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The Gender Side of Trade Shocks: Evidence from the Italian Labor Market*

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Abstract

This paper investigates the gendered effects of trade liberalization on local labor markets in Italy, a country marked by low female labor force participation. Building on recent evidence that trade shocks can exacerbate or mitigate gender inequalities depending on labor market segmentation and institutional context, we examine how exposure to Chinese and Eastern European import competition has affected the labor market in Italy, with a focus on the gender discrepancies. We construct a shift-share measure of import exposure, exploiting variation in pre-existing industry specialization across provinces. Using labor-force survey and trade data with detailed labor market indicators, we assess whether observed gender gaps result from asymmetric dynamics between women and men, and how these patterns vary by sector, contract type, and skills. By providing new empirical evidence and a theoretical framework to interpret these patterns, our findings indicate that trade shocks tend to reinforce existing gender disparities in Italy, with effects concentrated in sectors characterized by high female employment shares and precarious job arrangements.

Keywords: Import competition, labour market, gender inequality
JEL classification: F14, J21, J16

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1 Introduction

Since the early 2000s, WTO-driven trade liberalization (especially following China's accession) has triggered profound changes in labour markets worldwide. A growing literature has documented its impact on wage inequality and job polarization (Autor et al., 2013), firm productivity (Mion and Zhu (2013);Forlani et al. (2021)), R&D intensity (Bloom et al., 2016), and political preferences in affected regions (Autor et al. (2020); Colantone and Stanig (2018); Barone and Kreuter (2021)). Firms respond to global competition by reorganizing production, with far-reaching consequences for employment composition, wages, and career trajectories. However, while the aggregate labour market effects of trade shocks are well established, their gender-specific consequences remain less understood.

Recent studies have begun to uncover how trade-induced shocks may differentially affect male and female workers, due to persistent gender segregation in tasks, sectors, and contract types, as well as differing household constraints and labour supply responses. These impacts are especially relevant in countries like Italy, where female labour force participation (LFP) remains among the lowest in Europe, despite a relatively contained average gender pay gap. This apparent paradox underscores the need to better understand the mechanisms through which global economic forces interact with gender dynamics in the labour market.

An emerging body of evidence suggests that gendered impacts of trade are not merely country-specific anomalies, but reflect structural features of labour markets—such as segmentation, informality, and institutional constraints. For instance, Heckl (2024) documents how trade shocks in Mexico increased informal self-employment especially among women, pointing to a coping mechanism with long-term scarring effects. Ghosh et al. (2022) find that displaced women in the United States suffer disproportionately large and persistent earnings losses relative to men, despite similar pre-displacement profiles. Goes et al. (2023) highlight that in countries with high gender segmentation, like Tunisia, trade shocks can decrease the female-to-male employment ratio, particularly among married women—suggesting intra-household labour substitution in response to male-dominated sector expansion.

These findings show that while the direction and magnitude of trade's gender effects may differ across settings, a common pattern emerges: women tend to bear a disproportionate share of the adjustment costs. This is not only due to differences in industrial concentration or policy regimes, but also because of how trade interacts with pre-existing gendered frictions in the labour market. In this context, Italy offers a particularly relevant case for investigation. The Italian labor market is characterized by pronounced gender asymmetries in sectoral allocation, job characteristics (such as part-time and fixed-term work), and household responsibilities (Mussida and Picchio (2014); Castagnetti and Giorgetti (2019)). In addition, macroeconomic shocks—such as international migration flows and the COVID-19 pandemic—have disproportionately affected women's employment (Caselli et al. (2022)), raising concerns about cumulative disadvantage. Importantly, Italy also exhibits significant regional (and provincial) heterogeneity in both trade exposure and female LFP, which we exploit empirically.

Our research aims to fill this gap by studying how exposure to import competition (particularly from China and Eastern Europe) has shaped women's employment and earnings outcomes in Italy. We focus on a key period of structural change (2004-2019), using detailed ISTAT data across local labour markets/provinces and industries[1], and Comtrade data. We construct a Bartik-style shift-share index of import exposure, exploiting variation in pre-shock industry structure to identify exogenous changes in trade exposure at the local level (Autor et al. (2013); Acemoglu et al. (2016); Acemoglu et al. (2016)).

Our contribution is threefold. First, we provide robust empirical evidence on how international trade competition affects women's labor market outcomes in Italy. Second, we investigate the key mechanisms about this gender gap. Third, we complement the empirical analysis with a theoretical framework that clarifies the channels through which trade shocks affect gendered labor allocation. By linking trade shocks with gender-segmented labor market outcomes in a Southern European context, our study contributes to a broader understanding of how globalization interacts with persistent labor market inequalities. In doing so, we provide novel evidence that can inform the design of policies aimed at mitigating adverse effects and promoting gender equity in an increasingly globalized economy.

The rest of the paper is organized as follows. Section 2 presents the literature review. Section 3 presents a theoretical framework to interpret the findings. Section 4 describes the data and institutional context. Section 5 presents the empirical strategy and the main results. Section 6 shows the robustness checks. Section 7 concludes.

[1] Our project is based on data from the Adele Laboratory (ISTAT). For this reason, we are not allowed to disclose the data before the project is completed and validated, which is expected to happen by February 2026. This draft of the paper will therefore be incomplete, as we cannot present the values of the coefficients from our analyses (although they have already been computed) and will only be disclosed once the project is finalized.

2 Literature review

Trade liberalization (particularly following China's accession to the WTO) has been the focus of intense empirical scrutiny in the past two decades, with a growing consensus that rising import competition has had deep and uneven effects on labour markets. However, only recently have scholars begun to explore how these effects vary by gender. This section reviews both strands of the literature: the general consequences of trade shocks on labour market outcomes, and their gender-differentiated impacts, with a specific focus on the Italian context.

A substantial literature has investigated the labour market effects of rising import competition. A foundational contribution by Autor et al. (2013) exploits cross-market variation in exposure to Chinese imports across U.S. local labour markets, instrumenting U.S. import penetration using Chinese exports to other high-income countries. They find that trade-exposed areas experienced higher unemployment, lower labour force participation, and reduced wages, with import competition accounting for one-quarter of the aggregate decline in U.S. manufacturing employment. These adverse effects were accompanied by a significant rise in government transfer payments for unemployment, disability, retirement, and healthcare. Complementing this evidence, Bloom et al. (2016) show that in twelve European countries between 1996 and 2007, Chinese import competition fostered innovation, measured through patenting, IT adoption, and total factor productivity, particularly within affected firms. While employment decreased and the share of unskilled workers declined, technological upgrading occurred both within and across firms, accounting for 14% of technology improvements in the period.

Further evidence from Mion and Zhu (2013) using Belgian firm-level data shows that Chinese import competition reduced employment growth and induced skill upgrading in low-tech manufacturing industries. Manual workers were particularly affected, though firm survival was not impacted, and offshoring to China increased the probability of survival and contributed modestly to the rise in non-production workers. Keller and Utar (2022), analyzing Danish employer-employee matched data, find that rising import competition following the dismantling of import quotas led to job polarization: employment declined in mid-wage occupations but increased in both high- and low-wage jobs. Workers with higher education tended to move upward, while those in manual occupations were most negatively affected, independent of routine-task intensity—indicating a distinct mechanism from routine-biased technological change.

Several studies have linked trade shocks to political realignments. Barone and Kreuter (2021) find that exposure to Chinese imports across Italian municipalities from 1992 to 2013 increased support for populist parties, abstention, and invalid voting, mediated by deteriorating local labour market conditions such as higher unemployment and greater inequality. Similarly, Autor et al. (2020) show that trade-exposed U.S. districts experienced growing political polarization between 2000 and 2016, with rightward shifts in some areas and stronger ideological divides in others. These patterns were particularly evident in racially distinct districts: White-majority areas became more likely to elect Republicans, while majority-minority districts shifted towards more progressive Democrats, largely at the expense of moderates.

More recent research explores the gendered dimensions of trade liberalization. Heckl (2024), focusing on Mexico, shows that while both men and women experienced slower employment growth in trade-exposed areas, male labour force participation declined while female participation increased, driven by a shift to self-employment and informality, which served as a buffer. In the U.S., Ghosh et al. (2022) use detailed Trade Adjustment Assistance data to examine post-layoff outcomes and documents that while pre-layoff wage gaps between men and women were substantial, re-employment outcomes showed convergence due to wage compression, with women performing comparably to men in terms of reemployment rates. Brussevich (2018) uses a dynamic structural model to simulate trade shocks in the manufacturing sector, finding that men are more adversely affected due to higher exit costs and male-dominated employment structures, implying higher welfare gains from trade for women. Besedeš et al. (2021) also find that trade liberalization with China reduced the gender wage gap in U.S. metropolitan areas, with higher female labour market entry—especially among educated women—and exit of less-educated men. Nevertheless, residual wage gaps increased, indicating significant selection effects. Gendered intrahousehold adjustments emerged, with women entering the labour force to offset male job loss, though often in part-time or less stable roles. Goes et al. (2023) develop a theoretical model showing that the gendered effects of trade depend on the sectoral composition of shocks and the gender intensity of employment. Empirically, they find that in Tunisia—where male-dominated sectors were most exposed to foreign demand shocks—the female-to-male employment ratio declined as households substituted male for female labour supply. Gaddis and Pieters (2017), studying Brazil's trade liberalization, similarly show that although both men and women reduced participation in tradable sectors, effects were stronger for men in absolute terms, leading to a narrowing of the gender employment gap without proportional gains for women.

Despite a rich international literature, the gendered impacts of trade liberalization in Italy remain largely unexplored. Italy exhibits one of the lowest female labour force participation rates in Europe, especially in the South, alongside high gender segmentation, persistent occupational clustering, and a large informal sector (Mussida and Picchio (2014); Castagnetti and Giorgetti (2019)). Women are overrepresented in temporary, part-time, and low-paid jobs and tend to exit the labour force during downturns (Forlani et al. (2015); Caselli et al. (2022)). While Barone and Kreuter (2021) shed light on the political implications of trade shocks in Italy, little is known about how such shocks interact with existing gender inequalities in the labour market. This paper seeks to address this gap by offering the first empirical analysis of the gendered effects of Chinese import competition in Italian local labour markets, contributing to broader discussions on the distributional consequences of globalization.

3 Theoretical model

In this section, we propose a simple theoretical model to explore how trade competition may affect employment across genders. We could make different assumptions on the structure of commodity markets (perfect competition or monopolistic competition with differentiated products) and on the relevance of the country in the world economy (small versus large economy). However, the main point is that some outside shock affects the domestic market on the supply side, reducing the optimal supply of domestic firms. The details about how this happens do not matter much. To avoid to get bogged down in the details concerning the different possible market structures, we will take the shock to the *quantity* sold by domestic firm as exogenous and we will consider the adjustments of the labor markets to this shock. In Appendix 1, we will specialize the analysis to competitive markets outlining the *full* adjustment to a supply shock.[2]

Generally speaking, a trade shock may impact on the gendered labor market conditions through two distinct channels. The first is the first order effect of the direct impact of the supply shock due to the increased imports from China and Eastern European countries, our measure of import competition, on the conditions of the different good markets. Evidently, if the shock is stronger in an industry with a male (or female) intensive labor force, the quantitative impact could easily be stronger for the most common gender in that industry labor force. However, for a large class of production functions, the inputs expansion path are linear,[3] so that the *percentage* change

[2] As we will see, this will not change substantially our results.

[3] Conditions identifying production function with linear input expansion path are analyzed in Färe and Mitchell (1992). Essentially, the condition is that we can decompose the cost function as $C(w, Y) = c(w)\theta(Y)$, which requires the production function to be

in employment due to a decrease in output is, as a first order effect, identical across genders.

There is, however, a second channel which can produce gender differentiated effects even with homothetic production functions (hence, with linear labor expansion paths). It rests on the second order effect induced by the wage adjustments required to bring the labor markets back to the equilibrium after the shock. Here, the crucial property at play are the differences in the labor supply elasticities across genders, which are a well-established empirical fact.

Let's now sketch a simple partial equilibrium model, aiming to make the argument precise and thus allowing us to clarify the two different effects. For now, let's focus on the conditional demand functions, implicitly taking as given a shock to the quantity demanded of the output of the national firms and ignoring the feedback effects from the adjustments on the labor markets to the additional adjustments in the output markets, which obviously depends upon the specific structure of these markets.

In each industry j, and for each firm, the production function is described by

$$Y_j(.) = \prod_{\ell=1}^{\ell=L} Q_{j\ell}^{\theta_{j\ell}} \left[\left(\phi_M Q_{jM}^{\beta_j} + \phi_F Q_{jF}^{\beta_j} \right)^{\frac{1}{\beta_j}} \right]^{\theta_{jL}},$$

where $Q_{j\ell}$ is the input of commodity ℓ in the production of each firm in industry j. (Q_{jM}, Q_{jF}) are the inputs of male and female labor. For each firm, $Y_j(.)$ exhibits decreasing return to scale (alternatively, assume that we consider the short-run), i.e., we assume that $\sum_{\ell=1}^{\ell=L} \theta_{j\ell} \equiv \overline{\theta}_j < 1$, for each j. The term in square brackets is a (constant return to scale) *aggregator function* for the two types of labor inputs. Evidently, $(\phi_M, \phi_F) >> 0$ and $\beta_j \in (-\infty, 1]$. Let w_ℓ be the price of each input ℓ. .

We compute the (conditional) demand functions for the two types of labor using a two stage process. First, we compute the demand for the aggregator variable L, then we analyze the actual demand for male and female labor. For the time being, then assume that there is a well-defined input "aggregate labor" with quantity L and price w_L. Hence, the production function reduces to a Cobb-Douglas and, as standard, the demand and supply functions are (we omit the index j for simplicity)

$$Q_\ell(w, p) = \frac{\theta_\ell}{w_\ell} \left[\prod_{\ell=1}^{\ell=L} \left(\frac{\theta_\ell}{w_\ell} \right)^{\theta_\ell} \right]^{\frac{1}{1-\overline{\theta}}} p^{\frac{1}{1-\overline{\theta}}} \text{ and } Y(w, p) = \left[\prod_{\ell=1}^{\ell=L} \left(\frac{\theta_{j\ell}}{w_\ell} \right)^{\frac{\theta_\ell}{1-\overline{\theta}}} \right] p^{\frac{\overline{\theta}}{1-\overline{\theta}}}.$$

homotetic.

The conditional demand functions are

$$Q_\ell(w,Y) = \frac{\theta_\ell}{w_\ell}\left[\prod_{\ell=1}^{\ell=L}\left(\frac{w_\ell}{\theta_\ell}\right)^{\frac{\theta_\ell}{\overline{\theta}}}\right]Y^{\frac{1}{\overline{\theta}_j}} \equiv Q_\ell(w)\,Y^{\frac{1}{\overline{\theta}}},$$

with partial derivative

$$\frac{\partial Q_L(w,Y)}{\partial w_L} = -\frac{\overline{\theta}-\theta_L}{\overline{\theta}}\frac{1}{w_L}Q_L(w)\,Y^{\frac{1}{\overline{\theta}}}$$

Let's now take as given the desired value of the aggregator variable, Q_L, and compute the optimal value of (Q_M, Q_F). Evidently, profit maximization requires that each value of Q_L is obtained with the cost minimizing combination of the two inputs. Solving the cost minimization problem

$$\min_{(Q_M,Q_F)} w_M Q_M + w_F Q_F \quad \textit{subject to} \quad \left[\phi_M Q_M^\beta + \phi_F Q_F^\beta\right]^{\frac{1}{\beta}} = Q_L$$

we obtain the *unit* cost function

$$w_L(w_M,w_F) = \left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{w_F}{\phi_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{\beta-1}{\beta}},$$

which can be interpreted as the unit wage of the "optimally aggregated labor". The conditional demand functions for the two types of labor are

$$Q_g(w,Y) = \frac{\left(\frac{w_g}{\phi_g}\right)^{\frac{1}{\beta-1}} \times Q_L(w) \times Y^{\frac{1}{\overline{\theta}}}}{\left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{\phi_F}{w_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{1}{\beta}}} \equiv \widetilde{Q}_M(w) \times Q_L(w) \times Y^{\frac{1}{\overline{\theta}}}.$$

Given that $\widetilde{Q}_g(w,Y) = Q_g(w)\,Y^{\frac{1}{\overline{\theta}}}$, for each g, the first order effect of a trade shock is $\frac{\partial \widetilde{Q}_g(w,Y)}{\partial Y} = \frac{\widetilde{Q}_g(w,Y)}{\overline{\theta}Y}$ and the quantitative impact of a change in Y on the gendered labor demands just depends on the initial employment ratio $\widetilde{Q}_M(w,Y)/\widetilde{Q}_F(w,Y) = \left(\frac{w_M}{\phi_M}\frac{\phi_F}{W_F}\right)^{\frac{1}{\beta-1}}$. To see this in a different way: the two output elasticities of the labor demand functions are both equal to $\frac{1}{\overline{\theta}}$, so that the employment ratio is unaffected by a pure change in the value of Y, i.e., it is gender neutral.

The second order relative effect is, however, not gender neutral: The adjustment of employment to shocks also rests on the reactions of the two equilibrium wages. If the wage elasticities of the labor supplies differ across genders, an empirically well-established fact, the equilibrium adjustments of employment must also differ.

To carry out this argument, we need to formalize the supply side of the labor markets. Consider the simplest model, with all the workers characterized by quasi-linear utility functions, $U_g(C_g, H_g) = C_g - \frac{H^{1+\frac{1}{\gamma_g}}}{1+\frac{1}{\gamma_g}}$. Given the consumption price p, the associated demand and supply functions are $H_g(w_g) = \left(\frac{w_g}{p}\right)^{\gamma_g}$ and $C_g(w_g) = \left(\frac{w_g}{p}\right)^{1+\gamma_g}$. The real wage elasticity of the labor supply is γ_g. Assume that $\gamma_F > \gamma_M$, with γ_M close to zero. Also, assume that there exist a continuum $[0,1]$ of identical workers of a given gender, and an interval $[0,1]$ of identical firms.

As we have argued, the direct impact of the trade shock on employment is proportional to dY, with coefficients identical to the ones defining the relative employment for the two genders, in this well-defined sense, the first order effect is gender neutral.

Let's consider now the total effects on employment, taking into account the adjustments of the equilibrium wages. Take as fixed the output price, $p = 1$. Omitting the integrals, given $\overline{Y}$, the equilibrium on the labor markets is described by

$$\begin{aligned}\Psi(w,Y) &\equiv \begin{bmatrix} \Psi_M(w,Y) \\ \Psi_F(w,Y) \end{bmatrix} \equiv \begin{bmatrix} Q_M(w,Y) - H_M(w_M) \\ Q_F(w,Y) - H_F(w_F) \end{bmatrix} \\ &\equiv \begin{bmatrix} \widetilde{Q}_M(w)Q_L(w)Y^{\frac{1}{\theta}} - H_M(w_M) \\ \widetilde{Q}_F(w)Q_L(w)Y^{\frac{1}{\theta}} - H_F(w_F) \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix}.\end{aligned}$$

By the implicit function thm., the wage adjustments to an exogenous trade shock are given by

$$\nabla_Y w^T = -\left[D_w\Psi\right]^{-1} \nabla_Y \Psi^T,$$

or

$$\begin{bmatrix} \frac{\partial w_M}{\partial Y} \\ \frac{\partial w_F}{\partial Y} \end{bmatrix} = \frac{1}{\det D_w\Psi} \begin{bmatrix} -\frac{\partial Q_F}{\partial w_F} + \frac{\partial H_F}{\partial w_F} & \frac{\partial Q_F}{\partial w_M} \\ \frac{\partial Q_M}{\partial w_F} & -\frac{\partial Q_M}{\partial w_M} + \frac{\partial H_M}{\partial w_M} \end{bmatrix} \begin{bmatrix} \frac{\partial \Psi_M}{\partial Y} \\ \frac{\partial \Psi_F}{\partial Y} \end{bmatrix},$$

where

$$\begin{bmatrix} \frac{\partial \Psi_M}{\partial Y} \\ \frac{\partial \Psi_F}{\partial Y} \end{bmatrix} = \begin{bmatrix} \left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}} \\ \left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}} \end{bmatrix} \frac{Q_L(w)\,\overline{Y}^{\frac{1}{\theta}}}{\overline{\theta Y}\left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{\phi_F}{w_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{1}{\beta}}}.$$

In Appendix 1, we show that

$$\begin{aligned}
\frac{\partial w_M}{\partial Y} &= \frac{A}{w_F} + \left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}} \frac{\partial H_F\left(w_F\right)}{\partial w_F}, \\
\frac{\partial w_F}{\partial Y} &= \frac{A}{w_M} + \left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}} \frac{\partial H_M\left(w_M\right)}{\partial w_M},
\end{aligned}$$

where

$$A \equiv \left(\frac{\left(Q_L(w)Y^{\frac{1}{\theta}}\right)^2 \left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}} \left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}}}{\left(1-\beta\right)\overline{\theta Y} S^{1+\frac{1}{\beta}}} \right).$$

More interesting is to understand what happens when the male labor supply is very inelastic (i.e., when γ_M is very close to zero), while the female labor supply is fairly elastic (i.e., when γ_F is comparatively large). Consider the polar case where the male labor supply is perfectly inelastic. Define the wage-gap as $\Delta \equiv \frac{w_F}{w_M}$,

$$\begin{aligned}
\frac{\partial \Delta}{\partial Y} &= \frac{\left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}} \frac{\partial H_F(w_F)}{\partial w_F} w_F - \left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}} \frac{\partial H_M(w_M)}{\partial w_M} w_M}{w_F^2} \\
&= \frac{\left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}} H_F\left(w_F\right)\gamma_F - \left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}} H_M\left(w_F\right)\gamma_M}{w_F^2}.
\end{aligned}$$

Then, a negative trade shock will reduce the wage gap (when $\gamma_M = 0$). Conversely, the adjustment in employment will be nil (or relatively small) for male, and larger for female workers:

$$\begin{aligned}
\frac{\partial H_M}{\partial Y} &= \frac{\partial H_M}{\partial w_M}\frac{\partial w_M}{\partial Y} = \gamma_M \frac{H_M\left(w_M\right)}{w_M}\left(\frac{A}{w_F} + \left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}} \frac{\partial H_F\left(w_F\right)}{\partial w_F}\right), \\
\frac{\partial H_F}{\partial Y} &= \frac{\partial H_F}{\partial w_F}\frac{\partial w_F}{\partial Y} = \gamma_F \frac{H_F\left(w_M\right)}{wF}\left(\frac{A}{w_M} + \left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}} \frac{\partial H_M\left(w_M\right)}{\partial w_M}\right).
\end{aligned}$$

To consider a fully developed and formalized model of the output side would not significantly enhance the results or interpretation. The key point is that, due to an exogenous trade shock, the optimal quantity produced by a firm decreases. This is true in both large and small open economies. Here, in line with our empirical strategy, we focus on the effects of thisa change on the labor market.[4] To do this, we focus on the conditional demand functions.

[4] In the non-tradable sector, the qualitative results remain unchanged.

This allows us to simplify drastically the analysis without any substantive loss of analytical results, at least as long as we are willing to accept the (quite general) assumptions which guarantees that the input demand function are multiplicative in output.[5]

4 Data Description

Our analysis relies on three main data sources: i) the ISTAT Labour Force Survey, ii) the Censuses of industry and services for 1991, and iii) the BACI-ComTrade Database.

ISTAT - Labour Force Survey - The ISTAT Labour Force Survey (LFS) is the most timely and comprehensive source of statistical information on the Italian labor market. Conducted at the individual level, it provides crucial insights into labor supply by collecting data directly from the population. The LFS forms the basis for official estimates of employment and unemployment and offers detailed information on key labor market aggregates, including occupation, industry sector, hours worked, contract type and duration, and training. In addition, it reports demographic characteristics such as gender, educational attainment, geographic location as well as occupational status: employed, unemployed, and inactive individuals. Using sampling weights from ISTAT, we aggregate quarterly cross-sectional data to construct statistics for 103 province (NUTS-3 level) for the period 2004 to 2019. [6]

ISTAT - Census - The Census of Industry and Services (Censimento dell'Industria e dei Servizi [7]) is a decennial survey conducted by ISTAT to collect detailed structural data on the productive sectors of the Italian economy, including industry, commerce, services, and craftsmanship. Its origins date back to 1927, and from 1951 onwards, the census was conducted every ten years, progressively expanding in coverage and improving in methodology. The 1991 edition marked a significant turning point, introducing a modern classification of units of analysis into enterprises, institutions, and local units, and fully aligning the Italian statistical system with European and

[5]In Appendix 1, we consider also the case of perfectly competitive markets. Focusing on the effect of a supply price shock, we show that the results in the text are indeed robust.

[6]Details at https://www.istat.it/informazioni-sulla-rilevazione/forze-lavoro/. The LFS data accessed through LABORATORIO ADELE also include information at the city level, enabling the construction of local labor market indicators. However, these statistics suffer from limited representativeness, and approximately one-third of local labor markets lack sufficient data for reliable aggregation.

[7]See https://www.istat.it/statistiche-per-temi/censimenti/censimenti-storici/industria-e-servizi/).

international standards. In our analysis, we make use of the weights derived from this 1991 census. These weights provide a consistent reference point for evaluating structural characteristics of the Italian economy at a fine-grained geographic level. In 1996, a mid-decade intermediate census was conducted, followed by the 8th and 9th editions in 2001 and 2011, respectively. Over time, the census evolved to incorporate data from administrative sources, particularly through the development of the ASIA (Archivio Statistico delle Imprese Attive) statistical register, which integrates data from tax authorities, social security institutions, and chambers of commerce. Since 2016, the traditional decennial format has been replaced by a system of permanent economic censuses.

Baci - Comtrade - Trade data are retrieved from BACI - Comtrade database. We collect trade data at HS classification (6 digit) recombine with ATECO classification used by ISTAT using correspondence tables.

4.1 Descriptive Statistics

We rely on a balanced panel of 1,648 observations, consisting of 103 Italian provinces observed over a 16-year period from 2004 to 2019. On average, the female unemployment rate is 11%, compared to 8% for men. The female inactivity rate—defined as the ratio of inactive women to the female labor force—stands at a striking 40%, while the corresponding figure for men is 18%.

Turning to long-term differences between 2012 and 2019 (based on 206 observations), we observe a modest increase in male unemployment (1.5 p.p from 2004 to 2011 and 1.8.pp in 2019-2012), while female unemployment first decreases slightly in 2011 (–0.8pp) and then rises considerably by 2019 (+2.3pp). The share of men employed in manufacturing steadily declines (-1.2pp in 2011 and -1.5pp in 2019), as does that of women, although more mildly (-1.7pp in 2011, with virtually no change in 2019). Inactivity rates display more pronounced changes: male inactivity rises slightly in 2011 (+0.9pp) but decreases notably by 2019 (–2.5pp); female inactivity declines consistently (–1.6pp in 2011 and –5pp in 2019). Employment in services (education and health) increases slightly for women over time (6pp in 2011 and 1.1pp in 2019), whereas the male share remains almost flat. During the same period, the average increase in China's import competition for Italy is 0.04, while for other countries it is as high as 0.85 (see Eq.2).

5 Empirical strategy

In this section, we analyze the effect of trade shocks on the gender gap at the provincial level. We begin by defining the econometric model, discussing potential sources of endogeneity, and outlining the strategies used to address these issues, including tests for the validity of the identification strategy. We then present the main empirical results, examine the impact of trade shocks on the gender gap, and provide a series of robustness checks to support our findings.

Econometric model and identification

Our empirical strategy closely follows Autor et al. (2013). The identification strategy relies on cross-provincial variation in initial industry structure and uses Italian import growth from China in manufacturing industries over the period 2004–2019. We analyze the effect of import competition on changes in employment levels—both overall and by gender—at the provincial level by aggregating individual-level data on an annual basis. More specifically, we estimate the following first-difference equation:

$$\Delta Y_{pt} = \alpha \Delta IP_{pt} + X_{p,1991}\boldsymbol{\beta} + \gamma_p + \gamma_t + \varepsilon_{pt} \tag{1}$$

where $\Delta Y_{pt} = Y_{pt} - Y_{pt-\tau}$ denotes changes in labor outcome variables in the province p (e.g. Number of inactive individuals, Number of employed individuals, Number of individuals employed in manufacturing) between t and $t-\tau$, with $\tau = 8$ or 3 years: in the gender-specific analysis, labor market outcomes are analyzed separately for men and women. ΔIP_{pt} is the import shock assigned to province p. The variable $X_{p,1991}$ represents the controls: for example, we consider the the share of migrants in the total population in province in 1991. γ_p and γ_t are the province/region and time fixed effects, respectively, that account for differences within areas and periods that could potentially impact labor market outcomes. [8] Finally, standard errors are clustered at province level.

We consider both the short-run and long-run effects of import changes. For the short-run analysis, we use equally spaced three-year intervals between observations from 2004 to 2019. For the long-run analysis, we focus on two intervals, 2004-2011 and 2012-2019.

Import Competition Exposure In order to analyze the effect of import competition from China at province level, we measure for each industry j the

[8] If we use regional fixed effects, we add as control variable the share of people employed in manufacturing in 1991.

change in the import penetration ratio

$$\Delta IP_{j\tau}^{ITA} = \frac{\Delta M_{j\tau}^{ITA}}{Prod_{j95} + M_{j95} - X_{j95}} \tag{2}$$

where $\Delta M_{j\tau}$ is the change in the Italian import from China in industry j on the time horizon τ (three or eight years); $Prod_{j95}$, M_{j95}, and X_{j95} are the domestic production, import, and export for industry j in 1995. IP measures industry j's exposure to export competition resulting from a Chinese supply shock, defined as the share of Italian domestic consumption accounted for by Chinese exports.

To evaluate the effect of import shocks on provincial labor market outcomes, we redistribute shocks as in Eq.2 using the 1991 employment share of industry j within provinces. The goal is to investigate whether different exposure to international competition shocks lead to different labor market outcomes at the provincial level, with particular attention to gender differences.

The Bartik-type measure of exposure to import competition is defined as follows:

$$\Delta IP_{pt}^{ITA} = \sum_j \frac{L_{jp,1991}}{L_{p,1991}} \Delta IP_{j\tau}^{ITA} \tag{3}$$

where $\Delta IP_{j\tau}^{ITA}$ is the change in the import penetration from China in the period τ. $L_{jp,1991}$ is the employment level in industry j and province p in 1991, while $L_{p,1991}$ is the total private non-agricultural employment in 1991.[9]

The import penetration index, as defined in Eq.3, relies on two sources of identification: the industry employment shares and the industry-level import shocks (the shifts). In order to identify the causal effect of import competition in provincial labour market outcomes we need that both sources.

Identification Strategy

The estimation strategy outlined in Eq. 1 may be biased due to endogeneity arising from omitted variables and reverse causality. As previously mentioned, the exogeneity of both the initial industry shares and the trade shocks (i.e., shifts in Chinese exports) is crucial for identifying an unbiased estimate of the coefficient β_1 (Borusyak et al. (2025) and Goldsmith-Pinkham et al. (2020)).

[9] Notice that the weights at province level do not sum up to 1, given that we consider only exposure in manufacturing sectors. Industry are defined according to the Ateco91 classification at four digit.

A key source of concern is the potential correlation between $\Delta IP^{ITA}_{j\tau}$ and the error term, which could result from unobserved industry-specific or local demand shocks that also affect employment at the local level. To address this issue and isolate supply-driven trade shocks from demand-driven changes in Italian imports from China, we adopt an instrumental variables (IV) approach commonly used in the recent literature (Acemoglu et al. (2016), Autor et al. (2013), Citino and Linarello (2022)).

Specifically, we instrument the change in Italian imports from China (numerator of Eq.2) with the same change on Chinese exports to other high-income countries outside the EU (denoted as OC). [10] The instrument is defined as:

$$\Delta IP^{OC}_{j\tau} = \frac{\Delta M^{OC}_{j\tau}}{Y_{j,95} + M_{j,95} - X_{j,95}}, \tag{4}$$

where $\Delta M^{OC}_{j\tau}$ captures the change in Chinese exports to developed countries other than Italy, and the denominator remains the usual normalization term for the size of the industry. Thus, we instrument Eq.3 using the following variable

$$\Delta IP^{ITA}_{pt} = \sum_j \frac{L_{jp,1991}}{L_{p,1991}} \Delta IP^{OC}_{j\tau} \tag{5}$$

The underlying idea is that China's accession to the WTO and structural changes in its economy led to rapid expansion in sectors where it held a comparative advantage (Bloom et al., 2016). This growth translated into increased exports across multiple destinations. For the IV to be valid, it must be correlated with $\Delta IP^{ITA}_{j\tau}$ through China's rising productivity and export growth, but uncorrelated with Italian demand for Chinese goods or sector j technological shocks. In this way, the instrument isolates the component of Italian import growth that reflects supply-side developments in China, rather than shifts in domestic demand correlated with labor market dynamics. We are quite confident that the instrument captures changes in Italian import but it is not correlated with sector and demand shocks. [11]

Conditional on treating the shifts in Eq. 4 as exogenous, the unbiased identification of α hinges on the assumption that the labor market share is exogenous (Borusyak et al., 2025). Thus, cross-sectional heterogeneity is driven by differences in the manufacturing share of employment and by differences in sectoral composition within provinces. To account for this, we

[10] Countries are: Australia, Canada, Japan, New Zealand, and United States.

[11] The correlation between Italian imports and import for other countries is equal to 0.65 with a maximum value of 0.75 for 2009.

include the 1991 province-level share of manufacturing employment as a control variable (i.e., the sum of weights), while also using region fixed effects. Alternatively, we employ province fixed effects. In other words, we control for the fact that some provinces are more manufacturing-oriented—and therefore may receive higher values of the instrument—than others. This approach allows us to identify the effect from variation in industrial composition across provinces with similar levels of manufacturing specialization, thereby mitigating concerns about omitted variable bias.

In addition, to ensure the exogeneity of the employment share, we use the industrial distribution as of 1991. This helps minimize the risk that unobserved factors influencing current labor market dynamics are correlated with the industry mix across province. In Section 6, we formally discuss and test the exogeneity assumption of the share variable.

Results

We begin this section by presenting the main results from the estimation of Eq.1 over two time periods: τ = 2011–2004 and τ = 2019–2012. [12] The outcome variables, $\Delta Y_{p\tau}$, represent changes in the unemployment rate, the share of people employed in manufacturing, the share of inactive individuals, and the share of individuals employed in the health and education sectors (all measured relative to the labor force). These variables are reported separately for the total population, as well as by gender (female and male). The results are based on an instrumental variable approach and standard errors are clustered at province level.

First, we report and summarize (Table 1) the results obtained by using region and period fixed effects; we consider as control variables both the change in the skilled worker share at province level as well as the share of manufacturing workers in 1991. Moreover, to account for the possible confounding effect of migration on native women's labor market outcomes, we include as a control the share of migrants in the total population in province p in 1991 (Barone and Mocetti, 2011; Forlani et al., 2015). However, rather than using the observed number of migrants, we adopt a shift-share (or "predicted share") measure: the share is computed by re-allocating national inflows of extra-eu migrants according to the 1991 distribution of residence permitst of migrants across Italian provinces, following Barone et al. (2016).

Unemployment rate - The unemployment rate increases as a consequence of import competition. Provinces more exposed to Chinese trade

[12] We have 206 observations, 103 provinces by 2 time periods.

Table 1: Summary of estimated effects of import competition on labor market outcomes

	Total	Female	Male	Notes
Unemployment rate	$+^{***}$	$+^{***}$	$+^{**}$	Higher (+) effect for women.
Manufacturing employment	$-^{***}$	$-^{***}$	n.s.	Negative effect for women.
Inactivity rate	n.s.	n.s.	$-^{**}$	Decrease among men only.
Employment in health & education	$+^{*}$	$+^{**}$	n.s.	Increase for women only.

Notes: The table reports only the sign and statistical significance of estimated coefficients. $n.s.$ = not statistically significant. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

shocks experience a larger rise in unemployment compared to less exposed regions. Notably, the increase in unemployment is more pronounced among women, with nearly double the growth rate compared to men.

Manufacturing employment - Consistent with Citino and Linarello (2022), the share of individuals employed in manufacturing declines. This decline is primarily driven by a reduction in female employment in the sector, rather than by a drop in male employment.

Inactivity rate - The share of inactive individuals (relative to the labor force) decreases, but only among male workers. In provinces more exposed to Chinese import competition, the reduction in the inactivity rate tends to be larger (i.e., more negative).

Employment in Health and Instruction - We find a positive, though not strongly significant, effect of trade shocks on the share of women employed in the health and education sectors. In particular, we consider the education, healthcare, and residential social work services sector.

In all the models, the F-Statistics is above 100 making us confident of the predictive power of the instruments.

When we estimate Equation 1 including province and time fixed effects—and excluding the initial share of manufacturing workers in 1991—we obtain similar results. In particular, the statistical significance of import competition increases for both the female share of employment in manufacturing and the share of women employed in services. As above, the F-Stat remains above 100.

Short run analysis

We also estimate Equation 1 using shorter time differences, specifically five periods with three-year intervals from 2004 to 2019. When controlling for region-by-year fixed effects (to account for region-specific shocks), the initial share of manufacturing workers, and changes in the skilled labor force, we find similar results, albeit of smaller magnitude. The unemployment rate increases across all groups, the share of manufacturing employment slightly declines for the overall working population, and the inactivity rate decreases for both men and women, while no effect on the share of employed in services. These findings remain consistent when using province and year fixed effects instead.

6 Robustness checks

To ensure the robustness of our findings, we carry out a detailed set of additional analyses and alternative specifications.

First, we revisit our measure of import exposure by expanding beyond China to also include imports from Eastern European countries. This adjustment is meant to capture trade pressures stemming from post-enlargement EU dynamics, which could have affected Italian local labor markets. Specifically, we modify Equation 2 and its associated instrument to reflect exports from a group of new EU member states—namely Cyprus, Estonia, Latvia, Lithuania, Malta, Poland, Czech Republic, Slovakia, Slovenia, and Hungary—directed to both Italy and other high-income economies outside the EU.[13] The estimates from this expanded setup are highly consistent with those reported in Section 4, reinforcing the stability of our results to broader definitions of trade exposure.

Second, we incorporate a refinement accounting for bilateral trade flows between Italy and China, following Heckl (2024). We update Equation 2 into Equation 6, which subtracts the change in Italian exports to China from Chinese import growth—normalized by industry output in 1995—to better isolate the competitive effect of imports from China. We instrument using changes in Chinese exports to other non-EU high-income countries.

$$\Delta IP_{j\tau}^{ITA} = \frac{\Delta M_{j\tau}^{ITA} - \Delta X_{j\tau}^{ITA}}{Prod_{j95} + M_{j95} - X_{j95}} \tag{6}$$

where $\Delta X_{j\tau}^{ITA}$ denotes the change in Italian exports to China over a three- or eight-year period. Again, our main findings remain robust and qualitatively unchanged, supporting the validity of our causal interpretation.

Additional robustness:

- Results remain robust even when considering the competition from Chinese supply shock faced by Italian companies in the destination markets of exported products (Citino and Linarello, 2022).

[13] Again, the Countries are: Australia, Canada, Japan, New Zealand, and United States.

- Effects are driven by import competition in sectors that predominantly employed women in 1991, typically in textile industries.

- We refine our analysis by considering aggregates by gender and age group (above or below 40 years), as well as by gender and education level. We find that in provinces more exposed to international trade competition, female unskilled workers and young women experience a larger decrease in manufacturing employment. In addition, young women increase their share of employment in services, while the unemployment rate among young workers (both male and female) rises.

Instrument Validity

In this section, we aim to test the validity of our identification strategy, particularly the exogeneity of the initial industrial share by province. Following the approach proposed by Goldsmith-Pinkham et al. (2020), we estimate the Rotemberg weight for each of the 235 four-digit industries. These weights collectively account for approximately 44% (=0.68/1.53) of the absolute weight in the estimator (considering only positive weights). We examine whether the industrial share by province of the top five sectors is correlated with future labor market shocks (occurring in 2004–2011 and 2012–2019) or with other economic shocks at the provincial level. We do not find any strong statistical relationship between industry shares and labor market outcomes, with the exception of the shoe production sector. However, it is reassuring for our results that manufacturing industries with higher female employment are not the ones with the largest Rotemberg weights.

7 Conclusions

This study highlights the effects of increased import competition—particularly from China—on the Italian labor market. Our findings indicate that both in the short and long run, higher exposure to trade shocks leads to rising unemployment, especially among women, and a decline in manufacturing employment predominantly driven by female job losses. In contrast, inactivity rates decrease among men and show no significant change for women, while some evidence suggests a modest reallocation of female employment toward health and education sectors.

These results align with a growing body of literature documenting the uneven effects of globalization on advanced economies, but add new insight by emphasizing their gendered dimension. The stronger adverse effects on women likely reflect task-based gender segregation, institutional rigidities, and intra-household labor supply constraints. In Italy's segmented labor market—where women are overrepresented in part-time, fixed-term, or care-related jobs—trade shocks may reinforce existing inequalities rather than mitigate them.

Our findings underscore the importance of applying a gender-disaggregated lens when evaluating trade and labor market policies. Targeted interventions—such as active labor market programs, female re-skilling initiatives, and improved childcare support—could help cushion adjustment costs and promote more inclusive labor market transitions.

Finally, future research should explore the long-term implications of trade-induced employment shifts for gender wage gaps, occupational mobility, and educational choices, as well as the role of complementary institutions and policies in shaping these outcomes. As international trade continues to evolve, understanding how it intersects with persistent structural inequalities remains a critical challenge for policymakers and scholars alike.

Appendix 1

A1.1 The case of conditional demand functions

We consider first the Cobb-Douglas production function. As usual, the conditional demand function, for each ℓ, can be written as

$$Q_\ell\left(w\right) Y^{\frac{1}{\overline{\theta}}} = \frac{\theta_\ell}{w_\ell}\left[\prod_{\ell=1}^{\ell=L}\left(\frac{w_\ell}{\theta_\ell}\right)^{\frac{\theta_\ell}{\overline{\theta}}}\right] Y^{\frac{1}{\overline{\theta}}}, \quad \text{with } \frac{\partial Q_\ell\left(w\right)}{\partial w_\ell} = -\frac{\overline{\theta}-\theta_\ell}{\overline{\theta}}\frac{1}{w_\ell}Q_\ell\left(w\right).$$

Consider next the cost minimization problem for the two labor inputs

$$\min_{(Q_M,Q_F)} w_M Q_M + w_F Q_F \quad \textit{subject to} \quad \left[\phi_M Q_M^\beta + \phi_F Q_F^\beta\right]^{\frac{1}{\beta}} = Q_L,$$

with first order conditions

$$\left[\begin{array}{c} w_M - \lambda\phi_M\left[\phi_M Q_M^\beta + \phi_F Q_F^\beta\right]^{\frac{1-\beta}{\beta}} Q_M^{\beta-1} \\ w_F - \lambda\phi_F\left[\phi_M Q_M^\beta + \phi_F Q_F^\beta\right]^{\frac{1-\beta}{\beta}} Q_F^{\beta-1} \\ -\left[\left[\phi_M Q_M^\beta + \phi_F Q_F^\beta\right]^{\frac{1}{\beta}} - Q_L\right] \end{array}\right] = \left[\begin{array}{c} 0 \\ 0 \\ 0 \end{array}\right].$$

Taking the ratio of the first two conditions and simplifying, we obtain $Q_M = \left(\frac{w_M}{\phi_M}\frac{\phi_F}{w_F}\right)^{\frac{1}{\beta-1}} Q_F$. Replacing into the production function, we obtain the conditional demand functions

$$Q_g\left(w, Q_L\right) = \frac{\left(\frac{w_g}{\phi_g}\right)^{\frac{1}{\beta-1}} Q_L(w) Y^{\frac{1}{\theta}}}{\left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{w_F}{\phi_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{1}{\beta}}} \equiv \widetilde{Q}_M\left(w\right) \times Q_L\left(w\right) \times Y^{\frac{1}{\theta}}.$$

Finally, the *unit* cost function is

$$\begin{aligned} w_L\left(.\right) \equiv & \frac{w_M\left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}} + w_F\left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}}}{\left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{w_F}{\phi_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{1}{\beta}}} = \frac{\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{w_F}{\phi_F}\right)^{\frac{\beta}{\beta-1}}}{\left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{w_F}{\phi_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{1}{\beta}}} \\ = & \left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{w_F}{\phi_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{\beta-1}{\beta}}, \end{aligned}$$

with partial derivative

$$\frac{\partial w_L\left(.\right)}{\partial w_g} = \frac{\left(\frac{w_g}{\phi_g}\right)^{\frac{1}{\beta-1}}}{\left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{w_F}{\phi_F}\right)^{\frac{\beta}{\beta-1}}\right)^{\frac{1}{\beta}}}.$$

Let's now compute the terms required to apply the IFT in the main text. Consider

$$\begin{bmatrix} \frac{\partial w_M}{\partial Y} \\ \frac{\partial w_F}{\partial Y} \end{bmatrix} = \frac{-1}{\det D_w \Psi} \begin{bmatrix} \frac{\partial \Psi_F}{\partial w_F} & -\frac{\partial \Psi_F}{\partial w_M} \\ -\frac{\partial \Psi_M}{\partial w_F} & \frac{\partial \Psi_M}{\partial w_M} \end{bmatrix} \begin{bmatrix} \frac{\partial \Psi_F}{\partial Y} \\ \frac{\partial \Psi_F}{\partial Y} \end{bmatrix},$$

For notational convenience, let $S \equiv \left(\phi_M\left(\frac{w_M}{\phi_M}\right)^{\frac{\beta}{\beta-1}} + \phi_F\left(\frac{\phi_F}{w_F}\right)^{\frac{\beta}{\beta-1}}\right)$. Then,

$$\begin{bmatrix} \frac{\partial \Psi_M}{\partial Y} \\ \frac{\partial \Psi_F}{\partial Y} \end{bmatrix} = \begin{bmatrix} \left(\frac{w_M}{\phi_M}\right)^{\frac{1}{\beta-1}} \\ \left(\frac{w_F}{\phi_F}\right)^{\frac{1}{\beta-1}} \end{bmatrix} \frac{Q_L\left(w\right)\overline{Y}^{\frac{1}{\theta}}}{\overline{\theta Y} S^{\frac{1}{\beta}}}$$

and

$$
\begin{aligned}
\frac{\partial \Psi_g}{\partial w_g} &= Q_L(w) Y^{\frac{1}{\theta}} \left(\frac{\partial \widetilde{Q}_g}{\partial w_g} + \frac{\partial w_L}{\partial w_g} \frac{\partial Q_L}{\partial w_L} \widetilde{Q}_g(w) \right) - \frac{\partial H_g(w_g)}{\partial w_g} \\
&= \left(\frac{Q_L(w) Y^{\frac{1}{\theta}}}{(1-\beta) S^{\frac{1+\beta}{\beta}}} \right) \frac{\left(\frac{w_g}{\phi_g}\right)^{\frac{2-\beta}{\beta-1}}}{\phi_g} \left(\phi_{g'} \left(\frac{w_{g'}}{\phi_{g'}} \right)^{\frac{\beta}{\beta-1}} + \phi_g \left(\frac{w_g}{\phi_g} \right)^{\frac{\beta}{\beta-1}} (1-\beta) \left(\frac{\overline{\theta} - \theta_L}{\overline{\theta}} \right) \right) \\
&\quad - \frac{\partial H_g(w_g)}{\partial w_g}
\end{aligned}
$$

$$
\begin{aligned}
\frac{\partial \Psi_g}{\partial w_{g'}} &= Q_L(w) Y^{\frac{1}{\theta}} \left(\frac{\partial \widetilde{Q}_g}{\partial w_{g'}} + \frac{\partial w_L}{\partial w_{g'}} \frac{\partial Q_L(w)}{\partial w_L} \widetilde{Q}_g(w) \right) \\
&= \left(\frac{Q_L(w) Y^{\frac{1}{\theta}}}{(1-\beta) S^{\frac{1+\beta}{\beta}}} \right) \left(\frac{w_F}{\phi_F} \right)^{\frac{1}{\beta-1}} \left(\frac{w_M}{\phi_M} \right)^{\frac{1}{\beta-1}} \left(\beta + (1-\beta) \frac{\theta_L}{\overline{\theta}} \right),
\end{aligned}
$$

where we have exploited the explicit formulas for the various coefficients:

$$
\begin{aligned}
\frac{\partial \widetilde{Q}_g}{\partial w_g} &= \frac{\frac{\phi_{g'}}{\phi_g} \left(\frac{w_g}{\phi_g} \right)^{\frac{2-\beta}{\beta-1}} \left(\frac{w_{g'}}{\phi_{g'}} \right)^{\frac{\beta}{\beta-1}}}{(\beta-1) S^{\frac{1}{\beta}(\beta+1)}}, \\
\frac{\partial \widetilde{Q}_M}{\partial w_F} &= \frac{\partial \widetilde{Q}_F}{\partial w_M} = \frac{\left(\frac{w_M}{\phi_M} \right)^{\frac{1}{\beta-1}} \left(\frac{w_F}{\phi_F} \right)^{\frac{1}{\beta-1}}}{(1-\beta) S^{\frac{1+\beta}{\beta}}}.
\end{aligned}
$$

Applying the IFT formula, a tedious computation shows that

$$
\begin{aligned}
\frac{\partial w_M}{\partial Y} &\equiv \frac{A}{w_F} + \left(\frac{w_M}{\phi_M} \right)^{\frac{1}{\beta-1}} \frac{\partial H_F(w_F)}{\partial w_F}, \\
\frac{\partial w_F}{\partial Y} &\equiv \frac{A}{w_M} + \left(\frac{w_F}{\phi_F} \right)^{\frac{1}{\beta-1}} \frac{\partial H_M(w_M)}{\partial w_M},
\end{aligned}
$$

where

$$
A \equiv \left(\frac{\left(Q_L(w) Y^{\frac{1}{\theta}} \right)^2 \left(\frac{w_F}{\phi_F} \right)^{\frac{1}{\beta-1}} \left(\frac{w_M}{\phi_M} \right)^{\frac{1}{\beta-1}}}{(1-\beta) \overline{\theta Y} S^{1+\frac{1}{\beta}}} \right).
$$

A1.2 The case of perfect competition

The main change is that, now, we must consider the full impact of the external shock on the labor markets, i.e., we are now going to consider the actual demand functions instead of the conditional ones.

Apart from this, the main line of argument is as developed above. The only difference is that now Y is not any longer an exogenous variable, it is instead the value of the supply function. The exogenous trade shock can be identified with a reduction of the output price. Hence,

$$\Psi(w,p) \equiv \begin{bmatrix} \Psi_M(w,p) \\ \Psi_F(w,p) \end{bmatrix} \equiv \begin{bmatrix} \widetilde{Q}_M(w)Q_L(w)Y(w,p)^{\frac{1}{\bar{\theta}}} - H_M(w_M) \\ \widetilde{Q}_F(w)Q_L(w)Y(w,p)^{\frac{1}{\bar{\theta}}} - H_F(w_F) \end{bmatrix} = \begin{bmatrix} 0 \\ 0 \end{bmatrix}.$$

The wage adjustment to the exogenous trade shock is $\nabla_p w = -[D_w\Psi]^{-1}\nabla_p\Psi$, and

$$\begin{bmatrix} \frac{\partial w_M}{\partial p} \\ \frac{\partial w_F}{\partial p} \end{bmatrix} = \frac{-1}{\det D_w\Psi} \times \left[D_w\Psi^{-1}\right] \begin{bmatrix} Q_M(w)Y(w,p)^{\frac{1}{\bar{\theta}}}\frac{1}{(1-\bar{\theta})p} \\ Q_F(w)Y(w,p)^{\frac{1}{\bar{\theta}}}\frac{1}{(1-\bar{\theta})p} \end{bmatrix},$$

since $\frac{\partial Y}{\partial p} = \frac{\bar{\theta}}{(1-\bar{\theta})p}Y(w,p)$. Moreover,

$$D_{w_g}\Psi_g = \frac{\partial Q_g}{\partial w_g}Y(w,p)^{\frac{1}{\bar{\theta}}} + Q_g(.)\frac{1}{\bar{\theta}Y(w,p)}\frac{\partial Y}{\partial w_L}\frac{\partial w_L}{\partial w_g} - \frac{\partial H_g}{\partial w_g}.$$

Consider the effect of the external shock on the labor market. A simple computations, following basically the same steps as above, shows that the first order effect - due to the direct impact of the change in the output prices - is gender neutral, while the differential impacts on the two labor markets are asymmetric due to the differences in the wage elasticities of the labor supplies. The relative size has the same properties discussed above.

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