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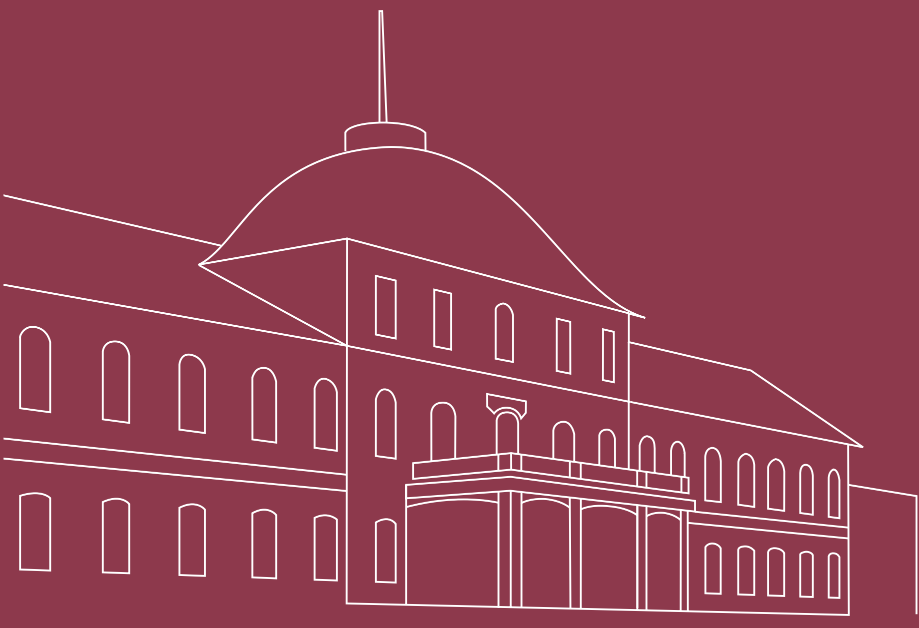

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# The Convergence of the Gender Pay Gap – An Alternative Estimation Approach –*

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**Abstract**

So far, little work has been done on directly estimating differences of wage gaps. Studies estimating pay differentials, generally compare them across different subsamples. This comparison does not allow to conduct any inference or, in the case of decompositions, to confront the respective decomposition components across subsamples. We propose an extension of an Oaxaca-Blinder type decomposition based on the omitted variable bias formula to directly estimate the change in pay gaps across subsamples. The method proposed can be made robust to the index-number problem of the standard Oaxaca-Blinder decomposition and to the indeterminacy problem of the intercept-shift approach. Using Italian micro data, we estimate the difference in the gender pay gap across time (2005 and 2014). By applying our proposed decomposition, we find that the convergence of the gender pay gap over time is only driven by the catching-up of women in terms of observable characteristics, while the impact of anti-discrimination legislation is found to be negligible.

**Keywords**: Pay Gap, Omitted Varibale Bias Formula, Oaxaca-Blinder Decomposition.

**JEL - Classification**: J7, J13, J310

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# 1 Introduction

Gender differentials in the labor market have obtained much attention from policy makers and researchers leading to the implementation of equal-pay legislation and the promotion of equal opportunities. Even though equal-pay legislation and equal opportunities have been promoted in Western industrialized countries for several decades, differences in pay between men and women persist (see for example Blau and Kahn, 1992, 2006; Goldin, 2014; Blau and Kahn, 2016). For example, in the European Union in 2014, women earned on average 16.7% less than men (Eurostat, 2017).

Typically, different Gender Pay Gaps (GPGs) are found across time. In particular, declining GPGs are observed with slower convergence in recent decades (see Blau and Kahn, 2006; England, 2006). The main reasons for the decline of the GPG over time are found to be the catching-up of women in terms of education and labour market experience (Goldin, 2006), technical development (Black and Spitz-Oener, 2010), changes in attitudes towards women in the labor market, less occupational segregation (Cotter, 2004; England, 2006; Mandel and Semyonov, 2014) and anti-discrimination laws (Fortin, 2015). Research has shown that the unexplained or coefficients effect of the GPG is reduced subsequently over time (e.g. Mandel and Semyonov, 2014). Differences in pay are revealed also across sectors and especially between the public and the private sector. The Public-Private Sector Wage Gap (PPWG) is found to differ significantly for men and women (Melly, 2005; Lucifora and Meurs, 2006; Arulampalam *et al.*, 2007). In fact, the difference in pay by gender is found to be smaller in the public compared to the private sector (see for example Melly, 2005; Arulampalam *et al.*, 2007). Regardless of gender, pay levels in the public sector are on average higher than in the private sector (Lucifora and Meurs, 2006). The public sector is generally the preferred sector of women due to its fairer recruitment, selection criteria and remuneration as well as better implementation of anti-discrimination laws (Gornick and Jacobs, 1998; Grimshaw, 2000).

However, studies examining changes in the wage gap over time and between groups do not directly estimate the difference of the GPG in year $t$ and year $t+1$ (or the wage gap by sector for men and women for instance), but rather compare the results of the pay gaps in the corresponding subsamples ex post (e.g. Christofides and Michael, 2013; Mandel and Semyonov, 2014). Studies estimating the difference of the pay gaps in different subsamples, often do not even provide standard errors for the decomposition. Hence, it is not possible to conduct statistical inference (Mandel and Semyonov, 2014; Bar *et al.*, 2015). Indeed, this does not allow to draw conclusions on which of the two wage gaps is more statistically significant, i.e. whether the difference between the two pay gaps under investigation is statistically significantly different from zero. Additionally, the conclusion about drivers of the change of pay gaps between groups may be different, when estimated directly compared to analyzing results estimated in different subsamples. The reason is that it is not possible to draw direct inference of the difference of the respective components in the latter case. Moreover, the standard method, i.e. ex-post comparison of the decomposition results, does not allow to catch time- (or sector-) and gender-specific effects that may exist simultaneously, i.e. interactions across gender and time or sector and gender (in the case of the GPG over time and the PPWG by gender, respectively). We slightly extend the method proposed by Gelbach (2016) that is based on the Omitted Variable Bias (OVB) formula to estimate directly the difference between two wage gaps. We are then able to draw inference on the changes of the pay gap by groups across subsamples and to compare the various contributors directly, i.e. we can test whether there has been a significant change of the explained or unexplained part of the gap. Moreover, we can draw conclusions on the relevance of interaction effects across subsamples and groups. The standard method in applied labor economics, when it comes to pay gaps between groups is the Oaxaca (1973) and Blinder (1973) decomposition method (Fortin *et al.*, 2011). The approach, however, suffers from non-invariance with respect to categorical variables and the index-number problem. The intercept-shift approach attempts to solve the latter but suffers, in particular, from the indeterminacy problem (Lee, 2015). We extend our proposed method based on the OVB formula and show that it can be made robust to the above mentioned problems.

We apply our model to two cases. First, we examine the evolution of the GPG over ten years, from 2005 to 2014 in Italy. Second, we analyze the PPWG between men and women in 2014 in Italy. We analyze each case with the standard Oaxaca-Blinder decomposition method and then repeat the examination with our proposed extension of the Gelbach decomposition. We expect to find a statistically significant change in differences in observable characteristics (such as educational attainment, labor market presence as well as job-, industry- or occupational-specific characteristics) by gender over time as well as a statistically significant change in differences in coefficients to these characteristics between men and women over time. In fact, the latter may indicate the effectiveness of anti-discrimination policies. For the second empirical application, the PPWG by gender, we expect, to find in line with the literature larger pay gaps for women between the public and the private sector than for men. Additionally, we expect to find a larger effect of the unexplained component in the PPWG for women; while differences in endowments may be the main driver of the pay differential for men, they may not explain equally the difference in the PPWG for women.

For the first case, the findings of the study reveal interesting differences in results when applying our proposed estimation methodology compared to the 'standard' approach.[1] Changes in gender differences of observable characteristics are found to be the only statistically significant driving force of the convergence of the GPG in the last decade in Italy. On the contrary, by comparing the different components of the GPGs following Oaxaca (1973) and Blinder (1973), differences in returns to observable characteristics, often referred to as the unexplained part of the GPG, seem to play a role in closing the gap over the last decade in Italy. In the second case, we can confirm the conclusions drawn from the estimation in the respective subsamples;

[1] i.e. the Oaxaca-Blinder decomposition and ex-post comparison of the decomposition results.

the higher PPWG for women than for men is due to both differences in the explained and unexplained component.

The paper is organized as follows. Section 2 presents the standard Oaxaca-Blinder decomposition. In Section 3, we outline the method by Gelbach (2016) as well as our proposed modification. Similarly, we discuss problems of the standard approach and show the robustness of our method to these problems. Next, in Section 4, we empirically apply the method proposed to the GPG over time as well as to the PPWG by gender and discuss the results obtained. Section 5 concludes.

## 2 Standard Estimation Strategy

The standard methodology to decompose pay differentials between two groups is the Oaxaca (1973) and Blinder (1973) decomposition. The methodology estimates Mincer-type wage regressions separately for a specific group (e.g. men or women, the public or the private sector) and then decomposes the wage differential in different components. We use the three-fold Oaxaca-Blinder approach and thus decompose the pay gap in three components; endowments, coefficients and interactions:[2]

$$\begin{aligned}\overline{ln(w_0)} - \overline{ln(w_1)} &= \hat{\alpha}_0 + \bar{X}_0\hat{\beta}_0 - \hat{\alpha}_1 - \bar{X}_1\hat{\beta}_1 \\ &= (\bar{X}_0 - \bar{X}_1)\hat{\beta}_1 + (\hat{\alpha}_0 - \hat{\alpha}_1) + \bar{X}_1(\hat{\beta}_0 - \hat{\beta}_1) \\ &+ (\bar{X}_0 - \bar{X}_1)(\hat{\beta}_0 - \hat{\beta}_1)\end{aligned}$$

where $\overline{ln(w_G)}$ is the logarithmic hourly wage of group $G$ evaluated at the mean, $\hat{\alpha}_G$ is the intercept of group $G$ and $\bar{X}'_G$ and $\hat{\beta}_G$ are $K \times 1$ vectors of average characteristics and estimated coefficients for $G \in \{0, 1\}$. The first term is the effect due to differences in observable characteristics. As different observed characteristics are expected to have different effects on earnings, the difference in observable characteristics is also referred to as the explained component, the quantity or endowments effect of the Oaxaca-Blinder decomposition. The second term is due to differences in the starting point, i.e. differences in the intercept. The third term is the effect due to differences in returns on the same set of observable characteristics. This component is generally referred to as the unexplained part, price or coefficients effect of the gap. Differences in the intercept are attributed to the coefficients component. In the case of the GPG, if the

[2]An alternative to the three-fold decomposition outlined here is the standard two-fold decomposition that decomposes the wage differential in an explained and an unexplained part;

$$\begin{aligned}\overline{ln(w_0)} - \overline{ln(w_1)} &= \hat{\alpha}_0 + \bar{X}_0\hat{\beta}_0 - \hat{\alpha}_1 - \bar{X}_1\hat{\beta}_1 \\ &= (\bar{X}_0 - \bar{X}_1)\hat{\beta}_0 + (\hat{\alpha}_0 - \hat{\alpha}_1) + \bar{X}_1(\hat{\beta}_0 - \hat{\beta}_1)\end{aligned}$$

We focus here on the three-fold decomposition, as we argue that interaction effects may be important when considering differences across pay gaps.

differential is mainly due to the price effect, this may indicate the presence of gender discrimination.[3] The last term is the so-called interaction term. The intuition behind is that differences in endowments and coefficients may exist simultaneously between groups (Jann, 2008).

# 3 Econometric Model

We propose a slight modification of the decomposition method by Gelbach (2016). The Gelbach-approach decomposes cross-specification differences in Ordinary Least Squares (OLS) estimates of the group-dummy coefficient from the wage model in a path-independent way yielding a Oaxaca-Blinder type decomposition. By using the OVB formula, the decomposition is consistently estimated conditional on all covariates used in the regression. This method, similar to the standard estimation approach outlined in Section 2, decomposes the sample mean difference in wages between different groups in an explained and an unexplained part (see Gelbach, 2016, for details).

## 3.1 Extension of Gelbach (2016)

The model outlined in the following allows not only to obtain information on whether the pay gap has decreased in a statistically significant way on aggregate but also to testify what are the main contributors to the change (if any) of the differential. Consider the case, when we estimate the wage equation separately by $G$ (group) and $Y$ (data wave or a group different from $G$, i.e. $Y \neq G$) for individual $i$, with $i = 1, 2, \ldots N$:

$$ln(w_{iGY}) = \alpha_{GY} + X_{iGY}\beta_{GY} + \epsilon_{iGY} \tag{1}$$

with $G \in \{0, 1\}$, $Y \in \{A, B\}$; and where $ln(w_{iGY})$ is individual $i$'s logarithmic wage of $G$ in $Y$, $\alpha_{GY}$ is a constant, $X_{iGY}$ is a $1 \times K$ vector of exogenous regressors, $\beta_{GY}$ is the corresponding $K \times 1$ vector of coefficients and $\epsilon_{iGY}$ is the error term.[4] When we evaluate the estimation at the mean given the OLS property that OLS estimates must go through the mean of the data, equation (1) becomes:

$$\overline{ln(w_{GY})} = \hat{\alpha}_{GY} + \bar{x}_{GY}\hat{\beta}_{GY} \tag{2}$$

[3]However, as pointed out by Blau and Kahn (2006), the unexplained portion of the GPG may include effects of unobserved characteristics such as individual productivity, motivation or educational quality.

[4]In the first empirical application in Section 4, we set the index $G$ equal to gender and the index $Y$ equal to different years or waves of the data set. Consequently, in case 1 of the empirical implementation, we have for $G \in \{0, 1\}$; $0$ = male and $1$ = female and for $Y \in \{A, B\}$; $A$ = starting period or 2005 and $B$ = ending period or 2014. In the second empirical example shown in Section 4, group $G$ represents different sectors and $Y$ men or women. Thus, in case 2 of the empirical part, we have for $G \in \{0, 1\}$; $0$ = public-sector employment and $1$ = private sector employment and for $Y \in \{A, B\}$; $A$ = female and $B$ = male.

where $\hat{\alpha}_{GY}$ is the constant, $\bar{x}_{GY}$ is the $1 \times K$ row vector of sample means of observable characteristics in $X$:

$$\bar{x}_{GY} = \left[\bar{x}_{k1}, \bar{x}_{k2}, \ldots, \bar{x}_{K}\right]$$

and $\hat{\beta}_{GY}$ is the corresponding $K \times 1$ vector of parameter estimates. Four different pairs of $(G, Y)$ and thus four regressions of equation (2) are possible; $(0, A)$, $(0, B)$, $(1, A)$, $(1, B)$. The corresponding regressions between $G$ and $Y$ are conducted by assuming the same set of regressors for all four cases.

Now, consider estimating the joint model. The first group index $G$ is added to the regression as a dummy variable $G_i$ among the controls on the right-hand side. Analogously, the second group index $Y$ is transformed in a dummy variable $Y_i$ controlling for group $Y$ membership. The indicator variable takes value one, if the observation corresponds to $A$ and takes value zero, if we observe $B$.[5] As in Gelbach (2016), we distinguish between two sets of regressors, $X_{i1}$ and $X_{i2}$, where the set of regressors $X_{i1}$, with dimension $1 \times 4$, is the base specification containing only (for each observation $i$) a constant, an interaction term between the group dummies, $G_i Y_i$, as well as the dummies, $G_i$ and $Y_i$, separately. The interaction of the dummies for group membership $G_i$ and $Y_i$ are contained in $G_i Y_i$. The base model is therefore defined as follows:

$$\begin{aligned} ln(w_{iGY}) &= X_{i1}\alpha^{base} + \epsilon_{iGY}^{base} \\ ln(w_{iGY}) &= \alpha_0^{base} + G_i Y_i \alpha_1^{base} + G_i \alpha_2^{base} + Y_i \alpha_3^{base} + \epsilon_{iGY}^{base} \end{aligned} \tag{3}$$

where $\alpha_0^{base}$ is the constant and $\alpha_1^{base}, \alpha_2^{base}, \alpha_3^{base}$ are the corresponding coefficients contained in the $4 \times 1$ column vector $\alpha^{base}$, $\epsilon_{iGY}^{base}$ is the corresponding error term. The second set of regressors, $X_{i2}$, has dimension $1 \times 4K$ and contains the $1 \times K$ vector of explanatory variables $X_i$ as well as the interactions of $X_i$ with $G_i$, $Y_i$ and $G_i Y_i$, respectively. The set of regressors $X_{i2}$ will be considered later as omitted variables in order to obtain a decomposition of the change of the wage gap between $G_i$ across $Y_i$. The full model is then defined as:

$$\begin{aligned} ln(w_{iGY}) &= X_{i1}\alpha^{full} + X_{i2}\beta + \epsilon_{iGY}^{full} \\ ln(w_{iGY}) &= \alpha_0^{full} + G_i Y_i \alpha_1^{full} + G_i \alpha_2^{full} + Y_i \alpha_3^{full} + X_i\beta_1 + G_i X_i \beta_2 + Y_i X_i \beta_3 + G_i Y_i X_i \beta_4 + \epsilon_{iGY}^{full} \end{aligned} \tag{4}$$

[5] We thus have the index $G \in \{0, 1\}$ and the dummy variable $G_i$, with

$$G_i = \begin{cases} 1 & \text{if the index of person } i \text{ is } G = 1 \\ 0 & \text{if the index of person } i \text{ is } G = 0 \end{cases}$$

For the second group, we have the index $Y \in \{A, B\}$ and the dummy variable $Y_i$, with

$$Y_i = \begin{cases} 1 & \text{if the index of person } i \text{ is } Y = A \\ 0 & \text{if the index of person } i \text{ is } Y = B \end{cases}$$

where $\alpha^{full}$ and $\beta$ are the $4 \times 1$ and $4K \times 1$ vectors of coefficients from $X_{i1}$ and $X_{i2}$, respectively. The error term is represented by $\epsilon_{iGY}^{full}$.

We can recast the parameters of the full model evaluated at the mean from the pair-wise regressions of (2):

1. When (the indices) G=1 and Y=A, we get:
   - $\hat{\alpha}_{1A} = \hat{\alpha}_0^{full} + \hat{\alpha}_1^{full} + \hat{\alpha}_2^{full} + \hat{\alpha}_3^{full}$
   - $\hat{\beta}_{1A} = \hat{\beta}_1 + \hat{\beta}_2 + \hat{\beta}_3 + \hat{\beta}_4$
2. When (the indices) G=0 and Y=A, we get:
   - $\hat{\alpha}_{0A} = \hat{\alpha}_0^{full} + \hat{\alpha}_3^{full}$
   - $\hat{\beta}_{0A} = \hat{\beta}_1 + \hat{\beta}_3$
3. When (the indices) G=1 and Y=B, we get:
   - $\hat{\alpha}_{1B} = \hat{\alpha}_0^{full} + \hat{\alpha}_2^{full}$
   - $\hat{\beta}_{1B} = \hat{\beta}_1 + \hat{\beta}_2$
4. When (the indices) G=0 and Y=B, we get:
   - $\hat{\alpha}_{0B} = \hat{\alpha}_0^{full}$
   - $\hat{\beta}_{0B} = \hat{\beta}_1$

Re-arranging the terms slightly, gives us:

$$\begin{aligned}
\hat{\alpha}_0^{full} &= \hat{\alpha}_{0B} \\
\hat{\alpha}_2^{full} &= \hat{\alpha}_{1B} - \hat{\alpha}_{0B} \\
\hat{\alpha}_3^{full} &= \hat{\alpha}_{0A} - \hat{\alpha}_{0B} \\
\hat{\alpha}_1^{full} &= \hat{\alpha}_{1A} - \hat{\alpha}_{0B} - \hat{\alpha}_{1B} + \hat{\alpha}_{0B} - \hat{\alpha}_{0A} + \hat{\alpha}_{0B} \\
&= (\hat{\alpha}_{0B} - \hat{\alpha}_{1B}) - (\hat{\alpha}_{0A} - \hat{\alpha}_{1A})
\end{aligned}$$

$$\begin{aligned}
\hat{\beta}_1 &= \hat{\beta}_{0B} \\
\hat{\beta}_2 &= \hat{\beta}_{1B} - \hat{\beta}_{0B} \\
\hat{\beta}_3 &= \hat{\beta}_{0A} - \hat{\beta}_{0B} \\
\hat{\beta}_4 &= \hat{\beta}_{0B} - \hat{\beta}_{1B} - \hat{\beta}_{0A} + \hat{\beta}_{1A} \\
&= (\hat{\beta}_{0B} - \hat{\beta}_{1B}) - (\hat{\beta}_{0A} - \hat{\beta}_{1A})
\end{aligned}$$

By estimating the base model and considering the set of regressors $X_{i2}$ as omitted variables, we obtain the following specification:

$$\hat{\alpha}^{base} = \hat{\alpha}^{full} + (X'_{i1}X_{i1})^{-1}X'_{i1}X_{i2}\hat{\beta}^{full} \tag{5}$$

where

- $(X'_{i1}X_{i1})^{-1}X'_{i1}X_{i2}\hat{\beta}^{full}$ is the OVB
- The parameter estimates from the base model (3) evaluated at the mean are:

  $$\hat{\alpha}^{base} = \left[\hat{\alpha}_0^{base}, \hat{\alpha}_1^{base}, \hat{\alpha}_2^{base}, \hat{\alpha}_3^{base}\right]^T$$

  being a $4 \times 1$ column vector.
- $\hat{\alpha}^{full}$ is the $4 \times 1$ column vector containing the coefficient estimates of $X_{i1}$ from the full model (4) evaluated at the mean.
- $(X'_{i1}X_{i1})^{-1}X'_{i1}X_{i2}$ is the linear projection of $X_{i2}$ on $X_{i1}$, with dimension $4 \times 4K$.
- $$\hat{\beta}^{full} = \left[\hat{\beta}_1, \hat{\beta}_2, \hat{\beta}_3, \hat{\beta}_4\right]^T$$

  is a $4K \times 1$ column vector of coefficients from the full model (4) evaluated at the mean. The model specification in equation (5) can be decomposed as follows:

  $$\hat{\alpha}^{base} = \hat{\alpha}^{full} + \hat{\delta}^1 + \hat{\delta}^2 + \hat{\delta}^3 + \hat{\delta}^4 \tag{6}$$

  with $\hat{\delta} \equiv \hat{\alpha}^{base} - \hat{\alpha}^{full} = (X'_{i1}X_{i1})^{-1} X'_{i1}X_{i2}\hat{\beta}^{full}$, where
  - $\hat{\delta}^q = \hat{\Gamma}^q\hat{\beta}_q^{full}$, with $\hat{\Gamma}^q = (X'_{i1}X_{i1})^{-1} X'_{i1}X_{i2q}$ of dimension $k_{X_{i1}} \times k_q$ and $X_{i2q}$ being the $q$th column of $X_{i2}$, for $q = 1, .., Q$. The column vector $\hat{\beta}_q^{full}$ has dimension $k_q \times 1$, thus $\hat{\delta}_q$ is a $k_{X_{i1}} \times 1$ column vector;
  - $k_{X_{i1}}$ is equal to the number of regressors from $X_{i1}$, i.e. 4 in our case ($X_{i1}$ has dimension $1 \times 4$);
  - $k_q$ is equal to the number of regressors in the $q$th column of $X_{i2}$.

## 3.2 Decomposition

Recall that we are interested in the estimation and decomposition of the change in the pay gap between group $G$ across group $Y$, i.e.[6]

$$\Delta^{\mathrm{B}} - \Delta^{\mathrm{A}} = \left(\overline{ln(w_{0B})} - \overline{ln(w_{1B})}\right) - \left(\overline{ln(w_{0A})} - \overline{ln(w_{1A})}\right)$$

[6] For example, the change of the GPG across two years.

with $\Delta^{\mathrm{B}}$ being the pay gap by group $G$ given that $Y = B$ and $\Delta^{\mathrm{A}}$ being the wage gap between $G$ given that $Y = A$. From equation (2), we know that:

$$\Delta^{\mathrm{B}} = \left(\overline{ln(w_{0B})} - \overline{ln(w_{1B})}\right) = -\hat{\alpha}_2^{base}$$

$$\Delta^{\mathrm{A}} = \left(\overline{ln(w_{0A})} - \overline{ln(w_{1A})}\right) = -\hat{\alpha}_1^{base} - \hat{\alpha}_2^{base}$$

and hence $\hat{\alpha}_1$ represents the difference of the two wage gaps:

$$\Delta^{\mathrm{B}} - \Delta^{\mathrm{A}} = \hat{\alpha}_1^{base}$$

Given the definition of $\hat{\alpha}^{base}$, we are interested in the second row of $\hat{\alpha}^{base}$, i.e. of equation (5), or $\hat{\alpha}_1^{base}$ in order to obtain the change of the wage gaps, $\Delta^{\mathrm{B}} - \Delta^{\mathrm{A}}$. Starting from equation (5), we calculate the second row of the $4 \times 4K$ matrix $(X'_{i1}X_{i1})^{-1}X'_{i1}X_{i2}$ considering average observable characteristics:

$$\kappa = \left[(\bar{x}_{0B} - \bar{x}_{1B}) - (\bar{x}_{0A} - \bar{x}_{1A}), (\bar{x}_{1A} - \bar{x}_{1B}), (\bar{x}_{1A} - \bar{x}_{0A}), \bar{x}_{1A}\right]$$

with dimension $1 \times 4K$. The second row of equation (5) or the difference of the respective wage gap evaluated at the mean is thus:

$$\hat{\alpha}_1^{base} = \hat{\alpha}_1^{full} + \kappa\hat{\beta}^{full} \tag{7}$$

and can be re-written as:

$$\begin{aligned}
\hat{\alpha_1}^{base} = &\underbrace{(\hat{\alpha}_{0B} - \hat{\alpha}_{1B}) - (\hat{\alpha}_{0A} - \hat{\alpha}_{1A})}_{\hat{\alpha}_1^{full}} + [(\bar{x}_{0B} - \bar{x}_{1B}) - (\bar{x}_{0A} - \bar{x}_{1A})]\underbrace{\hat{\beta}_{0B}}_{\hat{\beta}_1} \\
&+ (\bar{x}_{1A} - \bar{x}_{1B})\underbrace{(\hat{\beta}_{1B} - \hat{\beta}_{0B})}_{\hat{\beta}_2} \\
&+ (\bar{x}_{1A} - \bar{x}_{0A})\underbrace{(\hat{\beta}_{0A} - \hat{\beta}_{0B})}_{\hat{\beta}_3} \\
&+ \bar{x}_{1A}\underbrace{[(\hat{\beta}_{0B} - \hat{\beta}_{1B}) - (\hat{\beta}_{0A} - \hat{\beta}_{1A})]}_{\hat{\beta}_4} \\
&= \Delta^B - \Delta^A
\end{aligned} \tag{8}$$

where $\hat{\alpha_1}^{base}$ and $\hat{\alpha_1}^{full}$ are scalars and $\bar{x}'_{GY}$, $\hat{\beta}_1$, $\hat{\beta}_2$, $\hat{\beta}_3$, $\hat{\beta}_4$ are $K \times 1$ column vectors, respectively. The above expression can be re-written as a 'double' (two-fold) Oaxaca-Blinder decomposition:

$$\begin{aligned}\hat{\alpha_1}^{base} &= (\hat{\alpha}_{0B} - \hat{\alpha}_{1B}) + (\bar{x}_{0B} - \bar{x}_{1B})\hat{\beta}_{0B} + \bar{x}_{1B}(\hat{\beta}_{0B} - \hat{\beta}_{1B}) \\ &- [(\hat{\alpha}_{0A} - \hat{\alpha}_{1A}) + (\bar{x}_{0A} - \bar{x}_{1A})\hat{\beta}_{0A} + \bar{x}_{1A}(\hat{\beta}_{0A} - \hat{\beta}_{1A})]\end{aligned}$$

Decomposing the change in the wage gap between group $G$ across group $Y$ in the following way allows to better understand the elements that contribute to the earnings differences across $G$ and $Y$: $\Delta^{\mathrm{B}} - \Delta^{\mathrm{A}} = E + U + I1 + I2$, with

$$E = [(\bar{x}_{0B} - \bar{x}_{1B}) - (\bar{x}_{0A} - \bar{x}_{1A})]\hat{\beta}_{0B} \tag{9}$$

Here, the same prices, namely the ones of the respective base category, $\hat{\beta}_{0B}$, are assumed. Thus, $E$ measures the amount of the change of the gap attributable to differences in observed characteristics. It is the component referred to as differences in quantities, i.e. the explained part. The unexplained component becomes the following:

$$\begin{aligned}U &= \hat{\alpha}_1^{full} + \underbrace{\bar{x}_{1A}[(\hat{\beta}_{0B} - \hat{\beta}_{1B}) - (\hat{\beta}_{0A} - \hat{\beta}_{1A})]}_{u} \\ &= \hat{\alpha}_1^{full} + u \qquad (10)\end{aligned}$$

$U$ measures the change of differences in the intercepts, $\hat{\alpha}_1^{full}$, as well as the change over $Y$ of the differences in coefficients by $G$. Characteristics are hold fix at $\bar{x}_{1A}$. Additionally, we observe now two interaction terms, $I1$ and $I2$, accounting for the fact that differences in characteristics and parameters exist simultaneously between the four groups. The interaction effects are the following:

$$I1 = (\bar{x}_{1A} - \bar{x}_{1B})(\hat{\beta}_{1B} - \hat{\beta}_{0B}) \tag{11}$$

and

$$I2 = (\bar{x}_{1A} - \bar{x}_{0A})(\hat{\beta}_{0A} - \hat{\beta}_{0B}) \tag{12}$$

$I1$ accounts for differences in prices by $G$ given changes in the set of endowments across $Y$.[7] $I2$ catches changes in coefficients over $Y$ given that endowments between $G$ are different.[8]

[7] In the case of the GPG over time, $I1$ catches year-specific effects in endowments given gender-related differences in prices in the ending period. That is assuming that in the ending period differences in prices between men and women persist (compared to the starting period), it accounts for changing endowments of women over time.

[8] In the first case of the empirical application, $I2$ assumes different endowments between women and men in the starting period and asks how coefficients change over time given gender differences in quantities.

Despite using the decomposition approach based on the OVB formula, we can compare differences in pay gaps by estimating a system of Seemingly Unrelated Equations (SURE). Using the SURE method allows errors to be correlated across equations and is more efficient. However, we prefer the more intuitive or more familiar interpretation of the method outlined above. Furthermore, the model based on the OVB formula catches otherwise unobserved interaction effects.

### 3.3 Robustness of the Method Proposed and Problems of the Standard Approach

The Oaxaca-Blinder decomposition suffers from various problems. In particular, the method is not unique and its components may be unstable when different controls are added to the Mincer-type wage equation. As the Oaxaca-Blinder decomposition is not unique, the choice of the non-discriminatory wage structure matters and the results may change according to the reference category chosen (Reimers, 1983; Cotton, 1988; Neumark, 1988; Oaxaca and Ransom, 1994; Fortin, 2008). Several solutions have been proposed in the literature to solve the so-called index-number problem. Suggestions in the literature consist in esimating a pooled wage structure (Neumark, 1988; Oaxaca and Ransom, 1994) or assigning different weights to the two groups (Reimers, 1983; Cotton, 1988). The intercept-shift approach including the group indicator and parameter restrictions, re-writes the decomposition in terms of advantages of men and disadvantages of women (Fortin, 2008). Thereby, the decompostion does no longer depend on the choice of the non-discriminatory wage structure. In the empirical application in Section 3.2, we take men and the ending period as base category or non-discriminatory wage structure.[9] Indeed, the standard case of the Oaxaca-Blinder decomposition assumes positive discrimination against women, i.e. it takes men as the non-discriminatory wage structure. For a recent application, see for example Mandel and Semyonov (2014). We can easily change the reference category by imposing different weights across groups (following Reimers, 1983; Cotton, 1988) and show in Appendix A that the standard case of the GPG can be decomposed in the sense of the intercept-shift approach as proposed by Fortin (2008) based on the OVB formula. In the case of a detailed decomposition, the standard Oaxaca-Blinder decomposition varies with the choice of the left-out category of categorical variables included in the estimation. We show the invariance with respect to categorical variables of the decomposition aproach based on the OVB formula in Appendix B. The coefficients of the categorical variables are transformed making them invariant to the choice of the (omitted) base category (Gardeazabal and Ugidos, 2004; Fortin, 2008). Moreover, in Appendix C, we show that the decomposition based on the intercept-shift approach holds also for our proposed decomposition of pay gaps between groups $G$ and $Y$. In Appendix D, we show that the critique of Lee (2015) stating that the intercept-shift approach relies on second moments, while first moments should be considered, does not apply to

[9] In the second empirical application, men in the public sector are the non-discriminatory wage structure.

our proposed decomposition approach with gender dummies along with parameter restrictions.[10] We derive the results in the appendices based on the GPG. However, the derived results are not only valid for the case of the GPG but can be applied to a variety of decomposition problems.

# 4 Empirical Implementation

In this Section, we consider the change of the GPG over time (case 1) as well as the PPWG between men and women (case 2). By applying our proposed approach, we are able to draw inference on the diverse contributors to the GPG over time.[11] The results from the standard model are also shown for the sake of comparison.

## 4.1 Data and Descriptive Statistics

We use the 2014 and 2005 cross-sectional files of the survey ISFOL PLUS[12] from the Italian Institute for the Development of Vocational Training for Workers (ISFOL). The data was collected jointly with the Italian Ministry of Labor and Social Policy. Special characteristics of the survey are that it provides broad information on the intervieews' working profiles and motivation to work as well as on the demographic and family background of the participants. Data collection is conducted by Computer Assisted Telephone Interviewing (CATI) and the data set is based on subjective measures only.

In 2005, the original sample contains 38,940 observations. In the wave 2014, 54,961 individuals were interviewed. In our analysis, we focus on full-time employees aged 18-64 years. We include only individuals in the sample that work at least 36 hours per week and exclude self-employed workers from the analysis. The sample is further restricted to earnings from the main job only, i.e. from the job that yields the highest income. After dropping observations with missing data on other variables of interest, our sample contains 9,495 positive wage observations in 2005 and 8,423 in 2014. For the analysis of the evolution of the GPG over time, we pool together the two cross sections of 2005 and 2014. For the analysis of the PPWG between men and women, we use the latest release, i.e. the wave of 2014. In 2005, our sample contains 4,778 women (50.3%) and 4,717 men (49.7%). In the 2014-release, 3,828 (45.4%) individuals are female and 4,595 (54.6%) are male. In 2014, 1,799 women (52.8% of total public-sector employment) and 1,607 men (47.2% of total public-sector employment) are occupied in the public sector. Thus, slightly more women than men are employed in the public sector. The OLS estimates are based on the natural logarithm of net hourly wages as dependent variable. The data set includes also a variable for monthly gross earnings. However, 98% of all observations contain missing values.[13] Therefore, we prefer to use the monthly-based net income as dependent variable. Table 1

[10] That is the model outlined in Appendix C.

[11] In the second case, we draw inference on the components of the PPWG by gender.

[12] Participation, Labor, Unemployment Survey (PLUS)

[13] The survey contains also gross annual earnings. Unfortunately, gross annual earnings divided by the number

and 2 report mean and standard deviation for some of the variables included in the analysis for the two cases under consideration, respectively. Detailed information on the variables used in the analysis can be found in Appendix E.

### 4.1.1 Descriptive Statistics Case 1

Table 1 shows that women have on average higher educational attainment than men and that their human capital increased from 2005 to 2014 (*Schooling*). For men, the increase is less pronounced. Men still outperform women in terms of labor market characteristics (*Exper* and *Tenure*). However, while the average years of experience of women increased over the last decade, men's average years of experience decreased slightly. Nonetheless, the average level of labor-market experience is still higher for men than for women in 2014. On average, men hold more often an unlimited contract in both years (*Contract_Type*). The proportion of married women and men reduced slightly over the last decade (*Married*). The share of individuals employed in Northern Italy decreased slightly for both men and women (*North*). In 2014, more females than males are emplyoed in highly specialized occupations, while for the wave of 2005, the opposite holds (*Manager*).

Table 1: Descriptive Statistics Case 1

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | Women | | | | Men | | | |
| | 2005 | | 2014 | | 2005 | | 2014 | |
| Variables | Mean | Std.Dev. | Mean | Std.Dev. | Mean | Std.Dev. | Mean | Std.Dev. |
| Exper | 16.23 | 11.33 | 17.73 | 12.08 | 20.51 | 12.86 | 20.19 | 12.95 |
| Tenure | 10.42 | 9.822 | 13.52 | 11.46 | 14.10 | 11.70 | 15.41 | 12.34 |
| Schooling | 12.72 | 2.722 | 14.30 | 1.486 | 12.26 | 2.842 | 13.95 | 1.397 |
| Contract_Type | 0.838 | 0.369 | 0.862 | 0.345 | 0.879 | 0.327 | 0.884 | 0.321 |
| Married | 0.591 | 0.492 | 0.580 | 0.494 | 0.580 | 0.494 | 0.577 | 0.494 |
| Italian | 0.989 | 0.103 | 0.987 | 0.115 | 0.994 | 0.0768 | 0.993 | 0.0857 |
| North | 0.533 | 0.499 | 0.502 | 0.500 | 0.463 | 0.499 | 0.480 | 0.500 |
| Centre | 0.205 | 0.404 | 0.223 | 0.416 | 0.183 | 0.387 | 0.211 | 0.408 |
| Manager | 0.118 | 0.323 | 0.247 | 0.431 | 0.136 | 0.343 | 0.232 | 0.422 |
| Intermed_Prof | 0.617 | 0.486 | 0.609 | 0.488 | 0.465 | 0.499 | 0.499 | 0.500 |
| Observations | 4,778 | | 3,828 | | 4,717 | | 4,595 | |

of months in a calendar year (including a $13^{\text{th}}$ month), differ by more than 800 Euros (per month) from the reported monthly gross income.

#### 4.1.2 Descriptive Statistics Case 2

Table 2 shows that the average level of educational attainment is higher in the public compared to the private sector. Women have on average higher educational attainment than men in both sectors. Female civil servants are even better educated than their female colleagues in the private sector. Similarly, men in the public sector have higher educational performance compared to their male peers in the private sector. Men outperform women in both sectors in terms of labor market presence and job tenure. About the equal amount of male and female employees is married, yet, the proportion of married employees is higher in the public sector. In the public sector, men and women are more often employed in highly specialized jobs. The proportion of highly specialized females in public employment is higher than that of males.

Table 2: Descriptive Statistics Case 2

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | Women | | | | Men | | | |
| | Private Sector | | Public Sector | | Private Sector | | Public Sector | |
| Variables | Mean | Std.Dev. | Mean | Std.Dev. | Mean | Std.Dev. | Mean | Std.Dev. |
| Exper | 14.09 | 10.65 | 21.83 | 12.29 | 17.69 | 12.57 | 24.84 | 12.35 |
| Tenure | 9.766 | 9.442 | 17.75 | 12.04 | 12.57 | 11.47 | 20.68 | 12.18 |
| Schooling | 14.13 | 1.454 | 14.48 | 1.500 | 13.79 | 1.320 | 14.26 | 1.481 |
| Contract_Type | 0.819 | 0.385 | 0.911 | 0.286 | 0.859 | 0.348 | 0.928 | 0.258 |
| Married | 0.471 | 0.499 | 0.703 | 0.457 | 0.495 | 0.500 | 0.730 | 0.444 |
| Italian | 0.978 | 0.147 | 0.997 | 0.0577 | 0.991 | 0.0964 | 0.996 | 0.0610 |
| North | 0.555 | 0.497 | 0.442 | 0.497 | 0.553 | 0.497 | 0.343 | 0.475 |
| Centre | 0.218 | 0.413 | 0.228 | 0.420 | 0.210 | 0.407 | 0.214 | 0.410 |
| Manager | 0.140 | 0.347 | 0.367 | 0.482 | 0.180 | 0.384 | 0.327 | 0.469 |
| Intermed_Prof | 0.646 | 0.478 | 0.569 | 0.495 | 0.498 | 0.500 | 0.502 | 0.500 |
| Observations | 2,029 | | 1,799 | | 2,988 | | 1,607 | |

### 4.2 Empirical Results

We first present the decomposition results from the standard Oaxaca-Blinder approach and discuss the conclusions drawn on the change of the wage gap in this framework. Next, we apply the method derived in Section 3 in order to directly estimate changes of the wage gaps and in order to draw inference on the diverse contributors to the change of the gap.

#### 4.2.1 The Gender Pay Gap over Time

A general finding in the literature is that the gap in pay by gender was reduced over time (Blau and Kahn, 2006; Goldin, 2014; Mandel and Semyonov, 2014). The part attributable to observed characteristics and therefore referred to as explained component increased, while the unexplained part, i.e. the component due to differences in returns to wage-related characteristics and differences in the intercepts, decreased.

Indeed, by applying the traditional approach to our data, we also find a reduction of the GPG in hourly wages over time; 12.4% in 2005 and 9.5% in 2014.[14] Table 3 shows that the gaps are highly statistically significant in either case. The composition of the gap also changed across the decade. In 2005, the explained component does not play a role in determining the GPG (as it is not statistically significant), while in 2014, the endowments part becomes highly statistically significant and contributes to a narrowing of the GPG (negative term). Differences in the unexplained component are statistically significant in both years. The component in 2014 decreased slightly (86.2% in 2005 versus 84.3% in 2014). A relatively small decrease in the unexplained component of the GPG in 2014 is in line with results of other scholars (e.g. Fortin, 2008; Mandel and Semyonov, 2014). In 2005, differences in endowments and coefficients that exist simultaneously between men and women, have a statistically significant impact as well, what is no longer the case in 2014.

All in all, our data delivers results in line with the literature, when applying the standard estimation methodology. The GPG declined over the last decade, differences in endowments (in favor of women) have become statistically significant in 2014 and the part of the GPG due to differences in prices has declined.

[14]The estimated GPGs in this paper are larger than the pay gaps found by Eurostat (2017). Eurostat (2017) finds wage gaps amounting to 4.4% in 2006 (missing in 2005) and 6.1% in 2014 for Italy. These relatively larger gaps are due to our sample restriction of considering only employees working at least 36 hours per week.

Table 3: Standard Decomposition of the GPG in 2005 and 2014

| | (1) | (2) |
|---|---|---|
| Variables | 2005 | 2014 |
| *Differential* | | |
| $\overline{ln(w_M)}$ | 1.999*** | 2.134*** |
| | (0.006) | (0.007) |
| $\overline{ln(w_F)}$ | 1.875*** | 2.039*** |
| | (0.006) | (0.007) |
| Difference | 0.124*** | 0.095*** |
| | (0.008) | (0.009) |
| *Decomposition* | | |
| Endowments | 0.008 | -0.016*** |
| | (0.006) | (0.006) |
| Coefficients | 0.107*** | 0.107*** |
| | (0.008) | (0.009) |
| Interaction | 0.009* | 0.004 |
| | (0.006) | (0.006) |
| *%−Contribution* | | |
| Endowments | 6.5 | 12.6 |
| Coefficients | 86.2 | 84.3 |
| Interaction | 7.3 | 3.1 |
| Observations | 9,495 | 8,423 |

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

*Notes:* For the GPG in 2014, the %–contribution for the endowments effect is $\frac{|0.016|}{(|0.016|+0.107+0.004)} \times 100$.

Next, we directly estimate the change of the GPG between 2014 and 2005 and decompose that change in explained and unexplained components as well as interaction effects. Table 4, column (1), shows the base model of case 1. The coefficient estimate of *femyear* shows the change of the GPG from 2014 to 2005. The difference between the GPG in 2014 and 2005 amounts to $-0.03$ log points and is statistically significant. Given the negative sign, the GPG has decreased over time. The magnitude as well as the sign of the change is also visible by looking at the aggregate GPGs from the outcome of the standard estimation in Table 3. However, now we can also conclude that this reduction in the GPG is statistically significant. The full model is presented in column (2) of Table 4. We immediately see that the part of the price effect due to differences in the intercepts, $\hat{\alpha}_1^{full}$, is not statistically significant. Similarly, the effect of being a woman or in year 2005, all else equal, becomes statistically insignificant. The remaining coefficient estimates show the expected signs.[15]

[15]The full regression output is shown in Table F.1 in Appendix F.

Table 4: OLS Estimates of Log Hourly Wages – Case 1, Base and Full Specification

| | (1) | (2) |
|---|---|---|
| | Basic Specification | Full Specification |
| femyear | -0.028** | -0.051 |
| | (0.012) | (0.185) |
| female | -0.095*** | -0.148 |
| | (0.009) | (0.152) |
| year | -0.135*** | -0.010 |
| | (0.009) | (0.130) |
| *Groups of Covariates* | | |
| Labor Market Presence | No | Yes |
| Educational Attainment | No | Yes |
| Job Characteristics | No | Yes |
| Demographic and Family Background Characteristics | No | Yes |
| Industrial and Occupational Dummies | No | Yes |
| Interaction Terms | No | Yes |
| Observations | 17,918 | 17,918 |
| R-squared | 0.050 | 0.291 |

Robust standard errors in parentheses
$***p < 0.01, **p < 0.05, *p < 0.1$

Table 5 presents the results from our proposed decomposition. The results show that the change of the GPG is only explained by the quantity effect. The change of the GPG over time is explained by changes in observed characteristics between men and women (in favor of women) over time. We know from Table 1 that women's set of observable human capital and labor market characteristics (*Schooling*, *Exper*) is increasing over the last decade, while that of men is partly even decreasing (*Exper*) or remained lower than that of women (educational attainment). In fact, in educational matters, women have outpaced men (Goldin, 2006). The results from the standard method discussed in Section 4.2.1 suggest that the coefficients part of the GPGs, i.e. the part due to differences in returns on observable characteristics, was a main contributor to the GPG in either year with decreasing importance in the ending period. However, by estimating the difference of the GPG over time directly, we see that this so-called discriminatory part has not significantly changed over the last ten years in Italy. The decomposition shows that the only factor that contributes statistically significantly to the narrowing of the gap are better observable characteristics for women. Hence, the closing of the GPG is not explained by anti-discrimination laws, changes in attitudes towards women in the labor market or changes in the family structure and birth control. The latter is, apart from the unexplained part ($U$), caught by the interaction effects accounting for simultaneous differences in endowments over time and changing prices between men and women ($I1$) as well as variation in the set of endowments by gender and changing prices over time ($I2$). The components account for the effects of changes in institutional settings or attributes towards women on prices (given differences in endowments). Yet, the effects are not statistically significant.

Table 5: Decomposition of the Change in the GPG over time – Case 1

| | (1) |
|---|---|
| | Pooled Sample (2005 and 2014) |
| *Decomposition* | |
| E | -0.023*** |
| | (0.007) |
| I1 | 0.002 |
| | (0.013) |
| I2 | -0.006 |
| | (0.006) |
| u | 0.050 |
| | (0.179) |
| Total = E + I1 + I2 + u | 0.023 |
| | (0.185) |
| Observations | 17,918 |

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

#### 4.2.2 The Public-Private Sector Wage Gap between Men and Women

In the literature a positive wage gap between the public and the private sector is found (Melly, 2005; Lucifora and Meurs, 2006; Arulampalam *et al.*, 2007; Christofides and Michael, 2013; Mandel and Semyonov, 2014). Table 6 shows that also in our data for Italy, we find differences in earnings by sector, with higher wage levels in the public sector. A general result is that women are better-off in the public compared to the private sector, while for men the public-sector premia is less important (e.g. Melly, 2005). We find different PPWGs by gender as well; 23.2% for women and 19.8% for men (see Table 6). Both gaps are found to be highly statistically significant. Also, the composition of the PPWGs differs by gender. For women, the PPWG is mainly due to the unexplained part (54.3%). On the contrary, for men, the endowments effect is the main driver of the pay gap (59.9%). Interaction effects are rather small but more important for the wage gap in the female subsample (15.5% compared to 6.1% in the male subsample).

The decomposition outcome of the PPWG between men and women using our proposed model is provided in Tables 7–8. The results from the base model suggest that there is a positive and statistically significant difference in the PPWG between men and women equal to $-0.03$ log points.[16] The dummy variable for working in the private sector (*private*) negative and significant, tells us that there is a wage loss for working in the private sector compared to public-sector employment. As expected, the coefficient on the $female$-dummy shows that being a women has a significant and negative impact on labor income. In the full model, the effect of private-sector employment as well as being female on wages turns statistically insignificant. Yet, the interaction term *fempriv*, is statistically significant and strongly negative ($-0.72$). Hence, $\hat{\alpha}_1^{full}$, i.e. the part of the price or unexplained effect due to differences in the starting points is statistically significant. This implies that there is a premia for simply working in the public sector and that this premia is higher for women than for men. Again, the remaining parameter

[16] Indicated by the interaction of the dummies *female* and *private*; *fempriv*.

Table 6: Standard Decomposition of the PPWG for Women and Men in 2014

| | (1) | (2) |
|---|---|---|
| | Women | Men |
| *Differential* | | |
| $\overline{ln(w_{\text{Public\_Sector}})}$ | 2.162*** | 2.263*** |
| | (0.009) | (0.011) |
| $\overline{ln(w_{\text{Private\_Sector}})}$ | 1.930*** | 2.065*** |
| | (0.010) | (0.008) |
| Difference | 0.232*** | 0.198*** |
| | (0.013) | (0.013) |
| *Decomposition* | | |
| Endowments | 0.070*** | 0.118*** |
| | (0.015) | (0.015) |
| Coefficients | 0.126*** | 0.067*** |
| | (0.016) | (0.024) |
| Interaction | 0.036** | 0.012 |
| | (0.018) | (0.023) |
| *%–Contribution* | | |
| Endowments | 30.2 | 59.9 |
| Coefficients | 54.3 | 34.0 |
| Interaction | 15.5 | 6.1 |
| Observations | 3,828 | 4,595 |

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

estimates impact on wages as expected.[17]

By looking at the decomposition, we find that the difference in observable characteristics across sectors and gender, $E$, does play a statistically significant role in explaining the difference of the PPWG between men and women. In particular, the explained component drives the negative PPWG by gender as best-educated females are more often located in the public sector (Bordogna, 2012; Piazzalunga and Di Tommaso, 2015). The difference in the unexplained component, $u$, of the PPWG between men and women is significant as well and shows that the change works towards a positive PPWG between men and women. This implies that more egalitarian pay schemes in the public sector are ruled out by female discrimination in prices in both sectors. Moreover, we observe simultaneously differences in characteristics between women and men as well as difference in coefficients between the private and the public sector (for men; $I1$). Hence, more favorable endowments of men in the private sector compared to women in the private sector and higher pay schemes in the public sector narrow the (negative) PPWG between men and women. All in all, for case 2, the conclusions drawn from the standard estimation are confirmed; both quantity and price effects contribute to the difference in the PPWG between men and women. Yet, we gain the additional insight that the set-up or organization of the public sector does play a role as well. That is institutional norms of the public sector being relatively more gender-equal in combination with more discriminatory practices against women in the private sector lead to an increase of the significant difference in the PPWG between men

[17] The complete regression outcome of the full model is shown in Table F.2 in Appendix F.

and women in 2014 in Italy.

Table 7: OLS Estimates of Log Hourly Wages – Case 2, Base and Full Specification

| | (1) | (2) |
|---|---|---|
| | Basic Sepcification | Full Specification |
| fempriv | -0.034* | -0.724** |
| | (0.019) | (0.289) |
| female | -0.101*** | 0.278 |
| | (0.014) | (0.196) |
| private | -0.198*** | 0.309 |
| | (0.013) | (0.205) |
| *Groups of Covariates* | | |
| Labor Market Presence | No | Yes |
| Educational Attainment | No | Yes |
| Job Characteristics | No | Yes |
| Demographic and Family Background Characteristics | No | Yes |
| Industrial and Occupational Dummies | No | Yes |
| Interaction Terms | No | Yes |
| Observations | 8,423 | 8,423 |
| R-squared | 0.069 | 0.236 |

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

Table 8: Decomposition of the Change in the PPWG by Gender – Case 2

| | (1) |
|---|---|
| | Pooled Sample (Women and Men) |
| *Decomposition* | |
| E | -0.028*** |
| | (0.011) |
| I1 | 0.041* |
| | (0.022) |
| I2 | 0.002 |
| | (0.011) |
| u | 0.675* |
| | (0.357) |
| Total = E + I1 + I2 + u | 0.689* |
| | (0.360) |
| Observations | 8,423 |

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

## 5 Conclusion

Adding to the discussion of the convergence of the GPG over time and the persistence of a PPWG between men and women, we propose an alternative decomposition method allowing to draw inference on the difference of two wage gaps on aggregate as well as on its components. The model set-up bases on the OVB formula and the Gelbach decomposition. Despite additional insights on the composition of differences in gaps, the method can be made robust to the choice of the reference category (Reimers, 1983; Cotton, 1988; Neumark, 1988; Oaxaca and Ransom, 1994; Fortin, 2008) as well as to the indeterminacy problem (Lee, 2015). The method proposed can be applied to a variety of cases such as differences in the GPG and its drivers over time, across countries, sectors, occupations or unions. We empirically consider two cases; the change of the GPG over time as well as the PPWG between men and women in Italy.

The observed closing of the GPG over time is heavily discussed in the literature and the determination of the reasons of the narrowing is of huge interest, especially with regard to policy implications (Blau and Kahn, 2006; Goldin, 2014; Blau and Kahn, 2016). Similarly, the PPWG that is found to differ for men and women is a topic of on-going research (Melly, 2005). Yet, up to know, in the standard estimation framework, direct inference on the difference of pay gaps and changes in their components could not be drawn. Conclusions were rather drawn by estimating the pay gaps separately in different subsamples and comparing the results ex post. In this way, it is not possible to test the significance of the change in the estimated pay gaps on aggregate or the components of the decomposition. Besides the estimation of the change of the GPG over time on aggregate as well as of the explained and unexplained component, our method also catches otherwise unobserved interaction effects across the respective groups of interest.

We find a significant convergence of the GPG over the last decade in Italy. The convergence of the GPG over time was found to be only explained by a reduction in differences in observable characteristics by gender. On the contrary, by estimating the GPG separately for 2005 and 2014, i.e. following the standard approach in the literature, the relative decline in the contribution of the price component to the wage gap might have led to the conclusion that the implementation of anti-discrimination laws and changing attitudes towards women in the labor market have influenced the narrowing of the pay gap over time as well. Yet, these policies as well as changes in social norms seem to have been less effective than expected a priori. Thereby, we add to the literature on the covergence of the GPG over time for the case of Italy the finding that the closing of the pay differential by gender over the last decade was entirely explained by the catching-up of women in terms of endowments. The results for the second case we have examined, i.e. the PPWG between men and women, point the attention to differences in the structure of the public and private sector, which are found to be important to explain the differential. Better educated females are more often employed in the public sector given more egalitarian pay schemes as well as job stability (Bordogna, 2012; Piazzalunga and Di Tommaso, 2015). In this case, the results derived from the standard approach concerning the explained and unexplained part are confirmed in the sense that both components contribute significantly to the change of the PPWG between men and women.

All in all, the analysis with the proposed decomposition method offers a better understanding of what has led to the narrowing of the GPG in the last ten years and what drives the difference in the PPWG between men and women. Most importantly, we can infer what drives the difference in the respective pay gaps in a statistically significant manner. The model proposed offers an intuitive approach to directly estimate changes in wage gaps between groups and can be applied to various problems.

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# Appendix

The robustness of the decompostion is for simplicity shown for the case of the GPG. Deriving the robust model based on the GPG allows also for a better comparison of the method with the approaches in the literature (e.g. Fortin, 2008, uses the case of the GPG).[18] In Appendix C, when considering differences of gaps, we derive the model for the GPG changing over time. Notably, the methods can be applied to various other decomposition problems.

## A Solving the Index-Number Problem of Decompositions using the Intercept-Shift Approach

As is well known in the literature, the Oaxaca-Blinder decomposition is not unique. Therefore, the choice of the non-discriminatory wage structure (men or women) matters and leads to different results (Cotton, 1988; Oaxaca and Ransom, 1994). Several approaches have been proposed to circumvent this problem (Reimers, 1983; Cotton, 1988; Neumark, 1988; Oaxaca and Ransom, 1994; Fortin, 2008). We extend the method proposed by Gelbach (2016) in order to have a wage decomposition invariant to the reference category adopted. In particular, we adopt the decomposition proposed by Fortin (2008) that includes gender intercept shifts along with an identification restriction in the regression of females and males pooled together, when considering the standard case of the GPG for individual $i$:

$$ln(w_i) = \gamma_0 + \gamma_{0F} F_i + \gamma_{0M} M_i + X_i \gamma + \epsilon_i$$

subject to:

$$\gamma_{0F} + \gamma_{0M} = 0$$

where $F_i$ is equal to one if the individual is female and zero otherwise and $M_i$ equals one if the individual is male and zero otherwise, i.e. $F_i = (1 - M_i)$. Correspondingly, the index $F$ identifies women and the index $M$ identifies men. For the pooled regression with male and female dummies, respectively, evaluated at the mean, we have:

$$\overline{ln(w_M)} = \hat{\gamma_0} + \hat{\gamma}_{0M} M + \bar{X}_M \hat{\gamma}$$
$$\overline{ln(w_F)} = \hat{\gamma_0} + \hat{\gamma}_{0F} F + \bar{X}_F \hat{\gamma}$$

The identification restriction imposes that the pooled wage equation truly represents a non-discriminatory wage structure, i.e. a wage structure, where the advantage of men is equal to the

[18] The derived model is robust to the index-number problem and invariant with respect to categorical variables as well as robust to the indeterminacy problem.

disadvantage of women:

$$\overline{ln(w_M)} - \overline{ln(w_F)} = (\bar{X}_M - \bar{X}_F)\hat{\gamma} + (\hat{\gamma}_{0M} - \hat{\gamma}_{0F})$$

The first component on the right-hand side, $(\bar{X}_M - \bar{X}_F)\hat{\gamma}$, is the explained part, while $\hat{\gamma}_{0M}$ and $\hat{\gamma}_{0F}$ are the *advantage of men* and the *disadvantage of women*, respectively. In particular, from the difference of the wage regression separately for men and women and the pooled wage regression with a gender dummy, we have:

$$\begin{aligned} \hat{\gamma}_{0M} &= \bar{X}_M(\hat{\beta}_M - \hat{\gamma}) + (\hat{\beta}_{0M} - \hat{\gamma}_0) \qquad \textit{advantage of men} \\ \hat{\gamma}_{0F} &= \bar{X}_F(\hat{\beta}_F - \hat{\gamma}) + (\hat{\beta}_{0F} - \hat{\gamma}_0) \qquad \textit{disadvantage of women} \end{aligned}$$

where $\hat{\beta}_{0M}, \hat{\beta}_{0F}$ are the intercepts and $\hat{\beta}_M, \hat{\beta}_F$ are the estimated coefficients of wage equations estimated separately for men and women:

$$ln(w_{iM}) = \beta_{0M} + X_{iM}\beta_M + \epsilon_{iM} \tag{A.1}$$

$$ln(w_{iF}) = \beta_{0F} + X_{iF}\beta_F + \epsilon_{iF} \tag{A.2}$$

In order to adopt the above wage decomposition within the conditional decomposition framework proposed by Gelbach (2016), we estimate the following wage equation:

$$ln(w_i) = \gamma_0 + \gamma_{0F}F_i + \gamma_{0M}M_i + X_i\gamma + X_iF_i\gamma_{XF} + X_iM_i\gamma_{XM} + \nu_i \tag{A.3}$$

subject to:

$$\begin{aligned} \gamma_{0F} + \gamma_{0M} &= 0 \\ \gamma_{X_kF} + \gamma_{X_kM} &= 0 \quad \text{for } k = 1 \dots K \end{aligned}$$

where $\gamma_{X_kF}$ and $\gamma_{X_kM}$ are the parameters of the interaction term between the $k$th regressor $X_i$ and the dummy $F_i$ and $M_i$, respectively. The error term is represented by $v_i$. Then,

$$\begin{aligned} \overline{ln(w_M)} &= \hat{\gamma}_0 + \hat{\gamma}_{0M} + \bar{X}_M\hat{\gamma} + \bar{X}_M\hat{\gamma}_{XM} \\ \overline{ln(w_F)} &= \hat{\gamma}_0 + \hat{\gamma}_{0F} + \bar{X}_F\hat{\gamma} + \bar{X}_F\hat{\gamma}_{XF} \end{aligned}$$

Consequently, the GPG becomes:

$$\begin{aligned} \overline{ln(w_M)} - \overline{ln(w_F)} &= (\hat{\gamma}_{0M} - \hat{\gamma}_{0F}) + (\bar{X}_M - \bar{X}_F)\hat{\gamma} + \bar{X}_M\hat{\gamma}_{XM} - \bar{X}_F\hat{\gamma}_{XF} \\ &= -2\hat{\gamma}_{0F} + (\bar{X}_M - \bar{X}_F)\hat{\gamma} - (\bar{X}_M + \bar{X}_F)\hat{\gamma}_{XF} \end{aligned} \tag{A.4}$$

First, we observe that it can be easily shown that there exists the following relationship between the parameter estimates of equations (A.2)-(A.1) and (A.3):

$$\begin{aligned}
\hat{\gamma} + \hat{\gamma}_{XF} &= \hat{\beta}_F \\
\hat{\gamma}_0 + \hat{\gamma}_{0F} &= \hat{\beta}_{0F} \\
\hat{\gamma} - \hat{\gamma}_{XF} &= \hat{\beta}_M \\
\hat{\gamma}_0 - \hat{\gamma}_{0F} &= \hat{\beta}_{0M}
\end{aligned}$$

Therefore, the GPG of (A.4) can be re-written in terms of the Fortin decomposition as:

$$\begin{aligned}
\overline{ln(w_M)} - \overline{ln(w_F)} &= (\hat{\beta}_{0M} - \hat{\gamma}_0) - (\hat{\beta}_{0F} - \hat{\gamma}_0) + (\bar{X}_M - \bar{X}_F)\hat{\gamma} + \bar{X}_M(\hat{\beta}_M - \hat{\gamma}) - \bar{X}_F(\hat{\beta}_F - \hat{\gamma}) \\
&= (\bar{X}_M - \bar{X}_F)\hat{\gamma} + [\bar{X}_M(\hat{\beta}_M - \hat{\gamma}) + (\hat{\beta}_{0M} - \hat{\gamma}_0)] - [\bar{X}_F(\hat{\beta}_F - \hat{\gamma}) + (\hat{\beta}_{0F} - \hat{\gamma}_0)]
\end{aligned} \tag{A.5}$$

Second, the estimation can be recast in terms of the sequential decomposition of Gelbach by considering the following base model for individual $i$:

$$ln(w_i) = \gamma_0^{base} + (F_i - M_i)\gamma_{0F}^{base} + \epsilon_i^{base} \tag{A.6}$$

where the $1 \times 2$ vector of regressors $X_{i1}$ of the base specification contains for each obseravation $i$ a constant and the difference between the two dummy variables $F_i$ and $M_i$, $(F_i - M_i)$. The full model is defined as follows:

$$ln(w_i) = \gamma_0^{full} + (F_i - M_i)\gamma_{0F}^{full} + X_i\gamma + (F_i - M_i)X_i\gamma_{XF} + \epsilon_i^{full} \tag{A.7}$$

where the regressors $X_i$ as well as the interaction between $X_i$ and the difference between the two dummy variables $F_i$ and $M_i$ are contained in the $1 \times 2K$ vector $X_{i2}$. The regressors in $X_{i2}$ are the omitted variables. By the OVB formula the following relationship holds:

$$\begin{bmatrix} \hat{\gamma}_0^{base} \\ \hat{\gamma}_{0F}^{base} \end{bmatrix} = \begin{bmatrix} \hat{\gamma}_0^{full} \\ \hat{\gamma}_{0F}^{full} \end{bmatrix} + (X'_{i1}X_{i1})^{-1}X'_{i1}X_{i2} \begin{bmatrix} \hat{\gamma} \\ \hat{\gamma}_{XF} \end{bmatrix} \tag{A.8}$$

where $\left[\hat{\gamma}_0^{base}, \hat{\gamma}_{0F}^{base}\right]^T$ is the $2 \times 1$ vector of coefficient estimates of $X_1$ from the base model (A.6) evaluated at the mean; $\left[\hat{\gamma}_0^{full}, \hat{\gamma}_{0F}^{full}\right]^T$ is the $2 \times 1$ vector containing the coefficient estimates of $X_{i1}$ from the full model (A.7) evaluated at the mean and $\left[\hat{\gamma}, \hat{\gamma}_{XF}\right]^T$ is the vector of coefficients estimates of $X_{i2}$ from the full model (A.7) at the mean, i.e. $\hat{\gamma}^{full}$ with dimension $2K \times 1$. First observe that $\hat{\gamma}_{0F}^{base}$ is equal to $\frac{\overline{ln(w_F)} - \overline{ln(w_M)}}{2}$ and that $\hat{\gamma}_{0F}^{full}$ is equal to $\frac{\hat{\beta}_{0F} - \hat{\beta}_{0M}}{2}$. As in Section 3, we are interested in the second row of equation (A.8). Given the relationship in (A.8), we observe that:

$$\hat{\gamma}_{0F}^{base} = -\frac{\overline{ln(w_M)} - \overline{ln(w_F)}}{2} = -\frac{\Delta}{2} = \hat{\gamma}_{0F}^{full} + \eta\hat{\gamma}^{full} \quad \text{(A.9)}$$

where $\Delta$ is the GPG and $\eta = \left[\frac{(\bar{x}_F - \bar{x}_M)}{2}, \frac{(\bar{x}_F + \bar{x}_M)}{2}\right]$ contains the sample means of obervable characteristics in $X_i$ obtained from the linear projection of $X_i$ and $(F_i - M_i)X_i$ with respect to $X_{i1}$. The row vector $\eta$ has dimension $1 \times 2K$. Moreover, we have $\hat{\gamma}_{0F}^{full} = \frac{\hat{\beta}_{0F} - \hat{\beta}_{0M}}{2} = \frac{(\hat{\beta}_{0F} - \hat{\gamma}_0^{full}) - (\hat{\beta}_{0M} - \hat{\gamma}_0^{full})}{2}$. Consequently, the GPG can be written as:

$$-2\gamma_{0F}^{base} = -2\gamma_{0F}^{full} + (\bar{x}_M - \bar{x}_F)\gamma - (\bar{x}_M - \bar{x}_F)\gamma_{XF} \quad \text{(A.10)}$$

what completes the proof of decomposition equivalence.

# B Invariance Decomposition with respect to Categorical Variables

A second type of identification issue arises when dummy variables are considered in a detailed wage decomposition. Oaxaca and Ransom (1999) show that the assignment of the explained part of the GPG to specific variables is not invariant to the choice of reference groups. This problem can be easily solved by imposing the following parameter restrictions as proposed by Gardeazabal and Ugidos (2004), Yun (2005) and Fortin (2008):

$$\sum_{j=1}^{C_k} \gamma_{jk} = 0, \ k \in C \quad \text{(B.1)}$$

where $C$ denotes the set of categorical variables, and $C_k$ the number of categories for variable $k$. The neutral, i.e. non-sensitive to any left-out category, Oaxaca-Blinder decomposition follows. The zero-sum restriction (B.1) is applied to the wage equation, when female and male wages are estimated separately as well as to the pooled regression with gender dummies. The latter is additionally estimated with the identification restriction $\gamma_{0M} + \gamma_{0F} = 0$ on the gender parameters. Thereby, the intercepts, $\beta_{0M}$, $\beta_{0F}$ and $\gamma_0$, are no longer influenced by the choice of the reference category in the case of categorical variables.

The restriction (B.1) can also be applied to the method proposed in Section 3 leading to indicator variables that are invariant to the choice of the left-out category in the case of categorical variables.

## C Estimating Differences of Gaps with the Intercept-Shift Approach

The extension of the decomposition described in Appendix A to the case of the estimation of the difference of wage gaps follows straightforward. We consider, as in Section 3.2, the indicator variable $Y_i$ that takes values $\{0, 1\}$. Again, when the indicator variable $Y_i$ is used as an index $(Y)$, $Y_i = 0$ corresponds to $B$ and $Y_i = 1$ to $A$. Similarly, in order to circumvent confusion with the intercept (referred to as $\beta_0$ in coherence with Appendix A), the gender index is not numerical here, but $G \in \{F, M\}$ with $F =$ female and $M =$ male replacing the numerical index $\{1, 0\}$, respectively. The set of regressors considered in Section 3.2 are hence transformed as follows:

$$X_{i1} = [1, (F_i - M_i)Y_i, (F_i - M_i), Y_i]$$
$$X_{i2} = [X, (F_i - M_i)X_i, Y_iX_i, (F_i - M_i)Y_iX_i]$$

for each individual $i$, with $X_{i1}$ having dimension $1 \times 4$ and $X_{i2}$ having dimension $1 \times 4K$. $X_{i1}$ contains the interaction of $(F_i - M_i)$ with $Y_i$; $(F_i - M_i)Y_i$. The second set of regressors, $X_{i2}$ contains the $1 \times K$ vector of characteristics $X_i$ as well as the interaction of $X_i$ with $(F_i - M_i)$ and $Y_i$; $(F_i - M_i)X_i, Y_iX_i$ and $(F_i - M_i)Y_iX_i$, respectively. The base model is then:

$$ln(w_i) = \gamma_0^{base} + (F_i - M_i)Y_i\gamma_{FY}^{base} + (F_i - M_i)\gamma_F^{base} + Y_i\gamma_Y^{base} + \epsilon_i^{base} \quad \text{(C.1)}$$

while the full model is defined as follows:

$$\begin{aligned} ln(w_i) = \gamma_0^{full} &+ (F_i - M_i)Y_i\gamma_{F_iY_i}^{full} + (F_i - M_i)\gamma_F^{full} + Y_i\gamma_Y^{full} \\ &+ X_i\gamma + (F_i - M_i)X_i\gamma_{XF} + Y_iX_i\gamma_{XY} + (F_i - M_i)Y_iX_i\gamma_{XYF} + \epsilon_i^{full} \end{aligned} \quad \text{(C.2)}$$

where $\gamma_0^{base}$ is the constant and $\gamma_{FY}^{base}, \gamma_F^{base}, \gamma_Y^{base}$ are the coefficients of the the base model (C.1), $\gamma_0^{full}, \gamma_{FY}^{full}, \gamma_F^{full}, \gamma_Y^{full}$ are the corresponding constant and coefficients of $X_{i1}$ from the full model (C.2). $\gamma, \gamma_{XF}, \gamma_{XY}, \gamma_{XYF}$ are the $K \times 1$ coefficient vectors of $X_{i2}$ from the full model (C.2). The second row of the linear projection of $X_{i2}$ with respect to $X_{i1}$ is contained in the following $1 \times 4K$ vector:

$$\zeta = \left[\tfrac{(\bar{x}_{0A} - \bar{x}_{1A}) - (\bar{x}_{0B} - \bar{x}_{1B})}{2}, \tfrac{(\bar{x}_{0A} + \bar{x}_{1A}) - (\bar{x}_{0B} + \bar{x}_{1B})}{2}, \tfrac{(\bar{x}_{1A} - \bar{x}_{0A})}{2}, \tfrac{(\bar{x}_{1A} + \bar{x}_{0A})}{2}\right]$$

Consider the equivalence between the following parameter estimates evaluated at the mean:

$$\begin{aligned}
\hat{\gamma}_0^{full} - \hat{\gamma}_{FY}^{full} - \hat{\gamma}_F^{full} + \hat{\gamma}_Y^{full} &= \hat{\beta}_{0,MA} \\
\hat{\gamma}_0^{full} + \hat{\gamma}_{FY}^{full} + \hat{\gamma}_F^{full} + \gamma_Y^{full} &= \hat{\beta}_{0,FA} \\
\hat{\gamma}_0^{full} + \hat{\gamma}_F^{full} &= \hat{\beta}_{0,FB} \\
\hat{\gamma}_0^{full} - \hat{\gamma}_F^{full} &= \hat{\beta}_{0,MB} \\
\hat{\gamma} + \hat{\gamma}_{XF} + \hat{\gamma}_{XY} + \hat{\gamma}_{XYF} &= \hat{\beta}_{FA} \\
\hat{\gamma} - \hat{\gamma}_{XF} + \hat{\gamma}_{XY} - \hat{\gamma}_{XYF} &= \hat{\beta}_{MA} \\
\hat{\gamma} + \hat{\gamma}_{XF} &= \hat{\beta}_{FB} \\
\hat{\gamma} - \hat{\gamma}_{XF} &= \hat{\beta}_{MB}
\end{aligned}$$

Observe that $\hat{\gamma}_{FY}^{base}$ is equal to $\frac{\Delta GPG}{2}$ and $\hat{\gamma}_{FY}^{full}$ is equal to $\frac{(\hat{\beta}_{0,MB} - \hat{\beta}_{0,FB}) - (\hat{\beta}_{0,MA} - \hat{\beta}_{0,FA})}{2}$. Given the fact that

$$\begin{aligned}
\hat{\gamma}_{FY}^{base} &= \frac{\left(\overline{ln(w_{MB})} - \overline{ln(w_{FB})}\right) - \left(\overline{ln(w_{MA})} - \overline{ln(w_{FA})}\right)}{2} \\
&= \frac{\Delta GPG}{2}
\end{aligned}$$

The relationship:

$$\hat{\gamma}_{FY}^{base} = \hat{\gamma}_{FY}^{full} + \zeta \hat{\gamma}^{full}$$

can be re-written in terms of the $\Delta GPG$ as:

$$\begin{aligned}
2\hat{\gamma}_{FY}^{base} &= \Delta GPG = \\
&= \underbrace{[(\hat{\beta}_{0,MB} - \hat{\beta}_{0,FB}) - (\hat{\beta}_{0,MA} - \hat{\beta}_{0,FA})]}_{\hat{\gamma}_{FY}^{full}} + \underbrace{(\Delta \bar{x}^B - \Delta \bar{x}^A)\hat{\gamma}}_{\Lambda} \\
&+ \underbrace{(\sum \bar{x}^{\mathrm{A}} - \sum \bar{x}^B)\hat{\gamma}_{XF}}_{\Omega} - \underbrace{\Delta \bar{x}^A \hat{\gamma}_{XY}}_{\Theta} + \underbrace{\sum \bar{x}^A \hat{\gamma}_{XYF}}_{\Upsilon}
\end{aligned}$$

where $\Delta \bar{x}^Y$ is the difference between the average level of observed characteristics of men and women in a certain year, with $Y \in \{A, B\}$ and $\sum \bar{x}^Y$ represents the sum of observable labor market characteristics present for men and women in $Y$. Recall that the model can be re-written in terms of the OVB formula as follows:

$$\begin{aligned}
2\hat{\gamma}_{FY}^{base} &= \hat{\gamma}_{FY}^{full} + \hat{\delta}^{\Lambda} + \hat{\delta}^{\Omega} + \hat{\delta}^{\Theta} + \hat{\delta}^{\Upsilon} \\
\hat{P} + \hat{Q} &= \hat{\gamma}_{FY}^{full} + \hat{\delta}^{\Lambda} + \hat{\delta}^{\Omega} + \hat{\delta}^{\Theta} + \hat{\delta}^{\Upsilon}
\end{aligned}$$

with $P$ accounting for the price effect and $Q$ for the quantity effect. In particular,

$$\hat{P} = \hat{\gamma}_{FY}^{full} + \Upsilon$$
$$\hat{Q} = \Omega + \underbrace{\Theta}_{Y\text{-specific term}} + \underbrace{\Lambda}_{\text{gender-specific term}}$$

$\hat{\gamma}_{FY}^{full}$ represents the change in the disadvantage of women over time. Thereby, accounting for the relative improvement (or deterioration) of women's position in the labor market. $\Lambda$ measures the amount of the pay difference attributable to differences in observable characteristics assuming the same prices over time and gender. $\Omega$ accounts for differences in human capital and other observable labor market characteristics in the economy over time. The underlying prices are the coefficient estimates obtained when holding $F_i$ fixed at 1 given $X_i$. Equivalently, the prices could be expressed as the coefficient estimates obtained when holding $F_i$ fixed at 0 given $X_i$ thanks to the constraint imposed: $\gamma_{XF} = -\gamma_{XM}$. $\Theta$ accounts for differences in endowments by gender holding the second indicator variable fixed, i.e. setting the index $Y = A$. The component $\Upsilon$ can be re-written as:

$$\begin{aligned}
\Upsilon &= [\sum \bar{x}^A \hat{\gamma}_{XYF}] \\
&= [\bar{x}_{1A}\hat{\gamma}_{XYF} + \bar{x}_{0A}(-\hat{\gamma}_{XYM})] \\
&= \underbrace{\bar{x}_{FA}\hat{\gamma}_{XYF}}_{\text{disadvantage of women}} - \underbrace{\bar{x}_{MA}\hat{\gamma}_{XYM}}_{\text{advantage of men}}
\end{aligned}$$

For the component $\Upsilon$, the underlying set of characteristics are the average male and female endowments observed in $Y = A$, respectively. The prices can be expressed in terms of men's advantage or women's disadvantage given $X_i$.

Again, the pooled wage equation including the gender parameters and the male and female earnings equations are estimated separately using additional constraints for each categorical variable, i.e. under the zero-sum constraint (B.1).

# D Intercept-Shift Approach versus Pooled-Sample Approach

Lee (2015) shows that the intercept-shift approach proposed by Fortin (2008) presents two drawbacks. Firstly, the reference parameter for the Oaxaca-Blinder decomposition, i.e. the parameter that would prevail in a 'fair' world under no discrimination, relies on the variance difference among categories. Secondly, the reference intercept is arbitrary: the same Oaxaca-Blinder decomposition holds with vastly different reference intercepts.

However, it can be easily shown that our proposed decomposition does not suffer from any of these aspects. Our decomposition arises from a specification that allows different intercepts and slopes. In addition, the constraints imposed on the parameters that identify the counterfactual reference parameters are the parameters such that the advantage of men is equal to the disadvantage of women. In fact, in our model the slope that would prevail under *no discrimination*, $\gamma$, is the sample average of the group slopes; $\beta_{0M}$ and $\beta_{0F}$:

$$\gamma = 0.5\beta_{0M} + 0.5\beta_{0F}$$

i.e. it is equivalent to considering the weights proposed by Reimers (1983).[19] Moreover, the constraint:

$$\beta_{0F} - \gamma_{0F} = \beta_{0M} + \gamma_{0F}$$

prevents the indeterminacy problem shown by Lee (2015). It turns out, that in our model, the intercept indeterminacy problem highlighted by Lee (2015) is ruled out by imposing the constraint that the advantage of men should be equal to the disadvantage of women.

[19]See also Lee (2015).

# E Definition of Variables

Table E.1: Definition of Variables

| Variable Name | Definition |
|---|---|
| | Dependent Variables |
| Lhwage | Natural logarithm of net hourly wages<br>Hourly wages in Euros, net of taxes and social security contributions |
| | Independent Variable |
| | Group Dummies and Interaction Terms |
| female | One if the individual is a woman, zero otherwise |
| year | One if year is 2005, zero otherwise |
| private | One if individual is employed in the private sector |
| femyear | Interactive effect of *year* and *female*, i.e. one if employee is observed in 2005 and is female, zero otherwise |
| fempriv | Interactive effect of *private* and *female*, i.e. one if employee is employed in the private sector and is female, zero otherwise |
| Inter_female_$X$ | Interactive effect of *female* and the set of regresors $X$;<br>*Inter_female_Exper–Inter_female_Intermed_Prof* |
| Inter_year_$X$ | Interactive effect of *year* and the set of regresors $X$;<br>*Inter_year_Exper–Inter_year_Intermed_Prof* |
| Inter_femyear_$X$ | Interactive effect of *femyear* and the set of regresors $X$;<br>*Inter_femyear_Exper–Inter_femyear_Intermed_Prof* |
| Inter_private_$X$ | Interactive effect of *private* and the set of regresors $X$;<br>*Inter_private_Exper–Inter_private_Intermed_Prof* |
| Inter_fempriv_$X$ | Interactive effect of *fempriv* and the set of regresors $X$;<br>*Inter_fempriv_Exper–Inter_fempriv_Intermed_Prof* |
| | Labor Market Presence |
| Exper | Number of years of prior work experience |
| Exper2 | *Exper* squared |
| Tenure | Number of years worked for current employer |

Educational Attainment

| | |
|---|---|
| Schooling | Number of years of schooling completed |

Job Characteristics

| | |
|---|---|
| Work_Climate | Individual's level of statisfaction with the working climate at the individual's current job $\in (0, 4)$, where 4 is the highest level of satisfaction and 0 the lowest |
| Work_Stab | Individual's level of statisfaction with the stability of the individual's current job $\in (0, 4)$, where 4 is the highest level of satisfaction and 0 the lowest |
| Work_Time | Individual's level of statisfaction with the working time at the individual's current job, where 4 is the highest level of satisfaction and 0 the lowest |
| Work_Task | Individual's level of statisfaction with the tasks at the individual's current job $\in (0, 4)$, where 4 is the highest level of satisfaction and 0 the lowest |
| Contract_Type | One if the individual holds an unlimited contract, zero otherwise |

Demographic Background

| | |
|---|---|
| Italian | One if individual is Italian, zero otherwise |
| Homeowner | One if individual owns a house (including houses financed by bank loans), zero otherwise |
| North | One if the individual lives and works in the North of Italy, zero otherwise |
| Centre | One if the individual lives and works in the Centre of Italy, zero otherwise |

Family Background

| | |
|---|---|
| Married | One if individual is married, zero otherwise |
| Educ_Moth_Uni | One if mother's education is equal to *Laurea*, i.e. mother holds a university degree, zero otherwise |
| Educ_Fath_Uni | One if father's education is equal to *Laurea*, i.e. father holds a university degree, zero otherwise |

Industry and Occupations

| | |
|---|---|
| Sec_Ind | One if individual is engaged in the industrial sector, zero otherwise |
| Sec_Tour | One if individual is engaged in tourism, zero otherwise |
| Sec_Trans | One if individual is engaged in transport, zero otherwise |

| | |
|---|---|
| Sec_Comm | One if individual is engaged in communication, zero otherwise |
| Sec_Fina | One if individual is engaged in financial sector, zero otherwise |
| Sec_Serv | One if individual is engaged in firm services, zero otherwise |
| Sec_PA | One if individual is engaged in the public administration, zero otherwise |
| Sec_Heal | One if individual is engaged in health, zero otherwise |
| Sec_Prof | One if individual is engaged in science and other professional activities, zero otherwise |
| Manager | One if individual executes intellectual professions; scientific and highly specialized occupations, zero otherwise |
| Intermediate_Prof | One if individual executes intermediary positions in commercial, technical or administrative sectors, health services and technicians, zero otherwise |

# F Regression Output from the Full Specification

Table F.1: OLS Estimates of Log Hourly Wages – Case 1, Full Specification

| | (1) |
|---|---|
| Variables | |
| femyear | -0.051 |
| | (0.185) |
| female | -0.148 |
| | (0.152) |
| year | -0.010 |
| | (0.130) |
| Exper | 0.019*** |
| | (0.002) |
| Exper2 | -0.000*** |
| | (0.000) |
| Tenure | 0.004*** |
| | (0.001) |
| Schooling | 0.038*** |
| | (0.005) |
| Contract_Type | 0.080*** |
| | (0.023) |
| Work_Climate | 0.001 |
| | (0.008) |
| Work_Time | 0.009 |
| | (0.007) |
| Work_Task | -0.002 |
| | (0.008) |
| Work_Stab | -0.024*** |
| | (0.007) |
| North | 0.060*** |
| | (0.014) |
| Centre | 0.038** |
| | (0.015) |
| Italian | 0.004 |

| | |
|---|---|
| | (0.065) |
| Homeowner | -0.006 |
| | (0.018) |
| Married | 0.062*** |
| | (0.014) |
| Educ_Moth_Uni | -0.011 |
| | (0.033) |
| Educ_Fath_Uni | 0.069*** |
| | (0.027) |
| Manager | 0.136*** |
| | (0.020) |
| Intermed_Prof | 0.035*** |
| | (0.013) |
| Constant | 1.163*** |
| | (0.110) |
| Industrial Dummies | Yes |
| Interaction Terms | Yes |
| Observations | 17,918 |
| R-squared | 0.291 |

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

Table F.2: OLS Estimates of Log Hourly Wages – Case 2, Full Specification

| | (1) |
|---|---|
| Variables | |
| fempriv | -0.724** |
| | (0.289) |
| female | 0.278 |
| | (0.196) |
| private | 0.309 |
| | (0.205) |
| Exper | 0.019*** |
| | (0.004) |
| Exper2 | -0.000*** |
| | (0.000) |
| Tenure | 0.002 |
| | (0.001) |
| Schooling | 0.055*** |
| | (0.007) |
| Contract_Type | 0.182*** |
| | (0.057) |
| Work_Climate | 0.014 |
| | (0.012) |
| Work_Time | -0.001 |
| | (0.013) |
| Work_Task | -0.004 |
| | (0.014) |
| Work_Stab | -0.017 |

| | |
|---|---|
| | (0.013) |
| North | 0.049** |
| | (0.023) |
| Centre | 0.072*** |
| | (0.023) |
| Italian | -0.177*** |
| | (0.063) |
| Homeowner | 0.050 |
| | (0.032) |
| Married | 0.031 |
| | (0.026) |
| Educ_Moth_Uni | 0.074 |
| | (0.058) |
| Educ_Fath_Uni | 0.043 |
| | (0.046) |
| Manager | 0.118*** |
| | (0.032) |
| Intermed_Prof | -0.015 |
| | (0.024) |
| Constant | 1.046*** |
| | (0.147) |
| Industrial Dummies | Yes |
| Interaction Terms | Yes |
| Observations | 8,423 |
| R-squared | 0.236 |

Robust standard errors in parentheses
$***p < 0.01$, $**p < 0.05$, $*p < 0.1$

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