Trade Liberalization and Female Welfare: Dimensions of Adjustment in the United States

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Abstract

This paper provides empirical results that trade liberalization with China lowered gender gaps in the US labor market. However, the lower gender gaps in wage and labor participation are driven by underperformance of male workers, rather than the improvement of female ones. In MSAs with higher exposure to trade liberalization, male workers' wages and labor participation rates significantly decreased while female workers' outcomes were not affected. We also show that trade liberalization increased reliance of U.S. female workers on part-time jobs, while working fewer hours in the more affected MSAs. Hence, the lower gender gaps may not indicate female welfare gains from trade. Moreover, our results suggest that trade liberalization may provoke another type of discrimination. The impacts from trade liberalization were more severe for African-American female workers than for White-Americans.

JEL: F16, F66, J15, J16, J70

Key words: Trade liberalization, Gender wage gap, Gender gap in labor participation

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1. Introduction

This paper considers the impact of trade liberalization on gender inequality in the labor market. A number of previous papers show that episodes of trade liberalizations tend to be associated with lower gender gaps in wage and employment. Aguayo-Tellez et al. (2013) and Juhn et al. (2014) show that the formation of the North American Free Trade Agreement (NAFTA) improved the relative wage and employment of Mexican female workers. Ederington et al. (2009) report that tariff reductions in Chile, as the results of Chile's entry into GATT/WTO, raised the number of female workers (relative to male ones) in blue-collar jobs. Black and Brained (2004) and Brussevich (2018) show that import penetration is associated with greater reductions in the gender wage gap in the U.S. We also find empirical evidence that trade liberalization reduces the gender wage and labor participation gaps. However, we show that these reductions do not necessarily indicate female welfare gains as they are driven by underperformance of men. Furthermore, it can provoke another type of discrimination for racial minority groups, as our results indicate that African American female workers were more severely affected.

We use the Pierce and Schott (2016, 2019) approach to measure a local labor market's exposure to trade liberalization. Pierce and Schott (2016) show that granting Permanent Normal Trade Relations (PNTR) 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. 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 and U.S. firms to do more offshoring/outsourcing. In a follow-up paper, Pierce and Schott (2019) calculate the exposure level to PNTR for each US 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 to reflect local labor markets. We find that a Metropolitan Statistical Area (MSA) with higher PNTR exposure shows decreased gender gaps in wage and labor force participation rate (LFPR) after trade liberalization with China. In addition, we find that MSAs with higher PNTR exposure show larger increases in female relative wages and labor force participation rate. However, our estimates reveal that these reductions in gender gaps are mainly driven by the underperformance of male workers, rather than the improvement of female labor market outcomes. Our results show no significant changes in female wages and LFPR with respect to a MSA's exposure to PNTR, while male wages and LFPR significantly decreased. Furthermore, in MSAs with higher PNTR, female overall work hours tend to decrease, but their part-time work increased. These changes are indicative of reduced welfare for female workers, despite what one would usually associate with reductions in labor-market-related gaps. Firstly, female workers reduced the number of hours they spend working. Secondly, the increase in part-time hours indicate an even greater reduction in hours worked in full-time jobs. As is well known, part-time jobs do not come with benefits. Hence, trade liberalization with China has reduced female welfare by reducing overall hours worked as well as hours spent in full-time employment.

We also show that the increased competition in labor market from trade liberalization could have deepened racial inequalities. By comparing African-American and White-American female workers, we show that decreases in female work hours are driven by African-American women. The increase in part-time work with respect to PNTR exposure was larger for African-American women as well.

Other studies have found significant effects of trade liberalization on labor market outcomes and gender inequality using different methodologies. Autor, Dorn, and Hanson (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 adapted by a number of follow-up papers. Following Autor, Dorn, and Hanson (2013)'s identification strategy, Brussevich (2018) shows that U.S. commuting zones with higher import penetration show greater reduction in the gender wage gap and shows that wage and welfare gains from trade are higher for females since the import competition shock in manufacturing sector disproportionally affected the labor market outcomes of two gender groups. Benguria and Ederington (2017) show that increased competition from China lowered the gender wage gap in Brazil and show that this lower gender inequality is driven by the underperformance of male workers as we propose with U.S. data. Aguayo-Tellez et al. (2013) show that tariff reductions, accompanied by 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 females especially in blue collar tasks in Mexico. They explain that higher competition encouraged firms to modernize their technology, and thus their physical ability dependence became lower.

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, a 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 greater commitment from employees. Hence, female workers, who tend to have less flexible schedules, receive lower relative wages.

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. Our paper provides additional insights into the effects of trade liberalization in developed economies, and in particular, shows that with respect to the effect on gender gaps, trade liberalization has had similar effects in developed and developing countries. The latter conclusion is based 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. Unlike previous efforts, we also show that trade liberalization has had a differential impact on racial groups, with African-American women being affected more negatively than White-American women.

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 the 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 active duty military or disability.[1] Our sample in 2001 includes 104,962 individuals who resided in 272 MSAs with trade liberalization exposure data. Table 1 shows the demographics of our sample. Females account for 52 percent of the sample. In terms of racial distribution, White Americans account for 82 percent of the population while African-Americans represent 11 percent.

Table 1: CPS sample composition in 2001

	Female	Male	Total
By race			
White	42.51%	39.98%	82.49%
Black	6.10%	4.66%	10.76%
Other	3.53%	3.22%	6.75%
By employment			
No employment	11.23%	2.99%	14.22%
Manufacturing	4.28%	8.61%	12.89%
Nonmanufacturing	36.63%	36.25%	72.88%
Total	52.14%	47.86%	100.00%

N. of obs. 104,962

Our variables of interest include the average hourly wage, w_{it}, the number of total hours worked, h_{it}, and labor participation status, l_{it}, of 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." The hourly wage, w_{it}, is calculated as "Wage and salary income" divided by the number of hours, 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 labor force if she or he worked, was looking for a job, or was temporarily absent/laid-off during the reference period.

[1]For variables that require only those who are working, we further restrict the sample to include workers with stronger attachment to the labor market - worked for more than 20 weeks in the previous year, and more than 35 hours per week in the previous year (both inclusive).

Table 2 summarizes these variables. In 1990, the average hourly wage for the whole population was \$12.26 with the average female hourly wage was \$10.41, equivalent to 74 percent of the average male hourly wage. In 2010, the female worker's average wage increased to about 76 percent of the male worker's.[2] We observe similar patterns for other variables. In 1990 the labor force participation rate (LFPR) for females was 22 percentage points lower than for males. In 2010, the gap decreased to about 14 percentage points. In 1990 female workers' average work hours was equivalent to 83 percent of male average work hours. It increased to 87 percent of males' hours worked by 2010. Our data reveal the well-known patterns: while male workers tend to outperform female workers, gender haps have been on the decline.

Table 2: Summary Statistics on CPS variables (mean values)

	1990	2010	1990	2010	1990	2010
	All		Female		Male	
LFPR	0.779	0.858	0.693	0.773	0.890	0.904
	(0.111)	(0.093)	(0.101)	(0.077)	(0.056)	(0.060)
Wage	12.259	23.673	10.409	20.587	14.052	26.954
	(1.832)	(4.473)	(1.799)	(4.551)	(2.402)	(6.554)
Work Hours	1924	1911.58	1709.329	1792.516	2069.576	2071.078
	(96.696)	(106.648)	(146.988)	(141.480)	(122.382)	(130.838)

Standard errors in parentheses.

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

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

where m indicates an MSA, S refers to demographic groups, $N_{m,t}^{S}$ is the number of individuals who resided in MSA m in year t, and $\widehat{N}_{mt}^{S}$ is the number of individuals with 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.[3] Figure 1 summarizes regional differences of changes in female relative wages w_{mt}^{F}/w_{mt}^{M}.

[2]The median female wage increased form 74 percent of the male wage in 1990 to 81 percent in 2010.

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

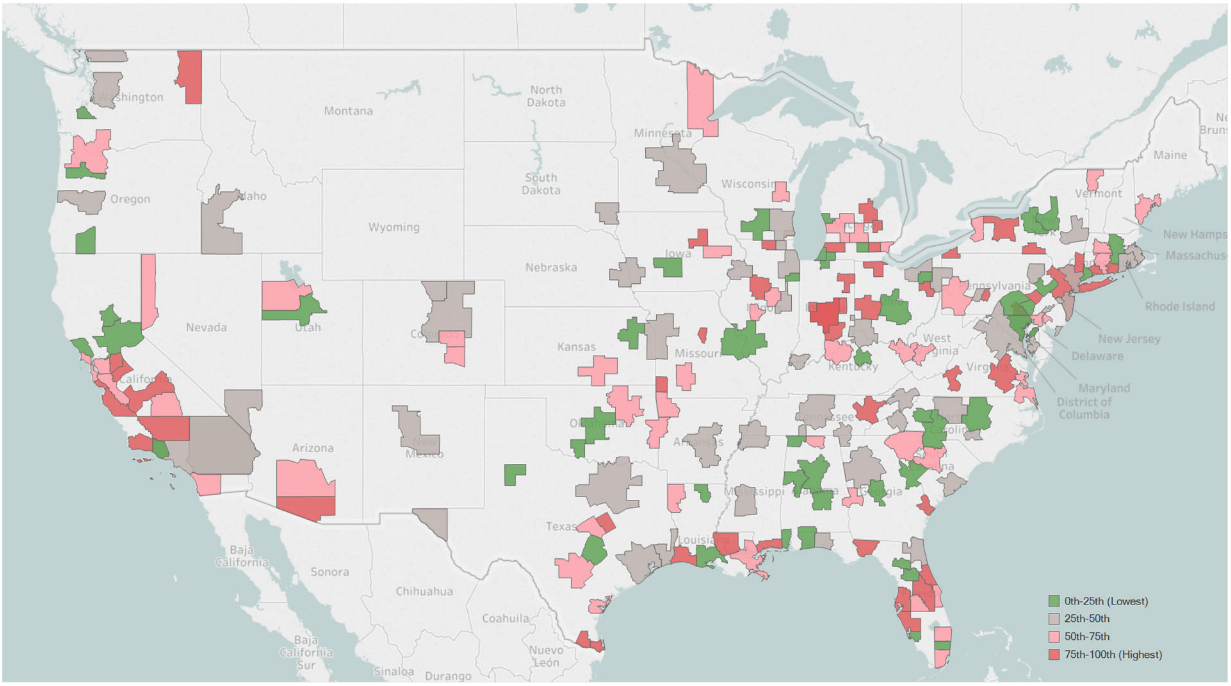

Figure 1: Female Relative Wage Changes

Note that our analysis uses MSA-level variables while Pierce and Schott (2017) use county-level variables. As we are interested in labor market outcomes, we conducted our analysis at the MSA level since they are defined by boundaries of local labor markets. Using Commuting Zones (CZs), which include MSAs, may be another option, but not a feasible one since CPS does not include identifier of county or Public Use Microdata Area (PUMA) over the sample period 1990-2013 in this study.[4]

2.2. NTR Gap

Our measure of exposure of an MSA to trade liberalization follows Pierce and Schott (2019). 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 (NTR) with the U.S. will be assessed tariff rates established by the Smoot-Hawley Tariff Act of 1930. These rates are significantly higher than the NTR tariffs rates which are assessed on

[4]CPS's county identifier is only available since 1996, and it does not provide PUMA as American Community Survey does.

imports from countries who are members of the World Trade Organization (WTO). China was first granted temporary NTR tariffs in 1980. China's temporary access to NTR rates was reaffirmed on an annual basis. As Pierce and Schott (2019) discuss the uncertainty associated with the renewal ebbed and flowed based on flashpoints 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 (2019)'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 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. Pierce and Schott (2016) show that the elimination of trade uncertainty explains the sharp drops in US manufacturing employment.

Using $NTR\ Gap_j$, Pierce and Schott (2019) 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 a 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 highest and lowest $NTR\ Gap_m$ while Figure 2 illustrates the geographical distribution of $NTR\ Gap_m$. MSAs in Midwest and East Coast are more exposed to trade liberalization.

Table 3: MSAs with highest and lowest NTR Gaps

Rank	Metropolitan Statistical Area	NTR Gap
1	Hickory-Morganton, NC	0.449
2	Greensboro-Winston Salem, NC	0.274
3	Anderson, IN	0.271
4	Poughkeepsie, NY	0.269
5	Mansfield, OH	0.266
268	Ocean City, NJ	0.057
269	Farmington, NM	0.05
270	Las Vegas, NV	0.049
271	Atlantic City, NJ	0.049
272	Anchorage, AK	0.047

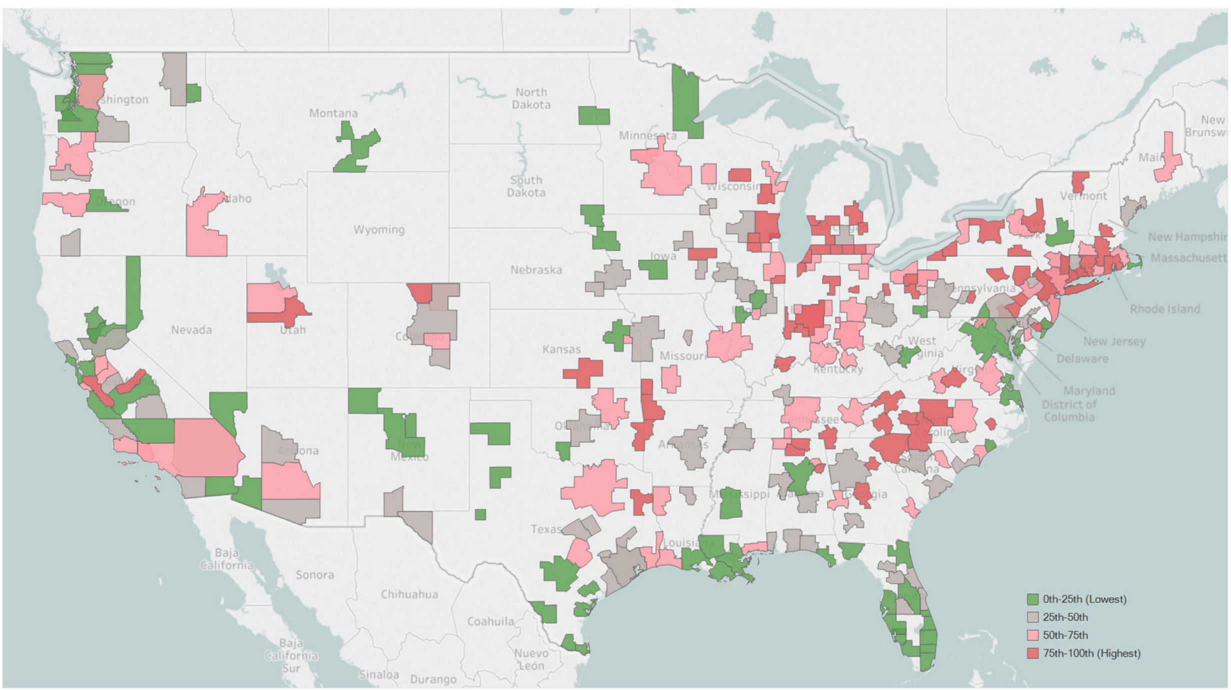

Figure 2: NTR Gaps by MSA

3. Estimation
3.1. DID identification Strategy

Our estimation strategy follows Pierce and Schott (2019). 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),

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

The left-hand-side (LHS) variable, either the gender wage or labor-force-participation-rate gap, is defined in year *t* and 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 rate associated with the industries active in the MSA. $\boldsymbol{Z}_m$ represents the initial-period MSA attributes, 1990 median household income, 1990 share of population without any college education, and 1990 share of population that are veterans. 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 (2019).

3.2. Estimates for Gender Gaps

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^F_{mt}/w^M_{mt}, and labor force participation ratio, l^F_{mt}/l^M_{mt} , as dependent variables and estimate the DID point estimates of interest θ. The first and second columns of Table 4 report the results for the female-male wage ratio. The third and fourth columns provide the estimation results with female-male labor force participation ratio. The first and third columns report coefficient estimates for a specification containing just the DID term of interest and fixed effects. The second and fourth columns controls for policy changes $\boldsymbol{X}_{mt}$ and demographic variables $\boldsymbol{Z}_m$. The DID point estimates of interest are positive ($\theta > 0$) and statistically significant at conventional levels across all columns.[5] Our empirical results suggest that higher import competition from China is associated with lower gender gaps in the U.S. labor market outcomes

[5]Our results are unchanged if we exclude employees in textile production sector from our sample. This shows the robustness of our results to the expiration of the global Multi-Fiber Arrangement (MFA).

as in Brussevich (2018).

Table 4: Relative wage and LFPR

	w_{mt}^F/w_{mt}^M		l_{mt}^F/l_{mt}^M	
Post * NTR Gap	0.149	0.291**	0.173***	0.269***
	(0.0985)	(0.112)	(0.0632)	(0.0681)
NTR		1.463*		1.103***
		(0.858)		(0.362)
Post * Median HHI		-0.0272		-0.0155
		(0.0255)		(0.0157)
Post * No College		-0.0500		-0.00225
		(0.0565)		(0.0323)
Post * Veteran		0.257**		0.0434
		(0.128)		(0.0749)
MSAs	272	272	272	272
Observations	5,540	5,467	5,469	5,469
R2	0.125	0.130	0.304	0.306
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

The main question of this paper is whether the reduction in gender gaps in labor market outcomes had brought welfare gains to female workers in the U.S. As a first step, we examine changes in wages and labor force participation rates for women and men. We estimate specification (1) by using the log value of the wage, $\log(w_{mt}^S)$, and labor force participation rate, l_{mt}^S, for each gender $S \in \{F, M\}$ as dependent variables. The estimated coefficients in Tables 5 and 6 suggest that the lower gender gaps, reported in Table 4, are driven by negative effects on male workers rather than improvement of female workers' outcomes. Table 5 shows that, in MSAs with higher exposure to PNTR, male wages significantly decreased, but female wages did not change in a statistically significant way. Table 6 shows higher labor force participation rates for female workers and lower labor force participation rates for male workers in the wake of China being granted permanent normal trade relations.

Table 5: Wage Changes

	$\log(w_{mt}^F)$		$\log(w_{mt}^M)$	
Post * NTR Gap	-0.0400	0.0905	-0.248*	-0.279*
	(0.102)	(0.114)	(0.131)	(0.142)
NTR		1.146		-0.554
		(0.711)		(0.732)
Post * Median HHI		0.0457*		0.0782***
		(0.0243)		(0.0258)
Post * No College		-0.191***		-0.126**
		(0.0572)		(0.0520)
Post * Veteran		0.277**		-0.00697
		(0.111)		(0.0990)
MSAs	272	272	272	272
Observations	5,540	5,467	5,542	5,469
R2	0.788	0.794	0.747	0.752
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

Table 6: Labor Force Participation Rate Changes

	l_{mt}^F		l_{mt}^M	
Post * NTR Gap	0.0632	0.113**	-0.0940***	-0.132***
	(0.0489)	(0.0504)	(0.0338)	(0.0415)
NTR		0.709***		-0.293
		(0.260)		(0.224)
Post * Median HHI		0.0112		0.0221**
		(0.0120)		(0.00929)
Post * No College		0.0148		0.0208
		(0.0279)		(0.0185)
Post * Veteran		-0.0599		-0.123***
		(0.0550)		(0.0430)
MSAs	272	272	272	272
Observations	5,469	5,469	5,469	5,469
R2	0.434	0.435	0.273	0.277
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

At first blush one might be tempted to conclude that the increase in female labor force participation rates coupled with no changes in their wages are indicative of female welfare gains. Such a conclusion is warranted only under the assumption that labor force participation rate reflects employment in full-time jobs which provide benefits. However, the labor force participation variable reflects employment in both full- and part-time jobs as well as instances where an individual is unemployed but actively seeking employment during the reference period. To get a better understanding of welfare consequences, we examine how many hours females in our sample worked in a year, h_{mt}^{F}, and how many weeks they worked in part time jobs, $\hat{h}_{mt}^{F}$.

Table 7: Hours Worked by Women

	$\log(h_{mt}^{F})$		$\log(\hat{h}_{mt}^{F})$	
Post * NTR Gap	-0.142***	-0.108*	0.780	0.953*
	(0.0510)	(0.0576)	(0.475)	(0.507)
NTR		0.543		1.734
		(0.367)		(1.840)
Post * Median HHI		0.0162		-0.0620
		(0.0108)		(0.0696)
Post * No College		0.0318		-0.185
		(0.0252)		(0.173)
Post * Veteran		-0.0335		-0.209
		(0.0540)		(0.370)
Sample Period	1990-13	1990-13	1990-13	1990-13
MSAs	272	272	272	272
Observations	5,542	5,469	5,531	5,459
R2	0.292	0.293	0.277	0.276
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The first and the second columns of Table 7 report the regression results from estimating specification (1) using $\log(h_{mt}^{F})$ as a dependent variable.[6] In these two columns, the DID point estimates are statistically significant and negative, implying that female workers in MSAs with higher PNTR exposure worked fewer hours. The third and fourth columns show the results for

[6]Our results are robust if we only use the number of work hours per week as the dependent variable.

weeks spent working in a part-time job. Though somewhat weakly estimated, the DID coefficients indicate an increase in part-time work in MSAs with greater PNTR exposure. Taken together, results presented in table 7 show that women spent fewer hours working in a year, but spent more weeks holding part-time jobs. Thus, the increase in female labor force participation is likely due to an increase in part-time employment as well as more women seeking jobs which together mask a reduction in full-time employment. All of these changes are unlikely to have improved the welfare of female workers.

Our results suggest that, in MSAs with higher exposure to trade liberalization with China, female workers experienced an increase in labor market participation even though the quality of their jobs could have worsened. Our findings are consistent with Pierce and Schott (2016) who show that granting PNTR to China in 2001 brought about a staggering reduction of employment in U.S. manufacturing sectors. Since manufacturing is more male-labor intensive, the direct effect from trade liberalization should be larger for male workers. The negative effects on male-intensive-labor markets might have increased the overall competitiveness in local labor markets by forcing more females to be engaged in economic activities. Our empirical results in this section may suggest that the lower gender gaps in MSAs with higher exposure to trade liberalization with China is driven by underperformance of male workers, rather than improved performance of female workers.

3.3. Estimates for Raicial Gaps

We now turn our attention to examining whether trade liberalization had a differential effect across racial groups. As in Pierce and Schott (2017), a region's higher exposure to PNTR may generate higher competition in the local labor market. This higher competition may differently affect different racial groups. Our hypothesis is that, if business managers tend to prefer White-American female workers over African-American ones, the higher competition in the labor market may hurt African-American workers more severely.

We run regression of (1) with work hours h_{mt}^{S} separately for White-American females, $S = FW$, and African-American females, $S = FA$. The first and second columns of table 8 report the estimated coefficients from specification (1) using White-American female's hours $\log(h_{mt}^{FW})$ as a dependent variable, while the third and fourth columns show the corresponding results with African-American females' $\log(h_{mt}^{FA})$. The DID estimates for African-American female workers are statistically significant in both columns, while the ones for White female workers are not statistically meaningful. Thus, the reduction in hours worked for all female workers, tends to be driven mostly by a reduction in hours worked for African-American female workers.

Table 8 : Work hours for different racial groups for women

	$\log(h_{mt}^{FW})$		$\log(h_{mt}^{FA})$	
Post * NTR Gap	-0.108	-0.0400	-0.578*	-0.538*
	(0.0671)	(0.0805)	(0.294)	(0.291)
NTR		0.676		1.710
		(0.430)		(1.497)
Post * Median HHI		0.0174		0.154***
		(0.0155)		(0.0522)
Post * No College		0.0130		0.354**
		(0.0316)		(0.137)
Post * Veteran		-0.0836		0.277
		(0.0745)		(0.205)
MSAs	252	252	252	252
Observations	4,353	4,315	4,353	4,315
R2	0.233	0.233	0.114	0.120
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

We also examine changes in the number of part-time-work weeks separately for White and African-American females. Table 9 reports estimation results with the number of weeks worked as part-time separately for each racial group. The first and second columns show the results for White female workers, and the third and fourth columns report the results with African-American female workers. The base-line specification with fixed effects reports the DID estimate for African-American female group as 2.155, about two and half times larger than the DID estimate for White female group. The specification with the controls shows similar magnitudes. We also looked at female workers' wage changes for the two racial groups.

Consistent with our findings in Table 5, the wage changes for both racial groups are not statistically significant.

Table 9 : Part-time work hours for different racial groups for women

	$\log(\hat{h}^{FW}_{mt})$		$\log(\hat{h}^{FA}_{mt})$	
Post * NTR Gap	1.016***	1.226**	2.155**	2.857***
	(0.359)	(0.509)	(0.920)	(1.093)
NTR		2.984		8.459
		(3.132)		(5.808)
Post * Median HHI		-0.124		-0.0733
		(0.139)		(0.179)
Post * No College		-0.0429		-0.671
		(0.294)		(0.450)
Post * Veteran		-0.321		-1.031
		(0.490)		(1.092)
MSAs	231	231	231	231
Observations	3,180	3,164	3,195	3,177
R2	0.281	0.284	0.176	0.177
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

Since there are differences across races for women, it is only natural to ask whether trade liberalization with China resulted in a differential impact across races for men as well. Tables 10 and 11 present the results for total hours worked and part-time hours worked for White (MW) and African-American (MA) men. For men, racial differences are more muted. If anything, White males experienced a larger reduction in total hours worked than African-American men, while African-American men experienced a larger increase in part-time work, though those results are not statistically very precise.

Our results in this section confirm our hypothesis. They also suggest that the competition in labor market caused by trade liberalization could even raise inequalities across racial groups. Increased racial inequality though that effect may be gender specific. Our results are somewhat surprising as they seem to suggest that among women, African-American women fared worse, while among men African-American men fared better than their White counterparts.

Table 10: Work hours for different racial groups for men

	$\log(h_{mt}^{MW})$		$\log(h_{mt}^{MA})$	
Post * NTR Gap	-0.0999***	-0.0864*	-0.0592	-0.178
	(0.0380)	(0.0495)	(0.321)	(0.252)
NTR		0.110		-1.760
		(0.308)		(1.913)
Post * Median HHI		0.0229**		0.0879*
		(0.00996)		(0.0501)
Post * No College		-0.0142		0.0718
		(0.0211)		(0.105)
Post * Veteran		0.0714		0.786**
		(0.0504)		(0.318)
MSAs	260	259	260	259
Observations	4,408	4,367	4,409	4,368
R2	0.302	0.308	0.119	0.127
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

Table 11: Part-time work hours for different racial groups for men

	$\log(\widehat{h_{mt}^{MW}})$		$\log(\widehat{h_{mt}^{MA}})$	
Post * NTR Gap	1.406**	1.336	2.109*	1.480
	(0.679)	(0.860)	(1.207)	(1.361)
NTR		1.654		-2.266
		(4.840)		(8.924)
Post * Median HHI		0.325*		0.148
		(0.165)		(0.272)
Post * No College		0.0512		0.669
		(0.478)		(0.658)
Post * Veteran		0.580		-2.791**
		(0.905)		(1.220)
MSAs	225	224	225	224
Observations	2,600	2,584	2,600	2,584
R2	0.247	0.247	0.190	0.188
Clustering	m	m	m	m
Fixed Effects	m,t	m,t	m,t	m,t

Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

4. Discussion

Our sample period includes the Great Recession of 2008, another large income shock to local economies. It would be hard to generalize our claims if our empirical results are driven by the coincidence of two large exogenous shocks. Hence, we separately estimate different time effects from the trade liberalization by running following specification:

$$\begin{aligned} LHS_{mt} = &\left(\theta_B \cdot 1_{01\le year\le 07} + \theta_M \cdot 1_{08\le year\le 10} + \theta_A \cdot 1_{11\le year\le 13}\right) \times NTR\ Gap_m \\ &+\left(\gamma_B \cdot 1_{01\le year\le 07} + \gamma_M \cdot 1_{07\le year\le 09} + \gamma_A \cdot 1_{10\le year\le 13}\right) \times \boldsymbol{Z_m} \\ &+\beta \boldsymbol{X}_{mt} + \delta_m + \delta_t + \varepsilon_{mt}. \end{aligned} \quad (2)$$

where $1_{01\le year\le 07}$ refers to an indicator variable for the pre-Great-Recession period, $1_{08\le year\le 10}$ is indicator for Great Recession years, and $1_{11\le year\le 13}$ refers to the post-Great-Recession period. Hence, the DID estimates on $\{\theta_B,\ \theta_M,\ \theta_A\}$ are of our interests. Table 10 shows the coefficients from (2) with the dependent variables considered above.

The point estimates for θ_B, reported in the 1st column of Table 11 confirms that our findings in Section 3 do not rely on the financial crisis. From 2001 to 2007, in MSAs with higher PNTR exposure, gender gaps in wages and labor participations decreased, but they are due to the underperformance of male workers rather than the improvement of female ones: male wages and LFPR significantly decreased, while female ones were not significantly affected. The same column also shows that female workers worked fewer hours but engaged in more part-time work, which may suggest welfare loss for female workers. Moreover, the negative effects from trade liberalization were larger for African-American females than for White-American counterparts. African-American females' work hours significantly decreased while the change of White-American females' work hours was not statistically significant.

We observe similar patterns in the 2nd and 3rd columns with point estimates for $\{\theta_M, \theta_B\}$. It is interesting that the impacts on labor participation lasted longer than the impacts on wages. Another interesting point is that the effects on the number of weeks, White-American females worked as part-time workers, disappeared in post-financial crisis period while the impacts on African-American females still remained.

Table 11 : Timing of labor market outcomes

	2001-2007	2008-2010	2011-2013
w^F_{mt}/w^M_{mt}	0.436***	0.466***	0.291
	(0.138)	(0.178)	(0.188)
$\log(w^F_{mt})$	0.215	0.158	0.114
	(0.138)	(0.141)	(0.169)
$\log(w^M_{mt})$	-0.322**	-0.455**	-0.346
	(0.159)	(0.22)	(0.211)
l^F_{mt}/l^M_{mt}	0.196***	0.340***	0.350***
	(0.057)	(0.131)	(0.111)
l^F_{mt}	0.0728	0.14	0.134*
	(0.0473)	(0.0957)	(0.0769)
l^M_{mt}	-0.111***	-0.150**	-0.178***
	(0.0373)	(0.0593)	(0.0599)
$\log(h^F_{mt})$	-0.169***	-0.0509	-0.0014
	(0.062)	(0.0958)	(0.0886)
$\log(\hat{h}^F_{mt})$	1.074**	1.098	1.214*
	(0.431)	(0.679)	(0.672)
$\log(h^{FW}_{mt})$	-0.0866	0.0567	0.118
	(0.0698)	(0.104)	(0.102)
$\log(h^{FA}_{mt})$	-0.720**	-0.874	-0.627
	(0.286)	(0.714)	(0.382)
$\log(\hat{h}^{FW}_{mt})$	0.826*	1.311**	0.894
	(0.452)	(0.655)	(0.651)
$\log(\hat{h}^{FA}_{mt})$	2.016*	4.510***	4.619***
	(1.162)	(1.294)	(1.373)

Standard errors in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.1

5. Conclusion

This paper provides empirical findings that the conferral of PNTR status to China lowered the gender gaps in the US labor market. However, this lower gender gaps do not imply female welfare gains since the lower gender gaps are driven by the underperformance of male workers. We also find that the impacts from the trade liberalization were higher for African-American female workers than for White-American ones.

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