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D3.3. The Role of Flexible Wage Components in Gender Wage Differences

WP3 Digital Transformation: Impact on skills and inequality

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Contents

Abstract

A main driver of the gender gap wage gap is the fact that women have a lower chance to enter high-paying firms. Also, even upon entering, they receive a lower share of the firm-specific wage premium than their male co-workers. We use a novel Hungarian linked employer-employee dataset and AKM decomposition to show that performance and overtime payments are main drivers of these gender differences in firm premia. One fifth of the total gender wage gap can be attributed to the fact that women receive a lower share of the firm specific wage premium at firms with overtime and performance payments. At the same time, labor productivity or firm size has a negligible effect on the gender difference in firm-specific wage premium conditional on the wage structure.

Keywords: wage inequality, bargaining, sorting, overtime, performance payments JEL codes: J31

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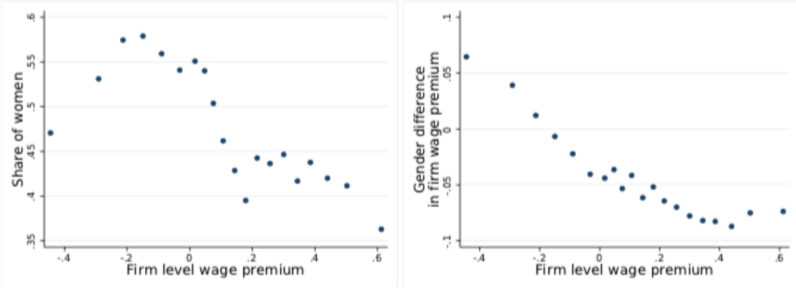
1. Introduction

The gender wage gap has been rapidly falling in industrialized countries since the Second World War (Olivetti & Petrongolo, 2016). One of the key driving forces behind this trend is technological progress. Due to technological progress, particularly computerization, the tasks executed by men and women converged, which also decreased occupational and workplace segregation (Black & Spitz-Oener, 2010). This trend, however, slowed down during the last two decades (England et al., 2020) as gender segregation at top-paying firms is not decreasing any more. Therefore, women still have a lower chance to enter firms with the highest wage premia, and even when they do, they tend to earn less than their male co-workers (Bruns, 2019).

For example, in the case of Hungary (Figure 1), the share of women is 10 percentage points lower at firms that have an above-median wage premium compared to firms that are below the median (see Figure 1). In the meantime, the higher the premium firms pay on average, the less they pay on average to women relative to men.

Figure 1: Gender differences in participation and wage premium by the quality of firms



Notes: The figure shows the share of women and the difference in gender-specific wage premia by firm-level average wage premium. See Section 3 for the estimation procedure.

It is debated in the literature why women have a weaker position at these firms. One strand of the literature (Card et al., 2016; Casarico & Lattanzio, 2019; Sin et al., 2022) argues that women have lower bargaining power, thus, they can extract a lower share of the firm-level wage premium than otherwise similar men. In contrast to these, there may be differences in preferences as the highest-paying firms have some non-pay characteristics which decrease the willingness of women to enter such firms, or even hinder their ability to obtain the same firm-level premium as men (Sorkin, 2017). Finally, the differences in wage structure may play a key role in gender selectivity at large firms. The reason for this is that high-paying firms offer performance payments (Bloom & Van Reenen, 2007, 2010) and require overtime hours (Reizer, 2022) more often, and women are shown to earn less under these types of work arrangements (Albanesi & Olivetti, 2009; Goldin, 2014). These differences in the wage structure are endangering further closure of the gender pay gap at high-paying firms as globalization (Bøler et al., 2018) and technological progress (Lemieux et al., 2009) are increasing the prevalence of these types of wage arrangements.

The aim of our paper is to quantify to what extent performance payments and overtime payments can explain gender participation and wage gaps at high-paying firms. We use a novel Hungarian administrative linked employer-employee database with information on the employment history of 50 percent of the Hungarian population between 2003 and 2017 and also on the prevalence of performance and overtime payments at private sector firms. Our analysis proceeds in two steps. First, we estimate firm and gender-specific wage premia and then we calculate the actual contribution of performance and overtime payments to the gender gap. In the first step, we estimate an Abowd, Kramarz, Margolis (1999) model with both individual fixed effects, which reflect individual earning potential, and gender (and time)-specific firm fixed effects. The latter terms reflect the wage premium of firms available to men and women. We show that the total gender wage gap in the private sector is 22.7 percent, which is somewhat larger compared to the national-level wage gap (14.7 percent) and the 20 percent OECD average (OECD, 2012).

Following the methodology of Card et al. (2016), we find that 9.5 percentage points of the total gender gap can be attributed to the gender difference in firm-specific wage premia. Using a Oaxaca-Blinder type decomposition, we show that 4.1 percentage points of this difference are due to the fact that women work at firms with a lower (overall) wage premium (*sorting effect*). The remaining 5.4 percentage points can be attributed to the *bargaining effect*, namely that women receive a lower share of the firm premium than their male co-workers, even within the same firm.

As the second step of the empirical analysis, we turn to the investigation of performance and overtime payments. We show that the allocation of performance and overtime payments are not random. Larger firms which innovate and participate in international trade are more likely paying flexible wages. Besides women are less likely to work in flexible wage jobs. For instance, only 63.4% of women in our data receive overtime payments, while for men this ratio is 68.4%. This difference is solely due to composition effects as women are more likely to work at occupations where flexible wage schemes are less prevalent. Turning to the wage effect of flexible wages, we show that the gender gap in firm premium is only 1 percent at firms where workers do not receive any flexible wage components, while the gender gap is linearly increasing in the share of workers with overtime or performance payments. This difference is not solely driven by composition effects as it remains significant once we control for differences in sector, size or the productivity of the firms. We find that the gender gap in firm premium is 5.1 percentage points larger at firms where every worker receives performance payments compared to firms where no worker receives performance payments. This difference is 3.9 percentage points in the case of overtime payments. The Oaxaca-Blinder decomposition reveals that 3-3.5 percentage points of this difference can be attributed to a sorting effect, as women are less likely to work at firms where they receive performance payments or where overtime is frequent, and the remaining 1-1.5 percentage points to the within-firm effect (*bargaining*). The estimated contribution of overtime and performance payments is significant in economic terms, as conditional on firm size and wage structure, a 10 percent increase in firm productivity corresponds only to a 0.09 percent increase in the gender gap in firm premium.

We contribute to several strands of the literature. First, we augment the literature on flexible wages. There is widespread evidence showing that flexible wages increase worker productivity, and firms that measure and reward worker effort are on average more productive and profitable (Bender et al., 2018; Bloom et al., 2016; Ichniowski et al., 1997). The cost of these work arrangements is that they increase income inequality within the firm (Bandiera et al., 2007; Bidwell et al., 2013; Lazear, 2000; Lemieux et al., 2009; Shearer, 2004). We add to the literature that this increase in inequality hurts women disproportionately. Second, we contribute to the literature on the glass ceiling, the phenomenon that the gender wage gap at the top of the wage distribution is much larger than the average gender wage gap

(Albrecht et al., 2003; Arulampalam et al., 2007; Christofides et al., 2013). There are many factors contributing to the glass ceiling effect, such as lower work experience because of motherhood (Bütikofer et al., 2018; England et al., 2016), less working hours (Azmat & Ferrer, 2017; Goldin, 2014) or gender differences in social interactions (Cullen & Perez-Truglia, 2019). We add to the literature by showing that the gender gap is much smaller at firms which do not use performance or overtime payments.

Finally, the results of Biasi & Parsons (2022) are the closest to our results. Using a policy reform in Wisconsin, their study provides causal evidence showing that flexible wages increase the gender gap among teachers. They show that an important mechanism contributing to the gender gap is that women negotiate for wages less often. We add to the paper by analysing not only one occupation, but showing instead that flexible wages increase the gender gap at the level of the whole economy as well.

The rest of the paper is structured as follows. Section 2 presents short facts about the Hungarian labor market and the two datasets we use. Section 3 introduces our methodological framework for decomposing the gender wage gap and quantifying the role of flexible wage components. Section 4 contains the results of our empirical exercises. Section 5 concludes.

2. Institutional background and data

2.1 Institutional background

Hungarian employment contracts have to classify whether the worker is salaried and paid monthly or by the hour. Even if the worker is salaried, firms can require additional working hours above the regular schedule, but in this case, the firm has to pay overtime payments on an hourly basis. Besides working time, wage contracts have to specify the monthly or hourly base wage which can be decreased only with the written consent of workers. On top of the base wage and overtime payments, firms can pay additional bonuses, premia or allowances. These additional side payments are entirely determined by the employer and are not regulated by the Labor Code. The prevalence of overtime payments and performance payments have a similar magnitude in Hungary as in other countries in the European Union (Druant et al., 2009; Kézdi & Kónya, 2011).

The wage-setting institutions in Hungary are similar to those in Anglo-Saxon countries. The share of union members are low (OECD, 2012), wage bargaining takes place at the individual level (Rigó, 2012) and it is relatively easy to lay off workers compared to other Western-European countries (Tonin et al., 2009).

2.2 Data

We use two main data sources for our empirical investigation. The first dataset is the ADMIN3, the administrative linked employer-employee dataset of the Centre for Economic and Regional Studies (Sebő́k, 2019). The ADMIN3 contains the work history of a 50 percent random sample of the Hungarian population between 2003 and 2017. This sample consists of approximately 5.4 million distinct individuals, of whom 3.4 million are observed for at least one month as working at a private firm or a public employer. It consists of the employment status of individuals on the 15th day of every month and monthly gross earnings based on social security contribution payments.

On the firm side, the ADMIN3 contains information on the balance sheet and income statement of the employing firm. The source of the balance sheet data is the yearly corporate income tax returns collected by the National Tax and Customs Administration. In Hungary, every firm has to report their financial data as part of the tax declaration forms, thus, we observe the balance sheet of every firm which employed at least one person in the worker sample of ADMIN3.

A shortcoming of the ADMIN3 is that it contains only the total salary without information on specific wage components. That is why we match ADMIN3 data with the Hungarian Structure of Earnings Survey (HSES). The Structure of Earnings Surveys are available in every country of the European Union and contain detailed information on the specific wage components earned by the workers in a specific month of the year. In contrast to most other countries, the Hungarian version is conducted yearly instead of every four years, and contains the balance sheet of the firm. The Hungarian version contains information on wage structure earned in May.

In the rest of the paper, we consider a worker to have performance payments if she received either monthly or occasional bonuses, premia or allowances. While we consider a worker to receive overtime payments if she received additional payments for overtime hours or for weekend and night shifts. For the sake of simplicity, we refer to a worker as one who receives flexible wages if she received either overtime payments or performance payments.

The HSES uses a stratified sampling design and contains yearly information on 10 thousand firms and 170 thousand workers. At firm level, the HSES has a panel structure. Firms with more than 50 employees have to participate every year while only a random sample of firms between 5 and 50 employees have to report wages on their workers. At worker level, the HSES has a repeated cross-sectional structure. We observe every worker whose firm has less than 50 employees and a 6-10 percent random sample of workers whose firm has more than 50 workers[1].

It is not possible to merge the two datasets at the individual level. That is why we match the two datasets at the firm-year-occupation level with probabilistic matching. We considered a firm in a given year the same in the two distinct datasets if they were both unique and identical with respect to reported financial data characteristics (including sales, reported size, total assets, etc.) in both datasets. Utilizing the panel structure of both datasets, we matched firms based on data from all fifteen observed years. This greatly decreased the number of cases where observable characteristics do not uniquely identify firms. To make the method robust for potential data errors, we allowed the variable vectors of firms to differ between the two datasets in one out of the fifteen years. This way, we could unambiguously match 99.1-99.9% of employers from the HSES wage survey to employers in the ADMIN3 dataset in the years between 2004 and 2016. For observations from 2003, match quality was 96.6%, and for 2017, it was only 85.7%. Therefore, we omit the latter year from our estimations presented in this paper. We cannot link the public sector part of the HSES to ADMIN3 due to administrative reasons, thus, we use the public sector to estimate the individual and firm fixed effect in the AKM model, then we restrict attention to the private sector in the main analysis.

After creating the link between the two datasets, we calculated the number of workers receiving performance payments[2] (bonus, premia) or overtime payments2 by firm-gender-occupation-year cells in the HSES. Then, we matched these cell-level numbers to ADMIN3, and calculated the firm-year and firm-gender-year level shares used in our analyses.

We use the Community Innovation Survey (CIS) to investigates the relationship between the prevalence of flexible wage structure and innovation. This survex is repeated every second year and it is conducted in every EU countries. The CIS contains information on specific types of innovation (e.g. introduction of a new product, a new process or an organization types). Every firm with more than 50 employees and a random sample of firms with less than 50 employees have participate in the survey. We

[1] The sample covers of 6.6% of physical workers and 9.9% of white-collar workers, based on their date of birth.

[2] We consider additional payments for night and weekend shifts as part of overtime payments as well.

can merge the CIS data base to the balance sheet data but we are not able to merge them to the administrative employment and wage data due to administrative restrictions.

Sample Selection

From the ADMIN3 dataset we kept all workers who had an employment contract for the full month, either in the private or in the public sector. For computational feasibility, we used only the monthly observations from January, April, June and October. As we aim to utilize AKM firm effects, we had to restrict the sample to firms in the largest connected mobility set of workers. The reason for this is that the estimated firm parameters are comparable on the same scale only across firms which are connected by movement of workers[3]. We identify these sets separately for male and female workers. Firms which are included in the giant components in both the male and female mobility networks formed the dual-connected set, as in Card et al. (2016). To decrease the potential effect of the limited mobility bias problem (Andrews et al., 2008), we removed employment spells at firms with less then 2 mobility events (hires or separations) on average per year throughout the entire period. Despite not having information on the wage schemes employed at public sector employers, we included such observations in the estimation of firm-level and individual-level wage components in the AKM model. This way, we observe more job-to-job mobility and we can estimate the wage premium of the private sector firms more precisely. However, we restrict attention to the private sector in the main analysis.

2.3 Descriptive analysis

Table 1 shows the average characteristics of men and women in the whole sample and of workers whose employer is observed in the HSES as well. Women in Hungary earn 13.5 log points (around 14.5%) less, while this difference is 21 log points (23.4%) in the sample in which we observe the HSES firms. We observe this difference because small firms are underrepresented in the HSES survey and it does not contain information on the public sector where the gender wage gap is smaller. In line with this, the gender

[3] As Torres et al. (2018) notes, if a third set of high-dimensional fixed effects such as occupation is included in the model, three-way connected sets have to be identified. We followed the approach of Weeks and Williams (1964) to obtain these sets.

difference in firm-level wage premium is smaller in the whole sample than at the firms which we observe in the HSES as well.

Women in Hungary work at slightly larger firms than men. The average firm size of women is around 323 and only around 282 for men. This is in contrast to Portugal (Card et al., 2016) or the United States (Papps, 2012) where women work at slightly smaller firms. As explained before, the average firm size in the subsample where we observe the HSES is larger because of the sampling design. Still, the average firm size of women is slightly larger. In contrast, the value added per worker (by 16.5 percent), the share of exporting firms (by 5.3 percentage points) and the firm-level wage premium (by 3.8 percent) is all larger among firms where men work. These differences are similar for the whole sample and for the firms which we also observe in the HSES.

Table 1: Average characteristics of male and female workers

	Whole sample				The firm is observed in the HSES			
	Women	Men	diff	t-score	Women	Men	diff	t-score
Log(wage)	6.505	6.64	0.135	23.21	6.617	6.827	0.210	23.84
	(0.46)	(0.504)			(0.435)	(0.437)		
Firm-unit level premium	0.176	0.215	0.039	9.45	0.247	0.334	0.086	15.18
	(0.291)	(0.347)			(0.262)	(0.296)		
Log(value added/worker)	8.184	8.37	0.185	14.29	8.417	8.607	0.190	9.95
	(1.095)	(1.007)			(0.989)	(0.901)		
Export	0.43	0.483	0.053	10.76	0.615	0.673	0.058	6.21
	(0.495)	(0.5)			(0.487)	(0.469)		
Average firm size	322.84	281.84	-41.00	-3.75	578.97	529.81	-49.16	-2.72
	(530.53)	(487.98)			(615.6)	(586.9)		
Individual observations	19979820	28847806			8904479	12728481		
Firm obs.	277044	305379			31270	32595		

Notes: The table shows the average characteristics of men and women in the whole sample, and at firms which we observe also in the HSES.

Table 2 shows that men are more likely to earn performance payments (68.4 vs 63.4 percent) and overtime payments (56.2 vs 54 percent) as well. In contrast, we do not find these differences within occupational categories. What is more, female managers are likely to earn overtime payments with a 9 percentage points higher probability than male managers. Thus, men are more likely to earn overtime payments and flexible payments because they have occupations where these wage components are more prevalent.

Table 2: Prevalence of overtime and performance payments by gender and occupation

	Received overtime payments		Received performance payments	
	Women	Men	Women	Men
Political/religious/NGO leader (Nace 1*)	53.1%	55.6%	46.9%	41.8%
Top manager (Nace 1*)	27.7%	26.9%	41.8%	40.5%
Other manager (Nace 1*)	60.5%	51.9%	61.4%	61.7%
Professional (Nace 2)	49.5%	51.5%	62.2%	62.2%
Other white-collar (Nace 3-4)	54.1%	57.6%	60.8%	60.1%
Skilled blue-collar (Nace 5-7)	71.0%	70.1%	46.2%	53.6%
Machine operators (Nace 8)	86.2%	84.7%	58.9%	60.2%
Unskilled laborer (Nace 9)	56.4%	60.4%	32.4%	37.4%
Total	63.4%	68.4%	54.0%	56.2%

Notes: The table shows the share of workers receiving performance payments and overtime payments by occupational categories.

As firms can decide about performance and overtime payments on their own, we investigate which firm characteristics are related to flexible wage structure. Table 3 presents results from using additional information on innovation activities from the CIS dataset. The table shows that conditional on firm size and the educational level of workers, women are more likely to work at firms which offer flexible payments. Besides the table highlights that higher quality firms are more likely to use flexible wage structure. Thus firms are more likely to pay flexible wages if they are innovating, participating in international trade, or simply are larger and or productive.

Table 3: Prevalence of overtime and performance payments by firm characteristics

	(1)	(2)	(3)	(4)	(5)	(6)
	Perf. pay	Perf. pay	Overtime pay	Overtime pay	Bonuses	Bonuses
Share of females	0.043***	0.052***	-0.026*	0.031*	0.013	0.019
	(0.013)	(0.015)	(0.015)	(0.017)	(0.017)	(0.020)
Share of vocational training	-0.046***	-0.005	0.053***	0.045***	-0.074***	-0.036**
	(0.015)	(0.014)	(0.019)	(0.015)	(0.018)	(0.018)
Share of high school diploma	-0.181***	-0.175***	-0.577***	-0.454***	-0.183**	-0.213***
	(0.068)	(0.061)	(0.167)	(0.060)	(0.076)	(0.069)
Share of college	-0.018	-0.045**	-0.488***	-0.402***	0.021	0.003
	(0.017)	(0.021)	(0.019)	(0.020)	(0.022)	(0.026)
Technological innovation	0.097***	0.007	0.091***	-0.004	0.108***	0.020**
	(0.006)	(0.006)	(0.007)	(0.007)	(0.009)	(0.009)
Organizational innovation	0.046***	-0.017***	0.048***	-0.010	0.063***	-0.006
	(0.006)	(0.006)	(0.007)	(0.007)	(0.008)	(0.008)
Exporter (dummy)	0.324***	0.333***	0.021	0.018	0.253***	0.220***
	(0.030)	(0.035)	(0.073)	(0.045)	(0.034)	(0.050)
Log(size)		0.098***		0.106***		0.107***
		(0.003)		(0.003)		(0.004)
Log(valua added/worker)		0.094***		0.064***		0.080***
		(0.005)		(0.006)		(0.007)
Constant	0.393***	-0.844***	0.541***	-0.526***	0.359***	-0.748***
	(0.038)	(0.066)	(0.086)	(0.073)	(0.044)	(0.087)
Observations	28,049	25,340	28,049	25,340	28,049	25,340
R-squared	0.051	0.270	0.135	0.357	0.051	0.210

Notes: Regressions on the number of workers with various wage components. Control variables include innovation measures from the CIS. Standard errors are in parantheses. *** p<0.01, ** p<0.05, * p<0.1

3. Method

In the following, we shortly summarize the framework we adapt. Building on the works of Abowd et al. (1999), Card et al. (2013), Gerard et al. (2021), and Lachowska et al. (2019), we estimate the following three-way AKM model:

$$ln\ w_{ijtg} = \boldsymbol{X_{ijtg}}\beta + \theta_i + \Psi_{jgt} + \lambda_{k(ijt)} + \varepsilon_{ijtg} \qquad (1)$$

where the dependent variable is the wage earned by worker *i* of gender *g* at firm *j* at year *t* . X_{ijtg} denotes the time-varying individual characteristics, θ_i the worker fixed effect, ψ_{jgt} the gender and year-specific firm fixed effects, and $\lambda_{k(ijt)}$ the occupation of worker *i* in year *t*.

Previous authors mostly assume that firm effects may vary across workers of different types in the same firm, but they do not change over time within the same firm unit. Relying on the work of Bruns (2019) and Lachowska et al. (2019), we relax this assumption and allow the effects to vary over years. This allows us to consider changes to the in firm specific wage premium or to the change of flexible wage components. Following Torres et al. (2018), we also augment the model with occupation fixed effects to capture the non-random selection of male and female workers into high or low-wage occupations. This way, differences in the firm-specific components will be devoid of the effects of occupational selection.

After defining the largest connected sets (in the mobility networks) separately for male and female workers in the estimation sample[4], we estimate the fixed effects model on all firms that are part of *either* the female or the male connected sets. The estimated gender-specific firm-year effects are initially comparable on the same scale only within the male or the female connected set – as there is no mobility between the two disjoint components of the labor mobility network. To overcome this issue, we follow the strategy put forward by Card et al., 2016, and assume that there should not be differences in firm-specific premia by gender among the set of firms with the lowest productivity.

[4] We followed the algorithm of Weeks and Williams (1964), as proposed by Torres et al. (2018) for the cases of more than two high-dimensional fixed effects.

The estimated and rescaled firm-group effects then can be decomposed, in a second stage, into a composite of an average gender difference in firm-specific wage premium and (baseline) firm effects:

$$\Psi_{jgt} = G_g \tilde{\beta}_G + \tilde{\psi}_{jt} + \varepsilon^G_{jgt} \tag{2}$$

$\tilde{\psi}_{jt}$ in this formulation is a firm-year specific effect that is already cleaned of the gender composition of the firm. $\tilde{\beta}_g$, the average within-firm effect of gender, is identified on the set of dual-connected firms, that is, firms that are included in *both* the male and female connected sets in the mobility network. ε^G_{jgt} has an expected value of zero, while its variation captures how much of the firm-gender specific effects are not captured by separate firm fixed effects and an assumed to be constant gender difference.

Using this second stage, an alternative to the decomposition of Card et al. (2016) can be provided by taking the difference of Equation 2 across gender groups G .

$$\frac{\partial \Psi_{jgt}}{\partial G} = \tilde{\beta}_G + \frac{\partial \tilde{\psi}_{jt}}{\partial G} \tag{3}$$

In practice, we estimate a set of simple regressions to obtain the above components. Specifically, the left hand side term, representing the overall difference in estimated firm-group effects between groups of any G_g can be obtained from the following single regression:

$$\Psi_{jgt} = G_g \tilde{\beta}_{\mathrm{OA}} + \varepsilon_{jg} \tag{4}$$

Then, repeating this regression by incorporating fixed effects for all firm-year pairs will yield an estimator of within-firm differences in gender-specific premia, $\tilde{\beta}_{WI} = \tilde{\beta}_G$

$$\Psi_{jgt} = G_g \tilde{\beta}_{\mathrm{WI}} + \tilde{\psi}_{jt} + \varepsilon^G_{jg} \tag{5}$$

Finally, a second stage regression ran on the stored firm effects of Equation 5 will provide the between-firm average difference of the firm-group parameters, capturing sorting in this specification.

$$\tilde{\psi}_{jt} = G_g \tilde{\beta}_{\text{BW}} + \epsilon_{jt} \tag{6}$$

We also note that $\tilde{\beta}_{OA} = \tilde{\beta}_{WI} + \tilde{\beta}_{BW}$, as it is reflected in Equation 3. As Boza (2022) discusses in detail, this specification is a close alternative to the decomposition of Card et al. (2016), with some convenient features[5].

[5] Card et al. (2016) proposes the following Oaxaca-Blinder style decomposition of the differences in observed firm-gender fixed effects.

$$E(\Psi_{j\text{M}}|\text{M}) - E(\Psi_{j\text{F}}|\text{F}) = \underbrace{E(\Psi_{j\text{M}}|\text{M}) - E(\Psi_{j\text{M}}|\text{F})}_{\text{Sorting}} + \underbrace{E(\Psi_{j\text{M}}|\text{F}) - E(\Psi_{j\text{F}}|\text{F})}_{\text{Bargaining}} = \tag{7}$$

.

$$\underbrace{E(\Psi_{j\text{M}}|\text{M}) - E(\Psi_{j\text{F}}|\text{M})}_{\text{Bargaining}} + \underbrace{E(\Psi_{j\text{F}}|\text{M}) - E(\Psi_{j\text{F}}|\text{F})}_{\text{Sorting}} \tag{8}$$

This decomposition utilizes two counterfactual states for the mid-point of the decompositions. It either allocates the firm premia experienced by female workers to the male workers of their firms, or it takes the female distribution over firms as a given and assumes that the latter group would receive the same premia as male workers. Bruns (2019) argues that the latter specification is more relevant, while Casarico and Lattanzio (2019) simply report the arithmetic mean of the elements in the two different possible decompositions. Boza (2022), on the other hand, provides a regression-based formulation of the problem, which provides an unambiguous decomposition of the firm-group-specific wage components into a within-firm (bargaining) and a between-firm (sorting) component. This approach also has the advantages of being easily generalizable into differences across multiple groups, and also allows for the inclusion of control variables or other parameters of interest.

4. Results

First, we will utilize the decomposition method introduced in Section 3 to illustrate the basic channels and patterns of the gender wage gap in Hungary. Then, we augment the baseline approach to magnify how flexible wage components contribute to the whole gender wage gap.

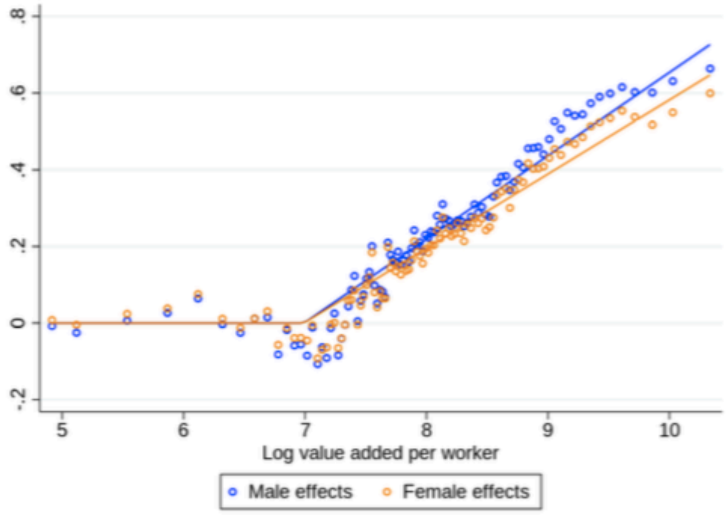
Bargaining and sorting

Since male and female workers cannot belong in the same connected set in the mobility network where firm-gender(-year) units are treated separately, we cannot identify directly comparable firm-gender(-year) effects (Abowd et al., 2002). Thus, we normalize firm-specific wage premia before the empirical investigation. We follow the strategy of Card et al. (2016) and Bruns (2019), and assume that the least productive firms provide zero firm-level wage premium to both men and women, while firms with a higher productivity provide positive premium to both gender.

In line with our normalization assumption, Figure 2 shows that the wage premium does not co-move with productivity along the set of low-productivity firms, while it increases among firms of a higher productivity. Similarly to the Portuguese (Card et al., 2016) and German (Bruns, 2019) estimates, we also find that the fitted linear regression line for women is flatter. This means that women receive a lower share of any firm premium than men.

Finally, for precise normalization, we have to find the kink point where firm-level wage premium starts to increase with productivity. For this purpose, we fitted several kinked regressions and chose the kink point from the specification producing the best fit (lowest RMSE across the separately fitted male and female regressions). Based on this criterion, we set the kink point at 6.99.

Figure 2: Re-scaled gender-firm-year effects versus log value added per worker of firms



Notes: Data points are means of estimated firm-gender fixed effects corresponding to a hundred percentiles of firm-year observations along the distribution of the logarithm of value added per worker for firms with balance sheet data available in the ADMIN3 dataset. Firm-gender-year effects are both normalized to have a zero mean value in both gender for observations below a log value added of 6.99 – the threshold that provided the best fit for the kinked function presented on the graphs.

As the second step of the empirical analysis, we calculate the contribution of firms to the gender wage gap. For this purpose, in Table 4, we report the parameters from Equations 4, 5 and 6 from three different sets of regressions: one utilizing the whole time period (row 1), one having a different parameter estimated for two distinct time periods (rows 2 and 3), and a specification capturing the difference between the latter two parameters in regression form (row 4).

Table 4: Decomposition of the gender wage gap - HSES sample

	Gender gap log(wage)	Overall firm premium Ψ_{jgt}	WI (bargaining) Ψ_{jgt}	BW (sorting) $\bar{\psi}_{jt}$
2003-2016	0.227***	0.095***	0.054***	0.041***
	(0.003)	(0.002)	(0.001)	(0.001)
2003-2009	0.234***	0.105***	0.061***	0.044***
	(0.004)	(0.002)	(0.001)	(0.002)
2010-2016	0.219***	0.086***	0.047***	0.038***
	(0.004)	(0.002)	(0.001)	(0.002)
Difference	-0.015**	-0.019***	-0.013***	-0.006*
	(0.005)	(0.003)	(0.001)	(0.003)
Observations	108,228	108,228	108,228	108,228
Firm effects	x	x	✓	x
Number of units	1	1	19703	1

Notes: The table shows the contribution of firm pay premia to the gender wage gap using the modified method of Card et al. (2016), presented in Section 3 and Boza (2022). Row 1 is based on regressions on the whole time period; rows 2 and 3 are regressions with distinct parameters for the two time periods. Row 4 contains the difference of rows 2 and 3 (from a regression formulation of different parametrization). See Section 3 for details. Standard errors are in parentheses. *** p<0.001, ** p<0.01, * p<0.05

The table shows that the gender wage gap in the whole period was 22.7 percent. During the 2000s, the gender gap was 23.4 percent, which decreased by 1.5 percentage points to the 2010s. The contribution of differences in firm premium explained 9.5 percentage points of the whole gender wage gap. 5.4 percentage points (around a quarter) of the gender gap is attributable to the fact that women earn less than men at the same firm, even after controlling for occupational selection (*bargaining channel*). Differences in *sorting*, namely the fact that women work at firms with a lower firm-specific wage premium explains 4.1 percentage points of the total wage gap. Finally, we see that not only the total gender gap decreased between the two decades but also the difference in firm-specific wage premium. The bargaining effect decreased by 1.3 percentage points, while the sorting effect decreased only by 0.6 percentage point, slightly increasing the relative importance of the latter channel.

The role of flexible wage components

To better understand the role of flexible wage components, we split our sample into three parts based on the share of flexible wage components at the firm level. In Panel A of Table 5, we show workers who work at a firm where more than 95 percent of workers receive flexible wage components. In Panel B, this ratio is between 5 and 95 percent, while in Panel C, we show the observations where less than 5 percent of the workers receive flexible wage components.

The main message of the table is that the higher the share of workers with flexible wages at the firm, the wider the gender wage gap. In Panel A, where the prevalence of flexible wage components is very high, the gender wage gap is 24.8 percentage points, while in Panel C, where flexible wage components are rare, this difference is only 12.3 percentage points. Furthermore, the table highlights that this difference in gender wage premium is driven mostly by gender-specific firm premia. At firms where almost everybody receives flexible wage components, the gender difference in firm premium is 11 percentage points. Half of this difference comes from different gender sorting across firms, and half of it comes from bargaining, namely from the gender gap in firm premium across workers who work at the same firm. In contrast, we find only a very small, 1 percentage point difference in gender-specific firm premium at firms which do not rely on flexible wage components. These findings are in line with Lemieux et al. (2009) who argue that flexible wages are the main driver of within-firm wage differences.

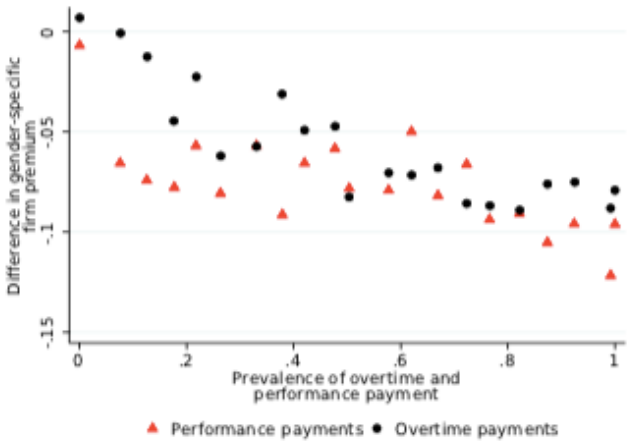
Table 5: Decomposition of the gender gap – by the share of flexible wages

	Gender gap lnw	Overall firm premium Ψ_{fg}	WI (bargaining) Ψ_{fg}	BW (sorting) ψ_f
Panel A. Flexible share ≥ 95%				
Male	0.248***	0.109***	0.057***	0.052***
	(0.004)	(0.002)	(0.001)	(0.002)
Observations	32,778	32,778	32,778	32,778
R-squared	0.114	0.058	0.873	0.016
Firm effects	x	x	✓	x
Number of units	1	1	6947	1
Panel B. Flexible share: 5-95%				
Male	0.209***	0.083***	0.055***	0.028***
	(0.004)	(0.002)	(0.001)	(0.002)
Observations	62,937	58,448	58,448	58,448
R-squared	0.052	0.025	0.891	0.003
Firm effects	x	x	✓	x
Number of units	1	1	12124	1
Panel C. Flexible share ≤ 5%				
Male	0.123***	0.010*	0.008***	0.002
	(0.006)	(0.004)	(0.002)	(0.004)
Observations	22,150	22,150	22,150	22,150
R-squared	0.017	0.000	0.875	0.000
Firm effects	x	x	✓	x
Number of units	1	1	8721	1

Notes: See Section 3 for the methodology of the decomposition. Flexible shares are defined as the share of workers receiving either overtime or performance payments at a given firm in a given year. Standard errors are in parentheses. *** p<0.001, ** p<0.01, * p<0.05.

In Figure 3 we investigate the effect of overtime payments and performance payments on gender-specific firm premium separately. For this purpose, we order firm-gender-occupation cells by the prevalence of overtime payments and the prevalence of performance payments. Then, we plot the difference in gender-specific firm premium by 0.05 unit bins. The figure suggests that the difference in gender-specific firm premium is linearly decreasing both in the share of overtime and performance payments. We do not find any difference in gender-specific firm premium across workers who do not receive overtime or flexible wages if we apply the normalization explained at the beginning of this section. Furthermore, if the share of overtime or performance payments increases, the firm premium of women becomes lower and lower compared to the firm premium of men. As in Table 5, we find that the gender gap in firm-specific wage premium is approximately 10 percentage points across workers where everybody receives either overtime payments or performance payments. This magnitude is significant in economic terms since the total gender wage gap is more than 22 percent in our sample.

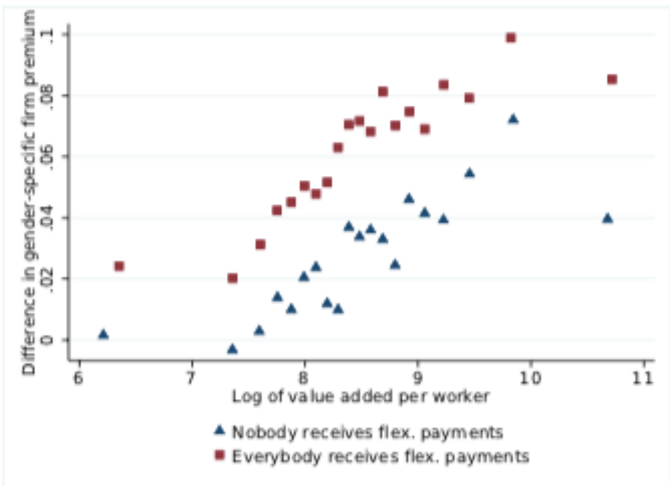
Figure 3: Wage premium and the prevalence of flexible wages



Notes: The horizontal axis shows the share of workers at the firm who receive performance or overtime payments and the vertical axis shows the gender difference in firm-gender premium.

It is important to note that the differences found in Table 5 and Figure 3 cannot be interpreted as a causal effect of flexible wage components on the gender-specific firm premium. The reason for this is that low-productivity firms which offer lower firm-specific wage premia are also less likely to offer flexible wage components. That is why we investigate the relationship between labor productivity, flexible wage components and the gender wage gap first in Figure 4. For this purpose, we ordered the firms by labor productivity and made twenty equally sized bins. Then, we plotted the average gender gap in firm-gender-specific wage premium for firms where every worker receives flexible wages and for firms where no worker receives flexible wage components.

Figure 4: Flexible wage payments and the gap in gender-specific wage premium by firm productivity



Notes: The figure shows the difference in gender-specific wage premium among workers who do and do not receive flexible wage components

The results confirm the previous ones and show that the gender gap in firm premium is increasing with labor productivity. This relationship holds for firms with and without flexible wage components alike. Most importantly, the results show that the firm-specific wage premium is always larger at flexible-wage firms than at firms of the same productivity but without flexible wage components. This relationship holds even among low-productivity firms were only lower rents are generated and therefore there is less possibility for an emerging gender gap.

To systematically investigate which factors contribute to the gender gap in firm-specific wage premia at flexible-wage firms, we extend the decomposition method introduced in Section 3. In practice this means that we extend Equations 4, 5 and 6 with control variables:

$$\Psi_{jg} = \beta_G^{OA} G_g + \beta_X^{OA} X_{jt} + \beta_{GX}^{OA} G_g X_{jt} + \theta_{s(j)} + \epsilon_{gjt} \qquad (9)$$

$$\Psi_{jg} = \beta_G^{WI} G_g + \beta_X^{WI} X_{jt} + \beta_{GX}^{WI} G_g X_{jt} + \tilde{\psi}_j + \varepsilon_{jgt}^{G} \qquad (10)$$

$$\tilde{\psi}_j = \beta_G^{BW} G_g + \beta_X^{BW} X_{jt} + \beta_{GX}^{BW} G_g X_{jt} + \theta_{s(j)} + \varepsilon_{jt} \qquad (11)$$

where G_g corresponds to a gender dummy, and X_{jt} to observable firm characteristics such as size, productivity or the share of workers with overtime and performance pay components.

We demean the control variables, thus the β_G^{OA}, β_G^{WI} and β_G^{BW} parameters correspond to the overall, within-firm (*bargaining*) and between-firm (*sorting)* gender differences in the firm-specific wage premium in an average firm after controlling for differences in worker composition and firm quality. The β_X^{OA}, β_X^{WI} and β_X^{BW} parameters will capture the relation between wage levels and productivity, firm size and the prevalence of flexible wage components. These control variables also have a within-firm (intertemporal) margin. For instance, the same firms could pay more in years when they become more productive.

Finally, the β_{GX}^{OA}, β_{GX}^{WI} and β_{GX}^{BW} parameters will capture whether a higher share *GXGX GX* of flexible components or a higher productivity comes with higher gender wage gaps when controlling for the other factors. Again, the gender differences could be generated within firms (e.g. if productivity rents are shared differently within the same firm) or between firms (e.g. if women sort into firms where the returns to productivity or flexible wages is higher), as noted by Boza (2022).

Table 6: Decomposition of the gender gap – with firm controls

VARIABLES	(1) OA $\ln w_{igt}$	(2) OA Ψ_{fgt}	(3) WI Ψ_{fgt}	(4) BW $\bar{\psi}_{ft}$
Female	-0.174***	-0.042***	-0.033***	-0.008***
	(0.002)	(0.001)	(0.001)	(0.001)
Log value added per worker	0.280***	0.149***	0.041***	0.108***
	(0.001)	(0.001)	(0.001)	(0.001)
Log(size)	0.018***	0.017***	0.009***	0.008***
	(0.001)	(0.000)	(0.001)	(0.000)
Share of performance payments	0.204***	0.138***	0.019***	0.119***
	(0.003)	(0.002)	(0.001)	(0.002)
Share of overtime payments	0.041***	0.096***	0.031***	0.065***
	(0.004)	(0.003)	(0.002)	(0.002)
Female#(labor prod.)	-0.008***	-0.009***	-0.015***	0.006***
	(0.002)	(0.001)	(0.001)	(0.001)
Female#log(size)	-0.018***	-0.005***	0.003***	-0.008***
	(0.001)	(0.001)	(0.000)	(0.001)
Female#(performance payments)	-0.034***	-0.051***	-0.016***	-0.035***
	(0.005)	(0.003)	(0.002)	(0.003)
Female#(overtime payments)	-0.090***	-0.038***	-0.011***	-0.027***
	(0.006)	(0.004)	(0.002)	(0.003)
Constant	6.760***	0.274***	0.324***	-0.050***
	(0.001)	(0.001)	(0.001)	(0.001)
Observations	97,690	97,690	97,690	97,690
R-squared	0.636	0.555	0.907	0.455
Firm effects	x	x	✓	x
Number of units	14	14	13737	14

Notes: The regressions are estimated based on Equations 9, 10 and 11, and include 1-digit sector and year fixed effects as well besides the presented variables. Standard errors are in parentheses. *** p<0.001, ** p<0.01, * p<0.05.

The results are presented in Table 6. The first column shows that, conditional on firm characteristics, the gender wage gap at an average firm is 17.4 percent. The second column shows that 4.2 percentage points of this difference can be attributed to the gender difference in the firm-specific wage premium. An important change compared to the raw differences is that, conditional on firm characteristics, the between-firm difference in the gender-specific wage premium is only 0.8 percent. In other words, we do not see evidence that women are more likely to work at firms with a substantially lower firm-specific wage premium. Previous differences in this sorting component are attributable instead to differences in the observable firm characteristics of Table 6.

Besides the average gender difference in the firm-specific wage premium, we also see important heterogeneity across firms. In line with previous findings in the literature, wage levels increase with the productivity and size of the firm. For example, men who work at a firm of a 1 percent higher productivity earn 0.28 percent more on average. Most importantly, men who work at firms where every worker receives performance payments earn 20.4 percentage points more than men who work at firms where nobody receives performance payments. In contrast to this, men working at firms where everybody receives overtime payments earn only 4.1 percent more than men who work at firms where nobody receives overtime payments.

We also see large gender differences in the wage effects of firm characteristics. For example, men earn a 0.18 percent higher salary at firms with 10 percent more employment, but we do not see that women earn more at larger firms, conditional on firm productivity and wage structure. Similarly, women gain less than men from working at firms where the share of performance payments is larger. What is more, women do not earn more at firms where overtime payments are more prevalent.

The most important results of our paper are presented in Column 2 of Table 6. It shows that the firm-specific wage premium of men increases by 0.138 percentage point if the share of workers with performance payments increases by 1 percentage point. This difference is almost exactly the same as the wage premium of a firm of 1 percent higher labor productivity. The wage premium of overtime payments has a similar magnitude. Men receive a 0.96 percentage point higher firm-specific wage premium at firms where the share of workers with overtime payments is higher with 10 percentage points.

The lower panel of Column 2 shows that women receive a much lower firm-specific wage premium than men if they work at firms where flexible wages are more prevalent. For example, if the share of workers with performance payments increases from 0 to 1, the firm-specific wage premium of women increases 5.1 percentage points less than the wage premium of men. This difference is significant in economic terms since it is more than a fifth of the total gender wage gap. We also find a significant difference in the returns to overtime payments. Women earn only a 5.8 percentage points larger firm-specific wage premium at firms where everybody receives overtime payments compared to a firm where no worker receives overtime payments, while for male workers this difference is 3.8 percentage points larger (9.6).

Finally, the comparison of Column (3) and (4) reveals that the main drivers of the gender gap in firm-specific wage premium are between-firm differences. Firms where men work pay 12% larger wages to everyone if they pay performance payments to everyone compared to firms offering no performance payments. This between-firm difference in wage premium is 3.5% smaller for firms where female workers tend to work. As opposed to this, the within-firm gender gap in wage premium is on average only 1.6 percentage points larger at firms where everybody receives performance payments[6].

[6] As an additional test of this hypothesis, we ran a simple regression on firm-level gender wage gaps as outcome variables. In such a simple formulation, we also found that the prevalence of performance pay and overtime, and the

As opposed to this, firm characteristics have a much lower effect on within-firm difference in the firm-specific wage premium. For instance, Column (3) reveals that the wage premium of men would increase by around 1.9% if the firm switched the share of workers with performance payments from 0 to 1. We see a gender gap in this margin as well because female wages would increase only by 0.3 percent if the firm would increase the share of performance payments from 0 to 1.

Analogously, the portion of the gender wage gap generated by the different reliance on overtime payments (-0.038) emerges mostly from between-firm differences in the intensity of overtime (-0.027), while within-firm differences play a much smaller role[7].

5. Discussion

It has been documented that gender differences in firm-specific wage premia are main drivers of the overall gender wage gap in many countries. This means that women earn less than men partially because women are less likely to work at firms with a high wage premium (*sorting effect*) and even if they can enter these firms, they tend to earn less than their male co-workers. In this paper we investigated the extent to which overtime and performance payments contribute to this undesirable phenomenon. We used a Hungarian administrative linked employer-employee dataset combined with a wage survey containing information on individual-level performance payments (including bonuses) and overtime payments.

We found that the firm-specific wage premium of women is 5.1 percentage points lower than the wage premium of men at firms where everybody receives performance payments compared to firms where nobody receives performance payments. This gender difference in firm-specific wage premium is 3.8 percentage points in the case of overtime payments. Furthermore, we showed that two-thirds of these differences come from between-firm differences, namely from the fact that women are less likely to work

productivity of the firm all contributed to larger within-firm gaps. (Table A2.) This was true even if we considered the composition-adjusted wage gaps that we had obtained by taking the difference in the AKM effects of the male and female parts of the same firm.

[7] In the case of productivity, within-firm rent-sharing differences – 1.5 percentage points lower returns to productivity for women than men in the same firm – are actually somewhat offset by women sorting into firms with higher rent-sharing elasticities, as noted by Boza (2022).

at firms which offer either overtime payments or performance payments. These differences are large in economic terms as the total gender wage gap at private sector firms is 22 percent in Hungary.

The results imply that policy interventions which regulate bonuses and overtime payments could decrease the gender pay gap. More specifically, a stricter regulation of overtime work (e.g higher taxes on overtime, or direct restrictions) would be an efficient for this purpose (Goldin, 2014). The regulation of paid overtime hours would be not sufficient without the restriction of unpaid working hours. The reason is that workers receiving bonuses work more unpaid overtime (Engellandt & Riphahn, 2011) and in some cases firms manipulate reported overtime hours for tax optimization purposes (Cahuc & Carcillo, 2014). Still, the introduction of such policies needs caution because they flexible wage components have incentive effects and increase worker productivity (Bloom & Van Reenen, 2011; Oyer & Schaefer, 2011).

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Appendices

Appendix A – Further figures and tables

Table A1: Decomposition of the gender wage gap - full sample

	OA lnw	OA Ψ_{fg}	WI Ψ_{fg}	BW $\tilde{\psi}_f$
2003-2016	0.146***	0.111***	0.026***	0.085***
	(0.001)	(0.001)	(0.000)	(0.001)
2003-2009	0.137***	0.114***	0.026***	0.087***
	(0.002)	(0.001)	(0.000)	(0.001)
2010-2016	0.156***	0.107***	0.025***	0.082***
	(0.002)	(0.001)	(0.001)	(0.001)
Difference	0.019***	-0.007***	-0.001	-0.006***
	(0.002)	(0.002)	(0.001)	(0.001)
Observations	475,538	475,538	475,538	475,538
Firm effects	x	x	✓	x
Number of units	1	1	73636	1

Notes: The table shows the contribution of firm pay premia to the gender wage gap using the modified method of Card et al. (2016), presented in Section 3 and Boza (2022).

Table A2: Regressions on wage gaps as outcomes

	(1)	(2)	(3)	(4)
		Wi sector		Wi sector
	D(lnw)	D(lnw)	$D(\Psi_{fgt})$	$D(\Psi_{fgt})$
Log value added per worker	0.017***	0.014***	0.016***	0.000
	(0.001)	(0.001)	(0.001)	(0.001)
Log(size)	-0.005***	-0.001**	-0.003***	0.010***
	(0.000)	(0.000)	(0.001)	(0.001)
Share of performance payments	0.029***	0.026***	0.008**	0.013***
	(0.001)	(0.001)	(0.002)	(0.002)
Share of overtime payments	0.016***	0.015***	0.051***	0.049***
	(0.002)	(0.002)	(0.003)	(0.003)
Observations	52,039	52,036	52,039	52,036
R-squared	0.046	0.073	0.016	0.125

Standard errors are in parentheses

*** p<0.001, ** p<0.01, * p<0.05

Notes: The parameters are coming from firm-level regressions in which the outcome variable is the firm-level gender wage gap, either estimated by the difference in log wages or by the difference of the AKM firm-gender-year effects of the firm. Besides the shown variables of interest, controls include year and sector dummies for specifications (2) and (4).

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Coordinator

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Consortium

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Centre for European Policy Studies (Belgium)
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