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

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

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Abstract

An emerging literature on the role of unions in international 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

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†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 capital and goods markets has fueled a lively debate about the merits and downsides of globalization. While skill biased technological changes and increased outsourcing opportunities contributed to the surge in high-skilled wages, earnings in the low-skilled sectors were stagnant, accompanied by a surge in inequality.[1] In line with that, there is also consensus in the established literature that exporters are significantly different from nonexporters. 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[2] in combination with the advancing global integration may have contributed to the rising wage inequality.

However, one aspect has been less prominent in the discussion so far, namely the role of wage bargaining institutions. 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. 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). 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 dataset is well suited for our purposes as it contains also information on the export participation and the type of bargaining regime a plant belongs to. Our results indicate that rent-sharing in exporting plants is lower if wages are bargained at the plant-level. This result is in line with the model of Montagna and Nocco (2011) and underlines the importance of the wage setting mechanism and labor market institutions in the globalization context.

The system of industrial relations in Germany is based on a dual system of representation by unions and work councils.[3] 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

[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] See also Schank et al. (2007) for a survey of different studies.

[3] 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.

parties. Moreover, firms normally extend this agreement also to all workers, even if they are not union 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. As mentioned before, 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), where it is shown that wages are lower in industries characterized by a larger plant-heterogeneity if wage are bargained over at the industry-level.

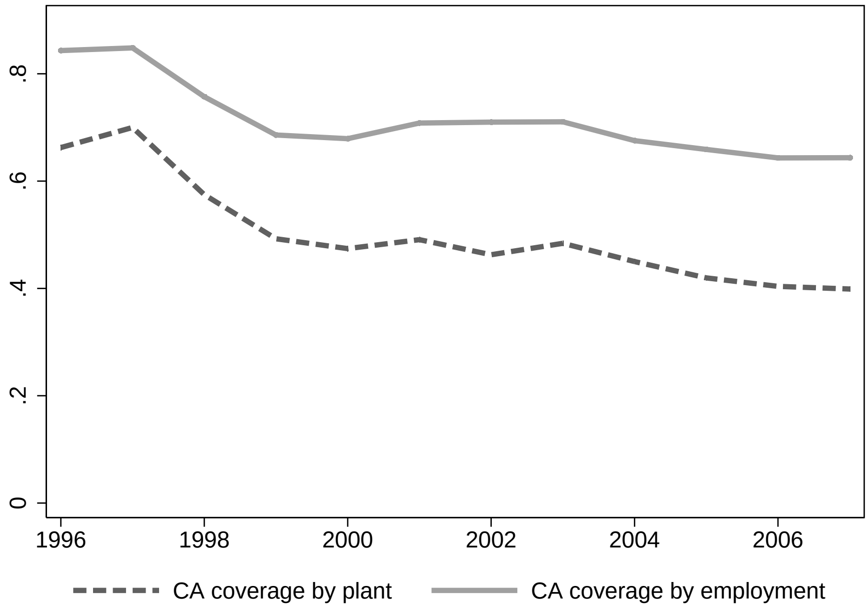

Figure 1: Collective agreement (CA) coverage,
German manufacturing, LIAB 1996-2007

Consistent with the existing literature we find that overall, the wage level is higher in plants more prone 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

[4] 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.

one-third of the Mexican exporter wage premium can be explained by unobservable differences in the workforce composition.

Moreover, we 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.

Our result indicate that rent-sharing is somewhat mitigated by more intensive trade on the plant level. This result can be rationalized by a number of recent contributions in international trade theory, where the competition that arises to the internationalization of the firm has negative implications for the bargaining-position of a union. Firm or plant-level unions are more cautious about employment-effects of globalization when changes in the firm's environment cause potential employment cuts. 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.

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 closest 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 (2010) 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

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 firms and the cleansing 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.

2 Data

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 structure and size of the plants in the sample. 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 which usually do not provide detailed information for both workers and plants. 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 comprises 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

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.

[8] For further information on the IAB establishment panel see Fischer et al. (2009) and Kölling (2000).

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 contributions 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. Our preferred measure of establishment productivity is total factor productivity (TFP). To account for possible endogeneity problems arising from unobserved productivity shocks we apply the approach of Levinsohn and Petrin (2003) and use intermediate inputs as proxy for those unobserved shocks.[12] We put much effort into tackling one remaining problem: total factor productivity measures the profitability of the firm as such, but the regressions might be spurious due to the firm's work-force composition. We thus have to purge the TFP measure from skill-compositional effects in order to avoid an upward bias towards more productive firm. 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.

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.

[9] 1996 was the first year the survey has been carried out also in Eastern Germany.
[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] In particular we use the Stata routine `levpet` provided by Petrin et al. (2004) for the estimation of the production function.

3 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 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.

3.1 Measuring human capital

Following Abowd et al. (1999) in general, and Andrews et al. (2008) as 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

$$w_{it} = \bar{w} + \beta(x_{it} - \bar{x}) + \gamma(y_{j(i)t} - \bar{y}) + \theta_i + \phi_{j(i)t} + \epsilon_{it} \quad , \tag{1}$$

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 firm dummy absorbs some of the unobserved heterogeneity on the firm level. However, the identification of the firm specific time-invariant effect hinges on the number of movers between firms. 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.[13] 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 suf-

[13] 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.

Table 1: FELSDV results

Dependent variable: Logarithm of individual daily wage *Variable of interest: Firm and person fixed effects*			
	(1)	(2)	(3)
Age	0.076***	0.075***	0.074***
	(0.001)	(0.001)	(0.001)
$Age^2/100$	−0.085***	−0.083***	−0.081***
	(0.001)	(0.001)	(0.001)
$Age^3/1000$	0.003***	0.003***	0.003***
	(0.000)	(0.000)	(0.000)
Employment (ln)		0.036***	0.014***
		(0.0005)	(0.0006)
Capital intensity (ln)			0.003***
			(0.0004)
Observations	10,108,813	10,108,813	7,571,076

Clustered 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}.

ficient number of movers between firms to increase the quality of the estimated firm fixed effect.[14] Table A1 gives an overview over the number of movers in the sample.

3.2 Production function estimations

The consistent estimates of the worker productivity measure h then 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

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

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

[14] 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.

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 testable solution of the production function in form of

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

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 worker productivity measures consistently estimated in equation (1).

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 $\tilde{L}$. The estimated δ reads as $\beta \frac{1}{2}(\rho - 1)$.

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.

Table 2: Production function estimates

Dependent variable: Value added (ln)					
	(1)	(2)	(3)	Non-exporter (4)	Exporter (5)
	OLS	FE	LP	LP	LP
Panel A: Without controlling for the workforce composition					
Employment (ln)	0.902***	0.595***	0.688***	0.694***	0.698***
	(0.012)	(0.030)	(0.015)	(0.019)	(0.021)
Capital (ln)	0.174***	0.133***	0.157***	0.116**	0.182***
	(0.009)	(0.026)	(0.028)	(0.057)	(0.052)
	[0.000]	[0.000]	[0.000]	[0.000]	[0.058]
	OLS	FE	LP	LP	LP
Panel B: Controlling for the workforce composition					
Employment×$\bar{h}_{jt}$ (ln)	0.854***	0.622***	0.692***	0.693***	0.692***
	(0.010)	(0.030)	(0.015)	(0.021)	(0.024)
Capital (ln)	0.157***	0.135***	0.167***	0.132**	0.196***
	(0.008)	(0.025)	(0.036)	(0.053)	(0.075)
VC$(h_{jt})^2$	0.252**	0.152	0.221**	0.461***	−0.291
	(0.126)	(0.132)	(0.109)	(0.140)	(0.188)
	[0.012]	[0.000]	[0.000]	[0.001]	[0.137]
Observations	21,771	21,771	21,771	9,566	12,011

Standard errors in parenthesis, * significant at 10%, ** significant at 5%, *** significant at 1%. All estimations include industry and time fixed effects. Estimation methods: OLS is ordinary least squares, FE is fixed effects and LP is Levinsohn and Petrin (2003). Standard errors are clustered at the plant level in columns (1)-(2) and bootstrapped in columns (3)-(5). 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.

3.3 Data descriptive statistics.

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.

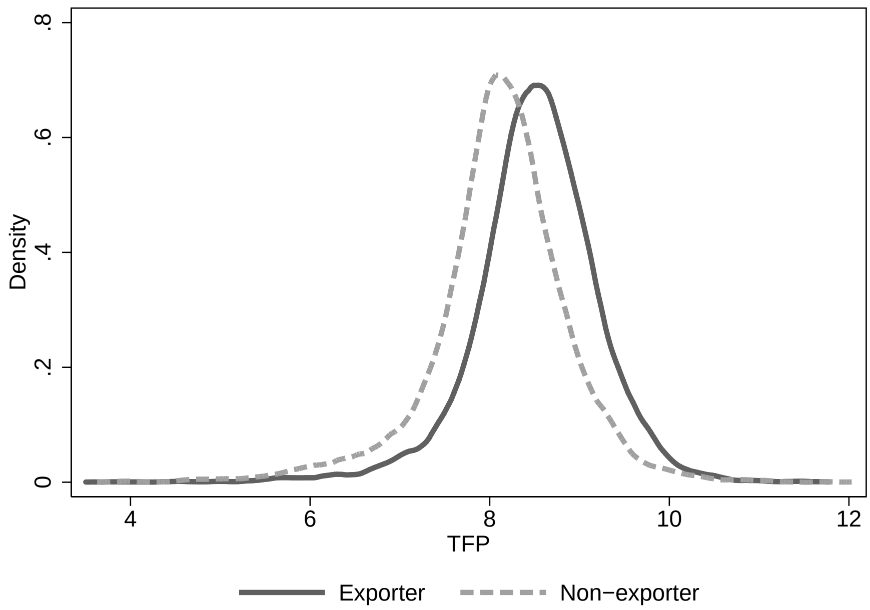

Figure 2: Kernel density plot of the profitability measure

Profitability measures. We argue that not controlling for the firm's workforce composition yields upward biased results when regressing firm profitability on wages. Assortative matching implies that more productive firms have a more skilled work force and thus higher profitability rates. We tackle this problem by applying the method proposed by Iranzo et al. (2008). Table 3 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, ranging from 7 to 14 percent difference in the 10th percentile to 30-40 percent difference in the 90th percentile in the standard Levinsohn and Petrin (2003) regressions (upper panel). This productivity gap between exporters and non-exporters decreases when controlling for the work force composition in the lower Panel B, where the gap ranges from 7-4 percent (10th percentile) to 20-7 percent (90th percentile). Thus especially for the more productive firms controlling for ability has a significant impact on the productivity estimates. As expected the gap between exporting and non-exporting firms is smaller when controlling for the work force composition.

Table 3: Total factor productivity distribution by export status

Panel A: Levinsohn and Petrin without workforce-composition controls					
	Mean	Std. Dev.	p10	p50	p90
Year 1996					
Non-exporter	64.3	51.9	23.6	53.5	127.5
Exporter	86.9	77.5	30.9	70.2	158.6
Year 2002					
Non-exporter	72.0	80.5	16.7	58.7	118.1
Exporter	97.1	81.7	28.0	78.3	172.7
Year 2007					
Non-exporter	68.5	80.8	22.7	50.9	118.7
Exporter	94.7	69.2	37.3	80.3	160.4
Panel B: Levinsohn and Petrin including workforce-composition controls					
	Mean	Std. Dev.	p10	p50	p90
Year 1996					
Non-exporter	74.1	56.4	29.1	60.6	122.1
Exporter	86.5	55.9	36.2	74.9	143.8
Year 2002					
Non-exporter	78.4	79.3	22.1	63.1	141.2
Exporter	91.9	72.0	27.3	79.4	148.9
Year 2007					
Non-exporter	79.9	84.1	32.0	61.9	140.4
Exporter	89.4	60.2	36.7	77.0	147.0

TFP is constructed following Levinsohn and Petrin (2003). The means, standard deviations, 10th, 50th, and 90th percentile of TFP 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.

Summary statistics. Table 4 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 summary reports.

Table 4: Summary statistics - unweighted

	Individual level		Establishment level	
Variable	Mean	Std. Dev.	Mean	Std. Dev.
Individual characteristics				
Daily imputed wage (ln)	4.586	0.388	4.218	0.373
Daily non-imputed wage (ln)	4.564	0.353	4.211	0.365
Female worker (dummy)	0.175	0.380	0.253	0.225
Foreign worker (dummy)	0.102	0.303	0.051	0.095
White-collar worker (dummy)	0.344	0.475	0.295	0.230
Low-skilled worker (dummy)	0.172	0.378	0.132	0.183
Medium-skilled worker (dummy)	0.703	0.457	0.787	0.201
High-skilled worker (dummy)	0.125	0.331	0.080	0.124
Age (years)	41.418	10.065	41.449	4.198
Tenure (years)	11.424	8.195	7.901	4.231
Experience (years)	16.854	8.332	14.050	4.862
Establishment characteristics				
Exporting plant (dummy)	0.896	0.306	0.557	0.497
Exports (share of total sales)	0.415	0.273	0.186	0.252
TFP (ln)	8.914	0.840	8.344	0.756
Labor productivity (ln)	11.169	0.857	10.789	0.792
Employment (ln)	7.380	1.874	4.085	1.773
Value added (ln)	18.549	2.158	14.874	2.135
Capital intensity (ln)	11.310	0.922	10.556	1.257
Female workers (share)	0.206	0.154	0.271	0.215
High qualified workers (share)	0.698	0.234	0.731	0.241
Part-time workers (share)	0.046	0.060	0.080	0.126
CA, industry-level (dummy)	0.764	0.425	0.467	0.499
CA, firm-level (dummy)	0.132	0.339	0.095	0.294
Importer of intermediates (dummy)	0.734	0.442	0.330	0.470
Industry-level characteristics				
Export orientation (dummy)	0.921	0.270	0.827	0.378
Trade openness	13.455	3.790	11.770	3.697

Summary statistics of benchmark regression sample. Source: German matched employer-employee data (LIAB), 1996-2007, manufacturing industries. Data comprise 5,630,117 observations, obtained by matching 1,625,294 individuals to 5,392 manufacturing establishments. 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 5: Summary statistics - weighted

	Individual level		Establishment level	
Variable	Mean	Std. Dev.	Mean	Std. Dev.
Individual characteristics				
Daily imputed wage (ln)	4.422	0.439	4.116	0.366
Daily non-imputed wage (ln)	4.408	0.416	4.112	0.361
Female worker (dummy)	0.216	0.412	0.278	0.265
Foreign worker (dummy)	0.090	0.286	0.058	0.126
White-collar worker (dummy)	0.320	0.466	0.278	0.270
Low-skilled worker (dummy)	0.187	0.390	0.132	0.224
Medium-skilled worker (dummy)	0.730	0.444	0.829	0.241
High-skilled worker (dummy)	0.083	0.276	0.040	0.114
Age (years)	41.121	10.468	40.061	5.825
Tenure (years)	9.433	7.755	7.271	4.590
Experience (years)	16.180	8.467	14.282	5.319
Establishment characteristics				
Exporting plant (dummy)	0.724	0.447	0.305	0.461
Exports (share of total sales)	0.276	0.277	0.077	0.174
TFP (ln)	8.595	0.777	8.143	0.718
Labor productivity (ln)	11.011	0.777	10.741	0.782
Employment (ln)	5.257	2.018	2.435	1.282
Value added (ln)	16.267	2.343	13.175	1.531
Capital intensity (ln)	10.895	1.142	10.482	1.289
Female workers (share)	0.245	0.191	0.298	0.231
High qualified workers (share)	0.677	0.243	0.710	0.251
Part-time workers (share)	0.071	0.106	0.144	0.180
CA, industry-level (dummy)	0.648	0.478	0.487	0.500
CA, firm-level (dummy)	0.096	0.295	0.055	0.228
Importer of intermediates (dummy)	0.486	0.500	0.146	0.353
Industry-level characteristics				
Export orientation (dummy)	0.848	0.359	0.753	0.431
Trade openness	12.338	3.798	11.138	3.842

Summary statistics of benchmark regression sample. Source: German matched employer-employee data (LIAB), 1996-2007, manufacturing industries. Data comprise 5,630,117 observations, obtained by matching 1,625,294 individuals to 5,392 manufacturing establishments. 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. Weighted by inverse probability weights.

4 Empirical strategy and results

4.1 Main regression setup

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

$$
\begin{aligned}
\ln w_{ijt} \quad &= \quad \gamma \times \ln \varphi_{jt} + \xi \times EXP_{jt} + \kappa \ln \varphi_{jt} \times EXP_{jt} \\
&\quad + \mathbf{ff}_1' \times Z_{it} + \alpha_2' \times Z_{jt} + \nu_t + \nu_i \times \nu_j + v_{ijt}
\end{aligned} \tag{7}
$$

as preferred regression setup. 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 international dependency 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 interested and it has the advantage that the identification is independent of the number of movers.[15] Standard errors are clustered at the plant level.

4.2 Regression results

The Exporter Wage Premium revisited. Results obtained from (7) are reported in Table (6). Worker and firm controls other than the variables of interest were omitted in the

[15] 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.

regression tables for the sake of clarity.[16]

Table 6: The export wage-premium and the role of TFP (I)

Dependent variable: Logarithm of individual daily wage

	(1) OLS	(2) FE-Person	(3) FE-Spell	(4) OLS	(5) FE-Person	(6) FE-Spell
Export share	0.0325*	−0.0344**	−0.0372**			
	(0.0172)	(0.0149)	(0.0150)			
TFP				0.0231**	0.0113***	0.0109***
				(0.0109)	(0.0035)	(0.0034)
R^2	0.5467	0.1999	0.1931	0.5484	0.2024	0.1953
Observations	5089729	5089729	5089729	5089729	5089729	5089729

Clustered standard errors in parentheses, *significant at 10%, ** significant at 5%, *** significant at 1%. Constant estimated but not reported. As openness measures we use an firm-level export dummy. Total factor productivity is constructed following Iranzo et al. (2008). We apply the Levinsohn Petrin method to control for unobserved productivity shocks. TFP obtained from regression (3) in the lower panel B of Table 2. Capital stocks are calculated using the perpetual inventory method following Mueller (2008).

The benchmark specification includes controls for worker characteristics as tenure 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, with person-fixed effects reported in the second, and spell-fixed effects reported in the third column. The latter purges the data from both firm and person fixed effects, which will be the standard in the remaining analysis.

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 32.5 percent higher wages. However, this result might be plagued by an omitted variable bias. Due to assortative matching plants more exposed to trade are more productive on average and thus have a more potential work force due to assortative matching. This may lead to a spurious correlation between export intensity and wages. Controlling for

[16] Detailed output tables are available upon request.

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. This might cause some problem given that we try to identify a premium which might be time inconsistent as well.[17] Moreover, wages are rigid and may react much slower to the change in the export behavior of a firm. On the collective bargaining level unions have to organize negotiations after realizing potential profitability increases within the firm. On the individual level, renegotiating the wage could be even more difficult. The random effects model has one advantage over the fixed effects model since it uses both the between and the within variation of the data, but the estimates are biased if the strong assumption on zero correlation between the random effect and the regressors is not met.

In our application, the inclusion of fixed effects without taking the plant's profitability into account reverses the sign of the export share measure. Plants that increase their export activities by 10 percentage points tend to pay 34.4 percent lower wages. We have serious doubts about the reliability of that result and using spell-fixed effects instead of person fixed effects does not change the picture by much.

However, 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.[18] If the export wage premium is driven by rent-sharing as in Egger and Kreickemeier (2009) then we would expect that the adjustment 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 (4) - (6), 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 0.2 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.

In a next step we investigate the link between the export-status of the firm and its performance. We are able to show that there is a non-monotonic relationship between TFP and export wage-premium which has to be taken into account 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

[17] 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.

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

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 an exporter-productivity driven by rent-sharing between firms and workers and that the premium gets smaller for more productive firms.

Table 7: The export wage-premium and the role of TFP (II)

Dependent variable: Logarithm of individual daily wage						
	(1) OLS	(2) FE-Person	(3) FE-Spell	(4) OLS	(5) FE-Person	(6) FE-Spell
TFP	0.052***	0.021***	0.020***	0.106***	0.046***	0.046***
	(0.007)	(0.006)	(0.006)	(0.017)	(0.013)	(0.013)
Export share	0.611***	0.176*	0.158			
	(0.173)	(0.097)	(0.097)			
TFP × Export	−0.056***	−0.019**	−0.017*			
	(0.017)	(0.010)	(0.010)			
Openness				0.062***	0.032***	0.032***
				(0.010)	(0.008)	(0.008)
TFP×Openness				−0.005***	−0.002***	−0.002***
				(0.001)	(0.001)	(0.001)
R^2	0.550	0.203	0.196	0.555	0.207	0.207
Observations	5089729	5089729	5089729	5345620	5345620	5345620

Clustered standard errors in parentheses, *significant at 10%, ** significant at 5%, *** significant at 1%. Constant estimated but not reported. As openness measures we use an firm-level export dummy. Total factor productivity is constructed following Iranzo et al. (2008). We apply the Levinsohn Petrin method to control for unobserved productivity shocks. TFP obtained from regression (3) in the lower panel B of Table 2. Capital stocks are calculated using the perpetual inventory method following Mueller (2008).

Regressions (1) to (6) in Table 7 include both export share and the profitability measure TFP, plus the interaction between both. Purpose of the interaction is to shed light on the non-monotonicity between export intensity and productivity. We then 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 relatively lower wage rate. The magnitude of the effect becomes lower when we include also person or spell dummies. However, stinkingly 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. In the next table we will see that the effect is driven by 50 percent of the plants in the sample, namely the plants that do not engage in centralized wage bargaining, which is a potential explanation for the bad performance in the full sample. Our OLS results indicate that plants which are 10 percent more productive pay 0.5 percent higher wages. Secondly, plants with 10 percentage points higher export intensity pay on average 6 percent higher wages. Evaluated at the mean export share of 0.41 the interaction translates into a marginal effect for TFP equal to 0.029. Thus, the magnitude of rent sharing between firms and workers reduces from 0.5 (non-exporters) to 0.29 percent (exporters). One could express the same in terms of the exporter premium. Overall the exporter wage premium is positiv. 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 close to the minimum we find an exporter wage premium around 0.331. That premium almost vanishes if we double TFP from the minimum level to 10. In that case we find an export wage premium equal to 0.051. 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 (4) to (6) 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. The marginal effect declines from 0.106 (closed economies) to 0.04 (open economy, evaluated at the mean).

The role of centralized and plant-level 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-employment trade off. To test that relationship we exploit the information about the type of collective agreements. On the firm-level we would expect that industry-wide agreements do not feedback into differences between exporting and non-exporting firms. The union sets an industry-wide wage by facing the tradeoff between industry labor demand and wages without taking the plant-level export share into consideration. We thus reduce the sample to plants that indicate that wages are bargained by unions on the industry level. Table 8 reports the results for the separate regressions. We 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 all OLS, regressions in the middle panel are all spell-fixed effects, and regressions in the lower panel are all spell random-effects models. 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. Regressions indicated by (1) always include labor productivity, (2) include Levinsohn and Petrin (2003) productivity measure, and regression (3) include skill-free

Levinsohn and Petrin (2003) productivity measures. 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 when using the subsample including plants that bargain collectively on the plant level, which supports theory in that rent-sharing is relatively lower in plants more dependent on the foreign market.

In a second step we focus on firms that indicate the use of industry-level collective agreements in regressions (4) - (6). The OLS, fixed-, and random-effects models in regression (5) and (7) support theory in that there is no evidence on a positive exporter-wage premium. Both export-share and interaction are insignificant which is surprising given the large number of observations. About 3000 firms report industry-level collective agreements, which is slightly less than 50 percent of all firms in our sample. However, we still find robust evidence on rent-sharing between workers and plants as indicated by the significant and positive coefficient for the included profitability measure. On average, plants that stick to centralized bargaining agreements on the industry level with higher profitability pay higher wages (rent sharing) but the export intensity of the plant has little net-effects. This result is not surprising. The individual plant has less power to manipulate bargaining decisions on the central level.

For the collective bargaining regime we find a positive exporter wage premium for the OLS and the random effects model. Rent sharing is significant for all estimators employed. The same holds for the interaction between export share and productivity. The results for our benchmark regressions translate into a magnitude of rents sharing which is approximately 0.086 when using labor productivity in regression (1), Table (8), panel 1. Evaluated by the average productivity we find that the degree of rent sharing decreases from 0.086 to 0.042. We can replicate the same findings based on the spell fixed effects estimator. As a robustness check we also perform spell random effects regressions, which yields results that are inbetween OLS and fixed effects regressions, as reported in panel 3 of Table (8).

For the larger fraction of plants covered by industry collective agreements we neither find evidence for the existence of an exporter wage premium, nor do find evidence for a negative impact of a plant's international engagement. This result holds irrespective the model used. However, we still find evidence for rent sharing between plants and workers as indicated by the positive and significant sign of the profitability measure *productivity*.

As a last step we want to isolate the effects of an increase in the export intensity on collective bargaining outcomes. In line with the channel highlighted above we expect that the plant-level union would accept wage-cuts as response to intensified competition abroad. This can be tested by exploiting the within-variation of the wage data

Table 8: The role of collective agreements

Dependent variable: Logarithm of individual daily wage						
	Collective bargaining plants			**Centralized bargaining plants**		
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
	Ordinary Least Squares Results			Ordinary Least Squares Results		
Export (share)	1.2808***	1.1357***	0.8797***	0.2409	0.2266	0.1878
	(0.1924)	(0.1709)	(0.1505)	(0.2281)	(0.2075)	(0.1804)
Productivity	0.0864***	0.0789***	0.0696***	0.0593***	0.0585***	0.0455***
	(0.0124)	(0.0114)	(0.0113)	(0.0097)	(0.0098)	(0.0095)
Productivity×Export	−0.1102***	−0.1049***	−0.0928***	−0.0200	−0.0204	−0.0190
	(0.0169)	(0.0160)	(0.0160)	(0.0203)	(0.0204)	(0.0204)
Constant	2.8281***	2.9318***	3.0954***	3.3383***	3.3648***	3.5344***
R-squared	0.6081	0.6075	0.6055	0.4850	0.4848	0.4818
	Spell Fixed-Effects results			Spell Fixed-Effects results		
Export (share)	0.2482	0.1800	0.1313	0.0745	0.1304	0.1257
	(0.1581)	(0.1562)	(0.1333)	(0.1816)	(0.1814)	(0.1570)
Productivity	0.0222**	0.0180*	0.0170*	0.0275***	0.0308***	0.0288***
	(0.0099)	(0.0093)	(0.0091)	(0.0090)	(0.0101)	(0.0099)
Productivity×export	−0.0291**	−0.0245*	−0.0231*	−0.0093	−0.0157	−0.0175
	(0.0145)	(0.0143)	(0.0138)	(0.0162)	(0.0178)	(0.0176)
Constant	2.7267***	2.7249***	2.7718***	3.8095***	3.7843***	3.8575***
R-squared	0.1116	0.1111	0.1109	0.1972	0.1975	0.1966
	Spell Random-Effects results			Spell Random-Effects results		
Export (share)	0.6546***	0.4620***	0.3479***	0.1710	0.2220	0.1860
	(0.1501)	(0.1287)	(0.1099)	(0.1835)	(0.1733)	(0.1505)
Productivity	0.0473***	0.0356***	0.0315***	0.0400***	0.0423***	0.0366***
	(0.0097)	(0.0077)	(0.0075)	(0.0084)	(0.0089)	(0.0087)
Productivity×Export	−0.0587***	−0.0439***	−0.0385***	−0.0168	−0.0233	−0.0228
	(0.0129)	(0.0119)	(0.0117)	(0.0162)	(0.0167)	(0.0166)
Constant	3.1945***	3.3049***	3.3772***	3.4983***	3.4858***	3.5791***
N	662450	662450	662431	3515854	3515854	3515659
No of firm-clusters	915	915	898	2996	2996	2928

Clustered standard errors in parentheses, *significant at 10%, ** at 5%, *** at 1%. Tenure, medium- and high-skill dummies, employment, capital intensity, council, female (share), part timer (share), region-, sector-, and time-dummies included but not reported in all regressions. We use the firm-level export share, profitability proxied by labor productivity in regression (1), LP-TFP in (2), and the skill-free LP-TFP in (3) using fixed effects regressions. Again we compare the outcome to our control group, the centralized bargaining plants, where we expect that changes in the export intensity have little or no effects at all. Results are reported in Table A2. All regressions include spell-fixed effects and again we find that an increase in the export share is associated with wage-cuts on the individual level in plants that engage in plant-wide collective agreements. The finding is robust and holds in all specifications. Regression (5) to (8) compare the outcome with our control group, other centralized bargaining plants. Again, the negative sign of the export share variable is mainly driven by plants that indicate plant-level collective agreements, which further supports the theoretical findings.

Interpretation Figure (3) gives an interpretation of the results by taking the interaction into account when computing the marginal-effect of TFP.

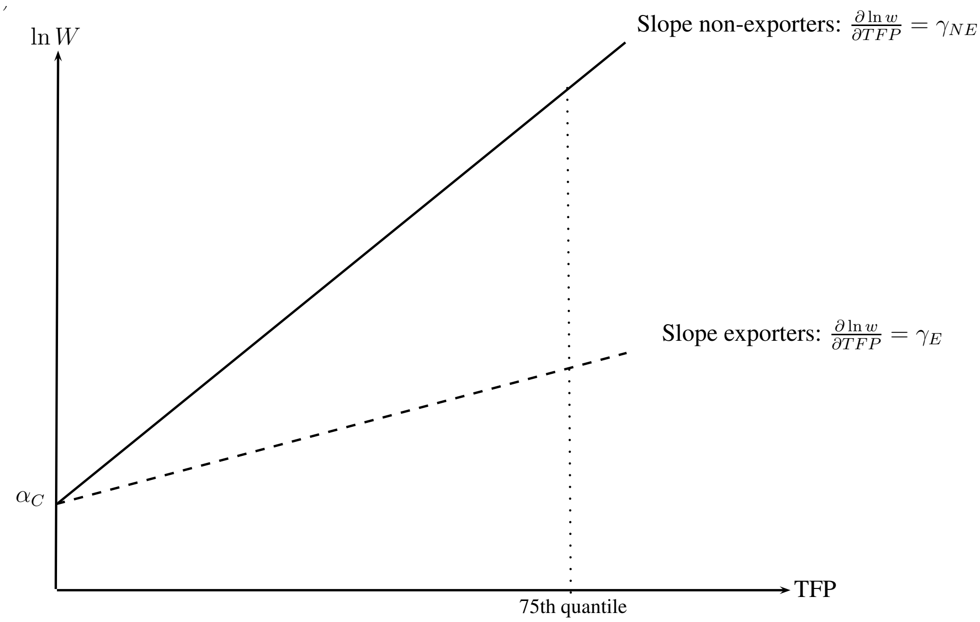

Figure 3: $\partial w / \partial TFP$ for plants with collective agreements on the plant level.

The two lines represent the magnitude of rent-sharing for exporters and non-exporters. The slope of the lines are determined by the marginal effects obtained from the Mincerian wage regression in (7). Rent-sharing of the non-exporters is equal to the coefficient of the profitability measure and thus translates into $\frac{\partial \ln w}{\partial TFP} = \gamma_{NE} = \gamma$.

Due to the negative interaction between export-intensity and plant-level profitability we obtain a lower slope $\gamma_E = \gamma - \kappa \times EXP$ for exporters with $EXP > 0$.

Comparing two firms with the same productivity we thus find that the share of revenues going to workers in plants with a higher export-intensity is lower. Put differently, comparing two firms at the 75th percentile of the profit distribution we find that those working for the plant with lower export intensity are better off.

However, this results holds only for plants that indicate the use of collective agreements on the plant level. The export-intensity of plants that set wages according to industry-wide collective agreements turns out insignificant.

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 are labor productivity and 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 if we base the identification on both the within and the between variation of the data and/or if we explicitly allow for non-monotonicity 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: Number of movers per establishment

	(1)		(2)		(3)	
Movers	Freq.	Perc.	Freq.	Perc.	Freq.	Perc.
0	4093	40.41	4080	40.34	2550	41.20
1-5	2296	22.67	2296	22.70	1384	22.36
6-10	825	8.15	825	8.16	546	8.82
11-20	753	7.43	753	7.44	499	8.06
21-30	379	3.74	379	3.75	290	4.69
31-50	477	4.71	477	4.72	278	4.49
51-100	446	4.40	446	4.41	272	4.39
$>$100	859	8.48	859	8.49	370	5.98
Total	10128	100.00	10115	100.00	6189	100.00

Number of movers per establishment in FELSDV regressions.

Table A2: Exporter wage-premia in collective bargaining plants

Dependent variable: Logarithm of individual daily wage

	Collective bargaining (plant-level)				**Central collective bargaining (industry level)**			
	(1)	**(2)**	**(3)**	**(4)**	**(1)**	**(2)**	**(3)**	**(4)**
Export (share)	−0.0758**	−0.0841***	−0.0767*	−0.1101***	−0.0088	−0.0165	−0.0292	−0.0226
	(0.0300)	(0.0258)	(0.0394)	(0.0418)	(0.0155)	(0.0159)	(0.0179)	(0.0195)
Lag export (share)		−0.0544*		−0.0631**		0.0012		−0.0175
		(0.0280)		(0.0285)		(0.0184)		(0.0255)
Labor productivity (ln)			0.0022	−0.0014			0.0234***	0.0272***
			(0.0034)	(0.0045)			(0.0060)	(0.0063)
Employment (ln)	0.0318	0.0441	0.0535*	0.0449	0.0107	0.0203	0.0077	0.0231
	(0.0283)	(0.0422)	(0.0313)	(0.0450)	(0.0277)	(0.0399)	(0.0310)	(0.0426)
Capital intensity (ln)	0.0349	0.0674*	0.0432	0.0693*	0.0015	0.0003	0.0031	0.0123
	(0.0261)	(0.0352)	(0.0319)	(0.0418)	(0.0268)	(0.0430)	(0.0275)	(0.0388)
Female (share)	−0.0350	0.1068	0.0452	0.1124	−0.0072	−0.0643	−0.0865*	−0.1022*
	(0.0514)	(0.0730)	(0.0827)	(0.0779)	(0.0494)	(0.0507)	(0.0469)	(0.0538)
Part timer (share)	−0.0885	−0.1536	−0.0703	−0.1592	−0.0802	0.1126*	0.1656***	0.1490**
	(0.1158)	(0.1930)	(0.1368)	(0.2176)	(0.1650)	(0.0621)	(0.0608)	(0.0683)
Tenure	0.0001***	0.0001***	0.0001***	0.0001***	0.0001***	0.0001***	0.0001***	0.0001***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Medium skill (dummy)	−0.0103	−0.0097*	−0.0073	−0.0098	0.0198*	−0.0025	0.0133	−0.0056
	(0.0076)	(0.0056)	(0.0078)	(0.0061)	(0.0107)	(0.0124)	(0.0094)	(0.0106)
High skill (dummy)	0.0473***	0.0194	0.0382**	0.0231*	0.1386***	0.0808***	0.1358***	0.0839***
	(0.0148)	(0.0130)	(0.0175)	(0.0139)	(0.0153)	(0.0137)	(0.0145)	(0.0115)
Worker council (dummy)	−0.0180	−0.0189*	−0.0142	−0.0169	0.0110	0.0127	0.0147	0.0180
	(0.0123)	(0.0105)	(0.0136)	(0.0118)	(0.0110)	(0.0120)	(0.0115)	(0.0119)
R-squared	0.1096	0.1045	0.1100	0.1159	0.2021	0.1824	0.1971	0.1715
N	827197	604705	662450	520423	4386982	3182692	3515854	2649678
No of firm-clusters	915	915	898	2996	2996	2928	2996	2928

Clustered standard errors in parentheses, *significant at 10%, ** at 5%, *** at 1%. We use the firm-level export share, profitability proxied by labor productivity.

Table A3: Further interactions

Dependent variable: Logarithm of individual daily wage, aggregated to the industry level

	(1) OLS	(2) FE	(3) RE	(4) FE	(5) RE	(6) OLS	(7) FE	(8) RE	(9) FE	(10) RE
Export share	10.175***	11.491***	11.614***	8.590**	8.590***	8.018***	10.198***	9.929***	8.226**	8.226***
	(3.006)	(3.881)	(3.645)	(3.103)	(3.262)	(2.896)	(3.098)	(3.247)	(2.868)	(3.036)
TFP × export share	−1.073***	−1.257**	−1.250***	−0.947**	−0.947***	−0.861***	−1.122***	−1.066***	−0.915**	−0.915***
	(0.340)	(0.432)	(0.412)	(0.339)	(0.357)	(0.326)	(0.347)	(0.367)	(0.317)	(0.335)
TFP	0.740***	0.748***	0.762***	0.523***	0.523***	0.581***	0.650***	0.659***	0.506***	0.506***
	(0.144)	(0.219)	(0.202)	(0.176)	(0.185)	(0.153)	(0.180)	(0.180)	(0.162)	(0.171)
Balance of trade dummy						0.110***	−0.032	0.026	0.002	0.002
						(0.028)	(0.033)	(0.038)	(0.031)	(0.033)
R&D						0.004***	−0.011***	−0.000	0.000	0.000
						(0.002)	(0.001)	(0.002)	(0.003)	(0.003)
Investment share						0.005	0.006	0.007	0.001	0.001
						(0.003)	(0.008)	(0.006)	(0.005)	(0.006)
Capital to labor ratio						−0.000	0.000	0.000	−0.000	−0.000
						(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Threshold export share E^*	0.689	0.595	0.610	0.552	0.552	0.675	0.579	0.618	0.553	0.553
Threshold productivity TFP^*	9.483	9.142	9.291	9.071	9.071	9.302	9.089	9.314	8.990	8.990
Industry dummies		x		x			x		x	
Year dummies	x	x	x	x	x	x	x	x	x	x
Industry × year dummies				x	x				x	x
R-squared	0.693	0.448		0.828		0.749	0.489		0.809	
Observations	188	188	188	188	188	172	172	172	172	172

Clustered standard errors in parentheses, *significant at 10%, ** significant at 5%, *** significant at 1%. As openness measure we use the firm-level export share, aggregated up to the industry level. Total factor productivity is constructed using the Levinsohn Petrin method, aggregated to the industry level. The minimum industry-level TFP is 6.972 and the maximum is 9.442. For the industry-level export share we find 0 as minimum and 0.663 as maximum value. The threshold levels are computed as $\frac{\partial \ln w}{\partial E} = TFP^*$ and $\frac{\partial \ln w}{\partial TFP} = E^*$.

Table A4: Summary statistics, by export status, plant level

	Non-Exporters		Exporters	
Variable	Mean	Std. Dev.	Mean	Std. Dev.
		Unweighted		
Exports (share of total sales)	0.000	0.000	0.334	0.255
TFP (ln)	8.116	0.766	8.518	0.697
Labor productivity (ln)	10.546	0.829	10.976	0.706
Employment (ln)	2.980	1.401	4.933	1.528
Value added (ln)	13.527	1.714	15.909	1.804
Capital intensity (ln)	10.179	1.356	10.844	1.088
Female workers (share)	0.292	0.242	0.255	0.188
High qualified workers (share)	0.754	0.239	0.714	0.241
Part-time workers (share)	0.105	0.156	0.061	0.091
CA, industry-level (dummy)	0.398	0.489	0.518	0.500
CA, firm-level (dummy)	0.079	0.270	0.108	0.310
Importer of intermediates (dummy)	0.101	0.302	0.511	0.500
		Weighted		
Exports (share of total sales)	0.000	0.000	0.251	0.235
TFP (ln)	8.071	0.726	8.304	0.670
Labor productivity (ln)	10.646	0.797	10.952	0.703
Employment (ln)	2.057	0.991	3.279	1.441
Value added (ln)	12.704	1.223	14.231	1.614
Capital intensity (ln)	10.372	1.313	10.730	1.197
Female workers (share)	0.312	0.243	0.265	0.198
High qualified workers (share)	0.713	0.251	0.702	0.251
Part-time workers (share)	0.157	0.191	0.114	0.146
CA, industry-level (dummy)	0.516	0.500	0.421	0.494
CA, firm-level (dummy)	0.042	0.201	0.083	0.276
Importer of intermediates (dummy)	0.071	0.257	0.342	0.474

Summary statistics of benchmark regression sample. Source: German matched employer-employee data (LIAB), 1996-2007, manufacturing industries. Data comprise 5,630,117 observations, obtained by matching 1,625,294 individuals to 5,392 manufacturing establishments. All monetary variables are expressed in real terms using a two-digit industry value added deflator. Weighted by inverse probability weights.

Table A5: Summary statistics, by worker types, individual level

	Blue-collar workers		White-collar workers	
Variable	Mean	Std. Dev.	Mean	Std. Dev.
		Unweighted		
Daily imputed wage (ln)	4.450	0.312	4.845	0.385
Daily non-imputed wage (ln)	4.447	0.307	4.786	0.328
Female worker (dummy)	0.133	0.339	0.256	0.437
Foreign worker (dummy)	0.139	0.346	0.032	0.175
White-collar worker (dummy)	0.000	0.000	1.000	0.000
Low-skilled worker (dummy)	0.241	0.427	0.044	0.206
Medium-skilled worker (dummy)	0.756	0.430	0.604	0.489
High-skilled worker (dummy)	0.004	0.061	0.352	0.478
Age (years)	40.940	10.152	42.328	9.832
Tenure (years)	11.562	8.096	11.159	8.375
Experience (years)	16.676	8.390	17.192	8.210
		Weighted		
Daily imputed wage (ln)	4.317	0.354	4.645	0.511
Daily non-imputed wage (ln)	4.315	0.351	4.607	0.470
Female worker (dummy)	0.158	0.365	0.338	0.473
Foreign worker (dummy)	0.121	0.326	0.025	0.155
White-collar worker (dummy)	0.000	0.000	1.000	0.000
Low-skilled worker (dummy)	0.256	0.436	0.042	0.200
Medium-skilled worker (dummy)	0.740	0.439	0.711	0.453
High-skilled worker (dummy)	0.004	0.067	0.247	0.431
Age (years)	40.693	10.520	42.033	10.297
Tenure (years)	9.455	7.691	9.388	7.890
Experience (years)	15.823	8.514	16.939	8.315

Summary statistics of benchmark regression sample. Source: German matched employer-employee data (LIAB), 1996-2007, manufacturing industries. Data comprise 5,630,117 observations, obtained by matching 1,625,294 individuals to 5,392 manufacturing establishments. All monetary variables are expressed in real terms using a two-digit industry value added deflator. Weighted by inverse probability weights.

Table A6: Production function estimates, separately for each industry

Dependent variable: Value added (ln)				
	Food	*Textiles*	*Printing*	*Wood*
Employment$\times\bar{h}_{jt}$ (ln)	0.626***	0.692***	0.891***	0.623***
	(0.036)	(0.072)	(0.060)	(0.068)
Capital (ln)	0.210	0.206	−0.080	0.241
	(0.200)	(0.384)	(0.182)	(0.171)
VC$(h_{jt})^2$	0.349	0.178	0.248	0.959**
	(0.358)	(0.504)	(0.466)	(0.482)
	[0.386]	[0.796]	[0.329]	[0.432]
Observations	2104	708	1213	1194
	Chemicals	*Plastic*	*Non-metallic*	*Metallic*
Employment$\times\bar{h}_{jt}$ (ln)	0.596***	0.618***	0.748***	0.707***
	(0.056)	(0.055)	(0.057)	(0.057)
Capital (ln)	0.195	0.312**	0.082	−0.002
	(0.141)	(0.157)	(0.156)	(0.121)
VC$(h_{jt})^2$	0.505	0.521	−0.262	−0.536
	(0.530)	(0.524)	(0.486)	(0.702)
	[0.142]	[0.636]	[0.298]	[0.014]
Observations	1286	1173	1204	1710
	Recycling	*Steel*	*Machinery*	*Vehicles, road*
Employment$\times\bar{h}_{jt}$ (ln)	0.596***	0.802***	0.764***	0.702***
	(0.095)	(0.036)	(0.038)	(0.049)
Capital (ln)	0.655**	0.226**	0.245***	0.291**
	(0.312)	(0.091)	(0.064)	(0.126)
VC$(h_{jt})^2$	−0.604	−0.542	−0.460	0.072
	(0.985)	(0.411)	(0.365)	(0.463)
	[0.385]	[0.742]	[0.914]	[0.949]
Observations	191	2718	3044	1176
	Vehicles, misc	*Electronic*	*Optic*	*Furniture*
Employment$\times\bar{h}_{jt}$ (ln)	0.628***	0.710***	0.786***	0.679***
	(0.117)	(0.057)	(0.044)	(0.111)
Capital (ln)	0.262	0.260**	0.459***	0.372
	(0.315)	(0.120)	(0.130)	(0.369)
VC$(h_{jt})^2$	0.623	−0.429	0.520	0.773
	(0.478)	(0.352)	(0.389)	(0.529)
	[0.795]	[0.810]	[0.083]	[0.897]
Observations	341	1821	1284	604

Bootstraped standard errors in parenthesis, * significant at 10%, ** significant at 5%, *** significant at 1%. All estimations include time fixed effects. Estimated by Levinsohn and Petrin (2003) and additionally controlling 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.

Table A7: Is TFP Pareto distributed?

	k-parameter	R^2	Obs.
	Pooled sample		
Total	1.135	0.737	21770
	By year		
1996	1.194	0.751	1025
1997	1.112	0.738	989
1998	1.046	0.695	1175
1999	1.131	0.709	1405
2000	1.087	0.720	2136
2001	1.109	0.727	2360
2002	1.053	0.706	2270
2003	1.073	0.697	2301
2004	1.123	0.731	2267
2005	1.110	0.739	2120
2006	1.312	0.833	1962
2007	1.315	0.815	1749
	By industry		
Food	0.972	0.820	2103
Textiles	1.034	0.686	707
Printing	1.039	0.709	1212
Wood	1.249	0.790	1193
Chemicals	1.121	0.750	1285
Plastic	1.095	0.611	1172
Non-metallic	1.177	0.714	1203
Metallic	1.173	0.711	1709
Recycling	1.006	0.699	190
Steel	1.251	0.687	2717
Machinery	1.201	0.703	3043
Vehicles,	1.018	0.714	1175
road	1.067	0.743	340
Vehicles,	1.180	0.770	1820
misc	1.226	0.714	1283
Electronic	0.994	0.629	603

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 R2 is close to one."

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