Discussion Paper Series

RIEB

Kobe University

DP2025-04

Trade Liberalization, Wage, and Unemployment: Theory and Evidence from Chile

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Revised April 23, 2026

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Data availability statement

The data that support the findings of this study are available in the Supplementary material. More detailed data are available upon request.

Funding statement

This work was supported by the JSPS KAKENHI [Grant numbers 20K13482, 23K01401, and 24K22610].

Conflict of interest disclosure

There are no potential conflicts of interest to declare.

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Acknowledgments

The authors are grateful to Hideki Esho, Hiroshi Goto, Nobuaki Hamaguchi, Isao Kamata, Keisuke Kondo, Mikio Kuwayama, Hiroyuki Nishiyama, and Takahiro Sato for their insightful comments and suggestions, and Yoshihiko Hashiguchi and Junsong Shi for their assistance with this research.

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We developed a theoretical model that incorporates a fair wage model into a trade model with heterogeneous firms to analyze the effects of trade liberalization on wages and unemployment rates. By introducing a unique setting for the determinant of average wage and different fixed production costs between exporting and domestic firms, the model predicts that lower trade costs lead to lower prices of intermediate inputs, which reduces the production costs for final-goods firms and stimulates their labor demand. Consequently, the model predicts that trade liberalization leads to higher average productivity, thereby raising wages and reducing unemployment rates. The empirical analysis employing exogenous variations in exposures to regional trade agreements (RTAs) across local labor markets in Chile from 2000 to 2006 supported our theoretical predictions. Exploiting industrial variations in tariff reductions derived from the RTA schemes, initial differences in industrial compositions across local labor markets, and the lack of inter-local labor mobility, we found that a reduction in input tariff rates expectedly increased local wages and decreased unemployment rates. These findings were robust to the inclusion of various controls, the control for the endogeneity of the input tariffs, and the use of different regional units of analysis.

Keywords: firm heterogeneity; unemployment rate; fair wage; regional trade agreements; local labor markets, Chile

JEL classification codes: F12; F16; F66; J3; J64; R23

1. Introduction

Globalization has affected labor markets in developing and emerging countries through various channels. An abundant body of literature has empirically and theoretically explored its impact, primarily focusing on the impact of trade liberalization on wages and wage inequality. Based on a comprehensive survey of the literature, Goldberg and Pavcnik (2007) concluded that trade liberalization was accompanied by an increase in wage inequality in most developing countries during the 1980s and 1990s; this opposes the traditional Heckscher-Ohlin theory, which predicts that trade liberalization would decrease wage inequality between skilled and unskilled workers in developing countries well-endowed with unskilled labor. However, theoretical developments have reconciled the contradiction. In particular, theoretical studies focusing on heterogeneous firms show that trade liberalization leads to an increase in wage inequality between workers employed in high-productivity firms and low-productivity firms (Helpman, Itskhoki, & Redding, 2010), as well as workers employed in exporting or importing firms and domestically oriented firms (Amiti & Davis, 2011).[1]

However, despite some exceptions—such as Helpman and Itskhoki (2010); Felbermayr, Prat, and Schmerer (2011); and de Pinto and Michaelis (2014, 2019)—few studies have theoretically analyzed the impact of trade liberalization on unemployment. Considering that unemployment is likely to disproportionately affect lower-income people and, thus, have significant impacts on income inequality, a theoretical and empirical analysis identifying the impact of trade liberalization on unemployment is necessary. Given that unemployment is of equal importance in terms of its impact on labor markets, there is a particular need for studies that incorporate theoretical models, such as fair wage hypothesis, which allows the analysis of the impacts of trade liberalization on both unemployment and wages under the same theoretical framework.

Helpman and Itskhoki (2010) developed Melitz's (2003) model with search and matching frictions in a two-country and two-sector framework. They argued that while trade liberalization improves welfare for both countries, it can increase the overall unemployment rate if labor is reallocated toward industries with higher sectoral unemployment rates. Similarly, Felbermayr et al. (2011) developed Melitz's (2003) model with search and matching frictions in a framework of symmetric countries, arguing that trade liberalization decreases the unemployment rate and increases wages as long as it leads to an improvement in average productivity. Using simulation analysis based on Felbermayr et al.'s (2011) model, Capuano and Schmerer (2015) assessed the effects of different wage-setting regimes on the relationship between trade and unemployment, without providing theoretical prediction based on comparative statics. However, these studies only capture the effects of trade liberalization on frictional unemployment, where workers must spend time and resources to find new jobs during the reallocation of labor.

In this respect, two studies are closely related to our study because they incorporated the fair wage hypothesis into a trade model with heterogeneous firms. First, Egger and Kreickemeier (2009) showed that trade liberalization leads to greater wage inequality between workers employed in higher-

[1] Helpman et al. (2010) developed Melitz's (2003) model with search and matching frictions. They additionally showed that trade liberalization can decrease wage inequality in the long run. Amiti and Davis (2011) also developed Melitz's (2003) model with trade in intermediate inputs.

and lower-productivity firms as well as higher unemployment rates. However, they did not analytically examine the impact of trade liberalization on unemployment rates under the assumption of different fixed production costs.[2] Second, Nishiyama and Gintani (2021) theoretically analyzed the impact of trade liberalization on wages, unemployment rates, and welfare through price fluctuations. However, they did not provide clear predictions of the effects of trade liberalization on unemployment rates and welfare; rather, they showed that these effects depend on the degree of trade costs. Moreover, they did not conduct a formal empirical analysis.

Additionally, de Pinto and Michaelis (2014) developed Melitz's (2003) model with a union bargaining model and worker heterogeneity, arguing that the impact of trade liberalization on unemployment rates depends on the distribution of worker abilities and labor market institutions. Similarly, de Pinto and Michaelis (2019) developed Melitz's (2003) model with union heterogeneity and a union bargaining model, arguing that the impact of trade liberalization on unemployment rates depends on the bargaining power of trade unions. However, they did not conduct a formal empirical analysis, making it unclear whether their theoretical predictions were supported.

Therefore, this study aimed to develop a theoretical model that integrates the fair wage hypothesis into Melitz's (2003) framework and empirically test the theoretical prediction. Specifically, by introducing a unique setting for the determinant of the average wage, our model provided clear predictions of the effects of trade liberalization on wages and unemployment rates under the assumption of different fixed production costs. Consequently, our model predicts that lower trade costs lead to higher average productivity, thereby increasing wages and reducing unemployment rates.

We tested our model's predictions by employing exogenous variations in exposure to regional trade agreements (RTAs) across local labor markets in Chile from 2000 to 2006. In particular, by exploiting industrial variations in tariff reductions derived from the RTA scheme, initial differences in industrial compositions across local labor markets, and the lack of inter-local labor mobility, we identified the causal effects of trade liberalization on wages and unemployment rates.

[2] Egger and Kreickemeier (2009) argued that trade liberalization increases the unemployment rate under the assumption of identical fixed costs across all firms. However, their model has difficulty in analytically deriving the results under the assumption of heterogeneous fixed costs (see Proposition 2 of Egger & Kreickemeier, 2009, p. 205; see also footnote 28, p.210).

The empirical analysis of this study is closely related to two previous studies that analyzed Chile during the same period: César, Falcone, and Gasparini (2021) and Murakami (2021). However, César et al. (2021) focused on trade shocks derived from Chinese import competition and demand for exports and found no significant effects on the unemployment rates. Furthermore, the effects of RTAs were beyond the scope of their analysis. Although Murakami (2021) examined the effects of tariff reductions derived from the RTA schemes on wages and wage inequality, he did not analyze the effect on unemployment. Moreover, these studies were not based on any particular theoretical predictions. Consequently, aligned with the clear predictions of our theoretical model, we provided new evidence on the effects of trade liberalization on both wages and unemployment rates, which is a novel contribution of this study to the literature.

The remainder of this paper is organized as follows. Section 2 presents the theoretical model. Section 3 presents the empirical specifications. Section 4 explains the data sources and definitions of the variables used and presents the descriptive statistics. Section 5 presents the estimation results, and the final section concludes the paper.

2. Theoretical framework

We developed our model based on Egger and Kreickemeier (2009), who incorporated a fair wage model into Melitz's (2003) framework. Egger and Kreickemeier (2009) focused on the impact of trade liberalization on wage inequality. However, due to the structure of their model, they did not analytically examine the impact of trade liberalization on the unemployment rate under the assumption of different fixed costs. By changing the average wage setting, our model can predict the effects of trade liberalization on unemployment rates and wages.

We assume that the world is composed of $n + 1$ symmetric countries. Workers are the only factor of production in inelastic labor supply L for each country and are immobile across countries. The final good is non-tradable, whereas intermediate goods are tradable.

2-1. Final-goods firms

The final good Y is produced under perfect competition using all available intermediate goods as inputs. The final-goods firm maximizes its profit, subject to the following production function.[3]

[3] The methodology for Y is based on Blanchard and Giavazzi (2003). For a more in-depth discussion of

$$Y = \left(M^{-\frac{1}{\sigma}} \int_{\omega \in \Omega} q_D(\omega)^{\frac{\sigma-1}{\sigma}} d\omega + nM^{-\frac{1}{\sigma}} \int_{\omega \in \Omega} q_{EX}(\omega)^{\frac{\sigma-1}{\sigma}} d\omega \right)^{\frac{\sigma}{\sigma-1}}, \tag{1}$$

where M is the mass of available intermediate goods; ω indexes varieties; Ω is the number of differentiated goods sets; $q_i(\omega)$ is the optimal quantity demanded of intermediate goods (i = D; EX); and $\sigma > 1$ is the elasticity of substitution among intermediate goods. We assume that the intermediate goods are differentiated. The price index P of Y is

$$P = \left(M^{-1} \int_{\omega \in \Omega} p_D(\omega)^{1-\sigma} d\omega + nM^{-1} \int_{\omega \in \Omega} p_{EX}(\omega)^{1-\sigma} d\omega \right)^{\frac{1}{1-\sigma}}, \tag{2}$$

where $p_i(\omega)$ is the price of variety ω. Following Egger and Kreickemeier (2009) and Felbermayr et al. (2011), we set $P = 1$, considering Y as a numeraire good. The individual demand quantity is given by

$$q_i(\omega) = \left(p_{i(\omega)}\right)^{-\sigma} \frac{Y}{M}, (i = D, EX). \tag{3}$$

2-2. Intermediate-goods firm

Firms producing intermediate goods aim to maximize their profits:

$$\max_{p_D(\varphi), p_{EX}(\varphi)} p_D(\varphi) q_D(\varphi) - w(\varphi) l_D(\varphi) - f_D + I(\varphi) n [p_{EX}(\varphi) q_{EX}(\varphi) - w(\varphi) l_{EX}(\varphi) - f_{EX}], \tag{4}$$

subject to the production functions of an intermediate goods sector given by

$$q_D(\varphi) = \varphi \varepsilon l_D(\varphi), \ \tau q_{EX}(\varphi) = \varphi \varepsilon l_{EX}(\varphi), \tag{5}$$

where $\varphi (> 0)$ is firm productivity; $w(\varphi)$ is the wage; f_i ($f_{EX} > f_D$) is a fixed production cost; $I(\varphi)$ is an indicator function that takes the value of 1 if a firm engages in exports and 0 otherwise; l_i is the unit of labor input; $\tau (> 1)$ is iceberg trade cost; and ε is the level of effort. The impact of trade liberalization on unemployment crucially depends on the relative magnitudes of domestic and foreign fixed costs. Allowing for $f_{EX} > f_D$ enables a more comprehensive analysis of how trade costs affect firms' entry and exit and labor market adjustments. The iceberg trade cost implies that only a fraction of the shipped goods reaches the foreign market. Specifically, for each unit that arrives abroad, τ units must be exported. Following Egger and Kreickemeier (2009), l_i is the efficiency level such that it corresponds to the number of workers and their effort level ε. We obtain the optimal price

$$p_D(\varphi) = \left(\frac{\sigma}{\sigma-1}\right) \frac{w(\varphi)}{\varphi \varepsilon}, p_{EX}(\varphi) = \tau \left(\frac{\sigma}{\sigma-1}\right) \frac{w(\varphi)}{\varphi \varepsilon}. \tag{6}$$

The respective profit functions of domestic and exporting firms are

$$\pi_D(\varphi) = \frac{r_D(\varphi)}{\sigma} - f_D, \ \pi_{EX}(\varphi) = \frac{r_{EX}(\varphi)}{\sigma} - f_{EX}. \tag{7}$$

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this approach, see Egger and Kreickemeier (2009) and Felbermayr et al. (2011).

Thus, the aggregate revenue is

$$R(l;\varphi) = p_D(\varphi)q_D(\varphi) + I(\varphi)n[p_{EX}(\varphi)q_{EX}(\varphi)] = \left(\frac{Y}{M}\right)^{\frac{1}{\sigma}}\left(\varphi l_D(\varphi)\right)^{\frac{\sigma-1}{\sigma}}(1 + I(\varphi)n\tau^{1-\sigma}). \quad (8)$$

2-3. Setting of the labor market

We assume that unemployment occurs based on the fair wage hypothesis. We denote $\widehat{w}$ as the fair wage and $\overline{w}$ as the average wage. Actual wages w depend on the level of effort ε. Following Akerlof and Yellen (1990), we have $\varepsilon = \min(w/\widehat{w}, 1)$. If $w \geq \widehat{w}$, we have $\varepsilon = 1$, indicating that workers offer their maximum efforts. If $w < \widehat{w}$, we have $\varepsilon = w/\widehat{w}$, indicating the decline of workers' efforts and productivity. Thus, given that profit-maximizing firms have no incentive to pay less than the fair wage, we assume that firms pay fair wages and $\varepsilon = 1$. The assumption eliminates the distinction between the number of workers and the efficiency level that depends on the worker's effort level ε. Although Egger and Kreickemeier (2009) assumed that the average wage is indexed to firms' paid wage $\overline{w} = w$, we assume that the average wage is equal to the aggregate wage paid by the operating firms in the economy. The setting of the average wage allows our model to incorporate shifts in market composition—such as the expansion of exporting firms and the exit of domestic firms—into the determinants of workers' fair wages and unemployment rates. Thus, we define the average wage as

$$\overline{w} = \frac{1}{M}\left(\int_{\varphi_D^*}^{\infty} w(\varphi)\, M_D(\varphi)\mu_D(\varphi)d\varphi + n\int_{\varphi_{EX}^*}^{\infty} w(\varphi)\, M_{EX}(\varphi)\mu_{EX}(\varphi)d\varphi\right). \quad (9)$$

We denote the rent-sharing parameter by ($0<\psi<1$) and unemployment rate by u. In line with Akerlof (1982) and Danthine and Kurmann (2006), the fair wage is determined as the geometric average of the firm's productivity and the average wage per worker including the unemployed. Thus, we have the fair wage as follows:

$$\widehat{w}(\varphi) = \varphi^{\psi}[(1-u)\overline{w}]^{1-\psi}. \quad (10)$$

Assuming the productivity of any two firms as $\varphi_1 \neq \varphi_2$, we obtain the ratio of wages, optimal prices, optimal quantity demanded, imports, and labor input:

$$\frac{w(\varphi_1)}{w(\varphi_2)} = \left(\frac{\varphi_1}{\varphi_2}\right)^{\psi}, \frac{p_i(\varphi_1)}{p_i(\varphi_2)} = \left(\frac{\varphi_1}{\varphi_2}\right)^{\psi-1}, \frac{q_i(\varphi_1)}{q_i(\varphi_1)} = \left(\frac{\varphi_1}{\varphi_2}\right)^{\sigma(1-\psi)}, \frac{r_i(\varphi_1)}{r_i(\varphi_2)} = \left(\frac{\varphi_1}{\varphi_2}\right)^{\xi}, \frac{l_i(\varphi_1)}{l_i(\varphi_2)} = \left(\frac{\varphi_1}{\varphi_2}\right)^{\xi-\psi}, \quad (11)$$

where $\xi = (\sigma-1)(1-\psi)$.

We assume an unbounded pool of prospective entrants to the intermediate goods market. The firms are identical prior to entry. To enter, they must pay a free entry cost f_e. Once f_e is paid, a firm draws its productivity $\varphi > 0$ from a common distribution $g(\varphi)$ that has a continuous cumulative distribution

$G(\varphi)$. When the firms produce, they then face a constant probability $\delta \in (0,1)$ of a bad shock in every period. We assume the following Pareto distribution:

$$G(\varphi) = 1 - (\varphi)^{-k},\ g(\varphi) = k\varphi^{-(k+1)},\ k > (\sigma - 1). \tag{12}$$

Thus, the share of operating firms is given by

$$\mu_i(\varphi) = \begin{cases} \frac{g(\varphi)}{1-G(\varphi_i^*)} = \frac{k}{\varphi}\left(\frac{\varphi_i^*}{\varphi_i}\right)^k & if\ \varphi \geq \varphi_i^*, \\ & otherwise \end{cases} \tag{13}$$

where k (>1) is the shape parameter; lower k means greater dispersion in φ. The number of firms operating in the economy is given by $M = M_D + nM_{EX} = M_D(1 + n\chi)$. The share of exporting firms can be expressed as $\chi \equiv \frac{[1-G(\varphi_{EX}^*)]}{[1-G(\varphi_D^*)]} = \tau^{\sigma-1}\frac{f_{EX}}{f_D}$. From Equation (7), a zero-profit condition is

$$\pi(\varphi_i^*) = 0\ or\ r_i(\varphi_i^*) = \sigma f_i\,. \tag{14}$$

We define a weighted average of the firm productivity levels $\tilde{\varphi}$ such that $q_D(\tilde{\varphi}) = Y/M.$ Hence, Equation (3) and $P = 1$ implies $p_D(\tilde{\varphi}) = 1.$

Next, rewriting Equation (2), we obtain

$$P = \left\{M^{-1}\int_{\varphi_D^*}^{\infty} p_D(\varphi)^{1-\sigma} M_D\mu_D(\varphi)d\varphi + nM^{-1}\int_{\varphi_{EX}^*}^{\infty} p_{EX}(\omega)^{1-\sigma} M_{EX}\mu_{EX}(\varphi)d\varphi\right\}^{\frac{\sigma}{\sigma-1}}, \tag{15}$$

where $\tilde{\varphi}_i^{\xi} \equiv \int_{\varphi_i}^{\infty} \varphi^{\xi}\mu_i(\varphi)d\varphi$. Rewriting Equation (15), we obtain

$$\tilde{\varphi} = \left[\frac{1}{1+n\chi}\left\{\tilde{\varphi}_D^{\xi} + n\chi\tau^{1-\sigma}\tilde{\varphi}_{EX}^{\xi}\right\}\right]^{\frac{1}{\xi}}. \tag{16}$$

Hence, $\tilde{\varphi}$ is obtained as the weighted average of the productivity of domestic and exporting firms:

$$\tilde{\varphi} = \left[\frac{k}{k-\xi}\cdot\frac{1+n\chi\left(\frac{f_{EX}}{f_D}\right)}{1+n\chi}\right]^{\frac{1}{\xi}}\varphi_D^*. \tag{17}$$

Additionally, we find the following relation between $\tilde{\varphi}_i$ and φ_i^*:

$$\tilde{\varphi}_i = \left(\frac{k}{k-\xi}\right)^{\frac{1}{\xi}}\varphi_i^*. \tag{18}$$

where $0 < \xi < k$.

A firm faces a bad shock with some probability $\delta \in (0,1)$ in every time point.[4] In the steady state, assuming no discount rate, the value function of each firm is

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[4] δ is exogenous and does not depend on productivity φ. We followed Egger and Kreickemeier (2009, p. 195) for the timing of a firm's decision.

$$V(\varphi) = max\{0,\ \sum_{T=0}^{\infty}(1-\delta)^{T}\pi\,(\varphi)\} = \left\{0, \sum_{T=0}^{\infty}\frac{\pi(\varphi)}{\delta}\right\}. \tag{19}$$

Firms enter as long as their average profit covers the entry costs. From Equations (14) and (19) and $\overline{\pi} = \pi(\tilde{\varphi})$, the free entry condition is

$$f_e = \int_0^{\infty} v(\varphi)g(\varphi)d\varphi \Leftrightarrow \bar{\pi} = \delta f_e(\varphi_D^*)^k. \tag{20}$$

The average profit $\overline{\pi}$ is equal to the profit of the whole firms divided by the number of firms. Thus, the zero-profit cutoff condition is

$$\overline{\pi} = \frac{1}{M}\left(\int_{\varphi_D^*}^{\infty}\pi_D(\varphi)M_D\mu_D(\varphi) + n\int_{\varphi_{Ex}^*}^{\infty}\pi_{EX}(\varphi)M_{EX}\mu_{EX}(\varphi)\right) \Leftrightarrow \overline{\Pi} = \frac{\xi(f_D+n\chi f_{EX})}{(1+n\chi)(k-\xi)}. \tag{21}$$

From Equations (20) and (21), the cutoff productivity for domestic firms is

$$\varphi_D^* = \left(\frac{\xi(f_D+n\chi f_{EX})}{(1+n\chi)(k-\xi)\delta f_e}\right)^{\frac{1}{k}}. \tag{22}$$

Thus, we obtain the average productivity as follows:

$$\tilde{\varphi} = \left(\frac{k}{k-\xi}\right)^{\frac{1}{\xi}}\left(\frac{\xi(f_D+n\chi f_{EX})}{(1+n\chi)(k-\xi)\delta f_e}\right)^{\frac{1}{k}}\left[\frac{1+n\tau^{\frac{-k}{1-\psi}}(f_{Ex}/f_D)^{\frac{\xi-k}{\xi}}}{1+n\tau^{\frac{-k}{1-\psi}}(f_{Ex}/f_D)^{\frac{-k}{\xi}}}\right]^{\frac{1}{\xi}}. \tag{23}$$

From Equation (6) and $p_D(\tilde{\varphi}) = 1$, we derive workers' wages:

$$w(\tilde{\varphi}) = \rho\tilde{\varphi}, \tag{24}$$

where $\rho = \frac{\sigma-1}{\sigma}$. From Equations (10), (23) and (24), we derive the unemployment rate as

$$u = 1 - \rho^{\frac{\psi}{\psi-1}}\left(\frac{\tilde{\varphi}}{\overline{w}}\right), \tag{25}$$

where $\left(\frac{\tilde{\varphi}}{\overline{w}}\right) \equiv \rho\chi^{\frac{1}{\xi}}\left[\frac{k}{k-\xi}\cdot\frac{1+n\chi\left(\frac{f_{EX}}{f_D}\right)}{1+n\chi}\right]^{-\frac{\psi}{\xi}}$. We call this ratio labor profitability. From Equation (25), we find that the unemployment rate is a function of market competitiveness and labor profitability. The labor demand of domestic firm is $l_D(\varphi) = \left(\frac{\varphi}{\tilde{\varphi}}\right)^{\xi-\psi} l_D(\tilde{\varphi})$.[5] We can derive $q_D(\tilde{\varphi})$ from $r_D(\tilde{\varphi}_D) = p_D(\tilde{\varphi}_D)q_D(\tilde{\varphi}_D)$ and the ratio of optimal quantity demanded (11) as follows:

$$q_D(\tilde{\varphi}) = \frac{\sigma k f_D}{k-\xi}\frac{\left(1+n\chi\frac{f_{EX}}{f_D}\right)}{(1+n\chi)}, \tag{26}$$

[5] See Appendix A1 in the Supplemental material for the derivation of labor demand $l_D(\tilde{\varphi}), l_D(\varphi_D^*), l_{EX}(\tilde{\varphi})\ and\ l_{EX}(\varphi_{EX}^*)$.

2-4. Aggregation

The aggregate employment, $(1-u)L$, can be determined as follows:

$$(1-u)L = \int_{\varphi_D^*}^{\infty} l_D(\varphi)\, M_D(\varphi)\mu_D(\varphi)d\varphi + n\int_{\varphi_{EX}^*}^{\infty} l_{EX}(\varphi)\, M_{EX}(\varphi)\mu_{EX}(\varphi)d\varphi. \tag{27}$$

The number of firms is

$$M = (1-u)L\frac{k+\psi-\xi}{k}(1+n\chi)[l_D(\varphi_D^*) + l_{EX}(\varphi_{EX}^*)]^{-1}. \tag{28}$$

We can derive the rest of aggregate variable from $q_D(\tilde{\varphi}) = Y/M$,

$$Y = M\, q_D(\tilde{\varphi}). \tag{29}$$

2-5. Comparative statics

As in Melitz (2003), our model predicts that trade liberalization (a decrease in trade cost τ) leads to an increase in φ_D^* and a decrease in φ_{EX}^*:

$$\frac{\partial\varphi_D^*}{\partial\tau} < 0, \frac{\partial\varphi_{EX}^*}{\partial\tau} > 0\,, \frac{\partial q_D(\tilde{\varphi})}{\partial\tau} < 0, \frac{\partial\tilde{\varphi}}{\partial\tau} < 0, \frac{\partial w(\tilde{\varphi})}{\partial\tau} < 0, \frac{\partial u}{\partial\tau} > 0. \qquad (30)[6]$$

When trade costs decrease, the productivity cutoff for domestic firms φ_D^* increases, leading to the exit of low-productivity firms. Meanwhile, the decrease in the cutoff for exporting firms φ_{EX}^* allows previously non-exporting firms to enter into export markets, increasing the number of exporting firms and further intensifying market competition. Consequently, the firm-level production $q_D(\tilde{\varphi})$ increases, and the average productivity of the economy improves. As wages are proportional to the average productivity as shown in Equation (24), the resulting productivity improvement translates into a higher average wage. Furthermore, this productivity improvement leads to a decrease in unemployment rate, as shown in Equation (25). Therefore, trade liberalization leads to increased wages and reduced unemployment.

Proposition: The impacts of trade liberalization on wage and unemployment rate:

Lower trade costs reallocate resources to productive firms, leading to higher average productivity, thereby raising wages and reducing unemployment.

[6] For more details on the comparative statics on the wage and unemployment, see Appendix A.2.1 in the Supplemental material. Additionally, it is possible to show the comparative statics on the impact of rent sharing on each variable. For more details, see Appendix A.2.2.

3. Empirical specification

As explained in Section 2, our theoretical model predicts that tariffs on intermediate goods (i.e., input tariffs) affect wages and unemployment rates. Following Chor and Li (2024), Erten, Leight, and Tregenna (2019), McCaig (2011), and Topalova (2010), we employed regional variations in the exposure to RTAs in Chile as our fundamental identification strategy. In addition to the application of a uniform tariff to almost all products, Chile enacted RTAs with its major trading partners in the early 2000s; free trade agreements (FTAs) with the European Union, the United States of America, and the Republic of Korea entered into force in 2003, 2004, and 2006, respectively. Consequently, the applied tariff rates levied in each industry diverged from the uniform most-favored-nation (MFN) rate. Given that regional exposure to these tariff changes depends on the initial industrial composition, the enforcement of the RTAs provides exogenous variations. Thus, we constructed a measure of the regional exposure to RTAs in Chile from 2000 to 2006. Furthermore, following César et al. (2021), we defined the regional unit of the analysis as functional labor market areas (FLMAs) proposed by Casado-Díaz, Martínez-Bernabéu, and Rowe (2017) and Rowe (2017) instead of administrative units.

Thus, we constructed the input tariff rate τ_{rt} in FLMA *r* in year *t* as follows:

$$\tau_{rt} = \sum_j Emshare_{jr}^{1998} \tau_{jt}, \tag{32}$$

where $Emshare_{jr}^{1998}$ is the employment share in industry *j* and FLMA *r* in 1998, and τ_{jt} is the input tariff rate in industry *j* in year *t*. The input tariff rate is given by

$$\tau_{jt} = \sum_k s_{jk}^{1996} \tau_{kt}^{O}, \tag{33}$$

where s_{jt} is the share of inputs purchased from industry *k* in the total inputs including non-traded inputs of industry *j* in 1996, and τ_{kt}^{O} is the output tariff rate of industry *k* measured by the applied tariff rates on final goods.[7] Thus, we used the time-invariant employment and input shares prior to the enforcement of major FTAs in the 2000s. Consequently, our identification strategy employed the combination of the two Bartik-style shift-share variables (Bartik, 1991).

Therefore, we estimated the following equation for FLMA *r* in year *t*:

[7] We assigned non-traded sectors a zero tariff. Thus, the sum of exposure shares varied across industries, which corresponds to the "incomplete shares" case recommended by Borusyak, Hull, and Jaravel (2022).

$$y_{rt} = \alpha_0 + \beta\tau_{rt-1} + \alpha_r + \alpha_t + \varepsilon_{rt}, \tag{34}$$

where $y_{rt} \in \{wp_{rt}, u_{rt}\}$: wp_{rt} and u_{rt} are local wage premium and unemployment rate, respectively, for FLMA *r* in year *t*; α_r is FLMA fixed effects; α_t is year fixed effects; and ε_{rt} is the error term. The input tariff is lagged by one year to address potential endogeneity.

However, it is possible that the initial employment share can be endogenous. Thus, following César et al. (2021), Chor and Li (2024), McCaig (2011), and Topalova (2010), we included initial local characteristics at the FLMA level interacted with year fixed effects as additional controls in Equation (34) when being the unemployment rate as the dependent variable.[8] The initial local characteristics in 1998 included the average years of schooling, the employment shares at a more aggregate level than those used in the construction of the input tariff rates in Equation (32)[9] (i.e., the shares of workers in agriculture, mining, and manufacturing, with non-traded sectors being the omitted category), the urban share of the population, and the share of formal workers measured by those with an indefinite-term contract. If the dependent variable in Equation (34) is the unemployment rate, following César et al. (2021), we estimated the equation using weighted least squares (WLS) with the employment share of each FLMA in 1998 as the weights. The standard errors of Equation (34) were clustered at the FLMA level irrespective of the dependent variables.

As for wages, following Kovak (2013), we directly controlled for the observable workers' characteristics for each year. In other words, we used local wage premiums rather than average wages as the dependent variable in Equation (34). Thus, the local wage premiums were defined as wage differentials attributable to workers' local affiliations after controlling for other observable workers' characteristics. Therefore, we estimated them separately from the following wage equation for each year:

$$\ln w_{ijrt} = \boldsymbol{X}_{it}'\boldsymbol{\gamma} + I_j + \textstyle\sum_r FLMA_{irt} * wp_{rt} + e_{ijrt}, \tag{35}$$

where w_{ijrt} is log hourly wages deflated by the national consumer price index (CPI) (December 2008 = 1)[10] of worker *i* affiliated with industry *j* and living FLMA *r* in year *t*; $\boldsymbol{X}$ represents the vector of

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[8] Thus, we assume the conditional exogeneity of initial shares: $\mathrm{E}(Emshare_{jr}^{1998}\varepsilon_{rt}|\alpha_r, \alpha_t, W_r^{1998} \times \alpha_t) = 0$, where W_r^{1998} indicates initial local characteristics in 1998.

[9] The inclusion of the employment shares at the same level as used for Equation (32) will cause multicollinearity with the input tariff rate.

[10] We sourced the data from the Central Bank of Chile (http://www.bcentral.cl/estadisticaseconomicas/

variables at the individual level including years of education, years of potential labor market experience (age–years of schooling–6), its squared term, a dummy for head of the household, a dummy for married workers, an urban dummy, a dummy for formal workers, and dummy for male workers; I_j is industry fixed effects; $FLMA_{irt}$ is a dummy taking a value of 1 if worker *i* lives in FLMA *r* in year *t*; its coefficient wp_{rt} represents the local wage premium for FLMA *r* in year *t*; and e_{ijrt} is the error term.

In Equation (35), we normalized the local wage premiums to express them as deviations from the employment-share-weighted average wage premiums, so that the weighted average of the normalized wage premiums is zero. Following Kovak (2013), we calculated the normalized local wage premiums and their exact standard errors using Haisken-DeNew and Schmidt's (1997) two-step restricted least squares procedure. Given that the local wage premiums were estimated, we estimated Equation (34) using WLS with the inverse of the standard errors as the weights.

However, in addition to the initial employment share, the nationwide industry-level input tariff rate τ_{jt} also can be endogenous. This is because the industry-level applied tariff rates on final goods τ_{kt}^{O} were determined by exceptions to tariff reductions or eliminations in RTAs as well as the choice of partners with which Chile enacts RTAs. Thus, it is possible that time-variant industry characteristics (e.g., lobbying power) are correlated with FLMA-level input tariff rates, unemployment rates, and wage premiums through the different industry compositions of the FLMAs.[11]

To address this concern, following Murakami (2021), we proposed the following instrumental variable $IV\tau_{jt}$ for the input tariff rate τ_{jt} for industry *j* in year *t*:

$$IV\tau_{jt} = \sum_k s_{jk}^{1996} MFN\tau_t^{O}, \tag{36}$$

where $MFN\tau_t^{O}$ is the uniform MFN tariff rate on final goods in year *t*. The uniform MFN tariff rates were predetermined to be reduced by one percentage point in January each year from 11% by Law No. 19.589 on November 14, 1998 (Ministerio de Hacienda, 1998). Thus, they became 10% in 1999, 7% in 2002, and 6% in 2005. Therefore, there was no room for political lobbying in these reductions. Consequently, the variations of $IV\tau_{jt}$ across industries depends on only their initial input shares

series-indicadores/index_p.htm, accessed on January 1, 2015).

[11] For example, industries with lobbying power for protection were likely to have higher tariffs and wages in a given year. As a result, it is possible that regions with a higher share of such industries had higher tariffs and wages.

purchased from traded sectors (note that we assigned non-traded sectors a zero tariff).[12] Thus, the instrumental variable for the input tariff rate for FLMA *r* in year *t* was constructed as

$$IV\tau_{rt} = \sum_j Emshare_{jr}^{1998} IV\tau_{jt}. \tag{37}$$

4. Data and descriptive statistics

4-1. Data used

We sourced the data on individuals' characteristics, including information on their geographical locations, from the National Socioeconomic Characterization Survey (*Encuesta de Caracterización Socioeconómica Nacional*; CASEN) for 2000, 2003, and 2006.[13] The CASEN survey is a cross-sectional household survey conducted every two or three years by the Ministry of Social Development and Family of Chile (formerly the Ministry of Social Development and the Ministry of Planning and Cooperation). The survey provides detailed information on demographic characteristics, education, health, housing, employment, and income sources. We limited the sample to individuals aged 15–65 years. We constructed the FLMA-level variables using expansion weights for a given year.

Casado-Díaz et al. (2017) and Rowe (2017) originally classified 304 communes (*comunas*, the smallest administrative unit in Chile) into 65 FLMAs based on a grouping evolutionary algorithm using commuting data from the Chilean Internal Migration (CHIM) database for 1982, 1992, and 2002. They argued that FLMAs are more appropriate for defining local labor markets than administrative units such as provinces. However, the CHIM database provided data for only 304 communes, which differ from the 346 communes on which the 2006 CASEN survey was based. For the remaining 42 communes, we assigned existing FLMA codes if a commune became independent from another commune whose FLMA code was uniquely identified or was adjacent to other communes that were geographically contiguous and shared the same FLMA code. However, the FLMA codes of several communes were not uniquely

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[12] Thus, although the instrumental variable is also a Bartik-style instrument, the shift portion is indexed by only *t*.

[13] We sourced the data from the Ministry of Social Development (http://observatorio.ministeriodesarrollosocial.gob.cl/casen-multidimensional/casen/basedatos.php, accessed on June 6, 2018). We additionally used the 1998 survey for the initial employment share and characteristics.

identified. In this case, these communes were assigned multiple codes, which were then aggregated into one. Thus, the number of FLMA codes were reduced to 61. Moreover, we excluded eight FLMAs that did not have any communes covered by the 1998 CASEN survey and did not have initial shares (see Table A1 in the Supplemental material). Consequently, we used 53 FLMAs in our analysis.

We sourced the data on the applied tariff rates on final goods from 1999 to 2005 from Annex 4 of Becerra (2006, pp. 21–26). Some agricultural products (as defined by the World Trade Organization) had been exempt from applying the uniform MFN rates. The exceptions were 12.5% on poultry products and a price band system (bound at 31.5%) applied to wheat and wheat flour, various dairy products, oilseeds and oleaginous fruits, vegetable fats and oils, and cane or beet sugar (World Trade Organization, 2009). Becerra (2006) provided a thorough calculation for monthly data, including not only preferential margins granted by the RTA schemes but also those specific duties including the price band system. However, the data from Becerra (2006) covered only January 2000 to December 2005. Considering that the survey was conducted from November to December 2000, we used the data from the first quarter of 2000 for 1999. For 2002 and 2005, we used the average of the first to fourth quarters of the data.

We sourced the input share in Equation (33) from the input coefficient matrix, including the domestic and imported inputs of the Chilean input-output (I-O) table for 1996.[14] The CASEN survey reported workers' industry affiliations based on the International Standard Industrial Classification (ISIC) Revision 2. Becerra (2006) also provided industry-level applied tariff rates based on the ISIC Revision 2. Thus, based on the correspondence tables between the classification of the Chilean I-O table for 1986 and ISIC Revision 2 (Venegas Morales, 1994, p. 87) and between the classifications of the I-O tables for 1986 and 1996 (Annex 1 of Central Bank of Chile, 2001, p. 207), we matched the ISIC Revision 2 to the 1996 I-O table classification. However, some industries in the ISIC Revision 2 do not have exact equivalence in the classification of the 1996 I-O table. In this case, those industries (i.e., codes 1 to 3, 64 and 66, and 68 to 71) were aggregated into new codes and assigned to the most similar classification of the ISIC Revision 2 (see Table A2 in the Supplemental material). For this aggregation, the number of industries in the I-O table for 1996 is reduced from 73 to 67. We applied this classification (67 industries) to the

[14] We sourced the data from the Central Bank of Chile (https://si3.bcentral.cl/estadisticas/Principal1/Excel/CCNN/cdr/excel.html, accessed on June 21, 2023).

industry fixed effects of Equation (35). Finally, following McCaig (2011) and Topalova (2010), we assigned non-traded sectors (codes from 48 to 73) a zero tariff, as explained in footnote 7.

4-2. Descriptive statistics

We found that output tariff rates were the highest for food manufacturing products (ISIC 311–312) and agricultural products (ISIC 11) in 1999, which included products exempt from applying the uniform MFN tariff rates (see Table A2 in the Supplemental material). However, as major FTAs with faster and more comprehensive tariff elimination came into effect in the 2000s, the output tariff rates including those sectors decreased substantially from 1999 to 2005. Consequently, although the FLMA-level input tariff rates were the highest in FLMAs located in southern regions (e.g., codes 7103, 8203, 10202/10205, 10504, and 9102) in 1999, where the share of workers employed in the agricultural sector was high, the input tariff rates in those FLMAs also reduced substantially from 1999 to 2005. The maximum input tariff rate at the FLMA level was only 1.16% in 2005 (see Table A3 in Supplemental material). Additionally, Table A4 in the Supplemental material presents unemployment rates by FLMAs. According to the questionnaire of the CASEN survey, we defined the unemployment rate by the share of people aged 15–65 years who had looked for paid work in the past two months but had not worked in the past week. We found that the average unemployment rates decreased from 10.54% in 2000 to 7.50% in 2006, which was especially pronounced among FLMAs located in southern regions (e.g., code 10301), where input tariff rates also decreased substantially.

Tables 1 and 2 present the descriptive statistics and correlation matrices for the variables used. Because we used different weights for Equation (34) according to the two dependent variables (wage premium and unemployment rate), we presented the descriptive statistics and correlation matrices using the corresponding weights.[15] As expected, input tariff rates were negatively correlated with local wage premiums and positively correlated with unemployment rates.

In this empirical analysis, the crucial assumption is that labor is mobile across industries within FLMAs but immobile across FLMAs (Erten et al., 2019; Kovak, 2013). Although the CASEN surveys for 2000 and 2003 did not include any questions related to individuals' migration status, the survey for 2006 asked where they lived in April 2002. We found that only 3.81% of the labor force population reported

[15] Additionally, we presented the descriptive statistics of variables to estimate Equation (35) for each year in Table A5 in the Supplemental material.

moving across FLMAs between 2002 and 2006 (see Table A6 in the Supplemental material). This remarkably low migration across FLMAs strongly supports the assumption of our empirical analysis. Additionally, the survey for 2009 asked about workplaces. We found that only 3.79% of workers commuted outside the FLMAs where they lived (see Table A6 in the Supplemental material). Considering that approximately 9% of workers commuted outside the provinces, the FLMAs appropriately captured the spatial distribution of local labor markets in Chile.

5. Estimation results

5-1. Baseline estimation results

Columns [1] and [2] of Table 3 present the estimation results of Equation (34) for local wage premiums.[16] We found that the input tariff rate was negative and significant, which supports our theoretical prediction. We also found that the coefficient of the input tariff rate was underestimated in the absolute value in the estimation without the instrumental variable. The Kleibergen–Paap rank Wald *F*-statistics rejected the null hypothesis of a weak instrument. This finding indicated that unobservable time-varying local characteristics (e.g., lobbying power for protection) were positively correlated with both input tariff rates and wage premiums. Given that column [2] shows that a 1% reduction in the input tariff rate leads to a 5.24% increase in the local wage premium, the impact was practically large. Therefore, we concluded that trade-induced productivity improvements expectedly led to increased local wages. This finding was consistent with Murakami (2021), who found a negative association between output tariff rates and industry wage premiums in Chile from 2000 to 2006. However, the channel through which trade liberalization affected wages identified in this study was input tariffs rather than output tariffs. Additionally, this finding was quite different from those of previous studies, such as Kovak (2013) and Topalova (2010), who found a positive association between regional tariffs and wages or consumption for Brazil and India, and Erten et al. (2019), who found no significant association for South Africa.

Columns [3] and [4] of Table 5 present the estimation results of Equation (34) for unemployment rates. The input tariff rate was positive and significant in the estimation with the instrumental variable, supporting our theoretical prediction. Note that the input tariff rate was also

[16] See Tables A7 and A8 in the Supplemental material for the first-stage estimation results of Equation (35) and the FLMA-level local wage premiums for each year.

positive and significant at the 15% level in the estimation without the instrumental variable, which underestimated the coefficient of the input tariff rate. This finding indicated that unobservable time-varying local characteristics (e.g., lobbying power for protection) were positively correlated with input tariff rates and negatively correlated with FLMA-level unemployment rates. However, after appropriately controlling for the unobservable time-varying local characteristics, we found that a reduction in the input tariff rate leads to a decrease in the unemployment rate. Given that column [4] shows that a 1% reduction in the input tariff rate leads to a 5.01% decrease in the unemployment rate, the impact was practically large. This finding was quite different from that of Erten et al. (2019), who found a negative association between regional tariffs and unemployment rates in South Africa.

Based on the estimated local wage premiums and calculated unemployment rates (see Tables A8 and A4 in the Supplemental material respectively), we found that the mechanism underlying the negative and positive effects of the input tariff rates was that FLMAs with initially lower wage premiums and higher unemployment rates, mainly located in southern regions (e.g., codes 10202/10205 and 10301), experienced larger tariff reduction as well as larger wage increase and unemployment reduction. Furthermore, the practically larger effects observed in this study compared to previous research were derived from the fact that, despite the relatively small reduction in the tariff rates, at most around 4% (see Table A3), the increase in wages and decrease in unemployment rates were substantial, at most 13.5% and 17.3%, respectively (see Tables A8 and A4).[17]

Consequently, we found that the reduction in input tariff rates derived from the enforcement of RTAs contributed to increasing local wages and decreasing unemployment rates in Chile, thereby supporting our theoretical predictions in Section 2.

5-2. Robustness checks

Although FLMAs, which we used as the regional unit of analysis, more appropriately captured the spatial distribution of local labor markets than administrative units, it can be useful to check whether our estimation results are robust to changing the regional unit. Considering that the number of provinces (50) is comparable to that of FLMAs (53), we repeated our estimation using provinces as the unit of

[17] For example, Kovak (2013) found that estimated the coefficient of tariff reductions on local wage premiums were around 0.4, with 15% tariff reductions at most.

analysis.[18] Table 5 presents the estimation results, which are remarkably similar to the baseline results of Table 4. The input tariffs were negatively associated with wage premiums and positively associated with unemployment rates. Therefore, we concluded that the baseline results were robust to using the different regional units of analysis.

Finally, to rule out other possible channels through which trade liberalization affected wages and unemployment rates, we repeated our estimation using output tariffs instead of input tariffs.[19] Table 5 presents the estimation results. We found that although output tariffs were negative for local wage premiums and positive for unemployment rates, the effects are insignificant. Therefore, we concluded that the main channel was through input tariffs rather than output tariffs, which is consistent with our theoretical model.

6. Concluding remarks

By introducing a unique setting for the determinant of the average wage and different fixed production costs between exporting and domestic firms, we developed a theoretical model that incorporates a fair wage model into a trade model with heterogeneous firms. This setting incorporates market composition changes through firms' entry and exit into the determinants of workers' fair wages and unemployment rates. The model predicts that trade liberalization lowers prices of intermediate inputs, which reduces the production costs for final-goods firms and stimulates their labor demand. Consequently, trade liberalization leads to higher average productivity, thereby raising wages and reducing unemployment rates. Thus, the model allows us to reveal how reductions in trade costs affect wages and unemployment rates within local labor markets through reductions in production costs.

The empirical analysis employing exogenous variations in exposures to RTAs across local labor markets in Chile from 2000 to 2006 supported our theoretical predictions. Exploiting industrial variations

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[18] Although Chile had 53 provinces in 2007, three provinces (Antártica Chilena, Isla de Pascua, and Palena) were not covered by the 1998 CASEN survey.

[19] In this case, we constructed the output tariff rate τ_{rt}^{O} in FLMA *r* in year *t* by $\tau_{rt}^{O} = \sum_j Emshare_{jr}^{1998}\tau_{jt}^{O}$ based on Equation (32). Additionally, we cannot construct the instrument variable for τ_{jt}^{O} based on Equation (36) because MFN tariff rates were uniform across industries a given year. Thus, we only presented the estimation results without the instrument in Table 5.

in tariff reductions derived from the RTA schemes, initial differences in industrial compositions across local labor markets, and the lack of inter-local labor mobility, we found that a reduction in input tariff rates expectedly increased local wages and decreased unemployment rates. These findings were robust to the inclusion of various controls, the control for the endogeneity of the input tariffs, and the use of different regional units of analysis.

The results of this study provide important policy implications for the ongoing debate on trade liberalization and labor markets. Our results indicate that protection policies based on the adverse effects of trade liberalization on labor markets are likely to decrease wages and increase unemployment rates, contrary to their intention. Consequently, our findings indicate that policies that commit to trade liberalization and boost productivity should be adopted owing to their effects on labor markets.

References

Akerlof, G. A. (1982). Labor contracts as partial gift exchange. *Quarterly Journal of Economics*, *97*(4), 543–569. https://doi.org/10.2307/1885099

Akerlof, G. A., & Yellen, J. L. (1990). The fair wage-effort hypothesis and unemployment. *Quarterly Journal of Economics*, *105*(2), 255–283. https://doi.org/10.2307/2937787

Amiti, M., & Davis, D. R. (2011). Trade, firms, and wages: theory and evidence. *Review of Economic Studies*, *79*(1), 1–36. https://doi.org/10.1093/restud/rdr016

Bartik, T.J. (1991). *Who benefits from state and local economic development policies?* W. E. Upjohn Institute for Employment Research.

Becerra, G. (2006). *Arancel efectivo de las importaciones Chilenas: 2000–2005* (Studies in Economic Statistics No. 50). Central Bank of Chile.

Blanchard, O., & Giavazzi, F. (2003). Macroeconomic effects of regulation and deregulation in goods and labor markets. *Quarterly Journal of Economics*, *118*(3), 879–907. https://doi.org/10.1162/00335530360698450

Borusyak, K., Hull, P., & Jaravel, X. (2022). Quasi-experimental shift-share research designs. *Review of Economic Studies*, *89*(1), 181–213. https://doi.org/10.1093/restud/rdab030

Capuano, S., & Schmerer, H.-J. (2015). Trade and unemployment revisited: Do institutions matter? *The World Economy*, *38*(7), 1037–1063. https://doi.org/10.1111/twec.12220

Casado-Díaz, J. M., Martínez-Bernabéu, L., & Rowe, F. (2017). An evolutionary approach to the delimitation of labour market areas: An empirical application for Chile. *Spatial Economic Analysis*, *12*(4), 379–403. https://doi.org/10.1080/17421772.2017.1273541

Central Bank of Chile. (2001). *Matriz de insumo-producto de la economía chilena 1996*. Central Bank of Chile.

César, A., Falcone, G., & Gasparini, L. (2021). Costs and benefits of trade shocks: Evidence from Chilean local labor markets. *Labour Economics*, *73*, 102075. https://doi.org/10.1016/j.labeco.2021.102075

Chor, D., & Li, B. (2024). Illuminating the effects of the US-China tariff war on China's economy. *Journal of International Economics*, 150, Article 103926. https://doi.org/10.1016/j.jinteco.2024.103926

Danthine, J.-P., & Kurmann, A. (2006). Efficiency wages revisited: The internal reference perspective. *Economics Letters*, *90*(2), 278–284. https://doi.org/10.1016/j.econlet.2005.08.015

de Pinto, M., & Michaelis, J. (2014). International trade and unemployment—The worker-selection effect. *Review of International Economics*, *22*(2), 226–252. https://doi.org/10.1111/roie.12111

de Pinto, M., & Michaelis, J. (2019). The labor market effects of trade union heterogeneity. *Economic Modelling*, *78*, 60–72. https://doi.org/10.1016/j.econmod.2018.08.017

Egger, H., & Kreickemeier, U. (2009). Firm heterogeneity and the labor market effects of trade liberalization. *International Economic Review*, *50*(1), 187–216. https://doi.org/10.1111/j.1468-2354.2008.00527.x

Erten, B., Leight, J., & Tregenna, F. (2019). Trade liberalization and local labor market adjustment in South Africa. *Journal of International Economics*, *118*, 448–467. https://doi.org/10.1016/j.jinteco.2019.02.006

Felbermayr, G., Prat, J., & Schmerer, H. J. (2011). Globalization and labor market outcomes: Wage bargaining, search frictions, and firm heterogeneity. *Journal of Economic Theory*, *146*(1), 39–73. https://doi.org/10.1016/j.jet.2010.07.004

Goldberg, P. K., & Pavcnik, N. (2007). Distributional effects of globalization in developing countries. *Journal of Economic Literature*, *45*(1), 39–82. https://doi.org/10.1257/jel.45.1.39

Haisken-DeNew, J. P., & Schmidt, C. M. (1997). Interindustry and interregion differentials: Mechanics and interpretation. *The Review of Economics and Statistics*, *79*(3), 516–521. https://doi.org/10.1162/rest.1997.79.3.516

Helpman, E., Itskhoki, O., & Redding, S. (2010). Inequality and unemployment in a global economy. *Econometrica*, *78*(4), 1239–1283. https://doi.org/10.3982/ECTA8640

Helpman, E., & Itskhoki, O. (2010). Labor market rigidities, trade and unemployment. *Review of Economic Studies*, *77*(3), 1100–1137. https://doi.org/10.1111/j.1467-937X.2010.00600.x

Kovak, B. K. (2013). Regional effects of trade reform: What is the correct measure of liberalization? *American Economic Review*, *103*(5), 1960–1976. https://doi.org/10.1257/aer.103.5.1960

McCaig, B. (2011). Exporting out of poverty: Provincial poverty in Vietnam and U.S. market access. *Journal of International Economics*, *85*(1), 102–113. https://doi.org/10.1016/j.jinteco.2011.05.007

Melitz, M. J. (2003). The impact of trade on intra-industry reallocations and aggregate industry productivity. *Econometrica, 71*(6), 1695–1725. https://doi.org/10.1111/1468-0262.00467

Ministerio de Hacienda. (1998). *Ley 19589*. https://www.bcn.cl/leychile/navegar?idNorma=127108

Murakami, Y. (2021). Trade liberalization and wage inequality: Evidence from Chile. *The Journal of International Trade & Economic Development*, *30*(3), 407–438. https://doi.org/10.1080/09638199.2020.1871502

Nishiyama, H., & Gintani, Y. (2021). Globalization trap? Trade and labor market interactions revisited. *The International Economy*, *24*, 166–186. https://doi.org/10.5652/internationaleconomy/ie2020.24.07.hn

Rowe, F. (2017). Functional labour market areas for Chile. *Region*, *4*(3), 7–9. https://doi.org/10.18335/region.v4i3.199

Topalova, P. (2010). Factor immobility and regional impacts of trade liberalization: Evidence on poverty from India. *American Economic Journal: Applied Economics*, *2*(4), 1–41. https://doi.org/10.1257/app.2.4.1

Venegas Morales, J. (1994). *Una matriz insumo-producto inversa de la economia Chilena 1986* (Studies in Economic Statistics No. 38). Central Bank of Chile.

World Trade Organization (WTO). (2009). *Trade policy review, report by the secretariat: Chile*. WTO.

Table 1 Descriptive statistics of used variables.

Variables	Observations	Mean	Standard deviation	Min	Max
Local wage premium	159	-0.0596	0.1118	-0.3087	0.3248
Lag tariff	159	0.0225	0.0121	0.0075	0.0545

Variables	Observations	Mean	Standard deviation	Min	Max
Unemployment rate	159	0.0931	0.0231	0.0228	0.2208
Lag tariff	159	0.0222	0.0114	0.0075	0.0545

Source: Authors' calculations based on the data sources presented in Section 4.
Note: We used the inverse of the standard errors of wage premiums of each functional labor market area (FLMA) presented in Table A8 in the Supplemental material as the weights for the calculations of the first table, and the employment share of each FLMA in 1998 as the weights for the calculations of the second table.

Table 2 Correlation matrix of used variables.

	Local wage premium	Lag tariff
Local wage premium	1.0000	
Lag tariff	-0.1227	1.0000

	Unemployment	Lag tariff
Unemployment rate	1.0000	
Lag tariff	0.5060	1.0000

Source: Authors' calculations based on the data sources presented in Section 4.
Note: We use the inverse of the standard errors of wage premiums of each functional labor market area (FLMA) presented in Table A8 in the Supplemental material as the weights for the calculations of the first table, and the employment share of each FLMA in 1998 as the weights for the calculations of the second table.

Table 3 Estimation results of Equation (34).

	Local wage premium		Unemployment rate	
	(1)	(2)	(3)	(4)
Lag tariff	-4.1874*	-5.0533**	4.4506	5.0170***
	(2.2877)	(2.0174)	(2.6948)	(1.7465)
Initial local characteristics×Year fixed effects	No	No	Yes	Yes
FLMA fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Instruments	No	Yes	No	Yes
Kleibergen-Paap rank Wald *F*-statistics		312.55		100.418
Stock-Yogo weak identification test critical values		16.38		16.38
Observations	159	159	159	159
R-squared	0.9278	0.9276	0.8106	0.8105

Note: ∗∗∗, ∗∗, and ∗ indicate significance at 1%, 5%, and 10%, and levels, respectively. Numbers in parentheses represent robust standard errors clustered by functional labor market area (FLMA). The inverse of the standard errors of the wage premiums of each FLMA and the employment share of each FLMA in 1998 were used as the weights for the estimations for local wage premiums and unemployment rates, respectively.

Table 4 Estimation results of Equation (34) using provinces as the unit of analysis.

	Province wage premium		Unemployment rate	
	(1)	(2)	(3)	(4)
Lag tariff	-5.3006**	-6.0579***	4.0194*	4.1221*
	(2.4903)	(2.2271)	(2.3971)	(2.1465)
Initial local characteristics×Year fixed effects	No	No	Yes	Yes
Province fixed effects	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes
Instruments	No	Yes	No	Yes
Kleibergen-Paap rank Wald F-statistics		349.462		74.756
Stock-Yogo weak identification test critical values		16.38		16.38
Observations	150	150	150	150
R-squared	0.9306	0.9304	0.8114	0.8114

Note: ∗∗∗, ∗∗, and ∗ indicate significance at 1%, 5%, and 10%, and levels, respectively. Numbers in parentheses represent robust standard errors clustered by province. The inverse of the standard errors of the wage premiums of each province and the employment share of each province in 1998 were used as the weights for the estimations for province wage premiums and unemployment rates, respectively.

Table 5 Estimation results of Equation (34) using output tariff rates.

	Local wage premium	Unemployment rate
	(1)	(4)
Lag tariff	-1.6357	0.4864
	(1.0140)	(2.1131)
FLMA fixed effects	Yes	Yes
Year fixed effects	Yes	Yes
Initial FLMA characteristics×Year fixed effects	No	Yes
Observations	159	159
R-squared	0.9273	0.8047

Note: Numbers in parentheses represent robust standard errors clustered by functional labor market area (FLMA). The inverse of the standard errors of the wage premiums of each FLMA and the employment share of each FLMA in 1998 were used as the weights for the estimations for local wage premiums and unemployment rates, respectively.