Measured Wage Inequality and Input Trade Liberalization: Evidence from Chinese Firms*

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July 24, 2014

Abstract

This paper investigates how input trade liberalization affects within-firm wage inequality between skilled and unskilled labor. A fall in input tariffs increases firm profits, which, in turn, widens wage inequality since skilled labor enjoys a larger proportion of the incremental profits. Using Chinese firm-level production data, we first develop econometric methods to estimate and calculate within-firm *measured* wage inequality. After controlling for possible endogeneity, we find evidence that input trade liberalization widens within-firm measured wage inequality. Such findings are robust to different measures of wage inequality, as well as different empirical specifications and data spans.

JEL Classifications: F10, F12, F14

Keywords: Wage Inequality, Input Trade Liberalization, Firm Evidence

*We thank Andrew Bernard, Binkai Chen, Meredith Crowley, Zhiyuan Li, Kunwan Li, Larry Qiu, Puyang Sun, Yang Yao, Chuanchuan Zhang, Xiaobo Zhang, and participants at the 2012 Columbia-Tsinghua conference and SSE-CCER international conference in Beijing, 2014 EITI conference in Phuket, Xiamen University, Nankai University, China's Central University of Finance and Economics for their constructive and helpful suggestions. However, all errors are ours.

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1 Introduction

Tariffs have declined dramatically worldwide as the result of many rounds of GATT/WTO trade negotiations (Bagwell and Staiger, 1999). The question of how trade liberalization affects wages and income distributions, especially for the developing countries, has been the subject of many studies in the international trade literature. Most previous works relied on the new-classical Heckscher-Ohlin model as guidance to test whether or not trade liberalization benefits the abundant factor and, therefore, affects income distribution between skilled and unskilled labor. If the Stopler-Samuelson theorem is supported by data, trade liberalization on imported capital-intensive goods would mitigate wage inequality in the developing countries.[1] However, recent empirical studies find that globalization leads to larger wage inequality. For instance, Feenstra and Hanson (1996, 1999) also find that, in the presence of vertical integration, outsourcing as a result of freer trade would increase wage inequality in both developed and developing countries.

However, most studies rely on industry-level wage data, household survey data and proxy wage inequality using the Gini coefficient – a standard indicator of income inequality (e.g., Beyer *et al.* , 1999). The absence of firm and worker heterogeneity makes wage inequality within firms a type of "black-box." Using production data from China, the current paper tries to fill this gap.[2]

Previous studies often concentrate on trade liberalization in final goods. For example, Han *et al.* (2012) finds that China's accession to the World Trade Organization since 2001 was strongly

[1] Previous works have an intense discussion on the validity of factor price equalization (FPE) in explaining wage inequality in developed countries. For example, Johnson and Stafford (1993) and Leamer (1993, 1996) argue that FPE can explain the wage gap between skilled and unskilled workers in the U.S. However, Lawrence and Slaughter (1993) reviewed historical data on the prices of labor-intensive and capital-intensive goods and found that the movement of the relative prices of these two types of goods may suggest wage equality according to FPE.

[2] An outstanding exception is that of Akerman *et al.* (2013), which finds that trade liberalization not only enhances the dispersion of revenues across heterogenous firms but also widens wage inequality across workers and firms. We are also in line with Groizard *et al.* (forthcoming), which explores the endogenous nexus between trade liberalization and job flow in California. Furusawa and Konishi (2014) proposed a model to interpret why international trade can widen a wage gap between top income owners and others and thus cause job polarization.

associated with widening wage inequality in China using urban industrial survey data.[3] However, imported intermediate inputs are found to be crucial to boosting firm productivity for many countries, such as the U.S. (Hanson *et al.*, 2005), Indonesia (Amiti and Konings, 2007), India (Topalova and Khandelwal, 2011) and China (Yu, forthcoming), which could also, in turn, affect wage inequality. In this paper, we focus on the impact on wage inequality of trade liberalization in intermediate input markets.

Helpman et al (2014) is a recent study on the issue of trade and wage inequality in a model of heterogeneous firms. Using Brazilian data, they also find that wage inequality does not mainly stem from cross sector-occupation difference. Amiti and Davis (2012) investigate the impact on wages, but not on wage inequality, of both output and input tariff reductions. In particular, they find that a reduction in input tariffs raises wages at import-using firms relative to those at firms that only use domestic intermediate inputs.

The present paper investigates the impact of trade liberalization on income distribution from a new angle. In particular, we seek to understand the impact of input trade liberalization on within-firm wage inequality between skilled and unskilled labor. The paper makes the following two contributions to the literature. First, it provides a methodology for constructing firm-level measured wage inequality from firm profits and labor shares, which can be applied to other research projects facing similar data constraints. Second, the paper provides direct evidence that input tariff reductions increase a firm's profitability and consequently within-firm wage inequality between skilled and unskilled workers.[4] We find that the magnitude of the relative measured wage inequality is 2.21 (i.e. wages of skilled labor are more than twice that of unskilled labor). This figure is much higher than the one found in the U.S. (approximately 1.75) during the same period of 2000 to 2006 (Feenstra, 2010).[5] The result is in line with the fact that China's Gini coefficient was 0.49 in 2012, which is

[3]Autor et al. (2013) stresses that China's exports to the American market significantly contribute to the aggregate decline in the U.S. manufacturing employment and causes the sharp increase in U.S. social benefit claims.

[4]In Appendix C, we have also provided a theoretical model for this result.

[5]It is also higher than the one in European countries, presumably due to the fact that European countries typically have much stronger labor unions (Kranz, 2006).

much higher than the U.S. (approximately 0.30). Therefore, this study enriches our understanding of the sources of China's growing income inequality because wage inequality is an important part of its components.[6]

In carrying out this analysis, we face an immediate econometric challenge when we try to estimate the impact of trade liberalization on within-firm wage inequality. The Chinese firm-level production data do not have direct information on skilled and unskilled wages. The data set only reports firm's total wage bills. We hence develop an econometric approach to estimate the measured wage inequality (defined as within-firm skilled-labor and unskilled-labor wage ratio) and the measured wage gap (defined as the within-firm difference between skilled-labor and unskilled-labor) by using employment information of a firm's skilled and unskilled labor. Since the firm-level data set only provides the employment information on skilled labor and unskilled labor in 2004 (i.e., the third China's industrial census year), to overcome these data limitations we compute a proxy of the skilled-labor share for all other years by multiplying the skilled-labor share in 2004 with a provincial skilled-labor share in all years using 2004 as the base year.

Our primary econometric approach to estimate the measured wage inequality is in the spirit of "fair wages", a là Egger and Kleickemier (2012). In particular, firm's wage bills (for both the skilled and unskilled) consist of two components. One component is determined by the market value in a narrowly-defined industry whereas the other is determined by the firm-specific profit which can thus be treated as a "fair wage" component. In this way, the within-firm wage inequality can be inferred by our available firm-level information such as total wage bills, skilled-labor share, and firm profit. We then estimate and calculate firm-level wage inequality in four different ways.

We first obtain the wage gap between skilled and unskilled labor using a firm's total profit as a proxy for its profitability. Since a firm's total profit is correlated with its size, we then adopt a firm's profit-sales ratio as an alternative proxy to estimate a firm's wage inequality.[7] Since all such measured

[6]For example, Khan and Riskin (1998) found that wage inequality contributed to half of the income inequality in China in 1995.

[7]Similar to the literature on measures of wage inequality, we take an additional step to denote wage inequality as the ratio between skilled and unskilled wages.

wage inequality is estimated rather than observed, it could be the case that some observations are estimated more precisely than others. Therefore, we compute the standard deviation of a firm's relative wages across firms within an industry and multiply by a firm's relative wages, which then is used as a new measure of a firm's wage inequality.[8] To avoid a concern that our measure on wage inequality heavily relies on firm profit, our last alternative approach uses industrial minimum wage as a proxy to estimate the within-firm wage inequality.

After the index of within-firm wage inequality is measured, our second-step estimation is to examine the role of input trade liberalization on measured wage inequality. After controlling for possible endogeneity issues from reverse causality and omitted variables, we find that input trade liberalization widens wage inequality within firms. Such findings are robust to different measures of wage inequality, as well as different empirical specifications and data spans.

The rest of the paper is organized as follows. Section 2 describes the data and carefully introduces our econometric methods to measure within-firm wage inequality. Section 3 presents the empirical evidence and robustness checks. Finally, Section 4 concludes.

2 Data and Measures

2.1 Data

To investigate the impact of trade liberalization (mainly in terms of input tariff reductions) on a firm's wages gap, in this paper we use the following two disaggregated panel data sets: firm-level production data complied by China's National Bureau of Statistics (NBS), and China's import tariffs (*ad valorem*) data at HS 6-digit level as maintained by the World Integrated Trade Solution (WITS) of the World Bank.

Firm-level Production Data. China's NBS conducts an annual survey on two types of manufacturing firms: all state-owned enterprises (SOEs) and non-SOEs whose annual sales exceed RMB 5 million (or equivalently $830,000). The sample used in this paper is approximately 230,000 manu-

[8] Feenstra *et al.* (forthcoming) also use this approach to handle trade uncertainty regarding Chinese firms' exports.

facturing firms per year varying from 162,885 firms in 2000 to 301,961 firms in 2006. On average, the sample accounts for more than 95% of China's total annual output in the manufacturing sectors.[9] The data set covers more than 100 accounting variables and contains all of the information from the main accounting sheets, which includes balance sheets, loss and profit sheets and cash flow statements.

Given its rich information, the firm-level production data set is now widely used in research, including, among others, Cai and Liu (2009), Brandt *et al.* (2012), and Feenstra *et al.* (forthcoming). However, it has two limitations in our research. The first one is common to all research topics: some unqualified firms are wrongly included in the data set largely because of mis-reporting by some firms. Thus, following Feenstra *et al.* (forthcoming), we keep the observations in our analysis according to the requirements of generally accepted accounting principles (GAAP).[10] Accordingly, the total number of firms covered in the data set is reduced from 615,951 to 438,165, and approximately one-third of the firms are removed from the sample after the rigorous filter is applied. The second limitation is specific to our present paper. The data set does not include skilled-labor and unskilled-labor wages. It only reports the number of skilled workers and unskilled workers in the census year (2004). To overcome these data limitations we compute a proxy of the skilled-labor share for all other years by multiplying the skilled-labor share in 2004 with a provincial skilled-labor share changes in other years using 2004 as the base year.

Notice that some Chinese firms in the data are pure trade intermediaries that do not have production activities. To ensure the precision of our estimates, we exclude these firms from the sample in all estimates. In particular, trade intermediaries are selected according to the same procedures as in Ahn *et al.* (2011).

[9] In 2006, the value added of above-sale firms in the survey is RMB 9,107 billion, which accounts for 99% of the value-added of all firms in the manufaturing sectors (RMB 9,131 billion) as reported by China's Statistics Yearbook (2007).

[10] We keep the observations if all of the following holds: (1) total assets exceed liquid assets; (2) total assets exceed total fixed assets; (3) the net value of fixed assets is smaller than total assets; (4) the firm's identification number exists and is unique and (5) the established time is valid.

2.2 Measures

Since our data sets do not directly provide firm-level wages for skilled-labor and unskilled-labor, in this section we develop a method for constructing firm-level measured wage inequality. We will also introduce the index of input trade liberalization.

2.2.1 Measures of Within-firm Wage Inequality

Supposed that skilled wages (w_{ijt}^s) paid by a firm i of industry j in year t can be decomposed to two components: industrial average skilled wages (w_{jt}^s) and a firm-specific term (ε_{ijt}^s): $w_{ijt}^s = w_{jt}^s + \varepsilon_{ijt}^s$. Similarly, a firm's unskilled wages are decomposed to industrial average unskilled wages and a firm-specific term: $w_{ijt}^u = w_{jt}^u + \varepsilon_{ijt}^u$. That is, wages can be determined by the market value of the specific type of labor and their firm-specific incentive pay. Furthermore, the incentive pay to the skilled labor is presumed to be positively related to the firm's profitability thanks to their bargaining power to the firm, but there is no bargaining power in the unskilled labor due to their perfect substitutability (i.e. the unskilled labor market is perfectly competitive). Therefore, a firm's wages inequality ($wgap1_{ijt}$) can be expressed as

$$wgap1_{ijt} \equiv w_{ijt}^s - w_{ijt}^u = (w_{jt}^s - w_{jt}^u) + (\varepsilon_{ijt}^s - \varepsilon_{ijt}^u), \tag{1}$$

where the first equality is by definition. In the second equality, the first term is the industry-level wage inequality (later refer to α_{jt}) whereas the second term is the firm-level difference in the skilled and unskilled wage residuals. As suggested by "fair wage" theory and confirmed by previous findings such as Cahuc *et al.* (2006), such wage residuals are a function of a firm's profitability.[11] A more profitable firm would allocate more dividends to skilled workers.[12] Since the average profitability varies by industry, we shall capture such diverges in our estimates. Thus, the linear approximation

[11] Although labor union exists in most of Chinese enterprises, it is well-known as a quasi-administative organization which typically does not bargain benefits for the unskilled labor.

[12] Since larger firms (i.e., with more sales) usually have more profits, we measure a firm's profitability as firm profits over sales. However, our main estimation results remain robust even if we use a firm's total profit as a proxy for its profitability, which will be discussed later.

of the within-firm wage residuals can be estimated by $\varepsilon^s_{ijt} - \varepsilon^u_{ijt} = \beta_{jt}\pi_{ijt}$ where π_{ijt} is a firm's profitability and β_{jt}, the bargaining power of the skilled labor, is the estimated coefficient for industry j in year t. Hence, a firm's wages inequality is given by:

$$wgap1_{ijt} = \alpha_{jt} + \beta_{jt}\pi_{ijt} \tag{2}$$

Thus, the measured wage gap can be identified once the parameters α_{jt} and β_{jt} are estimated. To this end, we first estimate the coefficients in Equ. (2) by industry and year using data on average wages and share of skilled labor. Denoting θ_{ijt} the skilled labor share, which is measured by the share of employees with at least college degree, a firm's average wage can be expressed as (See Appendix A for details):

$$\begin{aligned}\overline{w}_{ijt} &\equiv \theta_{ijt}w^s_{ijt} + (1-\theta_{ijt})w^u_{ijt} = \theta_{ijt}(w^s_{jt} + \varepsilon^s_{ijt}) + (1-\theta_{ijt})(w^u_{jt} + \varepsilon^u_{ijt}) &(3)\\ &= w^u_{jt} + \alpha_{jt}\theta_{ijt} + \beta_{jt}(\theta_{ijt}\pi_{ijt}) + \varepsilon^u_{ijt}, &(4)\end{aligned}$$

where the first equality is by definition whereas the second one is obtained by incorporating Equ. (1). The third equality is obtained by using $w^s_{jt} = w^u_{jt} + \alpha_{jt}$ and $\varepsilon^s_{ijt} - \varepsilon^u_{ijt} = \beta_{jt}\pi_{ijt}$. The error term ε^u_{ijt} is assumed to be i.i.d. With data on a firm's average wages ($\overline{w}_{ijt}$), skilled labor share(θ_{ijt}) and firm profitability(π_{ijt}) measured by dividing firm profit-sales ratio, we can estimate the coefficients $\widehat{\alpha}_{jt}$ and $\widehat{\beta}_{jt}$. Notice that since the residual, ε^u_{ijt}, is also (positively) related to profitability, the estimated $\widehat{\beta}_{jt}$ is supposed to be smaller than the actual β_{jt}.[13] Thus, our estimates here shall be interpreted as the lower bound of the "true" estimated coefficient. Note that since the term w^u_{jt} is varied by each industry-year pair, we estimate Equ. (3) by industry and by year so that w^u_{jt} is treated as a constant term in *each* regression. The measured within-firm wage inequality can be computed by backing up

[13] An alternative approach is to further decompose the error term to a combination of a linear function of firm probability and an idnsyncratic term $\varepsilon^u_{ijt} = \beta^u_{jt}\pi_{ijt} + \varsigma_{ijt}$ where ς_{ijt} is i.i.d. by assuming that unskilled workers also get some proportion firm profit as their wage compensation. All our estimation results are unchanged by using this approach. To save space we do not report those estimation results though available upon request.

the estimated coefficients of $\widehat{\alpha}_{jt}$ and $\widehat{\beta}_{jt}$:

$$\widehat{wgap1}_{ijt} = \widehat{\alpha}_{jt} + \widehat{\beta}_{jt}\pi_{ijt}. \tag{5}$$

Note that since $\widehat{\beta}_{jt}$ is a lower bound of β_{jt}, $\widehat{wgap1}_{ijt}$ is therefore a conservative estimate of the actual wage gap between the skilled and unskilled. Column (1) of Table 1A presents the year-average measured firm-level wage inequality ($\widehat{wgap1}_{ijt}$) by Chinese two-digit industry. The mean of firm wage inequality, measured by the wage gap between skilled labor and unskilled labor in *absolute* terms, is RMB 11,320 (or equivalently, $1400) with a relative large standard deviation. This large standard deviation is likely due to the inclusion of outlier industries such as tobacco (code: 16. See column (1) of Table 2), which has an extremely high measured wage inequality. To ensure that our estimates are not contaminated by such outliers, we drop firms from the tobacco industry from our estimations. Table 1A also provides some basic statistical information for the key variables used in the estimations.

[Insert Table 1 Here]

2.2.2 Measures of Input Tariffs

Inspired by Amiti-Konings (2007) and Topalova and Khandelwal (2011), we construct the industry-level input tariffs, IIT_{jt}, as follows:

$$IIT_{jt} = \sum\nolimits_n \left(\frac{input_{nj}^{2002}}{\sum_n input_{nj}^{2002}} \right) \tau_{nt}, \tag{6}$$

where IIT_{jt} denotes the industry-level input tariffs facing firms in industry j in year t. τ_{nt} is the tariff of input n in year t. The weight in the parenthesis is measured as the production cost share of input n in industry j.

We use China's Input-Output Table from 2002 to construct the weight since NBS reports the Input-Output Table every five years and our data spread from 2000 to 2006. The industrial input tariffs are obtained as follows. First, since there are 71 manufacturing sectors reported in China's Input-Output Table (2002) and only 40 manufacturing sectors reported in Chinese industrial classifications (CIC), we start by making a concordance between the Input-Output Table and the CIC sectors. Secondly, we match the CIC sectors with International Standard Industrial Classifications (ISIC, rev. 3).[14] Third, we make another concordance to link ISIC and the HS 6-digit trade data where we can find the corresponding tariffs from the WITS. Fourth, we calculate the industry-level tariffs that are aggregated to the CIC sectorial level.[15] In particular, the simple average tariffs are used to calculate industry-level tariffs as follows:

$$\tau_{nt} = \frac{1}{N} \sum\nolimits_{k \in n, k=1}^{N} \tau_{kt}, \tag{7}$$

where k denotes products (at the HS 6-digit level) in industry n. We use these simple average tariffs as a default measure in the main estimates that follow. Finally, we calculate the industry-level *input* tariffs using Equ. (6). Similarly, the industry-level *output* tariff for industry n in year t is also obtained from Equ. (7).

To see how the input tariff reductions affect a firm's wages inequality, it is worthwhile to examine the evolution of China's trade liberalization throughout the sample period. Table 1B reports the mean and standard deviation for this key variable. As shown in Table 1B, the average industry input tariffs were cut in half from 16.6% in 2000 to 8.6% in 2006, and their standard deviation also dropped by about two-thirds over the same period.

[Insert Table 2 Here]

[14] Note that China's government adjusted its CIC in 2003. Therefore we also make similar adjustments in our data.
[15] We do not report the input weights by industry to save space; these data are available upon request.

3 Estimation Results

3.1 Baseline Results

After obtaining both measured firm-level wage inequality from Equ.(5) and industry-level input tariffs from Equ.(6), we consider the following specification to explore the impact of input tariffs on measured wage inequality:

$$\widehat{wgap1}_{ijt} = \beta_0 + \beta_1 IIT_{jt} + \boldsymbol{\psi}\mathbf{X} + \varpi_i + \gamma_t + \mu_{it}, \tag{8}$$

where $\mathbf{X}$ includes other control variables such as industry output tariffs and other firm characteristics (e.g., type of ownership, size and productivity). The error term in Equ. (8) can be further decomposed to three terms: (1) a firm-specific fixed effects ϖ_i to control for time-invariant factors such as a firm's unobserved managerial ability (Qiu and Yu, 2014); (2) year-specific fixed effects η_t to control for firm-invariant factors such as Chinese RMB appreciation since 2005; and (3) an error term μ_{it} for other unspecified characteristics.

We start our estimations by running a simple regression. By abstracting away all the control variables, the fixed-effects estimates in column (1) of Table 3 show that a fall in industry input tariffs tends to result in more wage inequality. One may wonder whether such a cost-saving effect could be weakened by tougher import competition effects due to the inclusion of output tariffs (Amiti and Konings, 2007). Furthermore, other firm characteristics, such as a firm's type of ownership, size (measured by log of firm employment) or productivity (measured by the Olley-Pakes (1996) TFP), could also affect a firm's wage gap. Therefore, we include all such control variables in column (2) and we still see a negative and significant estimate for industry input tariffs. In addition, output trade liberalization tends to narrow wages inequality, possibly, due to the tougher import competition and consequent decline in firm profitability (Horn *et al.* 1995). Note that SOEs and foreign indicators (FIEs) are still present in the estimates after controlling for firm-specific and year-specific fixed effects.

This is merely because some SOEs (FIEs) could switch to non-SOEs (non-FIEs) or vice versus.[16]

It is worthwhile to stress our productivity measures. Following Amiti and Konings (2007), we first estimate and calculate the augmented Olley-Pakes (1996) TFP by including an extra WTO dummy (i.e., one after 2001 and zero otherwise) and its interaction terms with input coefficients in the first-step control function (See Yu (forthcoming) for detailed steps). Since different industries use different technology, the measured Olley-Pakes TFP are not comparable across industry. To address such a potential shortcoming, inspired by Arkolakis (2010) and Arkolakis and Muendler (2011)[17], we normalize our Olley-Pakes TFP to obtain firm's relative TFP. Specifically, following Groizard *et al.* (forthcoming), we first rank firm's Olley-Pakes TFP from the lowest to the highest *within* each CIC 2-digit industry, then we map this ranking to the [0,1] range. In our main estimates we will use relative TFP to measure firm productivity, but our results remain robust even using conventional Olley-Pakes TFP, as shown in Table 7.

3.2 The Role of Processing Trade

Similar to Feenstra *et al.* (forthcoming), approximately 4.5% of firms in Table 1A are pure exporters that sell all of their products abroad. An interesting observation is that most pure exporters are processing firms that enjoy the special tariffs treatment (i.e., free duty) for importing (e.g., Dai *et al.* , 2012). The existence of processing firms suggests a helpful clue for our identification. First, with the inclusion of processing firms, our estimates of input trade liberalization on wage inequality may be under-estimated since the presence of processing firms, which are already duty-free, would dilute the magnitude of our estimations. Ideally we need to remove from the sample those processing firms whose processing imports equal total imports. But, the NBS firm-level data set does not have the information. However, by definition, a processing firm is also a pure exporter (although it is not necessary that a pure exporter be a processing firm). Therefore, we can re-run the fixed-effects

[16] To save space, we do not report the transitional probability for SOEs and FIEs, but they are available upon request.

[17] Arkolakis (2010) and Arkolakis and Muendler (2011) suggest following this approach to estimate the shape parameter of the underlying Pareto productivity distribution.

estimates after removing the pure exporters in column (3) of Table 3. The coefficient of input tariffs is still negative and significant, with a relatively larger magnitude than its counterpart in column (2) when pure exporters are included. The coefficients of firm productivity (in a form of relative TFP) are positive and significant in columns (2)-(3), suggesting that more productive firms have wider wage gaps between skilled and unskilled labor.

[Insert Table 3 Here]

3.3 Endogeneity Issues

In our previous estimations, we treat input trade liberalization exogenous. However, tariff formation could be endogenous in the sense that wage inequality could reversely affect tariff changes. With widening wage inequality, unskilled workers could blame free-trade policy and form labor unions to lobby the government for temporary trade protection (Bagwell and Staiger, 1990, 1999; Grossman and Helpman, 1994; Bown and Crowley, 2013). Although this happens in developed countries like the U.S. (Goldberg and Maggi, 1999) and in some developing countries like Turkey (Gawande and Bandyopadhyay, 2000), it is less likely to happen in China given that labor unions in China are symbolic organizations. As well, if these types of political factors are time invariant, then our fixed-effect panel estimates in Table 3 have accounted and controlled for them (Goldberg and Pavcnik, 2005). However, if they are time variant, we need to use the instrument variables (IV)-approach to control for these types of endogeneity issues.

It is always challenging to find an ideal instrument for tariffs. Inspired by Amiti and Davis (2012), we use the one-year lag of industry input tariffs as the instrument of the first difference in industrial input tariffs. The economic rationale is that the lag input tariffs are less likely to influence the time difference of input tariffs (Trefler, 2004). The last two columns of Table 3 perform the two-stage least squares (2SLS) estimates by treating industry input tariffs as endogenous.

The estimated coefficient of industry input tariffs in the 2SLS estimates in column (4) of Table 3 is close to its counterpart in the OLS estimates in column (3). A 10 percent reduction in industry input tariffs leads to an approximately 15.4 percent in a firm's wage inequality, *ceteris paribus.* By dropping pure exporters from the sample in column (5), the estimated coefficient of input tariffs is slightly larger than that in column (4), which, in turn, suggests that the inclusion of pure processing firms could dilute the impact of input trade liberalization on a firm's wage inequality.[18]

We now perform related statistical tests to check for the validity of such an instrument. The bottom module of Table 3 provides the first-stage estimates for all specifications. The coefficients of the instruments are negative and highly statistically significant, suggesting that industries with initial high tariffs are more challenging to remove their tariff barriers. In addition, several tests were performed to verify the quality of the instruments. First, we use the Kleibergen–Paap LM χ^2 statistic to check whether the excluded instruments are correlated with the endogenous regressors. As shown in the upper module of Table 3, the null hypothesis that the model is under-identified is rejected at the one percent significance level. Second, the Kleibergen–Paap (2006) F-statistics provide strong evidence for rejecting the null hypothesis that the first stage is weakly identified at a highly significant level. All these tests suggest that our instrument is valid and that the specifications are well justified.

3.4 Cross-Firm versus Time-Series Variations

Because we do not have a firm's skilled and unskilled labor data for any year except 2004, we have to multiply a proportion of skilled labor at the province-year level using 2004 as a base year to construct the variable of firm-year wage inequality. This may raise a concern that whether our results are driven by provincial heterogeneity rather than firm heterogeneity.[19] To address this concern, we perform the following two placebo tests.

The first robustness check is to drop samples in all years except 2004. We perform the cross-section OLS estimates with data from 2004 only in column (1) of Table 4. We also include two-digit industry fixed effects to wash out unspecified industry characteristics. The coefficient of input tariffs is still

[18] Amiti and Davis (2011) adopt five-period difference estimations. Introducing a longer period (e.g., two-period) differenced equation does not change our estimation results in a quantitative way, although there is a cost as we lose much of the sample in such a short panel.

[19] Note that including provincial dummies in the regressions does not change our estimation results. Since firms do not change their locations, all province-level fixed effects are automatically absorbed by firm-level fixed effects.

negative and significant, indicating that input trade liberalization widens a firm's wage inequality. Still, we suspect that the OLS fixed-effects estimates may be biased due to possible endogeneity issues caused by omitted variables or reverse causality. Therefore, it is ideal to perform the 2SLS estimates. Since we only have one-year data here, it is inapplicable to run the first-difference estimates. Hence, we instead choose the one-year lag of industry input tariffs as the instrument. The economic rationale is that industries with strong trade protection are more likely to maintain relatively high tariffs than those with weak protection due to the role of special interest groups (Grossman and Helpman, 2001). Column (2) of Table 4 performs the 2SLS estimates using one-period lag of input tariffs as the instrument. It turns out that the coefficient of input tariffs is relatively close to its counterpart for the full-sample 2SLS estimate in column (4) of Table 3, which confirms that our full-sample estimates are not driven by the adoption of a relatively aggregated multiplier (i.e. the province-year skilled-labor share). The positive and significant coefficient of the instrument in the first-stage ascertains our presumption that industries with strong trade protection are more likely to retain high tariffs compared to other industries.

The second robustness check is to narrow down the time-series window. Since we only have data of one-year on skilled (and unskilled) labor, one may worry that running regression for a seven-year period (2000 to 2006) may generate some serial correlations or cause some concern of unit roots that may be prevalent in long-period estimates. To address this issue, we shut down the long time-series window and only focus on a three-years period (2003 to 2005). We then perform the fixed-effects OLS estimates in column (3) and 2SLS estimates in column (4) of Table 4 in which the regressand is the first-difference of measured wage gap and the instrument is the one-lag industry input tariffs. After controlling for the endogeneity, the coefficient of input tariffs in the 2SLS estimates in column (4) has an identical negative sign and is statistically significant. Thus, we can conclude that our results are insensitive to the adoption of the province-year skilled-labor share as a remedy to data restrictions.

[Insert Table 4 Here]

3.5 Estimates using Alternative Tariffs Measure

As usual, the industry input tariffs are calculated using simple-average tariffs within each Chinese Industrial Classification (CIC) 2-digit industry level as shown in Equ. (7). Although taking the simple average across product within an industry seems straightforward, it bears a cost as the import heterogeneity for products within the industry is ignored. For example, suppose a firm imports 70% of lumber and 30% of steel. Tariffs on lumber are apparently more important to the firm than those on steel. However, a simple-average tariffs cannot take such a difference into account. To address this issue, we consider the following weighted input tariffs:

$$\tau_{nt} = \sum\nolimits_{k \in n} \left(\frac{m_{kt}}{\sum_{k \in n} m_{kt}} \right) \tau_{kt}, \tag{9}$$

where m_{kt} is the import values for product k within a CIC 2-digit industry n in year t. Once we obtain these weighted input tariffs, we plug them back into Equ. (6) to obtain the weighted industry input tariffs ($wiit_{it}$).[20]

Table 5 reports the estimates using weighted industry input tariffs. The fixed-effects OLS estimates in column (1) show that the coefficient of input tariffs is still negative and significant after considering the importance of import heterogeneity within an industry. To rule out possible estimation bias due to the inclusion of processing imports, column (2) removes pure exporters from the sample and we still obtain the results similar to those in column (1). Columns (3)-(4) perform the 2SLS estimations to control for the possible endogeneity of the weighted input tariffs. A one-period lag of industry input tariffs is served as the IV of the first-difference 2SLS estimates with a consequent change. The simple average tariffs calculated in Equ. (7) is replaced with weighted average tariffs in Equ. (9). After controlling for the endogeneity, Column (3) shows that the coefficient of weighted

[20] There is a caveat to this. As pointed out by Topalova and Khandelwal (2011), such a weighted industry input tariff may understate the actual input tariff reduction since the imported inputs with lower tariffs may receive higher import volume and thus have higher weights in Eq.(9). Therefore, the calculation using weighted tariffs and the associated estimations in Table 5 should be treated as lower-bound estimates of the effects of input tariffs on wage inequality.

industry input tariffs is still negative, though insignificant. We then drop pure exporters in column (4) and find that the coefficient of input tariffs turns to be negative and significant, suggesting that industry input trade liberalization tends to widen firm-level wage inequality.

[Insert Table 5 Here]

3.6 Further Estimates with Measured Relative Wages

Thus far, a firm's wage inequality is measured in an absolute term as the wage difference between skilled and unskilled labor. It is worthwhile to check whether our estimates are robust when the wage inequality is measured as relative wages between skilled and unskilled labor a là Feenstra and Hanson (1996, 1999). Table 6 performs this task.[21] The regressand in all estimates except column (5) is the ratio of a firm's skilled wage over unskilled wage. As seen in Table 1A, the overall annual relative wages during the sample period is 2.21, which is significantly higher than that in the U.S.(1.75). Column (3) of Table 2 reports the relative wage ratios by industry. The OLS estimates in column (1) of Table 6 and the 2SLS estimates in column (2) of Table 6 cover the seven-year sample (2000 to 2006) and, once again, find that input tariff reductions widen a firm's wage inequality. However, the magnitude of the coefficient of input tariffs seems to be too small and the key coefficient of industry input tariffs in column (2) is insignificant. We suspect that this is largely because of the spread of the provincial share of skilled labor in such a long time window. Therefore, we run the 2SLS estimates in column (3) with data from 2003 to 2005. But the coefficient of input tariffs is still negative and significant.

Since the observations of measured wage inequality (in both absolute and relative terms) are estimated but not observed, it is worthwhile to control for the fact that some observations are estimated more precisely than others. Therefore, we compute the standard deviation of a firm's relative wages across firms within an industry and multiply it with a firm's relative wage as the

[21] Firm's measured relative wages can be estimated as follows. Since $w^s_{ijt} - w^u_{ijt} = \alpha_{jt} - \beta_{jt}\pi_{ijt}$ and $\overline{w}_{ijt} \equiv \theta_{ijt}w^s_{ijt} + (1-\theta_{ijt})w^u_{ijt}$, so we can solve these two equations to obtain: $w^s_{ijt} = \overline{w}_{ijt} + (1-\theta)(\alpha_{jt} + \beta_{jt}\pi_{ijt})$ and $w^u_{ijt} = \overline{w}_{ijt} - \theta(\alpha_{jt} + \beta_{jt}\pi_{ijt})$. Therefore, firm's measured relative wages w^s_{ijt}/w^u_{ijt} is able to be traced back once α_{jt} and β_{jt} are estimated.

regressand in Table 6 (refer to weighted relative wages $\widehat{wrwage}_{ijt}$ with a relatively large mean, 2.89, as reported in Table 1A). The last two columns of Table 2 report the standard deviation and the mean of weighted relative wages by two-digit level Chinese industry. The coefficient of input tariffs is negative and significant again.

Our last step is to offer a more intuitive economic interpretation for our estimation results. As shown in the first column of Table 6, the coefficient of the industry input tariffs is -0.033, implying that a 10 percentage point fall in input tariffs leads to a 0.3 point increase in relative wage inequality. Average input tariffs were cut by about 7.39 percentage points (from 15.76 percent in 2000 to 8.37 percent in 2006). Thus, this predicts a $0.033 \times 7.39 = 0.24$ percentage points' increase in a firm's relative wages and accounts for approximately 8% of the difference in firm's measured relative wages, 3.06, between 2000 and 2006. By controlling for the possible reverse causality and possible measurement errors using estimating sample of wage inequality, the coefficient of the first-difference in industry input tariffs is -0.205, as shown in column (4) of Table 6. Simultaneously, average input tariffs were cut by only 1.17 percentage points (from 9.75 percent in 2003 to 8.58 percent in 2005). Thus, this predicts a $0.205 \times 1.17 = 0.24$ point increase in a firm's relative wages and accounts for approximately 33.7% of the difference in firm's measured relative wages, 0.71 (from 2.80 in 2003 to 3.51 in 2005). Hence, input trade liberalization contributes to around 8-34% of measured wage inequality. These results are comparable to that in Feenstra and Hanson (1999), which posits the figure of 15-40% for the impact of outsourcing on wage inequality in the United States in the 1980s.

[Insert Table 6 Here]

3.7 Estimates with Alternative Measured Wage Inequality

To obtain within-firm wage inequality, we rely on the argument of fair wages. Firms tend to allocate part of the profits to its skilled workers. Since larger firms usually have more profits, we divide firm profits by firm sales to capture profitability and use this to estimate within-firm wage inequality as in Equ. (5). In order to know whether our main findings are sensitive to the measure of profitability, we replace a firm's profits-sales ratio with total profits and re-estimate Equ. (2) to obtain within-firm measured wage inequality (refer to $\widehat{wgap2}_{ijt}$). By using this alternative measured wage inequality as the regressand, Table 7 runs fixed-effects regressions with different specifications. Column (1) includes all samples during the period 2000 to 2006, whereas column (2) excludes pure exporters. We see that declining input tariffs lead to an increase in a firm's measured wage inequality. Column (3) takes a further step by replacing the level of the measured wage inequality with the first-difference in measured wage inequality and we obtain similar results. To rule out the possibility that such results are due to the adoption of provincial skilled share as the multiplier, estimates in column (4) include data in 2004 only, and those in column (5) include data in the shorter period over 2003-2005. To show that our main results are insensitive to different productivity measures, we use conventional Olley-Pakes TFP in all estimates. Nevertheless, all specifications confirm that input trade liberalization widens a firm's measured wage inequality.

[Insert Table 7 Here]

3.8 Estimates with Industry Minimum Wages

The within-firm measured wage inequality relies heavily on firm profits as proxy. In this section, we use another proxy–industrial minimum wages–to back up our alternative index of measured wage inequality. In this way, we even do not necessarily need to rely on the fair-wage argument.

Previous works usually treat unskilled wages as a premium on industrial minimum wages (Anwar and Sun, 2012). Hence, we construct industrial minimum wages as follows: The minimum firm-level average wages *within* each four-digit Chinese industry is labeled as its industrial minimum wages. The economic rationale is that every firm must pay wages which at least higher than the minimum wages imposed by the states. Of course, since in reality all firms (including the marginal firm in each industry) shall pay a premium over the minimum wages. Our proxy of industrial minimum wages shall be treated as a upper bound of the *de jure* minimum wages.

We consider the following specification for unskilled wages $w^u_{ijt} = w^{\min}_{jt}(1 + s_{ijt})$, where $w^{\min}_{jt}$ is the minimum wage of four-digit industry j in year t and s_{ijt} is the premium set by firm i of four-digit industry j in year t. Inserting such a wage premium equation into Equ. (3) of a firm's average wages, we have

$$\overline{w}_{ijt} = \theta_{ijt} w_{ijt}^{s} + (1-\theta_{ijt}) w_{jt}^{\min}(1+s_{jt}). \tag{10}$$

By allowing firm-level wage heterogeneity for both skilled (ε_{ijt}^{s}) and unskilled labor (ε_{ijt}^{u}) within each industry, we have

$$\begin{aligned}\overline{w}_{ijt} &= \theta_{ijt}(w_{jt}^{s} + \varepsilon_{ijt}^{s}) + (1-\theta_{ijt})(w_{jt}^{\min}(1+s_{jt}) + \varepsilon_{ijt}^{u}). &(11)\\ &= \theta_{ijt} w_{jt}^{s} + (1-\theta_{ijt})(1+s_{jt}) w_{jt}^{\min} + \xi_{ijt} &(12)\end{aligned}$$

By absorbing all terms with wage residuals $\xi_{ijt} \equiv \theta_{ijt}\varepsilon_{ijt}^{s} + (1-\theta_{ijt})\varepsilon_{ijt}^{u}$ into the error term, we can estimate the following equation for each four-digit industry j in different years t with data on skilled labor share θ_{ijt} and industry minimum wage $w_{jt}^{\min}$:

$$\widehat{\overline{w}}_{ijt} = \hat{\gamma}_{1jt}\theta_{ijt} + \hat{\gamma}_{2jt}\left((1-\theta_{ijt}) w_{jt}^{\min}\right) \tag{13}$$

where the estimated coefficient $\hat{\gamma}_{1jt}$ denotes industrial skilled wages (w_{jt}^{s}) and $\hat{\gamma}_{2jt}$ corresponds to the average industrial wage premium $(1 + s_{jt})$ for industry j at year t. Once the measured wage inequality is obtained by backing up the coefficients $\hat{\gamma}_{1jt}$ and $\hat{\gamma}_{2jt}$, we then obtain the CIC 4-digit industry level wages inequality, $\widehat{wgap3}_{jt}$ (See Appendix B for details).

$$\widehat{wgap3}_{jt} \equiv w_{jt}^{s} - w_{jt}^{u} = \hat{\gamma}_{1jt} - \hat{\gamma}_{2jt} w_{jt,}^{\min}. \tag{14}$$

Table 1 in the appendix reports the mean of industrial wages inequality at an aggregated CIC two-digit industry level. By way of comparison, the measured firm-level wage inequality ($\widehat{wgap1}_{ijt}$) is at a more disaggregated level, but it has to rely on firm's profitability (either measured by profit-sales ratio or total profit). The measured four-digit *industry* level wage inequality ($\widehat{wgap3}_{jt}$) is more flexible and needs not depending on the "fair wages" argument, although it is not able to capture the wages inequality across firms within an industry. Nevertheless, it is still worthwhile to serve as a robustness check for our main question: whether or not input trade liberalization widens wage inequality.

Table 8 presents the estimation results using such industry-level wage inequality as the regressand. Column (1) starts with the OLS estimates. Note that the regressand $\widehat{wgap3}_{jt}$ is measured at the CIC 4-digit level. As a result, the number of observations is reduced to 1,750. We include weighted industry input tariffs ($wiit_{jt}$) and industry output tariffs in all estimates. We also include industry-level average log TFP with one lag to see whether industrial productivity affects the industrial wage gap. It turns out that the coefficient of industry input tariffs is negative but insignificant. We suspect that this is due to the lack of controlling fixed effects. We therefore run the year-specific and industry-specific fixed effects in the rest of Table 8. Estimates in column (2) show that the coefficient of industry input tariffs now become significant. In columns (3) and (4) we include industry-average log employment to control for industry size and still find that input tariff reductions widen wage inequality. As shown in column (4), such a finding is still robust if we exclude pure exporters. Thus, our main findings are robust to different measures of wages inequality and industrial input tariffs.

[Insert Table 8 Here]

4 Concluding Remarks

China has experienced a dramatic tariff reduction since its accession to WTO in 2001. On the other hand, China's wage inequality and, more broadly, income inequality had also increased over the years. To our knowledge, so far there is no micro-level evidence to explore the linkage between the two. Since there is no actual firm-level data on wages for both skilled and unskilled labor, we have to develop different econometric approaches to estimate within-firm wage inequality based on imperfect Chinese firm-level data on wage information. As in other ambitious attempts to investigate important issues with imperfect data, we have to make a number of compromises to conduct our estimates. Nevertheless, our results that a fall in input trade costs leads to an increase in measured wage inequality are quite robust under different econometric specifications.

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Table 1A: Summary Statistics of Key Variables (2000-2006)

Variables	Mean	Std. Dev.
Measured Firm Wages Inequality (RMB 1,000: $\widehat{wgap1}_{ijt}$)	11.32	352.7
Measured Firm Wages Inequality (RMB 1,000: $\widehat{wgap2}_{ijt}$)	9.59	280.2
Measured Industry Wages Inequality (RMB 1,000: $\widehat{wgap3}_{ijt}$)	5.23	24.09
Measured Firm Relative Wages ($\widehat{rwage}_{ijt}$)	2.21	2.09
Measured Weighted Firm Relative Wages ($\widehat{wrwage}_{ijt}$)	2.89	4.81
Industry Input Tariffs (%)	9.72	2.97
Weighted Industry Input Tariffs (%)	9.15	3.22
Industry Output Tariffs (%)	11.07	8.19
Log of Firm Labor	4.90	1.10
SOEs Indicator	.055	.228
Foreign Indicator	.222	.415
Pure Exporters	.044	.207

Notes: RMB 1 is equivalent to \$0.125 during the sample period.

Table 1B: China's Industrial Input Tariffs

Year	2000	2001	2002	2003	2004	2005	2006	Average
Ind. Input Tariffs	15.76	14.38	11.09	9.75	8.87	8.58	8.37	9.72
Std. Dev.	4.01	3.35	2.64	2.13	1.76	1.63	1.54	2.97

Notes: This table reports the mean and standard deviation of 3-digit industry-level input tariffs.

Table 2: Measured Within-firm Wages Inequality

Adjusted Chinese Industrial Classifications	$\widehat{wgap1}_i$	$\widehat{wgap2}_i$	$\widehat{rwage}_i$	Std.Dev	$\widehat{wrwage}_i$
Processing of Foods (13)	4.546	4.658	2.025	1.765	3.573
Manufacturing of Foods (14)	11.58	8.980	2.907	2.861	8.316
Manufacture of Beverages (15)	7.704	7.445	2.617	2.756	7.214
Manufacture of Tobacco (16)	109.1	47.96	3.533	3.259	11.51
Manufacture of Textile (17)	4.803	5.535	1.859	1.609	2.992
Manufacture of Apparel, Footwear,Caps (18)	3.428	3.424	1.570	0.996	1.564
Manufacture of Leather, Fur, Feather (19)	3.768	2.880	1.849	1.767	3.267
Processing of Timber, Manufacture of Wood, Bamboo, Rattan, Palm,Straw Products (20)	2.107	4.340	1.975	1.856	3.667
Manufacture of Furniture (21)	2.933	6.308	1.606	1.299	2.086
Manufacture of Paper and Paper Products (22)	13.96	9.158	2.293	2.003	4.594
Printing, Reproduction of Recording Media (23)	4.581	4.575	1.715	1.211	2.076
Mfg. For Culture, Education, Sports (24)	9.356	14.08	2.151	2.141	4.604
Processing of Petroleum, Coking, Fuel (25)	11.71	11.51	3.002	3.090	9.274
Manufacture of Raw Chemical Materials (26)	13.03	11.88	3.029	2.825	8.558
Manufacture of Medicines (27)	12.56	11.36	2.957	3.124	9.236
Manufacture of Chemical Fibers (28)	11.41	9.449	2.315	2.160	5.000
Manufacture of Rubber (29)	5.305	5.311	1.835	1.322	2.426
Manufacture of Plastics (30)	6.788	6.836	2.050	1.611	3.303
Manufacture of Non-metallic Mineral goods (31)	4.244	4.125	1.834	1.429	2.622
Smelting Pressing of Ferrous Metals (32)	5.768	5.534	1.784	1.270	2.266
Smelting Pressing of Non-ferrous Metals (33)	6.590	6.577	2.035	1.697	3.454
Manufacture of Metal Products (34)	6.868	6.723	1.995	1.608	3.207
Manufacture of General Purpose Machinery (35)	8.005	7.609	2.236	2.035	4.552
Manufacture of Special Purpose Machinery (36)	11.92	11.05	2.770	2.613	7.239
Manufacture of Transport Equipment (37)	10.65	8.815	2.361	2.308	5.449
Electrical Machinery Equipment (39)	8.220	9.327	2.518	2.287	5.759
Computers Other Electronic Equipment (40)	15.98	16.43	3.120	2.997	9.352
Manufacture of Measuring Instruments (41)	15.25	13.93	2.966	2.911	8.634
Manufacture of Artwork (42)	21.12	10.22	2.172	1.994	4.331

Notes: Unit in columns (1) and (2) is RMB 1,000 (equivalent to $125). Standard errors for each coefficient are not reported to save space though available upon request. The wage inequality index $\widehat{wgap1}_i$ (and the alternative wage inequality index $\widehat{wgap2}_i$) is computed by Equ. (5) with profit-sales ratio (firm's total profit) as a proxy of firm's profitability. Firm's relative wages ($\widehat{rwage}_i$) is the ratio of firm's skilled wages over unskilled wages which are calculated by Equ. (5) with profit-sales ratio as a proxy of firm's profitability. Standard deviation of firm's relative wages across firms within an industry are reported in the second last column. The last column ($\widehat{wrwage}_i$) is obtained by using industrial standard deviation as in the second last column to multiply firm's relative wages by industry.

Table 3: 2SLS Estimates using Measured Within-firm Wage Inequality

Econometric Method:	OLS			2SLS	
Regressand:	(1)	(2)	(3)	(4)	(5)
Measured Firm's Wage Inequality	$\widehat{wgap1}_{ijt}$	$\widehat{wgap1}_{ijt}$	$\widehat{wgap1}_{ijt}$	$\Delta\widehat{wgap1}_{ijt}$	$\Delta\widehat{wgap1}_{ijt}$
Industry Input Tariffs	-0.264***	-1.065***	-1.146***	-1.539***	-1.869***
	(-3.29)	(-6.16)	(-6.22)	(-5.94)	(-7.10)
Industry Output Tariffs		0.114***	0.117***	0.162***	0.158***
		(8.68)	(8.66)	(12.33)	(11.82)
State-owned Enterprises		1.119	1.138	3.363***	3.383***
		(0.88)	(0.90)	(3.38)	(3.38)
Foreign Firms		0.024	0.097	-0.240	-0.256
		(0.09)	(0.37)	(-0.26)	(-0.26)
Log of Firm Employment		0.084	0.160	-0.159	-0.109
		(0.62)	(1.12)	(-0.66)	(-0.44)
Lag of Firm Relative TFP		0.793**	0.822**	0.245	0.195
		(2.53)	(2.50)	(0.42)	(0.32)
Kleibergen-Paap rk LM χ^2 statistic	–	–	–	30,626†	30,195†
Kleibergen-Paap Wald rk F statistic	–	–	–	41,094†	40,911†
Year-Specific Fixed Effects	Yes	Yes	Yes	Yes	Yes
Firm-Specific Fixed Effects	Yes	Yes	Yes	Yes	Yes
Pure Exporters Included	Yes	Yes	No	Yes	No
Observations	526,969	233,613	222,367	120,235	115,284
R-squared	0.02	0.02	0.01	0.01	0.01
		First-Stage Regressions			
IV: One-Lag Industry Input Tariffs	–	–	–	-.161***	-.163***
				(-202.7)	(-202.3)

Notes: Robust t-values corrected for clustering at the firm level in parentheses.*,**(***) indicates significance at the 10,5 and 1 percent level, respectively. $^{\dagger}(^{\ddagger})$ indicates significance of p-value at the 1(5) percent level. Regressands in the OLS estimates of columns (1)-(3) are levels of firm's wages inequality ($\widehat{wgap1}_{ijt}$) whereas those in the 2SLS estimates of columns (4)-(5) are the first difference in firm's wage inequality ($\Delta\widehat{wgap1}_{ijt}$). Correspondingly, regressors in columns (1)-(3) are in levels whereas those in columns (4)-(5) are in the first difference. IV reports the coefficient of one-lag industry input tariffs using the first difference in firm's wage inequality as the regressand. Columns (1), (2), and (4) include all sample. Columns (3) and (5) include all sample except pure exporters.

Table 4: Cross-Section and Shorter Panel Estimates

Econometric Methods:	OLS	2SLS	OLS	2SLS
Regressand: Measured Firm's Wage Gap	$\widehat{wgap1}_{ijt}$	$\widehat{wgap1}_{ijt}$	$\widehat{wgap1}_{ijt}$	$\Delta\widehat{wgap1}_{ijt}$
	(1)	(2)	(3)	(4)
Industry Input Tariffs	-0.844***	-1.520***	-1.913***	-3.274***
	(-7.32)	(-6.13)	(-8.21)	(-10.10)
Industry Output Tariffs	0.316***	0.300***	0.001	-0.054**
	(7.91)	(16.69)	(0.07)	(-2.56)
State-owned Enterprises	2.932**	2.987***	4.857*	4.979***
	(2.52)	(6.32)	(1.93)	(3.55)
Foreign Firms	0.202	0.217	0.030	-0.847
	(1.38)	(0.99)	(0.05)	(-0.54)
Log of Firm Employment	0.101	0.110	0.232	-0.048
	(1.32)	(1.33)	(0.86)	(-0.13)
Lag of Firm Relative TFP	1.064***	1.101*	4.690***	0.589
	(3.55)	(1.79)	(6.92)	(0.64)
Kleibergen-Paap rk LM χ^2 statistic	–	15,179†	–	17,198†
Kleibergen-Paap Wald rk F statistic		22,915†		24,334†
Year-specific Fixed Effects	No	No	Yes	Yes
Firm-specific Fixed Effects	No	No	Yes	Yes
2-digit Industry Fixed Effects	Yes	Yes	No	No
Years Coverage	2004		2003-2005	
Observations	45,731	44,967	136,643	58,650
R-squared	0.57	0.57	0.01	0.01
First-Stage Regressions				
IV: One-Lag Industry Input Tariffs	–	.354***	–	-.192***
		(151.3)		(-155.9)

Notes: Robust t-values corrected for clustering at the firm level in parentheses. *** (**,*) denotes the significance at 1% (5%, 10%) level. Columns (1)-(2) include sample in 2004 only. Columns (3)-(4) include sample during 2003-2005. In the first-stage estimates of column (2), IV reports the coefficient of one-lag industry input tariffs using current industry input tariffs as the regressand. The regressand in the 2SLS estimates of column (4) is the first difference in firm's wage inequality ($\Delta\widehat{wgap1}_{ijt}$) and all regressors are in the first difference. In the first-stage estimates of column (4), IV reports the coefficient of one-lag industry input tariffs using the first difference in firm's wage inequality as the regressand.

Table 5: More 2SLS Estimates using Alternative *Tariffs* Measure

Econometric Methods:	OLS		2SLS	
Measured Firm's Wage Inequality	$\widehat{wgap1}_{ijt}$	$\widehat{wgap1}_{ijt}$	$\Delta\widehat{wgap1}_{ijt}$	$\Delta\widehat{wgap1}_{ijt}$
	(1)	(2)	(3)	(4)
Weighted Industry Input Tariffs ($wiit_{jt}$)	-0.818***	-0.886***	-0.129	-0.648*
	(-4.78)	(-4.90)	(-0.57)	(-1.88)
Industry Output Tariffs	0.106***	0.108***	0.145***	-0.065***
	(8.46)	(8.47)	(10.62)	(-3.00)
State-owned Enterprises	1.007	1.020	3.521***	5.280***
	(0.79)	(0.80)	(3.49)	(3.67)
Foreign Firms	-0.036	0.036	-0.252	-0.903
	(-0.16)	(0.15)	(-0.25)	(-0.52)
Log of Firm Employment	-0.002	0.032	-0.087	0.065
	(-0.02)	(0.24)	(-0.35)	(0.16)
Lag of Firm Relative TFP	-0.488**	-0.503**	-0.233	-0.438
	(-2.32)	(-2.29)	(-0.92)	(-1.07)
Kleibergen-Paap rk LM χ^2 statistic	–	–	23,847†	23,848†
Kleibergen-Paap Wald rk F statistic			30,067†	300,68†
Year-specific Fixed Effects	Yes	Yes	Yes	Yes
Firm-specific Fixed Effects	Yes	Yes	Yes	Yes
Pure Exporter Included	Yes	No	Yes	No
Observations	233,613	222,367	115,284	56,043
R-squared	0.01	0.01	0.01	0.01
First-Stage Regressions				
IV: One-Lag Weighted Industry Input Tariffs	–	–	-0.172***	-0.172***
			(-173.3)	(-173.4)

Notes: Robust t-values corrected for clustering at the firm level in parentheses. *** (**,*) denotes the significance at 1% (5%, 10%) level. Columns (1) and (3) include the entire sample whereas columns (2) and (4) include the entire sample except pure exporters. The regressands in the 2SLS estimates of columns (3)-(4) are the first difference of firm's wage inequality ($\Delta\widehat{wgap1}_{ijt}$) and all regressors are in the first difference. In the first-stage estimates of columns (3) and (4), IV reports the coefficient of one-lag industry input tariffs using the first difference of firm's wage inequality as the regressand.

Table 6: 2SLS Estimates using Measured Within-firm *Relative* Wage Inequality

Econometric Method:	OLS	2SLS		
Regressand:	$\widehat{rwage}_{ijt}$	$\Delta \widehat{rwage}_{ijt}$	$\Delta \widehat{rwage}_{ijt}$	$\Delta \widehat{wrwage}_{ijt}$
	(1)	(2)	(3)	(4)
Industry Input Tariffs	-0.033***	-0.023	-0.053*	-0.205***
	(-3.56)	(-0.97)	(-1.93)	(-3.32)
Industry Output Tariffs	0.005***	0.006***	0.014***	0.027***
	(5.40)	(5.80)	(9.08)	(7.51)
State-owned Enterprises	-0.069	-0.097	-0.067	-0.094
	(-0.87)	(-1.10)	(-0.61)	(-0.38)
Foreign Firms	-0.130**	-0.070	0.081	0.525*
	(-2.02)	(-0.82)	(0.62)	(1.78)
Log of Firm Employment	0.279***	0.402***	0.200***	0.396***
	(13.11)	(18.12)	(6.35)	(5.59)
Lag of Firm Relative TFP	-0.020	0.075	0.286***	0.455***
	(-0.40)	(1.45)	(3.88)	(2.75)
Kleibergen-Paap rk LM χ^2 statistic		27,359†	14,621†	14,622†
Kleibergen-Paap Wald rk F statistic		37,777†	21,203†	21,204†
Year-Specific Fixed Effects	Yes	Yes	Yes	Yes
Firm-Specific Fixed Effects	Yes	Yes	Yes	Yes
Pure Exporters Included	Yes	No	No	No
Year Coveraged	2000-2006		2003-2005	
Observations	213,196	99,211	47,102	47,102
R-squared	0.02	0.02	0.03	0.02
First-Stage Regressions				
IV: One-Lag Industry Input Tariffs	–	-.167***	-.196***	-.196***
		(-194.3)	(-145.6)	(-145.6)

Notes: Robust t-values corrected for clustering at the firm level in parentheses.*,**(***) indicates significance at the 10,5 and 1 percent level, respectively. $^\dagger(^\ddagger)$ indicates significance of p-value at the 1(5) percent level. Regressands in Columns (1)-(3) are firm-level firm's relative wages, $\widehat{rwage}_{ijt}$, which is defined as computed skilled wages over unskilled wages whereas that in Column (4) is the firm's relative wages multiplied by its sectorial standard deviation ($\widehat{wrwage}_{ijt}$). IV reports the coefficient of one-lag industry input tariffs using current industry input tariffs as the regressand. Columns (1)-(2) include the entire sample. Columns (3)-(4) include the sample during 2003-2005. In the first-stage estimates of columns (2)-(4), IV reports the coefficient of one-lag industry input tariffs using the first difference of firm's wage inequality as the regressand.

Table 7: Estimates using Alternative Within-firm Wage Inequality

Regressand:	(1)	(2)	(3)	(4)	(5)
Measured Firm's Wage Inequality	$\widehat{wgap2}_{ijt}$	$\widehat{wgap2}_{ijt}$	$\Delta\widehat{wgap2}_{ijt}$	$\widehat{wgap2}_{ijt}$	$\widehat{wgap2}_{ijt}$
Industry Input Tariffs	-0.135***	-0.126***	-0.533***	-0.627***	-0.726***
	(-14.34)	(-12.71)	(-24.63)	(-27.69)	(-31.48)
Industry Output Tariffs	0.051***	0.050***	0.017***	0.234***	0.045***
	(33.41)	(32.08)	(11.18)	(68.63)	(18.05)
State-owned Enterprises	-0.100	-0.105	0.139	0.087	0.008
	(-0.98)	(-1.03)	(1.15)	(1.10)	(0.06)
Foreign Firms	-0.024	-0.086	-0.003	0.164***	-0.088
	(-0.26)	(-0.89)	(-0.03)	(5.24)	(-0.69)
Log of Firm Employment	0.236***	0.237***	0.166***	0.156***	0.123***
	(10.41)	(10.04)	(6.09)	(12.52)	(3.65)
Log of Firm TFP	0.195***	0.201***	0.124***	0.587***	0.144***
	(6.78)	(6.71)	(3.87)	(12.98)	(3.53)
Year-specific Fixed Effects	Yes	Yes	Yes	Yes	Yes
Firm-specific Fixed Effects	Yes	Yes	Yes	Yes	Yes
Industry-specific Fixed Effects	No	No	No	Yes	No
Pure Exporters Included	Yes	No	Yes	Yes	Yes
Year Coverage		2000-2006		2004	2003-2005
Observations	366,356	345,807	207,541	96,226	232,411
R-squared	0.18	0.18	0.14	0.33	0.25

Notes: Robust t-values corrected for clustering at the firm level in parentheses. *** (**,*) denotes the significance at 1% (5%, 10%) level. Columns (1) and (3) include the entire sample, whereas column (2) includes the entire sample except pure exporters. Column (4) includes data in 2004 only. Column (5) includes data in 2003-2005. Regressands in all columns except column (3) are levels of firm's wages inequality ($\widehat{wgap2}_{ijt}$), whereas that in column (3) is the first difference of firm's wage inequality ($\Delta\widehat{wgap2}_{ijt}$). Correspondingly, regressors in all columns except column (3) are in levels, whereas those in column (3) are in the first difference.

Table 8: More Estimates using Alternative Industrial Wage Inequality

Measured Industry Wage Inequality ($\widehat{wgap3}_{jt}$)	(1)	(2)	(3)	(4)
Weighted Industry Input Tariffs ($wiit_{jt}$)	-0.119	-0.548***	-0.546***	-0.625***
	(-1.59)	(-2.98)	(-2.96)	(-3.27)
Industry Output Tariffs	-0.001	0.051	0.051	0.065
	(-0.03)	(1.20)	(1.21)	(1.47)
Industry-Level Log Employment			0.094	0.211
			(0.39)	(0.83)
Industry-Level Log TFP with One-Lag	0.596			
	(0.49)			
Year-specific Fixed Effects	No	Yes	Yes	Yes
Firm-specific Fixed Effects	No	Yes	Yes	Yes
Pure Exporters Included	Yes	Yes	Yes	No
Observations	1,750	1,750	1,750	1,657
R-squared	0.01	0.03	0.03	0.04

Notes: Robust t-values corrected for clustering at the firm level in parentheses. *** (**,*) denotes the significance at 1% (5%, 10%) level. The regressand is measured industry-level wage gap as discussed in Equ. (14) in the text.

Appendix

Appendix A: Estimating Measured Wage Inequality from Firm Profits and Labor Shares

In this appendix we describe the method that we use to construct firm-level wage inequality from information of firm profits and labor shares. We start from the derivation of measured firm-level wage inequality. Notice that firm i's average wage in industry j at year t can be expressed as

$$\begin{aligned}
\overline{w}_{ijt} &= \theta_{ijt} w^s_{ijt} + (1-\theta_{ijt}) w^u_{ijt} \\
&= \theta_{ijt}(w^s_{jt} + \varepsilon^s_{ijt}) + (1-\theta_{ijt})(w^u_{jt} + \varepsilon^u_{ijt}) \\
&= \theta_{ijt}(w^u_{jt} + \alpha_{jt}) + \theta_{ijt}\varepsilon^s_{ijt} + (1-\theta_{ijt}) w^u_{jt} + (1-\theta_{ijt})\varepsilon^u_{ijt} \\
&= w^u_{jt} + \theta_{ijt}\alpha_{it} + \theta_{ijt}(\varepsilon^s_{ijt} - \varepsilon^u_{ijt}) + \varepsilon^u_{ijt} \\
&= w^u_{jt} + \alpha_{it}\theta_{ijt} + \beta_{jt}(\theta_{ijt}\pi_{ijt}) + \varepsilon^u_{ijt}.
\end{aligned} \tag{15}$$

The second equality follows the definition of $w^s_{ijt} = w^s_{jt} + \varepsilon^s_{ijt}$ and $w^u_{ijt} = w^u_{jt} + \varepsilon^u_{ijt}$. The third equality is due to within-industry wage differential $w^s_{jt} - w^u_{jt} = \alpha_{jt}$. Rearranging the fourth equality, we can easily obtain the last equality by using the equation of within-firm wage differential $\varepsilon^s_{ijt} - \varepsilon^u_{ijt} = \beta_{jt}\pi_{ijt}$. Finally, the error term is assumed to be orthogonal to firm profitability. Therefore, the firm-level wage inequality is calculated using the estimated coefficients $\widehat{\alpha}_{jt}$ and $\widehat{\beta}_{jt}$

$$\widehat{wgap1}_{ijt} = \widehat{\alpha}_{jt} + \widehat{\beta}_{jt}\pi_{ijt}.$$

Appendix B: The Measured Wage Inequality using Minimum Wages as a Proxy

Alternatively, we can estimate and calculate the industry-level wage inequality ($\widehat{wgap3}_{jt}$) as follows. Consider the following specification for unskilled wage $w^u_{ijt} = w^{\min}_{jt}(1 + s_{ijt})$, where $w^{\min}_{ijt}$ is the minimum wage and s_{ijt} is the premium set by firm i of four-digit industry j at year t. Inserting this equation of wage premium to Equ. (3) of firm i's average wage, we have:

$$\overline{w}_{ijt} = \theta_{ijt} w^s_{ijt} + (1-\theta_{ijt}) w^{\min}_{jt}(1 + s_{ijt}). \tag{16}$$

By allowing firm-level wage heterogeneity for both skilled (ε^s_{ijt}) and unskilled labor (ε^u_{ijt}) within each industry, we have

$$\overline{w}_{ijt} = \theta_{ijt}(w^s_{jt} + \varepsilon^s_{ijt}) + (1-\theta_{ijt})(w^{\min}_{jt}(1 + s_{ijt}) + \varepsilon^u_{ijt}). \tag{17}$$

Therefore, we can estimate the following equation for each four-digit industry j in different years t

$$\widehat{\overline{w}}_{ijt} = \hat{\gamma}_{1jt}\theta_{ijt} + \hat{\gamma}_{2jt}(1-\theta_{ijt}) w^{\min}_{jt}, \tag{18}$$

where the estimated coefficient $\hat{\gamma}_{1jt}$ denotes industrial skilled wage and $\hat{\gamma}_{2jt}$ is corresponding to the industrial wage premium $(1 + s_{it})$ for industry j at year t. Notice that

$$w^s_{ijt} - w^u_{ijt} = (w^s_{jt} - w^u_{jt}) + (\theta_{ijt}\varepsilon^s_{ijt} - (1-\theta_{ijt})\varepsilon^u_{ijt}). \tag{19}$$

After Equ. (18) is estimated, we combine it with Equ. (19) to obtain:

$$(w^s_{jt} - w^u_{jt}) = E(\boldsymbol{\beta}\mathbf{X}_{it}|\mathbf{X}_{it}) + (\epsilon_{it} - \theta_{ijt}\varepsilon^s_{ijt} - (1-\theta_{ijt})\varepsilon^u_{ijt}).$$

Therefore, the measured wage inequality can be computed as follows (using $\hat{\gamma}_{1jt}$ and $\hat{\gamma}_{2jt}$):

$$\begin{aligned} \widehat{wgap3}_{jt} &\equiv \hat{\gamma}_{1jt} - \hat{\gamma}_{2jt} w_{jt,}^{\min} \\ &= E(\boldsymbol{\beta}\mathbf{X}_{it}|\mathbf{X}_{it}) + (\epsilon_{it} - \theta_{ijt}\varepsilon_{it}^{s} - (1-\theta_{ijt})\varepsilon_{it}^{u}), \\ &= \phi_0 + \phi_1 IIT_{jt} + \varpi_j + \gamma_t + \mu_{it} \end{aligned} \tag{20}$$

where the error term in Equ. (20) can be decomposed into three terms as in Equ. (11): (i) a industry-specific fixed effect ϖ_i to control for time-invariant factors such as a firm's managerial ability; (ii) a year-specific fixed effect η_t to control for firm-invariant factors such as Chinese RMB appreciation; and (iii) an error term μ_{it} for other unspecified factors.

Appendix Table 1: Estimated Skilled and Unskilled Wages of Chinese Firms

Adjusted Chinese Industrial Classfications	Skilled Wages $\hat{\gamma}_{1jt}$	Unskilled Premium $\hat{\gamma}_{2jt}$	Unskilled Wages $\hat{\gamma}_{2jt}w_{jt}^{\min}$	Measured Wage Inequality $\hat{\gamma}_{1jt}-\hat{\gamma}_{2jt}w_{jt}^{\min}$
Processing of Foods (13)	16.43	7.73	3.58	12.85
Manufacturing of Foods (14)	17.74	6.37	5.48	12.26
Manufacture of Beverages (15)	16.06	7.78	4.64	11.42
Manufacture of Tobacco (16)	36.69	4.40	7.07	29.62
Manufacture of Textile (17)	19.48	8.71	4.21	15.27
Manufacture of Apparel, Footwear & Caps (18)	22.32	21.85	3.21	19.11
Manufacture of Leather, Fur, & Feather (19)	17.09	9.33	7.35	9.74
Processing of Timber, Manufacture of Wood, Bamboo, Rattan, Palm & Straw Products (20)	16.50	11.13	6.02	10.48
Manufacture of Furniture (21)	21.66	5.47	3.88	17.78
Manufacture of Paper & Paper Products (22)	19.46	14.25	4.38	15.07
Printing, Reproduction of Recording Media (23)	19.81	15.76	6.51	13.29
Mfg. For Culture, Education & Sport (24)	20.85	3.57	5.42	15.42
Processing of Petroleum, Coking, &Fuel (25)	21.84	4.73	3.60	18.23
Manufacture of Raw Chemical Materials (26)	20.95	6.38	5.07	15.87
Manufacture of Medicines (27)	17.63	16.58	5.91	11.72
Manufacture of Chemical Fibers (28)	17.98	4.31	7.26	10.72
Manufacture of Rubber (29)	17.67	7.89	6.85	10.81
Manufacture of Plastics (30)	19.24	10.86	7.46	11.77
Manufacture of Non-metallic Mineral goods (31)	18.69	7.02	4.14	14.54
Smelting & Pressing of Ferrous Metals (32)	18.02	22.37	8.15	9.86
Smelting & Pressing of Non-ferrous Metals (33)	20.53	4.30	5.16	15.36
Manufacture of Metal Products (34)	21.41	6.22	5.15	16.26
Manufacture of General Purpose Machinery (35)	20.45	7.28	5.75	14.69
Manufacture of Special Purpose Machinery (36)	20.94	4.04	6.03	14.90
Manufacture of Transport Equipment (37)	21.35	3.46	3.96	17.39
Electrical Machinery & Equipment (39)	22.26	5.017	5.274	16.992
Computers & Other Electronic Equipment (40)	23.16	5.246	5.151	18.018
Manufacture of Measuring Instruments & Machinery for Cultural Activity & Office Work (41)	23.50	3.538	5.059	18.446
Manufacture of Artwork (42)	20.49	5.110	5.850	14.646

Notes: Unit is RMB 1,000 (equivalent to \$125 during the period 2000-2007). The wage gap is computed by the difference between estimated industry-level skilled wages ($\hat{\gamma}_{1jt}$) and unskilled wages which is the product of $\hat{\gamma}_{2jt}$ and industry-year minimum wages $w_{jt}^{\min}$ by industry and by year.

Appendix C: A Theoretical Intepretation

In this section, we develop a simple theoretical framework to explain our key empirical finding that input trade liberalization increases wage inequality. In order to investigate the effect of trade liberalization on wage inequality, instead of focusing on homogeneous labor, we extend the $(n+1)$-country model in Amiti and Davis (2012) by introducing both skilled and unskilled workers into the final-goods production.

- **Consumption (of final goods)**

A representative consumer allocates her expenditure E across a continuum of available final-goods varieties v to

$$\underset{p(v)}{\text{Min}}\ E = \int p(v)q\,(v)\,dv \quad \text{s.t.} \quad \left[\int q\,(v)^{\frac{\sigma-1}{\sigma}}\,dv\right]^{\frac{\sigma}{\sigma-1}} = U \tag{21}$$

where p denotes the price, q the quantity for variety v, and $\sigma > 1$ is the elasticity of substitution between final-goods varieties. The demand curve for the final product v is $q(v) = Q[p(v)/P]^{-\sigma}$ and the corresponding revenue is $r(v) = R[p(v)/P]^{1-\sigma}$, where $Q = U$ and P is an aggregate price index given by $P = [\int p(v)^{1-\sigma}dv]^{\frac{1}{1-\sigma}}$ with $PQ = R$.

- **Production of final goods (and intermediate inputs)**

Each country has a sector of intermediate inputs that are available in a fixed measure of varieties on a unit interval, $[0,1]$.[22] These inputs are produced under constant return-to-scales, with one unit of unskilled labor producing one unit of the intermediate input. Therefore, under free entry, the local price of the domestic intermediate inputs is also equal to the unskilled wage $\underline{w}$.

To produce final goods, each potential entrant/firm has to incur a sunk cost f_e to obtain a random draw $\lambda_v = (\phi_v, \theta_v, t_{Mv}, t_{Xv})$. The respective elements are the firm's production technology (productivity ϕ_v), the required share of skilled labor in production θ_v, and the idiosyncratic components of marginal trade costs in imports and exports (t_{Mv} and t_{Xv}). That is, for a given technology ϕ_v, we assume that production requires each firm to employ a particular share of the skilled labor (presumably, ϕ_v and θ_v are positively correlated).

After learning their characteristics, some firms exit without producing, and the remaining mass of firms M will choose labor (both skilled and unskilled) and intermediate inputs to produce final outputs destined for each market to maximize profits. Steady state requires that new entries matches firm exits (at a constant hazard death rate).

Firm technology is represented by the following Cobb-Douglas production function with a composite intermediate input M and a composite labor input L:

$$q_v = \phi_v L^{\alpha} M^{1-\alpha} - f, \tag{22}$$

where ϕ_v is the firm-specific technology/productivity parameter and f is the fixed cost of production. We assume thereafter that all fixed costs are in units of domestic intermediates.[23]

The composite labor input L is given by,

$$L = \min\{\frac{l_s}{\theta_v}, \frac{l_u}{1-\theta_v}\} \tag{23}$$

[22] The assumption of a fixed measure for domestic intermediate inputs avoids the complication of multiple equilibria. See further discussion of this issue in Venables (1996) and Amiti and Davis (2012).

[23] This assumption is similar to that in Helpman, Itskhoki and Redding (2010), in which firm fixed costs are paid in a competitive outside good.

where l_s and l_u are skilled and unskilled labor inputs, and θ_v is the share of the skilled workers employed. Therefore, $\frac{\theta_v}{1-\theta_v}$ is the firm-specific skilled-unskilled labor ratio. The above Leontief specification allows us to explain our main result in the most transparent way. An alternative specification allowing for the substitution between the skilled and unskilled labor produces the same insight but complicates the model significantly.

Unlike unskilled labor, skilled labor receives a wage, w_v, that is related to the performance of the firm for which they work. Following the fair-wage argument in Amiti and Davis (2011), skilled-labor wage, $w_v = w(\pi_v)$, is a function of a firm's profit because. As inspired by Acemoglu and Pischke (1999) and later evident by Cahuc *et al.* (2006), we assume that skilled workers have more bargaining power in production than the unskilled, which normalized to zero as a benchmark. Specifically, we have $w(0) = \underline{w}$, $0 < w'(\pi_v) < \infty$, $\underline{w} \leq w(\pi_v) \leq \overline{w}$. Therefore, the wage for the composite labor in (23) becomes,

$$\begin{aligned} W_v(\pi_v) &= \theta_v w(\pi_v) + (1-\theta_v)\underline{w} \\ &= \theta_v[w(\pi_v) - \underline{w}] + \underline{w} \\ \text{or}, W_v(\pi_v) &= \theta_v \Delta w_v + \underline{w} \end{aligned} \tag{24}$$

where $\Delta w_v = w(\pi_v) - \underline{w}$ is the wage gap between skilled and unskilled labor. Furthermore, since $W_v' = \theta_v w'$, the relationship between W_v and π_v in (24) is illustrated in Figure 1.

[Insert Appendix Figure 1 Here]

Without loss of generality, we normalize the unskilled wage to unity.[24] Thus, the local price of the domestic intermediate inputs of each country is also equal to unity, and the price index of the composite intermediate inputs becomes,

$$P_{M_v} = [1 + n\tau_{Mv}^{1-\gamma}]^{\frac{1}{1-\gamma}} \leq 1 \tag{25}$$

where $\tau_{Mv} = \tau_M t_{Mv} > 1$ is the effective price (to firm v) of the intermediate inputs from a foreign country that consists of a common iceberg component $\tau_M > 1$ and a firm-specific component $t_{Mv} \geq 1$. Parameter $\gamma > 1$ is the elasticity of substitution between any two varieties of intermediates.

Therefore, the marginal cost corresponding to (22) is

$$\begin{aligned} c_v &= \frac{kW_v^\alpha P_{M_v}^{1-\alpha}}{\phi_v} \\ &= \frac{kW_v^\alpha [1 + n\tau_{Mv}^{1-\gamma}]^{\frac{1-\alpha}{1-\gamma}}}{\phi_v}, \end{aligned} \tag{26}$$

where $k \equiv \alpha^{-\alpha}(1-\alpha)^{-(1-\alpha)}$. Because of the mark-up pricing rule, the domestic price of a final-goods variety is $p_{vd} = c_v/\rho$. Thus, revenue for firm v in the domestic market becomes

$$\begin{aligned} r_{vd} &= RP^{\sigma-1} p_{vd}^{1-\sigma} \\ &= RP^{\sigma-1}[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma}[1 + n\tau_{Mv}^{1-\gamma}]^{\frac{(1-\alpha)(1-\sigma)}{1-\gamma}} \end{aligned} \tag{27}$$

[24] For simplicity, we do not model unskilled-labor wage is a function of firm profit, though our theoretical prediction still hold if we do so.

The total revenue is

$$\begin{aligned} r_v &= (1+n\tau_{Xv}^{1-\sigma})r_{vd} \\ &= (1+n\tau_{Xv}^{1-\sigma})RP^{\sigma-1}[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma}[1+n\tau_{Mv}^{1-\gamma}]^{\frac{(1-\alpha)(1-\sigma)}{1-\gamma}} \end{aligned} \quad (28)$$

where $\tau_{Xv} = \tau_X t_{Xv} > 1$ is firm v's idiosyncratic iceberg export cost to serve a foreign market, which consists of a common component $\tau_X > 1$ and a firm-specific component $t_{Xv} \geq 1$. Notice that (28) reflects the fact that, in addition to the domestic market, exporting gives a firm access to n additional foreign markets, each of which is $\tau_{Xv}^{1-\sigma} < 1$ times the size of the former.

Therefore, the profit for a firm, which exports final goods and imports intermediates, is

$$\begin{aligned} \pi_v(W_v) &= \frac{r_v}{\sigma} - [f + n(f_X + f_M)] \\ &= (1+n\tau_{Xv}^{1-\sigma})(\frac{RP^{\sigma-1}}{\sigma})[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma}[1+n\tau_{Mv}^{1-\gamma}]^{\frac{(1-\alpha)(1-\sigma)}{1-\gamma}} - [f + n(f_X + f_M)] \end{aligned} \quad (29)$$

where f is the fixed cost of production, f_X (*resp.* f_M) is the fixed cost of exporting to (*resp.* fixed cost of importing from) a foreign country. When a firm only exports final goods, its profit becomes

$$\pi_v(W_v) = (1+n\tau_{Xv}^{1-\sigma})(\frac{RP^{\sigma-1}}{\sigma})[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma} - (f + nf_X). \quad (30)$$

When a firm only imports intermediates, it profit becomes

$$\pi_v(W_v) = (\frac{RP^{\sigma-1}}{\sigma})[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma}[1+n\tau_{Mv}^{1-\gamma}]^{\frac{(1-\alpha)(1-\sigma)}{1-\gamma}} - (f + nf_M) \quad (31)$$

When a firm only serves the domestic market, its profit is

$$\pi_v(W_v) = (\frac{RP^{\sigma-1}}{\sigma})[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma} - f \quad (32)$$

Firms whose profits are negative exist the market completely.

For given macro variables (i.e., R and P), Equ. (24), together with the corresponding one in Equ.(29)-(32), can determine a firm's profit and wages for the composite labor (and, therefore, the wage gap or the skilled wage using Equ. (24)). Among these four modes, each firm chooses the one that maximizes its profit. Thus, firm wages, profits and all other variables are determined conditional on the macro variables.

Following Amiti and Davis (2012), since most firms neither export nor import, we assume that (i) $f_X \geq f$ and (ii) $f_M > (\frac{f}{n})[(1+n\tau_M^{1-\gamma})^{\frac{(1-\alpha)(1-\sigma)}{1-\gamma}} - 1]$. The first assumption ensures that zero-profit firms do not export and the second that a firm earning zero profit when it fails to import intermediates will not find it advantageous to import intermediates.[25] Together, these assumptions imply that there is an equilibrium cut-off such that a firm survives if and only if $\phi \geq \phi^*$. Therefore, the profits of a firm conditional on the cut-off can be written as $\pi_v = \pi(\lambda_v, \widehat{\phi}^*)$, where $\widehat{\phi}^*$ is the

[25] Notice that the net gains from importing intermediates are $[(1+n\tau_{Mv}^{1-\gamma})^{\frac{(1-\alpha)(1-\sigma)}{1-\gamma}} - 1](\frac{RP^{\sigma-1}}{\sigma})[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma} - nf_M$. For a zero-profit firm, $(\frac{RP^{\sigma-1}}{\sigma})[\frac{kW_v^\alpha}{\rho\phi_v}]^{1-\sigma} = f$. Therefore (setting $t_{Mv} = 1$), the condition $[(1+n\tau_M^{1-\gamma})^{\frac{(1-\alpha)(1-\sigma)}{1-\gamma}} - 1]f - nf_M < 0$ means that the maximum gain from importing intermediates is negative.

notional cut-off productivity because zero-profit firms have wages equal to unity (see Equ. (24)):

$$\pi(\widehat{\phi}^*, W_v(0)) = (\frac{RP^{\sigma-1}}{\sigma})[\frac{k}{\rho\widehat{\phi}^*}]^{1-\sigma} - f = 0. \tag{33}$$

From Equ.(33), we can obtain the macro values consistent with $\widehat{\phi}^*$:

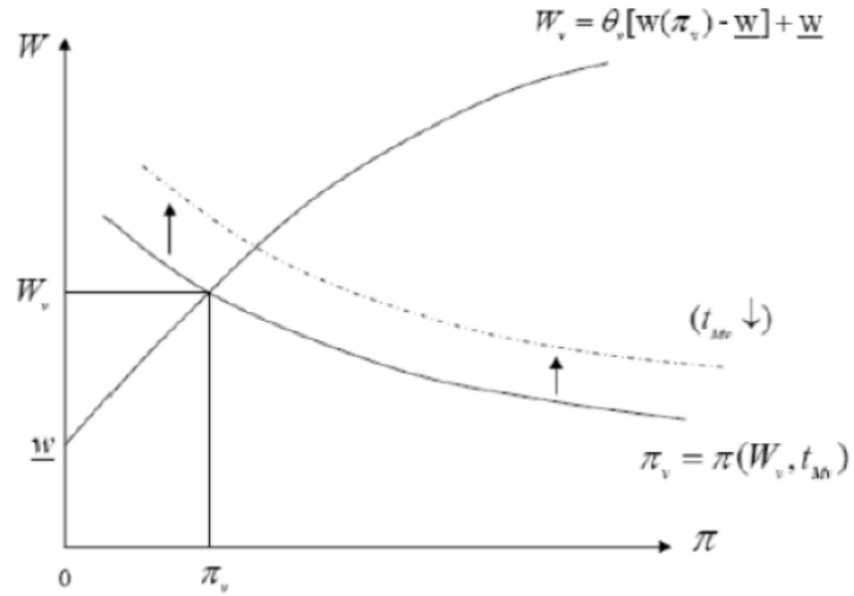
$$RP^{\sigma-1} = \sigma f(\frac{k}{\rho\widehat{\phi}^*})^{1-\sigma}. \tag{34}$$

With Equ.(34), from the previous firm's optimization problem we can obtain $\pi_v = \pi(\lambda_v, \widehat{\phi}^*)$, which is consistent with this notional cut-off and all other equilibrium variables.

Therefore, using Equ.(29) and Equ.(24), it is straightforward to obtain the following proposition.

Proposition: *A reduction of t_{Mv} increases the firm-level wage gap Δw_v between skilled and unskilled labor.*

This result can be illustrated in Appendix Figure 1. From Equ.(29) notice that $\pi'(W_v) < 0$ (i.e. higher wages reduce profits, *ceteris paribus*) and the intersection of $W_v(\pi_v)$-curve and $\pi_v(W_v)$-curve determines the equilibrium firm profit and wage (for a given mode). A reduction of t_{Mv} shifts the $\pi_v(W_v)$-curve up and, as a result, raises both π_v and W_v. Consequently, from Equ.(24), the wage gap increases.



Appendix Figure 1: Determination of Firm Average Wages and Profit