. We show that labor force participation rates increased among married women, while those of single women did not change. Thus, our results indicate that women entered the labor force to replace the lost family income due to husbands leaving the labor force or becoming unemployed.

Lower gender gaps in labor markets are often interpreted as female welfare improvement. However, if a negative shock in the labor market affected male workers more severely or forced more females to work as shown by this paper, this conclusion would be too hasty. While it is tempting to use our results to form conclusions about welfare implications, our approach and data limitations preclude us from doing so and are left for future work.

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