European Sociological Review, 2023, **39**, 280–300
https://doi.org/10.1093/esr/jcac046
Advance access publication 10 October 2022
Original Article

OXFORD

National work–family policies and the occupational segregation of women and mothers in European countries, 1999–2016

Jennifer L. Hook*, Meiying Li, Eunjeong Paek and Brigid Cotter

Department of Sociology, University of Southern California, 851 Downey Way, Hazel Stanley Hall 314, Los Angeles, CA, USA
*Corresponding author: E-mail: hook@usc.edu.

Abstract

Some scholars hypothesize that although work–family policies help incorporate women into the labour market, they do so by integrating women, and mothers specifically, into female-dominated occupations. Some suggest that although these policies are 'good' for lower educated women, they harm higher educated women by concentrating them in female-dominated professions. We revisit this debate using the highest quality data brought to bear on this question to date. We use the EU Labour Force Survey 1999–2016 (n = 21 countries, 235 country-years, 2.5 million men and women aged 20–44), combined with an original collection of country-year indicators. Specifically, we examine how the two most widely studied work–family policies—paid parental leave and early childhood education and care (ECEC)—and public sector size affect occupational segregation for men and women by educational attainment and parental status. We find no evidence that 'generous' welfare states promote segregation. Rather, a specific policy—parental leave in excess of 9 months—promotes segregation between men and women broadly, but most acutely for non-tertiary-educated mothers. Findings are generally null for paid leave of up to 9 months. ECEC is associated with greater integration, particularly for tertiary-educated women. Large public sectors are associated with segregation, with both tertiary-educated men and women more likely to work in feminized occupations. Public sector size, however, is not as tightly bundled with work–family policies as previous work suggests.

Introduction

Scholars generally agree that work–family policies—primarily early childhood education and care (ECEC) and moderate amounts of well-paid parental leave—help incorporate mothers into the labour force. They debate, however, whether these policies have deleterious consequences for women's or mothers' occupational integration, particularly for highly educated women (Mandel and Semyonov, 2005, 2006; Mandel and Shalev, 2009; Pettit and Hook, 2009; Mandel, 2011, 2012; Keck and Saraceno, 2013; Korpi et al., 2013; Hook and Pettit, 2016; Brady et al., 2020). One argument, the 'welfare state paradox', claims that well-developed welfare states increase women's employment at the expense of perpetuating occupational sex segregation and gender wage gaps (Mandel and Semyonov, 2006). Conflicting empirical findings, however, have left theoretical debates about the causes and remedies of occupational segregation unresolved. This is particularly vexing given its policy relevance.

Although vertical segregation—the extent to which women are under-represented in top jobs—is an important outcome and one we examine, we focus on horizontal occupational segregation[1]—the extent to which men and women are located in different occupations—for three reasons. First, it plays a central role in other labour market outcomes including gender inequality in wages (Blau and Kahn, 2017), promotion opportunities, and representation in 'dead-end jobs' (Bihagen and Ohls, 2007). It is nearly universally associated with inequality (Anker, 1998). Second, theoretically, horizontal segregation is central to the debate about whether generous work–family policies exacerbate occupational sex segregation (Mandel and Semyonov, 2006). Third, a focus only on vertical segregation obscures sex segregation, and its consequences, for the majority of people.

Received: July 2021; **revised:** July 2022; **accepted:** September 2022
© The Author(s) 2022. Published by Oxford University Press. All rights reserved. For permissions, please e-mail: journals.permissions@oup.com

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Yet, horizontal segregation is less examined in cross-national work–family policy literature, in part, because it is difficult to obtain detailed and harmonized occupational categories across countries. The Luxembourg Income Study Database (LIS) has been a reliable source for many employment outcomes, but contains only 10 harmonized occupational categories. Yet, research with more detailed occupational categories uniformly finds higher segregation (Reskin and Hartmann, 1986) and scholars in this debate have called for more detailed occupational categories (Mandel and Shalev, 2009; Steinmetz, 2011; Brady, Blome and Kmec, 2020).

Three additional limitations of the literature obscure our understanding of the link between work–family policies and occupational sex segregation. First, the most highly cited studies use single-year cross-sectional data from about 20 countries in the mid-1990s. This is potentially problematic because a small *N* at the macro-level may bias estimates of policy effects (Fairbrother, 2014; Schmidt-Catran et al., 2019). Second, most studies have examined outcomes for women overall, sometimes by parental status or education, but rarely the intersection of the two, despite theoretical reasons to suspect that they operate in interaction. Third, most studies do not explicitly address how selection into the labour market informs our understanding of policy effects on occupational segregation.

Drawing on the EU-LFS, we address these issues. We examine the association of ECEC, paid parental leave, and public sector size (the three components of Mandel and Semyonov's [2005, 2006] welfare state index) with sex segregation using three-digit occupation codes, yielding over 100 unique occupations, and vertical segregation by modelling attainment of management positions. We use repeated cross-sectional labour market data in a comparative longitudinal survey design (235 country-years from 21 countries) to partially address concerns of bias in small *N* designs and pay explicit attention to the role of selection and the interactive effects of gender, parenthood, and educational attainment. We find no support for a general welfare state paradox. Rather, a specific policy that encourages mothers to take longer separations from employment—parental leave in excess of 9 months—promotes segregation. And not just between parents, but between men and women more broadly. There is no evidence that moderate length parental leave or ECEC damages the occupational outcomes of tertiary-educated women. In fact, ECEC is associated with greater integration. Public sector size is associated with segregation; both tertiary-educated men and women are more likely to work in feminized occupations. Public sector size, however, is not as tightly bundled with work–family policies as previous work suggests. Contrary to the welfare state paradox, we conclude that countries can robustly fund ECEC and provide moderate amounts of paid parental leave without deleterious consequences for women.

Theory and empirics

How work–family policies affect women's occupational integration

Theoretically, the claim that work–family policies that incorporate mothers into the labour market could have deleterious consequences on gender inequality in other labour market outcomes, including occupational segregation, has solid footing. First, work–family policies may incorporate lower educated mothers who would otherwise not be in labour market into employment, clustering them in lower skilled female-dominated occupations. That is, as the pool of women in the labour market becomes less positively selected (e.g. on educational attainment) we should see more occupational segregation (Pettit and Hook, 2009). There is mixed evidence on whether work–family policies disproportionately incentivize lower educated women (Hook and Li, 2020).

Second, work–family policies may reduce women's human capital because they encourage breaks in employment (Mandel and Semyonov, 2005, 2006). Although most mothers who return to work after child-related leave return to the same or similar jobs, the risk of a downward occupational move increases with the length of leave and this risk varies by country and women's skill level (Aisenbrey et al., 2009). Furthermore, leaves may impact mothers' work commitment. Extensions of leave in Germany 'significantly and swiftly' changed norms about maternal employment and reduced mothers' work commitment (Gangl and Ziefle, 2015, p. 554).

Third, work–family policies may promote discrimination against mothers and women of child-rearing age more broadly. Experimental evidence suggests that mothers are discriminated against at the point of hire (Correll et al., 2007). Work–family policies may increase this discrimination by raising the costs of hiring mothers and reinforcing the stereotype that mothers are more committed to family responsibilities than fathers. Glass and Fodor (2011) find that state policies that make motherhood salient for employers—long parental leaves in Hungary—contribute to discrimination against mothers and to women of childbearing age more broadly.

The first mechanism, selection, is applicable to both ECEC and paid parental leave, but Pettit and Hook (2009) argue that the latter two mechanisms—lost human capital and discrimination—are only applicable to paid parental leave. Whereas Mandel and Semyonov

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

(2005, 2006) argue for a singular welfare state index—including paid leave, childcare, and the size of the public welfare sector—that reflects 'broad phenomenon', Pettit and Hook (2009) argue that each policy should be considered along two dimensions—whether it promotes labour market inclusion or exclusion and whether it discourages or promotes equality among those in the labour force. They agree with Mandel and Semyonov (2005, 2006) that lengthy paid parental leaves promote inequality, but they argue that ECEC promotes inclusion and equality because ECEC should help mothers maintain continuous employment, sustaining their human capital and lowering discrimination. Yet, Steinmetz (2011), while concurring that ECEC should lower vertical segregation (into management), argues that ECEC supports horizontal segregation by increasing the supply of typically women's jobs.

Furthermore, Shalev (2008) argues that generous family policy limits higher skilled women's achievements. Specifically, they face higher penalties for career interruptions (e.g. parental leave) resulting from decreased accumulation of human capital and employer discrimination. Mandel (2012) concurs that 'the absence of universal employment protection may in a sense benefit highly skilled women, because it lowers employers' motivation to practise statistical discrimination'. Yet, other researchers argue that because highly educated women are less likely to take long parental leaves, the policies have little influence on them (Korpi, Ferrarini and Englund, 2013).

Implicit in this discussion is who should experience the impacts of work–family policies; effects are generally anticipated to vary by women's parental status and educational attainment. Specifically, discussions of selection anticipate that work–family policies incentivize lower educated mothers to enter or stay in the labour force. Discussions of human capital generally anticipate that mothers, specifically mothers who use parental leave, will be impacted by work–family policy and that penalties are stronger for the highly educated. Discussions of discrimination generally focus on mothers and potential variation by educational attainment, but also acknowledge that discrimination may be applied to all women of childbearing age.

The role of public sector employment

Researchers also posit that generous work–family policies expand public sector employment in traditionally women's fields, fostering greater employment *and* horizontal segregation (Mandel and Semyonov, 2006; Shalev, 2008; Mandel and Shalev, 2009). Shalev (2008) refers to the public sector as a 'double pay-off' for lower educated women because the state offers both employment and a woman-friendly working environment. The public sector generally provides better work–life reconciliation and gender equality policies and better pay and pensions for women than in the private sector (Rubery, 2013, 2015). The 'woman-friendliness' of public employment, however, varies across countries and time periods. In general, this characterization does not apply as well to Central/Eastern European countries and when the public sector is subject to austerity measures (Rubery, 2013, 2015).

Although 'good' for lower educated women, large public sectors may be detrimental to the occupational integration of higher skilled women, concentrating them in feminized occupations and limiting attainment of top jobs (Mandel and Semyonov, 2006; Shalev, 2008; Mandel and Shalev, 2009). The argument couples discrimination in the private sector generated by 'social rights attached to women's employment' (Mandel and Semyonov, 2006, p. 1916) (i.e. paid leave) with favourable conditions in the public sector that funnel highly educated women into the public sector (Mandel and Semyonov, 2006). This argument, however, has bundled work–family policies (paid leave and ECEC) and public employment into a broad phenomenon both theoretically and empirically (i.e. welfare state index), even though how these factors cohere varies considerably.

Overall, we can anticipate that public sector size will impact horizontal occupational segregation, at least through selection. Women are more likely to be employed where public sector employment is robust (Fodor and Glass, 2018). Moreover, because countries with larger public sectors have more feminized and managerial occupations, tertiary-educated men and women are more likely to work in these occupations (Steinmetz, 2011). Beyond selection, however, the mechanisms are ambiguous. Without the bundling of large public sectors with 'generous' work–family policies (i.e. long paid leave), there is no clear reason why the existence of large public sectors should exacerbate inequality in the private sector.

Empirics

Although theoretical claims are clear, the empirical evidence about the link between work–family policies and occupational segregation is mixed. Largely due to data constraints, segregation is less studied in the cross-national, comparative literature on work–family policy than in other labour market outcomes. The evidence base comes from cross-sectional analyses of approximately 20 countries using 10–28 occupational categories. Furthermore, most studies have examined outcomes for women overall, sometimes by parental status or education, but rarely the intersection of the two. We organize this review into findings about women, mothers, and, finally, highly

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

educated women. Appendix A of the Supplementary Data details data sources, samples, and variables used in each study.

Mandel and Semyonov (2006) found that generosity on a welfare state index (paid leave, childcare, and public sector size) was associated with women's greater representation in female-typed occupations and lower representation in managerial and lucrative managerial positions. Pettit and Hook (2009), however, found that childcare supports both women's employment and their representation in professional (but not managerial) occupations. Conversely, Brady, Blome, and Kmec (2020), examining occupation per cent female and managerial positions, conclude that 'work-family policies are not associated with labor market outcomes for women or mothers' (p. 27).

Research focused on the occupational segregation of mothers, however, suggests that mothers are segregated from non-mothers into lower prestige occupations and this relationship varies within countries by educational attainment and across countries by policy design (Abendroth et al., 2014; Hook and Pettit, 2016). Hook and Pettit (2016) found occupational segregation clusters by motherhood in countries that enable mothers to devote substantial time to caregiving (e.g. long parental leaves), but clusters by educational attainment in countries with meagre family policy. They find less differentiation in countries that enable mothers to compete in the labour force (e.g. ECEC). Similarly, Abendroth, Huffman and Treas (2014) found smaller occupational prestige penalties for mothers in countries with higher spending on ECEC. Pettit and Hook (2009) found that ECEC increases women's representation in professional occupations more so for mothers than non-mothers.

Studies have not reached a consensus about how work–family policies affect the occupational integration of highly educated women. In a series of studies, Mandel and colleagues find that 'generous family policy' promotes gender equality for less advantaged women, whereas ungenerous policy promotes equality for more advantaged women (Mandel and Shalev, 2009; Mandel, 2011, 2012). Mandel (2011) finds that an index of advantaged women's vertical integration is negatively associated with a seven-item family policy index. In contrast, Korpi, Ferrarini and Englund (2013) found similar managerial gender gaps across countries with different family policy configurations. Steinmetz (2011) finds that childcare for young children (0–3) is associated with tertiary-educated women's *greater* likelihood of being in a managerial position and effective parental leave (weeks multiplied by payment percentage) is associated with their *greater* likelihood of being in male-dominated and managerial occupations. Pettit and Hook (2009) also found that publicly supported childcare supports higher educated women's attainment of professional occupations.

Overall, the debate on whether work–family policies have deleterious consequences for women's occupational integration is contentious. We revisit this question in the most robust framework to date. We use detailed occupational data from repeated cross-sectional labour market surveys in a comparative longitudinal survey design, paying careful attention to selection into the labour market and the intersection of parental status and educational attainment.

Research strategy

Micro-level data and sample

We used quarterly micro-data from the EU Labour Force Survey (EU-LFS) (Eurostat, 2018). The EU-LFS provides comparable, harmonized data, collected by each country's national statistical agency and subsequently provided to Eurostat. The EU-LFS uses a rotating sample of households. The EU-LFS aids comparability across countries and over time by using International Labour Organization guidelines and common classification systems.

The observation period is from 1999 to 2016. To accommodate a series break in OECD ECEC data, we began our macro-data series in 1998, corresponding to 1999 using a 1-year lag. We omit Poland and Slovenia because occupation is recorded with two-digit codes. We used the spring quarter sample because some countries had not transitioned to a quarterly continuous survey, but pulled in adjacent quarters for eight countries with small sample sizes, retaining respondents' first observation.

We restricted the analyses to adults aged 20–44, who were household heads or partners. Until the EU-LFS began collecting detailed household information (as late as 2009), the only way to derive partnership and parenthood variables is through each household member's relationship with the head. The vast majority of adults who are not heads or partners are adult children living in the parental home. We retained respondents through age 44 (e.g. Boeckmann et al., 2015) because we identified parents by the presence of their own household children. The presence of children underestimates women's number of children from age 40 onwards (Greulich and Dasré, 2018), which may narrow the differences we observe between parents and childless respondents. We omit those who are economically inactive due to retirement, disability, or education/training.

We omitted surveys with problematic patterns of missing data, often arising from large missingness on one variable. Retained surveys have missing data for fewer than 5 per cent of cases (mean = 0.5, median =

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

0). We deleted respondents with missing data listwise. The final sample size is N = 2,522,299 men and women in 235 country-years from 21 countries, including 2,060,474 currently employed respondents. Table A1 lists the countries, years of observation, and minimum sample size per country. Measures created from the data and descriptive statistics are weighted with EU-LFS provided weights, which account for sampling design and non-response.

Micro-level measures

Using three-digit occupational codes, we model the effect of family policy on the *percentage of women in a respondent's occupation*. Women's representation is calculated from each survey on a sample of respondents ages 20–59. Moving from one-digit to three-digit codes adds substantial variation. For example, in the Netherlands, in 2013 (median case), the one-digit category professional is 58.5 per cent women. Within this one-digit code, however, the share of women ranges from 6.8 per cent (software and applications developers) to 85.0 per cent (Nursing and midwifery professionals). The range spans at least 45 percentage points in all one-digit codes except for male-dominated skilled agricultural and forestry.

We use the percentage of women in a respondent's occupation to create three additional binary measures—whether a respondent is in a *female-dominated*, *male-dominated*, or *integrated occupation*. Following Steinmetz (2011), occupations are classified as female-dominated if 75–100 per cent of occupational incumbents are women, as male-dominated if 0–25 per cent are women, and integrated if 26–74 per cent are women. We use one-digit ISCO codes to identify *managerial occupation*.

The EU-LFS provides three-digit International Standard Classification of Occupations (ISCO) codes using ISCO-88 for 1999–2010 and ISCO-08 for 2011–2016. There are 114 codes for ISCO-88 and 134 for ISCO-08. Unfortunately, cross-walking occupations using Stata package 'ISCOGEN' (Jann, 2019) created a difference in the number of occupations just as large as the original difference and resulted in a loss 8,143 observations. Thus, we use original codes. Inspection of trendlines and inclusion of a dummy variable for 2011–2016 raised no concerns, and results are robust to using cross-walked data (in Supplementary Data).

Independent variables at the micro-level are parenthood and educational attainment. We code respondents as parents if they are living with children ages 0–19 (the EU-LFS reports age in 5-year bands). The results are robust to alternate age ranges including 0–4 and 0–9 (in Supplementary Data). Education includes two categories: with and without a tertiary degree (2011 ISCED levels 5–8 and 0–4, respectively). Most non-tertiary-educated respondents have completed upper secondary.

We controlled for age, partnership status, whether the respondent was foreign born, and the number of adults aged 65 and over in the household. Because the prevalence and meaning of cohabitation vary across countries (Hiekel et al., 2014), we coded respondents as partnered if they were married or cohabiting. The effect on household adults aged 65+ is ambiguous given that older adults may provide and/or consume caregiving, and both are linked to women's employment.

Macro-level data and measures

We built on our original collection of macro-level data (Hook and Paek, 2020) for each of the 235 country-years in the micro-data. Our indicators include ECEC spending, paid parental leave, and public sector size, as well as three control variables assessing part-time employment, gender egalitarianism, and women's share of employment. We also tested a control for unemployment, but its inclusion did not improve model fit nor change substantive results. Macro-level measures were lagged 1 year prior to the micro-level survey given that our micro-level data come from the spring quarter.

Spending on early childhood education and care.

Given the difficulty in obtaining valid measures of support for ECEC via publicly supported enrolment across a large set of countries and over time, we used a spending-based measure. Spending may include spending on public facilities, subsidies to private and non-profit providers of centre and home-based care, and transfers to families for purchasing care via tax relief, vouchers, and/or allowances (European Commission, 2014). We used ECEC spending per child under the primary school starting age. From the OECD Family Database (OECD, 2019b), we took spending in the millions in constant prices of national currency, adjusted it for purchasing power parities (PPP) (OECD, 2019d), and divided by the number of children (United Nations Statistics Division, 2019) under school starting age (OECD, 2018). In some—mostly federal—countries, the data do not fully capture local-government spending. Spending, however, encompasses the varied approaches governments take to ECEC, and there is a strong correlation between spending and enrolment (Van Belle, 2016).

Months of paid parental leave

Leave-policy design is complex and often combines periods of well-paid maternity or parental leave with

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

periods of low-paid parental leave or home-care allowance. We were theoretically interested in the maximum length of time that the state will pay a mother to care for her child at home; we refer to this time collectively as parental leave. Parental leave is taken from the OECD Family Database, which calculates leave for the first single birth and assumes that parents worked prior to birth and earned 100 per cent average gross earnings, met eligibility requirements, had a healthy birth, and made choices that maximized leave length (OECD, 2019b). Data for Estonia, Latvia, Lithuania, and Slovenia were constructed from multiple sources detailed in Appendix B of Supplementary Data. We used total weeks of paid parental leave for mothers, which includes weeks of payment from maternity, parental, and extended leave, regardless of employment protection. The measure is divided by 4.3 to convert weeks to months.

Given previous findings of a curvilinear relationship between parental leave and women's employment, we tested several specifications, including a linear measure, a curvilinear measure, and splines with one knot at either 3, 6, 9, 15, 18, or 21 months. Fit statistics preferred a spline at 9 months. Thus, one continuous variable quantified the effect of the number of months 0–9 and a second month greater than 9. This allows 'short' leave to differ from 'long' leave.

Public sector size

We use OECD's public employment in general government series 2007–2019 from the Government at a Glance database. It measures those directly employed 'in all levels of government (central, state, local, and social security funds) and includes core ministries, agencies, departments, and non-profit institutions that are controlled by public authorities' (OECD, 2019c). We supplement this series with data for 1995 and 2005 from Government at a Glance 2009 (OECD, 2009). We linearly interpolate years between 1995 and 2005 and between 2005 and the first year observed in the 2007–2019 database. This OECD series was the most complete for the country-years in this project and correlated well with other overlapping data sources.

Women's part-time employment rate

Part-time work's development, availability, use, characteristics, and quality vary greatly across European countries (Fagan et al., 2000) as do its consequences for women's employment (Barbieri et al., 2019). It is linked to the gender composition of occupations through multiple pathways (Bächmann et al., 2022). As the part-time employment rate is moderately correlated with work–family policies (Appendix of Supplementary Data), we include it as a control variable. We use data from the OECD Employment and Labour Market Statistics on the incidence of part-time employment among dependent employees as a percentage of women's total employment, using national definitions (OECD, 2019a).

Gender egalitarianism

Egalitarian gender norms are associated with both women's employment and work–family policies (Boeckmann, Misra and Budig, 2015; although see Uunk and Lersch, 2019). Gender norms may be linked to levels of occupational sex segregation through labour supply effects and/or norms shaping women's occupational location. Thus, we include gender norms as a possible confounder. Consistent with previous research, and due to data constraints, we used a measure gauging norms about maternal employment (Boeckmann, Misra and Budig, 2015).

We used data from the European Values Surveys (EVS; 1990–1992, 1999–2000, 2008–2009, 2017–2018) and the World Values Surveys (WVS; 1990, 2010–2012) (EVS, 2015, 2019; WVS, 2015). We used an item asking respondents how much they agreed with the statement 'A pre-school child is likely to suffer if his or her mother works' on a 4-point Likert scale: agree strongly, agree, disagree, and disagree strongly. We created a continuous measure of the country-year mean where higher scores indicated stronger support for egalitarianism. We used EVS/WVS sample weights and linear interpolation to calculate values between surveys.

Women's share of employment

We include a control variable measuring women's share of employment to account for cross-national differences and within-in country changes in the sex composition of the employed. Share was calculated from the microlevel sample.

Method

We used three-level linear regressions, with individuals nested within country-years (n = 235) within countries (n = 21), a common approach for analyzing repeated cross-sectional surveys (e.g. Hook and Paek, 2020; Taiji and Mills, 2020). We use the same linear model for binary outcomes (i.e. linear probability model, LPM). Although there are several advantages to LPM (Breen et al., 2018), we chose LPM because our models require three random slopes, which cause estimation issues in multilevel logit models. Substantive conclusions are robust to alternate modelling strategies as shown in Table OS-7 in Supplementary Data. We tested several specifications of time including a cross-classified model (country-years nested within countries and years). A linear specification was preferred. Level-2 results were consistent across approaches.

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Our aim is to test the association of family policy with occupational segregation for unique groups identified by the intersection of gender (women/men), educational attainment (tertiary/non-tertiary), and parenthood (yes/no). We accomplish this via four-way interactions, producing policy associations for eight groups. Given the importance of gender in conditioning all associations between variables and the outcome, we allow gender to interact with all variables in the model. The equation is presented in Appendix C of Supplementary Data.

At the country level, we included a random intercept. At the country-year level, we included a random intercept and random coefficients for gender, education, and the presence of children, which allow the slopes to vary for all level-1 variables included in interaction terms (Heisig and Schaeffer, 2019). Micro-level controls are constrained to fixed effects. Individual (level 1) variables were centred within country, which is preferred for testing cross-level interactions. Country-year (level 2) variables were grand-mean centred, which is preferred when testing level-2 interactions (Enders and Tofighi, 2007).

At the macro-level, there are two primary concerns about insufficient *N*. First, results are sensitive to outlying countries. We test the robustness of our results to outliers with the omission of (i) each individual country and (ii) seven country clusters. Second, multilevel models in small samples can produce standard errors that are anticonservative (Bryan and Jenkins, 2016; Schmidt-Catran, Fairbrother and Andreß, 2019). A widely cited paper by Bryan and Jenkins (2016) concludes that at least 25 countries are needed in linear models. To address this concern, we use repeated cross-sectional labour market data in a comparative longitudinal survey (Fairbrother, 2014; Schmidt-Catran, Fairbrother and Andreß, 2019), with REML estimation. Our macro-level variables vary by country-year (n = 235), which helps mitigate small sample problems. Parental leave, however, varies considerably less than other country-year measures, yielding 31 unique values. This is a core limitation of this and similar analyses.

Correlations are shown in the Table OS-1 in Supplementary Data. Contrary to the linking of large public sectors and work–family policy, public sector size is weakly correlated with leave over 9 months (.23) and moderately correlated with ECEC (.43). The highest correlation is between ECEC spending and gender egalitarianism (.64). We test for problems associated with multicollinearity in two ways. First, we run non-nested models to produce variance inflations factors (VIF). The mean VIF is 2.15 and no VIF exceeds 5. Second, in Supplementary Data, we report models omitting (i) controls for women's part-time employment and gender egalitarianism, (ii) public sector size, and (iii) ECEC spending. We discuss these models in the results section.

Selection

Although most cross-national studies of segregation do not explicitly deal with the selection and the consequences thereof (Mandel and Shalev, 2009; Korpi, Ferrarini and Englund, 2013; Brady, Blome and Kmec, 2020), selection is central to this topic (Pettit and Hook 2009). Common solutions used for wages are Heckman correction or imputation, both of which require strong assumptions (Blau and Kahn, 2017). A third option is extending the domain of segregation beyond the labour market by explicitly including those not in the labour force (Cohen, 2004; Hook and Pettit, 2016; Guinea-Martin et al., 2018).

This approach allows us to directly observe the consequences of selection without relying on assumptions that undergird other approaches. After examining our outcomes among the employed, we expand our sample to include the unemployed who report a prior occupation and those NILF. We consider the heterogeneity of economic inactivity (Guinea-Martin et al., 2018) and code NILF to the share of women in the inactivity category: (i) looking after children or incapacitated, (ii) other personal and family responsibilities, (iii) belief that no work is available, and (iv) other reasons.

Findings

Descriptive

Table 1 shows that across 235 country-years, women are employed in occupations that are approximately 66 per cent women, on average. The range spans from 56–77 per cent, indicating that women working in the least gender segregated country-year are in occupations that are, on average, 21 percentage points less segregated than women working in the most gender segregated country-year. This is further reflected in the wide variation across country-years in working in female-dominated, male-dominated, integrated, and management occupations. Employed men work in occupations that are approximately 28 per cent women, on average. There is less variability across country-years for men, although the range is similar, the standard deviation is smaller. Depending on the dependent variable, between 22 and 34 per cent of variation is within countries and 66–78 per cent is between countries (Table OS-6 in Supplementary Data).

Figure 1 displays variation in women's occupational segregation by educational attainment, country, and time. The line at the bottom of the graph shows women's share of employment for reference. The figure shows large variations across countries

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

in rates and trends as well as in the educational gap. There are three key takeaways. First, segregation declines in some countries, but increases in others. Second, in some countries, segregation moves in concert with employment share, whereas in others it does not. Third, in most countries, women without tertiary degrees are in more segregated occupations than women with tertiary degrees.

Table 1 also shows substantial variation across country-years in the presence of a child, educational attainment, and the demographic composition of women, men, and their households. The second half of Table 1 displays descriptive statistics for the country-year variables. There is sizable variation in all country-year variables.

Multilevel models

As the results of interest are four-way interactions, we present the results in Table 2 as marginal effects of country-year context for eight groups—gender × education × parenthood. Full tables appear in Table A2. We performed a series of robustness checks reported in the Online Supplement, including sensitivity to the removal of each individual country and of country clusters. We discuss these results with the main models where relevant.

ECEC: Generally integrating, but most integrating for tertiary-educated women

Column 1 of Table 2 shows the results for models predicting the percentage of women in respondents' occupation among the employed. Columns 3–5 show the component parts—working in female-dominated, male-dominated, or integrated occupations. Results for ECEC generally reveal an integrating effect. As spending increases, women are less likely to work with other women and are more likely to work in integrated occupations, with the exception of non-tertiary-educated mothers. Non-tertiary-educated men are more likely to work with women, showing a decline in working in male-dominated occupations. These results are highly robust to the removal of individual countries and country clusters. Tertiary-educated fathers, however, are less likely to work with women, suggesting greater segregation for this group. This result is sensitive to the exclusion of specific countries and clusters—it loses statistical significance with the exclusion of Greece (Table OS-2 in Supplementary Data) and the Southern or Ireland/U.K. clusters (Table OS-3). It is also not robust to include the unemployed and those not in the labour force (NILF), shown in column 2. Column 6 shows that tertiary-educated mothers and non-tertiary-educated women are more likely to attain management positions in contexts with greater ECEC spending.

To aid interpretation, predicted values for column 1 (of Table 2) are shown in column 1 of Figure 2 for all statistically significant marginal effects. A 1-SD increase in ECEC spending is associated with a 2.2 (mothers) to 2.4 (childless) percentage point decline in the share of women in a respondent's occupation for tertiary-educated women. Figure 2 shows that the effect of ECEC differs by educational attainment (the difference between mothers is statistically significant), and among non-tertiary-educated women by parental status (also stat. sig.), indicating the importance of examining the intersection of educational attainment and motherhood.

Paid leave less than 9 months: Findings generally null

Column 1 of Table 2 shows that the results for months of paid leave before 9 months are generally weak. Non-tertiary-educated men are more likely to work with women as leave increases up to 9 months. Women are more likely to work in male-dominated (and less likely to work in integrated occupations) as leave increases up to 9 months. Furthermore, once we expand to the unemployed/NILF sample, column 2, mothers without a tertiary degree are less segregated as leave increases up to 9 months. This likely reflects the positive association between short leave periods and mothers' labour force participation. For mothers, any occupation is less segregated than home-making. We do see evidence, however, that non-tertiary-educated women may be less likely to attain management positions as leave increases. There is little evidence that leave up to 9 months is segregating for tertiary-educated mothers, with the exception of one model that excludes the Southern cluster (Table OS-3 in Supplementary Data), and there are no detectable negative consequences for management.

Paid leave greater than 9 months: Segregating, most so for non-tertiary-educated mothers

In contrast, the pattern is clear that paid leave in excess of 9 months is associated with gender segregation; it is both feminizing for women—regardless of parental status— and de-feminizing for men. This result is robust to the omission of individual countries and country clusters. Women are more likely to be found in female-dominated occupations and men in male-dominated occupations. Column 3 of Figure 2 displays the predicted values. A 1-SD increase in months of paid parental leave in excess of 9 months is associated with an increase in the share of women in respondents' occupation of about 1.0 percentage point for tertiary-educated mothers, and 1.6 for non-tertiary-educated mothers (this difference is statistically significant).

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Table 1. Descriptive statistics at the county-year level (*N* = 235)

	Women				Men			
	Country-year				Country-year			
	Mean	SD	Min	Max	Mean	SD	Min	Max
Individual-level measures								
Among the employed								
% Women in R's occupation	66.2	3.9	56.5	76.9	28.4	2.9	22.5	43.0
Female-dominated occupation (%)	41.5	11.0	17.4	68.8	5.6	2.3	1.4	13.4
Male-dominated occupation (%)	6.8	2.0	1.5	12.3	54.4	5.9	24.5	66.2
Integrated occupation (%)	51.7	10.5	28.1	77.9	39.9	5.8	27.3	64.5
Manager (%)	5.0	2.5	0.3	13.9	9.1	3.9	1.2	21.5
Full sample								
% Women in R's occupation	68.8	3.2	62.8	77.6	30.3	2.9	24.7	44.2
Child (%)	67.6	9.9	40.3	84.9	56.7	12.1	25.5	82.5
Tertiary degree (%)	34.1	12.4	10.4	58.0	28.2	9.5	7.9	53.7
Age	34.3	1.1	31.6	36.4	34.8	1.2	31.5	37.1
Partner (%)	75.4	10.1	35.9	92.3	72.9	12.7	28.0	93.2
Number adults 65+	0.03	0.03	0.00	0.12	0.03	0.03	0.00	0.13
Foreign born (%)	14.4	11.0	0.5	64.7	13.6	10.6	0.4	63.1
Country-year measures								
ECEC Expenditure ($1,000s per child)[a]	3.2	1.9	0.04	8.9				
Months paid leave before 9 months	7.8	2.0	3.3	9.0				
Months paid leave after 9 months	9.0	12.1	0.0	40.8				
Public sector size	17.6	5.1	10.2	30.4				
Part-time prevalence (%)	28.0	18.3	4.2	75.5				
Gender egalitarianism (scale)	2.6	0.3	1.9	3.4				
Women's share of labor force (%)	47.5	2.5	38.3	53.3				
Number of observations								
N—Country				21				
N—country-year				235				
N—Individual (employed)				991,179				1,069,295
N—Individual (full sample)				1,357,153				1,165,146

[a]ECEC expenditure per child under primary school starting age in PPP adjusted US dollars, in thousands. In the full sample, respondents who are unemployed are coded to their prior occupation whereas those who are NILF are coded to the share of women in the corresponding category, economically inactive: (1) looking after children or incapacitated, (2) other personal and family responsibilities, (3) belief that no work is available, and (4) other reasons.

Values for childless women are slightly lower at .8 and 1.3 percentage points, respectively. Although all groups of women are more segregated with longer leave, it is most segregating for non-tertiary-educated mothers.

When we include respondents who are unemployed or NILF (column 2), the results are highly similar for men. For women, however, the feminizing effect of paid leave attenuates for childless women and intensifies for tertiary-educated mothers, who now show the largest negative effect. Consistent with recent research on the employment of mothers with young children (Hook and Paek, 2020), this suggests that long paid leave negatively impacts the employment of tertiary-educated women.

Size of the public sector: more feminized and managerial employment for tertiary-educated men and women

The results for public sector size are complex. It is associated with overall feminization (working with more women) for all tertiary-educated groups and defeminization (working with fewer women) for all

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

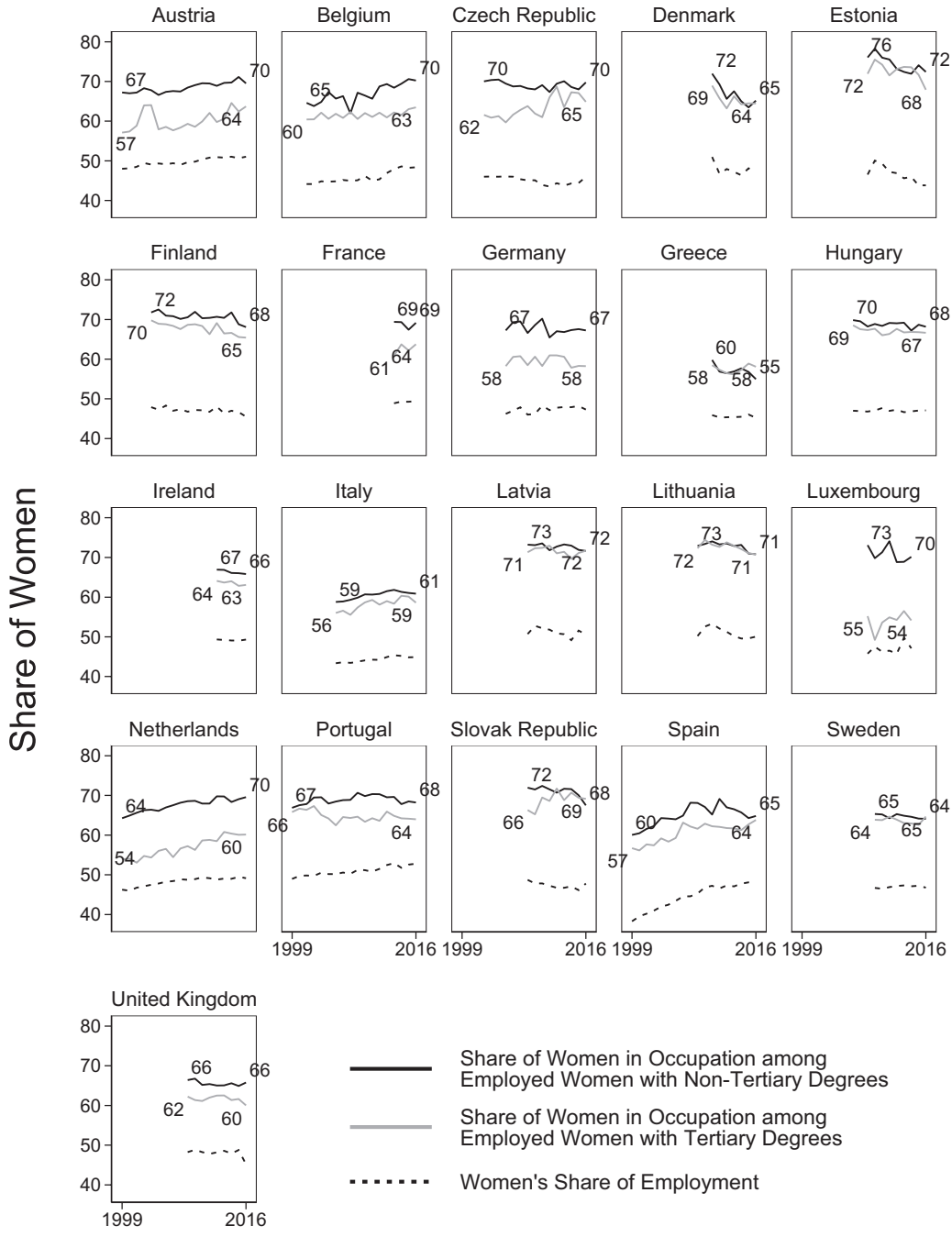

Figure 1. Mean share of women in occupation by educational attainment, year, and country, among employed women age 20–44 (N = 991,179 women; 235 country-years). Note: Unadjusted rates.

non-tertiary-educated groups. Furthermore, the feminizing effect among the tertiary educated is larger for fathers than for mothers. Column 4 of Figure 2 shows that a 1-SD increase in public sector size is associated with an increase in the share of women in respondents' occupation of about 1.4 percentage points for tertiary-educated mothers and 2.4 for tertiary-educated fathers; this difference is statistically significant. Results are robust to the omission of countries and country clusters. Public sector size is also positively associated

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Table 2. Marginal effects of country-year context on occupational segregation by gender, education, and parenthood (21 countries; 235 country-years; 2,060,474 individuals)

Marginal effect of…	…on			(1) Women in occ (%)		(2) Women in occ (%) w/ NILF		(3) Female-dominated		(4) Male-dominated		(5) Integrated		(6) Manager	
ECEC	Women	< Tertiary	Child	.06		.01		.000		−.008	**	.008		.010	***
Expenditure			No Child	−.43	**	−.27	*	−.012	**	−.006	*	.018	***	.010	***
Per Child		Tertiary	Child	−1.16	***	−.81	***	−.023	***	−.007	*	.030	***	.005	**
(in 1,000s)			No Child	−1.19	***	−.86	***	−.025	***	−.007	*	.032	***	.002	
	Men	< Tertiary	Child	.53	***	.45	**	.007		−.012	***	.006		.003	
			No Child	.87	***	.72	***	.011	*	−.017	***	.006		.001	
		Tertiary	Child	−.44	**	−.28		.004		.002		−.006		−.009	***
			No Child	−.18		−.11		.004		−.003		−.001		−.010	***
Months of	Women	< Tertiary	Child	−.18		−.25	*	.003		.006	**	−.009	**	−.003	**
Paid Leave			No Child	.01		−.12		.011	***	.005	*	−.016	***	−.002	*
<= 9 Months		Tertiary	Child	.00		.10		.006		.007	***	−.014	***	−.001	
			No Child	−.16		−.08		.005		.006	**	−.011	**	.002	
	Men	< Tertiary	Child	.67	***	.66	***	.006	*	−.002		−.004		−.003	*
			No Child	.52	***	.46	***	.003		−.002		−.002		.002	
		Tertiary	Child	−.05		−.08		−.003		.001		.002		.000	
			No Child	.04		−.03		−.003		−.003		.005		.005	***
Months of	Women	< Tertiary	Child	.11	***	.08	***	.004	***	−.001		−.003	***	.000	
Paid leave			No Child	.10	***	.04		.003	***	−.001		−.003	***	.000	
> 9 Months		Tertiary	Child	.07	**	.17	***	.003	***	−.001		−.003	***	.000	
			No Child	.05	*	.03		.003	***	.000		−.002	***	.000	
	Men	< Tertiary	Child	−.16	***	−.16	***	−.001		.002	***	−.001	*	.000	
			No Child	−.11	***	−.14	***	.000		.002	***	−.002	*	.000	
		Tertiary	Child	−.16	***	−.15	***	.000		.003	***	−.003	***	.001	***
			No Child	−.15	***	−.17	***	.000		.003	***	−.003	***	.000	
Public	Women	< Tertiary	Child	−.24	***	−.50	***	.007	***	.004	***	−.011	***	.000	
Sector			No Child	−.15	**	−.24	***	.008	***	.003	*	−.011	***	−.001	
(%)		Tertiary	Child	.30	***	.14	**	.016	***	.002	*	−.018	***	.002	*
			No Child	.35	***	.42	***	.016	***	.002		−.018	***	.001	
	Men	< Tertiary	Child	−.15	**	−.04		.003		.002		−.004	*	.002	**
			No Child	−.19	***	−.02		.003	*	.003	*	−.006	**	.001	
		Tertiary	Child	.49	***	.54	***	.007	***	−.007	***	.001		.006	***
			No Child	.35	***	.42	***	.008	***	−.004	**	−.004		.004	***

Notes: *N* for unemployed/NILF sample is 2,522,299. Table A2 presents full results, from which marginal effects are calculated. Marginal effects are based on a linear predictor that includes both the fixed and random effects.
*$P < .05$; **$P < .01$; ***$P < .001$ (two-tailed tests).

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

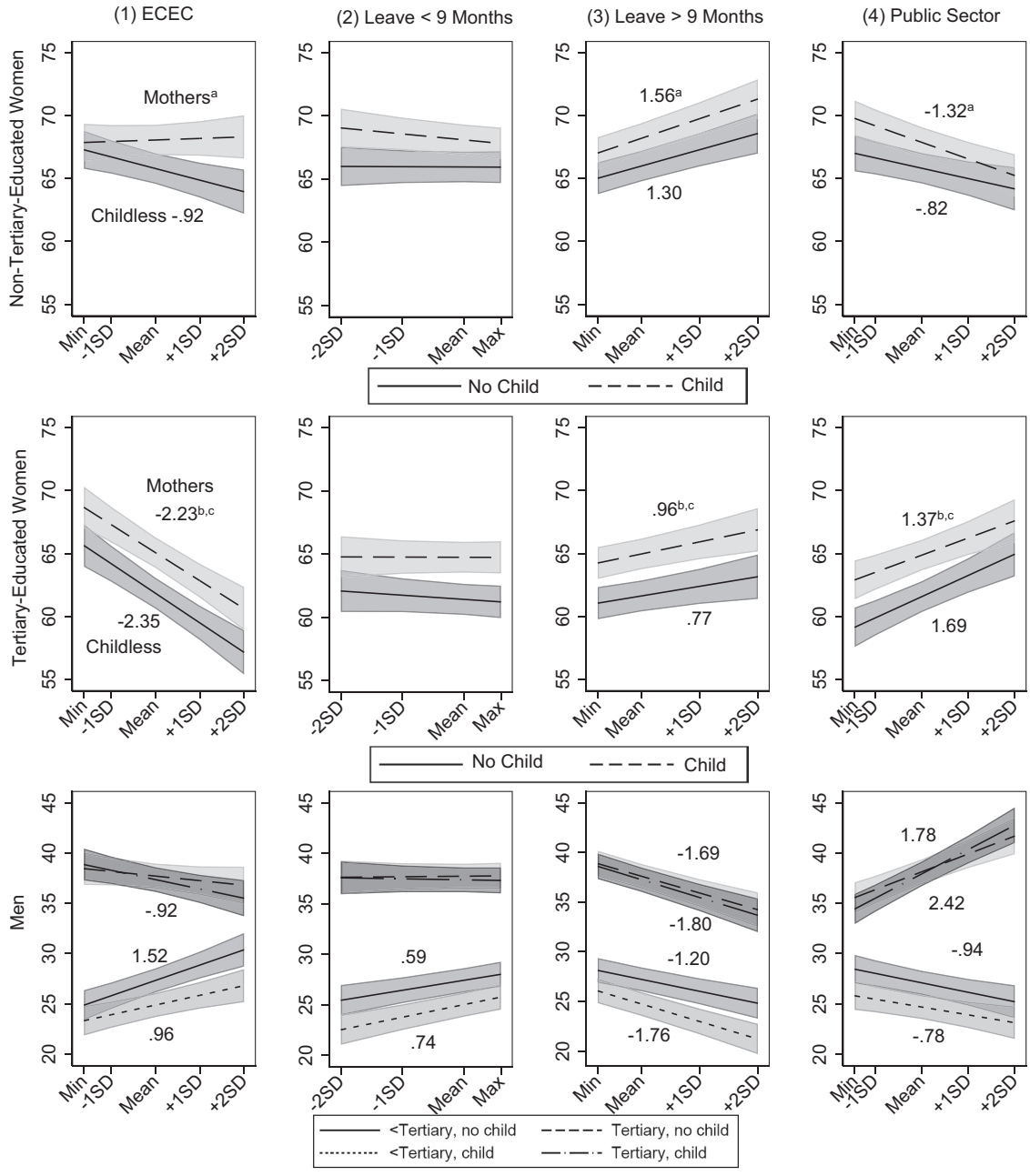

Figure 2. Predicted values—share of women in respondents' occupation—with 97.5 per cent confidence intervals, employed women and men by educational attainment and parenthood status, age 20–44. Notes: Predicted probabilities are based on results in Table A2. Values display the change in the predicted values for a 1-SD increase in X; predicted values are only shown for groups with statistically significant marginal effects (in Table 2). Confidence intervals are set to 97.5 per cent to adjust for pairwise comparisons. Marginal effects differ ($P < .01$) between (a) non-tertiary-educated women by motherhood status; (b) between mothers with and without a tertiary degree; (c) between tertiary-educated mothers and fathers. There are no statistically significant differences among tertiary-educated women by motherhood status.

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

with management for tertiary-educated men, tertiary-educated mothers, and non-tertiary-educated fathers. In general, large public sectors are associated with more feminized and management occupations for the tertiary educated.

Adding unemployed/NILF respondents further clarifies the effect of public sector size. The effect of public sector size becomes more de-feminizing for non-tertiary-educated mothers and becomes less feminizing for tertiary-educated mothers. This suggests that public sector employment draws women into employment regardless of educational attainment, and this employment is less feminized than home-making.

In sum, the welfare state paradox is only clearly supported for paid parental leave in excess of 9 months. Paid parental leave up to nine months is not generally associated with occupational gender segregation. ECEC spending is integrating. Public sector size is associated with occupational feminization for all tertiary-educated groups. Importantly, however, ECEC and public sector size as not as strongly correlated as previous work implies. Furthermore, the integrating effects of ECEC are larger than segregating effects of public sector size for tertiary-educated women. Finally, we never observe a negative effect for any policy measure on the attainment of management positions for tertiary-educated women, and the effect of ECEC is positive.

Robustness checks

In addition to removing each individual country and a series of country clusters, we performed additional robustness checks reported in Tables OS-4, OS-5, and OS-6 of Supplementary Data. Table OS-4 of Supplementary Data shows sensitivity to changes in macro-level variables: (i) the addition of months of paid leave for fathers, and the removal of (ii) macro-level control variables, (iii) ECEC, and (iv) public sector size. The addition of fathers' leave does not affect substantive conclusions and its own relationship to occupational segregation is weak. The exclusion of macro-level controls does not change the substantive conclusions. Nor does the exclusion of ECEC alter conclusions about long parental leave or public sector size. In its absence, however, paid leave less than 9 months is generally integrating for women. The removal of public sector size attenuates the effect sizes of ECEC somewhat, but does not alter substantive conclusions.

Table OS-5 of Supplementary Data shows substantive conclusions are robust to changes in micro-level variables (i) using a cross-walked version of three-digit occupational codes, (ii) defining parenthood as a child aged 0–9 or (iii) aged 0–4, and (iv) restricting the sample to respondents ages 25–44.

Finally, in Table OS-6 of Supplementary Data, we use aggregated data and nest country-years within countries predicting the percentage of workers in female-dominated, male-dominated, integrated, and managerial occupations. Aggregate analyses are consistent with micro-level findings. All files necessary for replication (except the EU-LFS which is obtained directly from Eurostat) are available via the Harvard Dataverse (Hook and Paek, 2022).

Discussion

We re-examined the welfare state paradox debate (Mandel and Semyonov, 2005, 2006), in what we believe is the best design to date, taking advantage of three-digit occupational codes, a comparative longitudinal design, and paying careful attention to selection and the intersection of educational attainment with parental status. Evidence of the welfare state paradox is limited to a policy that encourages mothers to pause employment for over 9 months and up to 3 years. There is little evidence of negative consequences of leave less than 9 months. ECEC is integrating. Contrary to the claims of negative effects emanating from large public sectors, there is no evidence that large public sectors cleave tertiary-educated men to male-dominated employment and women to female-dominated employment. We find that tertiary-educated men and women are more likely to work in feminized and managerial occupations in contexts with larger public sectors. In addition, public sector size is not as tightly bundled with work–family policies as previous work suggests, casting further doubt on a unified paradox.

Paid parental leave in excess of 9 months is segregating for women, regardless of educational attainment or parenthood. The magnitude differs across groups, however, with non-tertiary-educated mothers posting the largest association. Although some have argued that tertiary-educated mothers are spared negative consequences because they are less likely to use long parental leaves (Korpi, Ferrarini and Englund, 2013), we find no support for this. In fact, when we include those not in the labour force, the association is largest for tertiary-educated mothers. This highlights the importance of both selection into employment and heterogeneity by the intersection of women's educational attainment and parental status in modelling policy effects. Our findings are consistent with the theory that long leaves depress mothers' human capital, depress mothers' normative commitments to work, and/or increase employer discrimination against mothers or potential mothers. That we find penalties across all four groups

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

of women suggests that long leave simply exacerbates segregation, regardless of individual women's parental status or actual use of leave. This finding is most consistent with the mechanism of discrimination and supports qualitative work that suggests that all women of reproductive age are potentially coded as mothers by employers when state policies make motherhood salient (Glass and Fodor, 2011). An alternate explanation is that many childless women are future mothers and make decisions based on expectations of motherhood under the condition of long leave availability. Future work is needed to untangle these mechanisms.

The length of paid parental leave is only one measure of social rights to parental leave. It obscures complexities of parental leave design, such as wage replacement rates and eligibility, which impact the realization of these rights (Dobrotić and Blum, 2020; Otto et al., 2021). As advancements are made in conceptualizing and designing new indicators of parental leave there will be additional opportunities to explore these complexities. Although our measure is limited, modelling with splines presents two advantages. First, most countries that offer longer periods of leave combine a shorter well-paid period with a longer less well-paid period. Nine months serves as a rough proxy for shorter well-paid leave and longer less well-paid leave. Second, the common practice of using a curvilinear fit (a squared term) likely over- and under-fits the relationship between leave and labour market outcomes at various points of the curve. By using splines, we can better identify the point at which leave becomes 'too long' for promoting women's occupational integration. Although 9 months was the best fitting spline across country-years, there may be important variations across countries. In Sweden, Evertsson and Duvander (2011) argue that leave that deviates from the average is penalized, not the leave in itself.

There is robust support for the integrating effects of ECEC. Not all women, however, benefit equally. Tertiary-educated women post the biggest declines in occupational feminization, whereas non-tertiary-educated mothers show no association whatsoever. Scholars have posited seemingly contradictory expectations for ECEC. On one hand, it should be integrating because it facilitates women's employment, potentially sustaining their human capital and lowering discrimination (Pettit and Hook, 2009). On the other, it should be segregating because it increases the supply of typically women's jobs (Steinmetz, 2011). We posit that both explanations may be correct, but for different groups of women. The former may explain the integrating effect among tertiary-educated women, whereas the competing effects of the two may explain the null findings for non-tertiary-educated mothers. It is important to note, however, that ECEC spending increases employment among non-tertiary-educated mothers (Hook and Paek, 2020) and our results show that this increase does not cause greater segregation.

Although spending on ECEC correlates with enrolment (Van Belle, 2016), it conflates availability, quality, out-of-pocket costs for families, and hours of operation. As quality data series become available, this presents important avenues for research. For example, operating hours are strongly correlated with women's employment intensity (OECD, 2017), which may affect women's occupational integration, particularly into male-dominated occupations.

In contrast to strong claims about the gender-segregating effects of large public sectors (e.g. Shalev 2008), our results show a pattern defined by educational attainment, and less so by gender or parental status. Consistent with Steinmetz (2011), we find that larger public sectors are associated with the increased likelihood that tertiary-educated men *and* women will work in feminized and managerial occupations, that is, there is more feminized and managerial work for tertiary-educated workers in contexts with large public sectors. The feminizing effect among the tertiary educated, however, is larger for fathers than for mothers. We see the opposite pattern for non-tertiary-educated workers—the percentage of women in respondent's occupation declines with larger public sectors.

The effect of public sector size is an average effect across countries. Yet, the extent to which governments promote work–family reconciliation and gender equality within the public sector varies across countries (Rubery, 2013, 2015). Public sector employment may be more or less segregating in some contexts than others. Future work should explore variability in the quality of public sector employment and how this is linked to women's labour market outcomes.

Other limitations of this work provide opportunities for future research. First, due to data limitations, we only capture occupational outcomes from ages 20 to 44. If disadvantages accumulate or attenuate, we could find more or less evidence of the welfare state paradox later in life. Linking work–family policy to outcomes in later adulthood, however, would require a longer time series for ECEC than currently available. Second, the welfare state paradox should also be examined for wages. Unfortunately, this is not possible with the EU-LFS. Third, the analysis provides average effects across occupations, but the effects of work–family policies may differ across occupations. Future work should explore potential heterogeneity by examining occupational clusters, such as STEM.

In conclusion, we find no evidence that robust family policy damages the occupational outcomes of tertiary-educated women. Countries may provide well-funded ECEC and maternity/parental leaves up to 9 months without deleterious consequences for women's

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

occupational integration. Although large public sectors are associated with segregation (with both tertiary-educated men and women more likely to work in feminized occupations), public sector size is not tightly coupled with these work–family policies. Finally, although some studies fail to find evidence that leave can be too long (Keck and Saraceno, 2013), we find that parental leaves in excess of 9 months are segregating. To promote occupational integration policy-makers may wish to follow the path of several countries, such as Germany, that have reformed their parental leave programs to replace long parental leaves with shorter, well-paid leaves that incentivize women's returns to work.

Endnote

1 Although widely used, some scholars define horizontal segregation as the dimension of segregation that does not comprise pay inequalities and would refer to our measure as total segregation.`

Supplementary data

Supplementary data are available at *ESR* online.

Acknowledgements

Earlier versions of this work were presented at the annual meeting of the American Sociological Association in August 2020, the University of Pennsylvania, UCLA, and the University of Torino.

Funding

This work was partially supported by a Eunice Kennedy Shriver National Institute of Child Health and Human Development research infrastructure grant, P2C HD042828, to the Center for Studies in Demography & Ecology at the University of Washington. The content is solely the responsibility of the authors and does not necessarily represent the official views of the National Institutes of Health. This paper is based on data from Eurostat, EU Labour Force Survey 1999-2016. The responsibility for all conclusions drawn from the data lies entirely with the authors.

References

Abendroth, A. -K., Huffman, M. L. and Treas, J. (2014). The parity penalty in life course perspective: motherhood and occupational status in 13 European countries. *American Sociological Review*, **79**, 993–1014.

Aisenbrey, S., Evertsson, M. and Grunow, D. (2009). Is there a career penalty for mothers' time out? A comparison of Germany, Sweden and the United States. *Social Forces*, **88**, 573–605.

Anker, R. (1998). *Gender and Jobs: Sex Segregation of Occupations in the World*. Geneva: International Labour Organization.

Bächmann, A. -C. *et al*. (2022). Why do some occupations offer more part-time work than others? Reciprocal dynamics in occupational gender segregation and occupational part-time work in West Germany, 1976–2010. *Social Science Research*, **104**, 102685.

Barbieri, P. *et al*. (2019). Part-time employment as a way to increase women's employment: (where) does it work? *International Journal of Comparative Sociology*, **60**, 249–268.

Bihagen, E. and Ohls, M. (2007). Are women over-represented in dead-end jobs? A Swedish study using empirically derived measures of dead-end jobs. *Social Indicators Research*, **84**, 159–177.

Blau, F. D. and Kahn, L. M. (2017). The gender wage gap: extent, trends, and explanations. *Journal of Economic Literature*, **55**, 789–865.

Boeckmann, I., Misra, J. and Budig, M. J. (2015). Cultural and institutional factors shaping mothers' employment and working hours in postindustrial countries. *Social Forces*, **93**, 1301–1333.

Brady, D., Blome, A. and Kmec, J. A. (2020). Work-family reconciliation policies and women's and mothers' labor market outcomes in rich democracies. *Socio-Economic Review*, **18**, 125–161.

Breen, R., Karlson, K. B. and Holm, A. (2018). Interpreting and understanding logits, probits, and other nonlinear probability models. *Annual Review of Sociology*, **44**, 39–54.

Bryan, M. L. and Jenkins, S. P. (2016). Multilevel modelling of country effects: a cautionary tale. *European Sociological Review*, **32**, 3–22.

Cohen, P. N. (2004). The gender division of labor: "Keeping house" and occupational segregation in the United States. *Gender & Society*, **18**, 239–252.

Correll, S., Benard, S. and Paik, I. (2007). Getting a job: is there a motherhood penalty. *American Journal of Sociology*, **112**, 1297–1338.

Dobrotić, I. and Blum, S. (2020). Inclusiveness of parental-leave benefits in twenty-one European countries: measuring social and gender inequalities in leave eligibility. *Social Politics: International Studies in Gender, State & Society*, **27**, 588–614.

Enders, C. K. and Tofighi, D. (2007). Centering predictor variables in cross-sectional multilevel models: a new look at an old issue. *Psychological Methods*, **12**, 121.

European Commission. (2014). *Key Data on Early Childhood Education and Care in Europe*. Luxembourg: Publications Office of the European Union.

European Values Survey. (2015). *European Values Study 1981–2008, Longitudinal Data File*. Cologne: GESIS Data Archive.

European Values Survey. (2019). *European Values Study 2017: Integrated Dataset (evs 2017)*. Cologne: GESIS Data Archive.

Eurostat. (2018). *EU Labour Force Survey Database User Guide*. Luxembourg: European Commission.

Evertsson, M. and Duvander, A. -Z. (2011). Parental leave—possibility or trap? Does family leave length effect Swedish women's labour market opportunities? *European Sociological Review*, **27**, 435–450.

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Fagan, C., O'Reilly, J. and Rubery, J. (2000). Part-time work: challenging the "breadwinner" gender contract. In Jenson, J. *et al*. (Eds.), *The Gendering of Inequalities: Women, Men and Work*. London: Routledge, pp. 174–186.

Fairbrother, M. (2014). Two multilevel modeling techniques for analyzing comparative longitudinal survey datasets. *Political Science Research and Methods*, **2**, 119–140.

Fodor, E. and Glass, C. (2018). Labor market context, economic development, and family policy arrangements: explaining the gender gap in employment in Central and Eastern Europe. *Social Forces*, **96**, 1275–1302.

Gangl, M. and Ziefle, A. (2015). The making of a good woman: extended parental leave entitlements and mothers' work commitment in Germany. *American Journal of Sociology*, **121**, 511–563.

Glass, C. and Fodor, E. (2011). Public maternalism goes to market: recruitment, hiring, and promotion in postsocialist Hungary. *Gender & Society*, **25**, 5–26.

Greulich, A. and Dasré, A. (2018). Observing the number of children with EU-SILC: a quantification of biases. *Population*, **73**, 685–720.

Guinea-Martin, D., Mora, R. and Ruiz-Castillo, J. (2018). The evolution of gender segregation over the life course. *American Sociological Review*, **83**, 983–1019.

Heisig, J. P. and Schaeffer, M. (2019). Why you should always include a random slope for the lower-level variable involved in a cross-level interaction. *European Sociological Review*, **35**, 258–279.

Hiekel, N., Liefbroer, A. C. and Poortman, A. -R. (2014). Understanding diversity in the meaning of cohabitation across Europe. *European Journal of Population*, **30**, 391–410.

Hook, J. L. and Li, M. (2020). Gendered tradeoffs. In Nieuwenhuis, R. and Van Lancker, W. (Eds.), *Handbook on Family Policy*. Basingstoke: Palgrave, pp. 249–266.

Hook, J. L. and Paek, E. (2020). National family policies and mothers' employment: how earnings inequality shapes policy effects across and within countries. *American Sociological Review*, **85**, 381–416.

Hook, J. and Paek, E. (2022). Replication data for: national work-family policies and the occupational segregation of women and mothers in European countries, 1999–2016. https://doi.org/10.7910/DVN/XQ1ALG, Harvard Dataverse, V1, UNF:6:Wi4krtGLoYbYyCHbwrwPcg== [fileUNF].

Hook, J. L. and Pettit, B. (2016). Reproducing occupational inequality: motherhood and occupational segregation. *Social Politics: International Studies in Gender, State & Society*, **23**, 329–362.

Jann, B. (2019). *Iscogen: Stata Module to Translate ISCO Codes*. Boston: Boston College Department of Economics.

Keck, W. and Saraceno, C. (2013). The impact of different social-policy frameworks on social inequalities among women in the European Union: the labour-market participation of mothers. *Social Politics*, **20**, 297–328.

Korpi, W., Ferrarini, T. and Englund, S. (2013). Women's opportunities under different family policy constellations: gender, class, and inequality tradeoffs in western countries re-examined. *Social Politics*, **20**, 1–40.

Mandel, H. (2011). Rethinking the paradox: tradeoffs in work-family policy and patterns of gender inequality. *Community, Work & Family*, **14**, 159–176.

Mandel, H. (2012). Winners and losers: the consequences of welfare state policies for gender wage inequality. *European Sociological Review*, **28**, 241–262.

Mandel, H. and Semyonov, M. (2005). Family policies, wage structures, and gender gaps: sources of earnings inequality in 20 countries. *American Sociological Review*, **70**, 949–967.

Mandel, H. and Semyonov, M. (2006). A welfare state paradox: state interventions and women's employment opportunities in 22 countries. *American Journal of Sociology*, **111**, 1910–1949.

Mandel, H. and Shalev, M. (2009). How welfare states shape the gender pay gap: a theoretical and comparative analysis. *Social Forces*, **87**, 1873–1911.

OECD. (2009). *Government at a Glance 2009*. Paris: OECD Publishing.

OECD. (2017). *Starting Strong 2017*. Paris: OECD Publishing.

OECD. (2018). *Education at a Glance*. Paris: OECD Publishing.

OECD. (2019a). *OECD Employment and Labour Market Statistics*. Paris: OECD Publishing.

OECD. (2019b). *OECD Family Database*. Paris: OECD Publishing.

OECD. (2019c). *OECD Government at a Glance 2019 Database*. Paris: OECD Publishing.

OECD. (2019d). *Prices and Purchasing Power Parities*. Paris: OECD Publishing.

Otto, A., Bártová, A. and Van Lancker, W. (2021). Measuring the generosity of parental leave policies. *Social Inclusion*, **9**, 238–249.

Pettit, B. and Hook, J. L. (2009). *Gendered Tradeoffs: Family, Social Policy, and Economic Inequality in Twenty-one Countries*. New York: Russell Sage.

Reskin, B. F. and Hartmann, H. I. (1986). *Women's Work, Men's Work: Sex Segregation on the Job*. Washington, DC: National Academies Press.

Rubery, J. (2013). Public sector adjustment and the threat to gender equality. In Vaughan-Whitehead, D. (Ed.), *Public Sector Shock*. Cheltenham: Edward Elgar Publishing, pp. 43–83.

Rubery, J. (2015). Austerity, the public sector and the threat to gender equality-geary lecture 2014. *The Economic and Social Review*, **46**, 1–27.

Schmidt-Catran, A. W., Fairbrother, M. and Andreß, H. -J. (2019). Multilevel models for the analysis of comparative survey data: common problems and some solutions. *KZfSS Kölner Zeitschrift für Soziologie und Sozialpsychologie*, **71**, 99–128.

Shalev, M. (2008). Class divisions among women. *Politics & Society*, **36**, 421–444.

Steinmetz, S. (2011). *The Contextual Challenges of Occupational Sex Segregation*. Wiesbaden, Germany: Springer.

Taiji, R. and Mills, M. C. (2020). Non-standard schedules, work–family conflict, and the moderating role of national labour context: evidence from 32 European countries. *European Sociological Review*, **36**, 179–197.

United Nations Statistics Division. (2019). *Demographic Statistics Database*. New York: United Nations Statistics Division.

Uunk, W. and Lersch, P. M. (2019). The effect of regional gender-role attitudes on female labour supply: a longitudinal test using the BHPs, 1991–2007. *European Sociological Review*, **35**, 669–683.

Van Belle, J. (2016). *Early Childhood Education and Care (ECEC) and its Long-term Effects on Educational and Labour Market Outcomes*. RAND Europe.

World Values Survey. (2015). World values survey 1981-2014 longitudinal aggregate v. 20150418. Dataset.

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Jennifer L. Hook is Professor of Sociology at the University of Southern California. Her research examines how social contexts, particularly social policies and opportunities in the labor market, affect gendered divisions of work and care. Her work examines the influence of country context on women's employment, fathers' time with children, and the division of household labor. Her publications appear in the American Sociological Review, American Journal of Sociology, and Journal of Marriage and Family, amongst others. She is co-author of Gendered Tradeoffs: Family, Social Policy, and Economic Inequality in Twenty-One Countries (Russell Sage Foundation 2009).

Meiying Li is a PhD candidate in Sociology at University of Southern California. She is a quantitative scholar with research interests in gender, stratification, economic sociology, and family policies. Her dissertation examines changes in occupational sex segregation and its effect on wages. She investigates how social contexts—family policy, occupational gender stereotype, and skill level – alter the negative effect of occupational feminization on wages.

Eunjeong Paek is a Max Weber Postdoctoral Fellow at the European University Institute. Her research interests are stratification, gender, work, family, and social policy. Her dissertation investigated how changing work contexts affect family-friendly policies with a focus on technological advances, union decline, and demographic shifts in the labor force. Her work has been published in Work, Employment and Society, the American Sociological Review, and the Population and Development Review.

Brigid Cotter is a PhD Candidate in the Department of Sociology at the University of Southern California, specializing in stratification, policy, labor, and family. Her cross-national work analyzes how the characteristics of family policies affect women's employment.

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Appendix

Table A1. Surveys.

	First observed	Last observed	# of times observed	Mean *N*	Minimum *N*	Minimum group *N*
Austria	1999	2016	18	11,984	9,454	345
Belgium	2001	2016	16	6,068	4,921	340
Czechia	2002	2016	15	9,017	2,197	59
Denmark	2010	2016	7	4,645	1,656	95
Estonia	2008	2016	9	1,295	1,049	27
Finland	2003	2016	14	7,920	6,552	477
France	2013	2016	4	4,797	4,635	310
Germany	2005	2016	12	14,957	2,380	130
Greece	2010	2016	7	10,776	9,032	507
Hungary	2006	2016	11	20,137	17,448	466
Ireland	2012	2016	5	11,364	8,755	366
Italy	2005	2016	12	27,710	21,993	601
Latvia	2008	2016	9	4,519	3,990	108
Lithuania	2008	2016	9	6,723	6,263	173
Luxembourg	2008	2014	7	1,121	688	54
The Netherlands	1999	2016	18	14,738	4,010	317
Portugal	1999	2016	18	7,274	5,504	66
Slovakia	2008	2016	9	3,853	3,303	79
Spain	1999	2016	18	15,685	4,004	230
Sweden	2009	2016	8	10,098	5,218	421
United Kingdom	2008	2016	9	4,931	4,250	267

Note: 'Minimum *N*' lists the *N* from the survey with the smallest number of observations for each country.

Table A2. Multilevel linear regression, four-way interactions.

	% Occ Women		% Occ Women w/NILF		Female-dominated		Male-dominated		Integrated		Management	
Woman	37.95	***	38.80	***	.37	***	−.47	***	.10	***	−.05	***
	(.25)		(.25)		(.00)		(.00)		(.00)		(.00)	
Child	.23	**	2.69	***	.02	***	.01	***	−.03	***	.01	***
	(.07)		(.10)		(.00)		(.00)		(.00)		(.00)	
Child × Woman	4.43	***	8.73	***	.06	***	−.04	***	−.02	***	−.02	***
	(.09)		(.08)		(.00)		(.00)		(.00)		(.00)	
Tertiary degree	5.07	***	3.13	***	−.03	***	−.12	***	.15	***	.06	***
	(.17)		(.15)		(.00)		(.00)		(.00)		(.00)	
Tertiary degree × Woman	−16.34	***	−15.10	***	−.11	***	.23	***	−.11	***	−.05	***
	(.08)		(.07)		(.00)		(.00)		(.00)		(.00)	
Child × Tertiary degree	1.88	***	1.11	***	.01	***	−.03	***	.01	***	.02	***
	(.08)		(.08)		(.00)		(.00)		(.00)		(.00)	
Child × Tertiary × Woman	−1.01	***	−3.05	***	.01	*	.03	***	−.04	***	−.04	***
	(.16)		(.15)		(.00)		(.00)		(.00)		(.00)	
ECEC spending	−.04		.00		.00		−.01	***	.01	*	.00	**
	(.07)		(.09)		(.00)		(.00)		(.00)		(.00)	
ECEC spending × Child	.05		.00		.00	***	.00	*	−.01	***	.00	*
	(.05)		(.06)		(.00)		(.00)		(.00)		(.00)	
ECEC spending × Woman	−.76	***	−.61	**	−.02	***	.00		.01	***	.01	***
	(.20)		(.20)		(.00)		(.00)		(.00)		(.00)	
ECEC spending × Tertiary	−1.09	***	−.79	***	−.01	***	.01	***	.01	*	−.01	***
	(.11)		(.10)		(.00)		(.00)		(.00)		(.00)	
ECEC × Child × Woman	.70	***	.48	***	.01	***	−.01	***	−.01	***	.00	
	(.05)		(.05)		(.00)		(.00)		(.00)		(.00)	
ECEC × Child × Tertiary	−.23	***	−.07		.00	***	.00		.00	*	.00	**
	(.06)		(.05)		(.00)		(.00)		(.00)		(.00)	
ECEC × Woman × Tertiary	−.08		.01		−.02	***	−.01	***	.03	***	.01	***
	(.05)		(.05)		(.00)		(.00)		(.00)		(.00)	
ECEC × Child × Woman × Tertiary	−.57	***	−.37	***	−.02	***	.00		.01	***	.00	*
	(.11)		(.11)		(.00)		(.00)		(.00)		(.00)	
Months paid leave <9 months	.15	***	.12	*	.00		.00		−.01	*	.00	
	(.05)		(.06)		(.00)		(.00)		(.00)		(.00)	
Leave < 9 months × Child	−.02		.04		.00	*	.00	*	.00		.00	***

Downloaded from academic.oup.com/esr/article/39/2/280/6750529 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Table A2. Continued

	% Occ Women		% Occ Women w/NILF		Female-dominated		Male-dominated		Integrated		Management	
	(.04)		(.06)		(.00)		(.00)		(.00)		(.00)	
Leave <9 months × Woman	−.55	**	−.56	**	.00		.01	***	−.01	**	.00	*
	(.18)		(.18)		(.00)		(.00)		(.00)		(.00)	
Leave <9 months × Tertiary	−.28	**	−.18	*	.00		.00		.00		.00	*
	(.10)		(.09)		(.00)		(.00)		(.00)		(.00)	
Leave <9 months × Child × Woman	−.19	***	−.19	***	−.01	***	.00		.01	***	.00	***
	(.04)		(.04)		(.00)		(.00)		(.00)		(.00)	
Leave < 9 months × Child × Tertiary	.09	*	.06		.00	***	.00	**	−.01	***	.00	**
	(.04)		(.04)		(.00)		(.00)		(.00)		(.00)	
Leave < 9 months × Woman × Tertiary	.75	***	.95	***	.01	***	.00		−.01	***	.00	
	(.04)		(.04)		(.00)		(.00)		(.00)		(.00)	
Leave <9 months × Child × Woman × Tertiary	.64	***	.61	***	.01	***	.00	**	−.01	***	.00	**
	(.09)		(.08)		(.00)		(.00)		(.00)		(.00)	
Months paid leave > 9 months	−.02	*	−.03	*	.00	***	.00	*	.00	***	.00	
	(.01)		(.01)		(.00)		(.00)		(.00)		(.00)	
Leave> 9 months × Child	−.01		.03	***	.00		.00		.00		.00	
	(.01)		(.01)		(.00)		(.00)		(.00)		(.00)	
Leave > 9 months × Woman	.25	***	.24	***	.00	***	.00	***	.00	*	.00	
	(.03)		(.03)		(.00)		(.00)		(.00)		(.00)	
Leave >9 months × Tertiary	−.03		.03	*	.00		.00	*	.00		.00	***
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Leave >9 months × Child × Woman	.06	***	.09	***	.00	***	.00	*	.00	***	.00	**
	(.01)		(.01)		(.00)		(.00)		(.00)		(.00)	
Leave >9 months × Child × Tertiary	.02	*	.07	***	.00		.00	***	.00		.00	***
	(.01)		(.01)		(.00)		(.00)		(.00)		(.00)	
Leave >9 months × Woman × Tertiary	−.04	***	.06	***	.00	***	.00	***	.00	***	.00	***
	(.01)		(.01)		(.00)		(.00)		(.00)		(.00)	
Leave >9 months × Child × Woman × Tertiary	−.05	**	.05	**	.00		.00		.00		.00	
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Public sector employment	−.04		−.08	*	.01	***	.00		−.01	***	.00	
	(.03)		(.03)		(.00)		(.00)		(.00)		(.00)	
Public sector × Child	−.01		−.14	***	.00	*	.00		.00	**	.00	**
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Public sector × Woman	−.07		−.35	***	.01	***	.00	***	−.01	***	.00	***

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Table A2. Continued

	% Occ Women		% Occ Women w/NILF		Female-dominated		Male-dominated		Integrated		Management	
	(.07)		(.07)		(.00)		(.00)		(.00)		(.00)	
Public sector × Tertiary	.59	***	.62	***	.01	***	.00	***	.00	*	.00	***
	(.04)		(.04)		(.00)		(.00)		(.00)		(.00)	
Public × Child × Woman	−.15	***	−.30	***	.00		.00	***	.00	***	.00	*
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Public x Child x Tertiary	.07	***	.05	**	.00		.00	***	.00	***	.00	***
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Public × Woman × Tertiary	−.08	***	.12	***	.00	***	.01	***	−.01	***	.00	***
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Public × Child × Woman ×	−.07		−.17	***	.00		.00	***	.00	***	.00	
Tertiary	(.04)		(.04)		(.00)		(.00)		(.00)		(.00)	
Individual controls												
Age	−.03	***	−.14	***	.00	***	.00	***	.00	***	.00	***
	(.00)		(.00)		(.00)		(.00)		(.00)		(.00)	
Age × Woman	−.02	***	−.16	***	.00	***	.00	***	.00	***	.00	***
	(.01)		(.01)		(.00)		(.00)		(.00)		(.00)	
Partner	−1.18	***	−.07		−.01	***	.02	***	−.01	***	.01	***
	(.05)		(.04)		(.00)		(.00)		(.00)		(.00)	
Partner × Woman	1.84	***	5.82	***	.01	***	−.03	***	.03	***	−.01	***
	(.09)		(.09)		(.00)		(.00)		(.00)		(.00)	
Foreign–born	.76	***	2.17	***	.04	***	.00		−.04	***	−.02	***
	(.06)		(.05)		(.00)		(.00)		(.00)		(.00)	
Foreign born × Woman	1.43	***	1.06	***	.05	***	.00		−.05	***	.01	***
	(.11)		(.10)		(.00)		(.00)		(.00)		(.00)	
Country-year controls												
Women's PT employment	.00		−.06	***	.00	*	.00	**	.00	***	.00	*
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Women's PT × Woman	−.05	**	−.05	**	.00	*	.00		.00	***	.00	***
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Gender norms	2.83	***	2.31	***	−.04		−.02		.07	*	−.01	
	(.54)		(.60)		(.03)		(.02)		(.03)		(.01)	
Gender norms × Woman	1.97		.24		.06	*	.00		−.06	**	−.02	**
	(1.31)		(1.30)		(.03)		(.01)		(.02)		(.01)	
Year	−.04	*	−.07	**	.00		.00		.00		.00	**
	(.02)		(.02)		(.00)		(.00)		(.00)		(.00)	
Year × Woman	.03		.01		.00		.00	**	.00	**	.00	***

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026

Table A2. Continued

	% Occ Women		% Occ Women w/NILF		Female-dominated		Male-dominated		Integrated		Management	
	(.07)		(.07)		(.00)		(.00)		(.00)		(.00)	
Women's share of LF	.46	***	.13	***	.01	***	−.01	***	.00		.00	
	(.03)		(.04)		(.00)		(.00)		(.00)		(.00)	
Women's Share × Woman	.63	***	−.13		.02	***	.00	*	−.02	***	.00	***
	(.10)		(.10)		(.00)		(.00)		(.00)		(.00)	
2011–2016 dummy	−.31	**	−.27		−.01		.02	***	−.01		−.02	***
	(.11)		(.14)		(.01)		(.00)		(.01)		(.00)	
Constant	47.43	***	51.34	***	.24	***	.30	***	.46	***	.09	***
	(.51)		(.30)		(.01)		(.01)		(.02)		(.01)	
Variance components												
Country: Intercept	5.39	***	1.71		.00	***	.00	***	.00	***	.00	***
	(1.90)		(.67)		(.00)		(.00)		(.00)		(.00)	
Country-Year: Woman	9.90	***	9.88	***	.00	***	.00	***	.00	***	.00	***
	(.99)		(.97)		(.00)		(.00)		(.00)		(.00)	
Country-year: Tertiary	6.35	***	4.54	***	.00	***	.00	***	.00	***	.00	***
	(.64)		(.47)		(.00)		(.00)		(.00)		(.00)	
Country-year: Child	.63	**	1.70	***	.00	***	.00	***	.00	***	.00	***
	(.10)		(.20)		(.00)		(.00)		(.00)		(.00)	
Country-year: Intercept	.12	***	.25	***	.00	***	.00	***	.00	***	.00	***
	(.02)		(.03)		(.00)		(.00)		(.00)		(.00)	
var(Residual)	543.20	***	555.50	***	.13	***	.15	***	.23	***	.06	***
	(.54)		(.50)		(.00)		(.00)		(.00)		(.00)	
N	2,060,474		2,522,299		2,060,474		2,060,474		2,060,474		2,060,824	
AIC	18,825,502		23,101,203		1,670,812		1,993,911		2,839,272		181,770	
BIC	18,826,267		23,101,980		1,671,577		1,994,676		2,840,037		182,534	

Notes: Standard errors in second row.
*P < .05, **P < .01, ***P < .001.

Downloaded from academic.oup.com/esr/article/39/2/280/6755029 by EERP/BIBLIOTECA CENTRAL user on 07 November 2026