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Swinging female labor demand – How the public sector influences gender wage gaps in Europe

Paul Ramskogler
Aleksandra Riedl
Florian Schoiswohl

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Swinging female labor demand – How the public sector influences gender wage gaps in Europe

Paul Ramskogler[1], Aleksandra Riedl[2], Florian Schoiswohl[3]

Abstract

We incorporate an economy's sectoral structure into a standard theoretical framework to explain the influence of relative demand and supply effects on the gender wage gap. Using micro data covering 30 European countries over the 2003-2013 period, we construct a unique macro panel of gender wage gaps. We demonstrate that the public sector has causally determined half of the decrease in the gender wage gap over the period, thus acting as a 'swing demander' for female labor. We further prove that it is exclusively demand factors and not composition effects that are driving this result.

JEL classification: J3, J5, J7;
Keywords: Female net supply, Labor demand in the public sector, Remuneration effect, Wage inequality

[1]Oesterreichische Nationalbank, Otto-Wagner Platz 3, 1090 Vienna, Austria; E-Mail: paul.ramskogler@oenb.at.

[2]Corresponding author. Oesterreichische Nationalbank, Otto-Wagner Platz 3, 1090 Vienna, Austria; E-Mail: aleksandra.riedl@oenb.at.

[3]Bundesministerium für Finanzen, Johannesgasse 5, 1010 Vienna, Austria; E-Mail: florian.schoiswohl@bmf.gv.at

1 Introduction

The international oil market has very specific characteristics. It is dominated by a few large producers, though only one producer can alter supply by such significant quantities that she is able to influence prices substantially. This is the 'swing producer' Saudi Aramco. But of course, international oil markets are hardly comparable to European labor markets. Or are they? In this paper we demonstrate that indeed there is a 'swing demander' for female labor. Concretely, we show that in Europe the public sector is key in determining relative incomes being responsible for half of the significant reduction of the gender wage gap during the period 2003 to 2013.

One of the main motivations of our paper relates to the fact that in Europe a peculiar role accrues to the public sector exhibiting an extremely high female labor intensity (68%) and further being substantially larger than in other advanced economies (employing 23% of the labor force). The starting point of our considerations though, is the observation that in the post-war period in most industrialized economies female employment has increased and the gender wage gap has experienced a trend decline; albeit not a steady one (Olivetti and Petrongolo, 2016). This co-movement of quantities and prices into the same direction can be taken as materialization of a positive shift in demand.

Several explanations have been invoked for this development. There has been a shift from routine to non-routine and from physical to non-physical tasks (Beaudry and Lewis, 2014). Linked to this are the findings of Olivetti and Petrongolo (2014, 2016) who demonstrate that the bulk of the increase in female labor demand can be explained by the secular increase of the service sector in which female employment tends to increase far more rapidly than that of men (Olivetti, 2013). The fact that it is particularly women who specialize in service professions can possibly be explained by models based on tastes for discrimination (Blau and Kahn, 2016) and also by the social (gender) identity model that points at social norms of appropriate and inappropriate activities for women as well as child-rearing practices and psychological factors (Bertrand, 2011). Accordingly, women typically perform the bulk of home production, which goes together with a negative correlation between home hours and earnings across spouses (Albanesi and Olivetti, 2009). This development again is likely linked to the finding that a higher degree of substitution between market and non-market services can account for a large amount of the increase in female labor market participation in the US (Akbulut, 2011). Crucial for our purposes is the following implication: the elasticity of substitution between men and women across sectors is far from infinity and changes in the sectoral composition thus can causally drive the gender wage gap.

In this regard we should further note that the experience and education of the female workforce has largely caught up with that of men yielding a substantial decline in that part of the overall gender wage gap that can be explained by individual-specific factors. As the relevance of these effects declined, the relative importance of the remaining part of the gender wage gap increased steadily.[1] As a result the determinants of the 'corrected' gender wage gap (assuming homogenized human capital endowments in terms of experience and education) have become increasingly important. In Europe in some cases the 'corrected' gender wage gap even exceeds the uncorrected suggesting that female characteristics have become superior to those of men (Christofides et al., 2013a) Yet, – given sectoral segregation – it is exactly this part of the gender wage gap that reflects the effects of changes in the sectoral composition of the economy.

These findings decisively affect the choice of our research design. As we are interested in explaining the corrected gender wage gap, we start out estimating that part of the gender wage gap that is not linked to individual specific factors, like skills or experience. This results in a consistent and unique macro-panel of corrected gender wage gaps for 30 European countries in the period 2003 to 2013. Based on this data, we are able to analyze the immediate and direct effects of a change in female labor demand on relative wages over time. By quantifying the contribution of the public sector in shaping female labor demand, we can finally determine the role of the public sector in driving the gender wage gap in Europe. It is noteworthy that our setting is suitable to verify causality of our results and to rule out potential composition effects (i.e., a mechanical compression of the wage gap due to an

[1] See for instance Gallen et al. (2018), Blau and Kahn (2016) or Olivetti and Petrongolo (2016) for general reviews.

increasing weight of a more egalitarian sector).

Our results can directly be linked to the findings of Olivetti and Petrongolo (2014, 2016) on the interdependency between industry structure and gender outcomes and the decisive role of the service sector played in this regard. However, apart from applying a completely different methodology, we are the first to demonstrate that – within the European service industry – it is specifically the publicly influenced sector that is driving the gender wage gap. This also links us to the results of Piazzalunga and Di Tommaso (2019) published in this journal who show that pay freezes in the public sector had been responsible for the increase in the Italian gender wage gap in the period of 2009-2011. In contrast to their analysis, we have to rely on a broader concept of the public sector, i.e. typically highly publicly influenced sectors of the economy, to obtain a homogeneous definition across European countries.[2] This, however, allows us to show that public employment policies cannot only have an effect by freezing wages but by shifting equilibrium demand for female labor. These demand shifts have not only an immediate impact on relative wages in the public sector but spill-over to the private sector and hence influence the overall gender wage gap in the economy.

This paper is structured as follows. We start by setting out a simple accounting framework (section 2). This will serve as a basis for our empirical setting. We will then devote some space to the question how we estimate 'corrected' gender wage gaps and how we control for potential composition effects in order to isolate pure demand effects (section 3). Subsequently, we will introduce the empirical model to analyze the impact of demand and supply on the gender wage gap (section 4). This includes a discussion on the measurement of demand and supply and, more importantly, a sub-section where we quantify the contribution of the public sector in shaping female labor demand. Finally, we will present the results of our empirical model and their robustness (section 5) and conclude by pointing at some policy implications (section 6).

[2] Our definition of the publicly influenced sector is based on the NACE classification and comprises the sectors public administration and defence, education, human health and social work activities. The definition in Piazzalunga and Di Tommaso (2019) is based on a survey question that specifically asks the respondent if he/she is employed in the public or in the private sector (only available for Italy).

2 Accounting framework

When investigating the determinants of sectoral demand, we build on a reduced form of the model derived in Olivetti and Petrongolo (2014). This allows us to remain relatively brief. The starting point is a standard production function with constant elasticity of substitution (CES) between inputs over the entire production horizon, simplified by focusing solely on the labor force. We hereby diversify into male and female labor supply:

$$Q_s = [\beta_s L_{fs}^{\frac{\sigma-1}{\sigma}} + (1-\beta_s) L_{ms}^{\frac{\sigma-1}{\sigma}}]^{\frac{\sigma}{\sigma-1}} \theta_s \quad (1)$$

The quantity of output supplied by a respective sector s is denoted as Q_s and f and m are indexes for females and males, respectively. Further, σ denotes the elasticity of substitution between male and female labor. Given that men and women can be considered to constitute imperfect substitutes on labor markets we have $0 < \sigma < \infty$, β defines the share of activities performed in a respective sector by women and θ denotes demand.

Assuming perfect labor mobility across sectors (i.e. $W_{fs} = W_f$) and that production factors receive marginal productivity wages yields the demand-determined gender wage gap:

$$W_D : ln\frac{W_f}{W_m} = \frac{1}{\sigma} ln\frac{\Sigma_s(\beta_s^\sigma Q_s)}{\Sigma_s(1-\beta_s)^\sigma Q_s} - \frac{1}{\sigma} ln\frac{L_f}{L_m} = \frac{1}{\sigma}[ln\frac{\Sigma_s(\beta_s^\sigma Q_s)}{\Sigma_s(1-\beta_s)^\sigma Q_s} - ln\frac{L_f}{L_m}] \quad (2)$$

The framework closes with a simple labor supply function. Wages are set as a function of the elasticity of the labor supply $\eta > 0$ and an autonomous labor supply shift factor λ_s:

$$L_f^{\frac{1}{\eta}} = W_f \lambda_f \quad (3)$$

Calculating gaps, taking logs, setting $ln\frac{\lambda_f}{\lambda_m} = \lambda$ and further reformulating leaves us with the supply-side gender wage gap:

$$W_S : ln\frac{W_f}{W_m} = \frac{1}{\eta}ln\frac{L_f}{L_m} - \lambda \tag{4}$$

Equilibrium gender wage gaps are identified by the intersection of the supply and demand functions. Let us abbreviate the expression $ln\frac{\Sigma_s(\beta_s^\sigma Q_s)}{\Sigma_s(1-\beta_s)^\sigma Q_s}$ from (2) as βQ, which can be read as the relative intensity of production performed by women. Deriving the equilibrium gender employment gap, resubstituting, inverting and taking differences over time allows us to investigate the determinants of changes in the gender wage gap:

$$\Delta GAP = -(\frac{1}{\sigma + \eta})(\Delta\beta Q + \beta\Delta Q) + (\frac{\eta}{\sigma + \eta})\Delta\lambda \tag{5}$$

Thus there are demand factors (i.e. $\Delta\beta Q + \beta\Delta Q$) as well as supply factors (i.e. $\Delta\lambda$) that potentially explain changes in the gender wage gap over time. Of these factors, demand is determined in a twofold way. It might be driven by a change of relative inputs $\Delta\beta Q$, the within-industry effect. Alternatively demand can be determined by changes in relative sectoral demand $\beta\Delta Q$, the between-industry component.

Given our main interest, we focus exclusively on the between-industry effect here; in our case $\beta\Delta Q$. The between-industry dimension hereby captures general equilibrium effects of changes in the relative importance of certain sectors, including reactions to cyclical factors. As it is short to medium run perspective that we are primarily interested in we will concentrate on between-industry effects $\beta\Delta Q$ in the following. This is equivalent with assuming $\Delta\beta Q = 0$. The bias accruing out of this choice usually is regarded to be rather small and manageable (Katz and Murphy, 1992). Further, measuring labor demand by actual quantities, we proxy demand with employment L. This yields:

$$\Delta GAP = -(\frac{1}{\sigma + \eta})\beta\Delta L + (\frac{\eta}{\sigma + \eta})\Delta\lambda \tag{6}$$

In terms of the interpretation of $\beta\Delta L$ provided above, this means that we will be able to approximate our simplified version of between-industry changes by the change in relative employment of the respective industries.

3 Estimating the remuneration gap

Our major interest – as argued above – is on the immediate and direct demand effect of a change in the sectoral composition of the economy. More precisely, we aim to analyze whether one sector (the public sector in our case) has the power to influence equilibrium prices in the entire economy. This however, necessitates isolating that part of the gender pay gap that is not linked to individual specific effects. This is what we do in this chapter. We disentangle those parts of the gender pay gap that are driven by individual specific factors from those that are determined by other factors amongst which demand and supply figure prominently.

Hereby, we would like to note that we are among the first contributions to provide a macro panel of estimated gender wage gaps reflecting the relative remuneration of men and women. Among the few contributions that focus on a cross-country or time-varying perspective[3], only Blau and Kahn (2003) provide an international comparison of gender wage gaps over time.[4] The main aim of the paper by Blau and Kahn (2003) is to explain international differences in gender pay gaps by focusing on the role of institutions. Therefore, the authors estimate wage differentials on the assumption that men and women have the same average human endowment levels as U.S. men and women across all countries. Hence, their estimation strategy removes international differences in women's relative levels of measured characteristics, but still preserves gender differences in measured characteristics in each country, namely at U.S. levels. By doing so, they are able to isolate the impact of the wage structure on gender pay gaps. In contrast, given our question of interest, we remove all potential differences in

[3]E.g. Weichselbaumer and Winter-Ebmer (2005); Meulders et al. (2006); Arulampalam et al. (2007); Beblo et al. (2003); Brainerd (2000); Blau and Kahn (1996, 2003); Boeheim et al. (2007); Christofides et al. (2013b); Nicodemo (2009); Olivetti and Petrongolo (2008, 2014).

[4]Blau and Kahn (2003) construct a sample covering 15 European and 7 non-European countries in the period 1985 to 1994, where each country appears 4.5 times on average during this period yielding a panel of 100 observations.

men's and women's capital endowments and other job-related variables in order to isolate potential differences in prices, that is the remuneration of these endowments.

To tackle this issue, we rely on data from the EU-Survey of Income and Living Conditions (EU-SILC), which is frequently used to analyze income disparities along various dimensions (see e.g. Pittau et al. (2015) and Beblo and Knaus (2001)). EU-SILC is a household-survey that is representative for the countries' population. The data set comprises a wide range of individual and household characteristics including annual gross income and has been conducted at an annual frequency since 2004. This allows us to construct a consistent panel data set that comprehensively covers the European region for a considerable amount of time. Given certain sample restrictions, we are able to estimate gender wage gaps based on 283 individual country-year data sets covering 30 European countries over eleven years. For an overview of the participating countries and of specific restrictions regarding certain waves of the survey see table A1 in the appendix.

3.1 Sample selection

Before we move on to estimate the remuneration gap, we want to briefly elaborate on the issue of sample selection. In particular, the selection of an 'employed' sub-sample of the population, directly leads to the question if differences in the probability of selection between men and women significantly influence the estimation of average wage differentials. Prominently, Olivetti and Petrongolo (2008) have shown the importance of correcting for non-random sample selection in cross-country studies. In their study, uncorrected gaps are generally underestimated in countries with low female participation due to the fact that on average in these economies only high-wage females are in employment; a feature which is also found in Christofides et al. (2013b) for a cross-section of European countries, and reported in an overview of Beblo et al. (2003), in which many single-country studies report significant selection effects.

While it seems very reasonable to control for selection in order to compare gender wage gaps across countries[5], we opt not to correct for selection in our baseline model for two reasons. First, the focus of our paper is to explain changes in gender wage gaps over time rather than levels across countries. Hence, cross-country differences in gender pay gaps that are attributable to differences in employment gaps across countries (as shown by Olivetti and Petrongolo (2008)) are potentially wiped out in our setting. This is corroborated by the fact that the correlation coefficient between the remuneration gap and the employment gap is -0.51 in our sample, which is remarkably close to the figure found by Olivetti and Petrongolo (2008). However, if we take first differences of both variables we end up with a correlation coefficient of -0.15. Second, we explicitly model the gender wage gap as a function of (relative) female labor supply, i.e. we account for the fact that wages are only observed for those women that are employed (see also Blau and Kahn (2003), p. 120 for a related argument). However, we will compare our baseline results with a model specification that controls for selection[6] as this might unveil an interesting insight into the mechanism by which an increase in female labor supply can affect the gender wage gap. More precisely, an increase in labor supply can either lower average female wages (i) by changing the composition of low- and high-wage women in the economy, e.g. as more low-wage females enter the labor market or (ii) by lowering equilibrium wages (pure supply effect) induced by a right-ward shift of the labor supply curve (or both). Comparing our baseline results with a specification that controls for selection by assigning higher weights to underrepresented women can isolate the latter effect from the former.

3.2 Disentangling the total wage gap

We now proceed to estimate gender wage gaps for each country-year pair. We will start by calculating the total gender wage gap, i.e. the difference between mean log-hourly wages of

[5]Especially, European countries show substantial differences in employment rates (sample participation) of males and females. While in Nordic countries both participation-rates are above 90% for both men and women, in Southern European countries they are around 80% for males and 45% for females.

[6]More specifically, we use a two-stage Heckman (1979) selection model to account for non-random selection into employment. We estimate the first-stage probit-equation for females only, i.e. assuming random selection for males. The identification of the first-stage is based on a dummy if the woman is married, a variable in logs comprising the income of the spouse (if any) and other financial household income, and a dummy indicating the presence of young children in the woman's household.

males and females. Within each micro data file, we restrict our baseline sample to employees aged 16-64 years. We exclude the self-employed from the sample as income information for this group is measured only imprecisely, which might introduce a bias in our wage gap estimates. However, as a robustness check we will include the group of self-employed and present the respective results in section 5.3. Likewise, we will investigate the effect of restricting our sample to the group of employees aged 26-55 (see e.g., Christofides et al. (2013b)). As we are mainly interested in the relative remuneration between men and women, we have to single out the part of the total gender wage gap that reflects differences in endowments. Fortunately, the EU-SILC data allows us to estimate average wage differentials while correcting for individual-specific differences in characteristics. We thus can derive that part of the gender wage gap that is not determined by individual factors – commonly referred to as the remuneration effect, which is separated from the endowment effect. This is typically done by estimating separate wage equations for males and females with $w_{it,m/f}$ indicating the log-gross hourly wage of males and females (m/f) for country i of year t (skipping the subscript of the individuals). The average wage-differential (i.e. total gender wage gap) $\Delta\bar{w}_{it} = \bar{w}_{it,m} - \bar{w}_{it,f}$ can be decomposed employing a framework introduced by Oaxaca (1973) & Blinder (1973) and prominently extended by Oaxaca and Ransom (1994):

$$\Delta\bar{w}_{it} = \underbrace{(\hat{\gamma}_{it,m} - \gamma^*_{it})\bar{X}_{it,m} + (\gamma^*_{it} - \hat{\gamma}_{it,f})\bar{X}_{it,f}}_{\text{remuneration effect (GAP)}} + \underbrace{\gamma^*_{it}(\bar{X}_{it,m} - \bar{X}_{it,f})}_{\text{endowment effect}} \tag{7}$$

where γ^*_{it} is referred to as the non-discriminatory wage structure and $\hat{\gamma}_{it,m/f}$ as the coefficient from the separate wage equations that are estimated by OLS.[7]

The independent variable, $w_{it,m/f}$, is calculated from annual gross-income ("employee cash or near cash income") divided by total hours worked within the income year (including not only the main but any additional jobs). Note that all countries except the United Kingdom

[7]Note that γ^*_{it} can be defined as a weighted coefficient $\gamma^*_{it} = \Omega_{it}\gamma_{it,m} + (I - \Omega_{it})\gamma_{it,f}$, where γ^*_{it} can be set to the male-wage structure only $(\Omega = I)$, to the female wage-structure only $(\Omega = 0)$, or to the half of each as in Reimers (1983) $(\Omega = .5I)$. Oaxaca and Ransom (1994) and Neumark (1988) derived Ω from a pooled wage-equation. We opt for the first alternative, i.e. $\Omega = I$. Note however, that choosing the other alternatives do not qualitatively change the estimation results presented in the next section.

provide income data for the calendar year preceding the interview, while the UK publishes income statements for the current year[8]. Therefore, the gaps estimated from survey-year t refer to year $t-1$ (except for the UK).

To ensure high comparability across countries and over time, the set of explanatory variables $X_{it,m/f}$ is comprised of a minimum selection of important variables describing major compositional characteristics of the aggregated work-force. The set includes education dummies, a part-time-dummy (Erosa et al., 2016), as well as work-experience and its square.[9] Finally, estimating equation (7) gives us 283 country-year observations of both, the remuneration and the endowment effect of the total gender wage gap. The remuneration effect will serve as our baseline gender wage gap specification throughout the paper. From now on we will use remuneration effect and gender wage gap interchangeably. Note that we do not control for occupation in our baseline estimation as it is unclear whether a higher pay in male-dominated occupations is a result of women allocating themselves into lower-payed professions or whether it represents discrimination. However, as a robustness check we will re-calculate the remuneration gap by including occupational controls in order to see whether this affects our estimation results. We will discuss the respective results in the robustness section.[10]

[8] Ireland switched from a moving 12-month period to a fixed-calendar-year period in 2009. However, it cannot be exactly determined when the interview took place during the survey year, so that the income period cannot be located exactly before 2009. We therefore treat the income data for the years 2004-2008 as data from the period 2003-2007.

[9] We lump the pre-primary, primary, and the lower secondary group together. As a result we get three education categories: Primary/Lower Secondary, Upper Secondary (incl. post-secondary), and Tertiary. Note also, that in some countries work-experience is not available. In this case we add age and age-squared to the list of controls to proxy work experience (see also table A1).

[10] We use the ISCO-08 classification, which groups occupations in ten categories. Due to some missing observations in individual countries we form eight categories to yield a consistent set of occupational variables across countries and years. Note that we loose two country-year observations compared to our baseline specification. The occupation categories are (1) managers, (2) professionals, (3) technicians ans associate professionals, (4) clerical support workers, (5) services and sales workers, (6) skilled Agricultural, Forestry and Fishery Workers, craft and related trades workers, (7) plant and machine operators and assemblers, (8) elementary occupations. See also http://www.ilo.org/public/english/bureau/stat/isco/isco08/index.htm

3.3 Sectoral differences and the composition of the gender wage gap

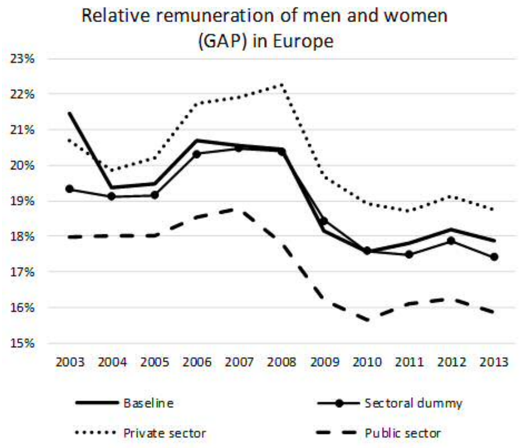


Figure 1: Remuneration gap, unweighted averages of 30 European countries

In this sub-section we discuss how to control for potential composition effects in our empirical model in order to isolate pure demand effects. We start by plotting the development of the overall gender wage gap in Europe for the years 2003 to 2013, depicted by the bold line in figure 1. We observe a decrease of the gender wage gap by roughly 3.5 percentage points during the observation period. In addition, we calculate the remuneration effect for the private and the public sector, i.e. we separately estimate equation (7) based on men's and women's wages in the respective sectors. The public sector in our definition hereby consists of public administration and defense, education, health and social work; sectors that in Europe can be reasonably assumed to be primarily publicly influenced. The results, again unweighted averages, are depicted by the dashed lines, which show that the remuneration gap of men and women is on average considerably smaller in the public sector. Similar results have been observed in the literature, e.g. by Stewart (2014) for the UK or by Cheng (2005) for Canada. Yet, this finding is not surprising given that the total (unadjusted) gender wag gap is lower in the public sector in most of the European countries for which data are available[11]. One reason might be that in most European countries, public sector employees are protected by collective pay agreements and other similar contracts establishing pay structures.

Interestingly, our estimation results also indicate that the difference in the gender wage gap across the two sectors seems to be quite persistent over time and their evolution further coincides with that of the total gender wage gap. As we will show in the next section, the importance of the public sector increased over time. Given the smaller gender wage gap in

[11] Source: Eurostat, http://ec.europa.eu/eurostat/statistics-explained/index.php/Gender_pay_gap_statistics#By_economic_activity

this sector, it is thus not obvious which part of any given reduction in the total gender wage gap is driven by changes of relative demand (a pure demand effect) and which is determined by a mechanical compression driven by the fact that the public sector has become relatively more important in the calculation of the total gender wage gap (composition effect).

In order to overcome this problem we re-calculate our baseline estimation of the remuneration effect by additionally controlling for potential sectoral effects.[12] More precisely, we include a public sector dummy in the wage equations of men and women in order to net out that part of the wage that is attributable to the sector. The result is depicted by the solid line with dots and again represents unweighted averages of all countries. What stands out is the quite high overlap of this series with the baseline gender wage gap (with the exception of 2003 data). This is an important first indication that the decrease in the remuneration effect in Europe is not the result of a composition effect. Of course, however, the graph masks developments across individual countries and the small but clearly visible deviations between both series might still reflect compositional effects. Therefore, we will consider both series in the panel estimation model in order to see whether labor demand effects change when compositional effects are wiped out in the estimation of remuneration effects.

In table A2 we present descriptive statistics of the total gender wage gap, the baseline remuneration gap and several other specifications of the remuneration gap including the one where selection is controlled for.

4 The empirical model

Given our accounting framework introduced in section 2 and the empirical specification of the gender wage gap in the previous section, we are now in the position to derive the empirical model. This will allow us to estimate the impact of supply and demand factors on the gender wage gap in Europe. In line with equation (6) the change in the gender wage gap $\Delta GAP_{i,t}$

[12] Be aware that the total GAP cannot be decomposed in a linear fashion, i.e. it is not simply a weighted average of the public and the private gender wage gap. If this were the case, we could simply isolate the demand effect by holding the weight constant over time.

is a linear function of changes in relative demand $\Delta D_{i,t}$ and relative supply $\Delta S_{i,t}$ of female labor. Hence, our empirical specification can be written as

$$\Delta GAP_{i,t} = \alpha_1 + \alpha_2 \Delta D_{i,t} + \alpha_3 \Delta S_{i,t} + \epsilon_{i,t} \tag{8}$$

Based on this equation, we will verify the derived hypothesis for 30 European countries in the period from 2003 to 2013.[13] In particular, we expect demand to impact the gender wage gap negatively ($\alpha_2 < 0$) while supply should have a positive influence ($\alpha_3 > 0$). Note that, we do not include unobservable country fixed effects into our empirical model, as we are estimating a first difference equation, where time constant variables are wiped out by construction. Hence, we will employ the pooled OLS estimator to estimate equation (8), which avoids a potential bias due to time invariant omitted variables (Baltagi, 2005). Also, we do not include time fixed effects as they are not significant in our setting. We will demonstrate this in the course of our robustness analysis in section 5.3.

By implementing the OLS estimator for our baseline specification, we assume that gender wage gaps do not influence relative labor demand and supply. Although we think that reverse causality is not a problem in our context[14], we will explicitly address this issue in the next section, where we will employ a GMM estimator to control for potential endogeneity issues (Blundell and Bond, 1998).

4.1 Measuring demand and supply effects

In order to estimate equation (8) we need to first specify the two main determinants of the gender wage gap, namely between-industry changes in relative labor demand $D_{i,t}$ and relative supply $S_{i,t}$. With regard to the first mentioned variable, we follow Katz and Murphy (1992) and Blau and Kahn (2003) who construct an index that indicates whether a change

[13] Since our estimates of gender wage gaps are not available since 2003 for all countries, our panel data set is comprised of 283 observations.

[14] In order to have reverse causality, a relative reduction of female remuneration would have to yield higher demand for products from female sectors via lower relative prices. This situation would require a high elasticity of substitution between the respective products (and a fixed mark-up). However, at the rather high level of sectoral aggregation that we are observing, the elasticity of substitution is likely to be extremely low, so that we can rule out this situation at least for the within-country perspective.

in the composition of output by industry is favorable for women. The demand index $D_{i,t}$ is composed of the relative gender intensity in a respective industry at a specific point in time and the change in employment of that industry over time with respect to a base year $t_0 = 2003$. Note that Blau and Kahn (2003) measure the change in sectoral employment relative to a base year *and* a specific country (the US) as they aim to explain international differences of gender wage gaps across countries. However, given that we want to specifically exploit changes of the gender wage gap over time (i.e., the within-country dimension of the data) we fix employment to a common base year only. Notating employment as L and the fraction of female employees in sector s as b, the demand index of country i in year t can be written as:

$$D_{i,t} = \Sigma_s b_{i,s} \Delta L_{i,s,t} = \Sigma_s b_{i,s} (\frac{L_{i,s,t}}{L_{i,t}} / \frac{L_{i,s,t0}}{L_{i,t0}} - 1) \tag{9}$$

The weighting factor $b_{i,s}$ represents the relative abundance of female employees in the respective sectors at a specific point in time, i.e.

$$b_{i,s} = \frac{L^f_{i,s}}{L_i} \tag{10}$$

where L_i is the sum of male (m) and female (f) employment, i.e. $L_i = L^m_i + L^f_i$.[15] In accordance with Blau and Kahn (2003) we fix gender intensity $b_{i,s}$ to a base year (2003) in order to have an exogenous measure of demand changes that is not influenced by price changes caused by the demand shifts.[16]

In figure 2 we display the shares of female employment in total employment for each sector. In three out of eight sectors (finance, public and arts) the share of female employment is higher compared to men's. However, what clearly stands out is that the share of female employment is the highest in the public sector. Moreover, total employment in the public sector is

[15] Note that, at first glance, our measure deviates from the one used by Blau and Kahn (2003) who proxy the relative gender intensity as $L^f_{i,s}/L^f_i$. However, row-normalisation of this proxy yields the expression given in equation (10), i.e., $b_{i,s}/\Sigma_s(L^f_{i,s,t0}/L_{i,t0}) = L^f_{i,s,t0}/L^f_{i,t0}$.

[16] Note that the choice of the base year does not qualitatively influence our estimation results, as we will show in section 5.3.

among the highest compared to other sectors. Hence, given this first inspection of sectoral segregation on the labor market, it is presumable that the public sector plays a crucial role in determining the relative demand for female labor. We will return to this issue in the next subsection, where we will assess the impact of public sector employment on the demand index.

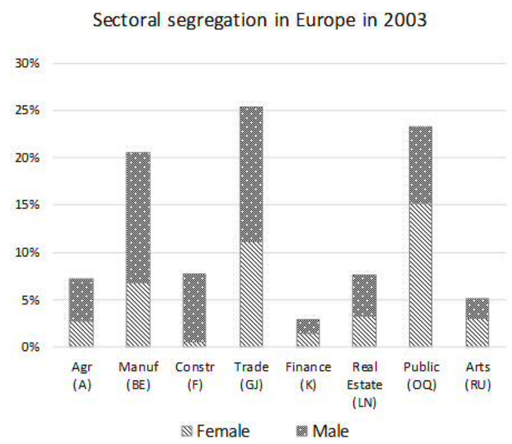


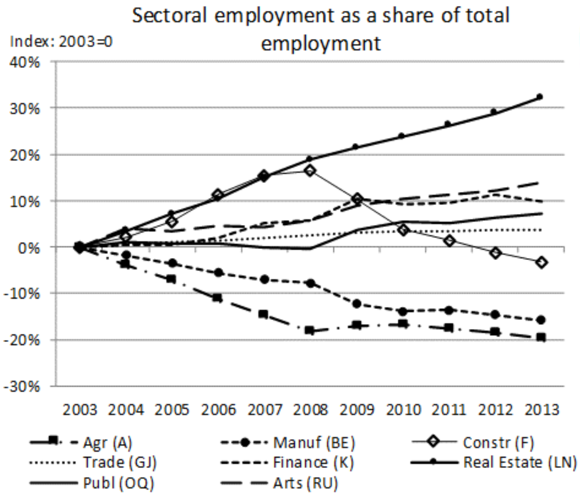
Figure 2: Sectoral employment, unweighted averages of 30 European countries. Source: Eurostat.

In addition to relative demand, we specify a variable reflecting changes in the relative labor supply of women compared to that of men, hereinafter referred to as $S_{i,t}$. It would be ideal to have an indirect measure of a female labor supply shifter. Albanesi and Olivetti (2009) for instance have shown the effect of home production decisions on relative incomes. Thus, time used by females in home production would be an ideal measure for female labor supply. However, in absence of the availability of time series for any such variables we resort to the avenue taken by Blau and Kahn (2003), by measuring female labor supply by the relative abundance of active women in a country's work force. Hereby, λ denotes the fraction of active women (i.e. either employed or job seeking) as a share of the total working age population (i.e. men and women):

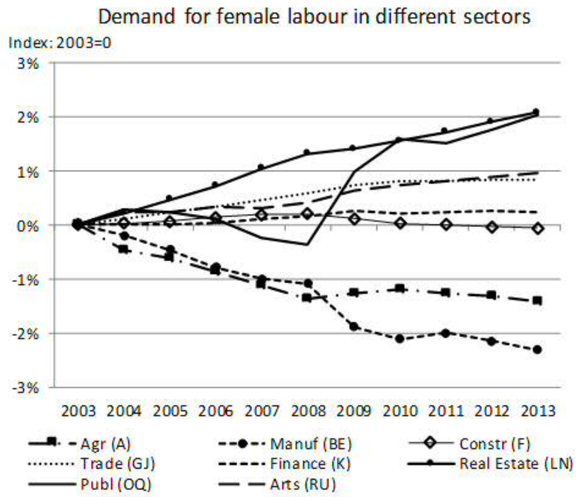
$$S_{i,t} = \frac{\lambda_{i,t}}{\lambda_{t0}} - 1 \tag{11}$$

To calculate $S_{i,t}$ and $D_{i,t}$ we use data from Eurostat, where comprehensive and comparable statistics for all European countries are provided.[17] More precisely, sectoral employment data to calculate the demand index are obtained from national accounts, which cover the whole economy, i.e. all persons engaged in some productive activity. We use the finest

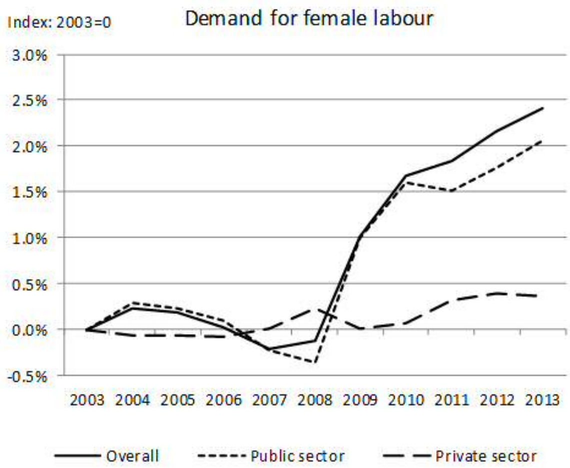
[17] Available at http://ec.europa.eu/eurostat (codes: $lfsa\_egana$ to calculate gender intensity weights, $nama\_10\_a10\_e$ for sectoral employment data, $lfsa\_agan$ and $lfsi\_emp\_a$ to measure relative supply).



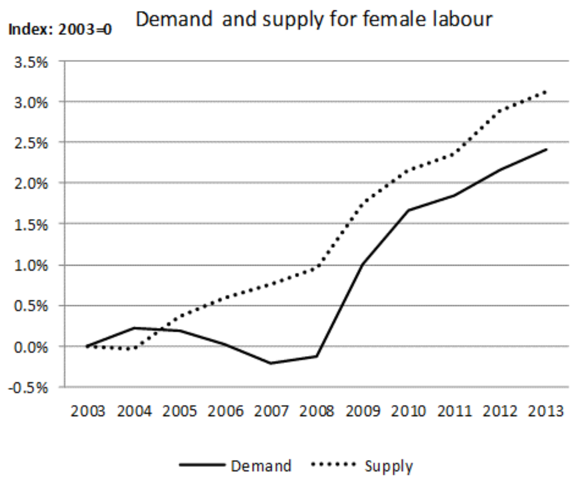
(a) $\Delta \overline{L}_{s,t}$



(b) $\overline{b}_{s,2001} \Delta \overline{L}_{s,t}$



(c) $D_t = \Sigma_s b_{s,2001} \Delta \overline{L}_{s,t}$, D_t^{pub}, D_t^{priv}



(d) D_t, S_t

Figure 3: Demand and supply for female labor in Europe, unweighted averages of 30 European countries

possible sector classification that is available for our country-year panel (NACE Rev.2 sector classification, one-digit), which results in eight sectors.[18] Note that we also use data from the labor force survey (also obtained from Eurostat) to calculate an alternative measure for

[18]The sectors are: Agriculture, forestry and fishing (A), Industry (B-E), Construction (F), Wholesale and retail trade, transport, accommodation, food service activities and information and communication (G-J), Financial and insurance activities (K), Real estate activities, Professional, scientific and technical activities; administrative and support service activities (L-N), Public administration, defence, education, human health and social work activities (O-Q), Arts, entertainment and recreation; other service activities; activities of household and extra-territorial organizations and bodies (R-U). For a detailed description of the statistical classification see http://ec.europa.eu/eurostat/documents/3859598/5902521/KS-RA-07-015-EN.PDF.

the demand index. This allows us to use a finer sector classification (11 instead of eight sectors) but covers only an incomplete subset of total demand, as it includes only workers aged 15-64. We will show in the robustness section 5.3 that using a different data source does not have any significant impact on the estimation results.[19]

In figure 3 we aim to visualize the development of the specified variables in Europe since 2003. Plotting simple unweighted averages of the demand and supply variables across the 30 European countries reveals a quite interesting pattern regarding the role of the public sector for female labor demand. Sub-figure 3a shows the employment trend in eight different industries as a share of total employment, i.e. the part of equation (9) that corresponds to $\Delta L_{i,s,t}$. As expected, employment in the service sector, especially activities included in subcategory LN (i.e., real estate, professional and technical activities as well as administrative and support service activities) registered a relative increase over the observation period, while employment in the manufacturing (BE) and agricultural (A) sector has become less important in relation to total employment. In contrast to these quite substantial employment shifts across sectors over the last decade, the development of public sector employment (OQ) does not really stand out as the share of public sector employment has increased by only 10% over the same time period. However, if we weight the employment share of the respective sectors by gender intensity (i.e., $\overline{b}_s$) we see that the public sector is one of the most important contributors to female labor demand (see sub-figure 3b). More importantly, the overall relative demand for female labor (sum of all sectors), which is depicted in sub-figure 3c (continuous line), is predominantly shaped by the variability of public sector employment. This observation is even more visible when we split the demand index into two parts, such that we can distinguish between female labor demand from the private (D_t^{priv}) and the public sector (D_t^{pub}).[20] Both indexes are added to sub-figure 3c and are marked by dashed lines. Finally, sub-figure 3d presents the development of both relative demand and supply for female labor in Europe. On average, the demand for female labor has increased by 2.4

[19] Note that the labor force survey provides sectoral employment data at the two-digit level as well. However, due to a lot of missing observations, we could not use this data to obtain a finer demand index for our sample.

[20] Note that $D_t^{priv} = D_t - D_t^{pub}$ where D_t^{pub} is just the line 'Publ OQ' drawn in subfigure 3b, i.e. $b_{s,2001} L_{s,t}$ with $s =$OQ.

percentage points, while the supply of female labor has risen by 3.1 percentage points over the entire observation period.

4.2 The impact of public employment on female labor demand

In the previous subsection, we have seen that the development of female labor demand is very much shaped by the variability of public sector demand. But to which extent has employment in the public sector contributed to the rise in female labor demand? Since the next section will present estimates of a 1 percentage point increase in overall relative demand on the gender wage gap, knowing the public sector's employment growth contribution to the development of the overall demand index would finally allow assessing the direct influence stemming from public sector employment on the gender wage gap. Yet, this is not a straightforward task, as the demand index is a non-linear function of eight unknown variables influencing each other.[21] As we are not able to derive an analytical expression for the growth contribution, we perform a counter-factual analysis, quantifying how overall female labor demand would have developed if employment in individual sectors had evolved differently. By comparing the outcomes with the actual development of the overall demand index ΔD, which increased by 2.4 percentage points on average, we can infer the contribution of each sector.

In table 1 we summarize the results from three different exercises. The left column (memo item) presents actual annual employment growth rates of each sector as an unweighted average of all countries' growth rates. Scenario (1) and (2) hereby are executed in order to get a feeling how the demand index reflects different conditions and particularly what the role of the public sector is for the determination of this variable. The counter-factual analysis whose results are displayed at the right side of the table deduces the growth contribution of each sector to the development of overall demand over the estimation period.

In scenario 1 we assume that employment in all sectors grows by 1% annually, except for the observed sector, whose employment growth rate is set to zero. In order to additionally

[21] Note that public sector employment enters each of the non-linear components of the demand function through the total employment term $L_{i,t}$.

eliminate the impact of sector size we further assume that each sector has 1,000 tsd. employees each in 2003. Hence, the numbers depicted in column (1) show the importance of each sector assuming that all remaining sectors are equal with the exception of the level of occupational segregation, i.e. the only influence stems from the gender intensity parameter $b_{i,s}$. Not surprisingly, we see that the highest (negative) impact on the demand index arises if public sector employment were held constant. In this setting, the demand for female labor would decrease by 2 percentage points. This mirrors the fact that the share of female employees is highest in the public sector (see also figure 2). In the second scenario, the sectors are allowed to be of different (actual) size and employment is allowed to grow at its actual rates with the exception of one sector, where the annual growth rate is assumed to be 1 percentage point lower compared to actual employment growth. For example, the number for the public sector in column (2) shows how the demand index would have evolved, if employment in the public sector had grown by 0.1% annually (instead of 1.1%, see memo item).[22] Again, we observe that lower employment growth in this sector impacts most on the demand index. While the actual demand index has increased by 2.4 percentage points, it would have increased by only 1.3 percentage points if public sector employment growth were reduced by 1 percentage points annually.

Finally, the counter-factual exercise assumes that employment in all sectors grows at the actual rate but fixes employment in the observed sector at the initial (2003) level, i.e. a zero employment growth rate is assumed. This is equivalent to reversely deriving the contribution of this sector to the growth of overall relative demand for female labor. Looking at the counter-factual demand indexes of this analysis provides a quite similar picture compared to the second scenario. The highest deviations from actual demand are again recorded in the public sector. If employment in this sector had been constant since 2003 the increase in overall demand of female labor would have reduced to 1.2 percentage points. This simultaneously implies that overall demand would only have increased by 1.2 percentage points. This is displayed in the column on the very right in which we show each sector's individual

[22]Note that all numbers presented here are unweighted averages of all countries. However, the scenarios are calculated on an individual country basis, i.e. if Austria has an annual actual employment rate of 2% in the public sector, we reduce the growth rate to 1%. This applies also for the counter-factual analysis.

Table 1: The impact of employment growth in single sectors on ΔD

	memo item	alt. scenarios		counter-factual analysis	
	average ΔL 2003-2013	scen. (1)	scen. (2)	ΔD if ΔL in the sector is 0	contribution to actual ΔD (=2.4)
Agriculture (A)	-2.2	0.6	2.6	2.0	0.4
Manufacturing (BE)	-1.7	-0.3	2.8	2.0	0.4
Construction (F)	-0.2	1.0	2.7	2.1	0.3
Trade (GJ)	0.8	-1.2	2.6	2.5	-0.1
Finance (K)	1.3	0.8	2.3	2.3	0.1
Real Estate (LN)	3.6	0.5	2.4	2.9	-0.4
Public (OQ)	1.1	-2.0	1.3	1.2	1.2
Arts (RU)	1.8	0.5	2.1	1.9	0.5

The actual demand index (D) grew by 2.4 percentage points. Numbers represent unweighted averages of 30 European countries. The first column shows how employment in individual sectors has evolved over time (average annual growth rates in %). Columns 1 to 2 show female labor demand effects from 2003 to 2013 (in percentage points) under alternative scenarios. Scenario 1 assumes equal sectors' size in 2003 (1,000 tsd. persons) and a growth rate of 1% annual employment growth (ΔL) for each but the observed sector (which has a growth rate of zero). Scenario 2 assumes actual employment growth in all sectors but reduces the annual growth rate of the observed sector by 1 percentage point (compared to actual employment growth). Column 3-4 – the counter-factual – is used to calculate contributions to the growth of demand. It assumes actual employment growth in all sectors but fixes employment in the observed sector at the initial (2003) level.

contribution to growth of demand, the sum of which, obviously, aggregates to 2.4. Hence, we can conclude that the growth contribution stemming from the public sector has accounted for half of the change in the overall demand index throughout the observation period.

Yet, from the counter-factual analysis we can also infer the role of the service sector as a whole in shaping overall labor demand. This can be done by summing up the individual contributions of the relevant sectors (i.e. trade, finance, real estate, public and arts) which results in an overall growth contribution of 1.3 percentage points. Hence, if employment in the service sector had stagnated in the period 2003 to 2013, female labor demand would have increased only by 1.1 percentage points. Obviously, nearly the entire contribution stemming from the service sector can be attributed to the public sector. At this point, it is worthwhile to recall the important insights of Olivetti and Petrongolo (2016) who demonstrated – applying a shift-share analysis already pioneered in Olivetti and Petrongolo (2014) – that the secular increase of the service sector accounts for the major part of the structural increase in female labor demand. According to our analysis we can hence confirm their findings. Moreover, we can add that, at least in the European context looking at a more recent period, it is particularly publicly influenced services – i.e. the public sector – that are responsible for the increase in female labor demand.

5 Results

5.1 Baseline Estimations

Table 2 presents the results of our empirical model introduced in equation (8) of section 4. Columns (1) to (5) show the estimated demand and supply effects for different specifications of the remuneration gap that have been discussed in detail in section 3. In column (1) we start out to show coefficient estimates for our baseline gender wage gap specification. The results reveal a negative impact of female labor demand on the gender wage gap while the impact of supply is positive. Moreover, both coefficient estimates are highly significant and strongly support the hypothesis predicted by our accounting framework. This result supports the empirical analysis of Blau and Kahn (2003) who – though in a different empirical setup – find some indication of female net supply effects.[23] At this point, we can hence conclude that changes in demand and supply of female labor seem to have significantly affected the development of gender wage gaps in European countries over the period 2003 to 2013.

According to the point estimates, a one percentage point increase in the demand for female labor has reduced gender wage gaps by 1.5 percentage points. If we recall that female labor demand has increased by 2.4 percentage points on average throughout the observation period, the gender wage gap would have been reduced by 3.6 percentage points all else equal. More importantly, as we have shown in the previous section, half of the increase in overall female labor demand can be attributed to public sector employment. Hence, if employment in the public sector were constant in all countries over the entire period, the gender wage gap would have been higher by 1.8 percentage points on average. This result is also economically significant, given that the gender wage gap in Europe has decreased by 3.5 percentage points

[23] In particular, Blau and Kahn (2003) analyze the relationship between wage-setting institutions and gender wage gaps across 15 European and 7 non-European countries in the period 1985 to 1994. They find strong evidence that institutions are driving international differences in gender wage gaps. Also, they find some evidence that female net supply matters, at least in specifications where country fixed effects are not controlled for. However, given their different empirical setting (e.g. different measurement of gender wage gaps), we have to stress that a comparison to our results has its limitations.

Table 2: Impact of demand and supply on the gender wage gap

	baseline	different estimations of ΔGap			
	ΔGap	selection	sector dummy	private sector	public sector
	(1)	(2)	(3)	(4)	(5)
ΔD	-1.49***	-1.41***	-1.39***	-1.43***	-1.18***
	(-5.05)	(-3.32)	(-4.97)	(-3.79)	(-3.25)
ΔS	0.73***	0.67*	0.77***	0.65**	0.82***
	(3.59)	(1.86)	(3.69)	(2.38)	(2.65)
Cons	0.00	0.00	0.00	0.00	-0.00
	(0.62)	(0.70)	(0.10)	(0.31)	(-0.02)
Model	OLS	OLS	OLS	OLS	OLS
N	252	252	252	252	252

Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis.

throughout the entire observation period (i.e., from 2003 to 2013). Turning to the supply side of the labor market, a one percentage point increase in female labor supply leads to a rise in the gender wage gap by roughly 0.7 percentage points. Hence, from 2003 to 2013, we would have observed an increase in the remuneration gap by 2.2 percentage points all else equal. Thus, compared to the estimated impact of labor demand, the extent of supply effects seems to be somewhat smaller.

In column (2) we present estimates of demand and supply effects on gender wage gaps considering non-random sample selection. Notably, controlling for selection has hardly any influence on the magnitude of both demand and supply effects. As far as the demand effect is concerned, this is not a surprising result. Yet, the fact that the supply effect is still significant indicates that an increase in labor supply indeed lowers equilibrium wages and therefore leads to a rise in the gender wage gap. However, as indicated by the t-statistic, the significance of the supply effect becomes weaker when sample selection is controlled for. Hence, we cannot fully exclude the possibility that the supply effect estimated in the baseline specification partly captures a composition effect that might arise with an increasing share of low-wage females entering the labor market.

Related to this argument, the estimated demand effect might also be influenced by a potential composition effect. As discussed in section 3.3, a sector with more compressed wage differentials could simply mechanically change an economy's total remuneration gap when it expands disproportionately without inducing demand shifts. Specifications (3) to (5) aim to shed light on this ambiguity. First, in column (3) we re-estimate our model by employing a gender wage gap specification that controls for the composition effect by netting out those part of the estimated wages that stem from the respective sector. To put it differently, we eliminate that part of wages from the gender wage gap calculation, which is determined by the sector in which men and women are employed. Interestingly, this alteration does hardly change our baseline results. The demand effect is reduced by 0.1 percentage points, which is a negligibly small change, while the supply effect is even less affected. The statistical significance of both coefficients remain broadly the same. Overall, this is a first indication that the measured impact of female labor demand is indeed attributable to labor demand shifts rather than to composition effects.

In columns (4) to (5) we want to explain each sector's gender wage gap separately. Column (4) presents demand and supply effects on the private sector gender wage gap, which again are nearly unaffected compared to our baseline specification. On the one hand, this outcome is not surprising given that the private sector is the far larger sector of each country's economy and thus is supposed to drive the baseline results, which cover the whole economy. On the other hand, however, as public sector wages do not enter the gender wage gap calculation in this specification, the demand effect should render insignificant if composition effects were dominant in the baseline model. We do not observe any changes in the estimated demand effect. From that we can deduce that our results are not driven by composition effects, at least not by composition effects determined by the public sector. Obviously, changes in female labor demand that are to a large extent driven by the public sector in our sample seem to spill over to the private sector and to influence relative wages there. Interestingly, the demand effect is less pronounced when demand and supply are regressed on the gender wage gap in the public sector, as shown in column (5). This might be related to the generally lower level of the wage gap in this sector. At the same time however, it should also be pointed out that the observed difference is small and statistically not significant.

5.2 Placing public sector effects in context

In section 4.2 we have shown that employment growth in the public sector has been the largest contributor to female labor demand among the eight sectors entering the demand index. In this section we take a different approach to assess the role of the public sector in determining gender wage gaps. In particular, we split the demand index in sub-indexes, which allows us to investigate the relative importance of public sector demand as compared to demand from other key segments of the economy such as for instance the goods producing sector. By taking this avenue, we cannot isolate the contribution of each sector's employment growth to total female labor demand[24] (in contrast to section 4.2), but we can estimate which of the sub-indexes explains most of the observed variation in gender wage gaps and therefore assess the relative importance of individual sectors in a more indirect way.

Table 3: Demand and supply effects of ΔGAP: standardized (β-) coefficients

	(1)	(2)
ΔS	0.19***	0.19***
	(3.59)	(3.62)
$\Delta D_{private}$	-0.12**	
	(-2.03)	
ΔD_{public}	-0.36***	-0.37***
	(-5.16)	(-5.19)
$\Delta D_{service\ sector\ (without\ public)}$		-0.14*
		(-1.82)
$\Delta D_{goods\ producing\ sector}$		-0.16**
		(-2.18)
Cons	0.00	0.00
	(0.54)	(0.20)
Model	OLS	OLS
N	252	252

Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis.

In the first column, we re-investigate the relative importance between the public and the private sector in order to compare the results with the findings obtained by the counter-factual analysis. Hence, we simply split the demand variable into two sub-components, D_{public} and

[24]Recall that each sector's demand (i.e. the sub-indexes entering the total demand index additively) is measured as the change in the share of employment in sector s to total employment. Hence, no implications can be made to what extent employment growth in one particular sector can influence the gender wage gap.

$D_{private}$. By re-estimating our baseline model with the two different sub-components, we can asses the relative importance of the two sectors based on standardized (beta) coefficients (Johnston and DiNardo, 2007). The estimates in column (1) reflect the effect of a one-standard-deviation change in the respective demand and supply variables on the dependent variable (also measured in standard deviations). Hence, the estimated magnitudes are comparable across variables as the latter are measured by the same metric. The results clearly confirm the prominent role that accrues to the public sector. While a one-standard-deviation increase in private sector demand decreases gender wage gaps by 0.12 standard deviations, public sector demand exerts a three times higher impact on gender wage gaps. This result strongly supports the findings presented so far and implies that the contribution of the public sector, which was assessed to be one-half of the overall demand effect, is not over-estimated.

In the second column, we further split private sector demand into demand from the service and demand from the goods producing sector. Given that female intensity is quite high in service oriented industries and that a substantial portion of the rise in female hours in industrialized economies is associated with the reallocation of labor from manufacturing into services (Olivetti and Petrongolo, 2016), it might be plausible that private sector demand masks important differences across the two industries. According to the estimates in column (2) however there is no difference between the impact of the goods producing sector and the service sector. Moreover, we also observe that the public sector still remains the most relevant determinant. Hence, we can confirm our findings from the counter-factual analysis presented in section 4.2, namely that within the service sector, it is public sector demand that explains most of the variation in gender wage gaps.

At first sight, this result stands in contrast to Olivetti and Petrongolo (2014) who find that the public sector does not determine the cross-country variation in wage and hours gap (but rather the service sector). However, this finding is not necessarily inconsistent with our results. Apart from the different sample and methodology applied[25], Olivetti and

[25] In particular, Olivetti and Petrongolo (2014) exploit the cross-country variation in gender gaps and hours based on a shift-share analysis, where they have to assume a value for the elasticity of substitution between male and female labor. Their empirical findings are based on 15 industrialized countries, including the US and Canada. Gender gaps (in wages and hours) are calculated using micro data in the period from 1994-2001.

Petrongolo (2014) use a much narrower definition of the public sector. In particular, they look at public sector administration and defense only, while the public sector in our definition also includes education, health and social work (see also section 3.3), i.e. sectors that are primarily publicly influenced in Europe. In our sample this difference is very significant as the public administration and defense sector in Europe is neither a female intense sector nor does it account for a considerable part of total employment.[26]

5.3 Robustness analysis

This sub-section discusses the robustness of our main findings, which are summarized in Table 3. The first column (0) is a memorandum item and reports our baseline results including 95%-confidence intervals for the estimated effects (in square brackets). The subsequent columns present robustness checks with respect to alternative gender wage gap estimations (1-3), alternative specifications of the demand variable (4-6) and alternative estimation techniques (7-8). Coefficient estimates that lie inside the confidence interval of the baseline results will be classified as insignificant deviations from the baseline model.

As discussed in section 3, our baseline gender wage gap calculation does not control for men's and women's occupations as it is unclear whether a higher pay in male-dominated occupations is a result of preferences or discrimination (see also Blau and Kahn, 2000). If it were discrimination, we would not want to wipe out differences in wages that arise due to different occupations. However, given this ambiguity, we want to assess whether controlling for occupations makes a difference in our setting. If we look at the descriptive statistics in table A2 in the appendix, we see that controlling for occupations as expected decreases the remuneration gap in Europe on average. Comparing the demand and supply effects in specification (1) to our baseline model (0) reveals that the influence from labor demand on gender wage gaps is somewhat mitigated. Still, the demand effect remains statistically and economically significant. Moreover, the difference in the demand effect between the baseline model (0) and specification (1) is statistically not significant, as the point estimate of -1.18

[26] In Europe, employment in the public administration and defense sector accounts only for 7% of total employment and the share of women working in this sector amounts to only 42%.

Table 4: Impact of demand and supply on the gender wage gap: Robustness checks

	baseline	alternative specifications of ΔGap			alternative specifications of ΔD			different estimation methods	
		occupational controls	age 26-55	incl. self-employed	gender intensity	different data source	crisis impact	time dummies	endog. check
	(0)	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ΔD	-1.49***	-1.18***	-1.59***	-1.31***	-1.15***	-1.07***	-1.62***	-1.21***	-2.30***
	[-2.1,-0.9]	(-4.30)	(-4.67)	(-3.87)	(-4.40)	(-3.80)	(-4.31)	(-4.04)	(-6.27)
ΔS	0.73***	0.70***	0.65***	0.54**	0.66***	0.64***	0.81***	0.70***	0.90**
	[0.5,0.9]	(3.66)	(3.06)	(2.47)	(3.17)	(2.92)	(3.05)	(3.36)	(2.54)
$\Delta D \times Crisis$							0.35		
							(1.13)		
$\Delta S \times Crisis$							-0.16		
							(-0.66)		
Crisis							-0.00		
							(-0.59)		
ΔGAP_{t-1}									-0.36***
									(-4.34)
Cons	0.00	-0.00	0.00	0.00	0.00	0.00	0.00	0.00**	0.00
	(0.62)	(-0.04)	(1.03)	(0.55)	(1.55)	(0.27)	(0.84)	(2.34)	(0.69)
Model	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS	GMM
TimeD	excl.	excl.	excl.	excl.	excl.	excl.	excl.	incl.	excl.
P-val (TimeD)	-	-	-	-	-	-	-	0.66	-
N	252	250	252	252	252	262	252	252	223

Two-tailed significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. T-statistics are reported in parenthesis. Column (1) uses gender wage gap estimates when occupation is controlled for (ISCO classification using 8 major groups). Column (2) restricts the sample to employees aged 26 to 55. Column (3) expands the sample by including self-employed persons. Column (4) varies the demand index by employing female intensity ($b_{i,s}$) of the year 2013 (instead of 2003). Column (5) uses employment data from the labor force survey (as opposed to national accounts) to construct the demand index. Column (6) introduces a crisis dummy (1 for the period 2008-2013) and interaction terms to check for differences in the slope parameters of demand and supply. Column (7) displays results for a specification with time dummies (which are not significant with a p-value 0.66). Column (9) implements the system GMM estimator. The corresponding test statistics (Arrelano-Bond and overidentifying restrictions) confirm that disturbances are auto-correlated of order one but not of order two and that the set of instruments is valid.

still lies in the 95% confidence interval of the demand effect from the baseline specification, which is [-2.1,-0.9]. However, if we were to assume that gender differences in occupations are fully explained by preferences, we would infer that the increase in female labor demand has reduced remuneration gaps by 2.8 percentage points on average over the entire observation period (compared to 3.6 derived from the baseline model).

In the next two columns we vary the sample that is used to calculate gender wage gaps. Specification (2) presents estimation results when remuneration gaps of age group 26-55 are applied (instead of 16 to 65) as especially in the older group category country differences could cause measurement errors (e.g. due to different pension laws). In specification (3) we extend the sample by the group of self-employed (instead of using employees only). Again, we see some slight variation in the magnitude of the coefficients. Notably, however, none of the alterations significantly change our main findings.

Next, we want to see whether the obtained results are sensitive with respect to the weighting factor ($b_{i,s}$) that is calculated for a specific point in time (in our baseline specification for the year 2003). We have seen that the relative abundance of female employees (i.e., the gender intensity) in the respective sectors does hardly change over time when we look at the European sample average (see figure xy). Still, in order to assess the influence from fixing this parameter to a specific base, we re-calculate the demand variable by using the weighting factor for the last year of our observation period. Hence, specification (4) reports the results with gender intensities of 2013. Again, demand and supply effects remain nearly unaffected by this alteration. The magnitude of the demand effect is slightly lower compared to our baseline specification but again the difference is statistically not significant.

In the next specification we show that our results remain qualitatively the same if we use a different data source to calculate the demand variable. As discussed in section 4.1, we have employed data from national accounts to calculate the demand index for our baseline specification as it covers all persons engaged in some productive activity. Here, we use data from the labor force survey, which allows to re-calculate the demand variable based on a somewhat finer sector classification (11 instead of 8 sectors). As shown in specification (5) the demand effect based on data from the alternative data source is somewhat smaller, but the difference between the coefficients derived from this exercise and our baseline is not statistically significant.

In specification (6) we want to see whether our empirical findings are affected by the global financial and economic crisis that started in 2008. Therefore, we re-estimate our baseline model allowing for different slope parameters of the demand and supply variable before and after 2008. Obviously, the estimated effects are not different across the two sample periods as both interaction terms ($\Delta D \times Crisis$, $\Delta S \times Crisis$) are statistically insignificant. This indicates that the estimated parameters of our baseline model are not biased due to structural breaks triggered by the crisis. In our next exercise we re-estimate our model by including time dummies, which were assumed to be zero in our baseline model. As we see in specification (7) this is indeed the case. At the bottom of the table we report the p-value from testing the joint significance of time dummies, which clearly indicates that time effects have to be excluded from the model equation as they have no explanatory power.

Finally, in specification (8) we employ instrumental variable techniques to address the potential endogeneity issue mentioned in the previous subsection. In particular, we apply the feasible system GMM estimator introduced by Blundell and Bond (1998) which uses internal instruments (i.e. time lags) for potentially endogenous variables (in our case the demand and supply variable).[27] The estimation results once again strengthen our baseline findings and strongly support the significance of demand and supply effects in the determination of gender wage gaps. Interestingly, in the short-run (a period of less than one year), the effects seem to be higher compared to our baseline estimates. However, if we calculate the long-run impact of both variables, the magnitude of the effects remain broadly the same. More specifically, taken into account the parameter estimate of the lagged dependent variable ΔGAP_{t-1}, we arrive at a point estimate for the demand effect of -1.69 and an estimate for the supply effect of 0.66.[28]

[27] Furthermore, we also add a lagged dependent variable given that the Arellano-Bond test confirms that the disturbances are autocorrelated of order one (but not of order two). The estimation includes country fixed effects. Testing for overidentifying restrictions indicates that the used set of instruments is valid across all specifications. Test statistics can be obtained from the authors upon request.

[28] Note that we obtain long-run coefficients $\hat{\beta}$ by $\hat{\beta} = \alpha/(1-\beta)$, where α is the parameter on the lagged

Taken together, the wide range of specifications has shown that the impact of demand and supply factors on gender wage gaps is highly robust across European countries. Parameter estimates, which are highly significant in each specification, vary from -1.1 to -1.7 for the demand effect and from 0.5 to 0.8 for the supply effect and are thus not only statistically but also economically significant.

6 Conclusions

In this paper, we have demonstrated that the public sector is a 'swing demander' for female labor. That is, we have shown that changes in the growth of employment in highly publicly influenced sectors have the potential to trigger economically significant fluctuations of the gender wage gap.

Our calculations show that the increase in female labor demand in Europe has reduced the corrected gender wage gap by 3.6 percentage points (being equivalent to a compression of 17%) in the period 2003 to 2013, all else equal. Furthermore, half of the increase in female labor demand can be traced back to the expansion of employment in sectors that are highly publicly influenced. Hence, had employment in the public sector remained constant in all countries over the entire period, the corrected gender wage gap would have been higher by 1.8 percentage points on average (again, all else equal). This is an economically significant effect, given that the corrected gender wage gap in Europe (accounting also for supply factors) has been compressed by a total of 3.5 percentage points in the same period.

Our findings do have quite substantial implications that have a cyclical as well as a more structural dimension. From a cyclical point of view, our results imply that consolidation policies – as far as they affect demand for public sector employment – will have an adverse impact on the gender wage gap. This is not to say that consolidation has to be counter-productive with regard to equal pay policies in any case. However, *if* consolidation is unaccompanied by policies specifically tailored at reducing sectoral segregation it will likely exhibit unwarranted

dependent variable and β the short-run coefficient on the respective independent variable.

effects on the gender wage gap. This implies that active labor market policies and retraining activities become even more important during consolidation episodes.

From a more structural policy point of view, our results raise important questions for further research. There are counter-veiling trends with regard to employment in the public sector. On the one hand, demographics in all likelihood will trigger a substantial hike of employment in the care sector in the future, likely increasing relative demand for female workers at the medium-skilled segment. On the other hand however, the growth of employment in other areas of the public sector will be probably more constrained amidst of over-stretched public finances thereby arguably slowing the growth of relative demand for female workers at the high skilled segment. In order to develop policies that aim at high employment and high gender equality, it will be crucial to learn more about the likelihood of these developments and their potential effects on relative demand and relative incomes at different skill-levels.

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7 Appendix

Table A1: Data-coverage and availability of certain variables

	Year of survey										
	2004	2005	2006	2007	2008	2009	2010	2011	2012	2013	2014
AT	▣	▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
BE	■	▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
BG				□	□	□	□	□	□	□	□
CH					▣	▣	▣	▣	▣	▣	▣
CY		▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
CZ		▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
DE		□	□	□	□	□	□	□	□	□	□
DK	□	□	□	□	□	□	□	□	□	□	□
EE	▣	▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
ES	■	■	▣	▣	▣	▣	▣	▣	▣	▣	▣
FI	□	□	□	□	□	□	□	□	□	□	□
FR	□	□	□	□	□	□	□	□	□	□	□
GR	■	■	■	□	□	□	□	□	□	□	□
HR							▣	▣	▣	▣	▣
HU		□	□	□	□	□	□	□	□	□	□
IE*	□	□	□	□	□	□	□	□	□	□	□
IT	■	■	■	▣	▣	▣	▣	▣	▣	▣	▣
LT		□	□	□	□	□	□	□	□	□	□
LU	▣	▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
LV		■	■	▣	▣	▣	▣	▣	▣	▣	▣
MT					▣	▣	▣	▣	▣	▣	▣
NL		▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
NO	□	□	□	□	□	□	□	□	□	□	□
PL		▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
PT	■	■	■	▣	▣	▣	▣	▣	▣	▣	▣
RO*				▣	▣	▣	▣	▣	▣	▣	▣
SE	□	□	□	□	□	□	□	□	□	□	□
SI		▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
SK		▣	▣	▣	▣	▣	▣	▣	▣	▣	▣
UK		□	□	□	□	□	□	□	□	□	□

(■, ▣, □) = Available EU-SILC cross-section datasets by country.

(■) = Dataset available but not used in analysis, (▣, □) = Data used in analysis.

(▣) = Work-experience variable available, (□) = Age and education-dummies used as work-experience proxies (potential experience).

Note that Spain, Greece, Italy, Latvia, and Portugal deliver gross income information only since 2007 (2006 in the case of Spain). Belgium 2004 is excluded as the variable on work-experience shows an exceptionally high fraction of missing data.

Table A2: Descriptive statistics of gender wage gap estimations, sample mean over time

	Total GAP	Total GAP_{OB}	Different estimations of the remuneration gap							
			base-line	sector dummy	private sector	public sector	occup. controls	selec-tion	self-empl.	age 26-55
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
AT	0.27	0.20	0.13	0.14	0.14	0.12	0.14	0.07	0.09	0.11
BT	0.10	0.09	0.13	0.13	0.14	0.13	0.12	0.13	0.13	0.14
BG	0.17	0.16	0.23	0.23	0.22	0.26	0.22	0.26	0.21	0.25
HR	0.11	0.10	0.15	0.16	0.18	0.10	0.15	0.09	0.15	0.16
CY	0.27	0.35	0.35	0.37	0.42	0.23	0.30	0.52	0.34	0.36
CZ	0.27	0.24	0.23	0.24	0.25	0.23	0.24	0.22	0.23	0.25
DK	0.13	0.10	0.17	0.11	0.14	0.08	0.15	0.12	0.13	0.16
EE	0.32	0.30	0.36	0.36	0.39	0.30	0.33	0.36	0.35	0.39
FI	0.24	0.18	0.25	0.23	0.22	0.25	0.18	0.24	0.05	0.25
FR	0.16	0.17	0.22	0.21	0.20	0.21	0.18	0.15	0.19	0.22
DE	0.30	0.24	0.17	0.16	0.19	0.13	0.20	0.14	0.16	0.13
GR	0.15	0.13	0.17	0.18	0.21	0.15	0.16	0.30	0.16	0.16
HU	0.10	0.07	0.14	0.15	0.14	0.15	0.16	0.14	0.14	0.15
IE	0.09	0.11	0.12	0.15	0.15	0.17	0.14	0.17	0.12	0.12
IT	0.15	0.09	0.12	0.14	0.14	0.12	0.13	0.16	0.09	0.13
LV	0.16	0.15	0.26	0.26	0.28	0.23	0.25	0.33	0.23	0.27
LT	0.10	0.11	0.20	0.21	0.25	0.16	0.20	0.11	0.17	0.22
LU	0.17	0.15	0.16	0.18	0.16	0.23	0.16	0.14	0.16	0.21
MT	0.14	0.10	0.12	0.11	0.12	0.11	0.12	0.04	0.10	0.11
NL	0.26	0.18	0.14	0.14	0.17	0.14	0.14	0.13	0.14	0.12
NO	0.25	0.21	0.24	0.17	0.23	0.10	0.22	0.14	0.18	0.25
PL	0.06	0.04	0.12	0.15	0.15	0.15	0.12	0.10	0.12	0.14
PT	0.06	0.11	0.21	0.24	0.25	0.22	0.21	0.37	0.21	0.22
RO	0.12	0.13	0.16	0.17	0.16	0.17	0.16	0.10	0.16	0.16
SK	0.20	0.19	0.21	0.21	0.23	0.19	0.22	0.17	0.21	0.23
SI	0.08	0.06	0.15	0.17	0.17	0.16	0.17	0.12	0.14	0.15
ES	0.15	0.15	0.16	0.19	0.19	0.18	0.16	0.26	0.15	0.16
SE	0.19	0.18	0.23	0.15	0.19	0.14	0.20	0.00	0.18	0.24
CH	0.30	0.23	0.10	0.09	0.12	0.04	0.10	0.08	0.05	0.08
UK	0.23	0.21	0.19	0.19	0.17	0.21	0.18	0.22	0.18	0.17
Total	0.18	0.16	0.19	0.19	0.20	0.17	0.18	0.18	0.17	0.19

Notes: Total GAP (1) represents the difference in (mean) log-hourly wages between male and female. Column (2) includes the same measure as column (1) but is based on a restricted sample. This sample consists of only those observations that enter the Oaxaca-Blinder decomposition, i.e. for which all control variables are available. Column (3) displays the remuneration gap, which is the difference in log-hourly wages excluding differences in characteristics, i.e. the gender wage gap after controlling for individual factors (baseline set of controls). Columns (4) to (10) present alternative estimations of the remuneration gap: Column (4) controls for the sector (public sector dummy), column (5) uses the baseline set of controls but restricts the sample to private sector employees, (6) restricts the sample to public sector employees, column (7) includes occupational controls, column (8) controls for non-random selection into employment, column (9) extends the sample by including self-employed persons and column (10) restricts the sample to 26-55 aged employees.