You're the one that I want! Understanding the over-representation of women in the public sector*

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Abstract

The public sector hires dis-proportionally more women than men. Using microdata for the United States, the United Kingdom, France, and Spain, we document gender differences in employment, transition probabilities, hours, and wages in the public and private sector. We use the data to calibrate a search and matching model where men and women decide if to participate and whether to enter public or private-sector labor markets. We then quantify how much of the selection of women into the public sector is driven by: (i) lower gender wage gaps, (ii) fewer hours, (iii) greater job security, or (iv) intrinsic preferences for public sector occupations. Preferences and wages explain most of the over-representation, with significant variations across countries and educational groups. We calculate the monetary value of public sector job security and hours premia, finding that women value fewer hours more while men value job security more. We also show how public-sector wage and employment policies dis-proportionally affect female unemployment.

JEL Classification: J21, J16, J45, H10, E60.

Keywords: public-sector employment, public-sector wages, female labor force participation, gender wage gap.

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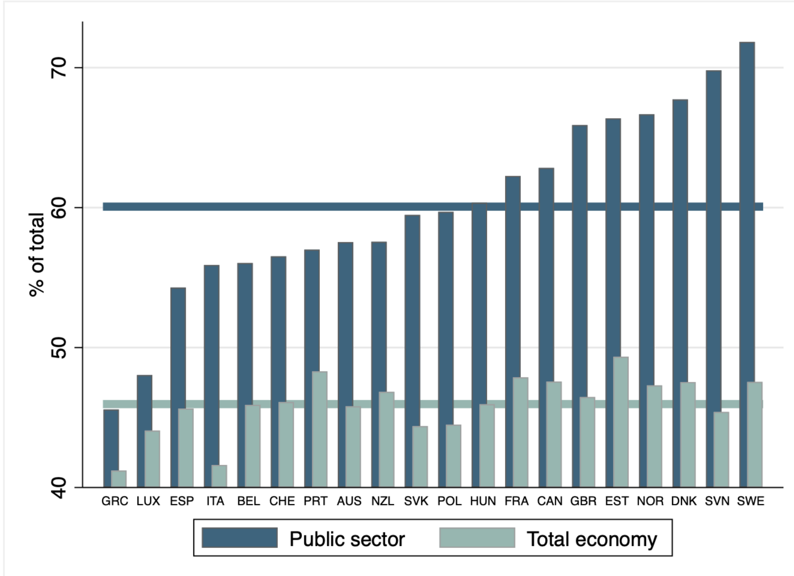
1 Introduction

Public employment in OECD countries accounts for 10 to 35 percent of total employment, and the public sector hires significantly more women than men. Although women only represent 46 percent of employment, they make up 60 percent of public-sector workers, see Figure 1. Despite its importance for women, there is no quantitative study aimed at understanding their over-representation in the public sector.

To fill this gap in the literature we first use microdata to document facts regarding gender differences in employment, transition probabilities, hours, and wages in the public and private sector in the United States, the United Kingdom, France, and Spain. We show that the over-representation of women in public employment persists across age groups, regions, levels of education, as well as over time. While less pronounced, women are also over-represented when excluding health care and education from public employment. In the public sector, gender wage gaps are smaller, full-time workers work fewer hours, and there is more job security reflected in a lower probability of moving to unemployment. Motivated by these empirical findings, we build a search and matching model where men and women decide whether to participate and whether to enter public or private-sector labor markets. We view the over-representation of women as driven by supply, meaning that the government does not explicitly hire more women, but it is women who choose the public sector more so than men. While more women might have a public-service motivation or prefer to be teachers, we argue that part of the explanation lies with other characteristics of public-sector jobs.

We calibrate our model separately to the United States, the United Kingdom, France and Spain, using statistics from our empirical analysis to identify key parameters. Running counterfactual experiments, we quantify how much of the selection of women into the public sector is driven by: (i) lower gender wage gaps and thus relatively higher wages for women in the public sector (estimated directly from the data), (ii) better reconciliation of work and family life in the public sector due to fewer hours of full time workers (estimated directly from the data), (iii) greater job security (derived from differences in flows from private and public employment to unemployment), or (iv) intrinsic preferences for public-sector activities (identified as a residual). We find that women's preferences are a key determinant for their over-representation but do not explain everything. In France, Spain, and the United States, relatively higher wages for women in the public sector explain 10, 51, and 70 percent of their over-representation. Fewer hours in the public sector play a role in France and Spain, accounting for 5-12 percent of women's over-representation. When considering individuals with different levels of education, we find that the female over-representation among college educated is mostly due to lower gender wage gaps and fewer working hours in the public

Figure 1: Share of women in the public sector and total economy



Source: OECD [2015]; this data does not include the United States; 56 percent of US public-sector workers are women compared to 48 of all workers, see Hammouya [1999].

sector. For individuals without a college degree preferences matter most. In all countries, greater job security in the public sector reduces the over-representation of women because it is valued more by men than by women.

This last result is not opposed to both men and women valuing the job security offered by the public sector; something we quantify within our framework. We calculate how much of their wages private-sector workers would be willing to sacrifice for lower job-separation rates as well as fewer working hours. Job security and hours premia range from 0.8-1.4 percent in the United States and the United Kingdom to 1.5-3.3 percent in France and Spain, countries with higher unemployment rates and fewer working hours in the public sector. In all countries, women are willing to pay more for fewer hours, and men are willing to pay more for job security.

Even though female labor force participation has increased remarkably over the past decades, gender gaps in participation and employment still persist. In 2022, in the United States, the United Kingdom, France, and Spain employment rates of women were between 7-10 percentage points lower than men's, the gap increasing to 13-20 percentage points when considering full-time equivalent (FTE) rates, see OECD [2010-2021]. Women also continue to earn lower wages. Many explanations for persisting gender gaps in the labor market have been proposed and tested.[1] However, one aspect that has been continuously overlooked is

[1] Ranging from education choices which lead to men and women working in differently paid industries

the role of public employment for female labor market outcomes and hence for gender gaps in employment, participation, and wages.[2] Incorporating a public sector into a search and matching model, we show that public employment slightly reduces the aggregate gender wage gap, and more importantly that government wage and employment policies have 2 to 5 times larger effects on female compared to male unemployment.

Literature on the intersection of public employment and female labor market outcomes mostly consists of descriptive empirical studies.[3] For instance, Gornick and Jacobs [1998] establish that women face lower gender wage gaps in the public compared to the private sector and attribute it to a more compressed public-sector wage distribution, while Anghel et al. [2011] find that unemployed and inactive women are more likely to search for public-sector jobs than men. Rosen's [1996] study on the expansion of the Swedish public sector reveals that between 1963 and 1993 employment of women in local government increased fourfold while that of men only doubled. Kolberg [1991] stresses that the expansion of the Scandinavian welfare state and increased public employment has led to more women participating in the labor market, and according to Adserà [2004] the higher share of women in stable public-sector jobs partially accounts for higher fertility rates in Scandinavian countries, a dimension we abstract from in our model.

Regarding theoretical contributions, the only related paper we are aware of is Bradley et al. [2017] who set up a model of public employment calibrated for markets segmented by gender and education, hence abstracting from the interaction of men and women in the labor market. Furthermore, while studying the effects of public-sector hirings on private employment, the authors do not model individuals' participation decisions. Given sizable gender differences regarding inactivity rates and transitions into non-employment, it seems

(Gemici and Wiswall [2014]), to maternity and institutional aspects like availability of child care and possibility of working part time (Del Boca [2002]), to maternity and selection away from industries with inflexible working hours and higher pay (Goldin [2014]), to behavioral gender differences regarding competition (Manning and Saidi [2010]), to differences in time spent on household chores (Albanesi and Olivetti [2009]).

[2] Two of the most influential surveys on female employment by Killingsworth and Heckman [1986] and Blundell and Macurdy [1999] do not even mention the public sector. This is a serious omission because unlike any other sector the public sector does not face the same competitive forces and constraints as private firms, and its employment and wage policies are tailored to different objectives; for instance, attaining budgetary targets (Poterba and Rueben [1995], Gyourko and Tracy [1989]), redistributing resources (Alesina et al. [2000] and [2001]) or satisfying interest groups for electoral gains (Borjas [1984], Matschke [2003]).

[3] Separately the topics (i) female labor market outcomes and (ii) public employment have generated a large body of theoretical literature. Regarding (i) the focus has been on aspects such as child care costs (García-Morán and Kuehn [2017]), parental leaves (Erosa et al [2010]) divorce risk (Fernández and Wong [2014]), or welfare states (Greenwood et al [2000]), but none considers the effects of public employment. Regarding (ii), theoretical models tend to emphasize aggregate labor market effects, in particular the effects on private wages and the crowding out of private employment (see Finn [1998] in an RBC model or Gomes [2015] in a search and matching model). Recent contributions include Navarro and Tejada [2022], Chassamboulli and Gomes [2023], Boeing-Reicher and Caponi [2024], or Geromichalos and Kospentaris [2022].

natural to account for non-participation when explicitly incorporating women into a search and matching model.[4] Hence, to the best of our knowledge, we are the first to propose a theoretical framework that combines both public employment and participation decisions, and where men and women interact. The latter is key for understanding how government wage and employment policies affect the over-representation of women.

The paper is organized as follows. The next section presents our empirical analysis. Sections 3 and 4 present the model and its calibration. In Section 5 we use the model to carry out counterfactual experiments, calculate compensating differentials for public-sector job characteristics, and conduct policy experiments. Section 6 concludes.

2 Empirical analysis

We use microdata to document gender differences in employment, transition probabilities, hours, and wages in the public (g) and private (p) sector in the US, the UK, France, and Spain. These four countries have sizable public sectors which encompass different industries and employ distinct hiring processes. Common findings are likely to be intrinsic characteristics of the public sector and unlikely due to country specificities. For employment and transition probabilities we consider data from representative labor force surveys from each country, extracted by Fontaine et al. [2020].[5] For estimating public-sector wage and hours premia we use data from the Structure of Earnings Survey [2002, 2006, 2010, 2014] for Europe and the CPS March Supplement [2003-2018] for the US; see the Appendix B for details on the data.

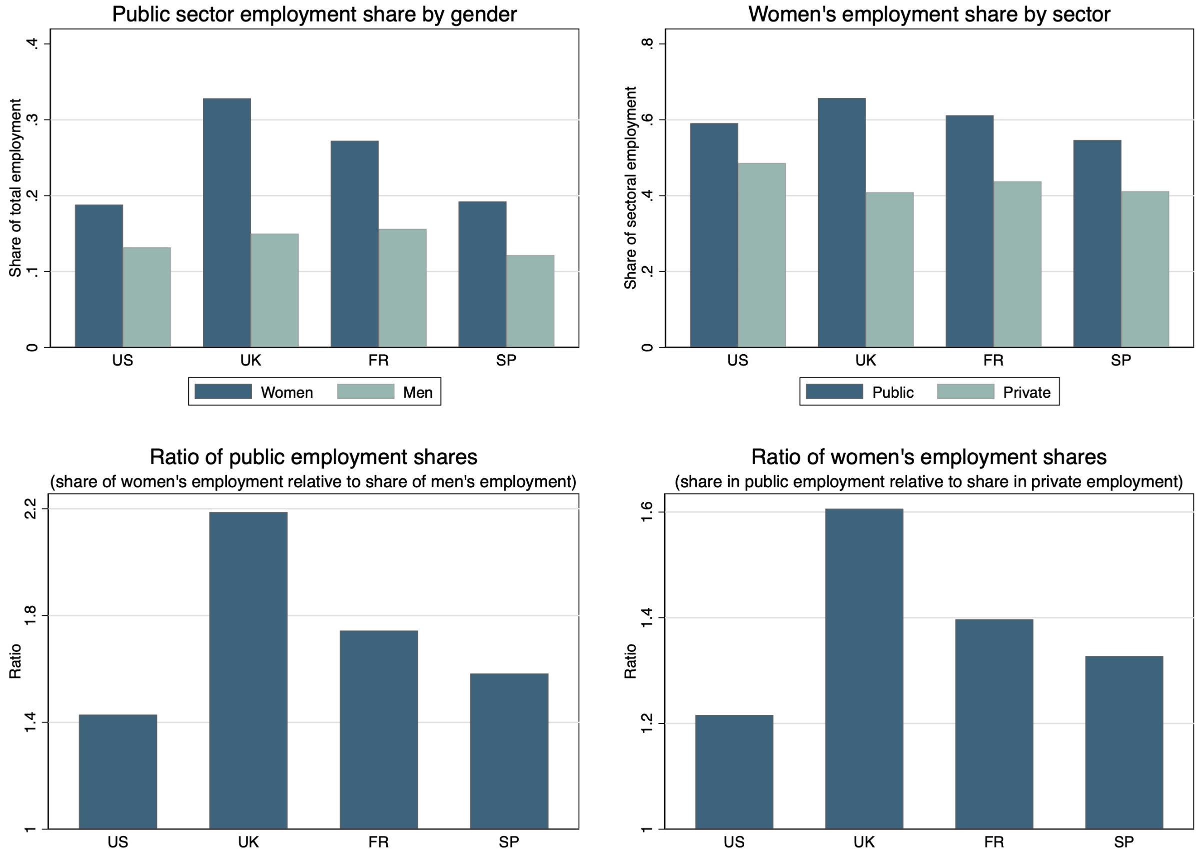
2.1 Over-representation of women in the public sector

While the size of the public sector varies across countries, being larger in the UK and France and smaller in the US and Spain, the share of public employment is always larger for women who represent the majority of public-sector workers, see the top graphs of Figure 2. The two bottom graphs display two size-independent indicators for the over-representation of women. The first is the ratio of public employment shares, $(\frac{e_{g,f}}{e_f}/\frac{e_{g,m}}{e_m})$, where $e_{g,f}$ ($e_{g,m}$) is the number of women (men) employed in the public sector, and e_f (e_m) denotes all employed women (men). The second indicator is the ratio of women's employment shares, defined as $\frac{e_{g,f}}{e_g}/\frac{e_{p,f}}{e_p}$,

[4] Modeling non-participation goes back to Pissarides [1990] Chapter 6, and has been advanced by Garibaldi and Wasmer [2005], Pries and Rogerson [2009], Krusell et al. [2011], Haefke and Reiter [2011] and Albanesi and Şahin [2018]; none of which specify a public sector.

[5] See Fontaine et al. [2020] for the methodological details. The extracted data is available on the authors' websites

Figure 2: Different measures for the over-representation of women in public employment



Note: Data is from the CPS [2003-2018], the UK Labour Force Survey [2003-2018], the French Labour Force Survey [2003-2017], and the Spanish Labour Force Survey [2003-2018], extracted by Fontaine et al. [2020].

where e_g (e_p) denotes employment in the public (private) sector. In case of perfect gender symmetry across sectors, both indicators would take on a value of 1. However, across the four countries the ratio of public employment shares lies between 1.4 and 2.2, and the ratio of women's employment shares varies between 1.2 and 1.6.

One explanation is that certain types of jobs predominantly carried out by the government are preferred by women. As the top two graphs of Figure A.1 in Appendix A reveal, for the US, the UK, and France, once we exclude health care and education, while lower, women's public employment is still 20-50 percent higher than men's. Although within health care and education men and women are similarly likely to work in the public or private sector, both industries contribute to the female over-representation because women make up a larger fraction of educators and health care professionals overall. In the UK, for instance, disregarding health care and education lowers the ratio of public employment shares from 2.2 to 1.5. For the US, we also analyze the gender composition of public-sector jobs based on a 3-digit ISCO-08 occupational classification, see bottom graphs of Figure A.1 in Appendix A. Within two-thirds of occupations women's relative representation in the public sector exceeds that of men, more so among *Radiation Therapists* or *Software Developers*, and less so among *Library Assistants* or *Human Resources Specialists*. This indicates that women's preferences matter, but that they cannot account for their entire over-representation. In our model, we include gender differences in preferences for public-sector jobs but also additional explanations related to job characteristics (wages, hours, job security), and we test for the importance of each.

We also consider how our size-independent indicators vary over time, by workers' age, level of education, region, and hours worked. Figures A.2 and A.3 show that both indicators are persistent over time, even though they fell around the time of the great recession, due to large changes in private employment. Regarding workers' age groups, both ratios are close to 1 for very young workers, but they increase around age 20 and remain relatively constant over the life-cycle (see Figures A.4 and A.5). The over-representation of women is present across all levels of education and particularly strong among primary and tertiary educated workers, see Figure A.6. We also check for regional variation. The ratio of public employment shares is larger than one in all US states, ranging from 1.1 to 1.7. The picture is similar for the other three countries, with the exception of two regions in Spain – Ceuta and Melilla – characterized by a strong presence of the armed forces due to their location on the African continent. Finally, we also verify that the over-representation of women is not explained by their potentially higher part-time incidence in the public sector. As numbers in Table A.1 show, women in all three European countries are 2-3 times more likely to work part time compared to men. However, with the exception of Spain, this difference is similar across the two sectors. Hence, when computing full-time equivalent public employment shares ratios, we find that this does not affect numbers in the US, the UK, or France and slightly increases women's over-representation in Spain, by 8%.[6]

2.2 Transition rates by sector and gender

Public-sector workers have a much lower probability of becoming inactive or unemployed compared to private-sector workers as shown by the transition rates between private (P) and public (G) employment and unemployment (U) and inactivity (I) in Table 1. While the probability of dropping out of the labor force is higher for women compared to men, it is 35

[6] To arrive at these statistics, we consider part-time employment among men and women in the public and private sector, as reported in Table A.1 in Appendix A. We then calculate full time equivalent employment in each sector by weighting the respective part-time shares by 0.5 given that part-time is often done through job sharing where one job is split in two and filled by two individuals working part-time.

Table 1: Transition rates by sector and gender

	US			**UK**		
	All	Men	Women	All	Men	Women
$P \to U$	0.014	0.016	0.012	0.014	0.015	0.013
$G \to U$	0.007	0.006	0.008	0.005	0.006	0.005
$E \to U$	0.013	0.015	0.011	0.012	0.014	0.010
$P \to I$	0.023	0.019	0.028	0.021	0.016	0.030
$G \to I$	0.018	0.015	0.020	0.017	0.014	0.018
$E \to I$	0.023	0.019	0.026	0.020	0.016	0.026
	France			**Spain**		
	All	Men	Women	All	Men	Women
$P \to U$	0.021	0.020	0.022	0.041	0.040	0.043
$G \to U$	0.008	0.008	0.008	0.021	0.019	0.022
$E \to U$	0.018	0.018	0.018	0.038	0.037	0.038
$P \to I$	0.022	0.019	0.027	0.030	0.021	0.043
$G \to I$	0.017	0.014	0.018	0.022	0.016	0.026
$E \to I$	0.021	0.018	0.024	0.029	0.020	0.040

Note: Data is from the CPS [2003-2018], the UK Labour Force Survey [2003-2018], the French Labour Force Survey [2003-2017], and the Spanish Labour Force Survey [2003-2018], extracted by Fontaine et al. [2020]. Transition rates indicate the probability of an employed worker becoming unemployed or inactive in the following quarter (month in the US).

to 67 percent higher if she works in the private compared to the public sector. In our model such differences in employment to inactivity flows by gender and sector arise endogenously. In all countries separation rates for both genders are between 2 to 3 times higher in the private compared to the public sector. To capture differences in job security across the two sectors we use these last numbers as inputs in our model.

We also estimate probabilities of leaving employment conditional on observable characteristics such as age, education, and occupation to account for the different composition of public and private employment. Appendix B provides details on this estimation, and Figure B.1 shows the results. The conditional probability of dropping out of the labor force is higher for women than for men. For women in the three European countries this probability is lower if they work in the public compared to the private sector. In the US this difference is insignificant. The conditional probability of becoming unemployed is lower for public compared to private-sector workers, with the difference in job security being highest in France, followed by the UK, Spain, and the US.

2.3 Wage premium and working hours

To estimate public-sector wage premia we consider individuals' annual earnings instead of hourly wages to be able to separately identify higher wages from fewer working hours in the public sector.[7] We hence estimate the following regression

$$log(y_i) = \beta_0 + \beta_1 f + \beta_2 X_i + \beta_3 m \times pub + \beta_4 f \times pub + \beta_5 C_i + d_r + d_y + \epsilon_i, \quad (1)$$

where $log(y_i)$ is the log of individual i's gross yearly earnings, f is an indicator for female, X_i denotes other individual characteristics such as age, education, and race, $m \times pub$ ($f \times pub$) is an indicator for a man (woman) working in the public sector. We also control for other job characteristics (C_i) such as occupation, tenure, tenure squared, part time, as well as region and year fixed effects. Our coefficients of interest are β_1, β_3 and β_4. The first indicates how women's earnings in the private sector differ with respect to men's – the private-sector gender wage gap. The second and third coefficient measure the public-sector wage premia for men and women respectively.

Panel A in Table 2 displays the results. In column (1) all workers are considered, and we control for part-time status, while in column (2) – our preferred specification aligning more closely with our model – we limit the sample to full-time workers. In all countries but France, women working full time in the public sector have 4-7 percent higher gross yearly earnings. For men these premia are smaller, ranging from 4 percent in the UK to less than 1 percent in Spain, to being negative in the US and France. While in France women also face lower earnings in the public compared to the private sector, their discount is smaller. Higher public-sector premia for women compared to men are equivalent to lower gender wage gaps in the public compared to the private sector. Regarding private-sector earnings, these are 16-32 percent lower for women. We estimate the largest private-sector gender wage gap in the US, followed by Spain, the UK, and France.[8]

Our estimated coefficients are in line with findings in literature on relatively higher public-sector wages premia for women compared to men. Lucifora and Meurs [2006] estimate 8 (3)

[7] While this contrasts with most of the literature, consider a case of equivalent full time workers who earn the same monthly wage in both private and public sectors. If, in the private sector, individuals work more than the maximum weekly hours, then their hourly wage would be lower, and we would estimate a positive public-sector wage premium, potentially due to problems of enforcing working-time legislation in the private sector leading to longer hours. To include the two dimensions explicitly in the model we separately estimate an hours premium and a wage premium using annual earnings.

[8] Unlike data for European countries, data for the US includes information on individuals' race but unfortunately not on tenure which being on average longer in the public sector might imply that US estimates for the public sector wage premia could be somewhat upward biased. However, this is unlikely to directly affect gender differences in these premia.

Table 2: Public-sector wage and hours premium and private-sector gender wage gap

	US		**UK**		**France**		**Spain**	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Panel A: wage regressions								
Public-sector wage premium								
Men	-0.019***	-0.029***	0.031***	0.041***	-0.109***	-0.109***	0.0002	0.007***
	(-6.38)	(-10.02)	(9.29)	(13.56)	(-46.35)	(-50.56)	(0.10)	(3.02)
Women	0.057***	0.054***	0.048***	0.052***	-0.047***	-0.098***	0.069***	0.073***
	(21.68)	(19.76)	(18.07)	(17.52)	(-19.88)	(-40.63)	(28.58)	(32.33)
Gender wage gap								
Private	-0.290***	-0.284***	-0.187***	-0.219***	-0.167***	-0.186***	-0.214***	-0.247***
	(-165.21)	(-160.96)	(-81.29)	(-94.92)	(-100.36)	(-117.03)	(-163.17)	(-196.71)
Controls								
Demographics	X	X	X	X	X	X	X	X
Region Year FE	X	X	X	X	X	X	X	X
Job characteristics	X	X	X	X	X	X	X	X
Part time dummy	X		X		X		X	
Only full time wkr		X		X		X		X
Obs.	1,071,617	837,088	622,013	443,801	722,571	612,527	876,348	747,302
R-squared	0.516	0.403	0.561	0.387	0.496	0.453	0.599	0.538
Panel B: hours regressions								
Public-sector hours premium								
Public	-0.013***	-0.028***	-0.042***	-0.036***	-0.056***	-0.082***	-0.054***	-0.056***
	(-13.46)	(-39.54)	(-38.49)	(-73.80)	(-73.67)	(-213.63)	(-64.11)	(-230.85)
Controls								
Demographics	X	X	X	X	X	X	X	X
Region Year FE	X	X	X	X	X	X	X	X
Job characteristics	X	X	X	X	X	X	X	X
Part time dummy	X		X		X		X	
Only full time wkr		X		X		X		X
Obs.	1,071,617	837,088	622,013	443,801	722,571	612,527	876,348	747,302
R-squared	0.377	0.073	0.556	0.201	0.418	0.188	0.548	0.303

Note: Estimated by OLS regressions. In Panel A which estimates Equation 1 the dependent variable is the log of gross yearly earnings. The public-sector wage premium for men (women) corresponds to the coefficient β_3 (β_4). The private-sector gender wage gap corresponds to the coefficient β_1. Panel B estimates Equation 2. The dependent variable is the log of annual hours worked. The public sector hours premium correspond to the coefficient α_3. Data for the UK, France, and Spain from the Structure of Earnings Surveys [2002, 2006, 2010, 2014] and for the US from the CPS March Supplement [2003-2018]; for descriptive statistics see Table B.1 in Appendix B. Demographic controls include age, education and race (only for the US). Job characteristics include occupation and for European countries tenure and tenure squared.

percent higher returns in the UK public sector for the median woman (man). Bargain et al. [2018] also find negative public-sector wage premia for France (between 0 and -5 percent) which are slightly smaller for women. Controlling for selection, Hospido and Moral-Bonito [2016] estimate for Spain an average 10 percent public sector wage premium, somewhat higher for women.

Using the same data, we estimate if and by how much, individuals in the public sector work fewer hours compared to those in the private sector. To obtain a broad measure of working hours we combine information on holidays and usual hours worked per week to construct annual hours worked, and we run the following regression

$$log(hours_i) = \alpha_0 + \alpha_1 f + \alpha_2 X_i + \alpha_3 pub + \alpha_4 C_i + d_r + d_y + \epsilon_i \quad (2)$$

controlling for the same individual and job characteristics as before together with year and region fixed effects. Results are displayed in Panel B of Table 2. Focusing on our preferred specification in columns (2), we observe that individuals holding full-time jobs in the public sector work between 4-8 percent fewer hours compared to similar individuals in the private sector. The hours discount is lowest in the US and the UK (around 3-4%), and highest in Spain (6%) and France (8%). These differences might be due to higher unionization rates in the public sector in some countries leading to better collective bargaining outcomes for weekly hours or vacations, or different legislation on (or enforcement of) working hours across the public and private sector. However, fewer working hours for full-time employees are just one aspect of a better work-life balance, alongside additional sick days, flexibility to work from home, or employer-provided child care. Under the assumption that such practices are more extended in the public sector our estimated numbers on hours discounts capture a lower bound on the difference in work-life balance between the two sectors.

Results from our empirical analysis suggest that the over-representation of women in the public sector could potentially be due to fewer working hours, lower gender wage gaps, or greater job security. Alternatively women could simply have a preference for public sector occupations. To quantify the importance of each of these factors we set up a model economy.

3 Model

We build a search and matching model with a public sector that hires a fixed number of workers. At each instant τ, individuals are born (enter the labor market) and die (retire), such that the working population is constant and normalized to unity. Agents are risk-neutral and discount the future at rate $r > 0$. Time is continuous. Individuals – men and women – can either be employed or non-employed. All variables are indexed by two subscripts: $i = [g, p]$, where g refers to the public (government) sector and p to the private sector, and $j = [m, f]$, where m refers to male and f to female.

We define individuals' flow utilities for employment (E) and non-employment (NE) in each sector as:

$$v_{i,j}^{E} = (1 - \xi_i)x + w_{i,j}, \tag{3}$$

$$v_{i,j}^{NE} = x, \tag{4}$$

where x denotes the stochastic value of home production – the opportunity cost of working – distributed in $X = [0, \infty)$, and with cumulative distribution function $F_j(x)$ and density $f_j(x)$ which differ by gender. Non-employed individuals enjoy the utility of home production, but depending on the value of x, they would accept or not a job if offered, generating an endogenous split between unemployed and inactive individuals. Workers receive a wage payment $w_{i,j}$ and spend a fraction ξ_i, of their time at work, normalized to 1 in the private and set to $\xi_g \leq 1$ in the public sector.

Prior to entering the labor market individuals draw a preference ϵ, which reflects a taste for working in the public sector and/or individuals' entry costs (e.g. due to entrance exams for civil servants). We assume that for men and women individual preferences are distributed according to cumulative distribution functions $\Xi_m(\cdot)$ and $\Xi_f(\cdot)$. In the spirit of a generalized Roy model, an endogenous proportion of the population (those whose preferences are sufficiently high) enters the public sector, while the other fraction joins the private sector. This can be interpreted as an occupational choice given that certain jobs, such as teacher or police officer, require training that is specific to the public sector. This assumption that rules out any direct (job-to-job) or indirect (via unemployment or inactivity) transitions between the public and private sector is supported by two facts. First, the majority of inflows into and outflows from public employment are from and to non-employment.[9] Second, even after unemployment or inactivity spells, workers are more likely to find a job in their sector of previous employment.[10] Nevertheless, there will be interaction between the public and private sector labor markets. If, for instance, the public sector offers shorter working hours $\xi_g < 1$, it will attract more women who have on average higher outside options x. As the number

[9] As shown by Chassamboulli et al. [2020], in France and Spain workers employed in the private sector in the previous quarter represent only 10 to 15 percent of inflows into public employment (around 30 percent in the US and the UK). Similar magnitudes hold for outflows. Dickson et al [2014] report five-year transition matrices between private and public sector employment for workers in Spain (France) where more than 88% (98%) of workers are observed in their previous sector five years later.

[10] In the US, the unconditional job-finding rate in the public sector is 1.8 percent, but conditional on having been employed in the public sector in the month preceding unemployment it is close to 30 percent. The public sector job-finding rate conditional on being previously employed in the private sector is 1.4 percent, roughly equal to having been unemployed or inactive. For the private sector, the job-finding rate conditional on previous private-sector employment is higher than 40 percent. Being previously employed in the public sector does not raise the job-finding rate in the private sector relative to having been unemployed or inactive (with job-finding rates of around 16 percent); see Fontaine et al. [2020] (Appendix V).

of jobs is fixed, this lowers the probability of finding jobs for both men and women. Men, who value work-life balance less, will hence turn to the private sector, further reinforcing the gender imbalance in the public sector.[11]

3.1 Value functions

Given individuals' flow utilities and transition probabilities, the value functions for employment and non-employment for men and women in the two sectors are as follows:

$$(r+\tau+\lambda)E_{i,j}(x) = v^{E}_{i,j}(x)+\delta_i[NE_{i,j}(x)-E_{i,j}(x)]+\lambda\int_0^\infty \max(E_{i,j}(x'),NE_{i,j}(x'))dF_j(x'), \quad (5)$$

$$(r+\tau+\lambda)NE_{i,j}(x) = v^{NE}_{i,j}(x)+m(\theta_i)[\max(E_{i,j}(x),NE_{i,j}(x))-NE_{i,j}(x)]+\lambda\int_0^\infty NE_{i,j}(x')dF_j(x') \quad (6)$$

where λ is the arrival rate of i.i.d. shocks to the opportunity costs of working common to both sectors and genders, and δ_i is the separation rate. We interpret distinct separation rates by sector as reflecting differences in job security. The conditional job-finding rate in sector i is $m(\theta_i)$, which is endogenous and is assumed to be the same for men and women. When firms or the government are matched with a worker, they hire him or her independently of their gender.

The value of employment encompasses the flow utility of being employed, the loss suffered when separated, and the change whenever individuals draw a new x. The value of non-employment sums the flow utility of non-employment, the gains when finding a job, plus the change due to a new draw of x.

3.2 Threshold

Individuals' values of x are a main determinant of their labor market state. We can implicitly define a threshold for the marginal individual, indifferent between being employed or non-employed:

$$E_{i,j}(\hat{x}_{i,j}) = NE_{i,j}(\hat{x}_{i,j}). \quad (7)$$

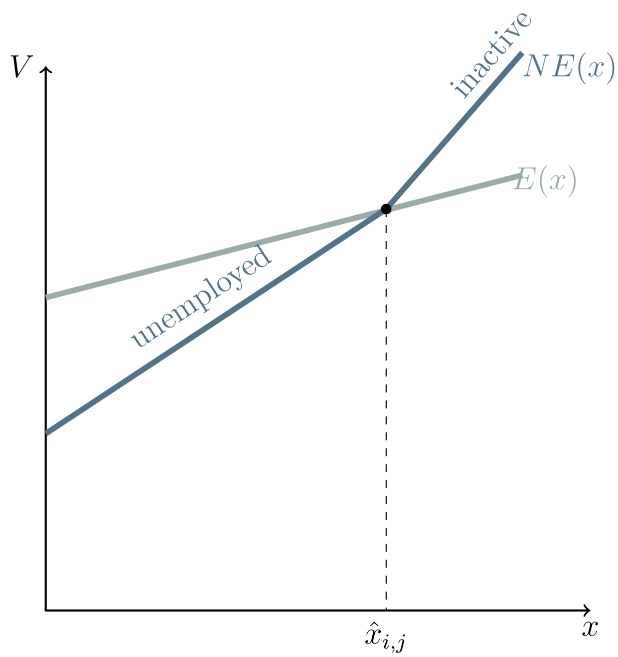
Individuals with a low enough x work while those with x above $\hat{x}_{i,j}$ prefer inactivity. Figure 3 displays the value functions for employment and non-employment together with the threshold, which can be expressed as:

[11] Allowing agents to switch between sectors would require us to take a stance on aspects such as reinstatement policies which allow individuals in the US who have worked for the public sector to return more easily, or loss of pension rights when giving up civil servant status (as in France or Spain).

$$\hat{x}_{i,j} = \frac{w_{i,j}}{\xi_i} + \frac{\lambda}{\xi_i}[A_{i,j} - B_{i,j}], \tag{8}$$

where $A_{i,j} = \int_0^\infty \max(NE_{i,j}(x'), E_{i,j}(x'))dF_j(x')$ and $B_{i,j} = \int_0^\infty NE_{i,j}(x')dF_j(x')$. A higher wage moves the threshold to the right, and fewer individuals quit their jobs for inactivity. This implies a direct link between a gender wage gap and higher inactivity rates of women compared to those of men.

Figure 3: Decision Threshold



Note that in Figure 3, the slope of the value function for non-employment is discontinuous, due to the difference between unemployed and inactive individuals. If $x > \hat{x}_{i,j}$, upon separation workers move into inactivity, while if $x \leq \hat{x}_{i,j}$ they become unemployed (if offered a job they would accept it). Instead of one value function for non-employed individuals we could thus define two values functions, one for unemployed and another one for inactive individuals, see Equations C.1 to C.3 in Appendix C.

3.3 Flows in and out of each state

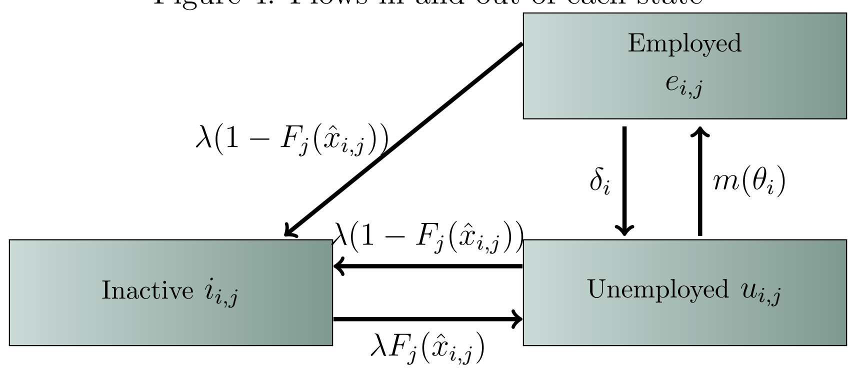
For men and women in each sector, there are three labor market states: inactive ($i_{i,j}$), unemployed ($u_{i,j}$), and employed ($e_{i,j}$). Figure 4 shows the hazard rates between the states, abstracting from labor force entries or retirements. In steady state, the flows in and out of each stock must be equal:

$$i_{i,j}(\lambda F_j(\hat{x}_{i,j}) + \tau) = \lambda(1 - F_j(\hat{x}_{i,j}))[e_{i,j} + u_{i,j}] + \tau(1 - F_j(\hat{x}_{i,j})) \tag{9}$$

$$u_{i,j}(\lambda(1 - F_j(\hat{x}_{i,j})) + \tau + m(\theta_i)) = \delta_i e_{i,j} + \lambda F_j(\hat{x}_{i,j}) i_{i,j} + \tau F_j(\hat{x}_{i,j}) \tag{10}$$

$$e_{i,j}(\lambda(1 - F_j(\hat{x}_{i,j})) + \tau + \delta_i) = m(\theta_i) u_{i,j}. \tag{11}$$

Figure 4: Flows in and out of each state



3.4 Private sector

To limit the complexity of the model, we abstain from explicitly modeling bargaining over the surplus of the match. Instead, we assume that male private-sector wages are a constant fraction β of workers' productivity

$$w_{p,m} = \beta y. \tag{12}$$

Male private-sector wages are hence completely isolated from policy changes which allows us to focus on the first-order effects of such policies on female labor market outcomes. Female wages, on the other hand, are determined as in Albanesi and Şahin [2018]. Given male wages and free entry for firms, the value of a job filled by a man is

$$rJ_m = (1 - \beta)y - \delta_p + \tau + \lambda(1 - F_m(\hat{x}_{p,m}))J_m, \tag{13}$$

which solving for J_m gives

$$J_m = \frac{(1 - \beta)y}{r + \delta_p + \tau + \lambda(1 - F_j(\hat{x}_{p,m}))}. \tag{14}$$

We set the female private-sector wage such that the value of a job for a firm is the same for male and female workers, $J_m = J_f$, and hence we arrive at:

$$w^*_{p,f} = y\left(1 - (1-\beta)\frac{r + \delta_p + \tau + \lambda(1 - F_j(\hat{x}_{p,f}))}{r + \delta_p + \tau + \lambda(1 - F_j(\hat{x}_{p,m}))}\right). \tag{15}$$

As long as there are more inactive women compared to men, i.e. $F_j(\hat{x}_{p,f}) > F_j(\hat{x}_{p,m})$, then $w_{p,m} > w^*_{p,f}$. Women receive lower wages because they have higher quit rates into inactivity which is anticipated by employers. However, this mechanism alone will not allow us to replicate the gender wage gap observed in the data as our model does not encompass important drivers of gender wage inequality such as differences in occupations or "normal" and "odd" hours; see Goldin [2014].[12] Nevertheless, as one of our explanations for the over-representation of women relates to lower gender wage gaps in the public sector, we have to make sure to reproduce the private-sector gender wage gap. Otherwise, our model would load gender differences in activity rates solely onto differences in distributions of outside options. To avoid this, we assume that the private-sector wage received by women is only a fraction of what is paid by firms. Similar to the literature on the misallocation of resources, we introduce an exogenous wedge, $0 \leq \alpha \leq 1$, that acts as a tax on women's wages. We hence define female private-sector wages as follows:[13]

$$w_{p,f} = (1-\alpha)w^*_{p,f}. \tag{16}$$

Finally, the value of a vacancy for a firm is given by

$$rV_p = -\kappa + q(\theta_p)[\psi^p J_m + (1-\psi^p)J_f], \tag{17}$$

where κ is the cost of creating a vacancy, $q(\theta_p)$ the probability of finding a worker, and ψ^p the fraction of men among the unemployed in the private sector. Given $J_m = J_f$, firms do not discriminate between hiring a man or a woman. Hence, the free-entry condition that pins down tightness in the private sector is

$$\frac{\kappa}{q(\theta_p)} = \frac{y(1-\beta)}{r + \delta_p + \tau + \lambda(1 - F_j(\hat{x}_{p,m}))}. \tag{18}$$

[12] In Albanesi and Sahin [2018] the gender wage gap disappears in the 1996 calibration of the model. The authors point out that this "is due to the fact that the rise in women's labor force attachment causes their quit rates to get closer to men's. In the model, when quit rates are similar, the value associated to hiring male and female workers also converges, causing the gender wage gap to decrease. In the data, a substantial gender wage gap still remains, suggesting that the remaining gap is most likely due to factors absent in our model." (pg 61).

[13] Alternative approaches assume that women are less productive than men (see Erosa et al [2022]) or that they have lower bargaining power than men. Both would result in firms discriminating between male and female hires, and hence, different from our model, would lead to segregated labor markets by gender.

We assume a Cobb-Douglas matching function for the private sector, $m(\theta_p) = \theta_p q(\theta_p) = \zeta \theta_p^\eta$.

3.5 Government

The government employs $\bar{e}_g$ workers and must hire enough individuals to compensate for those who retire, exogenously separate into unemployment or inactivity, or endogenously separate into inactivity. We assume the following matching function for the government $M_g = \min\{v_g, u_g\}$, with its precise functional form being irrelevant as we set vacancies to match the observed employment level.

The number of vacancies in the public sector is given by those who retire or separate:

$$v_g = \bar{e}_g[\tau + \varphi_g(\delta_g + \lambda(1 - F_m(\hat{x}_{g,m}))) + (1 - \varphi_g)(\delta_g + \lambda(1 - F_f(\hat{x}_{g,f})))], \tag{19}$$

where φ_g is the fraction of men in public employment, which is endogenous. The most important difference across the two sectors relates to the creation of jobs. Private-sector job creation is governed by a free-entry condition that pins down market tightness, and hence vacancies posted respond to the number of unemployed searching. Vacancies in the public sector, on the other hand, although they are endogenous, are determined by the government's employment target and do not respond to market tightness, i.e. $m(\theta_g) = p_g$.

Following our empirical findings we assume the government pays an exogenous premium, π_j over private-sector wages, different for men and women

$$w_{g,m} = \pi_m w_{p,m}, \tag{20}$$

$$w_{g,f} = \pi_f w_{p,f}. \tag{21}$$

Although the wage premium is exogenous, female public-sector wages are endogenous as they are a function of female private-sector wages which are partly determined by differences in men's and women's participation decisions (Equation 15).[14] The government exogenously sets hours for public-sector workers, $\xi_g \leq 1$.

3.6 Initial choice of sector

Once born, at rate τ, men and women chose which sector to enter. They compare the expected values of non-employment in the two sectors: $\max\{NE_{p,j}(x); NE_{g,j}(x) + \epsilon_j\}$, where

[14] Exogenous public-sector wage premia are a common modeling choice in the literature on public employment. An extensive literature from the 1970s documents how public-sector wages are used to satisfy unions or other interest groups, or to perform redistribution or to win elections, all aspects exogenous to the labor market; see the discussion in Garibaldi et al. [2021].

ϵ_j denotes individuals' preferences for public-sector jobs. Thresholds for the choice of sector, different for men and women, are hence given by

$$\epsilon_j^* = NE_{p,j}(x) - NE_{g,j}(x),\ j = m, f, \tag{22}$$

and the fraction of men and women entering the public-sector labor market are

$$1 - \Xi_m(\epsilon_m^*), \tag{23}$$

$$1 - \Xi_f(\epsilon_f^*). \tag{24}$$

Without heterogeneity in preferences for public-sector jobs the selection of workers into each sector would only be driven by aggregate variables. As a result, in a world of gender and sector symmetry, the share of women in the public sector would be undetermined. If the public sector increased wages for women slightly, more women would join, lowering the job-finding probability for all workers. Hence, all men would then prefer the private sector. The only possible equilibrium would be one where only women would queue in the public sector. Heterogeneity in preferences allows us to generate equilibria where both men and women enter the public sector.

3.7 Definition of steady-state equilibrium

Definition 1. *A steady-state equilibrium is a set of thresholds* $\{\bar{\epsilon}_f, \bar{\epsilon}_m, \hat{x}_{g,m}, \hat{x}_{p,m}, \hat{x}_{g,f}, \hat{x}_{p,f}\}$*, job-finding probabilities* $\{m(\theta_p), p_g\}$*, stocks of inactive* $\{i_{p,m}, i_{p,f}, i_{g,m}, i_{g,f}\}$*, unemployed* $\{u_{p,m}, u_{p,f}, u_{g,m}, u_{g,f}\}$*, employed* $\{e_{p,m}, e_{p,f}, e_{g,m}, e_{g,f}\}$*, and private and public-sector wages* $\{w_{p,f}, w_{p,m}, w_{g,f}, w_{f,m},\}$*, such that, for given government policies* $\{\pi_m, \pi_f, \bar{e}_g, \xi_g\}$ *and an exogenous "wedge" for female private-sector wages* $\{\alpha\}$*:*

1. *Private sector firms post vacancies according to Equation 18.*
2. *Male private-sector wages are set according to Equation 12.*
3. *Female private-sector wages prior to a"wedge" are determined by Equation 15.*
4. *Newborn men and women decide optimally which sector to join (Equation 22).*
5. *Workers decide optimally the threshold value of x according to Equation 8.*
6. *Worker flows in and out of the three stocks are constant; Equations 9 to 11 hold.*
7. *The total population adds up to 1 (0.5 men, 0.5, women):*

- $\frac{1}{2}(1-\Xi_m(\bar{\epsilon}_m)) = i_{g,m} + u_{g,m} + e_{g,m}$
- $\frac{1}{2}\Xi_m(\bar{\epsilon}_m)) = i_{p,m} + u_{p,m} + e_{p,m}$
- $\frac{1}{2}(1-\Xi_f(\bar{\epsilon}_f)) = i_{g,f} + u_{g,f} + e_{g,f}$
- $\frac{1}{2}\Xi_f(\bar{\epsilon}_f)) = i_{p,f} + u_{p,f} + e_{p,f}$.

3.8 Mechanisms behind the over-representation of women

A lower gender wage gap in the public sector, reflected in a higher wage premium for women, $\pi_f > \pi_m$, contributes to their over-representation. Suppose π_f increases. Then women's wages and their flow utility from working in the public sector increase, raising the value of employment $E_{g,f}$, but also the option value of non-employment $NE_{g,f}$. Hence, when choosing which sector to enter, more women (with lower public-sector motivation) will turn to the public sector, lowering the threshold ϵ^*_f. Furthermore, a higher value of employment $E_{g,f}$, raises the threshold for women to become inactive, $\hat{x}_{g,f}$. More women in the public sector and fewer of them quitting voluntarily into inactivity, means that the government needs to open fewer vacancies to replace them. More women queuing for jobs and fewer open vacancies make it harder to find a job, which dampens the initial increase in the value of non-employment in the public sector for women and reduces it for men. Suddenly, and without any change to their own wages, men find public-sector jobs harder to find, which feeds back into a higher threshold ϵ^*_m; (i.e. only men with a passion for the public sector will tolerate the lower job-finding rate). Crowding out of men, in turn, shortens the queues, further attracting more women and amplifying the original effect. How many more women enter and how many men leave the sector in equilibrium depends on the distribution of their preferences. A higher variance implies that fewer individuals will change sectors.

Better reconciliation of work and family life. Consider a reduction in working hours in the public sector ξ_g, which benefits men and women alike. This increases the flow utility from working and values of employment and non-employment in the public sector. Whether this will attract more men or women depends on the distributions of outside options. If employed women have on average higher opportunity costs of working the desirability of work-life balance will be stronger for them. Furthermore, as queues in the public sector increase, lowering the job-finding rate, the appeal of the public sector for men is reduced. If this crowding-out effect is strong men might prefer the private sector despite longer working hours, amplifying the over-representation of women.

Greater job security. Safer public sector jobs, namely a lower separation rate δ_g, has two effects. First, it increases the value of employment, and indirectly the value of non-employment, raising the thresholds to become inactive in the public sector, $\hat{x}_{g,j}$. Second, fewer separations imply a lower turnover, with the government having to hire fewer replacements. More individuals in the public sector and fewer of them inactive, together with fewer vacancies, decrease the job-finding rate in the public sector, partially offsetting the initial increase in the value of non-employment. Whether the effects are asymmetric across genders depends how job separation interacts with the values of employment and non-employment. If employed women have higher opportunity costs of working and lower wages than men, they might benefit less from safer jobs. In this case, and against our initial intuition, greater job security could reduce the over-representation of women.

Differences in intrinsic preferences for the public sector. If men and women had equal preferences for working in the public sector, the over-representation of women could only be driven by different job characteristics across the two sectors. Assuming preference distributions with different means for men and women will mechanically affect the gender composition of the public sector. The variance of these distributions, assumed equal across genders, determines the strength of the crowding-out effects.

4 Calibration

We calibrate our model separately to the four countries from our empirical analysis. Some parameters are set exogenously based on outside information while the remaining are calibrated to match data moments. We assume men and women draw values of home production x from cumulative exponential distribution functions $F_m(\mu_{x,m})$ and $F_f(\mu_{x,f})$ with different means. We assume normally distributed preferences for working in the public sector, $\Xi_m(\tilde{\epsilon}_m, \sigma^\epsilon)$ and $\Xi_f(\tilde{\epsilon}_f, \sigma^\epsilon)$, with different means by gender but a common standard deviation.

Table 3 displays all parameters for each country. We set the parameter r to match a 4 percent annual interest rate and τ to match a working life of 40 years. As we have monthly data for the US and quarterly data for the other countries, parameters r and τ are three times smaller for the US. For public-sector wage premia and hours discounts, we use results for full-time workers (see columns (2) of Table 2). Numbers for public employment are from our empirical analysis. We normalize the matching efficiency and the time cost of working in the private sector to 1, and we set the elasticity of the matching function with respect to unemployment to the typical value of 0.5; see Petrongolo and Pissarides [2001]. We also

Table 3: Baseline calibration

	US	UK	France	Spain	
Parameters set exogenously					**Source**
Discounting					
Interest rate (r)	0.004	0.012	0.012	0.012	Annual interest rate of 4%
Death rate (τ)	0.002	0.006	0.006	0.006	Working life of 40 years
Public sector policies					
Wage (men) (π_m)	0.971	1.041	0.891	1.007	Wage regressions
Wage (women) (π_f)	1.054	1.052	0.902	1.073	Wage regressions
Employment - (e_g)	0.107	0.157	0.128	0.094	Census Data
Labor market parameters					
Matching efficiency (ζ)	1	1	1	1	Normalization
Matching elasticity (η)	0.5	0.5	0.5	0.5	Petrongolo and Pissarides [2001]
Time cost of labor force					
Private (ξ_p)	1	1	1	1	Normalization
Public (ξ_g)	0.972	0.964	0.918	0.944	Hours regressions
Arrival rate of shocks					
Separation - private (δ_p)	0.015	0.014	0.021	0.042	P-U flow, aggregate
Separation - public (δ_g)	0.007	0.005	0.008	0.021	G-U flow, aggregate
Calibrated parameters					**Target**
Bargaining power men (β)	0.929	0.971	0.964	0.949	Overall unemployment
Labor market parameters					
Vacancy costs (κ)	3.704	1.080	1.514	1.929	Equivalent to 8 weekly wages
"Wedge" (α)	0.265	0.208	0.180	0.236	Private sector gender wage gap
Outside option distribution: Exponential					
Mean men: $\mu_{x,m}$	0.635	0.573	0.736	0.686	Non-employment rate, men
Mean women: $\mu_{x,f}$	0.706	0.897	0.915	0.853	Non-employment rate, women
Arrival rate of shocks					
Outside option (λ)	0.084	0.082	0.066	0.102	E-I flow, aggregate
Preference distribution: Normal					
Mean - men ($\tilde{\epsilon}_m$)	-98.975	-58.369	-14.182	-14.034	Job finding public/private sector
Mean - women ($\tilde{\epsilon}_f$)	-87.125	-25.598	-8.582	-12.183	% women in public sector
Std. - all (σ_ϵ)	89.961	55.047	16.867	10.989	Slope of regional variation in over-representation

Note: The model is calibrated at monthly frequency for the US and at a quarterly frequency for the remaining countries.

take the separation rates in the private and public sector directly from the data.[15]

We calibrate the remaining nine parameters to match nine data moments. In our model, all parameters affect all targets but some calibrated parameters are directly related to data moments. The bargaining power of men (β) is set to match the overall unemployment rate in the economy. Following Gomes [2015] we identify the cost of posting vacancies κ, by matching firms' expected vacancy costs equal to eight weekly wages. The different US value is due to the difference in data frequency. The exogenous "wedge" on female private-

[15] We compute the continuous arrival rate, $\lambda^* = -\ln(1-\lambda)$, which generates slightly different numbers compared to those in Table 1.

sector wages α, is linked to the resulting private-sector gender wage gap, and hence it is calibrated to a lower (higher) value in France (the US) where the gap is smallest (largest). Our calibrated values for α are close to the observed gender wage gaps, indicating that, as in Albanesi and Sahin [2018], the model's endogenous mechanism to generate these gaps only explains a small fraction.

We target the means of the opportunity cost distribution to match non-employment rates for men and women. Given stark differences in part-time work between genders, we consider full-time equivalent non-employment rates which better capture gender differences in the valuation of time spent not working. These numbers are averages of data from the OECD [2010-2021]. For coherency this requires adjustment in all employment numbers, slightly reducing (increasing) public employment (the unemployment rate). The calibrated means of the exponential distributions are higher for women than for men, ranging from 11% higher in the US to 24% in France and Spain, up to 57% in the UK. With the exception of France, means for men are higher in countries with higher non-employment rates, potentially reflecting cross-country differences in preferences for leisure.[16] Figure C.2 in Appendix C displays the distributions of outside options together with the thresholds for men and women in the two sectors. For the US, the two distributions are almost indistinguishable while they are more distinct in France and Spain and especially in the UK. Differences in the distributions for women are also a reflection of differences in child care costs and availability, paid parental leaves, but also the prevalence of single motherhood and welfare systems. For instance, according to the OECD [2024] single motherhood is most prevalent in the US followed by the UK, France, and Spain, but only the UK and somewhat less so France provide fairly generous benefits to single mothers working part-time which could in part explain the low full-time participation rates by women in these countries.[17] Thresholds in both sectors are lower for women compared to men in all countries. With the exception of France, the only country in our sample where both men and women earn less in the public compared to the private sector, thresholds for moving into non-employment are higher in the public compared to the private sector. We set the arrival rates of shocks to the opportunity costs of working λ to match aggregate flows from employment to inactivity.

Finally, there are three parameters related to the distribution of preferences for working in the public sector: the two means and the standard deviation which is assumed to be

[16] In France the fact that the wedge on women's wages is relatively low implies that matching fairly large non-employment rates requires higher means for both genders.

[17] Other related statistics display a less clear relationship to the calibrated distributions for women. Parental leaves are longest in the UK but with lower average payments (39 weeks at 30%) compared to France and Spain (16 weeks at 100% each) and the US (none), while enrollment in early childcare is most prevalent in France, followed by the UK, Spain, and the US; see OECD [2024].

Table 4: Targets: model vs. data

Targets	**US**		**UK**		**France**		**Spain**	
	Data	Model	Data	Model	Data	Model	Data	Model
Unemployment rate								
$(u_m + u_f)/(1 - i_m) + (1 - i_f))$	0.071	0.071	0.065	0.065	0.099	0.099	0.186	0.186
Non-employment rates, FTE								
Male $(i_m + u_m)$	0.252	0.252	0.200	0.200	0.315	0.315	0.338	0.340
Female $(i_f + u_f)$	0.418	0.414	0.450	0.451	0.473	0.473	0.522	0.518
Private-sector wage gap								
$w_f^p/w_m^p - 1$	-0.284	-0.285	-0.219	-0.219	-0.186	-0.186	-0.247	-0.246
Nr. of weekly wages- exp. cost vacancy								
$\kappa\Theta^{1-\eta}/(W_{mp}/12)$	8.000	8.004	8.000	8.009	8.000	8.009	8.000	8.005
Flow rate								
$E \rightarrow I$, aggregate	0.023	0.023	0.021	0.021	0.021	0.021	0.029	0.029
Public-sector employment shares ratio								
$(e_f^g/(e_f^p + e_f^g))/(e_m^g/(e_m^p + e_m^g))$	1.427	1.427	2.187	2.187	1.744	1.744	1.504	1.503
Ratio probability job finding public/private								
$p_g/m(\theta_p)$	1.066	1.065	0.743	0.742	0.847	0.846	0.878	0.878
Regional variation in public sector size								
and women's over-representation	0.004	0.004	0.002	0.002	0.007	0.007	0.011	0.011
Non-targeted moments								
Unemployment rates								
Male $(u_m/(1 - i_m))$	0.071	0.061	0.063	0.052	0.091	0.090	0.165	0.170
Female $(u_f/(1 - i_f))$	0.071	0.083	0.067	0.084	0.109	0.110	0.214	0.207
Inactivity rates								
Male $(i_m))$	0.195	0.203	0.146	0.156	0.246	0.247	0.207	0.205
Female $(i_f))$	0.373	0.361	0.410	0.401	0.409	0.408	0.392	0.392
Aggregate gender wage gap	-0.268	-0.275	-0.215	-0.210	-0.183	-0.194	-0.237	-0.236
Flow rates								
$P \rightarrow I$,men	0.019	0.017	0.016	0.013	0.019	0.016	0.021	0.021
$P \rightarrow I$, women	0.028	0.031	0.030	0.034	0.027	0.027	0.044	0.041
$G \rightarrow I$, men	0.015	0.017	0.015	0.011	0.014	0.017	0.016	0.019
$G \rightarrow I$, women	0.020	0.028	0.018	0.031	0.018	0.027	0.027	0.036

equal for both genders. First, the mean preference for men helps us to match the ratio of (conditional) job finding rates in the public compared to the private sector (equal to $p_g/m(\theta_p)$). In the data, this ratio is equivalent to the ratio of unemployment duration of new hires in the private over that of new hires in the public sector, which we observe in our microdata. This statistic is smallest in the UK, followed by France, Spain, and the US, the only country where it is larger than one meaning that the unemployment duration is lower in the public sector. Second, we set the mean preference for women to match the female over-representation in the public sector as described by the ratio of public employment shares. For all countries, mean preferences for public-sector jobs are higher for women. To pin down the standard deviation we need an additional target related to the mass of individuals being moved by changes in government policies. Ideally we would target some causal effects of policies but we do not have suitable data to do so. Furthermore, the empirical literature is scarce and, as far as we are aware, there are no natural experiments we can use. Instead we consider regional variation and quantify how the share of women changes with the size of the public sector.[18] We regress the share of female public employment by region on each region's total size of the public sector (number of public-sector workers over working-age population). We find negative correlations in all four countries, with coefficients ranging from -0.0016 in the UK to -0.0114 in Spain, see Figure C.1 and Table C.1 in Appendix C. In the model we then use these coefficients targeting the change in the share of women in the public sector caused by a 1 percent increase in public employment.

Table 4 displays our model statistics next to the targeted data moments. Data moments are matched well with an average absolute percentage deviation of less than 0.15 percent. Regarding non-targeted moments, the model generates higher unemployment rates for women than for men which is true in France and Spain but not for the UK and the US. Inactivity rates by gender, on the other hand, are close to the data. Including public-sector wages leads to slightly lower aggregate gender wage gaps in all countries in model and data, with the exception of France where wages in the public sector are lower for both men and women, and where the model generates a somewhat larger aggregate gender wage gap.[19] Finally, regarding flows from private and public employment to inactivity the model captures fairly well the ranking of flows, being larger for women than men and lower in the public compared to the private sector.

5 Examining public sector policies

5.1 Counterfactual Experiments

We run counterfactual experiments shutting down differences across sectors and genders, and comparing the resulting over-representation of women to the one in our benchmark economy. The first experiment focuses on sector differences. To quantify the contribution of each characteristic of the public sector, we shut down one by one: (i) public-sector wage premia ($\pi_f = \pi_m = 1$), (iii) differences in hours ($\xi_g = \xi_p = 1$), and (iii) differences in job security ($\delta_g = \delta_p$), to finally eliminate (iv) all sector differences. In a second experiment we

[18] While the political autonomy to carry out independent public employment policies in these regions differs across countries, with US states being more independent than French regions, we are not interested in the size of the public sector by region per se. Instead we consider how the share of female workers holding public-sector jobs changes with the size of the public sector, linking this to the dispersion in preferences.

[19] Aggregate gender wage gaps are the female coefficient in wage regressions without the interaction effects between gender and public sector, see Table A.2 in the Appendix.

Table 5: Gender composition in public sector under different scenarios

Panel A: Sector differences

	Benchmark	No wage difference $\pi_f = \pi_m = 1$	No hours difference $\xi_g = \xi_p = 1$	No job security difference $\delta_g = \delta_p$	No sector difference $\pi_f = \pi_m = 1$ $\xi_g = \xi_p$ $\delta_g = \delta_p$
Public employment shares ratio					
US	1.427	1.210(50.7%)	1.420(1.5%)	1.432(-1.3%)	1.211*(50.5%)
UK	2.187	2.189*(-0.2%)	2.164(1.9%)	2.201(-1.2%)	2.201*(-1.2%)
France	1.744	1.667(10.4%)	1.708(4.9%)	1.758(-1.8%)	1.625(16.0%)
Spain	1.503	1.152(69.8%)	1.445(11.6%)	1.532(-5.6%)	1.130(74.2%)

Panel B: Gender differences

	Benchmark	No preference difference $\epsilon_f = \epsilon_m$	No x distribution difference $\mu_f = \mu_m$	No wedge $\alpha = 1$	No gender difference $\mu_f = \mu_m$ $\alpha = 1$ $\epsilon_f = \epsilon_m$
Public employment shares ratio					
US	1.427	1.167*(60.8%)	1.449(-5.2%)	1.482(-13.0%)	1.216(49.4%)
UK	2.187	1.007*(99.4%)	2.419(-19.5%)	2.273(-7.3%)	1.015*(98.7%)
France	1.744	1.084*(88.8%)	1.751(-0.9%)	1.708(4.8%)	1.039(94.8%)
Spain	1.503	1.123(75.5%)	1.660(-31.0%)	1.802(-59.2%)	1.402(20.1%)

Note: Model simulations. Public employment shares ratios defined as $\frac{e_{g,f}}{e_f} / \frac{e_{g,m}}{e_m}$. *Eliminating all sector and gender differences leads to public employment shares ratios of 1. In brackets we report the % of over-representation explained; e.g. US wage differences explain (0.427-0.210)/0.427= 50.7% of women's public employment shares ratio deviations from 1. Percentages do not necessarily add up because of interaction effects. In some counterfactuals (marked with *) the government cannot fill all its vacancies so the size of the public sector is adjusted.*

check how much of women's over-representation is due to gender differences, i.e. regarding preferences for working in the public sector, as well as differences in outside options and the "wedge" on women's wages.

Table 5 displays the results from these experiments. In our empirical analysis, with the exception of France, we estimated positive public-sector wage premia for women that were higher than those for men. But even in France the wage discount was lower for women. When we eliminate the wage premia in all countries the over-representation of women in the public sector is reduced, except in the UK where the difference in public-sector wage premia between men and women is smallest. The fact that women earn relatively higher wages in the public sector explains 51 and 70 percent of their over-representation in the US and Spain, and it explains around 10 percent in France.

Our empirical analysis also revealed fewer average annual hours worked in the public compared to the private sector, ranging from discounts of 8.2 percent in France, 5.6 percent in Spain, 3.6 percent in the UK, to 2.8 percent in the US. In the third column where we consider a counterfactural scenario without differences in hours worked between sectors we observe little changes to the gender composition of the public sector in the UK and the US. But in Spain and France fewer working hours in the public sector account for 5 to 12 percent of women's over-representation. We then equate job-separation rates between sectors. Eliminating differences in job security increases the over-representation of women in all countries, albeit marginally. Removing all sector differences in the last column of Panel A accounts for 74, 51 and 16 percent of the over-presentation in Spain, the US, and France respectively. The remaining over-representation of women is due to gender differences which we disentangle in Panel B. For the UK the first experiment which imposes the same preferences for men and women comes close to generating perfect gender symmetry across sectors. This reveals that gender differences in preferences for public-sector jobs almost entirely explain the female over-representation in the UK. In France, Spain, and the US, on the other hand, preferences play a smaller role. Eliminating the "wedge" on women's wages or gender differences in outside options on the other hand makes women more similar to men in terms of participation and wages. Hence, for given gender differences in preferences these experiments lead to more women joining the public sector.

In each country different driving forces matter for the over-representation of women. The fact that in the UK gender differences in preferences explain almost all of it is consistent with our empirical observation of an important reduction in the over-representation of women when excluding health care and education. In the US both preferences and relatively higher wages matter almost equally. In France, next to an important role for preferences, sector differences in wage premia and hours are clearly relevant. The large hours discount of 8.2% in the French public sector contributes half as much as the lower gender wage gap. For Spain, we find that the lower gender wage gap together with better reconciliation of work and family life in the public sector contribute as much as gender differences in preferences. This importance of work-life balance is not surprising. For instance, during the summer weekly hours in the Spanish public sector are effectively reduced from 37 to 32.5 hours.[20]

Perhaps the most surprising result is that public-sector job security contributes negatively to the over-representation of women. In a setting where individuals are risk neutral, as men earn higher wages and have, on average, a lower opportunity cost of working, unemployment is more costly for them. Hence men are more attracted by the job-security aspect of the

[20] Between mid July and mid September, instead of working 8 hours from Monday to Thursday and 5 hours on Friday, many can opt to work 6.5 hours between 8.00-15.00, Monday through Friday (see law BOE-A-2019-2861).

public sector.

5.2 Sensitivity Analysis

A key parameter determining the gender composition of the public sector under different scenarios is the variance of the preference distribution for public-sector jobs. We identified this parameter using coefficients obtained from a cross-regional regression of the share of women in public employment on the size of the public sector. To assess the sensitivity of our quantitative results with respect to this parameter we re-calibrate our model using these estimated coefficients plus and minus one standard error, see Tables C.6 and C.7 in Appendix C. A higher coefficient requires a smaller variance of the preference distribution, implying that more individuals react to a policy change. We repeat the previous experiments, shutting down different model features and assessing the resulting over-representation of women in the public sector. Table 6 displays the results for the counterfactuals regarding sector differences. While the magnitude of each driver changes – becoming larger with a larger slope coefficient and corresponding lower variance – the order of importance of each factor is maintained.[21]

5.3 Heterogeneity by education

There exists ample evidence that the wage benefits of working in the public sector depend on workers' education (see e.g. Gomes [2018], Chassamboulli and Gomes [2023] or Garibaldi et al. [2021]). In line with this evidence we find that college-educated men have lower wage premia in the public sector compared to men without a college degree. The same holds true for women, except for the US where the premium is about the same across education categories (Table A.3 in Appendix A.)[22] Perhaps more important, the gap between men's and women's wage premia in the public sector is always larger among college compared to non-college educated workers. This suggests that wage differentials might be particularly important to explain the over-representation of college-educated women.

Regarding hours worked, the empirical evidence is less clear. While college educated full-time workers have lower hours discounts in the US and Spain, these are higher in France and the UK. To understand how differences in public-sector working conditions across educational

[21] We also provide a sensitivity analysis for the matching elasticity (η), set to 0.5 in our benchmark calibration. We re-calibrate our model for two alternative values, of 0.3 and 0.4, found in the literature. The results displayed in Table C.4 are almost indistinguishable from our benchmark results. As can be seen in Tables C.8 and C.9 in the Appendix which display these alternative calibrations, lower values for the matching elasticity only affect the value of the vacancy costs. All other parameters hardly change.

[22] Our estimates for the two education groups for Spain are in line with those by Couceiro de León and Dolado [2023] who find public-sector wage premia for women at the 25 percentile to be more than twice as large as those at the 75 percentile.

Table 6: Sensitivity Analysis: Gender composition in public sector under different scenarios

Panel A: Sector differences

	Benchmark	No wage difference $\pi_w = \pi_m = 1$	No hours difference $\xi_g = \xi_p = 1$	No job security difference $\delta_g = \delta_p$	No sector difference $\pi_w = \pi_m = 1$ $\xi_g = \xi_p$ $\delta_g = \delta_p$
Public employment shares ratio					
+std	1.428	1.169(60.5%)	1.420(1.9%)	1.434(-1.5%)	1.168*(60.7%)
US	1.427	1.210(50.7%)	1.420(1.5%)	1.432(-1.3%)	1.211*(50.5%)
-std	1.427	1.228(46.5%)	1.421(1.3%)	1.433(-1.5%)	1.230*(46.1%)
+std	2.187	2.187*(-0.0%)	2.156(2.6%)	2.201(-1.2%)	2.206*(-1.6%)
UK	2.187	2.189*(-0.2%)	2.164(1.9%)	2.201(-1.2%)	2.201*(-1.2%)
-std	2.185	2.190*(-0.4%)	2.169(1.3%)	2.200(-1.3%)	2.197*(-1.0%)
+std	1.742	1.546(26.4%)	1.604(18.6%)	1.798(-7.5%)	1.415(44.1%)
France	1.744	1.667(10.4%)	1.708(4.9%)	1.758(-1.8%)	1.625(16.0%)
-std	1.744	1.689(7.5%)	1.729(2.0%)	1.755(-1.5%)	1.673(9.5%)
+std	1.509	0.505(197.1%)	1.252(50.6%)	1.687(-34.9%)	0.471(203.8%)
Spain	1.503	1.152(69.8%)	1.445(11.6%)	1.532(-5.6%)	1.130(74.2%)
-std	1.504	1.376(25.4%)	1.483(4.1%)	1.522(-3.5%)	1.371(26.4%)

Note: Model simulations. Public employment shares ratios defined as $\frac{e_{g,f}}{e_f} / \frac{e_{g,m}}{e_m}$. *Eliminating all sector and gender differences leads to public employment shares ratios of 1. In brackets we report the % of over-representation explained. Percentages do not necessarily add up because of interaction effects. In some counterfactuals (marked with *) the government cannot fill all its vacancies so the size of the public sector is adjusted. +std (-std) refers to a model re-calibrated to a larger (smaller) slope of regional variation in over-representation; for Panel B on gender differences see Table C.3.*

groups affect the driving forces for the over-representation of women, we re-calibrate our model separately for college and non-college educated individuals for all four countries.[23]

Table 7 displays the results from our experiments for college and non-college educated workers. Across the four countries, the over-representation of women among college educated is mainly driven by relatively higher wages and fewer working hours. With the exception of the UK, sector differences account for a much higher fraction of the over-representation than gender differences. On the other hand, for non-college educated individuals differential characteristics of the public sector play a less important role, with gender differences in preferences mattering most.

[23] Tables C.10 to C.13 in the Appendix show these calibrations and Figure C.3 displays the calibrated distributions for individuals' outside options. Following Gomes [2018] we adjust vacancy costs to be lower for non-college compared to college educated individuals; 5 and 10 weekly wages respectively.

Table 7: Gender composition of the public sector for different education groups

Panel A: Sector differences

	Benchmark	No wage difference $\pi_w = \pi_m = 1$	No hours difference $\xi_g = \xi_p$	No job security difference $\delta_g = \delta_p$	No sector difference $\pi_w = \pi_m = 1$ $\xi_g = \xi_p$ $\delta_g = \delta_p$
Public employment shares ratio					
US					
College	1.550	1.187(66.1%)	1.173*(68.6%)	1.551(-0.2%)	1.172(68.6%)
Non-college	1.296	1.292*(1.4%)	1.293(0.8%)	1.299(-1.1%)	1.298*(-0.9%)
UK					
College	1.977	1.949*(2.9%)	1.947*(3.1%)	1.981(-0.4%)	1.958*(2.0%)
Non-college	1.098	1.106*(-8.3%)	1.094(4.0%)	1.103(-5.0%)	1.111*(-13.1%)
France					
College	1.693	2.582(-128.3%)	1.332*(52.1%)	1.716 (-3.4%)	1.246(64.5%)
Non-college	1.676	1.632(6.6%)	1.660(2.4%)	1.690(-2.1%)	1.628(7.2%)
Spain					
College	1.544	0.210(245.1%)	1.422(22.6%)	1.523(3.9%)	0.181(250.5%)
Non-college	1.306	1.197(35.6%)	1.273(10.8%)	1.328(-7.0%)	1.187(38.8%)

*Note: Model simulations. Public employment shares ratios defined as $\frac{e_{g,f}}{e_f} / \frac{e_{g,m}}{e_m}$. Eliminating all sector and gender differences leads to public employment shares ratios of 1. In brackets we report the % of over-representation explained. Percentages do not necessarily add up because of interaction effects. In some counterfactuals (marked with *) the government cannot fill all its vacancies so the size of the public sector is adjusted. Panel B eliminating gender differences is shown in Table C.5 in Appendix C*

5.4 Quantifying the value of public sector characteristics

The fact that certain characteristics of public-sector jobs such greater job security or better work-life balance do not turn out to be the main drivers behind women's over-representation in the public sector does not imply that women do not value these aspects. It merely indicates that individuals' valuation for them does not differ as much across gender as other factors such as preferences or wages. We use our model to calculate individuals' valuation, asking how much of their earnings male and female private-sector workers would be willing to sacrifice to: (i) work the same number of hours and (ii) enjoy the same job security as public-sector workers.

Regarding (i), a private-sector worker with opportunity costs x would be willing to sacrifice in terms of his wage the additional time gained, evaluated at his opportunity cost, $(1 - \xi_g)x/w_{p,j}$. To obtain the aggregate compensating differential for hours we consider the average individual who prefers working to inactivity, by taking the expected value of x, conditional on being employed, in percentage of private-sector wages

$$PremiumH_j^p = \frac{(1-\xi_g)\int_0^{\hat{x}_{p,j}} x f_j(x)dx}{F_j(\hat{x_{p,j}})}\frac{1}{w_{p,j}} \times 100, j = [m, f].$$

For estimating (ii), the job security compensating differential, consider a private-sector worker with wage $w_1 = w_{p,j}$, job-separation rate δ_p and opportunity cost of working $x < \hat{x}_{p,j}$. If offered a job with separation rate δ_g, to maintain the value of employment, the worker would only be willing to accept if paid a wage of at least w_2, where $w_2 = w_1 + \delta_p(U_{p,j}(x|\delta_g) - E_{p,j}(x|\delta_p)) - \delta_g(U_{p,j}(x|\delta_g) - E_{p,j}(x|\delta_g))$. After manipulating and integrating over x we calculate the conditional expected value:

$$\begin{aligned} PremiumS_j^p = & \left(\frac{\delta_p - \delta_g}{r+\tau+\lambda+\delta_p+m(\theta_p)}\right) \\ & \times \left[F(\hat{x}_{p,j})\hat{x}_{p,j} - \int_0^{\hat{x}_{p,j}} x f_{p,j}(x)dx\right]\frac{1}{F(\hat{x}_{p,j})w_{p,j}} \times 100, \end{aligned} \tag{25}$$

for $j = [m, f]$. Different compensating differentials for men and women in the private sector regarding working hours or job security are thus driven by gender differences in (i) the distribution of the opportunity costs of working, f_j, (ii) the threshold in the private sector defining a worker who is indifferent between working or being inactive, $\hat{x}_{p,j}$, and (iii) private-sector wages $w_{p,j}$.

Table 8 displays the compensating differentials for female and male private-sector workers. In each country women are willing to pay more for fewer working hours, and men are willing to pay more for job security. In the UK and France, female workers value fewer hours more than job security. The opposite is true in the US and Spain but for different reasons. In the US, the calibrated distributions of outside options for men and women are very similar, and hence we estimate a low hours premium for women, while in Spain where unemployment is highest, women's estimated job security premium is relatively high. In the US and the UK, hours and job security premia are equal to 0.8 to 1.4 percent of wages while numbers for Spain and France are higher, 1.5 to 3.3 percent. Numbers are similar when evaluating these differentials through the lens of a public-sector worker (see Table C.14 in Appendix C).

Aggregating the estimated numbers we find that these additional benefits of the public sector are significant. They represent a premium equivalent to around 2 to 3 percent of wages in the UK and the US, and close to 5 percent in Spain and France.

Table 8: Value of public sector job characteristics

	Perspective of a private sector worker			
	Hours premium $[\xi_p = \xi_g]$		**Job security** $[\delta_{p,j} = \delta_{g,j}]$	
	Women	Men	Women	Men
US	0.839	0.798	0.994	1.041
UK	1.446	0.946	0.806	1.023
France	3.335	2.692	1.541	1.785
Spain	2.203	1.844	2.843	3.293

Notes: Model simulations; percentages of private-sector wages that men and women are willing to give up for working fewer hours or greater higher job security (as in the public sector).

5.5 Effects of public-sector wages and employment

The over-representation of women in the public sector implies that government wage and employment policies have different effects for male and female labor market outcomes. To quantify this, we consider separately an increase in public wages of 1 percent and an increase in public employment of 1 percent. Table 9 displays the changes in male and female unemployment, inactivity rates, and in the aggregate gender wage gap compared to our benchmark economy.

Higher wages increase male and female unemployment, as more individuals and in particular more women decide to search for public-sector jobs. More individuals queuing in a sector where job creation does not respond to labor market conditions, increases the unemployment rate. The negative effect on the unemployment rate is between 2 to 5 times larger for women than for men. Inactivity rates particularly for women decrease as threshold values for non-employment in the public sector increase. Higher public-sector wages reduce the aggregate gender wage gap, and more in countries like the UK (around 0.14 percentage points) with larger public sectors and more pronounced over-representation of women.

Increasing public employment on the other hand reduces unemployment, for men and even more so for women because the probability to find a job increases. Similar to a wage raise increasing public employment reduces the size of the private sector. However, unlike a wage raise, additional jobs have a direct job-creation effect which is larger than the crowding-out effect on private employment, and hence unemployment falls. Again the magnitude of the effect is between 2 to 4 times larger for women compared to men. Effects on inactivity rates and the aggregate gender wage gap are rather small. Overall, our findings are in line with evidence regarding different effects on men's and women's labor market outcomes of

Table 9: Effects of public sector policies for different countries

Policy	**US**	**UK**	**France**	**Spain**
Panel A: Increase of wages by 1 percent				
Δ unemployment rate male	0.11	0.08	0.11	0.09
Δ unemployment rate female	0.27	0.39	0.34	0.21
Δ inactivity rate male	-0.11	-0.09	-0.10	-0.08
Δ inactivity rate female	-0.15	-0.21	-0.15	-0.10
Δ aggregate wage gap	-0.05	-0.14	-0.10	-0.03
Panel B: Increase of employment by 1 percent				
Δ unemployment rate male	-0.06	-0.07	-0.05	-0.03
Δ unemployment rate female	-0.13	-0.31	-0.13	-0.05
Δ inactivity rate male	0.03	0.04	0.02	0.00
Δ inactivity rate female	0.04	0.06	0.02	0.00
Δ aggregate wage gap	-0.01	-0.01	0.00	-0.01

Notes: Model simulations; percentage point changes due to increases in public sector wages (Panel A) and employment (Panel B).

US fiscal policies, see Bonk and Simon [2022].

5.6 Discussion of alternative modeling assumptions

To keep the model tractable we abstract from potentially important dimensions. We briefly discuss how our simplifying assumptions can be reconciled with data, and how they condition our findings.

Risk aversion We consider agents with linear utility. We conjecture that introducing risk aversion would lead to larger estimates of how men and women value public-sector job security. Potential differences in risk aversion between men and women are currently captured by differences in preferences for working in the public sector. Explicitly including gender differences in risk aversion would most likely reduce the role for preferences, and it could potentially reverse our result on the negative role of greater public-sector job security for explaining the over-representation of women. However, while there exists some experimental evidence that women are more risk averse than men (see e.g. Eckel and Grossman [2008]), these findings are not conclusive (see e.g. Filippin and Crosetto [2016]).

Wage earnings profiles We model unique wages for men and women in the public and private sector, rather than wage-tenure profiles. While Postel-Vinay and Turon [2007] emphasize that lifetime earning in the public sector might be lower than static wage comparisons suggest, Bradley et al. [2017] find that differences in lifetime earnings and static wages across sectors are rather similar for both men and women. Given the relatively constant gap in public employment between men and women along the life cycle – see Figure A.4 in Appendix

A – we conjecture that explicitly introducing wage-tenure profiles into our model would not significantly alter our findings regarding drivers behind the over-representation of women.

Sector switching We rule out direct or indirect transitions between the public and private sector which are uncommon in Spain and France but somewhat more common in the US and the UK. Moreover, switches from public to private employment are similar for men and women, but women are more likely than men to switch from private to public employment. Though these last hazard rates are small, if we allowed individuals to switch sectors, in combination with gender differences in the opportunity costs of working, such a model might imply a larger role for work-life balance for explaining the over-representation of women.

6 Conclusion

Women are over-represented in public employment. To understand why, we build a model where men and women decide if to participate and whether to enter private or public-sector labor markets. We calibrate our model separately to the United States, the United Kingdom, France, and Spain to quantify how much different characteristics of public employment contribute to the over-representation of women in the public sector.

Sector differences explain a significant part of the over-representation of women, ranging from 16% in France, 51% in the United States to 74% in Spain. The most important determinant are relatively higher public-sector wages for women. Work-life balance, as captured by fewer working hours in the public sector only matter in France and Spain. In the United Kingdom, preferences for public-sector jobs play the most important role. Maybe surprisingly, greater job security attracts more men than women, something we confirm when calculating how much private-sector workers are willing to sacrifice for fewer working hours and greater job security. When considering individuals with different levels of education, we find for college (non-college) educated that wages (preferences) are the main driver behind the over-representation of women.

Our estimated compensating differentials for public-sector job characteristics and our findings on the gendered impacts of public-sector wage and employment policies are important for policy makers. First, governments should be aware that wage or employment policies have 2 to 5 times larger effects on female compared to male unemployment. Second, when discussing increases or cuts to public-sector wages it is commonly argued that job security and better work-life balance provide compensating differentials. However, to the best of our knowledge only few have calculated these forms of compensation. Our results indicate that, depending on the country, private sector workers would be willing to give up between 2 to 5 percent of their wages to enjoy such benefits.[24]

Our findings open up a variety of questions for future research. In light of empirical findings on sector switches upon child birth (see Pertold-Gebicka et al. [2016]), explicitly modeling women's participation and fertility choices would allow for the study of public-sector wage and employment policies on fertility. Another interesting question, from a micro rather than a macro perspective, would be to disentangle women's preferences for public service from their preferences to work in public-sector occupations. Given that the latter is closely linked to individuals' specialization choices, incorporating this aspect into our model would require modeling education decisions prior to entering private or public-sector labor markets.

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[24] Fontaine et al. [2020] report similar numbers for job security considering risk-neutral agents under different unemployment scenarios. Numbers in Gomes and Wellschmied [2020] who propose a model with assets and risk aversion are slightly larger. More has been done in terms of calculating compensating differentials including pensions. Danzer and Dolton [2012] use UK survey data to estimate total reward differentials, including current earnings, pensions, hours of work, paid holidays, employer provided health care and probability of unemployment for highly educated individuals. Gomes and Wellschmied [2020] estimate the value of pension premia over the life-cycle for public sector workers with different levels of education. Bradley et al. [2017] incorporate differences in career progression into the comparison between public and private sector compensation.

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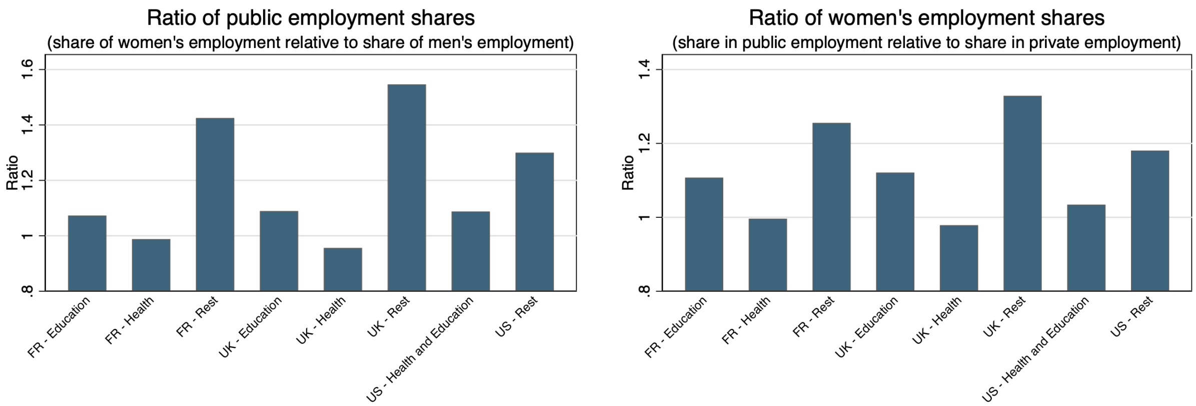
COMPANION APPENDIX

You're the one that I want! Understanding the Over-Representation of Women in the Public Sector by Pedro Gomes and Zoë Kuehn

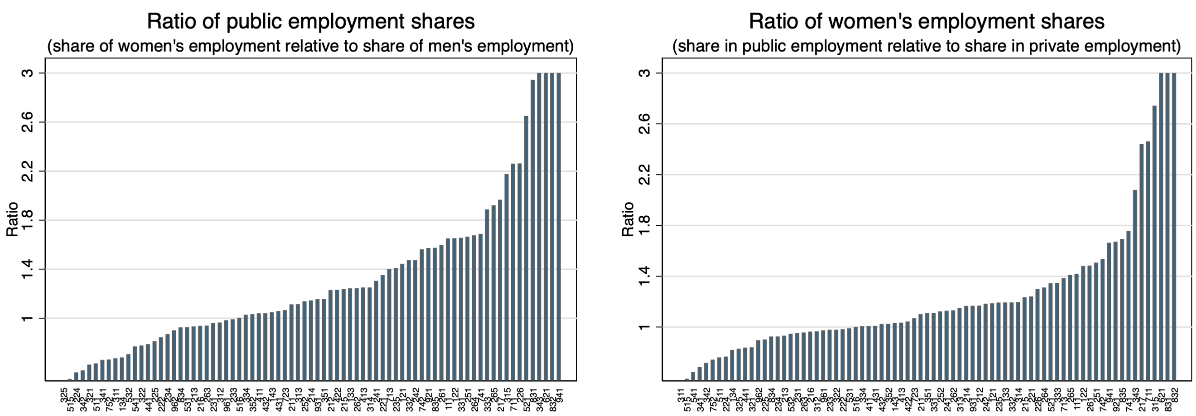
A Empirical Analysis, additional results

Figure A.1: Over-representation of women in public employment by industry and occupation

Industries

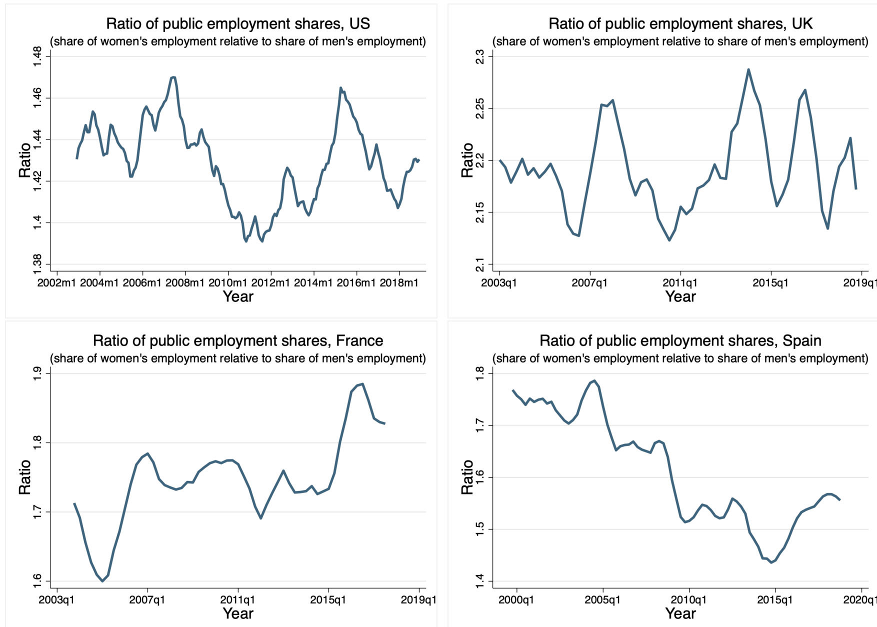


3-digit Occupations



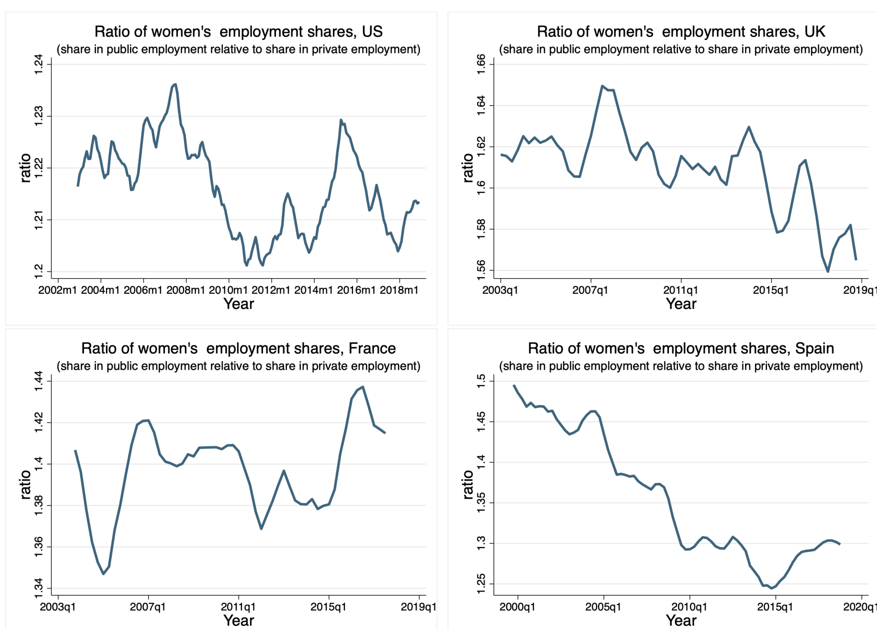
Note: the 1st panel uses data from the French Labor Force Survey [2003-2017] and the UK Labour Force Survey [2003-2018] and the CPS [2003-2018], extracted by Fontaine et al. [2020]. The Spanish Labor Force Survey does not allow for a disaggregation of public employment by industry. The 2nd panel shows CPS data, averages over 1996-2017. 3-digit occupations that have an overall share of public-sector employment between 0.05 and 0.95. The ratios were capped at 3 for readability.

Figure A.2: Public employment shares ratio, time variation



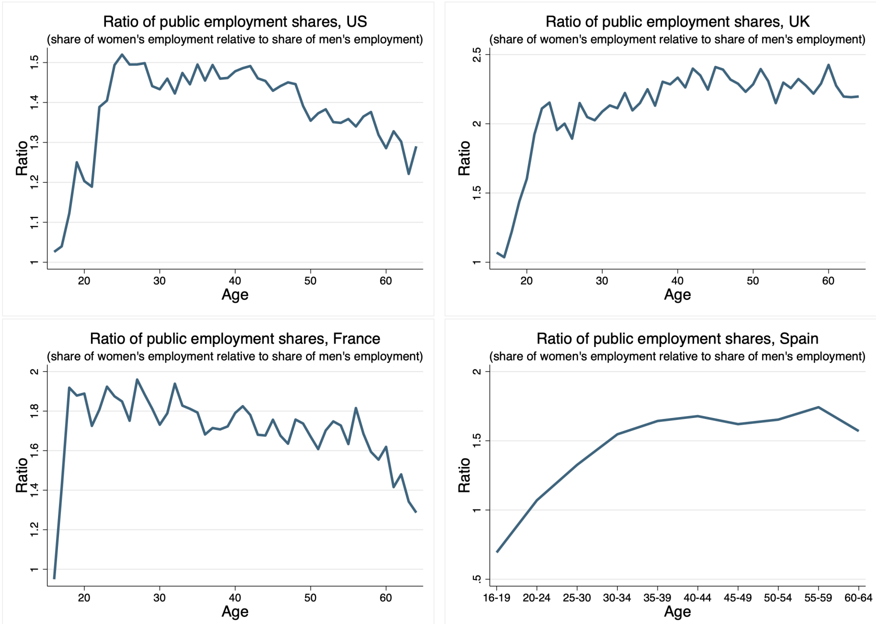
Note: CPS [2003-2018], UK LFS [2003-2018], French LFS [2003-2017], and Spanish LFS [2003-2018], extracted by Fontaine et al. [2020].

Figure A.3: Ratio of women's employment shares, time variation



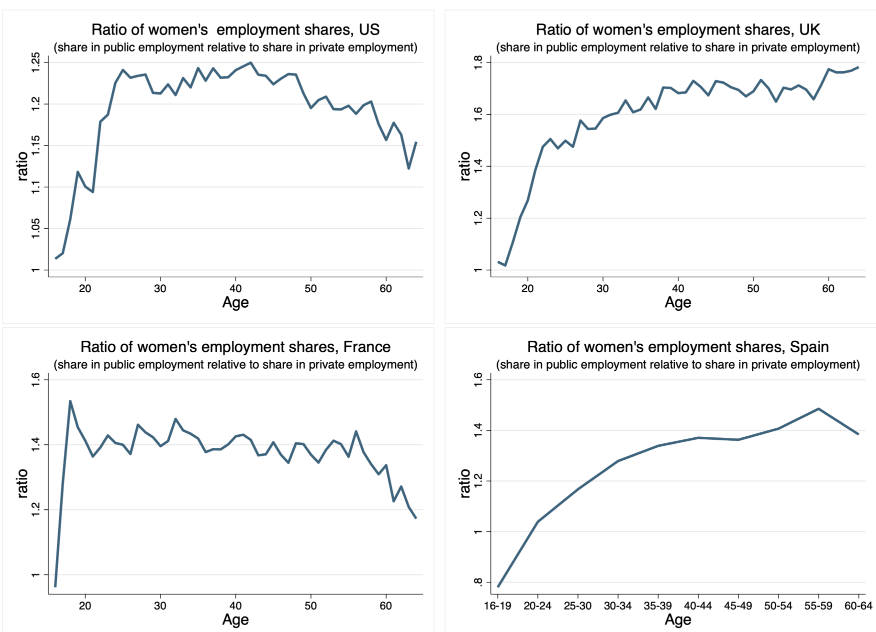
Note: CPS [2003-2018], UK LFS [2003-2018], French LFS [2003-2017], and Spanish LFS [2003-2018], extracted by Fontaine et al. [2020].

Figure A.4: Public employment shares ratio, variation over age groups



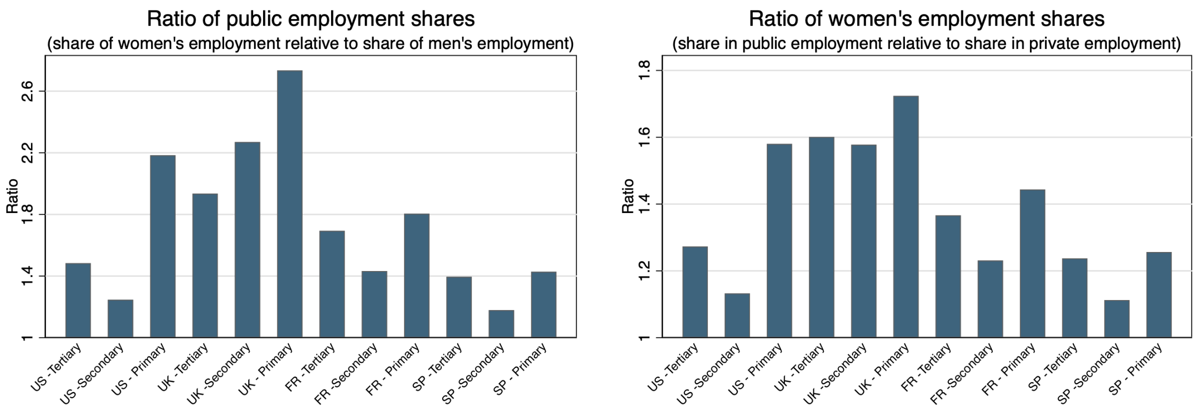
Note: CPS [2003-2018], UK LFS [2003-2018], French LFS [2003-2017], and Spanish LFS [2003-2018], extracted by Fontaine et al. [2020].

Figure A.5: Ratio of women's employment shares, variation over age groups



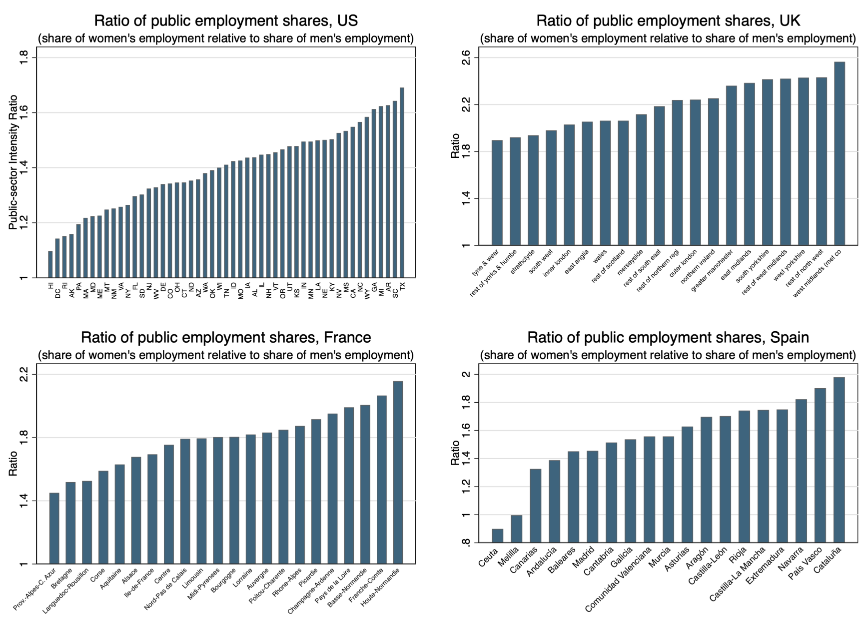
Note: Note: CPS [2003-2018], UK LFS [2003-2018], French LFS [2003-2017], and Spanish LFS [2003-2018], extracted by Fontaine et al. [2020].

Figure A.6: Over-representation of women in public employment, by education



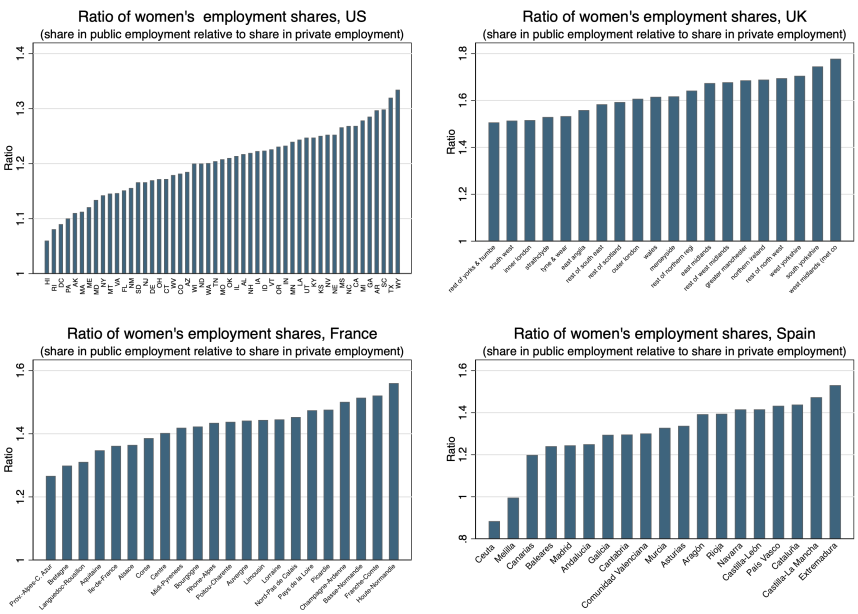
Note: CPS [2003-2018], UK LFS [2003-2018], French LFS [2003-2017], and Spanish LFS [2003-2018], extracted by Fontaine et al. [2020].

Figure A.7: Public Employment Share Ratios, regional variation



Note: CPS [2003-2018], UK LFS [2003-2018], French LFS [2003-2017], and Spanish LFS [2003-2018], extracted by Fontaine et al. [2020].

Figure A.8: Ratio of women's employment shares, regional variation



Note: CPS [2003-2018], UK LFS [2003-2018], French LFS [2003-2017], and Spanish LFS [2003-2018], extracted by Fontaine et al. [2020].

Table A.1: Incidence of part time by sector, gender and education

	US	**UK**	**France**	**Spain**
	(1)	(2)	(3)	(4)
Incidence part-time	0.219	0.287	0.152	0.147
Public sector	0.194	0.325	0.174	0.098
Private sector	0.224	0.272	0.146	0.154
Men	0.151	0.132	0.074	0.074
Women	0.288	0.435	0.251	0.258
Men in public sector	0.134	0.122	0.081	0.081
Men in private sector	0.153	0.135	0.073	0.073
Women in public sector	0.238	0.427	0.26	0.113
Women in private sector	0.301	0.44	0.247	0.286
Among college educated	0.158	0.216	0.123	0.116
Public sector	0.161	0.263	0.163	0.11
Private sector	0.157	0.19	0.11	0.117
Men	0.095	0.093	0.065	0.075
Women	0.218	0.332	0.2	0.162
Men in public sector	0.118	0.113	0.094	0.109
Men in private sector	0.088	0.086	0.057	0.068
Women in public sector	0.189	0.338	0.229	0.11
Women in private sector	0.232	0.327	0.188	0.181
Among non-college educated	0.249	0.321	0.17	0.16
Public sector	0.228	0.371	0.182	0.086
Private sector	0.251	0.306	0.167	0.167
Men	0.177	0.151	0.08	0.074
Women	0.323	0.486	0.28	0.306
Men in public sector	0.149	0.129	0.071	0.058
Men in private sector	0.18	0.156	0.082	0.075
Women in public sector	0.293	0.493	0.281	0.116
Women in private sector	0.329	0.484	0.28	0.33
Nr. observations	1,071,617	622,013	722,571	876,348

Note: Structure of Earnings Survey [2002, 2006, 2010, 2014] for France, UK, and Spain; CPS March Supplement [2003-2018] for US, see Appendix B for details.

Table A.2: Aggregate gender wage gap

	US		**UK**		**France**		**Spain**	
	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Panel A: wage regressions								
Aggregate gender wage gap								
	-0.276***	-0.268***	-0.182***	-0.216***	-0.152***	-0.183***	-0.205***	-0.237***
	(-170.08)	(-165.17)	(-90.96)	(-110.12)	(-103.31)	(-131.46)	(-167.16)	(-203.68)
Controls								
Demographics	X	X	X	X	X	X	X	X
Region year FE	X	X	X	X	X	X	X	X
Reg., year	X	X	X	X	X	X	X	X
Job characteristics	X	X	X	X	X	X	X	X
Part time dummy	X		X		X		X	
Only fll time wkr		X		X		X		X
Obs.	1,071,617	837,088	622,013	443,801	722,571	612,527	876,348	747,302
R-squared	0.516	0.403	0.561	0.387	0.496	0.453	0.599	0.538

Note: Estimated by OLS regressions of the log of gross yearly earnings on a female dummy, a dummy for working in the public sector, controlling for region, year, occupation, education, age groups, part-time, tenure and tenure squared. The aggregate gender wage gap corresponds to the coefficient on the female dummy. Data for UK, France, and Spain from the Structure of Earnings Survey [2002, 2006, 2010, 2014]; for the US from the CPS March Supplement [2003-2018]. Demographic controls include age, education, and race (only for the US). Job characteristics include occupation and for European countries tenure and tenure squared.

Table A.3: Public-sector wage and hours premium and private-sector gender wage gap: For College (C) and Non-college (NC) educated individuals.

	US		**UK**		**France**		**Spain**	
	(C)	(NC)	(C)	(NC)	(C)	(NC)	(C)	(NC)
				Panel A: wage regressions				
Public sector wag premium								
Men	-0.088***	0.034***	0.01**	0.061***	-0.207***	-0.044***	-0.061***	0.048***
	(-20.35)	(8.43)	(2.06)	(15.28)	(-57.22)	(-16.52)	(-16.53)	(16.94)
Women	0.043***	0.042***	0.031***	0.067***	-0.188***	-0.036***	0.038***	0.107***
	(10.79)	(11.24)	(6.97)	(16.76)	(-46.56)	(-12.20)	(11.06)	(34.26)
Gender wage gap								
Private	-0.264***	-0.287***	-0.215***	-0.219***	-0.199***	-0.162***	-0.245***	-0.246***
	(-90.69)	(-129.42)	(-56.84)	(-75.37)	(-75.43)	(-83.02)	(-102.88)	(-167.41)
Controls								
Demographics	X	X	X	X	X	X	X	X
Region Year FE	X	X	X	X	X	X	X	X
Job Characteristics	X	X	X	X	X	X	X	X
Only fll time wkr	X	X	X	X	X	X	X	X
Obs.	296,111	540,977	159,293	284,508	242,034	370,493	222,147	525,155
R-squared	0.331	0.307	0.330	0.365	0.382	0.399	0.481	0.482
				Panel B: hours regressions				
Public	-0.027***	-0.029***	-0.042***	-0.033***	-0.123***	-0.055***	-0.038***	-0.069***
	(-25.41)	(-31.63)	(-57.48)	(-50.07)	(-183.96)	(-121.19)	(-80.15)	(-256.22)
Controls								
Demographics	X	X	X	X	X	X	X	X
Region Year FE	X	X	X	X	X	X	X	X
Job Characteristics	X	X	X	X	X	X	X	X
Only fll time wkr	X	X	X	X	X	X	X	X
Obs.	296,111	540,977	159,293	284,508	242,034	370,493	222,147	525,155
R-squared	0.072	0.066	0.173	0.181	0.261	0.126	0.218	0.335
				Panel C: wage regressions				
Aggregate Gender Wage Gap								
Women	-0.229***	-0.286***	-0.207***	-0.218***	-0.194***	-0.161***	-0.222***	-0.239***
	(-90.29)	(-135.16)	(-67.92)	(-85.30)	(-85.29)	(-92.40)	(-105.69)	(-171.73)
Controls								
Demographics	X	X	X	X	X	X	X	X
Region Year FE	X	X	X	X	X	X	X	X
Job Characteristics	X	X	X	X	X	X	X	X
Only fll time wkr	X	X	X	X	X	X	X	X
Obs.	296,111	540,977	159,293	284,508	242,034	370,493	222,147	525,155
R-squared	0.330	0.307	0.330	0.365	0.382	0.399	0.480	0.482

Note: Estimated by OLS regressions. Panel A regresses the log of gross yearly earnings on a female dummy, a female and male dummy interacted with a dummy for working in the public sector, controlling for region, year, occupation, education, age groups, part-time, tenure and tenure squared. Panel B (panel A) regresses log hours worked on a female dummy, a dummy for working in the public sector, controlling for region, year, occupation, education, age groups, part-time, tenure and tenure squared. In panel A, the public-sector wage premium for men (women) corresponds to the coefficient β_3 (β_4). The private-sector gender wage gap corresponds to the coefficient β_1 from Equation 1. In panel B, the public sector hours premium correspond to the coefficient α_3 from Equation 2. Data for UK, France, and Spain from the Structure of Earnings Survey [2002, 2006, 2010, 2014]; for the US from the CPS March Supplement [2003-2018]. Demographic controls include age, education and race (only for the US). Job characteristics include occupation and for European countries tenure and tenure squared.

B Data description, details

Data sources

CPS and Labor Force Surveys The CPS is conducted on a monthly basis while the other surveys are conducted quarterly. The surveys include individuals' demographic characteristics, as well as information on their labor force status, sector of employment, occupation, industry of employment, weeks worked, and hours per week worked. We restrict our sample to individuals aged 16 to 64. For calculating stocks of unemployed, employed, and inactive individuals we use averages from 2003 to 2018. We define public employment in line with each country's official statistics. For the US, the public sector includes individuals who work for the government (further disaggregated into Federal, State or Local government). In the UK, we include the following categories: i) Central Government, Civil Service; ii) Local government or council (including police, fire services and local authority controlled schools or colleges); iii) University or other grant-funded educational establishments; iv) Health authority or NHS trust; and v) Armed forces. We exclude from our definition every private organization, as well as: i) Public company; ii) Nationalised industry or state corporation; iii) Charity, voluntary organisation or trust; and iv) other organisation. As in Fontaine et al. [2020], we exclude publicly-owned companies because those sell their goods and services and thus face market forces. Including them into private employment, together with non-profit institutions tends to reduce the observed differences between the two sectors. A similar definition is used for France. For Spain, the survey asks directly whether respondents work for the public or the private sector. For the US, we also use CPS data to analyze the gender composition of public sector jobs based on a 3-digit ISCO-08 occupational classification. To this end, we consider only occupations with non-trivial public and private-sector employment, i.e. occupations where the share of the public sector in total employment is larger than 5 percent and smaller than 95 percent. This implies that some top-paid occupations are excluded (i.e. as manufacturing, mining, construction, and distribution managers) as well as some low-paid jobs (i.e. domestic, hotel and office cleaners and helpers or waiters and bartenders).

Structure of Earnings Survey and CPS March Supplement We impose the following sample restrictions for the CPS March Supplement. We eliminate all individuals older than 65 and those currently not working. We also exclude self-employed and unpaid family workers as well as those in agriculture, fishing and forestry as our European data does not extent to these sectors. In particular our measure of annual hours worked is constructed the following way. We consider 260 working days and subtract the number of holidays and then multiply those by daily hours which we obtain dividing usual hours per month by 20. Note that US data does not include information on holidays which is why we use data from the Bureau of Labor Statistics [2022a] on Employee Benefits in the US which provides information on paid vacation for government and private industry workers. This data is provided by workers' fulltime and part time status and tenure. Information on the latter variable is not available in the CPS March Supplement. We do however have data on individuals' age and hence using additional information from the BLS [2022b] on employee tenure by age we assign tenure to individuals of different ages and then assign paid vacations by tenure separately to public and private-sector workers of full and part-time status. Note that in any case for the US, there is very little difference in the estimated public sector hours discount when using actual hours worked or annual hours worked (2.5% vs. 2.8%).

Table B.1: Descriptive Statistics: Samples for wage and hours regressions

	US		**UK**		**France**		**Spain**	
	Mean	**Std.**	**Mean**	**Std.**	**Mean**	**Std..**	**Mean**	**Std.**
Columns (1) - All workers								
Yearly earnings	44684.27	52107.26	22188.46	33097.16	36800.44	36245.00	21800.25	19336.08
Annual hours	2048.95	2613.54	1713.05	593.83	1719.37	368.33	1915.30	402.70
Public sector	0.176	0.381	0.272	0.445	0.231	0.421	0.123	0.328
Women	0.498	0.500	0.509	0.500	0.443	0.497	0.399	0.490
Women public	0.102	0.303	0.181	0.385	0.121	0.326	0.065	0.247
Men public	0.074	0.262	0.091	0.287	0.110	0.313	0.058	0.233
Age	40.03	12.20	40.53	12.48	42.71	10.90	40.04	10.88
College	0.328	0.470	0.327	0.469	0.382	0.486	0.287	0.452
Part time	0.219	0.413	0.287	0.452	0.152	0.359	0.147	0.354
Tenure	–	–	7.17	8.15	11.75	10.59	8.55	9.69
Non-white	0.200	0.400	–	–	–	–	–	
Nr of observations	1,071,617		622,013		722,571		876,348	
Columns (2)- Full time workers only								
Yearly earnings	50444.03	54583.05	27608.57	36245.45	40014.76	37703.08	24106.44	19832.53
Annual hours	2176.01	2224.49	2011.22	328.82	1822.70	239.25	2048.06	153.87
Public sector	0.182	0.386	0.257	0.437	0.225	0.418	0.130	0.336
Women	0.454	0.498	0.402	0.490	0.391	0.488	0.348	0.476
Women public	0.100	0.300	0.145	0.352	0.105	0.307	0.068	0.251
Men public	0.082	0.274	0.112	0.315	0.120	0.325	0.062	0.241
Age	41.10	11.41	40.42	11.83	42.59	10.83	40.13	10.68
College	0.354	0.478	0.359	0.480	0.395	0.489	0.297	0.457
Tenure			7.75	8.49	12.05	10.67	9.05	9.81
Non-white	0.204	0.403						
Nr of observations	837,088		443,801		612,527		747,302	

Note: Data for the US from CPS March Supplement [2003-2018], for the UK, France and Spain from the Structure of Earnings Survey [2002, 2006, 2010, 2014]; see description above for details.

Estimation of conditional transition probabilities

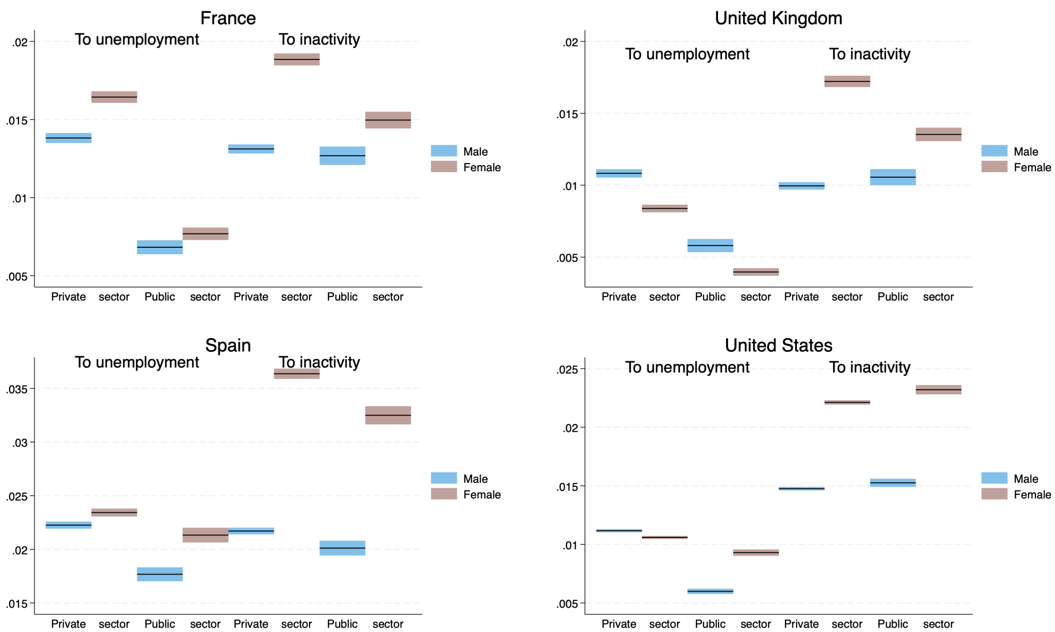
Conditional on being employed, a worker can keep his job, become unemployed or become inactive. We consider staying employed as the base outcome and compute the probabilities of becoming unemployed or inactive as:

$$\lambda_i^U = \frac{\exp(x_i\beta_U)}{1+\exp(x_i\beta_U)+\exp(x_i\beta_I)} \tag{B.1}$$

$$\lambda_i^I = \frac{\exp(x_i \beta_I)}{1 + \exp(x_i \beta_U) + \exp(x_i \beta_I)}, \tag{B.2}$$

where x_i denotes the control variables age and age squared, as well as indicator variables for education, region, year, occupation, and age between 60 and 64 to capture increasing flows into retirement. The estimation also includes a female dummy, a public sector dummy, and an interaction term between the two. These estimates then allow us to predict transition probabilities for the average female and male employee in both public and private sector.

Figure B.1: Conditional transition probabilities out of employment



Note: Based on the estimation of equations B.1 and B.2 using a multinomial logit regression. For France the number of observations is 1,421,243 and the pseudo R-squared is 0.092. For the UK the number of observations is 1,393,928 and the pseudo R-squared is 0.071. For Spain the number of observations is 1,989,672 and the pseudo R-squared is 0.090. For the US the number of observations is 6,479,457 and the pseudo R-squared is 0.068. For France, the UK, and Spain, transition rates are quarterly, while they are monthly for the US. Included as controls are regional and year fixed effects, education and occupation dummies as well as age and age squared and a dummy for age 60-64. The predicted probability is calculated based on an individual with the average characteristics of the employed population. Data is for 2003-2016 (2005-2016 for Spain). The boxes report the 95 percent confidence interval on the prediction.

C Further model results and inputs

C.1 Dis-aggregated value functions

Our two value functions for employment and non-employment can be disaggregated into three value functions for employment, unemployment and inactivity as follows:

$$(r+\tau+\lambda)E_{i,j} = (1-\xi_i)x + w_{i,j} + \delta_i[U_{i,j} - E_{i,j}] + \lambda[A^1_{i,j} + A^2_{i,j}], \tag{C.1}$$

$$(r+\tau+\lambda)U_{i,j} = x + m(\theta_i)[E_{i,j} - U_{i,j}] + \lambda[B^1_{i,j} + A^2_{i,j}], \quad \text{if } x \leq \hat{x}_{i,j} \tag{C.2}$$

$$(r+\tau+\lambda)I_{i,j} = x + \lambda[B^1_{i,j} + A^2_{i,j}], \quad \text{if } x > \hat{x}_{i,j} \tag{C.3}$$

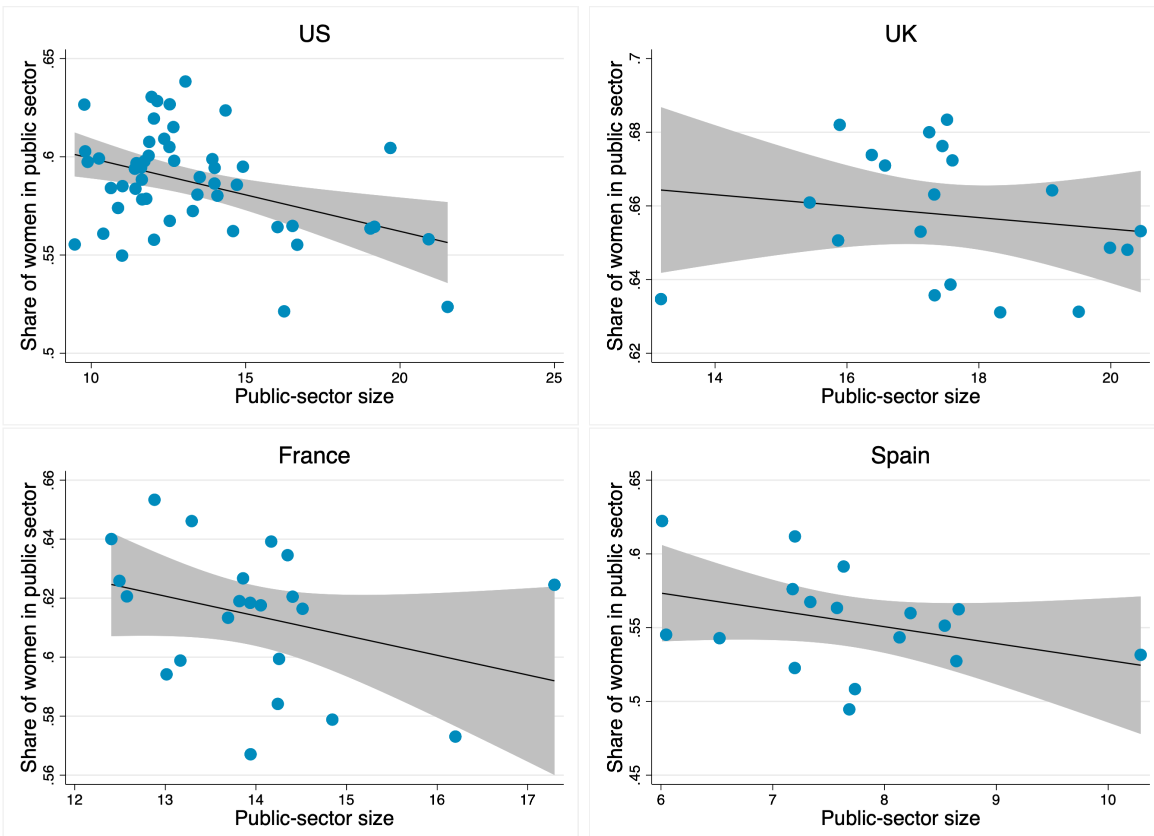
where $A^1_{i,j} = \int_0^{\hat{x}_{i,j}} E_{i,j}(x')dF_j(x')$, $A^2_{i,j} = \int_{\hat{x}_{i,j}}^{\infty} I_{i,j}(x')dF_j(x')$, $A_{i,j} = A^1_{i,j} + A^2_{i,j}$, $B^1_{i,j} = \int_0^{\hat{x}_{i,j}} U_{i,j}(x')dF_j(x')$, and $B_{i,j} = B^1_{i,j} + A^2_{i,j}$.

Table C.1: Cross-regional relationship between share of women in public sector and its size

	US	**UK**	**France**	**Spain**
% public sector employment	-0.0037***	-0.0016	-0.0067	-0.0114
	(0.0012)	(0.0023)	(0.0044)	(0.0078)
Constant	0.6363***	0.6848***	0.7075***	0.6418***
	(0.0159)	(0.0404)	(0.0612)	(0.0606)
Obs.	51	19	22	17
R-squared	0.1701	0.0248	0.1048	0.1242

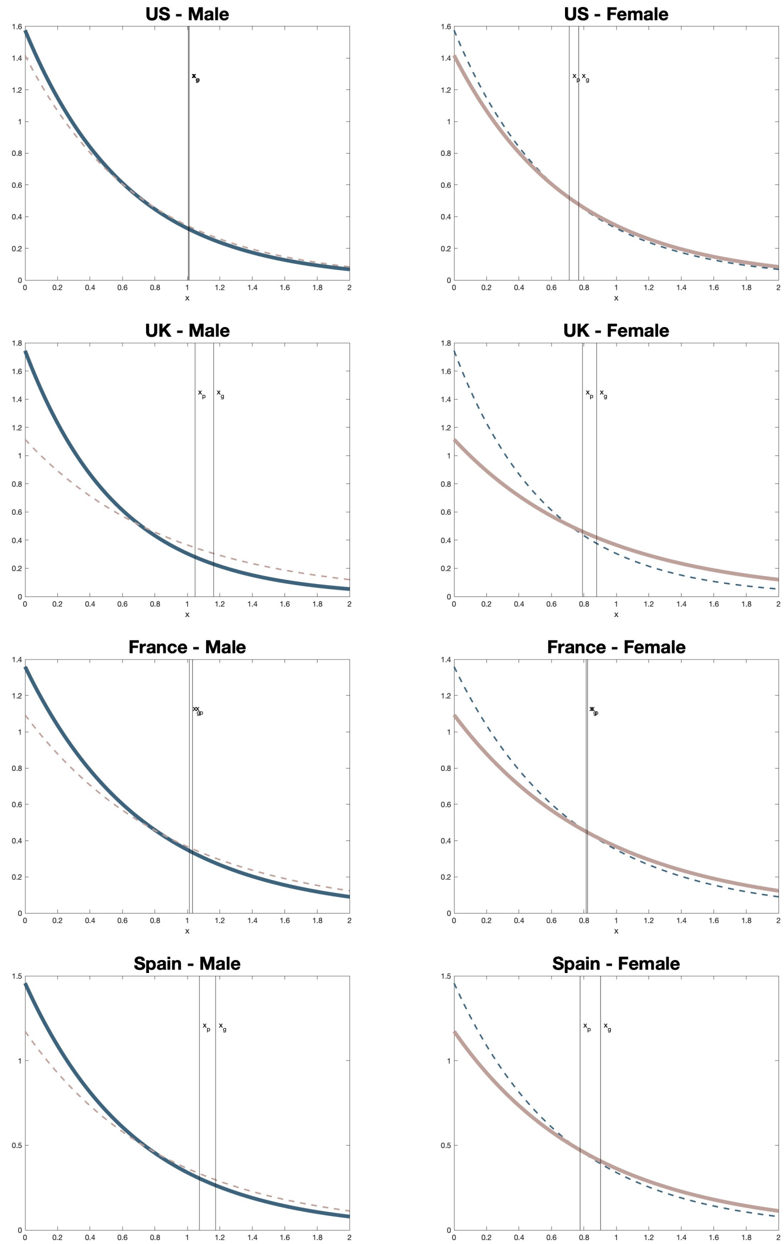
Notes: Estimated by OLS regressions. Data for UK LFS [2003-2018], French LFS [2003-2017] and Spanish LFS [2003-2018]; for US, CPS data [2003-2018], extracted by Fontaine et al. [2020]. For Spain, we exclude two regions – Ceuta and Melilla – characterized by a strong presence of the armed forces due to their location on the African continent.

Figure C.1: Share of women in public sector and the size of government



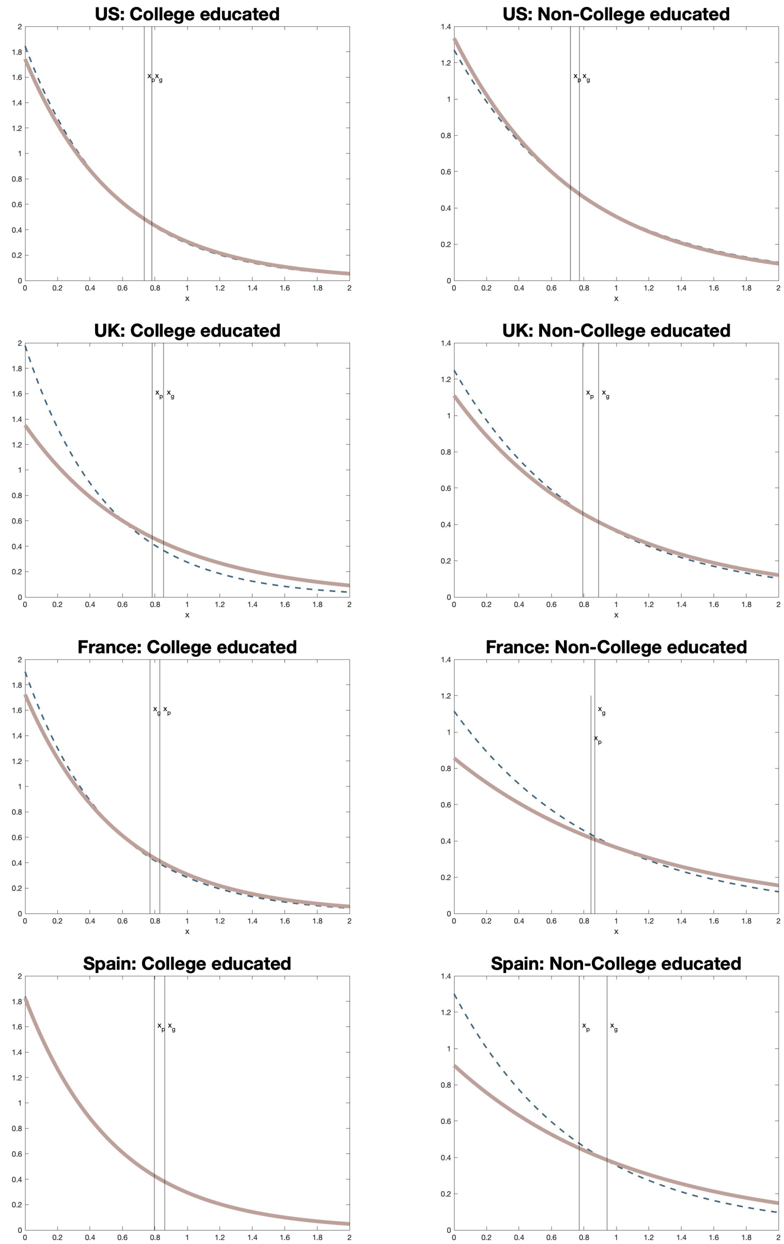
Note: Data for UK LFS [2003-2018], French LFS [2003-2017] and Spanish LFS [2003-2018]; for US, CPS data [2003-2018], extracted by Fontaine et al. [2020]. For Spain, we exclude two regions – Ceuta and Melilla – characterized by a strong presence of the armed forces due to their location on the African continent.

Figure C.2: Calibrated distributions for individuals' outside options, $F(\mu_x^j)$, $j = [m, f]$



Note: The left-hand graphs show the distributions of individuals' outside options together with the different thresholds for men (for comparison the distributions for women are plotted as dashed lines). The right-hand graphs show the distributions of individuals' outside options together with the different thresholds for women (for comparison the distributions for men are plotted as dashed lines). Means of these distributions for men (women) in each country are 0.635 (0.706) for the US, 0.573(0.897) for the UK, 0.736(0.915) for France and 0.686(0.853) for Spain. Thresholds for the public and private sector for men (women) are 1.006 (0.768) and 1.011 (0.709) in the US, 1.162 (0.879) and 1.047 (0.792) for the UK, 1.013 (0.814) and 1.032 (0.823) for France and 1.174 (0.903) 1.075(0.777) for Spain.

Figure C.3: Calibrated distributions for individuals' outside options, $F(\mu_x^j)$, $j = [m, f]$for college and non-college educated individuals



Note: The left-hand graphs show the distributions for college-educated individuals' outside options together with the different thresholds for women (for comparison the distributions for men are plotted as dashed lines). The right-hand graphs show these distributions for non-college educated individuals. Means of these distributions for men (women) with college education in the US are 0.542(0.574) for those without college education 0.787(0.749), in the UK 0.506(0.740) and 0.800(0.901), in France 0.525(0.580) and 0.897(1.169), and in Spain 0.544(0.548) and 0.769(1.103). Thresholds for the public and private sector for men (women) are 0.936 (0.781) and 1.011 (0.734) for college and 1.085 (0.773) and 1.015 (0.717) for non college educated in the US, 1.086(0.853) and 1.014 (0.783) for college and 1.156 (0.891) and 1.031 (0.794) for non college educated in the UK, 0.952 (0.770) and 1.049 (0.831) for college and 1.057 (0.867) and 1.030(0.843) for non college educated in France and 1.062 (0.861) and 1.092 (0.797) for college and 1.275 (0.943) and 1.089 (0.771) for non college educated in Spain.

Table C.2: Gender composition of the public sector under different scenarios, alternative and raw measures for the over-representation of women

Panel A: Sector differences					
	Benchmark	No wage difference $\pi_f = \pi_m = 1$	No hours diff. $\xi_g = \xi_p$	No job security diff $\delta_g = \delta_p$	No sector diff. $\pi_f = \pi_m = 1$ $\xi_g = \xi_p$ $\delta_g = \delta_p$
Women's employment shares ratio					
US	1.250	1.132(47.1%)	1.246(1.3%)	1.252(-1.1%)	1.132*(47.2%)
UK	1.723	1.721*(0.3%)	1.713(1.4%)	1.729(-0.9%)	1.717*(0.8%)
France	1.440	1.404(8.2%)	1.423(3.8%)	1.447(-1.5%)	1.384(12.7%)
Spain	1.301	1.101(66.5%)	1.271(10.0%)	1.316(-4.8%)	1.087(71.1%)
Raw Measures					
Share of public sector in women's employment					
US	0.193	0.178	0.192	0.193	0.176*
UK	0.343	0.340*	0.342	0.344	0.329*
France	0.278	0.274	0.277	0.280	0.274
Spain	0.204	0.178	0.201	0.206	0.177
Share of women in public-sector employment					
US	0.528	0.485	0.527	0.529	0.485*
UK	0.600	0.600*	0.597	0.602	0.600*
France	0.573	0.560	0.567	0.575	0.554
Spain	0.524	0.455	0.513	0.528	0.450
Panel B: Gender differences					
	Benchmark	No preference difference $\epsilon_f = \epsilon_m$	No x distrib. $\mu_f = \mu_m$	No wedge diff $\alpha = 1$	No gender diff. $\mu_f = \mu_m$ $\alpha = 1$ $\epsilon_f = \epsilon_m$
Women's employment shares ratio					
US	1.250	1.105*(58.1%)	1.248(0.4%)	1.243(2.4%)	1.115(53.8%)
UK	1.723	1.005*(99.3%)	1.641(11.4%)	1.671(7.2%)	1.009*(98.7%)
France	1.440	1.057*(87.0%)	1.394(10.4%)	1.381(13.4%)	1.024(94.7%)
Spain	1.301	1.082(72.7%)	1.338(-12.1%)	1.386(-28.2%)	1.200(33.7%)
Raw Measures					
Share of public sector in women's employment					
US	0.193	0.165*	0.187	0.177	0.157
UK	0.343	0.166*	0.302	0.321	0.168*
France	0.278	0.203*	0.256	0.256	0.189
Spain	0.204	0.175	0.196	0.199	0.166
Share of women in public-sector employment					
US	0.528	0.478*	0.550	0.585	0.550
UK	0.600	0.407*	0.685	0.642	0.504
France	0.573	0.456*	0.610	0.600	0.510*
Spain	0.524	0.450	0.588	0.614	0.586

*Note: Model simulations; Women's employment shares ratios defined as $\frac{e_{g,f}}{e_g} / \frac{e_{p,f}}{e_p}$. Eliminating all sector and gender differences leads to women's employment shares ratios of 1. In brackets we report the % of over-representation explained. Percentages do not necessarily add up because of interaction effects. In some counterfactuals (marked with *), the size of the public sector has to be adjusted when the government cannot fill all its vacancies.*

Table C.3: Sensitivity Analysis: Gender composition of the public sector under different scenarios

Panel B: Gender differences

	Benchmark	No preference difference	No x distrib.	No wedge diff	No gender diff.
			$\mu_f = \mu_m$		$\mu_f = \mu_m$
				$\alpha = 1$	$\alpha = 1$
		$\epsilon_f = \epsilon_m$			$\epsilon_f = \epsilon_m$
Public-sector employment shares ratio					
+std	1.428	1.200*(53.2%)	1.457 (-6.7%)	1.504(17.7%)	1.272(36.4%)
US	1.427	1.167*(60.8%)	1.449(-5.2%)	1.482(-13.0%)	1.216(49.4%)
-std	1.427	1.155*(63.7%)	1.452(-5.9%)	1.475(-11.3%)	1.195(54.4%)
+std	2.187	0.997*(100.3%)	2.451(-22.3%)	2.304(-9.9%)	1.020*(98.3%)
UK	2.187	1.007*(99.4%)	2.419(-19.5%)	2.273(-7.3%)	1.015*(98.7%)
-std	2.185	1.014*(98.8%)	2.394(-17.7%)	2.246(-5.2%)	1.012*(99.0%)
+std	1.742	1.738(0.5%)	1.529(28.8%)	1.401(45.9%)	1.191(74.2%)
France	1.744	1.084*(88.8%)	1.751(-0.9%)	1.708(4.8%)	1.039(94.8%)
-std	1.744	1.045*(93.9%)	1.769(-3.4%)	1.740(0.5%)	1.020*(97.3%)
+std	1.509	1.509(0.0%)	2.961(-285.0%)	4.173(-523.0%)	5.176(-719.8%)
Spain	1.503	1.123(75.5%)	1.660(-31.0%)	1.802(-59.2%)	1.402(20.1%)
-std	1.504	1.059(88.3%)	1.556(-10.2%)	1.572(-13.5%)	1.112(77.7%)

Note: Model simulations. Public employment shares ratios defined as $\frac{e_{g,f}}{e_f} / \frac{e_{g,m}}{e_m}$. *Eliminating all sector and gender differences leads to public employment shares ratios of 1. In brackets we report the % of over-representation explained. Percentages do not necessarily add up because of interaction effects. In some counterfactuals (marked with *) the government cannot fill all its vacancies so the size of the public sector is adjusted.*

Table C.4: Sensitivity Analysis: Gender composition of the public sector, different values for matching elasticity (η)

Panel A: Sector differences

	Benchmark	No wage difference $\pi_f = \pi_m = 1$	No hours diff. $\xi_g = \xi_p$	No job security diff $\delta_g = \delta_p$	No sector diff. $\pi_f = \pi_m = 1$ $\xi_g = \xi_p$ $\delta_g = \delta_p$
Public-sector employment shares ratio					
US					
$\eta = 0.3$	1.427	1.212(50.5%)	1.421(1.5%)	1.433(-1.3%)	1.212*(50.3%)
$\eta = 0.4$	1.426	1.211(50.5%)	1.420(1.5%)	1.432(-1.3%)	1.211*(50.4%)
$\eta = 0.5$	1.427	1.210(50.7%)	1.420(1.5%)	1.432(-1.3%)	1.211*(50.5%)
UK					
$\eta = 0.3$	2.188	2.190*(-0.2%)	2.165(1.9%)	2.202(-1.2%)	2.202*(-1.2%)
$\eta = 0.4$	2.187	2.189*(-0.2%)	2.164(1.9%)	2.201(-1.2%)	2.201*(-1.2%)
$\eta = 0.5$	2.187	2.189*(-0.2%)	2.164(1.9%)	2.201(-1.2%)	2.201*(-1.2%)
FR					
$\eta = 0.3$	1.744	1.667(10.4%)	1.708(4.9%)	1.758(-1.9%)	1.625(16.0%)
$\eta = 0.4$	1.744	1.666(10.5%)	1.707(5.0%)	1.757(-1.8%)	1.623(16.2%)
$\eta = 0.5$	1.744	1.667(10.4%)	1.708(4.9%)	1.758(-1.8%)	1.625(16.0%)
ES					
$\eta = 0.3$	1.504	1.180(64.4%)	1.448(11.0%)	1.532(-5.6%)	1.159(68.5%)
$\eta = 0.4$	1.504	1.180(64.3%)	1.448(11.0%)	1.532(-5.6%)	1.159(68.5%)
$\eta = 0.5$	1.503	1.152(69.8%)	1.445(11.6%)	1.532(-5.6%)	1.130(74.2%)

Panel B: Gender differences

	Benchmark	No preference difference $\epsilon_f = \epsilon_m$	No x distrib. $\mu_f = \mu_m$	No wedge diff $\alpha = 1$	No gender diff. $\mu_f = \mu_m$ $\alpha = 1$ $\epsilon_f = \epsilon_m$
Public-sector employment shares ratio					
US					
$\eta = 0.3$	1.427	1.167*(61.0%)	1.449(-5.1%)	1.483(-13.0%)	1.215(49.7%)
$\eta = 0.4$	1.426	1.167*(60.9%)	1.448(-5.1%)	1.481(-12.9%)	1.215(49.6%)
$\eta = 0.5$	1.427	1.167*(60.8%)	1.449(-5.2%)	1.482(-13.0%)	1.216(49.4%)
UK					
$\eta = 0.3$	2.188	1.007*(99.4%)	2.419(-19.5%)	2.274(-7.3%)	1.015*(98.7%)
$\eta = 0.4$	2.187	1.007*(99.4%)	2.418(-19.5%)	2.273(-7.3%)	1.015*(98.7%)
$\eta = 0.5$	2.187	1.007*(99.4%)	2.419(-19.5%)	2.273(-7.3%)	1.015*(98.7%)
France					
$\eta = 0.3$	1.744	1.083*(88.8%)	1.751(-0.9%)	1.708(4.8%)	1.039(94.8%)
$\eta = 0.4$	1.744	1.084*(88.6%)	1.749(-0.8%)	1.706(5.1%)	1.040(94.7%)
$\eta = 0.5$	1.744	1.084*(88.8%)	1.751(-0.9%)	1.708(4.8%)	1.039(94.8%)
Spain					
$\eta = 0.3$	1.504	1.107(78.7%)	1.660(-31.1%)	1.778(-54.5%)	1.363(28.0%)
$\eta = 0.4$	1.504	1.107(78.7%)	1.660(-31.1%)	1.778(-54.5%)	1.363(28.0%)
$\eta = 0.5$	1.503	1.123(75.5%)	1.660(-31.0%)	1.802(-59.2%)	1.402(20.1%)

Note: Model simulations. Public employment shares ratios defined as $\frac{e_{g,f}}{e_f} / \frac{e_{g,m}}{e_m}$. *Eliminating all sector and gender differences leads to public employment shares ratios of 1. In brackets we report the % of over-representation explained. Percentages do not necessarily add up because of interaction effects. In some counterfactuals (marked with *) the government cannot fill all its vacancies so the size of the public sector is adjusted.*

Table C.5: Gender composition of the public sector for different education groups

Panel B: Gender differences

	Benchmark	No preference difference	No x x distribution difference	No wedge	No gender difference
			$\mu_f = \mu_m$		$\mu_f = \mu_m$
			$\bar{x}_{x,f} = \bar{x}_{x,m}$		$\bar{x}_{x,f} = \bar{x}_{x,m}$
				$\alpha = 1$	$\alpha = 1$
		$\epsilon_f = \epsilon_m$			$\epsilon_f = \epsilon_m$
Public employment shares ratio					
US					
College	1.550	1.329*(40.2%)	1.562(-2.1%)	1.583(-6.0%)	1.357(35.0%)
Non-college	1.296	0.991*(102.9%)	1.292(1.3%)	1.331(-12.0%)	1.014(95.1%)
UK					
College	1.977	1.029*(97.1%)	2.135(-16.2%)	2.044(-6.8%)	1.023*(97.6%)
Non-college	1.098	0.982(118.0%)	1.109(-10.9%)	1.126(-28.1%)	1.009(90.7%)
FR					
College	1.693	15.5(-1992.2%)	1.276(60.1%)	0.554 (164.4%)	1.920(-32.7%)
Non-college	1.676	1.003(99.5%)	1.717(-6.0%)	1.702(-3.8%)	1.020(97.1%)
ES					
College	1.544	12.8(-2076.9%)	1.556(-2.0%)	2.343(-146.7%)	13.433(-2184.1%)
Non-college	1.306	1.039(87.2%)	1.385(-25.7%)	1.428(-39.6%)	1.128(58.2%)

*Note: Model simulations. Public employment shares ratios defined as $\frac{e_{g,f}}{e_f} / \frac{e_{g,m}}{e_m}$. Eliminating all sector and gender differences leads to public employment share ratios of 1. In brackets we report the % of over-representation explained. Percentages do not necessarily add up because of interaction effects. In some counterfactuals (marked with *) the government cannot fill all its vacancies so the size of the public sector is adjusted.*

Table C.6: Alternative calibrations: +/- std. error on slope coefficient

	US	UK	FR	ES	
Calibrated parameters					**Target**
+std on slope coefficient					
Bargaining power, men (β)	0.929	0.971	0.964	0.949	Unemployment
Labor market parameters					
Vacancy costs (κ)	3.707	1.081	1.513	1.938	Equivalent to 8 weekly wages
"Wedge" (α)	0.264	0.208	0.180	0.237	Private sector gender wage gap
Outside option distribution: Exponential					
Mean men: $\mu_{x,m}$	0.634	0.573	0.737	0.672	Non-employment rate, men
Mean women: $\mu_{x,f}$	0.708	0.894	0.905	0.944	Non-employment rate, women
Arrival rate of shocks					
Outside option(λ)	0.084	0.082	0.067	0.099	E I flow, aggregate
Preference distribution: Normal					
Mean - men ($\tilde{\epsilon}_m$)	-76.717	-42.487	-0.481	-4.010	Job finding public/private sector
Mean - women ($\tilde{\epsilon}_f$)	-68.898	-18.845	-0.476	-4.010	% women in public sector
Std. - all (σ_ϵ)	69.945	39.575	3.150	2.227	Slope of regional variation in over-representation; + std error
- std on slope coefficient					
Bargaining power, men (β)	0.929	0.971	0.964	0.949	Overall unemployment
Labor market parameters					
Vacancy costs (κ)	3.696	1.077	1.514	1.932	8 weekly wages
"Wedge" (α)	0.261	0.207	0.180	0.237	Gender wage gap
Outside option distribution: Exponential					
Mean men: $\mu_{x,m}$	0.633	0.573	0.736	0.682	Non-employment rate, men
Mean women: $\mu_{x,f}$	0.724	0.911	0.915	0.862	Non-employment rate, women
Arrival rate of shocks					
Outside option (λ)	0.084	0.082	0.066	0.102	E I flow, aggregate
Preference distribution: Normal					
Mean - men ($\tilde{\epsilon}_m$)	-110.562	-77.309	-35.884	-51.430	Job finding public/private sector
Mean - women ($\tilde{\epsilon}_f$)	-96.422	-33.422	-21.400	-41.661	% women in public sector
Std. - all (σ_ϵ)	100.799	73.761	38.574	43.789	Slope of regional variation in over-representation; - std. error.

Note: The model is calibrated at monthly frequency for the US and at quarterly frequency for the other countries.

Table C.7: Alternative calibrations: +/- std. error on slope coefficient - model vs. data

Targets	**US**		**UK**		**FR**		**ES**	
	Data	Model	Data	Model	Data	Model	Data	Model
+std on slope coefficient								
Unemployment rate								
$(u_m + u_f)/(1 - i_m) + (1 - i_f))$	0.071	0.071	0.065	0.065	0.099	0.099	0.186	0.186
Non-employment rates, FTE								
Male $(i_m + u_m)$	0.252	0.252	0.200	0.200	0.315	0.316	0.338	0.332
Female $(i_f + u_f)$	0.418	0.415	0.450	0.450	0.473	0.469	0.522	0.555
Private-sector wage gap								
$w_f^p/w_m^p - 1$	-0.284	-0.285	-0.219	-0.219	-0.186	-0.185	-0.247	-0.249
Nr. of weekly wages- exp. cost vacancy								
$\kappa\Theta^{1-\eta}/(W_{mp}/12)$	8.000	8.001	8.000	8.007	8.000	7.992	8.000	8.001
Flow rate								
$E \to I$, aggregate	0.023	0.023	0.021	0.021	0.021	0.021	0.029	0.029
Public-sector employment shares ratio								
$(e_f^g/(e_f^p + e_f^g))/(e_m^g/(e_m^p + e_m^g))$	1.427	1.428	2.187	2.187	1.744	1.742	1.504	1.509
Ratio probability job finding public/private								
$p_g/m(\theta_p)$	1.066	1.066	0.743	0.743	0.847	0.847	0.878	0.881
Regional variation in public sector size								
and women's over-representation	0.005	0.005	0.004	0.004	0.011	0.011	0.019	0.019
-std on slope coefficient								
Unemployment rate								
$(u_m + u_f)/(1 - i_m) + (1 - i_f))$	0.071	0.071	0.065	0.065	0.099	0.099	0.186	0.186
Non-employment rates, FTE								
Male $(i_m + u_m)$	0.252	0.252	0.200	0.200	0.315	0.315	0.338	0.338
Female $(i_f + u_f)$	0.418	0.423	0.450	0.457	0.473	0.473	0.522	0.522
Private-sector wage gap								
$w_f^p/w_m^p - 1$	-0.284	-0.283	-0.219	-0.218	-0.186	-0.186	-0.247	-0.247
Nr. of weekly wages- exp. cost vacancy								
$\kappa\Theta^{1-\eta}/(W_{mp}/12)$	8.000	8.007	8.000	7.989	8.000	8.009	8.000	7.998
Flow rate								
$E \to I$, aggregate	0.023	0.023	0.021	0.021	0.021	0.021	0.029	0.029
Public-sector employment shares ratio								
$(e_f^g/(e_f^p + e_f^g))/(e_m^g/(e_m^p + e_m^g))$	1.427	1.427	2.187	2.185	1.744	1.744	1.504	1.504
Ratio probability job finding public/private								
$p_g/m(\theta_p)$	1.066	1.066	0.743	0.744	0.847	0.846	0.878	0.878
Regional variation in public sector size								
and women's over-representation	0.003	0.003	-0.001	-0.001	0.002	0.002	0.004	0.004

Table C.8: Alternative calibrations: Different matching elasticities: $\eta = 0.3$, $\eta = 0.4$

	US	UK	FR	ES	
Calibrated parameters					**Target**
$\boldsymbol{\eta = 0.3}$					
Bargaining power, men (β)	0.929	0.971	0.964	0.949	Unemployment
Labor market parameters					
Vacancy costs (κ)	9.289	2.139	4.744	8.568	Equivalent to 8 weekly wages
"Wedge" (α)	0.265	0.208	0.180	0.237	Private sector gender wage gap
Outside option distribution: Exponential					
Mean men: $\mu_{x,m}$	0.635	0.573	0.736	0.682	Non-employment rate, men
Mean women: $\mu_{x,f}$	0.705	0.897	0.915	0.862	Non-employment rate, women
Arrival rate of shocks					
Outside option (λ)	0.084	0.082	0.066	0.102	E- I flow, aggregate
Preference distribution: Normal					
Mean - men ($\tilde{\epsilon}_m$)	-99.451	-58.423	-14.172	-15.281	Job finding public/private sector
Mean - women ($\tilde{\epsilon}_f$)	-87.492	-25.610	-8.576	-13.131	% women in public sector
Std. - all (σ_ϵ)	90.360	55.094	16.856	12.099	Slope of regional variation in over-representation
$\boldsymbol{\eta = 0.4}$					
Bargaining power, men (β)	0.929	0.971	0.964	0.949	Unemployment
Labor market parameters					
Vacancy costs (κ)	5.231	1.395	2.322	3.378	Equivalent to 8 weekly wages
"Wedge" (α)	0.265	0.208	0.180	0.237	Private sector gender wage gap
Outside option distribution: Exponential					
Mean men: $\mu_{x,m}$	0.635	0.573	0.736	0.682	Non-employment rate, men
Mean women: $\mu_{x,f}$	0.705	0.896	0.913	0.862	Non-employment rate, women
Arrival rate of shocks					
Outside option (λ)	0.084	0.082	0.066	0.102	E- I flow, aggregate
Preference distribution: Normal					
Mean - men ($\tilde{\epsilon}_m$)	-99.550	-58.459	-13.848	-15.282	Job finding public/private sector
Mean - women ($\tilde{\epsilon}_f$)	-87.626	-25.651	-8.389	-13.132	% women in public sector
Std. - all (σ_ϵ)	90.466	55.125	16.525	12.100	Slope of regional variation in over-representation

Note: The model is calibrated at monthly frequency for the US and at quarterly frequency for the other countries.

Table C.9: Alternative calibrations with $\eta = 0.3$, $\eta = 0.4$: model vs. data

Targets	**US**		**UK**		**FR**		**ES**	
	Data	Model	Data	Model	Data	Model	Data	Model
$\boldsymbol{\eta = 0.3}$								
Unemployment rate								
$(u_m + u_f)/(1 - i_m) + (1 - i_f))$	0.071	0.071	0.065	0.065	0.099	0.099	0.186	0.186
Non-employment rates, FTE								
Male $(i_m + u_m)$	0.252	0.252	0.200	0.200	0.315	0.315	0.338	0.338
Female $(i_f + u_f)$	0.418	0.414	0.450	0.451	0.473	0.473	0.522	0.522
Private-sector wage gap								
$w_f^p/w_m^p - 1$	-0.284	-0.285	-0.219	-0.219	-0.186	-0.186	-0.247	-0.247
Nr. of weekly wages- exp. cost vacancy								
$\kappa\Theta^{1-\eta}/(W_{mp}/12)$	8.000	8.001	8.000	8.008	8.000	8.009	8.000	7.998
Flow rate								
$E \rightarrow I$, aggregate	0.023	0.023	0.021	0.021	0.021	0.021	0.029	0.029
Public-sector employment shares ratio								
$(e_f^g/(e_f^p + e_f^g))/(e_m^g/(e_m^p + e_m^g))$	1.427	1.427	2.187	2.188	1.744	1.744	1.504	1.504
Ratio probability job finding public/private								
$p_g/m(\theta_p)$	1.066	1.066	0.743	0.743	0.847	0.847	0.878	0.878
Regional variation in public sector size								
and women's over-representation	0.004	0.004	0.002	0.002	0.007	0.007	0.011	0.011
$\boldsymbol{\eta = 0.4}$								
Unemployment rate								
$(u_m + u_f)/(1 - i_m) + (1 - i_f))$	0.071	0.071	0.065	0.065	0.099	0.099	0.186	0.186
Non-employment rates, FTE								
Male $(i_m + u_m)$	0.252	0.252	0.200	0.200	0.315	0.315	0.338	0.338
Female $(i_f + u_f)$	0.418	0.413	0.450	0.451	0.473	0.472	0.522	0.522
Private-sector wage gap								
$w_f^p/w_m^p - 1$	-0.284	-0.285	-0.219	-0.219	-0.186	-0.185	-0.247	-0.247
Nr. of weekly wages- exp. cost vacancy								
$\kappa\Theta^{1-\eta}/(W_{mp}/12)$	8.000	8.001	8.000	8.009	8.000	8.005	8.000	7.998
Flow rate								
$E \rightarrow I$, aggregate	0.023	0.023	0.021	0.021	0.021	0.021	0.029	0.029
Public-sector employment shares ratio								
$(e_f^g/(e_f^p + e_f^g))/(e_m^g/(e_m^p + e_m^g))$	1.427	1.426	2.187	2.187	1.744	1.744	1.504	1.504
Ratio probability job finding public/private								
$p_g/m(\theta_p)$	1.066	1.066	0.743	0.743	0.847	0.847	0.878	0.878
Regional variation in public sector size								
and women's over-representation	0.004	0.004	0.002	0.002	0.007	0.007	0.011	0.011

Table C.10: Additional calibration: US and UK for college and non -college educated

	US		UK		
	College	Non College	College	Non-College	**Source**
Parameters set exogenously					
Discounting					
Interest rate (r)	0.004	0.004	0.012	0.012	Annual interest rate of 4%
Death rate (τ)	0.002	0.002	0.006	0.006	Working life of 40 years
Public sector policies					
Wage (men) (π_m)	0.912	1.034	1.010	1.061	Wage regressions
Wage (women) (π_f)	1.043	1.042	1.031	1.067	Wage regressions
Employment - (e_g)	0.191	0.073	0.269	0.102	Census Data
Labor market parameters					
Matching efficiency (ζ)	1	1	1	1	Normalization
Matching elasticity (η)	0.5	0.5	0.5	0.5	Literature
Share of women	0.546	0.519	0.510	0.509	Census Data
Time cost of labor force					
Private (ξ_p)	1	1	1	1	Normalization
Public (ξ_g)	0.973	0.971	0.958	0.967	Hours regressions
Arrival rate of shocks					
Separation - private (δ_p)	0.007	0.018	0.012	0.017	P-U flow, aggregate
Separation - public (δ_g)	0.005	0.009	0.004	0.006	G-U flow, aggregate
Calibrated parameters					Target
Bargaining power, men (β)	0.948	0.945	0.968	0.978	Unemployment
Labor market parameters					
Vacancy costs (κ)	4.387	2.489	0.856	0.753	10 (college), 5 (non-college) weekly wages
"Wedge" (α)	0.253	0.279	0.206	0.217	Private sector gender wage gap
Arrival rate of shocks					
Outside option (λ)	0.058	0.083	0.069	0.069	E- I flow, aggregate
Outside option distribution:exponential					
Mean men: $\mu_{x,m}$	0.542	0.787	0.506	0.800	Non-employment rate, men
Mean women: $\mu_{x,f}$	0.574	0.749	0.740	0.901	Non-employment rate, women
Preference distribution: Normal					
Mean - men ($\bar{\epsilon}_m$)	-57.755	-197.805	-43.300	-49.879	Job finding public/private sector
Mean - women ($\bar{\epsilon}_f$)	-47.622	-174.266	-2.713	-46.568	% women in public sector
Std. - all (σ_ϵ)	80.454	150.003	63.152	46.307	Slope of regional variation in over-representation

Note: The model is calibrated at monthly frequency for the US and at quarterly frequency for the UK.

Table C.11: Additional calibration: France and Spain for college and non-college educated

	France		Spain		
	College	Non College	College	Non-College	**Source**
Parameters set exogenously					
Discounting					
Interest rate (r)	0.012	0.012	0.012	0.012	Annual interest rate of 4%
Death rate (τ)	0.006	0.006	0.006	0.006	Working life 40 years
Public sector policies					
Wage (men) (π_m)	0.793	0.956	0.939	1.048	Wage regressions
Wage (women) (π_f)	0.811	0.964	1.038	1.107	Wage regressions
Employment - (e_g)	0.208	0.094	0.201	0.050	Census Data
Labor market parameters					
Matching efficiency (ζ)	1	1	1	1	Normalization
Matching elasticity (η)	0.5	0.5	0.5	0.5	Literature
Share of women	0.544	0.496	0.526	0.483	Census Data
Time cost of labor force					
Private (ξ_p)	1	1	1	1	Normalization
Public (ξ_g)	0.877	0.945	0.962	0.931	Hours regressions
Arrival rate of shocks					
Separation - private (δ_p)	0.016	0.023	0.029	0.048	P-U flow, aggregate
Separation - public (δ_g)	0.005	0.010	0.014	0.033	G-U flow, aggregate
Calibrated parameters					Target
Bargaining power, men (β)	0.964	0.975	0.952	0.963	Unemployment
Labor market parameters					
Vacancy costs (κ)	1.529	1.030	2.041	1.301	10 (college), 5 (non-college) weekly wages
"Wedge" (α)	0.194	0.159	0.254	0.237	Private sector gender wage gap
Arrival rate of shocks					
Outside option (λ)	0.077	0.062	0.104	0.109	E- I flow, aggregate
Outside option distribution:exponential					
Mean men: $\mu_{x,m}$	0.525	0.897	0.544	0.769	Non-employment rate, men
Mean women: $\mu_{x,f}$	0.580	1.169	0.548	1.103	Non-employment rate, women
Preference distribution: Normal					
Mean - men ($\bar{\epsilon}_m$)	5.740	-26.126	-0.782	-48.016	Job finding public/private sector
Mean - women ($\bar{\epsilon}_f$)	3.954	-17.966	-3.156	-43.983	% women in public sector
Std. - all (σ_ϵ)	1.500	24.318	2.000	33.499	Slope of regional variation in over-representation

Table C.12: US and UK for college and non-college educated: model vs. data

	US				UK			
Targets	**College**		**Non-college**		**College**		**Non-college**	
	Data	Model	Data	Model	Data	Model	Data	Model
Unemployment rate								
$(u_m + u_f/(1-i_m) + (1-i_f))$	0.037	0.036	0.088	0.088	0.034	0.034	0.078	0.078
Non-employment rates								
Male $(i_m + u_m)$	0.158	0.185	0.331	0.332	0.154	0.154	0.321	0.322
Female $(i_f + u_f)$	0.289	0.301	0.441	0.440	0.363	0.360	0.466	0.459
Private-sector wage gap								
$w^p_f/w^p_m - 1$	-0.265	-0.267	-0.287	-0.287	-0.215	-0.215	-0.219	-0.220
Nr. of weekly wages- exp. cost vacancy								
$\kappa\Theta^{1-\eta}/(W_{mp}/12)$	10.000	10.013	5.000	4.998	10.000	9.984	5.000	4.997
Flow rate								
$E \to I$, aggregate	0.013	0.013	0.027	0.027	0.015	0.015	0.023	0.023
Public-sector employment shares ratio								
$(e^g_f/(e^p_f + e^g_f))/(e^g_m/(e^p_m + e^g_m))$	1.533	1.550	1.296	1.296	1.976	1.977	1.097	1.098
Ratio probability job finding public/private								
$p_g/m(\theta_p)$	1.248	1.238	1.009	1.008	0.742	0.743	0.735	0.736
Regional variation in public sector size								
and women's over-representation	0.004	0.004	0.004	0.004	0.002	0.002	0.002	0.002
Non-targeted moments								
Unemployment rates								
Male $(u_m/(1-i_m))$	0.037	0.031	0.091	0.081	0.035	0.027	0.076	0.071
Female $(u_f/(1-i_f))$	0.037	0.041	0.085	0.096	0.033	0.043	0.082	0.087
Inactivity rates								
Male (i_m))	0.126	0.160	0.263	0.273	0.123	0.130	0.266	0.270
Female (i_f))	0.261	0.272	0.389	0.381	0.342	0.332	0.418	0.408
Aggregate gender wage gap	-0.230	-0.244	-0.286	-0.286	-0.207	-0.205	-0.217	-0.218
Flow rates								
$P \to I$, men	0.009	0.009	0.024	0.023	0.011	0.009	0.023	0.019
$P \to I$, women	0.017	0.016	0.033	0.032	0.024	0.024	0.026	0.029
$G \to I$, men	0.011	0.010	0.019	0.021	0.012	0.008	0.016	0.016
$G \to I$, women	0.016	0.015	0.026	0.030	0.015	0.022	0.017	0.026

Table C.13: France and Spain for college and non-college educated: model vs. data

	France				Spain			
Targets	**College**		**Non-college**		**College**		**Non-college**	
	Data	Model	Data	Model	Data	Model	Data	Model
Unemployment rates $(u_m + u_f/(1-i_m)+(1-i_f))$	0.061	0.061	0.117	0.117	0.114	0.112	0.227	0.227
Non-employment rates								
Male $(i_m + u_m)$	0.189	0.189	0.391	0.390	0.214	0.226	0.394	0.395
Female $(i_f + u_f)$	0.298	0.298	0.549	0.550	0.316	0.318	0.623	0.623
Private-sector wage gap $w_f^p/w_m^p - 1$	-0.199	-0.199	-0.162	-0.162	-0.245	-0.261	-0.246	-0.246
Nr. of weekly wages- exp. cost vacancy $\kappa\Theta^{1-\eta}/(W_{mp}/12)$	10.000	10.004	5.000	5.009	10.000	9.999	5.000	4.999
Flow rate								
$E \to I$, aggregate	0.015	0.015	0.024	0.024	0.019	0.019	0.036	0.036
Public-sector employment shares ratio $(e_f^g/(e_f^p+e_f^g))/(e_m^g/(e_m^p+e_m^g))$	1.692	1.693	1.675	1.676	1.524	1.544	1.306	1.306
Ratio probability job finding public/private $p_g/m(\theta_p)$	0.955	0.955	0.807	0.806	1.064	1.064	0.775	0.775
Regional variation in public sector size and women's over-representation	0.007	0.007	0.007	0.007	0.011	0.008	0.011	0.011
Non-targeted moments								
Unemployment rates								
Male $(u_m/(1-i_m))$	0.059	0.055	0.106	0.108	0.097	0.105	0.199	0.206
Female $(u_f/(1-i_f))$	0.065	0.067	0.132	0.129	0.131	0.119	0.274	0.260
Inactivity rates								
Male $(i_m))$	0.139	0.141	0.319	0.316	0.130	0.136	0.244	0.238
Female $(i_f))$	0.250	0.248	0.480	0.484	0.213	0.226	0.481	0.490
Aggregate gender wage gap	-0.194	-0.219	-0.161	-0.164	-0.222	-0.241	-0.239	-0.240
Flow rates								
$P \to I$, men	0.013	0.010	0.021	0.020	0.013	0.014	0.025	0.026
$P \to I$, women	0.021	0.018	0.031	0.030	0.026	0.024	0.055	0.054
$G \to I$, men	0.011	0.013	0.017	0.019	0.013	0.015	0.021	0.021
$G \to I$, women	0.013	0.020	0.022	0.030	0.019	0.022	0.046	0.046

Calculation of compensating differentials for public-sector workers

Alternatively, we can measure compensating differential as the additional wage needed for a public sector worker to accept the same job characteristics as workers in the private sector. Hence the hours premium changes to:

$$PremiumH_j^g = \frac{(\xi_p - \xi_g)\int_0^{\bar{x}_{g,j}^{na}} xf(x)dx}{F(x_{g,j}^{\bar{n}a})}\frac{1}{w_{g,j}} \times 100, j = [m, f]. \tag{C.4}$$

Regarding job security one obtains:

$$\begin{aligned} PremiumS_j^g = \xi_{g,j} & \left(\frac{\delta_p}{r + \tau + \lambda + \delta_g + pg} - \frac{\delta_g}{r + \tau + \lambda + \delta_g + pg}\right) \\ & \times \left[F(\hat{x}_{g,j})\hat{x}_{g,j} - \int_0^{\hat{x}_{g,j}} xf_{g,j}(x)dx\right]\frac{1}{F(\hat{x}_{g,j})w_{g,j}} \times 100, \end{aligned} \tag{C.5}$$

$j = [m, f]$.

Table C.14: Value of public sector job characteristics

	Perspective of a public sector worker			
	Hours premium $[\xi_p = \xi_g]$		**Job security** $[\delta_{p,j} = \delta_{g,j}]$	
	Women	Men	Women	Men
US	0.870	0.817	0.941	0.985
UK	1.514	1.037	1.049	1.326
France	3.664	2.962	1.767	2.057
Spain	2.373	2.023	3.159	3.639

Notes: Model simulations; percentages of public sector wages that men and women are willing to give up for continuing to work fewer hours (compared to the private sector) and for keeping their greater job security (compared to the private sector).