International Trade and Collective Bargaining Outcomes: Evidence from German Employer-Employee Data*

Gabriel Felbermayr[†], Andreas Hauptmann[‡], and Hans-Jörg Schmerer[§]

December 2011

Abstract

An emerging literature on the role of unions in internationally active firms mitigates the general perception that exporting firms pay higher wages. In theory, fiercer competition due to the internationalization of a firm can have negative feedback effects into a union's bargaining position. We propose an empirical test of that prediction using German linked employer-employee data, where the information about *plant-* and *industry-level* collective agreements enable us to partition plants into different bargaining regimes. To test the rent-sharing argument we exploit the individual worker information of our data and construct profitability measures that are free of the plant's skill composition. Our results indicate that the relative bargaining position of the union is weakened by trade if wages are bargained collectively at the plant-level. In line with the theoretical prediction we also show that a surge in those plants' export intensity is negatively associated with wages.

Keywords: trade, unions, collective bargaining, employer-employee data.
JEL codes: F16, J51, E24, J3

---

*We are very grateful to Bernhard Boockmann, Hartmut Egger, Peter Egger, Daniel Etzel, Oleg Itskhoki, Elke Jahn, Christian Keuschnigg, Catia Montagna, Christoph Moser, Steffen Müller, Jakob Munch, Achim Schmillen, James Tybout, Klaus Wälde, as well as participants at workshops and conferences in Hohenheim, Mannheim, and Tübingen.

[†] Ifo Institute for Economic Research, University of Munich, felbermayr@ifo.de
[‡] Institute for Employment Research (IAB), andreas.hauptmann@iab.de
[§] Institute for Employment Research (IAB), hans-joerg.schmerer@iab.de

1 Introduction

The ongoing integration of global markets sparked a political and academic debate about the causes and consequences of the observable income-inequality. While skill biased technological change, increased outsourcing opportunities, and the exporter wage premium contributed to the surge in high-skilled wages[1], earnings of the low-skilled were stagnant.[2] Beyond its positive effects on the high-skilled, globalization may also have contributed to the stagnating low-skilled earnings by magnifying the decline of the bargaining position of the unions. From a rent-sharing point of view it may well be that export participation leads to an increase in domestic wages in exporting firms due to additional revenues earned abroad (Egger and Kreickemeier, 2009; Helpman et al., 2010; Egger et al., 2011). However, increasing international activities of firms may also weaken the relative bargaining position of local unions and therefore have a negative impact on wages (Montagna and Nocco, 2011; Eckel and Egger, 2009). In this paper we address the relevance of international interdependencies in the presence of different bargaining regimes for wages using linked employer-employee data for the German manufacturing industries between 1996 and 2007. This rich data set is well suited for our purposes as it contains information on the export participation and the type of bargaining regime a plant belongs to.

In Germany, as in other countries, collective agreements still play an important role in the wage determination process. Collective agreements are conducted either at the firm-level or the *industry-level*. *Firm-level* agreements are typically better suited to account for local economic conditions, such as increasing international integration. [3] We expect plants covered by local agreements can or have to respond to changes in local conditions, whereas for *industry-level* bargaining both parties have to meet the needs for all or most of their members. Gürtzgen (2009b) supports this view by showing that wages in plants covered by *firm-level* agreements are positively associated with quasi-rents, which may be furthermore interpreted as evidence for rent-sharing. This view is also supported by Gürtzgen (2009a), who shows that wages are lower in industries characterized by a larger plant-heterogeneity if wages are bargained at the *industry-level*. Our results indicate that rent-sharing in exporting plants is lower if wages are either bargained at the *plant-* or the *industry-level*, which is in line with the model of Montagna

[1] For Germany, the evolution of wages is documented by Dustmann et al. (2009). Attanasio et al. (2004) find a similar pattern for Columbia and they are able to link the rise in wage inequality partly to a tariff reform enforced in the 80's and 90's.

[2] Exporting firms are larger, more productive, invest more intensively, and - most important in our context - pay higher wages to their employees. Based on the seminal work of Bernard et al. (1995), the so called exporter wage premium in combination with the advancing global integration may have contributed to the rising wage inequality. See also Schank et al. (2007) for a survey of different studies.

[3] The system of industrial relations in Germany is based on a dual system of representation by unions and work councils. For a brief description of the German system see Schnabel et al. (2006). Addison et al. (2010, 2011) provide an overview of the structure and developments in the German collective bargaining system.

and Nocco (2011), Montagna and Nocco (2011), and Eckel and Egger (2009). Moreover, it underlines the importance of the wage setting mechanism and labor market institutions in the context of globalization.

Consistent with the existing literature, we also show that wages are higher in plants more open to globalization. However, once controlling for observed and unobserved worker and workplace characteristics the (residual) exporter wage premium decreases significantly (see also Schank et al., 2007), indicating that the positive premium is to a large extend driven by assortative matching.[4,5] In other words, differences in wages are at least partly driven by differences in workforce characteristics. Based on linked employer-employee data from Mexico, Frias et al. (2009) however find that only one-third of the Mexican exporter wage premium can be explained by unobservable differences in the workforce composition.

We also pay special attention to the interaction between export intensity and productivity. This goes beyond most of the Melitz (2003) applications, where firms either pay the same wages due to constant mark-ups as it is standard in a CES environment, or proportional shares of their profits, and where firms sort into an exporting regime according to their productivity. The descriptives for our profitability measure do not reveal a clear sorting of plants into domestic and export regimes as proposed by Melitz (2003). Firms that export are on average more productive, but we also observe profitable non-exporters and unprofitable exporters (Powell and Wagner, 2011). Opromolla and Irarrazabal (2005) model the evolution of productivity in a dynamic Melitz (2003) framework and show that firms can endure negative profits in the short run when productivity stochastically increases over time. Chaney (2005) sketches the dynamic forces in a short run Melitz (2003) model where firms that got hit by the exogenous death rate can go on hold if their expected future profits are high enough so that they become profitable again. Thus, short-run dynamics are an important and realistic but - for the sake of simplicity - to a large extend ignored feature in most of the established heterogeneous firm models. More important, both approaches can explain why a clear sorting of firms into different regimes is not supported by the data. A firm's export intensity can thus be a spurious measure for productivity. Moreover, it is also likely that firms that start to export have to bear additional foreign beachhead costs in order to establish new foreign distribution facilities, which could lead to a decrease in profitability in the short-run.

---

[4] Differences in the workforce composition are also in line with the models by, e.g. Helpman et al. (2010), Davidson et al. (2008), or Yeaple (2005). Krishna et al. (2011) and Davidson et al. (2010) also find empirical evidence for matching effects and sorting. In a similar context Krishna et al. (2011) show for Brazil that the impact of trade openness on wages turns insignificant if sorting effects are simultaneously considered.

[5] Klein et al. (2010) provide robust evidence on the existence of a negative exporter wage premium for low skilled workers for Germany.Based on the same data Schmillen (2011) demonstrates that the exporter wage premium shows up only in plants that export to more remote markets.

Theoretical considerations. Our result indicate that rent-sharing is somewhat mitigated by more intensive trade on the plant level. This result can be rationalized by an intensified competition due to the internationalization of the firm. Firm or plant-level unions are more cautious about employment-effects of globalization when changes in the firm's environment cause potential employment cuts.

Egger and Etzel (2009) analyze the effect of international competition on the relative position of the firm in the bargaining process between firms and the collective of workers in a oligopolistic competition model with unions in the labor market. Intensified competition due to the opening up of the country to international trade negatively affects wages in their oligopolistic continuum of industries framework. Firms in industries with higher labor productivity always pay higher wages. Intensified trade however reduces the rent-extracting ability of the union, which has a negative effect on wages. The intuition behind that result is that there are three countervailing effects. As standard in oligopolistic models going from autarky to free trade increases firms' labor demand and output, which has a positive impact on the wage rate demanded by the union. However, Egger and Etzel (2009) show that this positive effect is outweighed by *i*) lower firm profits due to more competition, and *ii*) a higher labor demand elasticity. A higher labor demand elasticity implies that unions are more cautious about the negative employment effects and therefore moderate their wage claims. The authors also extend their model by showing that centralized bargaining at the industry level yields qualitatively the same results. However, in their centralized bargaining environment unions still face the wage to employment trade-off due to the assumption of efficient wage bargaining about wages and industry-wide employment. This contrasts with Braun (2011), where centralized bargaining is modeled as wage floor above the reservation wage. The finding that centralized bargaining has even stronger effects on the rent-extracting ability of the union only holds on the industry level where industries with higher exposure to trade should exhibit lower bargaining outcomes for homogeneous workers and homogeneous firms. We test this prediction by *i*) taking industry openness on the firm level into consideration and *ii*) by performing regressions on the industry level. The latter is most closely related to Egger and Etzel (2009). Industries with higher average productivity should pay higher wages but increased competition due to international trade weakens the unions wage claims in favor of labor demand.[6,7]

[6] It is well documented that unions care about the well-being of their members. Donado and Wälde (2011) for instance show that unions play an important role in setting workplace safety standards. Plant-level unions are able to gather information about the health condition of the respective firm's workforce. Improvements in safety conditions not only improve the individual worker's well being, the firms are also better off due to the reduction of temporary shortfalls in its workforce caused by illness.

[7] From an empirical perspective our study is also closely related to Blien et al. (2009). The authors propose to take the type of wage setting mechanism into account when testing the wage curve. Based on the same data as our study, they find point estimates in line with Blanchflower and Oswald (1994) for firms that bargain wages collectively on the plant level.

Montagna and Nocco (2011) analyze how competition and variable markups in a heterogeneous firm framework affect bargaining. One of the crucial points in their model is the distinction between domestic and export profit-centers within a firm. Competition from abroad can reduce the bargaining position of the firm- (plant-) level union during wage negotiations and the separation of workers into plants with different export intensities leads to different outcomes for exporting and non-exporting firms. Their model extends the Melitz and Ottaviano (2008) framework by allowing for collective bargaining between firm-wide worker coalitions and the firm's decision makers. Exporting firms supply both the domestic and the foreign market. The clear distinction between domestic and export profit centers is consistent with firms consisting of different plants that supply the domestic or the foreign markets. Plant-level negotiations about wages and employment feedback into lower wage claims by the unions when international competition negatively affect firms' labor demand. Unions in the domestic supply center bargain wages above those bargained by worker-coalitions in the export supply center where the union takes the negative employment effects due to a higher competition on the export market into account. Exporting plants' price elasticity of demand is higher than the domestic supply plants' price elasticity, which reduces their monopoly price setting power in the foreign market and thus leads to more moderate wage claims of unions located in the foreign profit center.

Eckel and Egger (2009) or Skaksen (2004) both focus on the consequences of outsourcing on collective bargaining outcomes. Both papers show that the ability to outsource parts of the production chain to foreign affiliates reduces the bargaining position of the union by improving the multinational's fallback profit in case of disagreement during wage negotiations. Strengthening of the unions raises the multinational firm's incentive to invest abroad as reaction to the higher union's wage claims. Intensified international engagements by the firm is thus a potential threat for the union, which disciplines the wage claims.

Apart from the union papers discussed above, there is also a growing literature on potential labor market effects of trade on inequality and labor demand in heterogeneous firm models. Egger and Kreickemeier (2009) were the first to relax the full employment condition in the Melitz model by incorporating a fair wage constraint. Felbermayr et al. (2011a) highlight a channel through which trade liberalization reduces equilibrium unemployment through the selection of unproductive firms in an economy. The paper is closely related to the papers by Helpman and Itskhoki (2010), and Helpman et al. (2008, 2010) which focus on wage inequality, search unemployment, and the role of labor market institutions when firms are heterogeneous with respect to productivity. Felbermayr et al. (2011b) and Dutt et al. (2009) provide empirical evidence on the trade and unemployment nexus.

The structure of the paper is organized as follows. The second section outlines the data used for our empirical analysis and identifies some estimation problems and potential solutions. The main estimation strategy is discussed in the third section, followed by the results in section 4 and some concluding remarks.

2 Data and empirical strategy

We use German linked employer-employee data (LIAB) provided by the Institute of Employment Research (IAB) to test the link between export intensity and the role of union in plant-level collective wage agreements. The LIAB is a combination of the IAB establishment panel and the employment statistics of the Federal Employment Agency (Alda et al., 2005). Beginning in 1993, the IAB establishment panel is an annual survey of plants that employ at least one employee. The panel includes a variety of detailed information on the plant's structure and size. Variables include measures on the individual plant's labor force, revenues, usage of intermediate goods, the monthly wage bill, or export intensity.[8] Most important for our research is detailed plant-level information about collective agreements, which is unique for matched employer-employee data that usually do not provide detailed information for both workers and plants. Collective agreements are still widely applied and predominantly conducted at the industry- or regional-level but also at the *firm-level*. Those agreements constitute a legally binding wage floor between the two bargaining parties. Moreover, firms normally extend this agreement also to all workers, even to the non-members. Therefore the bargaining coverage is a better indicator than union density for our purposes. Figure 1 shows that, although declining over time, in 2007 about 70% of all employees in German manufacturing are still covered by collective agreements.

The employment statistics cover all employees subject to social security contributions which represents about 80% of all employed persons in Western Germany and 86% in Eastern Germany (Bender et al., 2000). Employees with no obligation to pay social security contributions, such as civil servants, workers in marginal employment and family workers, are excluded from the sample. The firms' social security contribution reports at the end of each year and additionally at the beginning and end of each employment spell are compulsory for the employer. The employment statistics also comprise detailed information on several individual characteristics such as age, gender, nationality, tenure and gross wage. Both data sets are merged by a common establishment identifier.

To include both west and east German manufacturing plants we focus on the period 1996-2007.[9] All Euro values are deflated for the base year 2000 using *industry-level* deflators from the OECD STAN database. To be consistent with the information from the individual data we use the total number of employees subject to social security contri-

[8] For further information on the IAB establishment panel see Fischer et al. (2009) and Kölling (2000).
[9] 1996 was the first year the survey has been carried out also in Eastern Germany.

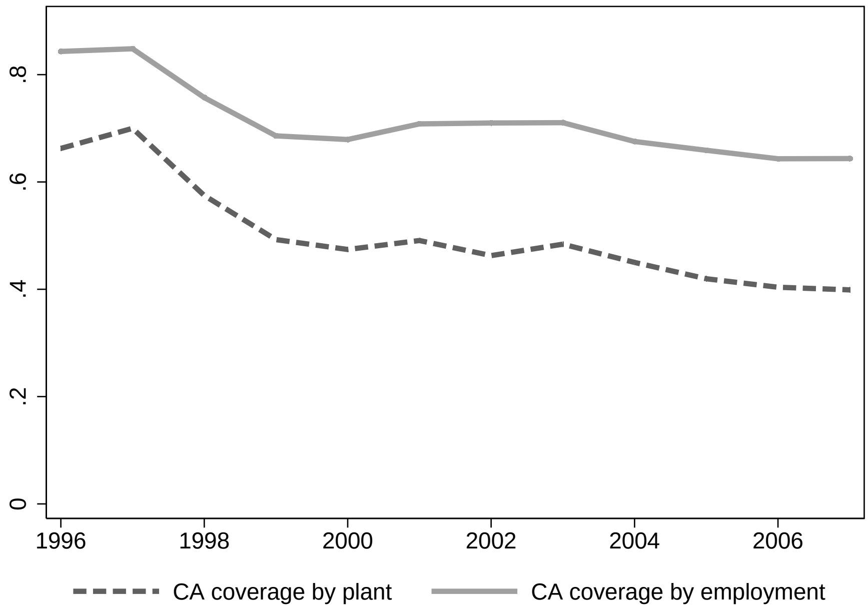

Figure 1: Collective agreement (CA) coverage,
German manufacturing, LIAB 1996-2007

butions as firm size control. Establishment output is measured by value added, i.e. total revenues minus intermediate inputs and external costs.[10] The firm's capital stock is constructed using the perpetual inventory method as proposed by Müller (2008, 2010).[11] In order to avoid outliers to bias our results, we compute the capital intensity and capital output ratio and drop all observations below the 5th and above the 95th percentile of the respective distribution. Furthermore we keep only observations with valid information on capital for two consecutive years.

Productivity as measure for rent-sharing. Our preferred proxy for rent-sharing is total factor productivity (TFP). From a theoretical point of view rent-sharing is directly linked to productivity through the positive productivity/profits relationship.[12] The total factor productivity measure is superior since it allows to account for assortative matching and possible endogeneity problems arising from unobserved productivity shocks. The latter is addressed using the approach of Levinsohn and Petrin (2003), which suggests to use intermediate inputs as proxy for those unobserved shocks.[13] The first problem is more complex. Without accounting for the work-force composition, the measured link be-

[10] We exclude establishments which do not report revenues as their business volume such as banks, financial institutions and insurance companies.

[11] Plants in the sample report investment volumes and type of investment, which allows to proxy the capital stock by summing per-period investments and taking investment specific depreciation rates into account.

[12] This standard outcome of heterogeneous firm models as Melitz (2003) can translate into a positive productivity/wage relationship. See Egger and Kreickemeier (2009) for instance.

[13] In particular we use the Stata routine `levpet` provided by Petrin et al. (2004) for the estimation of the production function.

tween profits and wages can be spurious due to assortative matching. We follow Iranzo et al. (2008) and tackle this problem by controlling for the firm's workforce composition (the average worker's ability) obtained from Mincerian wage regressions on the worker-level. Moreover, total factor productivity allows to estimate the different parameters as input-shares and elasticities simultaneously within one regression.

How to measure global interdependency? On the firm-level our data comprise information about the export intensity of the plant, measured as the share of goods produced for the export markets. Unfortunately we cannot address outsourcing directly on the firm level due to missing information about imported intermediates. Moreover, there is little to no information about the export destination available. However, we argue that the international engagement is already a threat for the unions during negotiations. Plants that are already active on international markets might find it easier to outsource parts of the production through foreign affiliates, which is already a threat for the union. Besides the plant level information about exports we also use industry-level openness measure taken from the OECD in order to tie our analysis closer to Egger and Etzel (2009).

With respect to the individual data, we focus on full-time employees only, as wages are reported as gross daily wages without any information on working hours. Therefore we exclude all observations for part-time workers, apprentices, interns and persons working at home. As the real gross daily wage will be of particular interest, we also have to deal with an additional issue concerning the wage information. Due to a reporting ceiling in social security system, wages are right-censored at the contribution limit. We impute wages by running Tobit regressions following the method proposed in Gartner (2005). For each year we run a separate regression using age, age squared, tenure, tenure squared, gender, foreign nationality as well as a full set of industry dummies as controls. The censored daily wages are replaced by predicted values obtained from the Tobit regression.

3 Empirical strategy and results

3.1 Main regression setup

To shed light on the interaction between rent-sharing and international engagement of the plant we estimate

$$\begin{aligned} \ln w_{ijt} &= \gamma \times \ln \varphi_{jt} + \xi \times EXP_{jt} + \kappa \ln \varphi_{jt} \times EXP_{jt} \\ &\quad + \alpha_1' \times Z_{it} + \alpha_2' \times Z_{jt} + \nu_t + \nu_i \times \nu_j + v_{ijt} \end{aligned} \tag{1}$$

as the preferred regression model. The dependent variable is the imputed log wage observed for individuum i employed in plant j at time t. As variables of interest we include the plant's export share to proxy exposure to international competition and TFP to proxy its profitability. Besides the identification of the exporter wage-premium and the magnitude of rent-sharing between plants and workers, our focus is also on the interaction between both. Controls for individual and plant characteristics purge the data from observable worker and plant heterogeneity. On the individual level we control for the worker's tenure measuring her time of employment within the plant and her observable level of skill. Unobservable differences in skill or ability are controlled for by including fixed-effects. On the plant-level we include a wide array of controls gathered in the vector Z_{jt}. Controls include for instance the plant's capital intensity, employment as size-control, the share of female and part-time workers employed, a dummy that takes the value one if the plant has a work-council, and dummies that indicate whether the plant bargains collectively on the firm/plant level and a dummy that indicates the use of centralized *industry-level* collective agreements. In a first step we compare OLS, person-, and spell-fixed effects regressions based on the whole set of observations. Coefficients in the spell-fixed effects regressions are identified using the within-variation in a certain plant-worker combination. A spell ends either because of a successful switch of a worker from one to another plant or due to a layout. Spell-fixed effects are preferred over person fixed effects as long as the decomposition of the time invariant effect into its worker- and plant-specific component is not a separate object of interest and it has the advantage that the identification is independent of the number of movers.[14] Standard errors are clustered at the plant level. For the main part of the analysis we also report random-effects regression results. We argue random-effects have the advantage that the identification relies on both the within- and the between variation of the data, which is important for our analysis since the export intensity relatively little variation over time.

3.2 Productivity measures

As argued in the introduction we are mainly interested in rent-sharing between firms and workers and to what extend the rent-sharing intensity hinges on the export behaviour of the plant. For that purpose we need a profitability measure on the plant-level which is not plagued by the firm's workforce composition. Assortative matching implies that more productive firms have workers with a higher ability and that has to be taken into account when analyzing the degree of rent-sharing between plants and workers. We construct the firm's profitability measure according to a method proposed by

[14] In regression (1) we were primarily interested in the worker component of the spell-fixed effect in order to purge the productivity measures from the work-force composition. Thus, we had to include both person and plant dummies in our Abowd et al. (1999) wage regression.

Iranzo et al. (2008) who suggest to use the decomposed unobserved heterogeneity from Mincerian wage regressions as additional control for the firm's workforce composition when estimating total factor productivity. Therefore we first discuss how the human capital measures are computed, followed by a discussion of the total factor productivity estimation in a subsequent step.

3.2.1 Production function estimations

The consistent estimates of the worker productivity measure h allows us to estimate a skill-free firm productivity measure according to Iranzo et al. (2008) by estimating the production function

$$Y_{jt} = A_{jt} \cdot K_{jt}^{\alpha} \cdot \tilde{L}_{jt}^{\beta} \quad , \tag{2}$$

where capital and a weighted labor-aggregate is used as inputs for the production. The labor-aggregate weights workers by its average productivity as

$$\tilde{L}_{jt} = L_{jt} \cdot E\left(h_1, ..., h_{L_{jt}}\right) \tag{3}$$

$$E = \left(1/L_{jt} \cdot \textstyle\sum_{i=1}^{L_{jt}} h_i^{\rho}\right)^{1/\rho} \quad . \tag{4}$$

Iranzo et al. (2008) use a second-order Taylor series expansion around the firm's mean ability in order to derive a testable production function in form of

$$\ln Y_{jt} \simeq \alpha \ln K_{jt} + \beta \ln L_{jt} + \beta \ln \left[\bar{h}_{jt} + \frac{1}{2}(\rho - 1)\left(\frac{\sigma_{jt}^2}{\bar{h}_{jt}}\right)\right] + \varepsilon_{jt} \tag{5}$$

We use $ln(x+y) = lnx + ln(1+y/x)$ and $ln(1+y/x) \approx y/x$ in order to derive a log-linear form of the production function that can be estimated

$$\ln Y_{jt} \simeq \alpha \ln K_{jt} + \beta \ln \left(L_{jt}\bar{h}_{jt}\right) + \delta \left(\frac{\sigma_{jt}}{\bar{h}_{jt}}\right)^2 + \varepsilon_{jt} \quad , \tag{6}$$

where $\delta = \beta\frac{1}{2}(\rho - 1)$. The average ability of the workforce, $\bar{h}_{jt}$, and the firm's standard deviation in its workers ability, σ_{jt}, are constructed using the consistently estimated worker productivity measures from equation (7).

The advantage of the second-order Taylor approximation is that it allows us to estimate the elasticity of substitution between different workers denoted by ρ. Iranzo et al. (2008) allow for substitutability between the workers within firms and estimate it instead of simply weighting the workers by its average ability when aggregating up the firm's input of workers $\tilde{L}$. Olley and Pakes (1996) or Levinsohn and Petrin (2003) stress the importance of controlling for unobservable short-run productivity shocks when estimating total factor productivity. Olley and Pakes (1996) use firms' investment as a proxy, whereas Levinsohn and Petrin (2003) use information about the firms' input of intermediate goods to weed out the simultaneity bias caused by omitting the unobserved productivity shocks. The authors are able to show that the main advantage of using intermediate inputs as proxy is that it allows to tackle another bias caused by zero investment flows reported by the firms simply because firms more likely report the use of intermediate inputs but not necessarily invest in their capital stock every period. We use the Levinsohn and Petrin (2003) method and estimate equation (5) in order to obtain an ability-free estimate for firms' total factor productivity.

3.2.2 Measuring human capital

Following Abowd et al. (1999) in general, and Andrews et al. (2008) as a particular application for the German data, we estimate unbiased worker-productivity measures by including firm fixed effects in the Mincerian wage regression. Abowd et al. (1999) suggest that the superior identification strategy is "person first and firms second". We thus estimate

$$\ln w_{it} = \bar{w} + \beta(x_{it} - \bar{x}) + \gamma(y_{j(i)t} - \bar{y}) + \theta_i + \phi_{j(i)t} + \epsilon_{it} \quad , \tag{7}$$

where w_{it} is the imputed daily compensation of individual worker i in time t and $\bar{w}$ is the grand mean of the imputed wage rate averaged over time. To reduce the omitted variable bias we also include person and firm characteristics gathered in the vectors x_{it} and $y_{j(i)t}$, where the latter is a weighted average control for firm j that employs worker i in time t. The larger the number of workers it employs, the higher the weight of the firm j.

The model we employ for constructing the human-capital index is different from (1) for two reasons. First of all we have to decompose the spell-fixed effect into its firm- and its worker component. Moreover, we also use a different set of control variables in order to maximize the number of movers in the sample. The identification of the firm fixed-effect hinges on the number of movers between firms. The sample size decreases rapidly in the number of firm-controls. The higher the total number of plants in the sample, the more likely it gets that plants are connected through workers switching jobs between two plants that are both observed in the sample. In order to reduce the number of plants that drop out of the sample we follow Abowd et al. (1999) by treating small firms as one group.

The firm dummy absorbs some of the unobserved heterogeneity on the firm level. Not controlling for the firm fixed effects would yield a biased estimator of the person fixed effects including both person and firm time-invariant components.[15] As Abowd et al. (1999) demonstrate, neglecting the firm fixed effect would yield estimates for $\phi_{j(i)t}$ which would also include the "employment-duration weighted average firm effect ϕ_j", provided that the other assumptions are not violated. Andrews et al. (2008) use their estimation strategy and analyze the importance of a sufficient number of movers between firms to increase the quality of the estimated firm fixed effect.[16]

Results for the human capital estimates. FELSDV regression results are reported in Table A1. To construct the human capital index as

$$\hat{h}_{it} = \hat{\eta} x_{it} + \hat{\theta}_i \tag{8}$$

we compute the worker fixed-effects from regression (7) as $\hat{\theta}_i$. The human capital index thus comprises observable and unobservable components. The first is measured as level of skill attained by the individual, and the latter captures the estimated worker-ability. The predicted $\hat{h}_{it}$ allows us to construct the first and second moments of the human-capital distribution within the plant, which facilitates the estimation of regression (5).

3.3 Results for the production function estimates

Table (1) reports the results of estimating regression (5) using the semiparametric method of Levinsohn and Petrin (2003) denoted by LP. Only the regressions in the lower panel do control for the workforce composition. Regression (1) in both panels is the benchmark including all firms. Regression (2) and (3) estimate the production function separately for non-exporters and for exporters. P-values obtained from the test on constant returns to scale are reported in squared brackets. The test does not reject the null that the sum of labor and capital coefficients sums up to unity. Total factor productivity is constructed as the predicted residuals of regression (1) including the workforce composition controls. All regressions yield comparable coefficients for capital around 0.2 - 0.4, and for labor ranging from 0.7 - 0.75.

[15] Especially for our application we have to disentangle the worker from the firm effects in order to test for assortative matching between firms and workers.

[16] Their focus lies on identifying the firm fixed effects in Abowd et al. (1999), which allows them to maximize the number of movers by using the full-sample of workers. Our sample is smaller and relies on information about the firm. We thus need matched employer-employee data, which also reduces the number of movers inside the firm. We therefore also propose a different identification strategy which relies more on the *firm-level* information when we estimate the firm-component.

Table 1: Production function estimates

Dependent variable: Value added (ln)			
		Non-exporter	Exporter
	(1)	(2)	(3)
	LP	LP	LP
Panel A: Without controlling for the workforce composition			
Employment (ln)	0.698***	0.688***	0.728***
	(0.016)	(0.022)	(0.021)
Capital (ln)	0.200***	0.155*	0.200*
	(0.056)	(0.088)	(0.109)
CRS-Test (p-value)	[0.065]	[0.093]	[0.515]
Panel B: Controlling for the workforce composition			
Employment$\times\bar{h}_{jt}$ (ln)	0.733***	0.727***	0.755***
	(0.017)	(0.022)	(0.017)
Capital (ln)	0.189***	0.153*	0.357***
	(0.061)	(0.091)	(0.094)
VC$(h_{jt})^2$	2.866***	3.237***	1.453
	(0.948)	(0.989)	(1.674)
CRS-Test (p-value)	[0.221]	[0.214]	[0.234]
Observations	20581	9273	11308

Standard errors in parentheses, * significant at 10%, ** significant at 5%, *** significant at 1%. All estimations include industry and time fixed effects. Estimation method: LP refers to Levinsohn and Petrin (2003). Standard errors are bootstrapped in columns (1)-(3). The second panel controls for the plant-level workforce composition by including the mean and the squared variance coefficient of the human capital index. Probability of the sum of parameter estimates on labor and capital to be equal to one in brackets.

3.4 Data descriptive statistics.

Profitability measures. We argue that not controlling for the firm's workforce composition yields upward biased results when regressing firm profitability on wages. Table 2 compares the standard Levinsohn and Petrin (2003) productivity measure and the skill-free Iranzo et al. (2008) productivity measure for the years 1996, 2002, and 2007. As expected the gap between exporting and non-exporting firms is smaller when controlling for the work force composition. However, the gap between non-exporter and exporter productivity increases over time and across different percentiles of the productivity distribution. This productivity gap between exporters and non-exporters decreases when controlling for the work force composition in the lower Panel B, where the gap declines by 3 to 6 percent on average. Kernel density plots on the productivity distribution, reported in Table (2), reveal the well-known stylized fact that exporting firms are more productive. Following Del Gatto et al. (2008) we also test whether TFP is pareto-distributed. However, the estimated shape-parameter is at a rather low $k = 1.14$ and the R-squared is lower than the proposed threshold reported in Del Gatto et al. (2008). See Table (A3) for more details.

4 Regression results

The Exporter Wage Premium revisited. Results obtained from regression (1) are reported in Table (3). Worker and firm controls other than the variables of interest were omitted in the regression tables for the sake of clarity.[17]

The benchmark specification includes controls for worker characteristics as tenure, age, a white collar dummy, and the level of skill attained by the respective employee. The low-skill dummy is the reference group and thus omitted in all regressions. The coefficients for medium and high skill dummies are all positive and have the expected ranking. Higher level of education is associated with a higher average wage rate. Our standard firm controls are log-employment to capture the firm's size, capital intensity measuring the relative capital to labor ratio on the plant-level, shares on the relative amount of females and part timers employed by the respective plant. The variables denoted by CA are dummy variables that indicate whether a plant bargains collectively on the plant level (Collective agreements on the plant level), and/or whether the plant sticks to industry-wide collective agreements. Council is a dummy that takes the value one if the plant has a worker-council. We compare standard OLS reported in the first column, and spell-fixed effects reported in the second column. The latter purges the data from both firm and person fixed effects, which will be the standard in the remaining analysis.

[17] Detailed output tables are available upon request.

Table 2: Total factor productivity distribution by export status

Panel A: Levinsohn and Petrin without workforce-composition controls					
	Mean	Std. Dev.	p10	p50	p90
1996					
Non-exporter	74.6	53.2	27.3	63.0	142.3
Exporter	104.0	93.1	44.0	85.7	170.5
2000					
Non-exporter	82.8	86.7	19.9	66.9	140.9
Exporter	103.4	89.4	31.8	86.2	176.0
2007					
Non-exporter	75.4	63.6	28.4	58.0	139.3
Exporter	102.6	92.3	42.1	81.5	163.8

Panel B: Levinsohn and Petrin including workforce-composition controls					
	Mean	Std. Dev.	p10	p50	p90
1996					
Non-exporter	78.3	53.2	31.4	65.9	131.9
Exporter	101.5	69.0	48.3	84.3	171.7
2000					
Non-exporter	83.3	77.3	21.5	67.7	145.4
Exporter	98.9	69.9	36.9	85.9	159.9
2007					
Non-exporter	78.5	60.7	34.3	63.0	139.8
Exporter	102.3	90.0	44.2	81.4	166.8

TFP is constructed following Levinsohn and Petrin (2003). The means, standard deviations, 10th, 50th, and 90th percentile of TFP are separately reported for non-exporters and exporters in the years 1996, 2002, and 2007. All values are expressed as percentage of the yearly-industry average, weighted by inverse drawing probability weights.

Table 3: The export wage-premium and the role of TFP (I)

Dependent variable: Logarithm of individual daily wage						
	(1) OLS	(2) FE-Spell	(3) OLS	(4) FE-Spell	(5) OLS	(6) FE-Spell
Exports (share)	0.043***	−0.016			0.049***	0.001
	(0.014)	(0.018)			(0.014)	(0.016)
TFP (ln)			0.025**	0.011***	0.026***	0.011***
			(0.010)	(0.003)	(0.009)	(0.004)
R^2	0.618	0.177	0.620	0.180	0.621	0.180
Plants	5040	5040	5040	5040	5040	5040
Observations	4658595	4658595	4658595	4658595	4658595	4658595

Standard errors in parentheses clustered at plant-level, * significant at 10%, ** significant at 5%, *** significant at 1%. Controls included but not reported are age, age squared, tenure, tenure squared, medium-,high-skill and white-collar dummies, plant size, capital intensity, the share of females and part timers and dummies for the existence of a worker council and collective agreements at the firm- or industry-level. Additionally, all estimations include a full set of region-, sector-, and time-dummies. Total factor productivity (TFP) is constructed following Iranzo et al. (2008). We apply the Levinsohn and Petrin (2003) method to control for unobserved productivity shocks.

Regression (1) confirms the general perception that plants more exposed to trade pay higher wages. Plants with a 10 percentage points higher export intensity pay on average 43 percent higher wages. However, this result might be driven by the worker's unobserved ability, resulting in a spurious correlation between export intensity and wages. Controlling for the unobserved worker heterogeneity is rather demanding and the standard procedure is to include fixed effects. The major drawback of this solution is however that the identification of the export premium then solely relies on the within variation of the data. The between component is completely absorbed by the fixed effects. The time invariant exporter-premium might by purged by the fixed-effect.[18] Moreover, the link between profitability and export status of a firm is less obvious. The inclusion of fixed effects without taking the plant's profitability into account reverses the sign of the export share measure. Regression (2) indicates that plants that increase their export activities by 10 percentage points tend to pay 16 percent lower wages. The effect is however not statistically different from zero. Nevertheless, we have serious doubts about the reliability of that result.

Export intensity is a kind of proxy for productivity or profitability which is in fact less variable then productivity itself. As in Opromolla and Irarrazabal (2005) it is likely that a change in a firm's exports is followed by a sluggish adjustment in productivity and profits towards its new steady state.[19] If the export wage premium is driven by rent-sharing as in Egger and Kreickemeier (2009) then we would expect that the adjustment

[18] Fixed effects regression can help to identify a causal effect by investigating how changes in the export behavior feed back into wage changes. We would expect that an increase in a firm's export intensity is associated with a higher profitability which in turn increases wages due to rent sharing.

[19] In their model the evolution of productivity is model by a Brownian motion with drift.

of wages is determined by the adjustment in the plant's profitability measure, which is in fact more variant over time than the export intensity. For the same sample we obtain a positive and highly significant coefficient for the profitability measure TFP in regression (3) - (4), which confirms our perception that the time invariant export intensity is not the appropriate measure to identify the export premium based on within variation of the data. The coefficient in (4) translates into 0.25 percent wage increase for a worker that switches to a 10 percent more productive firm. Including fixed effects reduces the magnitude of the effect to a 0.11 percent increase in the wage rate.

Most interestingly, the negative coefficient of the export share vanishes once we include the productivity measure TFP in the regressions. The coefficient of the export share is positive, but the magnitude is small and the effect is not significant. The coefficients of TFP do not change by much. This is a first hint that controlling for productivity is important for the identification of the exporter wage premium.

Based on that outcome we investigate the link between the export-status of the firm and its profitability by including the interaction between both.[20] We are able to show that there is some interaction between export-intensity and rent sharing. This interaction effect has to be taken into consideration in order to avoid the counterfactual result of a negative export premium. Powell and Wagner (2011) already showed that the exporter productivity-premium is largest at the lowest quantile. Employing quantile regressions they are able to show that the gap between exporting and non-exporting firms' productivity is largest for lower quantiles of the firms' productivity distribution.

Our results suggest that the export wage-premium is in fact dependent on the productivity of the plant. Rent-sharing between firms and workers gets smaller in plants more exposed to trade. Regressions (1) to (4) in Table 4 include both export share and the profitability measure TFP, plus the interaction between both. We obtain positive coefficients for both the export share and the profitability measure in all regressions. Both the coefficient for TFP and the coefficient for the export share variable are larger when including the interaction. To compute the marginal effects for both variables of interest one has to take the interaction into account. The negative interaction translates into a lower marginal effect for productivity for firms more exposed to trade, which can be interpreted as lower rent-sharing between firms and workers. Comparing two firms with the same productivity we find that the exporting firm pays a relatively lower wage rate. The magnitude of the effect becomes lower when we include also person or spell dummies. However, strikingly the results are significant but only for OLS and person fixed-effects regressions. For the spell fixed-effect regressions we find that the export-share measure is insignificant and that the interaction is significant only at the 10 percent level. Our OLS results indicate that plants which are 10 percent more productive pay

[20] Both measures are positively correlated. However, the corelation is at a rather low 0.11 so that colinearity is not a severe problem in our regressions.

0.7 percent higher wages.[21] Secondly, plants with 10 percentage points higher export intensity pay on average 8 percent higher wages.[22] Evaluated at the mean export share of 0.41 the interaction translates into a marginal effect for TFP equal to 0.03. Thus, the magnitude of rent sharing between firms and workers reduces from 0.6 (non-exporters) to 0.3 percent (exporters).

Table 4: The export wage-premium and the role of TFP (II)

Dependent variable: Logarithm of individual daily wage				
	(1) OLS	(2) FE-Spell	(3) OLS	(4) FE-Spell
TFP (ln)	0.071***	0.029***	0.108***	0.053**
	(0.007)	(0.006)	(0.011)	(0.021)
Exports (share)	0.785***	0.243***		
	(0.111)	(0.074)		
Exports × TFP	−0.089***	−0.029***		
	(0.013)	(0.009)		
Openness			0.056***	0.033
			(0.018)	(0.021)
Openness × TFP			−0.005***	−0.002**
			(0.001)	(0.001)
R^2	0.623	0.181	0.622	0.188
Plants	5040	5040	5003	5003
Observations	4658595	4658595	4654547	4654547

Standard errors in parentheses clustered at the plant-level in (1)-(2) and at the industry-level in (3)-(4), * significant at 10%, ** significant at 5%, *** significant at 1%. Controls included but not reported are age, age squared, tenure, tenure squared, medium-,high-skill and white-collar dummies, plant size, capital intensity, the share of females and part timers and dummies for the existence of a worker council and collective agreements at the firm- or industry-level. Additionally, all estimations include a full set of region-, sector-, and time-dummies. Total factor productivity (TFP) is constructed following Iranzo et al. (2008). We apply the Levinsohn and Petrin (2003) method to control for unobserved productivity shocks.

Overall the exporter wage premium is positif. However, if we compare to plants with the same export intensity but different productivity levels, the premium gets smaller the more profitable the firms is. For plants with a productivity 5 (close to the minimum) we find an marginal effect equal to 0.3. Evaluated at the mean the premium is around 0.048. As a last check we will also consider regressions with *industry-level* openness measures in order to tie our empirics closer to Egger and Etzel (2009). The results confirm the regressions based on the export intensity. Regression (3) to (4) indicate that wages in more open economies tend to be higher overall. The rent sharing between firms and workers is also positive. On the firm level we also find that the magnitude of rent sharing tends to be much more pronounced in industries which are less open.

[21] For zero export intensity.
[22] For zero productivity.

The role of collective bargaining. One of the explanations why exporting firms may pay relatively lower wages than non-exporting firms is the presence of unions that might be threatened by international competition and the wage-to-employment trade off. To test that relationship we exploit the information about the type of collective agreements. On the *firm-level* we expect that export intensity has some feedback effects into the bargaining outcome of exporting and non-exporting firms. The union sets an industry-wide wage by facing the tradeoff between industry labor demand and wages by taking the plant-level export share into consideration. According to Egger and Etzel (2009), openness on the *industry-level* should have similar effects on plants in both collective bargaining regimes.

Table (5) reports coefficients obtained from regressions either including observations for plants without collective bargaining in column (1) to (3), or plants that either set wages according to (*plant-* or *centralized-bargaining agreements*) in (4) to (6). The upper panel employs the information in the plant-level export share, whereas the lower panel exploits *industry-level* data as globalization proxy. We compare pooled OLS, spell fixed- and spell random-effects estimators. Both regimes are comparable due to the same number of plants included in both regressions.[23] In line with the rent-sharing argument we find a positive correlation between a plant's productivity and wages payed to their employees.

However, the exporter wage premium and the interaction is significant only for firms that bargain collectively. The positive productivity premium in the collective agreement regime can be explained by an efficiency wage approach. Firms can always depart from the union wage by paying wages above the industry-level agreements. Supporting Egger and Etzel (2009), Eckel and Egger (2009), and Montagna and Nocco (2011), we find that the negative interaction on the plant-level only holds only in the collective agreement regime. In line with Egger and Etzel (2009) we also find similar results employing industry-level openness measures, but again only for the collective bargaining regime. Our data allows the distinction between *plant-* and *industry-level* agreements so that we can go one step further by disentangling the collective bargaining regime into a plant- and an industry-level regime in a subsequent step.

Table 6 reports the results for the separate *firm-level* regressions. We again employ different regression models as OLS, spell fixed- and random effects and we also try different productivity measures as robustness checks. Regressions reported in the first panel are include the export-share as openness measure, whereas industry-level openness was used in the lower panel. Regression (1) - (3) in each panel focus on plants that indicate the use of plant-level collective agreements, whereas regressions (4) to (6) in each panel are based on the subsample of centralized collective bargaining plants.

[23] Though we have different number of observations the results are comparable since we cluster standard errors on the plant level.

Table 5: The role of collective agreements

Dependent variable: Logarithm of individual daily wage

	No collective agreement coverage			Collective agreement coverage		
	OLS	FE-Spell	RE-Spell	OLS	FE-Spell	RE-Spell
TFP (ln)	0.083***	0.031***	0.045***	0.066***	0.028***	0.041***
	(0.010)	(0.010)	(0.010)	(0.008)	(0.008)	(0.007)
Exports (share)	0.287	−0.100	0.018	0.726***	0.244***	0.423***
	(0.207)	(0.183)	(0.164)	(0.124)	(0.088)	(0.079)
Exports × TFP	−0.037	0.008	−0.004	−0.081***	−0.029***	−0.049***
	(0.026)	(0.023)	(0.020)	(0.015)	(0.011)	(0.009)
R^2	0.590	0.126		0.597	0.192	
Plants	2626	2626	2626	3302	3302	3302
Observations	491828	491828	491828	4166767	4166767	4166767

	No collective agreement coverage			Collective agreement coverage		
	OLS	FE-Spell	RE-Spell	OLS	FE-Spell	RE-Spell
TFP (ln)	0.101***	0.058	0.078**	0.104***	0.050**	0.073***
	(0.027)	(0.044)	(0.039)	(0.013)	(0.020)	(0.014)
Openness	0.053	0.048	0.055	0.052**	0.030	0.039**
	(0.037)	(0.042)	(0.040)	(0.018)	(0.020)	(0.018)
Openness × TFP	−0.003	−0.002	−0.003	−0.005***	−0.002**	−0.004***
	(0.002)	(0.003)	(0.003)	(0.001)	(0.001)	(0.001)
R^2	0.592	0.152		0.596	0.196	
Plants	2594	2594	2594	3284	3284	3284
Observations	489410	489410	489410	4165137	4165137	4165137

Standard errors in parentheses clustered at the plant-level in the upper panel and the industry-level in the lower panel, * significant at 10%, ** significant at 5%, *** significant at 1%. Controls included but not reported are age, age squared, tenure, tenure squared, medium-,high-skill and white-collar dummies, plant size, capital intensity, the share of females and part timers and a dummy for the existence of a worker council. Additionally, all estimations include a full set of region-, sector-, and time-dummies. Total factor productivity (TFP) is constructed following Iranzo et al. (2008). We apply the Levinsohn and Petrin (2003) method to control for unobserved productivity shocks.

Regressions indicated by (1) use an OLS estimator, (2) run fixed-effects regressions, and (3) the spell-random effects model. All regressions still reveal a positive relationship between plant profitability and wages paid to the workers. Additionally, the export-share and the interaction between export-share and the plant-level profitability measure are negative and significant for OLS and random-effects.

Table 6: The role of collective agreements

Dependent variable: Logarithm of individual daily wage

	Firm-level agreement			Industry-level agreement		
	OLS	FE-Spell	RE-Spell	OLS	FE-Spell	RE-Spell
TFP (ln)	0.068***	0.019**	0.039***	0.055***	0.032***	0.043***
	(0.013)	(0.010)	(0.009)	(0.009)	(0.010)	(0.008)
Exports (share)	0.789***	0.129	0.399***	0.347**	0.186	0.248**
	(0.157)	(0.142)	(0.113)	(0.164)	(0.135)	(0.123)
Exports × TFP	−0.089***	−0.017	−0.047***	−0.037*	−0.022	−0.029**
	(0.018)	(0.015)	(0.012)	(0.020)	(0.016)	(0.015)
R^2	0.685	0.156		0.584	0.206	
Plants	845	845	845	2804	2804	2804
Observations	654761	654761	654761	3512006	3512006	3512006

	Firm-level agreement			Industry-level agreement		
	OLS	FE-Spell	RE-Spell	OLS	FE-Spell	RE-Spell
TFP (ln)	0.109***	0.033	0.070*	0.075***	0.041*	0.050***
	(0.034)	(0.034)	(0.038)	(0.015)	(0.021)	(0.016)
Openness	0.072***	0.032	0.050*	0.032	0.024	0.023
	(0.024)	(0.032)	(0.030)	(0.019)	(0.021)	(0.020)
Openness × TFP	−0.005***	−0.001	−0.003*	−0.003**	−0.001	−0.001
	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
R^2	0.684	0.160		0.584	0.210	
Plants	838	838	838	2790	2790	2790
Observations	654524	654524	654524	3510613	3510613	3510613

Standard errors in parentheses clustered at the plant-level in the upper panel and the industry-level in the lower panel, * significant at 10%, ** significant at 5%, *** significant at 1%. Controls included but not reported are age, age squared, tenure, tenure squared, medium-,high-skill and white-collar dummies, plant size, capital intensity, the share of females and part timers and a dummy for the existence of a worker council. Additionally, all estimations include a full set of region-, sector-, and time-dummies. Total factor productivity (TFP) is constructed following Iranzo et al. (2008). We apply the Levinsohn and Petrin (2003) method to control for unobserved productivity shocks.

5 Conclusion

This paper sheds light on the implications of global competition for the wage setting mechanism in the presence of unions. Quite to the contrary of common beliefs, our results indicate a weakening of the unions bargaining position when firms go global. Our analysis is based upon numerous theoretical contributions that demonstrate through which channels outsourcing or intensified dependency on foreign markets affect collective bargaining outcomes. A benevolent union responds to fiercer competition generated through outsourcing or intensified trade relations by lowering its wage claims in order to protect their members' work places. As a result unions claim a lower share of the rents generated within the plant. Our preferred measures for rent-sharing is a profitability measure that is purged from the plant's skill-composition. In line with the theoretical predictions outlined in the introduction we are able to show that a surge in collective bargaining plants' export intensity is negatively associated with wages. The well-known exporter wage premium shows up in our regressions when the identification is based on both the within and the between variation of the data and/or if we explicitly allow for interactions between exports and productivity by taking a plant's profitability into account. Moreover, the export-share turns out significant only in plants that either bargain wages collectively or individually on the plant level. To the best of our knowledge, this paper is the first connecting different wage bargaining regimes to the exporter wage premium based on matched employer-employee data.

Appendix A. Additional Tables

Table A1: FELSDV results

Dependent variable: Logarithm of individual daily wage
Variables of interest: Firm and person fixed effects

	(1)	(2)	(3)
Age	0.076***	0.075***	0.073***
	(0.001)	(0.001)	(0.001)
$Age^2/100$	$-.084$***	$-.082$***	$-.079$***
	(0.001)	(0.001)	(0.001)
$Age^3/1000$	0.003***	0.003***	0.003***
	(0.000)	(0.000)	(0.000)
Employment (ln)		0.039***	0.034***
		(0.001)	(0.001)
Capital intensity (ln)			0.023***
			(0.001)
Observations	10107425	10107382	7611812

Rubust standard errors in parenthesis, * significant at 10%, ** significant at 5%, *** significant at 1%. Person, firm, year, and industry dummies included in all regressions. Person fixed effects of specification (2) are used to construct human capital measures consisting of observed and unobserved characteristics. These human capital measures are in turn used to construct firm-level human capital index variables such as the mean $\bar{h}_{jt}$ and the standard deviation σ_{jt}.

Exporter vs. non-exporter. Our later analysis hinges on the constructed total factor productivity measure which is our preferred proxy for firm profitability. The kernel density plot indicates that exporters in our sample are on average more productive. Moreover, the plots also reveal that productivity is normal distributed around the mean. Thus, there is no clear cutoff as predicted by Melitz (2003) and as indicated by the density plot and the test statistics presented in Table 2, firm profitability is not Pareto distributed.

Summary statistics. Table A2 reports further information about the variables used in the regressions covering unweighted and weighted means and standard deviation measures. The former are for interpretation of the regression results reported in the next section and the latter are weighted by an inverse drawing probability, which increases the representation-power of the data. The weighting matrixes have to be treated with caution. We refrain from using them in the main regressions because of the matched employer-employee setup, where the firm dimension is inflated due to the matching of the person data. We also distinguish between individual- and establishment-level, where variables are collapsed to the establishment-year dimension for the establishment-level

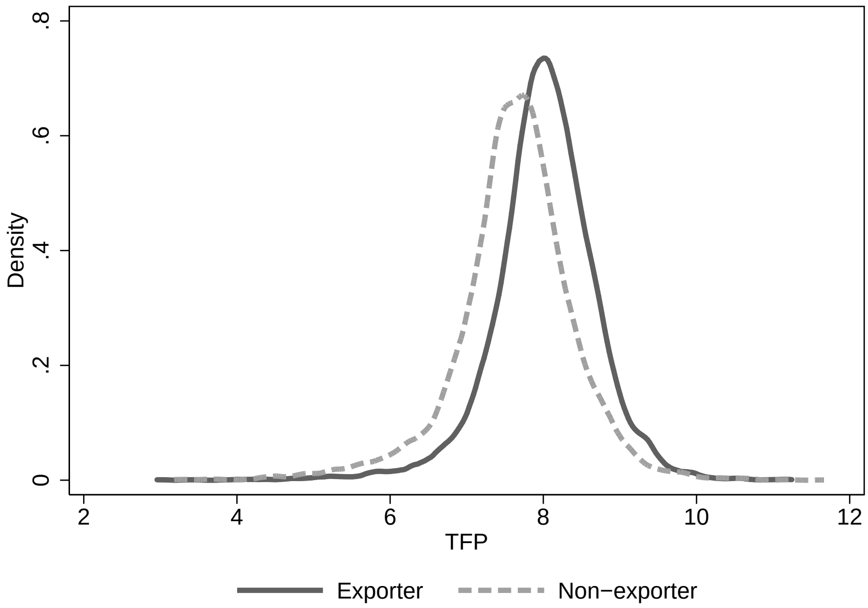

Figure 2: Kernel density plot of the profitability measure

summary reports.

Pareto test for the TFP estimates. Del Gatto et al. (2008): "Formally, consider a random variable X (e.g., our TFP) with observed cumulative distribution F(X). If the variable is distributed as a Pareto with shape parameter ks, then the OLS estimate of the slope parameter in the regression of ln(1 - F(X)) on ln (X) plus a constant is a consistent estimator of - ks and the corresponding R^2 is close to one."

Table A2: Summary statistics - unweighted

	Individual level		Plant level	
	Mean	Std. Dev.	Mean	Std. Dev.
Individual characteristics				
Daily imputed wage (ln)	4.585	0.390	4.214	0.377
Daily non-imputed wage (ln)	4.562	0.353	4.206	0.369
Female worker (dummy)	0.176	0.381	0.251	0.225
Foreign worker (dummy)	0.102	0.302	0.051	0.095
White-collar worker (dummy)	0.344	0.475	0.293	0.230
Low-skilled worker (dummy)	0.173	0.378	0.130	0.182
Medium-skilled worker (dummy)	0.701	0.458	0.789	0.202
High-skilled worker (dummy)	0.126	0.332	0.081	0.126
Age (years)	41.413	10.075	41.391	4.231
Tenure (years)	11.340	8.164	7.823	4.216
Experience (years)	16.830	8.335	13.996	4.852
Establishment characteristics				
Exporting plant (dummy)	0.890	0.313	0.549	0.498
Exports (share of total sales)	0.408	0.271	0.182	0.250
TFP (ln)	8.275	0.823	7.843	0.748
Labor productivity (ln)	11.160	0.861	10.785	0.788
Employment (ln)	7.359	1.858	4.063	1.807
Value added (ln)	18.518	2.132	14.848	2.170
Capital intensity (ln)	11.385	0.930	10.641	1.279
Female workers (share)	0.206	0.154	0.270	0.213
Part-time workers (share)	0.046	0.059	0.079	0.125
CA, industry-level (dummy)	0.762	0.426	0.465	0.499
CA, firm-level (dummy)	0.133	0.340	0.094	0.292
Existence worker council (dummy)	0.930	0.255	0.463	0.499
Industry-level characteristics				
Export orientation (dummy)	0.920	0.271	0.829	0.376
Sectoral trade openness (share)	13.448	3.802	11.812	3.706

Note: German matched employer-employee data (LIAB), 1996-2007, manufacturing industries. All monetary variables are expressed in real terms using a two-digit industry value added deflator. All industry-level variables are taken from the OECD STAN database.

Table A3: Is TFP Pareto distributed?

	k-parameter	R^2	Obs.
	Pooled sample		
Total	1.144	0.734	20580
	By year		
1996	1.204	0.741	955
1997	1.114	0.724	936
1998	1.059	0.692	1093
1999	1.130	0.714	1309
2000	1.103	0.718	2008
2001	1.128	0.724	2213
2002	1.058	0.700	2145
2003	1.079	0.700	2158
2004	1.138	0.734	2134
2005	1.119	0.740	1990
2006	1.307	0.820	1839
2007	1.309	0.808	1789
	By industry		
Textiles	1.032	0.698	664
Printing	1.036	0.695	1093
Wood	1.225	0.779	1138
Chemicals	1.134	0.766	1198
Plastic	1.083	0.596	1122
Non-metallic	1.192	0.725	1116
Metallic	1.199	0.695	1636
Recycling	1.073	0.766	178
Steel	1.273	0.678	2599
Machinery	1.206	0.695	2947
Vehicles a	1.076	0.722	1124
Vehicles b	1.066	0.733	324
Electronic	1.179	0.758	1730
Optic	1.229	0.712	1190
Furniture	1.006	0.627	570

Del Gatto et al. (2008): "Formally, consider a random variable X (e.g., our TFP) with observed cumulative distribution F(X). If the variable is distributed as a Pareto with shape parameter ks, then the OLS estimate of the slope parameter in the regression of ln(1 - F(X)) on ln (X) plus a constant is a consistent estimator of - ks and the corresponding R^2 is close to one."

References

Abowd, J. M., Kramarz, F., and Margolis, D. N. (1999). High wage workers and high wage firms. *Econometrica*, 67(2):251–334.

Addison, J. T., Bryson, A., Teixeira, P., Pahnke, A., and Bellmann, L. (2010). The state of collective bargaining and worker representation in germany: The erosion continues. IZA Discussion Papers 5030, Institute for the Study of Labor (IZA).

Addison, J. T., Teixeira, P., Bryson, A., and Pahnke, A. (2011). The structure of collective bargaining and worker representation: Change and persistence in the german model. IZA Discussion Papers 5987, Institute for the Study of Labor (IZA).

Alda, H., Bender, S., and Gartner, H. (2005). The linked employer-employee dataset created from the iab establishment panel and the process-produced data of the iab (liab). *Schmollers Jahrbuch*, 125(2):327–336.

Andrews, M. J., Gill, L., Schank, T., and Upward, R. (2008). High wage workers and low wage firms: negative assortative matching or limited mobility bias? *Journal Of The Royal Statistical Society Series A*, 171(3):673–697.

Attanasio, O., Goldberg, P. K., and Pavcnik, N. (2004). Trade reforms and wage inequality in colombia. *Journal of Development Economics*, 74(2):331–366.

Bender, S., Haas, A., and Klose, C. (2000). The iab employment subsample 1975-1995. *Schmollers Jahrbuch*, 120(4):649–662.

Bernard, A. B., Jensen, J. B., and Lawrence, R. Z. (1995). Exporters, jobs, and wages in u.s. manufacturing: 1976-1987. *Brookings Papers on Economic Activity. Microeconomics*, 1995:pp. 67–119.

Blanchflower, D. G. and Oswald, A. J. (1994). Estimating a wage curve for britain: 1973-90. *Economic Journal*, 104(426):1025–43.

Blien, U., Dauth, W., Schank, T., and Schnabel, C. (2009). The institutional context of an "empirical law": The wage curve under different regimes of collective bargaining. IZA Discussion Papers 4488, Institute for the Study of Labor (IZA).

Braun, S. (2011). Unionisation structures, productivity and firm performance: New insights from a heterogeneous firm model. *Labour Economics*, 18(1):120–129.

Chaney, T. (2005). Productivity overshooting: The dynamic impact of trade opening with heterogeneous firms. mimeo.

Davidson, C., Heyman, F., Matusz, S., Sjöholm, F., and Chun Zhu, S. (2010). Globalization and imperfect labor market sorting. Working Paper Series 856, Research Institute of Industrial Economics.

Davidson, C., Matusz, S. J., and Shevchenko, A. (2008). Globalization and firm level adjustment with imperfect labor markets. *Journal of International Economics*, 75(2):295–309.

Del Gatto, M., Ottaviano, G. I. P., and Pagnini, M. (2008). Openness to trade and industry cost dispersion: Evidence from a panel of italian firms. *Journal of Regional Science*, 48(1):97–129.

Donado, A. and Wälde, K. (2011). How trade unions increase welfare. *Economic Journal*. forthcoming.

Dustmann, C., Ludsteck, J., and Schönberg, U. (2009). Revisiting the german wage structure. *The Quarterly Journal of Economics*, 124(2):843–881.

Dutt, P., Mitra, D., and Ranjan, P. (2009). International trade and unemployment: Theory and cross-national evidence. *Journal of International Economics*, 78(1):32–44.

Eckel, C. and Egger, H. (2009). Wage bargaining and multinational firms. *Journal of International Economics*, 77(2):206–214.

Egger, H. and Etzel, D. (2009). The impact of trade on employment, welfare, and income distribution in unionized general oligopolistic equilibrium. CESifo Working Paper Series 2895, CESifo Group Munich.

Egger, H. and Kreickemeier, U. (2009). Firm heterogeneity and the labor market effects of trade liberalization. *International Economic Review*, 50(1):187–216.

Egger, P., Egger, H., and Kreickemeier, U. (2011). Trade, wages, and profits. CEPR working paper DP8727, CEPR.

Felbermayr, G., Prat, J., and Schmerer, H.-J. (2011a). Globalization and labor market outcomes: Wage bargaining, search frictions, and firm heterogeneity. *Journal of Economic Theory*, 146(1):39–73.

Felbermayr, G., Prat, J., and Schmerer, H.-J. (2011b). Trade and unemployment: What do the data say? *European Economic Review*, 55(6):741–758.

Fischer, G., Janik, F., Müller, D., and Schmucker, A. (2009). European data watch: The iab establishment panel - things users should know. *Schmollers Jahrbuch*, 129(1):133–148.

Frias, J. A., Kaplan, D. S., and Verhoogen, E. A. (2009). Exports and wage premia: Evidence from mexican employer-employee data. mimeo.

Gartner, H. (2005). The imputation of wages above the contribution limit with the german iab employment sample. FDZ Methodenreport 02/2005, Institute for Employment Research (IAB), Nuremberg.

Gürtzgen, N. (2009a). Firm heterogeneity and wages under different bargaining regimes: Does a centralised union care for low-productivity firms? *Journal of Economics and Statistics (Jahrbücher fuer Nationalökonomie und Statistik)*, 229(2-3):239–253.

Gürtzgen, N. (2009b). Rent-sharing and collective bargaining coverage: Evidence from linked employer-employee data. *Scandinavian Journal of Economics*, 111(2):323–349.

Helpman, E. and Itskhoki, O. (2010). Labour market rigidities, trade and unemployment. *Review of Economic Studies*, 77(3):1100–1137.

Helpman, E., Itskhoki, O., and Redding, S. (2008). Wages, unemployment and inequality with heterogeneous firms and workers. NBER Working Papers 14122, National Bureau of Economic Research, Inc.

Helpman, E., Itskhoki, O., and Redding, S. (2010). Inequality and unemployment in a global economy. *Econometrica*, 78(4):1239–1283.

Iranzo, S., Schivardi, F., and Tosetti, E. (2008). Skill dispersion and firm productivity: An analysis with employer-employee matched data. *Journal of Labor Economics*, 26(2):247–285.

Klein, M. W., Moser, C., and Urban, D. M. (2010). The contribution of trade to wage inequality: The role of skill, gender, and nationality. NBER Working Papers 15985, National Bureau of Economic Research, Inc.

Kölling, A. (2000). The iab-establishment panel. *Schmollers Jahrbuch*, 120(2):291–300.

Krishna, P., Poole, J. P., and Senses, M. Z. (2011). Wage effects of trade reform with endogenous worker mobility. NBER Working Papers 17256, National Bureau of Economic Research, Inc.

Levinsohn, J. and Petrin, A. (2003). Estimating production functions using inputs to control for unobservables. *Review of Economic Studies*, 70(2):317–341.

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

Melitz, M. J. and Ottaviano, G. I. P. (2008). Market size, trade, and productivity. *Review of Economic Studies*, 75(1):295–316.

Montagna, C. and Nocco, A. (2011). Unionisation, international integration and selection. Discussion Papers 257, University of Dundee, Economic Studies.

Müller, S. (2008). Capital stock approximation using firm level panel data, a modified perpetual inventory approach. *Journal of Economics and Statistics (Jahrbücher fuer Nationalökonomie und Statistik)*, 228(4):357–371.

Müller, S. (2010). Capital stock approximation with the perpetual inventory method : stata code for the iab establishment panel. Fdz methodenreport, Institute for Employment Research (IAB), Nuremberg.

Olley, G. S. and Pakes, A. (1996). The dynamics of productivity in the telecommunications equipment industry. *Econometrica*, 64(6):1263–97.

Opromolla, L. D. and Irarrazabal, A. (2005). Hysteresis in export markets. International Trade 0512003, EconWPA.

Petrin, A., Poi, B. P., and Levinsohn, J. (2004). Production function estimation in stata using inputs to control for unobservables. *Stata Journal*, 4(2):113–123.

Powell, D. and Wagner, J. (2011). The exporter productivity premium along the productivity distribution: Evidence from unconditional quantile regression with firm fixed effects. Working Papers 837, RAND Corporation Publications Department.

Schank, T., Schnabel, C., and Wagner, J. (2007). Do exporters really pay higher wages? first evidence from german linked employer-employee data. *Journal of International Economics*, 72(1):52–74.

Schmillen, A. (2011). The exporter wage premium reconsidered destinations, distances and linked employer-employee data. Working Papers 305, Osteuropa-Institut, Regensburg (Institut for East European Studies).

Schnabel, C., Zagelmeyer, S., and Kohaut, S. (2006). Collective bargaining structure and its determinants: An empirical analysis with british and german establishment data. *European Journal of Industrial Relations*, 12(2):165–188.

Skaksen (2004). International outsourcing when labour markets are unionized. *Canadian Journal of Economics*, 37(1):78–94.

Yeaple, S. R. (2005). A simple model of firm heterogeneity, international trade, and wages. *Journal of International Economics*, 65(1):1–20.