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Daniel Halvarsson
Olga Lark
Patrik Tingvall
Josefin Videnord

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LUND UNIVERSITY

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Daniel Halvarsson* Olga Lark† Patrik Tingvall‡ Josefin Videnord§

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Abstract

In this paper we study the link between globalization of firms and gender inequality. Specifically, we examine how the need for interpersonal contacts in trade and gender-specific differences in negotiations are related to the gender wage gap. Our key finding is that export of goods that are intensive in interpersonal contacts widens the gender wage gap. The effect is robust across various specifications and is most pronounced for domestic exporting firms, which do not trade within multinational corporations but with external foreign partners, where the contracting problem is most distinct. We ascribe this result to a male comparative advantage in bargaining.

Keywords: Export · Gender wage gap · Gender inequality · Contract intensity · Interpersonal contacts · International trade

JEL: J16, J31, F16, F66

*The Ratio Institute

†Lund University, the Research Institute of Industrial Economics. Corresponding author: olga.lark@nek.lu.se

‡The National Board of Trade Sweden, Södertörn University

§Uppsala University

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1 Introduction

A recent report by WTO reveals that trade policies, although designed as gender-neutral, exert differential impact on men and women. Despite joint attempts across countries at making trade more inclusive, available evidence suggests that globalization might, in certain circumstances, worsen female opportunities on the labor market (The World Bank, 2020). This illustrates a concerning gap between intended policies and the actual outcomes, and the fact that globalization and its effect on gender inequality is yet not fully understood.

In this paper we study how exports, and the associated need for communication with foreign partners, shapes the gender wage gap. An emerging literature on the topic has highlighted two personal attributes that are deemed important for exports and hence can explain the asymmetric remuneration of men and women in globalized firms. The first one is flexibility in working hours, which is needed when operating across different time zones. If women are less time-flexible, or perceived as such, they may face a relative wage penalty in exporting firms (Bøler *et al.*, 2018). Another personal attribute relates to interpersonal skills, which are arguably important for communication with foreign partners. Since women are found to have a comparative advantage in such skills, exporters could, on the contrary, generate better opportunities for female workers, and especially so in white-collar occupations (Bonfiglioli and De Pace, 2021).[1] Notably, earlier lines of research have also documented the male comparative advantage in negotiations—a valuable social skill when dealing with a diversity of cultures and social norms, intrinsic to foreign partners around the world.[2] Hence, the question of which type of gender-specific skills that are most needed and rewarded by exporters remains open. We add to the literature by documenting how the type of exported goods shapes firms' demand for particular skills and thereby drives a disproportionate wage impact across genders. Dealing with foreign contractors across the globe requires a certain degree of bargaining and relationship building. If intensified trade increases firms' demand for bargaining skills, this could offset the positive effect of the female comparative advantage in interpersonal skills on the relative wages.

To study which personal attribute that is most advantageous for exporters, we connect bargaining skills to the type of goods exported and examine how these relate to the remuneration of men and women. The connection between trade and the need for buyer–seller interaction dates back to Nunn (2007), who has shown that trade of some goods is more

[1]For the related literature on the female comparative advantage in interpersonal skills see Spitz-Oener (2006); Black and Spitz-Oener (2010); Borghans *et al.* (2014); Ngai and Petrongolo (2017); Cortes *et al.* (2021), among others.

[2]Walters *et al.* (1998); Stuhlmacher and Walters (1999); Kray and Thompson (2004); Hederos Eriksson and Sandberg (2012).

dependent on bargaining and interpersonal contacts compared to others, and therefore requires better contract enforcement and judicial quality. In this paper, we gauge the extent to which bargaining skills are crucial for exporting firms by using the Nunn (2007) measure of contract intensity (CI index), which captures the importance of buyer–seller interaction in an industry. To answer the research question, we estimate an empirical wage equation using detailed matched employer–employee data from Sweden. In addition to standard variables, such as individual and firm characteristics, we augment the wage equation with a three-way interaction term of export intensity, a female dummy, and the CI index. This allows us to scrutinize the link between goods exports, firm's need for interpersonal contact, and the gender wage gap. Inference is made by looking at how variation in firms' export intensity relates to employer–employee match dynamics. Specifically, we estimate an employer–employee match fixed effects model, which holds the within-firm gender composition constant and thereby deals with assortative matching and endogenous match quality issues. Added to that, the identification strategy we exploit mitigates concerns regarding export endogeneity with respect to wages, as well as the reverse causality between the two.

Our main finding is that when goods export intensifies, the gender wage gap widens and does more so for firms in high contract-intensive industries, where buyer–seller interaction is necessary for trade to occur. This result empathizes the role of male comparative advantage in bargaining when serving foreign markets, and that bargaining skills are important for the remuneration of workers. To further connect this finding to firms' demand for negotiation skills, we show that the observed negative effect is primarily driven by white-collar workers, and in particular, by managers and sales workers. Based on our findings, we conclude that doing business with a variety of partners across the globe changes the job skill demands of exporters.[3] In support, we demonstrate, via occupational structure analysis, that exports of contract-intensive goods shifts the composition of tasks in the respective firms, making a larger fraction of the labor force engage in the selling and bargaining operations.

An additional contribution we make is to separate the effect of foreign ownership from the effect of exporting—the two distinct but related aspects of firm globalization. As suggested by recent literature (Kodama *et al.*, 2018; Tang and Zhang, 2021; Halvarsson *et al.*, 2022), multinational enterprises (MNE) are able to transfer their corporate culture across international borders and thereby affect gender-specific labor market outcomes in the subsidiaries in the host countries. Moreover, Lanz and Miroudot (2011) document that the share of intra-firm export to total export in Swedish manufacturing sector is substantial (ca 51 percent),

[3] The occupational task-content literature was first spurred by Autor *et al.* (2003) and developed further in Acemoglu and Autor (2011). Acemoglu and Autor (2011) propose a task-based framework building on Autor *et al.* (2003),Acemoglu and Zilibotti (2001), and Costinot and Vogel (2010), among others.

implying that product contractability and negotiations is of less relevance for trade among a handful of Swedish (multinational) firms. To account for differential impact of foreign ownership and to avoid contamination of the effect of exporting, we exclude foreign-owned firms from the main sample and split the effect of interest into that exerted by foreign-owned versus domestic exporters.

In high contract-intensive industries, we find that the negative impact of exports on the gender wage gap becomes less pronounced when foreign subsidiaries are included in the sample. A closer look at the two types of exporting firms reveals a relatively strong impact on the gender wage gap from exports undertaken by domestic (non-multinational) firms as compared to multinational firms, for which neither contract intensity nor export intensity appear to exert any statistically significant effect on the relative female wages. Along these lines, we also note that domestic sales of high contract-intensive goods do not impact the gender wage gap, which further ties our findings to exporting activities. Our results are robust to a series of specification tests, including tests related to the flexibility hypothesis. We find that the lack of temporal flexibility by female workers does not appear to drive our estimates for export intensity and contract intensity. Lastly, the results indicate that male ability to negotiate is rewarded when there are rents to compete for, and that is in high contract-intensive industries. Taken together, our findings highlight a novel and important interplay between firms' demand for interpersonal and bargaining skills, the type of goods they export, and gender inequality.

The rest of the paper unfolds as follows. Next section discusses the related literature, followed by a section where we outline the basic empirical strategy and the main empirical challenges we face. The forth section describes the data, sample, and measurements used in the empirical analysis. Results, robustness checks, and discussion of potential mechanisms come in section 5, while section 6 concludes the study and provides some suggestions for future research.

2 Related literature

In a broad sense, our study connects to a vast body of work in economics on gender inequality in the labor market (Altonji and Blank, 1999; Blau and Kahn, 2000; Goldin, 2014). This literature highlights a number of factors that are considered important for explaining the gender gap in earnings, such as differences in human capital, occupational and industry segregation, temporal flexibility, as well as discrimination. Olivetti and Petrongolo (2016), Card *et al.* (2016), and Blau and Kahn (2017) constitute the more recent contributions in the field, where the focus has shifted towards the differences in psychological attributes and bargaining power across genders as potential explanations for the observed gender gaps.

Our study also relates to experimental and behavioral studies that point out three important differences in preferences and attitudes across genders. First, extensive literature reviews by Croson and Gneezy (2009) and Eckel and Grossman (2002) document that women are consistently more risk-averse compared to men. Added to that, men also show a higher willingness to lead, irrespectively of the gender-composition of the team (Born *et al.*, 2022). Second, women are also found to be reluctant to engage in competitive interactions and bargaining, and their performance and participation decrease compared to that of men, when competitive pressure increases (Gneezy *et al.*, 2003; Bowles *et al.*, 2005; Niederle and Vesterlund, 2007, 2011; Hederos Eriksson and Sandberg, 2012). The conclusions with respect to gender differences in competitiveness, however, remain context specific and generally less univocal (Marianne, 2011). Finally, by initially claiming a lower surplus, women appear to be less assertive in negotiations and as a consequence benefit less in the end (Kray *et al.*, 2001; Kray and Thompson, 2004). Altogether, this strand of literature suggests that men are better negotiators and appear to be superior to women in more risky and competitive environments. Related to this literature, our paper also connects to studies on face-to-face communication, which can enhance the transmission of information and thereby facilitates workers in their work (Battiston *et al.*, 2020).

Although this paper relates to diverse strands of the literature, we speak primarily to the debate on the role of globalization in shaping gender wage inequality.[4] In this regard, Juhn *et al.* (2014) exploit tariff reductions as an exogenous shock and analyze its effect on the gender-specific outcomes in Mexican firms. Their findings show that, by virtue of technological upgrading, trade liberalization improves labor market outcomes for female workers involved in blue-collar, but not white-collar tasks. Bonfiglioli and De Pace (2021) provide more recent evidence on the nexus between trade, female labor, and tasks and find no average effect of export intensity on the relative female wages in Germany. However, when the sample is split by occupation, the authors detect a reduction (an increase) in the gender wage gap for white-collar (blue-collar) workers as export surges. They argue that the mechanism behind their findings relates to the female comparative advantage in interpersonal skills needed to serve foreign markets. On the other hand, Bøler *et al.* (2018) who rely on matched employer–employee data from Norway, find that export increases the wage differential between college-educated men and women. A suggested explanation for this result is that female workers lack flexibility in working hours (Goldin, 2014), which is argued to be particularly important for firms trading across different time zones.

[4]Related to our question, Helpman (2018) provides a recent and extensive review of the literature on globalization and inequality in general.

Taken together, the existing literature, closest to our study, has investigated two potential mechanisms through which exporting might affect gender-specific labor outcomes: firms' temporal flexibility requirement and the role of female advantage in interpersonal skills. We contribute to the discussion with the finding that firms, tightly involved in international commerce of high contract-intensive goods, exhibit a larger gender wage gap. Our findings align with the theory suggesting that females are relatively disadvantaged in negotiations—a phenomenon underlined by experimental and behavioral studies. Hence, while changing the set of job tasks necessary for a firm to serve a foreign market, exporting tends to widen the gender wage gap by reinforcing male comparative advantage in bargaining. These findings increase our knowledge of firms' demand for gender-specific human capital and allow us to better understand the role of globalization in affecting gender wage inequality.

3 Empirical strategy

To study the nexus between exporting and wage inequality on the one hand, and the role of contract intensity on the other hand, we rely on an empirical wage equation that includes a three-way interaction term of a female dummy, firm export intensity (defined as total goods export through total sales), and the measure of a firm's contract intensity by industry. Formally, the wage of person i, employed at firm j in industry k at time t, can be written as follows:

$$\begin{aligned} ln(Wage)_{ijkt} =& \beta_1[Female_i \times (Export/Sales)_{jt} \times CI_k] \\ &+ \beta_2[Female_i \times (Export/Sales)_{jt}] \\ &+ \mathbf{X}_{it}\gamma + \mathbf{F}_{jt}\phi + \eta_{ij} + \eta_{jt} + \varepsilon_{ijkt} \end{aligned} \tag{1}$$

where $\mathbf{X}_{it}$ and $\mathbf{F}_{jt}$ are the two vectors, capturing worker and firm control variables[5], while η_{ij} and η_{jt} denote employer–employee match fixed effects and firm–year fixed effects, respectively—the two central parts of our identification strategy.[6]

Firstly, worker and firm fixed effects, which are embedded in match fixed effects, remove omitted variable biases associated with individual and firm characteristics that are constant over time, such as individual worker ability or wage-setting practices of a particular firm.

[5] For individual workers, the model controls for the potential labor market experience (*Experience*) and its square (*Experience*^{2}), the dummy variable for having children in household under 18 years old (*Children*), a dummy variable for college education (*College*), and a dummy variable for a white-collar occupation (*White collar*). As for the firm level controls, we use (log) sales interacted with the female dummy (*Female*$\times$*ln(Sales)*).

[6] In Table A2 in the Appendix, we present several versions of model (1) with different levels of fixed effects. Table A2 also shows the estimates of all control variables that are included in all specifications.

They also control for match differences between individuals that may arise due to worker labor market sorting. If, for example, exporting firms have a higher propensity of recruiting individuals with certain characteristics (including gender), or alternatively, if certain individuals seek to work for exporting firms, estimates of the gender wage gap will be biased. In a similar vein, if poorly matched workers exit firms first, when a negative shock hits the firm, it would also bias the estimated effects of interest. When making the inference, within-match identification allows us to exploit a finer source of variation stemming from firms' export activity and to hold within-firm gender composition constant. Moreover, match fixed effects reduce concerns associated with possible reverse causality, since an individual worker is unlikely to exert a sizable effect on firm's export decisions.

Secondly, the interaction of firm and year fixed effects accounts for both aggregate confounders (through year dummies) and for firm-specific unobserved time-varying heterogeneity. The latter might, for example, encompass concurrent changes to firm-specific labor demand and/or changes to its workforce composition. Most importantly, firm-level productivity shocks, which might simultaneously affect both the exporting behavior of firms and workers' wages and therefore make exports endogenous, are captured by firm–year fixed effects. In the most stringent specification, we also augment model (1) with firm–occupation–year fixed effects to control for contemporaneous shocks to firm productivity that might differently affect occupations and lead to changes in firm occupational composition, as well as individual–firm–occupation fixed effects to ensure the effects we find are not driven by workers switching occupation during the period of study.

Our identification strategy thus relies on the assumption that, after controlling for match fixed effects, observable worker and firm characteristics, and unobservable time-varying firm characteristics, firm export decisions are orthogonal to other shocks that may impact workers' wages trajectories. The empirical approach applied in this paper is reminiscent of earlier studies on globalization and gender wage inequality, in particular Bøler *et al.* (2018) and Bonfiglioli and De Pace (2021). The large number of fixed effects included in the model prevents us from using standard transformations to handle fixed effects in panel data estimation. With high-dimensional fixed effects as in model (1), particular algorithms must be used to handle the dimensionality problem and we rely on the algorithm by Correia (2016), allowing to include multiple levels of fixed effects. The standard errors are clustered at firm level in all specifications.[7]

[7] Our main results hold if we cluster standard errors at the industry or regional level. These results are available upon request.

4 Data, sample, and measurements

4.1 Data

For our investigation of how contract intensity affects the gender wage gap in internationally active firms, we use matched employer–employee data, provided by Statistics Sweden. The data span the period from 1997 to 2015 and offer detailed information on both individuals and firms, as well as customs data for exporting firms. In addition to offering high-quality data, Sweden is, much like its Nordic neighbours, a trade-dependent economy, with around 90 percent of its GDP traded.

Our main source of information on wages is the annual labor force survey, the Wage Structure Statistics (WSS), conducted by the Swedish National Mediation Office (*Medlingsinstitutet*). The survey provides full-time equivalent monthly earnings and contracted work hours, which are comparable to hourly wage rates. The survey data are available for all public sector employees with positive hours worked in the survey month (usually September). For the private sector, the survey covers all workers in firms with at least 500 employees and at least 50 percent of the remaining workforce. Specifically, private firms included in the structural business statistics (FEK) form the sampling units of the survey, which are stratified according to industry affiliation and firm size. Since we are specifically interested in the effect of exporting, public sector firms are excluded from the analysis. As a result, our sample includes approximately 50 percent of all private-sector workers (approximately two million individual workers) in any given year. Due to the stratification of smaller private-sector employers in WSS, there exist gaps for the dependent variable in some years. While it is possible to impute individual wages for the missing years, we have opted for not doing so due to the possibility of individuals to be temporarily employed elsewhere.

To the above-mentioned sample, we have merged a number of data sets: (i) the longitudinal integrated database for health insurance and labor market studies (LISA), covering all individuals in the labor force and their detailed socio-demographic characteristics; (ii) the FEK dataset, containing information on profits, sales, value added, and industry affiliation for all private non-financial companies; (iii) the labor statistics based on administrative sources (RAMS), providing information on the location and the number of employees across all plants in Sweden, whereas the plant data are aggregated to the firm level; and, finally, (iv) the Swedish Foreign Trade Statistics covering Sweden's export of goods broken down by country of destination and type of goods classification.[8] Table A1 in the Appendix contains

[8] Due to compulsory registration at Swedish Customs, the data cover all transactions in goods with countries outside the EU (Extrastat). Trade data for EU countries have been collected via a total population survey subject to a threshold, implying that the smallest transactions are excluded from the collection. In

detailed descriptions of all variables included in the analysis together with information on the data sources.

In the merged data set, all employed individuals are linked using a unique identifier to the firm, where they have earned their highest yearly income. As mentioned previously, we only consider domestic (non-multinational) exporters for the main sample. We further restrict the sample by excluding all part-time workers to avoid biases associated with part-time penalties (Manning and Petrongolo, 2008; Albrecht *et al.*, 2018), and also individuals below 18 and above 67 years of age.[9] In the final data set used for the analysis, we have at our disposal a sample of 4,886,752 worker–firm–year observations, represented by 5,166 private sector exporting firms.

4.2 Measures of contract intensity

Measuring the demand for interpersonal and bargaining skills across industries poses a considerable empirical challenge. We meet this challenge by exploiting the CI index developed by Nunn (2007), higher values of which indicate industries that rely more heavily on differentiated, or relationship-specific, input goods (Rauch, 1999).[10] Building on the Rauch (1999) commodity classification, Nunn (2007) constructs a measure which quantifies the relationship specificity of intermediate inputs used in the production of a particular final good. He argues that industries, heavily relying on differentiated intermediate inputs, are characterized by higher degree of interpersonal contact, needed between a buyer and a seller to complete a deal. Such industries are called contract-intensive. According to the theoretical framework behind the CI index, contract intensity is an exogenous industry characteristic since it stems from the peculiarities of a production process and the importance of certain relationship-specific inputs therein. In a seminal paper, Nunn (2007) demonstrates that countries, characterized by well-developed institutions, exhibit a comparative advantage and specialize in goods intensive in buyer–seller interactions, ultimately leading to a higher volume of trade in contract-intensive goods in such countries.

In this paper, the CI index is matched to firms based on the Swedish industry classification SNI2007 (equivalent to NACE Rev.2) at the 4-digit level. To take care of the firms changing industry classification during the years 1997–2015, we assign a firm to the same industry

addition to the submitted values, Statistics Sweden complements the data using information from VAT declarations to the Swedish Tax Agency.

[9]Our main results are robust to the inclusion of part-time workers. These results are available upon request.

[10]The CI index has been extensively used in the literature to answer a variety of research questions. See e.g. Altomonte and Békés (2009); Casaburi and Gattai (2009); Ferguson and Formai (2013); Bartel *et al.* (2005); Söderlund and Tingvall (2014); Strieborny and Kukenova (2016), among others.

throughout the sample period. Examples of the 4-digit SNI2007 industries with a high CI index include manufacturing of computers, graphical services before print, and breweries. In contrast, low CI index industries are, for instance, production of malt, production of meat, and manufacturing of electrical cables.

In addition to the CI index used for our main analysis, we have constructed two alternatives measures of firm contract intensity: the SPIN CI index and the Export CI index.[11] The SPIN CI index is obtained by matching the original CI index to firms based on the SPIN2007 product classification instead of SNI2007. Specifically, the SPIN2007 product classification in our data refers to the firm's main export product, as opposed to the SNI2007 industry classification, which refers to the firm's main, self-reported economic activity at large. As for the Export CI index, it reflects the ratio of all exported differentiated goods to the total firm–year export value. To make this measure exogenous with respect to current firm operations, we obtain an average Export CI index over the first three years since 1997 (or since the first year of export) and use this value for all subsequent years when we observe the firm. The first three exporting years are thus treated as pre-sample observations and are excluded from the subsequent analysis. The two alternative measures of contact intensity are firm-specific and are constructed in order to focus on the firm's export activity, rather than its domestic operations, when evaluating the importance of contract intensity for female labor market outcomes.

4.3 Descriptive statistics

In Table 1, we divide firms based on industry-level contract intensity and show sample means and medians for several firm-level variables. We define a firm as high (low) contract-intensive if it operates in an industry that lies above (below) the median CI index value in the firm-level sample.[12] Notably, the mean female share of the labor force of 28 percent is exactly identical for high and low contract-intensive firms, suggesting that both types of firms exhibit similar gender workforce composition. Firms with high CI index employ a larger number of employees on average and their mean sales are somewhat higher compared to their low CI counterparts. Also the mean export intensity is 23 percent for high contract-intensive firms and only 18 percent for low contract-intensive firms, suggesting that the latter engage less in international as opposed to domestic trade. Notably, when comparing export activity of the two types of firms across the median, they appear identical in that dimension. In sum, the division of exporting firms according to the Nunn (2007) CI index does not appear

[11] SPIN2007 is an acronym for Standard for Swedish product classification by industry. The consistent SPIN2007 product classification is available from 2000 and onwards.

[12] We follow a similar categorization of firms by contract intensity in all subsequent parts of the paper.

to generate any substantial differences in the observable firm-level characteristics across the two subsamples. To highlight possible difference in the type of workforce employed in these firms, we proceed by analyzing the individual attributes of the workers.

Table 1. Firm descriptive statistics: High versus low CI index firms

	Means			Medians		
	High CI	Low CI	p	High CI	Low CI	p
Firm size (number of employees)	294	237	0.00	64	59	0.00
Sales (mln €)	8,498	6,809	0.00	975	959	0.02
Export/Sales	0.23	0.18	0.00	0.06	0.06	0.00
CI index	0.64	0.37	0.00	0.64	0.40	0.00
Female share of labor force	0.28	0.28	0.01	0.23	0.24	0.00
Number of firms	2,619	2,547		2,619	2,547	

Notes: All numbers are based on the panel of firm-level data of domestic exporting firms for 1997–2015. Firms are classified as high (low) contract-intensive if their CI index is above (below) the median CI index in the sample. The p-value corresponds to a t-test of the null that the means/medians of the two groups are equal against the alternative that the means/medians are significantly different.

Table 2 displays individual characteristics of an average female and male worker in high versus low CI index firms. According to the observable characteristics of the workforce, high contract-intensive firms appear fairly similar to low contract-intensive firms. Specifically, while the two types of firms hire workers of similar age and having similar labor market experience, the share of workers with children remains 3 percentage points higher in firms more dependent on tight buyer–seller interaction. Another notable feature of firms in high contract-intensive industries is that they exhibit higher degree of skill intensity by way of employing more college-educated and white-collar workers. Finally, while high CI index firms pay slightly higher wages on average, the ratio of female to male wages is 0.91, which is the same for the two types of firms.

Table 2. Individual descriptive statistics: High versus low CI index firms

	High CI			Low CI		
	All	Female	Male	All	Female	Male
Monthly Wage (€)	3,476	3,238	3,542	3,041	2,837	3,112
Monthly Wage (log)	8.09	8.02	8.11	7.97	7.91	7.99
Experience	20.49	19.09	20.87	21.79	20.53	22.22
Age	42.11	41.42	42.30	42.10	41.71	42.23
Share with children	0.44	0.44	0.43	0.41	0.42	0.41
Education						
Share with college education	0.24	0.29	0.22	0.13	0.15	0.12
Occupation						
Share of white-collar workers	0.52	0.60	0.50	0.33	0.35	0.33
Share of blue-collar workers	0.36	0.27	0.38	0.53	0.50	0.54
Number of individuals	490,255	119,406	370,849	365,413	105,604	259,809
Number of individual–year obs	2,886,829	622,617	2,264,212	1,999,923	517,799	1,482,124

Notes: All numbers refer to average values of the indicated variables for the panel of worker-level data for 1997–2015. Workers belong to high (low) contract-intensive industry if the CI index of their employer is above (below) the median CI index in the sample.

Taken together, the summary statistics in Tables 1 and 2 indicate some differences between high and low CI index firms, but the average characteristics of their male and female workers still remain comparable, with the exception of skill intensity. One important finding for our analysis is that both types of firms exhibit similar gender workforce composition. Our empirical strategy, outlined in detail in previous section, controls for the potential biases that average differences presented above may bring to the estimated effects of interest.

5 Results

5.1 Contract intensity, goods exports, and the gender wage gap

Before delving into the main results of interest, we first examine the impact of variation in firms' export on the relative female wages in general. To this end, we estimate model (1) with employer–employee match fixed effects and firm–year fixed effects, but without the interaction with contract intensity. In column (1) of Table 3, we document an estimate of -0.029 for Female×Export/Sales, which is statistically significant at the 5 percent level. That is, if export intensity increases by ten percentage point, the relative wage of female workers decreases by 2.9 percent on average.[13] Our initial finding suggests that changes to

[13]See Table A3 for the effect of export intensity on the gender wage gap using versions of model (1) with different sets of fixed effects. The estimate of Female×Export switches from a positive to a negative sign when match fixed effects are included in the model—a finding discovered earlier and explained in detail by

firm export activity appear to exert a negative impact on female workers in the respective firms.

Table 3. Export, contract intensity, and the gender wage gap

Dep. var: ln(Wage)	(1)	(2)	(3)	(4)
Female×Export/Sales×CI		-0.118***	-0.109***	-0.093***
		(0.037)	(0.026)	(0.019)
Female×Export/Sales	-0.029**	-0.016**	-0.011***	-0.009**
	(0.014)	(0.007)	(0.004)	(0.004)
Match FE	yes	yes	yes	no
Firm×Year FE	yes	yes	no	no
Firm×Year×Occup. FE	no	no	yes	yes
Match×Occup. FE	no	no	no	yes
Observations	4,886,752	4,886,752	4,306,607	4,048,976
Adj R^2	0.930	0.930	0.937	0.943

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College, White collar, and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

Column (2) of Table 3 shows the key result of estimating model (1) with the interaction between the female dummy, firm export operations, and the CI index. As before, we find a negative association between goods export intensity and relative female wages. In particular, the gender wage gap widens as export intensity goes up, and it does more so in high contract-intensive industries. At the mean level of the CI index (0.54), we find a negative estimate of -0.016 for the Female×Export/Sales interaction.[14] Considering an extreme case, when goods export intensity changes from zero to one for firms with mean level contract intensity, the relative female wages would, on average, decrease by 1.6 percent. We can also interpret the result through the lens of a one standard deviation increase in export intensity, which would yield an approximately 0.5 percent decrease in the relative female wages.[15] As for the interaction Female×Export/Sales×CI, the estimate of -0.118 implies that the gender wage gap increases more when firm contract intensity is high.

To ensure that the effects in column (2) are not driven by shocks to any particular occupation, we augment model (1) with Firm×Year×Occupation fixed effects. Occupations are defined according to the Swedish Standard Classification of Occupations (SSYK96) at a detailed three-digit level. As seen in column (3) of Table 3, the extended fixed effects do not alter the main conclusions, but make the estimates slightly less negative. The Female×Export/Sales interaction also becomes less precisely estimated. Notably, the results are robust to the

Bøler *et al.* (2018).

[14]In all specifications, we choose to use a demeaned version of the CI index for its easier interpretation.

[15]The standard deviation of export intensity in the sample is 0.31.

inclusion of both Firm×Year×Occupation and Firm×Individual×Occupation fixed effects in the same specification, as in column (4). Results in columns (3) and (4) suggest that neither shocks to particular occupations, nor employees switching an occupation within the firm due to increased export intensity appear to drive our main findings. When comparing the obtained estimates in Table 3, we observe a persistent and negative sign throughout, as well as estimates similar in sign, magnitude, and precision. We interpret the stability of our findings as conclusive evidence of that increased export intensity yields, on average, a larger gender wage gap, whereas the effect appears to be consistently stronger for firms in high contract-intensive industries.

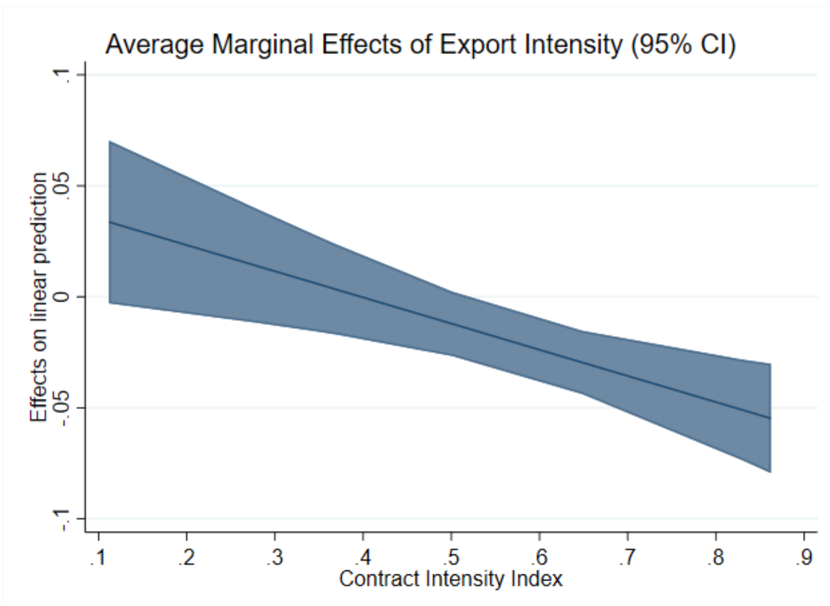

Figure 1. Marginal effects of goods export intensity

Figure 1 corresponds to the estimates in column (2) of Table 3. It shows average marginal effect of goods export intensity (y-axis) for different levels of the CI index (x-axis). The linear prediction is surrounded by 95 percent confidence bounds. The estimates are statistically significant at the 5 percent level as long as the confidence bounds do not intersect the zero horizontal line. As expected, we observe in Figure 1 that the estimate at the mean value of the CI index (0.54) corresponds to the estimate for Female×Export/Sales of -0.016, found in Table 3. The figure also clearly illustrates a negative slope of the interaction with contract intensity, indicating that the effect of increased export intensity on the gender wage gap becomes more negative with higher values of the CI index. When the CI index is below the mean value, the estimated effects are not statistically significant. If we interpret the magnitude of the estimate for the most contract-intensive firms with the CI index above 0.85, we observe a coefficient of approximately -0.05. In these firms, the gender wage gap would increase by 1.6 percent from a one standard deviation increase in export intensity, a number which highlights the economic importance of our findings. Figure 1 provides visual evidence of our key finding that export in high contract-intensive industries induces a wider gender wage gap.

5.2 Robustness

5.2.1 Robustness of main results

To validate our main findings, we apply a series of modifications to the baseline model. The results of this analysis are presented in Table 4, where the restrictions are indicated by the column heading. Due to data noisiness for smaller firms and due to convention in the literature, we exclude firms with less than 50 employees in column (1).[16] Another potential concern relates to recently hired employees, who might be differently affected by increased export intensity compared to their tenured colleagues. To address this issue, we exclude all workers with less than three years of tenure in column (2). The estimates for the two interaction terms, Female×(Export/Sales)×CI and Female×(Export/Sales), hardly change when we consider firms with more than 50 employees, whereas both estimates become slightly more negative when we restrict our focus to tenured workers. Hence, the sign and the size of the main interaction effects found in column (2) of Table 3 survive both of these restrictions.

Table 4. Robustness I

Dep. var: ln(Wage)	> 50 employees (1)	3+ yrs tenure (2)	Manufacturing (3)	Incl. fgn-owned (4)	Only fgn-owned (5)
Female×Export/Sales×CI	-0.119***	-0.123***	-0.110***	-0.082**	-0.001
	(0.039)	(0.040)	(0.042)	(0.036)	(0.025)
Female×Export/Sales	-0.016**	-0.015*	-0.021**	-0.013**	-0.002
	(0.008)	(0.008)	(0.008)	(0.005)	(0.005)
Match FE	yes	yes	yes	yes	yes
Firm×Year FE	yes	yes	yes	yes	yes
Observations	4,627,318	2,968,108	2,575,261	9,094,119	4,055,687
Adj R^2	0.929	0.939	0.939	0.932	0.938

Notes: Estimates are based on the worker-level panel data over 1997–2015. Workers of the following exporting firms are considered: (i) domestic in columns (1)–(3), (ii) all in column (4), (iii) only foreign-owned in column (5). Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College, White collar, and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

So far, we have included in the analysis all sectors of the economy. It is however common

[16] The potential source of errors with smaller firms has to do with the changes in firm identifiers in the event of a merger or an acquisition. To circumvent this problem, Statistics Sweden has created FAD-identification numbers that hold a firm ID constant as long as the majority of its employees is present in the two consecutive years. The method becomes less reliable when the total number of employees is small, hence the increased risk of errors and possible inconsistencies in firm IDs of small firms.

in the international trade literature in general, and in the recent papers on globalization and the gender wage gap (e.g. Bøler *et al.* (2018)), to consider only manufacturing sector, where the most of export activity occurs. When focusing on manufacturing firms, we find negligible changes to the estimates, compared to the benchmark estimates in Table 3. Next, when we add foreign-owned firms to the sample in column (4), our main finding still holds—increased export intensity negatively affects female workers and more so in high contract-intensive industries. The estimated effects are now smaller in absolute value and less significant, which can possibly be explained by intra-firm trade that foreign MNEs are involved in. Notably, in column (5) when we only consider foreign-owned exporters, neither of the estimates remain significant. This result alludes to the fact that firm contract intensity does not play a role in intra-firm trade.

We proceed by examining whether our main results are robust to measurement of contract intensity and firm assignment into industries. In the main part of the analysis, we use the Nunn (2007) CI index to measure firm contract intensity, which we merge to firm data based on industry (SNI2007 classification). As a first robustness check, we rely on the same CI index but we match it to firm-level data using a different industry classification. Specifically, we reassign all firms to an industry based on their main exported product (SPIN2007 product classification). The results from the new SPIN industry mapping are shown in column (1) of Table 5. Although we now connect firm contract intensity to its export operations more directly, we find quite similar estimates. For example, the estimate of the interaction term Female$\times$(Export/Sales)$\times$CI using SPIN industry matching is -0.086, compared to -0.118 using the standard SNI2007 industry mapping. The results indicate that our main finding appears to be robust to the alternative definition of firm industry affiliation.

Next, we construct a firm-specific measure of contract intensity, Export CI, which reflects the fraction of differentiated exported goods at the firm–year level. We obtain the new measure by matching firm export products from the customs data to the Rauch (1999) classification of products. Then the fraction of differentiated good of a firm is determined by its export value of all differentiated products weighted by the total export value in each year. The new measure of contract intensity confirms the negative association between exports and relative female wages found previously, as seen in column (2) of Table 5. Evaluated at the mean level of Export CI, we obtain a larger estimate in absolute value (-0.026), compared to our baseline estimate of -0.016 in column (2) of Table 3.

Table 5. Robustness II

	(1) CI SPIN	(2) Export CI Time-varying	(3) Export CI Fixed	(4) PPML	(5) Dom. sales
Female×Export/Sales×CI	-0.086**	-0.045***	-0.061***	-0.139***	-0.125***
	(0.035)	(0.017)	(0.022)	(0.043)	(0.043)
Female×Export/Sales	-0.019**	-0.026**	-0.028**	-0.019**	-0.017**
	(0.008)	(0.012)	(0.012)	(0.008)	(0.008)
Female×Dom.Sales×CI					-0.007
					(0.012)
Female×Dom.Sales					-0.001
					(0.002)
Match FE	yes	yes	yes	yes	yes
Firm×Year FE	yes	yes	yes	yes	yes
Observations	4,065,202	4,814,550	3,608,677	4,886,752	4,886,752
Adj. R^2 / Psuedo R^2	0.936	0.930	0.937	0.934	0.930

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. The dependent variable is deflated and annualized wage in log form in columns (1),(2),(4), and (5) and in levels in column (4). Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College, White collar, and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

In column (3), we make the Export CI measure exogenous to current firm operations by using a predetermined share of differentiated export goods, which is fixed across the years. Although the exogenous Export CI measure yields fewer observations and less variation due to exclusion of the pre-sample years, we find reassuringly similar estimates in column (3) and column (2), where we allow Export CI index to vary over time.

To test for the sensitivity of the results with respect to the choice of estimator, we in column (4) employ a Pseudo Poisson Maximum Likelihood (PPML) estimation method, which allows to dispense with the log-linear form of equation (1). The results correspond to the same specification as before but the wages are now estimated in levels. According to our findings, the main conclusion continue to hold irrespective of the functional form of the empirical wage equation we use.

A potential concern could be that the effects we find with respect to firm contract intensity are not driven by its international operations and sales abroad, but rather to an increase in sales in general, be it domestic or foreign. We have partly already addressed this concern since we have factored in domestic sales in our measure of export intensity (Export/Sales). To make an additional check we augment the baseline specification with the Female dummy interacted with domestic sales and a triple interaction Female×Domestic sales×CI. According to column (5) of Table 5, neither domestic sales nor domestic sales interacted with the

CI index exert any statistically significant effect on the gender wage gap.[17] Hence, the inclusion of domestic sales as a separate control does not appear to affect our findings and further strengthens our claim that its specifically globalization and exports that are driving the results.

5.3 Heterogeneity analysis

Our key finding that in high contract-intensive industries, the gender wage gap widens when firms are exposed to intensified export is robust to alternative measures of contract intensity and pertains to international, rather than domestic sales. The effect we find might not, however, be homogeneous across different groups of workers. If we consider bargaining with partners in other countries, for example, it seems likely that the tasks associated with negotiations are carried out by workers in particular occupations and/or with particular educational background. To explore the heterogeneous effects of intensified export across employee subgroups, we divide the full sample of workers into subsamples with respect to their educational attainment and observed occupation.

Table 6. Heterogeneity: Education and Occupation

	Education		Occupation	
Dep. var: ln(Wage)	College (1)	No college (2)	White-collar (3)	Blue-collar (4)
Female×Export/Sales×CI	-0.102***	-0.100***	-0.146***	0.006
	(0.030)	(0.028)	(0.035)	(0.025)
Female×Export/Sales	-0.020***	-0.012**	-0.016**	-0.002
	(0.007)	(0.006)	(0.007)	(0.006)
Match FE	yes	yes	yes	yes
Firm×Year FE	yes	yes	yes	yes
Observations	805,962	4,060,382	2,446,447	2,401,198
Adj R^2	0.949	0.904	0.946	0.807

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College (columns (3) and (4)), White collar (columns (1) and (2)), and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

In the first two columns of Table 6, the sample is split based on workers' highest attained educational level. In column (1) and (2), we observe that the estimates for workers with and without college education almost mirror each other and also align with the previous estimates in column (2) of Table 3 for all employees. In a firm with average level of contract

[17] We also test our model on a sample of non-exporting firms. The results show no indication of domestic sales exerting an effect on the relative female wages. The results are available upon request.

intensity, the wage of a college-educated woman appears to decrease by 0.6 percent relative to the wage of a college-educated man, once export intensity increases by one standard deviation. Hence, the negative association between exports and relative female wages, which is increasingly negative in contract intensity, appears to hold for all workers, irrespective of their educational level.

Next, we run the baseline specification separately for white-collar and blue-collar workers in columns (3) and (4), respectively. Despite a close similarity in sample size across the two subgroups, with around 2.4 million observations each, we only find statistically significant effects for white-collar workers. The result implies that the widening of the gender wage gap due to increased export exposure is concentrated among workers performing white-collar tasks, with no notable impact on their colleagues in blue-collar tasks. Focusing on the white-collar workers and holding contract intensity fixed at the mean, the estimated effect of a one standard deviation increase in export intensity on the relative female wages amounts to approximately -0.5 percent. Moreover, the coefficient of the triple interaction with contract intensity is -0.146, implying that the observed negative effect is increasing in the degree of contract intensity. This finding aligns with our main research hypothesis, which pertains mostly to white-collar workers since white collars are more likely to be active in tasks directly connected to bargaining.

As stated earlier, we expect that doing business with a variety of customers across the world changes the labor demand of firms and presumably shifts their occupational composition. To investigate how increase in exports might affect employment composition of firms, we calculate the weighted average employment shares across four white-collar occupation categories and firm export intensity. To adjust for compositional differences across industries, we divide a firm's employment share in a particular category by its weighted industry average employment share. When obtaining weighted industry average employment shares, we consider all firms in the sample. We also divide firms into 4 subgroups, by contract intensity (firms in high VS low contract-intensive industries) and export intensity (firms below VS above median export intensity). A value of 1 in Table 7 indicates that the firm-level and the industry-level weighted averages are equal or, in other words, firms of a particular type display a similar employment share as their counterparts operating in the same industry.[18]

When considering the results in Table 7, we notice at least two striking features in the employment composition of the two types of firms. Notably, both high and low contract-intensive exporting firms employ relatively more managers compared to their industry av-

[18]Note that the figures in Table 7 might happen to be all above or below 1 for a particular occupational category. If firm's occupational share is substantially above that of the industry, it can generate outlier observations affecting the resulting figures. To circumvent this problem, we apply weighted average employment shares across industries and different firm types.

erage. Specifically, exporting firms with low CI index use 13 percent more management workers compared to firms in the same industry. At the same time, a corresponding figure for exporters in high contract-intensive industries remains around 1.10, meaning that these firms employ 10 percent more managers compared to an average firm in the same industry.

Table 7. Firm occupational structure by contract intensity and firm export intensity

	High CI		Low CI	
Export intensity	Below median	Above median	Below median	Above median
Managers	1.09	1.10	1.13	1.13
Sales workers	1.17	1.23	1.12	1.07
Tech workers	0.94	0.87	1.25	1.05
Support workers	1.03	0.99	1.03	0.98

Notes: The numbers represent weighted average shares of employment in firms divided by the weighted industry average employment shares, by occupation category. The firms are divided according to their industry-level contract intensity and their export activity (above and below median export intensity in the sample). The sample is a panel of all firms over the years 2001–2015.

The most important finding for our paper, however, is that high contract-intensive firms rely more heavily on sales workers compared to their low contract-intensive counterparts. Namely, if we consider high CI firms with above median export intensity, we notice that the relative share of sales workers in these firms remains at 1.23, meaning that such firms employ almost a quarter more sales personnel compared to other firms in the same industry. For high CI firms with below median export intensity, the relative share of sales workers is lower and is equal to 1.17, suggesting that shifts in occupational structure are likely to be triggered by both the intensity of export operations and the type of goods sold (differentiated or homogeneous). Overall, our findings lend support to the idea that firms, trading differentiated products on the international market, require a different set of skills of their workers and therefore exhibit different occupational composition. Such changes in employment structure of high contract-intensive firms are consistent with them re-orienting operations towards more extensive trading activity, necessary to serve foreign markets.

To probe deeper into the heterogeneous effects of exports across white-collar occupations, we divide workers by four occupational categories outlined above and rerun the baseline specification separately for each subsample. The results in column (1) of Table 8 show that female managers, on average, are negatively affected by increased export intensity. The gender wage gap among managers is also increasing in contract intensity, as given by the negative estimate of the Female×Export/Sales×CI interaction term. Our results therefore suggest that changes in the task content associated with higher export intensity are more beneficial for male managers relative to female managers. We argue that this finding can largely be attributed to the male comparative advantage in tasks required by firms with a global reach. Negotiations, as discussed earlier, is one area where previous literature has established that men tend to outperform women. An increasing importance of activities related to negotiations serves as a plausible explanation for our findings for managers.

Table 8. Occupations

	White-collar occupations			
	Managers	Sales	Tech	Support
Dep. var: ln(Wage)	(1)	(2)	(3)	(4)
Female×Export/Sales×CI	-0.144**	-0.131	-0.092***	-0.010
	(0.071)	(0.084)	(0.017)	(0.028)
Female×Export/Sales	-0.028	-0.024	-0.012***	-0.025***
	(0.017)	(0.018)	(0.004)	(0.008)
Match FE	yes	yes	yes	yes
Firm×Year FE	yes	yes	yes	yes
Observations	280,367	320,259	800,611	661,043
Adj. R^2	0.959	0.901	0.946	0.955

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College, and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

For sales personnel, we also observe an increasing gender wage gap for firms exposed to trade in high contract-intensive industries, with the estimated effect of -0.131. For this subgroup, we do not observe a significant effect of either the double or the triple interaction on the gender wage gap at the mean CI level, but the size of the coefficients are around the same magnitude as for managers. One possible explanation for the lack of precision in column (2) may be that the category of sales workers is too broadly defined, encompassing both domestic and international sales workers. In column (3), we also find that the relative wages of female tech workers decrease when export intensity increases, and it does more so when a firm is operating in more contract-intensive industry. The estimates of Female×Export/Sales×CI and Female×Export/Sales are -0.092 and -0.012, respectively, and are statistically significant. Finally, for support workers, as seen in column (4), the main effect of interest is small and less precise. We do, however, find an increasing gender wage gap among support workers when export intensity surges.

Taken together, the in-depth subgroup analysis reveals two important empirical facts about globalization and relative female wages. First, the relationship between export intensity and the gender wage gap, which is negative and increasing in contract intensity, tends to hold for workers of all educational levels. More interestingly, we find that intensified export in high contract-intensive industries appears to negatively affect relative wages of white-collar female workers, while exerting no effect on blue-collars female workers. Our results further indicate that the adverse effect on the relative female wages is most pronounced among managers and sales workers, although also negative for tech and support female workers.

5.4 Mechanisms

5.4.1 Gender-biased rent sharing

Our results suggest that globalized firms, depending on their need for bargaining skills and the degree of contract intensity, tend to treat their employees differently. If that is the case, the male comparative advantage in negotiations is likely to be reflected in firm rent-sharing behavior too. That is, the two types of firms will share rents in a different manner, depending on the nature of goods they export. To explore this mechanism, we perform a rent-sharing analysis following Card *et al.* (2016) and Bruns (2019), where we estimate a within-firm rent-sharing model across genders.[19] The exercise is reminiscent of the design employed in the preceding rent-sharing literature (Guiso *et al.*, 2005; Card *et al.*, 2014; Carlsson *et al.*, 2016), but the estimation is performed on two disjoint samples of male and female workers. In this analysis, we therefore no longer rely on individual level data and within-match identification, but instead consider long-term differences in firm performance and average wages.

In essence, we examine within-firm variation in productivity (measured by excess value added per worker) and wages over time. The idea is to purge rent-sharing estimates from all time-invariant firm attributes by focusing on job stayers, i.e. workers who remain in the firm over a certain period of time (three years). This approach helps to eliminate biases generated by permanent firm heterogeneity and indicates whether relative rent-sharing elasticity of female wages might vary across different types of firms. We believe that this exercise offers convincing evidence on gender bargaining differentials and its association with the degree of firm contract intensity.

Similar to Card *et al.* (2016) and Bruns (2019), we report both male and female rent-sharing coefficients and the bargaining ratios. The latter are obtained by running a two-stage IV regression of average wage changes of female stayers on average wage changes of male stayers, instrumented by excess log value added (the normalized measure of firm surplus), separately for the two types of firms. When obtaining average wage changes, we also estimate two specifications: a basic model, which includes year fixed effect, and an extended model, which includes year, 2-digit industry, and region fixed effects. The results are summarized

[19] A detailed discussion of the model and the estimation procedure are presented in the Appendix.

Table 9. Rent-sharing models for male and female three-year stayers (1997–2015, excess log value added)

	Basic model			Extended model		
	Rent-sharing coef-s			Rent-sharing coefs		
	Male	Female	Ratio M/F	Male	Female	Ratio M/F
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. High CI firms, excess log value added per worker, 1997–2015, three-year stayers						
Three-year change, winsorized at $+/-0.75$	0.033	0.029**	0.895**	0.033*	0.028***	0.863***
	(0.022)	(0.012)	(0.367)	(0.017)	(0.010)	(0.199)
Three-year change, trimmed at $+/-0.75$	0.034	0.030**	0.897**	0.033*	0.029**	0.867***
	(0.025)	(0.013)	(0.396)	(0.019)	(0.011)	(0.224)
To restrictions	0.030	0.027**	0.896**	0.031**	0.027***	0.864***
	(0.020)	(0.011)	(0.351)	(0.016)	(0.009)	(0.188)
Panel B. Low CI firms, excess log value added per worker, 1997–2015, three-year stayers						
Three-year change, winsorized at $+/-0.75$	-0.004	0.010	-1.171	-0.007	0.000	-0.006
	(0.010)	(0.004)	(5.643)	(0.009)	(0.010)	(1.464)
Three-year change, trimmed at $+/-0.75$	0.004	0.009	2.231	0.002	0.007	4.257
	(0.010)	(0.010)	(3.877)	(0.009)	(0.010)	(17.005)
To restrictions	-0.003	0.003	-0.989	-0.006	-0.002	0.339
	(0.010)	(0.010)	(6.205)	(0.007)	(0.009)	(1.267)

Notes: The entries show coefficients of three-year wage changes of male and female stayers on three-year changes in excess log value added per worker. Wage changes are adjusted for a quadratic polynomial in age. Three-year changes in excess value added are adjusted for year fixed effects (the basic model) or, alternatively, for year, 2-digit industry, and region fixed effects (the extended model). Ratios in columns (3) and (6) are obtained via two-stage least squares, instrumenting the male firm effect by log excess value added. All models are estimated at the firm-year level (domestic exporting firms only), weighted by the total number of person years in the base year. Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

in Table 9. Panel A and B report estimates for firms in high and low contract-intensive industries, respectively.

Already at first glance, we detect some differences in the rent-sharing behavior of firms. As seen in columns (1), (2), (4), and (5) of Panel A, high contract-intensive firms tend to share rents with their workers, and they do unevenly so with men versus women, as indicated by ratios in columns (3) and (6) of Panel A. More specifically, females wages in high contract-intensive firms are 86-89 percent as responsive to changes in firm productivity as male wages are. Notably, unequal rent-sharing across genders in the Swedish labor market context has been earlier outlined by Nekby (2003). On the other hand, the estimated coefficients in Panel B for firms in low contract-intensive industries are statistically insignificant in almost all specifications and also small in magnitude in columns (1), (2), (4), and (5), suggesting that these firms do not appear to share rents with their employees.

One complication with the rent-sharing analysis is that firm surplus tends to vary drastically over time, let alone its measurement difficulty. To probe whether large variability in firm productivity might introduce attenuation bias to the estimates, we use three versions of the three-year changes in the excess log value added per employee, as stated in the row names of Table 9. Winsorizing and trimming the variable does not affect either the rent-sharing coefficients, or the bargaining ratios in Panel A. The obtained estimates are, therefore, robust to alternative restrictions on the surplus measure and inclusion of industry and region fixed effects. Taken together, the results suggest that firms in high contract-intensive industries share rents with their employees, whereas male workers benefit more from the increased firm surplus compared to female workers. On the contrary, we detect no rent-sharing by low contract-intensive firms with either of the genders. Apart from facilitating men in export-related job tasks, the documented male comparative advantage in negotiations appears to also allow them claiming higher rents compared to female workers. As expected, we find no evidence of rent sharing in low contract-intensive firms, since bargaining skills and other similar attributes are less needed in the industries where buyer and seller do not interact.

5.4.2 Flexibility in working hours

As proposed by Goldin (2014) and tested empirically in the international trade context by Bøler *et al.* (2018), the female lack of temporal flexibility could contribute to the widening of the gender wage gap. Arguably, there is a possibility that high contract-intensive firms require of their employees more commitment when their export operations intensify. In other words, the two types of firms may differ in their demand for temporal flexibility. To investigate this possibility, we examine how the flexibility hypothesis relates to firm contract intensity in shaping the gender wage gap in exporting firms. To this end, we subsequently exclude from the initial sample workers that are considered to be less flexible in time. In particular, we exclude workers: i) who have children between 0 and 6 years old, and ii) who are below 45 years old. We deem that both of these criteria allow us to identify workers who either have or plan for children and thus are more likely to be time-constrained.

In column (1) of Table 10, the baseline results from estimating model (1) for all workers are presented. These findings correspond to column (2) of Table 3, with an estimate of -0.118 for the interaction between female, export intensity, and the CI index, and -0.016 for the interaction between female and export intensity. Around one million of observations in our sample belongs to workers with young children, and these are excluded from the sample in column (2). When excluding workers with young children, the baseline estimates stay intact. Similarly, the exclusion of workers under 45 years old in column (3) does not challenge the baseline results either. The exclusion of workers under the age of 45 substantially shrinks the number of observations from around 4.9 million to 2 million, as people aged 18 to 44 constitute a vast majority of the workforce.

Table 10. Robustness: Temporal flexibility

Dep. var: ln(Wage)	Baseline (1)	No child 0-6 (2)	Age>44 (3)	High CI (4)	Low CI (5)
Female×Export/Sales×CI	-0.118***	-0.126***	-0.132***		
	(0.037)	(0.038)	(0.045)		
Female×Export/Sales	-0.016**	-0.014**	-0.015**		
	(0.007)	(0.007)	(0.007)		
Female×ln(BusHours)				-0.005	-0.002
				(0.003)	(0.004)
Match FE	yes	yes	yes	yes	yes
Firm×Year	yes	yes	yes	yes	yes
Observations	4,886,752	3,877,889	2,058,797	2,096,393	2,719,692
Adj. R^2	0.930	0.936	0.960	0.946	0.911

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College, White collar, and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

Apart from excluding workers considered less time-flexible, we do an additional check and interact the female dummy with a firm's business hours overlap variable in columns (4) and (5) of Table 10. Following Bøler *et al.* (2018), the variable is constructed as an average business hour overlap across the exported country–product combinations relevant for a firm in a given year. Also in line with Bøler *et al.* (2018), the business hours overlap is calculated assuming standard office hours between 9.00 and 17.00 and using the average values for countries with multiple time zones. In a specification similar to model (1), the main interaction of interest is now replaced with Female×ln(Business hours overlap) variable. The results displayed in columns (4) and (5) indicate no difference between high versus low contract-intensive firms in the estimated effect of interest. In particular, the estimates are -0.005 and -0.002 for high and low CI firms, respectively, and neither of them reaches conventional significance levels. To sum up, the tests of the temporal flexibility hypothesis in this section provide no evidence of firms' contract intensity and the demand for worker commitment being necessarily related.

6 Conclusions

We add to the literature on globalization and gender inequality by analyzing how the nature of exported goods matters for the gender wage gap. We establish that export of goods in contract-intensive industries disproportionately benefits men, leading to a widening of the gender wage gap. We ascribe this result to men being better in negotiations compared to women. The negative impact on the gender wage gap is, however, limited to white-collar workers. Furthermore, we find the strongest effect among managers and sales workers, as compared to workers in other white-collar occupations, indicating that trade-related bargaining skills are particularly valuable in these occupations.

Arguably, cross-border transactions are subject to greater contracting and communication problems, and especially so if they occur with a foreign partner outside of the same company group. In line with this assumption, we find that there is no effect on the gender wage gap of domestic sales, whereas the effect is strongest in the sample of domestic exporters. The latter result suggests that the need for, and the remuneration of, bargaining skills is especially large when firms operate with foreign contractors, external to the firm.

For the empirical analysis, we rely on matched employer–employee data from Sweden and estimate a multi-level fixed effects wage regression, subject to a range of specification tests. Our results are robust to alternative definitions of contract intensity, alternative estimation methods and model specifications, and to the inclusion of firm–year–occupation, and firm–individual–occupation fixed effects. When exploring the mechanisms behind our results, we find that men tend to claim higher rents relative to their female colleagues in industries where more buyer–seller interaction is needed. Male negotiation style appears to benefit them when there are rents to compete for in high contract-intensive firms. As an extension of the analysis, we also examine the complementary mechanism behind our findings and establish no evidence of our results being driven by firms' demand for temporal flexibility.

To sum up, we attribute the increased gender wage gap in exporting firms to the male comparative advantage in bargaining. On a similar note, an interesting finding in psychology and behavioural literature is that women tend to be more cooperative, while men remain more competitive in negotiation situations. This suggests that firm's trading partners are also likely to exert a certain impact on its wage-setting practices. We leave this intriguing question for future research—a question, we believe, can cast additional light on how globalization facilitates or hampers female labor market opportunities.

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A Tables

Table A1. Variable definitions

Variable	Source	Notes
Individual level		
Individual identifier	WSS	All individual identifiers are anonymized by Statistics Sweden using a serial number. The actual social security numbers come originally from the population register.
Wage	WSS	Full-time equivalent monthly earnings (comparable to hourly wage rates) relate to the survey month, usually September. The monthly wage data corresponds to the agreed-upon wage on top of amenities and variable incomes, absent over-time payments. The information is available for all employees in the public sector. In the private sector, the data are available for all workers in firms with at least 500 employees; for the rest, a stratified sample based on the industry affiliation and firm size is used. As a result, roughly 50 percent of private-sector workers are included in the sample in any given year. We use annualized and deflated data of wages expressed in EUR.
Occupation	WSS	Detailed information on worker occupation codes (up to 4 digits) is available from the salary structure statistics. The Swedish Standard Classification of Occupations (SSYK96) is based on the International Standard Classification of Occupations (ISCO-88). We define workers as white-collar workers if their one-digit occupation code is 1,2,3, or 4, and as blue-collar workers otherwise (occupation codes 5,6,7,8, or 9). A dummy variable for white-collar workers is included as a control variable in our main empirical specification.
Age	LISA	Individual's age, original source: the population register.
Gender	LISA	Individual's gender, original source: the population register.
Education	LISA	We define a worker as having university education (as being skilled) if he or she has at least two years of post-secondary education. To derive the variable, we rely on the information on the highest completed level of education (original source: the education register). The original education variable takes on the following values: (1) Primary and lower secondary education, less than 9 years; (2) Primary and lower secondary education, 9 or 10 years; (3) Upper secondary education; (4) Post-secondary education, less than two years; (5) Post-secondary education, two years or longer; and (6) Postgraduate education.

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Table A1 – continued from previous page

Variable	Source	Notes
Experience	LISA	The actual labor market experience of individuals is not available in our data. Instead we construct potential labor market experience as a difference between individual's age and (1) the year of obtaining highest level of education, or (2) years of education (based on the education variable). If neither of the variables are available, we subtract 16 to obtain potential labor market experience.
Children	LISA	The number of children below 18 years old is available from LISA database (original source: population register). We define a dummy taking a value of 1 if there is at least one child below 18 years old in the household.
Firm-level		
Firm identifier	FAD	The FAD-identifier is obtained from The Dynamics of Enterprises and Establishments Database (FAD, Statistics Sweden) and is developed to correct for administrative changes in firm legal identifiers over time and. The principle behind the FAD-identifier is that it remains the constant from one year to another if the firm's actual identifier has changed, but a majority of workers stay in the firm between the two consecutive years. More details on FAD-identifiers are available at Statistics Sweden website.
Export of goods	Swedish Foreign Trade Statistics (SFTS)	The Swedish Foreign Trade Statistics covers Swedish export (and import) of goods broken down by destination country and (8-digit) type of goods classification. The data on goods trade cover the years 1997–2015. To measure goods export intensity, we use goods export sales over total sales (FEK) for each firm and year. This yields the export intensity measure varying between 0 and 1 for each firm–year observation.
Firm size	RAMS	To measure firm size, we aggregate the number of employees from the establishment level data (RAMS) to the firm level.
Sales	FEK	Sales is the firm's total turnover from goods and services in million EUR, deflated using CPI and turned from SEK into EUR using the Swedish central back annual average exchange rate. Excise duty is excluded, and the measure is also adjusted for merchanting.
Value added	FEK	Value added is defined as the firm's production value minus the cost of intermediate inputs. The measure does not include wages, social fees, or the cost of goods resold without processing, for which only the trade margin is included in the production value.
Export CI index	SFTS & Rauch (1999)	The export CI index reflects the ratio of all exported differentiated goods to the total firm–year export value. It builds on export data at the product-firm-year level combined with the Rauch (1999) classification of goods. The Export CI index is constructed as an average over the first three years since 1997 (or since the first year of export).
Industry-level		

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Table A1 – continued from previous page

Variable	Source	Notes
CI index	FEK & Nunn (2007)	The Nunn (2007) contract intensity index is matched to firms based on the Swedish industry classification SNI2007 (equivalent to NACE Rev.2) at the 4-digit level, which refers to the firm's main, self-reported economic activity.
SPIN CI index	SFTS & Nunn (2007)	The Nunn (2007) contract intensity index is matched to firms based on the Swedish product classification by industry (SPIN2007). SPIN2007 industry classification reflects the firm's main export product.

Table A2. Export, contract intensity, and the gender wage gap: Alternative specifications

Dep. var: ln(Wage)	(1)	(2)	(3)	(4)	(5)
Female×Export/Sales×CI	-0.014	0.023	0.024	-0.109**	-0.118***
	(0.066)	(0.067)	(0.067)	(0.043)	(0.037)
Female×Export/Sales	0.034***	0.038***	0.037***	-0.014*	-0.016**
	(0.012)	(0.012)	(0.012)	(0.008)	(0.007)
Female×CI	-0.022	-0.048*	-0.048*		
	(0.029)	(0.028)	(0.028)		
Export/Sales×CI	-0.139**	-0.070		-0.054	
	(0.055)	(0.043)		(0.042)	
Female	-0.186***	-0.188***	-0.190***		
	(0.037)	(0.036)	(0.037)		
Export/Sales	0.031***	-0.000		0.005	
	(0.009)	(0.009)		(0.008)	
CI	0.094***				
	(0.024)				
ln(Sales)	0.026***	0.013***		0.011***	
	(0.003)	(0.003)		(0.003)	
Female×ln(Sales)	0.002	0.002	0.002	-0.001	-0.001
	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)
Experience	0.017***	0.017***	0.017***	0.009***	0.009***
	(0.001)	(0.001)	(0.001)	(0.000)	(0.000)
Experience^2/100	-0.029***	-0.028***	-0.028***	-0.030***	-0.029***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Children	0.027***	0.027***	0.027***	0.001	0.001
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
College	0.254***	0.246***	0.245***	0.081***	0.077***
	(0.006)	(0.006)	(0.006)	(0.003)	(0.003)
White collar	0.237***	0.236***	0.237***	0.034***	0.035***
	(0.007)	(0.007)	(0.007)	(0.005)	(0.005)
ln(Firm size)	-0.013***	-0.015**		0.013***	
	(0.004)	(0.007)		(0.005)	
Industry×Year FE	yes	yes	yes	yes	no
Firm FE	no	yes	no	no	no
Firm×Year FE	no	no	yes	no	yes
Match FE	no	no	no	yes	yes
Observations	4,886,752	4,886,752	4,886,752	4,886,752	4,886,752
Adj R^2	0.532	0.561	0.565	0.926	0.930

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

Table A3. Export and the gender wage gap: Alternative specifications

Dep. var: ln(Wage)	(1)	(2)	(3)	(4)	(5)
Female×Export/Sales	0.034***	0.039***	0.039***	-0.025	-0.029**
	(0.013)	(0.013)	(0.014)	(0.015)	(0.014)
Female	-0.173***	-0.168***	-0.169***		
	(0.035)	(0.036)	(0.036)		
Export/Sales	0.032***	-0.003		0.005	
	(0.010)	(0.010)		(0.009)	
Industry×Year FE	yes	yes	yes	yes	no
Firm FE	no	yes	no	no	no
Firm×YearFE	no	no	yes	no	yes
Match FE	no	no	no	yes	yes
Observations	4,886,752	4,886,752	4,886,752	4,886,752	4,886,752
Adj R^2	0.532	0.561	0.565	0.926	0.930

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College, White collar, and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

Table A4. Export and the gender wage gap: Heterogeneity by education and occupation

	Education		Occupation	
Dep. var: ln(Wage)	College	No college	White-collar	Blue-collar
Female×Export/Sales	-0.038***	-0.020*	-0.039**	-0.002
	(0.012)	(0.011)	(0.018)	(0.005)
Match FE	yes	yes	yes	yes
Firm×Year FE	yes	yes	yes	yes
Observations	805,962	4,060,382	2,446,447	2,401,198
Adj R^2	0.949	0.904	0.946	0.807

Notes: Estimates are based on the worker-level panel data over 1997–2015, whereas workers employed in domestic exporting firms are considered. Additional control variables included in all specifications are: Experience, Experience^2/100, Children, College (columns (3) and (4)), White collar (columns (1) and (2)), and Female×ln(Sales). Standard errors clustered at the firm level are in parentheses. Significance levels: *** ($p < 0.01$), ** ($p < 0.05$), and * ($p < 0.1$).

B Rent sharing

To evaluate the relative rent-sharing elasticity of female wages across high and low contract-intensive firms, we examine how variation in firm productivity relates to its average wages over time. In this analysis, we focus on job stayers, i.e. workers who remain in the firm over a certain period of time. That allows to adjust the rent-sharing estimates for time-invariant firm attributes and eliminate the associated bias.

Following Card *et al.* (2016), denote individual worker by $i \in \{1, \ldots N\}$ over period $t \in \{1, \ldots T\}$, worker i's gender as $G(i) \in \{M, F\}$, and his or her employer in year t by $J(i,t) \in \{1, J\}$. Assume further that

$$S_{jt} = \lambda \max\left\{0, S^0_{jt} - \tau\right\} + \varsigma_{jt} \equiv \lambda NS_{jt} + \varsigma_{jt} \tag{2}$$

where S_{jt} is an actual surplus per worker for firm j in year t, also defined as net surplus; S^0_{jt} is an observed surplus measure for firm j in year t; τ is a threshold value under which the net surplus shared by firms is assumed to be zero, as explained in more detail below; and ς_{jt} is an error term with mean zero.

Using a theoretical model outlined in Card *et al.* (2016), one can derive the following equation:

$$\begin{aligned}&\mathbb{E}\left[w_{iT} - w_{i1} | NS_{J(i,1)1}, NS_{J(i,1)T}, X_{i1}, X_{iT}, G(i), stayer\right] \\ &\quad = (X_{iT} - X_{i1})'\beta^{G(i)} + \theta^{G(i)}\left[NS_{J(i,1)T} - NS_{J(i,1)1}\right]\end{aligned} \tag{3}$$

where $\theta^{G(i)} = \lambda\gamma^{G(i)}$ and *stayer* is shorthand for conditioning on worker i being continuously employed at the same firm j throughout the sample period.

By running an OLS regression of model (3), we obtain estimates of θ^M and θ^F, ultimately providing the estimate of the relative rent-sharing elasticity across genders, or the so called *bargaining ratio*, $\frac{\gamma^F}{\gamma^M}$. Finally, according to the Card *et al.* (2016) model, the bargaining ratio relates to the average wage changes of male and female stayers. Formally:

$$\frac{\mathbb{E}\left[w_{iT} - w_{i1} - (X_{iT} - X_{i1})'\beta^F | female, stayer, J(i,1) = j\right]}{\mathbb{E}\left[w_{iT} - w_{i1} - (X_{iT} - X_{i1})'\beta^M | male, stayer, J(i,1) = j\right]} = \frac{\gamma^F}{\gamma^M} \tag{4}$$

In the analysis, equation (4) is estimated using a two-step IV approach. To this end, a regression of change in wages on covariates and firm dummies is run separately by gender to produce residualized average firm wage changes. In the second step, the adjusted average change in female wages is regressed upon the corresponding average male change, instrumented by the change in measured surplus and weighted by the total number of stayers at the firm.

Getting back to the threshold τ, when obtaining the residualized wage changes from the two disjoint samples, one has to normalize firm effects to make them comparable across samples. It means we need to impose a linear restriction on the firm fixed effects such that they are identified with respect to the same reference firm or firms. According to the Card *et al.* (2016) model, the true firm effects are non-negative and are equal to zero for firms offering no surplus. Hence, imposing the following restriction provides a set of normalized firm effects coinciding with their true counterparts:

$$\mathbb{E}\big[\psi^{G(i)}_{J(i,t)}|\bar{S}^0_{jt} \leq \tau\big] = 0 \qquad G(i) \in \{M, F\}, \tag{5}$$

where $\psi^{G(i)}_{J(i,t)}$ are gender-specific firm effects. The coefficient τ is then identified by simultaneously estimating the two equations of labor productivity and firm effects across genders. The threshold for no-surplus firms is derived at the point where the coefficient of determination is the highest. The estimated threshold is used to obtain the *excess log value added* – a normalized version of the firm surplus.

C Female share of employees

We have found a negative effect of export intensity on the relative female wages that gets more pronounced with contract intensity. In this section, we would like to examine whether firms in high contract-intensive industries change gender workforce composition to a larger extent than their low contract-intensive counterparts when export intensity surges. The firm-level yearly female shares are calculated as the number of female employees divided by the total number of employees in the respective category, indicated in the column heading of Table A5. In this analysis, we consider both part-time and full-time workers, unlike other parts of the paper where only full-time workers are examined. A firm-level model can be written as follows:

$$\begin{aligned} FemaleShare_{jkt} = &\beta_1 Export_{jt} + \beta_2[Export_{jt} \times CI_k] \\ &+ \beta_3 DomSales_{jt} + \beta_4[DomSales_{jt} \times CI_k] \\ &+ \eta_{kt} + \eta_j + \varepsilon_{jkt} \end{aligned} \tag{6}$$

where $FemaleShare_{jkt}$ is the share of female workers in firm j in industry k observed in year t and other variables are defined as before. We choose to include domestic sales, $DomSales_{jt}$, and its interaction with the CI index, $DomSales_{jt} \times CI_k$, to control for its potential effects on the female labor share. In addition to that, we consider two sets of fixed effects in model (4): (i) industry-year fixed effects (η_{kt}), and (ii) firm fixed effects (η_j), and thereby adjust for systematic variations in workforce composition across industries, as well as for unobserved time-invariant firm characteristics.

Table A5. Female share of employees

	All (1)	White Collar (2)	Blue Collar (3)	College (4)	New Hire (5)
Export×CI	-0.019	0.012	0.033	-0.058	-0.000
	(0.036)	(0.060)	(0.278)	(0.095)	(0.136)
Export	0.012*	0.030***	-0.007	0.017	-0.006
	(0.006)	(0.011)	(0.061)	(0.017)	(0.024)
Domestic Sales×CI	0.001	0.026	0.183***	-0.004	0.097
	(0.026)	(0.021)	(0.070)	(0.042)	(0.072)
Domestic Sales	0.001	0.007*	-0.029***	-0.003	0.010
	(0.005)	(0.004)	(0.011)	(0.006)	(0.012)
Industry×Year FE	yes	yes	yes	yes	yes
Firm FE	yes	yes	yes	yes	yes
Observations	24,761	19,138	19,138	19,240	21,787
Adj R2	0.961	0.839	0.599	0.798	0.459

Notes: The dependent variable is female share of all employees in column (1), female share of white collar employees in column (2), female share of blue collar employees in column (3), female share of college educated employees in column (4), and female share of new hires in column (5). Robust standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

As shown in Table A5, the interaction between export intensity and the CI-index show no significant correlation with the female share of employees in any of the studied groups. That is, we do not find any evidence of high contract-intensive firms hiring females in a different proportion compared to low contract-intensive firms when their export intensity increases. Instead, we find that higher export intensity appears to be correlated with an increase in the overall share of female employees, as stated in column (1), and also with the shares of white-collar females (column (2)). This could indicate higher skill-intensity of exporting firms. An increased export intensity does not, however, seem to significantly correlate with the female share of blue-collar workers in column (3), college educated workers in column (4), or newly-hired workers in column (5). Hence, export intensity seems to correlate with the gender composition of the firm, particularly when it pertains to workers in white-collar tasks, but the correlations do not seem to differ across firms with different degrees of contract intensity.