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National Work-Family Policies and Gender Earnings Inequality in 26 OECD Countries, 1999–2019

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National Work-Family Policies and Gender Earnings Inequality in 26 OECD Countries, 1999–2019

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Abstract

We investigate whether work-family policies help incorporate women into the labor market, but exacerbate the gender earnings gap and motherhood penalty, especially for mothers and/or tertiary-educated women. We use repeated cross-sectional income data from the Luxembourg Income Study Database (LIS) 1999–2019 (n = 26 countries, 280 country-years, 2.9 million employees) combined with an original collection of indicators on work-family policies, labor market conditions, and gender norms. We find only one work-family policy, long paid parental leave (> 6 months), is associated with a larger gender earnings gap for mothers and tertiary-educated women. The negative relationship between long paid leave and women's earning percentile is not well explained by selection, full-time status, work hours, experience, occupation, or sector, suggesting discrimination mechanisms. Our findings add to the growing evidence that long paid leave specifically, as opposed to work-family policies more generally, cleave the labor market outcomes of women from men.

Keywords: motherhood penalty, gender wage gap, work-family policy, ECEC, parental leave

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INTRODUCTION

Over the last half-century, mothers' paid employment has become normative across Western countries and national-level work-family policies have proliferated and evolved, although not always in tandem (Ferragina and Seeleib-Kaiser 2014). There is a large literature on the effects of these policies – such as paid parental leave and early childhood education and care (ECEC) - on women's and mothers' employment and labor market outcomes (for recent reviews see Ferragina 2019, Rubery and Figueiredo 2018). One highly cited argument, the "welfare state paradox," posits that while well-developed work-family policies may increase women's employment, they also perpetuate gender earnings gaps, particularly for higher-skilled women (Mandel and Semyonov 2005, Mandel and Semyonov 2006, Mandel 2012). Work-family policies may affect both the gender earnings gap and the motherhood penalty through a variety of pathways, including selection into the labor market, the accumulation of human capital, occupational and sectoral segregation, and employer discrimination. Despite relatively clear theoretical linkages between work-family policies and earnings gaps, the evidence is mixed, especially regarding gaps by educational attainment (for recent review see Hook and Li 2020).

Although the gender and motherhood earnings gaps have received significant attention in the comparative literature, there are several limitations to our understanding of the relationships between work-family policies and earnings gaps. First, the gender and motherhood gaps are rarely theorized and analyzed within the same framework (for exceptions see Cha, Weeden and Schnabel 2023 in the US, Cukrowska-Torzewska and Lovasz 2020). That is, most studies are limited to comparisons of women and men or variation among women, which hinders our understanding of how policy may impact each gap differently. For example, if long periods of parental leave make motherhood a more salient concern for employers, it may promote statistical

discrimination against all women of childrearing age (Glass and Fodor 2011), widening the gender earnings gap, but closing the motherhood gap (Halldén, Levanon and Kricheli-Katz 2015). Second, although scholars have examined skill-differentiated policy effects in single countries or case studies (see Cooke, Hägglund and Icardi 2022, Grönlund and Magnusson 2016), there is little recent comparative work that directly examines policy effects on earnings by skill-level (see Halldén, Levanon and Kricheli-Katz 2015).

Third, due to limitations in comparable microlevel data, there has been a lack of testing in a cross-national context of detailed individual-level mechanisms, such as occupation, through which policy may affect earnings, leaving our understanding of the importance of these individual-level mechanisms for policy associations weak. And as others have noted (Brady, Blome and Kmec 2020, Cooke, Hägglund and Icardi 2022), many highly cited studies rely on cross-sectional data from approximately 20 countries, and often from the mid-1990s. Thus, there are concerns about measuring “policy effects” with a small N at the macro-level, as well as basing strong conclusions on a particular snapshot in time, while both work-family policy and the gender revolution have continued to evolve (Barth, Reisel and Misje Østbakken 2023).

We address two questions: (1) Do work-family policies explain variation in the size of the gender earnings gap and/or the motherhood penalty and does this vary by educational attainment? (2) Do commonly hypothesized micro-level mechanisms explain observed relationships between work-family policies and the size of earnings gaps? To answer these questions, we use repeated cross-sectional income and labor market data from the Luxembourg Income Study Database (LIS) 1999–2019 ($n = 26$ countries, 280 country-years, 2.9 million employees) in a comparative longitudinal survey design. The data span two decades of the 21st century, allowing us to leverage variation between countries and within countries over time. We

combine this data with an original collection of country-year indicators on early childhood education and care (ECEC) spending, paid parental leave length, public sector size, part-time employment rates, and gender norms. We predict earnings percentile to control for the effects of wage structure in conditioning work-family policy effects. We advance the literature by theorizing and analyzing the relationship between work-family policies and the gender and motherhood earnings gaps in same framework, while focusing on differences by educational attainment and testing individual-level pathways through which policies are hypothesized to affect earnings.

We find no evidence that work-family policies that promote women's inclusion into the labor market -- ECEC and short paid leave (< 6 months) -- have negative consequences on earnings for any group of women. They do not explain variation in the motherhood penalty or the gender earnings gap. Longer paid leave (> 6 months), however, is consistently associated with larger gender earnings gaps between mothers and fathers, and between tertiary-educated childless women and men, but not non-tertiary educated childless women and men. The association of longer paid leave with the gender earnings gap is not well explained by selection, full-time status, work hours, experience, occupation, or sector. We conclude that findings are consistent with the argument that long periods of leave make (the capacity for) motherhood salient for employers, worsening gender-based discrimination (Glass and Fodor 2011). Our findings add to the growing evidence that long paid leave specifically, as opposed to work-family policies more generally, cleave the labor market outcomes of women from men.

THEORY AND EMPIRICS

We begin with a review of theory and findings on how the two most studied work-family policies -- ECEC and paid parental leave -- are associated with the gender earnings gap and

motherhood penalties. We distinguish “short” leave, which generally refers to a period of well-paid leave around the birth of a child within the range of 3-9 months (Cukrowska-Torzewska 2017), from “long” leave up to three years. Central to understanding the nuances of this research is that work on the motherhood penalty posits that work-family policies impact the earnings gap between mothers and childless women (or women before/after they become mothers). Whereas research on gender earnings gap posits that work-family policies impact the earnings gap between women and men (many of whom are parents). We detail the mechanisms implicated in potential policy effects, including selection into employment, micro-economic explanations focused on the characteristics of workers and the work they do (i.e., occupation and sector), and discrimination.

Another central concern is whether earnings differentials vary by skill-level and whether any effects of work-family policies on earnings gaps also vary by skill-level. On the one hand, the motherhood penalty may be exacerbated among highly skilled mothers because highly skilled work imposes a larger penalty for human capital depreciation, requires greater work effort and time, and employers have more latitude in wage setting given larger variation in wages.

Relatedly, discrimination may be most acute for women with high human capital, who are more costly for employers to replace (England et al. 2016, Halldén, Levanon and Kricheli-Katz 2015, Mandel 2012). On the other hand, the motherhood penalty may be buffered among highly skilled mothers because they have the financial means to outsource care, their education signals work commitment and affords greater bargaining power, they are less likely to spend long periods out of the labor market, and they have jobs providing greater flexibility and autonomy, enabling greater work-family balance (England et al. 2016, Halldén, Levanon and Kricheli-Katz 2015). Recent research finds that lower educated women tend to face larger motherhood penalties (Cukrowska-Torzewska and Matysiak 2020, Deming 2022, Doren 2019, Glauber 2018). Yet,

findings are mixed about whether work-family policies affect earnings more strongly for higher- versus lower-skilled women (Hook & Li 2020).

Selection into Employment

Short maternity/parental leave and publicly funded childcare increase mothers' employment. As Mandel and Shalev (2009) argue, if mothers with low skills and little career motivation enter the labor force, they are likely to earn less, potentially widening both gender and motherhood gaps. Job-protected leave boosts participation both prior and after leave, speeding return to work if leaves are not too long (Ruhm 1998). Childcare provision is associated with higher levels of employment among mothers and mothers with young children (Boeckmann, Misra and Budig 2015, Nieuwenhuis, Need and Van Der Kolk 2012, Pettit and Hook 2005, Pettit and Hook 2009, Uunk, Kalmijn and Muffels 2005) and lower educated mothers of young children (Hook and Paek 2020, Scherer and Pavolini 2023).

Long parental leave, in contrast, depresses mothers' employment. Thus, potential selection effects could narrow both gender and motherhood gaps if long leave disproportionately incentivizes lower skilled mothers to leave (or stay out of) the labor market. Recent research, however, finds that leaves over six months suppress the employment of higher educated mothers of young children, more so than lower educated mothers (Hook and Paek 2020). Overall, researchers generally find that leaves anywhere from six months (Hook and Paek 2020) to over two plus years depress rather than facilitate maternal employment (Nieuwenhuis, Need and Van der Kolk 2017, Thévenon and Solaz 2013), although some estimates approach three years (Pettit and Hook 2009), and others find that leaves cannot be too long (Keck and Saraceno 2013). Results for "long" leave are difficult to compare given different measures and methods.

Micro-Economic Explanations

Our understanding of potential policy effects on earnings is closely linked to micro-economic explanations of earnings -- where earnings are determined by the characteristics of workers and the work they do. Short leave and childcare are expected to impact the characteristics of workers (i.e., work hours and work experience) by allowing mothers to work more continuously and longer hours, narrowing both gender and motherhood gaps. Job-protected leave makes it more likely that mothers will return to full-time work instead of seeking part-time work (Akgunduz and Plantenga 2012). Childcare also has positive effects on mothers' working hours (Andringa, Nieuwenhuis and Van Gerven 2015, Boeckmann, Misra and Budig 2015), preferences for longer working hours (Pollmann-Schult 2016), and full-time employment for lower educated mothers of young children (Hook and Paek 2020). Greater involvement in full-time work and greater work continuity create a net gain in work experience (Ruhm 1998), potentially narrowing both gender and motherhood gaps.

Short leave and childcare are also expected to affect the work mothers do (i.e., occupation and sector) by helping mothers maintain attachment to employers and outsource some of mothers' childcare responsibilities, both of which should lower mothers' need to trade lower wages for better amenities (e.g., flexibility, security); although this explanation for the motherhood wage penalty has received little support (Glauber 2012). ECEC has been associated with highly educated women working in less feminized occupations (Hook et al. 2023), and their greater representation in professional (Pettit and Hook 2009) and managerial positions (Steinmetz 2011). ECEC is also associated with smaller occupational prestige penalties for mothers (Abendroth, Huffman and Treas 2014).

One potential consequence of ECEC, however, is the creation of a large workforce employed in feminized care occupations, in many countries within the public sector (Mandel and

Shalev 2009, Steinmetz 2011). Although “good” for lower educated women who benefit from higher wage floors in the public sector, large public sectors may be detrimental to higher skilled women’s wages, concentrating them in feminized occupations and limiting attainment of high wages found in the private sector (Mandel and Semyonov 2006, Mandel and Shalev 2009, Shalev 2008). Whether higher skilled women’s wages are penalized for public sector employment is contentious and findings vary across countries (Christofides, Polycarpou and Vrachimis 2013). Both tertiary-educated men and women are more likely to work in feminized occupations where the public sector is larger (Hook et al. 2023, Steinmetz 2011). Yet overall, public sector size is associated with *less* occupational sex segregation (Barth, Reisel and Misje Østbakken 2023). Despite contradictory findings about the “goodness” of public sector employment for highly educated women, short leave and ECEC should allow mothers to better compete with both men and childless women in the private sector.

In contrast, long leave makes mothers less similar to men and childless women in their work hours and work experience, potentially increasing both gender and motherhood gaps. The length of paid parental leave over six months is associated with a lower likelihood of full-time employment for higher educated mothers of young children (Hook and Paek 2020). During long periods of leave, mothers fail to accumulate (paid) work experience and their human capital depreciates, damaging wages (Ruhm 1998). Long leave is also anticipated to affect the work mothers do. Long periods out of the labor market may change mothers’ work commitment (Gangl and Ziefle 2015) and mothers may shift occupations after long periods out of employment (Lovejoy and Stone 2012). In Sweden, taking 16 months of leave or more has a negative effect on mothers’ upward occupational mobility upon return to work (Evertsson and Duvander 2011). Research in Norway suggests that leave may be particularly damaging for

women in highly competitive, client-driven, environments (Halrynjo and Mangset 2022).

Ultimately, even highly educated women may seek shelter in the public sector, to the detriment of their wages as discussed above (Mandel and Semyonov 2006).

In micro-economic explanations, the mechanisms are about how policy affects the behavior of individual women who use - or expect to use - work-family policies. Thus, entering these micro-level mechanisms into our models should explain much of any work-family policy effects. Short periods paid leave and ECEC should narrow both the gender and motherhood gaps by increasing mothers' full-time work, work hours, work experience, and employer continuity (negating downward occupational moves or shifts from the private to public sector). Conversely, long leave should have the opposite effect. We anticipate any policy effects will be stronger for the higher educated because of greater returns to human capital. This implies that micro-economic explanations may be especially helpful for explaining policy effects among the highly educated but may be less helpful in explaining policy effects for those with lower levels of education for whom the returns to human capital are lower.

Discrimination

Another strand of research focuses on discrimination mechanisms. There is robust evidence that employers discriminate against mothers. A meta-regression of the motherhood penalty by de Linde Leonard and Stanely (2020) concludes, "employer discrimination or a [employer] perception of lost productivity is the best explanation for the motherhood wage penalty that remains after intensive study by dozens of research studies and controlling for alternative explanations." Even when hired into the same firms, mothers may be offered lower starting salaries and fewer opportunities for training and advancement (Correll, Benard and Paik 2007, Glass and Fodor 2011), contributing to motherhood and gender wage gaps.

In discrimination-based explanations, the mechanism is about how policy affects the behavior of employers toward women who use - or who employers expect to use - work-family policies. Contrary to the argument that generous family policies increase discrimination in private sector employment (Mandel 2012), Cooke and colleagues (2022) argue that dual-earner-carer policies – such as ECEC - decrease gender differences in productivity among parents, thereby reducing statistical discrimination against women more broadly.

Long leave, however, may contribute to greater propensities to discriminate against mothers, and potentially, to all women of childbearing age as employers are incentivized to evaluate potential fertility or “future childbearing risk” (Glass and Fodor 2011, Halldén, Levanon and Kricheli-Katz 2015, Hutt Cabello 2023), contributing to occupational and sectoral segregation as employers funnel women into occupations where absences are less costly (Ruhm 1998). Research finds that parental leave in excess of 9 month promotes occupational segregation between men and women broadly, but most acutely for less educated mothers (Hook et al. 2023). In Hungary where leaves are up to three years, discrimination against women of childbearing age is strong (Glass and Fodor 2011). Discrimination may be most acute for women with high human capital, who are more costly for employers to replace (Mandel 2012), although Cooke and colleagues (2022) counter that lower-skilled women are likely to be seen by employers as the least reliable absent work-family policy supports.

From a discrimination perspective, entering micro-economic mechanisms into our models may not explain much of any work-family policy effects as the mechanism is not about the individual characteristics of women and the work they do, but what employers expect mothers, and thus women of childbearing age, to do given prevailing policies. We anticipate that ECEC and short leave decrease discrimination by reducing *perceived* gender differences in productivity

among parents, and long leave increases it. We anticipate any policy effects will be stronger for the higher educated because of greater incentives to discriminate against mothers and potential mothers with high human capital. This implies larger policy effects for the higher educated would persist after controlling for micro-economic mechanisms.

Prior Research Findings

Empirical findings on how short leave and childcare affect gender and motherhood earnings gaps are mixed. Several studies find that ECEC and moderate amounts of paid leave narrow the motherhood wage gap (Budig, Misra and Boeckmann 2016, Halldén, Levanon and Kricheli-Katz 2015, Misra, Budig and Boeckmann 2011). But, others find no association between ECEC and the motherhood wage gap (Brady, Blome and Kmec 2020) or the gender wage gap (Brady, Blome and Kmec 2020, Mandel and Semyonov 2005)². The limited research that examines whether policy associations differ by mothers' or women's skill-level find no support for differences in associations with ECEC, but mixed findings on paid leave measured by full-time equivalent weeks (weeks x replacement rates) (Halldén, Levanon and Kricheli-Katz 2015, no difference, Mandel 2012, detrimental to women earning above the median).

Long leave has been linked to larger motherhood wage gaps (Cukrowska-Torzewska 2017, Misra, Budig and Boeckmann 2011), and larger gender gaps (Christofides, Polycarpou and Vrachimis 2013) particularly for highly educated women (Pettit and Hook 2009). Cukrowska-Torzewska and Lovasz (2020) link high motherhood gaps in Eastern European countries to parental leave policies that encourage long separations from work, over one year. They also find that the gender wage gap among the childless is strongly correlated with leave length. This is

² The age range of 25-54 analyzed by Brady et al. (2020) and Mandel & Semyonov (2005) likely categorized many mothers with grown children as non-mothers, potentially underestimating differences between the two groups; see Greulich & Dasré (2018).

consistent with other work that suggests work-family policies affect gender wage inequality through generalized effects on women and men (Cooke, Hägglund and Icardi 2022). Yet, it is important to note that some studies find no negative effects of long leave on earnings (Brady, Blome and Kmec 2020). Although some studies controlled for aspects of workers' characteristics (e.g., work hours, experience) and their work (e.g., occupation and sector), they did not focus on examining these microlevel pathways that may explain policy associations.

RESEARCH STRATEGY

Microlevel Data

We use microdata from LIS (LIS 2024). LIS serves as a repository of nationally-representative, cross-sectional income surveys collected in over 50 countries, spanning 5 decades. Through rigorous data harmonization, LIS facilitates comparisons across countries and time, offering standardized, high-quality income data with large sample sizes. The observation period is 1999 to 2019. Due to a series break in OECD ECEC data, we began our macro-data series in 1998, corresponding to 1999 using a one-year lag. Country and survey selection was based on data availability constraints, including: (a) baseline micro-level data, (c) thresholds for missing data described below, (c) available comparable macro-data, and (d) observations at two or more time points. We include data from 26 OECD countries: Australia, Austria, Belgium, Czechia, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Lithuania, Luxembourg, Netherlands, Norway, Poland, Slovakia, Slovenia, Spain, Sweden, Switzerland, the UK, and the US.

We restricted the analyses to employees aged 25 to 44, omitting the self-employed consistent with prior research (Budig, Misra and Boeckmann 2016). We retained respondents through age 44 (e.g., Boeckmann, Misra and Budig 2015) because we identified parents by the

presence of own household children. The presence of children underestimates women's number of children from age 40 onward (Greulich and Dasré 2018), which may narrow the differences we observe between parents and childless respondents. Our data is dominated by those born in the 1970s, who we observe in their 20s through 40s. Birth cohorts from the 1960s and 1980s are well represented in the data. Table 1 of the Online Supplement shows the progression of cohorts.

We omitted surveys (n=11) with problematic patterns of missing data often arising from large missingness on one variable. Retained surveys have missing data for fewer than 7.1 percent of cases (mean = 0.85%, median = 0.27%); after 7.1 missingness jumps to 44 percent providing a natural break point for inclusion. We deleted respondents with missing data listwise. The final sample size is 2,957,461 currently employed respondents nested within 280 country-years from 26 countries. Appendix Table 1 lists the countries, years, and sample size per country. Measures created from the data and descriptive statistics are weighted (*ppopwgt*). Weights are provided by each data producer and inflated to the total national population; they typically correct for sampling bias, but not always for non-response (LIS 2019).

Microlevel Measures

Our dependent variable is annual earnings (*pill wage income*) converted to earnings percentile to standardize earnings for overall wage inequality and cross-national differences in currency and inflation (Bar-Haim et al. 2023, Brady, Blome and Kmec 2020, Mandel and Semyonov 2005). It measures each respondents' relative rank within each country-year's distribution of employees. As argued by Mandel and Semyonov (2005), countries with more compressed earnings distributions tend to have smaller gender pay gaps *and* more developed work-family policies. So, to understand how family policies affect the gender earnings gap, we need to remove differences in how earnings are distributed.

Independent variables at the micro-level are gender (*sex*), parenthood (*nchildren*), and educational attainment (*educ*). We code respondents as parents if they have own children living in the household. Although the motherhood penalty literature often analyses a per child penalty, we use a binary variable for parenthood. We report variation in the mean number of children per mother in Appendix Table 1. We also replicated the analyses using the number of children and it did not change our results for the motherhood penalty. LIS provides three educational categories: less than upper secondary education (ISCED 0-2), upper secondary and postsecondary non-tertiary (ISCED 3-4), and tertiary education (ISCED 5-8). We use all three categories in our selection equation (described below), but for our earnings equation we collapsed education into with and without a tertiary degree because of small cell sizes for the interaction between education and parenthood if low is separated from medium attainment. Although most non-tertiary respondents have medium levels of education, particularly in later years when it is rare not to finish secondary education, this amalgamation likely obscures meaningful heterogeneity.

Additionally, we controlled for age (*age*) and its square, and partnership status (*marital*). Because the prevalence and meaning of cohabitation varies across countries (Hiekel, Liefbroer and Poortman 2014), we coded respondents as partnered if they were married or cohabiting.

We include additional micro-level variables for subsamples as available. We include full-time status (*fyft*) for subsample of 23 countries, work hours measured by total weekly hours worked (*hourstot*) or weekly hours worked at main job (*hours1*) for 21 countries; years of total work experience (*wexptl*) for 12 countries; occupation at main job in 10 International Standard Classification of Occupations (ISCO) categories (*occb1*) for 22 countries; and a public sector dummy for main job (*public1*) for 15 countries. Sample sizes are in Table 1.

Macrolevel Data

We build on an original collection of macrolevel data (Hook and Paek 2020; Hook et al. 2023) for each of the 280 country-years in the microdata. Our focus is on ECEC spending and paid parental leave length. Given that we rely on variation between countries, we are concerned that ECEC spending and parental leave length may be a general proxy for how work and care are more generally arranged and gendered. Thus, we include three control variables assessing public sector size, part-time employment rates, and gender egalitarianism. Wage structure is addressed by using earnings percentile as the dependent variable. Our macro level variables vary by country-year ($n = 280$). Parental leave, however, varies less than other country-year measures, with 17 countries changing leave length at least once for a total of 59 unique observations. Macrolevel measures were lagged one year prior to the microlevel survey. We used interpolation to fill in unavailable data. Year is included as a linear time trend. Country-year variables were grand-mean centered.

Spending on early childhood education and care (ECEC). We use ECEC spending per child under the primary school starting age as it provides a standardized measure of public support for ECEC across a large set of countries and over time. Spending may include spending on public facilities, subsidies to providers of center and home-based care, and transfers to families for purchasing care (European Commission 2014). We construct this variable using the public expenditure on ECEC as a percentage of GDP (OECD 2019a) converted into GDP using the purchasing power parities (PPP) adjusted it to constant international dollars (World Bank 2019), and then divide it by the number of children (United Nations Statistics Division 2019) under school starting age (OECD 2018). In some countries, the data do not fully capture local-government spending. Spending and enrollment are strongly correlated (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 periods of low-paid parental leave or home-care allowance. We are theoretically interested in the maximum length of time that the state will pay a mother to care for her child at home, regardless of the replacement rate; we refer to this time collectively as parental leave. From the OECD Family Database (OECD 2019a), we take the maximized leave length in weeks for mothers and divide it by 4.3 to convert to months.

Theoretically it is unclear at which point national leave would be too long to trigger adverse impacts on earnings. Previous work has used a quadratic model to capture non-linearity. This model, however, imposes a specific functional form which may over- or under-fit parts of the curve and can falsely identify the turning point, leading to spurious conclusions. Cubic or spline models allow for more flexible curvature. We tested various specifications including linear, quadratic, cubic, 4-order polynomial, restricted cubic splines (with 3-5 knots), splines with one knot at either 3, 6, 9, and 12 months, and based on visualizations from cubic spline models, splines with two knots – one with knots at 6 and 21 and one with knots at 9 and 21, which would capture diminishing declines at greater lengths. A spline knot at 6 months best captured non-linearity and produces best model fit (AIC and BIC), but overall, there was little difference in substantive interpretation between 6 and 9 months (Online Supplement Appendix A). Six months is longer than the ILO's recommendation of at least 18 weeks, but consistent with UNICEF's recommendation of at least six months to promote children's healthy development (UNICEF 2019). Thus, one continuous variable quantified the effect of the number of months 0 to 6, and a second months greater than 6.

Public sector size. The correlation of public sector size with ECEC and its potential implications for earnings gaps is addressed above. 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 2019b). We supplement this series with data for 1995 and 2005 from Government at a Glance 2009 (OECD 2009).

Women's part-time employment rate. Part-time work's development, availability, use, characteristics, and quality vary greatly across European countries (Fagan, O'Reilly and Rubery 2000) as do its consequences for women's employment (Barbieri et al. 2019). As the part-time employment rate is moderately correlated with work-family policies 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 2019c).

Gender egalitarianism. Gender norms are associated with work-family policies and may be linked to earnings gaps through labor supply effects and/or earnings premia/penalties for gender and parenthood (Boeckmann, Misra and Budig 2015). Thus, we include gender norms as a possible confounder. We used data from the European Values Surveys (EVS; 1990-1992, 1999-2000, 2008-2009, 2017-2018). We supplement the data for the United States using the General Social Survey (GSS; 1998-2019) and for Australia using the World Values Surveys (WVS; 2012–2018) (EVS 2015, EVS 2019, General Social Survey 2020, WVS 2015). Consistent with previous research, and due to data constraints, we use a measure gauging norms about maternal employment (Boeckmann, Misra and Budig 2015). The item asks respondents how much they agreed with the statement "A pre-school child is likely to suffer if his or her mother works" on a four-point Likert scale. We create a continuous measure of the country-year mean where higher scores indicated stronger support for egalitarianism.

Method

We use a two-step, estimated dependent variable (EDV) approach. First, using the individual-level data (level 1) we run 280 models predicting earnings percentile in each country-year (level 2) and output these marginal effects by gender, parenthood status, and tertiary education. Second, we use these marginal effects as dependent variables, nesting them within countries (level 3) to model the association between work-family policies and the size of estimated gender and motherhood gaps.

We use an EDV approach because it accommodates complex models at the individual-level, here a Heckman correction model. Multilevel models could not simultaneously estimate a 3-level (individuals nested within country-years and countries) Heckman model with random slopes required for cross-level interactions (Heisig and Schaeffer 2019). EDV allows all level-1 variables to be freely estimated (i.e., random slopes for gender, education, parenthood, and control variables). EDV approaches are routinely used in similar research (e.g., Fauser and Gebel 2023, Pettit and Hook 2009).

At the individual-level, we first use OLS to model earnings percentile in each country-year as a function of gender x parenthood x education, controlling for age and its square, and partnered x gender. This initial base model is used to produce the predicted marginal effects of gender and motherhood (while other covariates are set to country-year means), which provides the size of the gender earnings gap and motherhood penalty for each country-year. From these models we also estimate gender earnings gaps by parenthood and educational attainment and motherhood penalties by educational attainment.

To examine selection into the employee sample we use a Heckman correction. Consistent with previous research we include five variables and one interaction term in our selection

equation; Descriptive statistics appear in Table 1. Following Budig and colleagues (2012), we use a probit model to predict employment (*emp*) in each country-year using (1) transfer income (*hitransfer*: pensions, public social benefits, and private transfers), (2) other household labor market earnings (*hilabour* – *pilabour*, household earnings from employment minus respondent's earnings), and (3) the presence of a preschool-aged child (age 0-5). We add (4) education (Mandel and Semyonov 2005, Mandel 2012, Misra, Budig and Boeckmann 2011) in three levels (low, medium, and high). Moreover, because sample includes men and women, we include (5) gender and its interaction with a preschool child (Mandel 2012). We tested models including additional interactions, but model fit statistics worsened. The presence of a preschooler is the instrumental variable. Theoretically, we anticipate that a young child constrains women's employment in a way that older children do not; and that conditional on being employed, a younger child should not affect women's earnings more than an older child affects earnings. Models (using pooled data on the female sample) indicate that the presence of a preschooler negatively affects the probability of employment (relevance) without negatively affecting earnings among the employed (exclusion restriction), controlling for either the presence or number of children. Although this variable is theoretically justifiable and performs as anticipated, it may not be entirely exogenous and thus, estimated coefficients may still suffer from bias. We used the *heckman* command in STATA, which corrects for underestimated SEs in second stage and incorporates this information in calculation of marginal effects. This is essential because gender appears in both the selection and outcome equations. These marginal effects are then used as outcomes in the next set of models.

To model the association between the estimated marginal effects and work-family policies, we use OLS with cluster-robust standard errors with a degrees-of-freedom adjustment

(*hc2 cluster, dfadjust*). Marginal effects of country-years ($n = 280$) are nested within countries ($n = 26$). This approach is generally superior, and more conservative, than weighted least squares (WLS) (Lewis and Linzer 2005). We find that our method is more conservative, producing slightly higher standard errors. Results, however, are robust to WLS and wild cluster bootstrapping (MacKinnon, Nielsen and Webb 2023) as shown in Table 2 of the Online Supplement. Equations are shown in Appendix B of the Online Supplement.

All files necessary for replication are publicly available via the Harvard Dataverse (Hook and Li 2025). Secure access to the LIS microdata is available by submitting code through LISSY, a web-based Job Submission Interface.

RESULTS

Descriptive

Table 1 shows descriptive statistics for micro-level variables and observed minimum and maximum survey-level means ($n = 280$), separately for employed women and men aged 25-44. The mean annual earnings percentile for men is 57.3 compared to women's 42.6. Men's advantage varies, however, from a minimum of near equality (51.4) to nearly 2/3 on average (64.8). Table 1 also shows that women are about 8 percentage points more likely to be tertiary-educated, but there is wide variation in the proportion of the population that is tertiary-educated across time and place. On average women, ages 25-44, are about 9 percentage points more likely than men to live in a household with a child, but again this varies widely. Partnership and age are, on average, similar for men and women. Because we are interested in the intersection of gender, parenthood, and education, Table 1 reports the group size for parenthood by education, separately for women and men. For both, non-tertiary parents are the largest group on average, although this varies across surveys. Table 1 also shows the variables included in the Heckman

selection equation, predicting employment. On average, men's employment rate is 14.7 percentage points higher than women's at ages 25-44.

[TABLE 1 ABOUT HERE]

The second half of Table 1 shows additional measures of work and workers' characteristics, along with information on the smaller subsamples where these measures are available. The full-time sample is restricted to 23 countries. On average, employed men are 18.6 percentage points more likely to be working full-time than women. In the work hours sample, restricted to 21 countries, men work 6.5 hours per week more than women on average, and women have a larger SD revealing greater variation in their work hours. In the experience sample, limited to 12 countries, men have 1.2 years more experience, on average than women. In the public sector sample, 15 countries, women are 7.8 percentage points more likely to be working in the public sector. In the occupation sample, restricted to 22 countries, men are more likely than women to be employed as managers, craft/trade workers, plant/machine operators, agricultural workers, and in armed forces occupations.

Table 2 shows macro-level descriptive statistics for 280 surveys, beginning with our estimates as dependent variables (EDVs). We show the marginal effect of gender and parenthood on annual earnings percentile as estimated from 280 OLS models (unadjusted; observed earnings gaps among the employed) and 280 Heckman selection models (adjusted; estimated earnings gaps among the population). The correction for selection reduces the gender earnings gap by about 9 percent and the motherhood penalty by 33 percent. The role of selection varies substantially across surveys. At the mean, adjusting for selection reduces the gender earnings gap -1.8 percentage points, but across the 280 surveys this ranges from -5.4 to +1.8. For the motherhood gap the mean is -3.7, with a range of -11.1 to +1.3 (results not shown).

[TABLE 2 ABOUT HERE]

Focusing on the adjusted results, the gender earnings gap is estimated to be -17.3 indicating that women, on average, are 17.3 percentage points lower in the earnings distribution than men. First, we notice the gender earnings gap (comparing women to men) is much larger, on average, than the motherhood penalty (comparing mothers to childless women). This is partly mechanistic, as the gender earnings gap includes motherhood penalties and/or fatherhood premiums. Second, although smaller, variation in the motherhood penalty is larger with the range spanning positive values. Figure 1 shows trends in both gaps by country, revealing substantial variation across and within countries. Figure 1 of the Online Supplement displays this variation as a histogram. Variance within country accounts for 28 percent of the variation in the gender wage gap and 19 percent in the motherhood penalty.

[FIGURE 1 ABOUT HERE]

The next four estimates of Table 2 show the gender earnings gap by parental status and educational attainment. The adjusted gender gap is -19.3 for tertiary-educated parents. Although these marginal effects were directly estimated, this gap can be calculated from other estimates in the table. It is the sum of the gender gap for the tertiary-educated childless (-8.8), the motherhood penalty for tertiary-educated women (-5.7) and the fatherhood premium for tertiary educated men -(4.9). Thus, not surprisingly, gender gaps are larger for parents than for the childless. Yet, the gender gaps we observe among the childless are still larger than the motherhood penalty and are roughly half the size of those we observed among parents. We also note that selection correction does little to alter the estimated gaps among the childless. Mean gender gaps and the motherhood penalty are also slightly larger for those without tertiary degrees compared to those with degrees.

The second half of Table 2 details the independent work-family policy and contextual measures we use to predict the EDVs. We observe wide variation in all macro-level measures. Appendix Table 1 reports variation between and within countries on all indicators.

Multivariate Models

(1) Do work-family policies explain variation in the size of the gender earnings gap and/or the motherhood penalty and does this vary by educational attainment?

Table 3 shows the effects of country-year level variables on the marginal effects of interest – the motherhood penalty and the gender earnings gap. Models 1a and 2a use marginal effects from OLS as the estimated dependent variable, whereas Models 1b and 2b use marginal effects from the Heckman specification. Motherhood penalty models (1a and 1b) show that measured work-family policies do not explain variation in the size of the motherhood penalty. Gender earnings gap models (2a and 2b), however, show that only one policy, long paid parental leave (> 6 months), is associated with a larger gender earnings gap. Comparing models with and without Heckman correction (2a and 2b), the coefficient for long paid leave decreases by 10 percent moving from the OLS model (2a) to the Heckman correction (2b), similar in size to the reduction of the overall gender gap reported in Table 2.

Next, we examine how robust our macro-level results are across the main and subsamples used below. Models 3-7 of Table 3 replicate Model 2b on the subsamples. Despite the large difference in sample sizes which range from 26 countries (280 country-years, about 2.96 million employees) to 12 countries (113 surveys; about 395,000 employees), the results are consistent across samples. The association of long paid leave and the gender earnings gap is robust. It is also robust to the exclusion of any one country as reported in the methods section.

A brief comment on control variables: We do not find a robust association between

gender egalitarianism or public sector size and the motherhood or gender earnings gaps. We find that women's part-time employment (as a percentage of women's employment) is associated with larger gaps. We also find that gender, but not motherhood, gaps are smaller, on average, over time. We do not attempt to disentangle whether this trend reflects period or cohort effects, as we expect both to vary across countries due to differences in national context and policy environments. These temporal dynamics are likely captured indirectly through the observed changes in average gender and motherhood gaps as policies evolve.

[TABLE 3 ABOUT HERE]

Table 4 replicates Model 1b (motherhood penalty with Heckman correction) and 2b (gender earnings gap with correction) but uses estimated marginal effects by education and parenthood. Even when examining effects by education, results remain null for the motherhood penalty. There is no evidence that measured work-family policies explain variation in the motherhood penalty.

Results for the gender earnings gap, however, continue to show that long paid leave is associated with a larger gender gap among all groups, except for the non-tertiary-educated childless (17.8 percent of employed women). We used seemingly unrelated regression (SUR) with postestimation Wald tests to examine differences across groups. The effect of long paid leave on the gender earnings gap is larger in all groups compared to the non-tertiary childless. And the effect is larger for tertiary-educated parents than non-tertiary-educated parents. At +1SD (an additional 10.5 months) of long parental leave, the predicted gender gap is 2.75 percentage points larger than the when long parental leave is at the mean (7.5 months), this estimate is 1.65 for the non-tertiary-educated parents. This suggests that long leave is more strongly associated with earnings inequality by gender among parents with higher levels of educational attainment.

Comparing the minimum (0 months) to maximum (43 months) of long leave amounts to an 11 percentage point difference in the estimated earnings gap among tertiary-educated parents. We note that ECEC is associated with closing the gender earnings gap for non-tertiary-educated parents. Although evidence is suggestive it is not robust across subsamples.

Given null findings for the motherhood penalty and other work-family policies on the gender earnings gap, we proceed to examine the relationship between long paid leave and the gender earnings gap. All analyses were repeated for the motherhood penalty and fatherhood bonus and continued to be null.

[TABLE 4 ABOUT HERE]

(2) Do commonly hypothesized micro-level mechanisms (characteristics of workers and work) explain observed relationships between long paid leave and the size of the gender earnings gap?

The next step of our analysis is to test how the addition of each work variable affects the estimates for long paid leave (over 6 months) reported in Table 4. Figure 2 summarizes the coefficients for long paid leave on the gender earnings percentile gap by education and parenthood across 12 models. These models examine the effects of paid leave over 6 months before and after the inclusion of each variable: (1) full-time status, (2) work hours, (3) occupation, (4) public sector employment, (5) work experience, and (6) a final model that includes full-time status and occupation. For example, the first coefficient $-.30$ (Full-Time) shown in panel A (tertiary-educated parents) is the same model shown in Table 4 for the gender earnings gap among tertiary-educated parents but using the full-time status subsample. The second coefficient $-.27$ (Full-Time+) shows the effect of paid leave over 6 months on the gender gap *after* controlling for full-time employment status at the individual-level. Full models are shown in Tables 2-5 of the Online Supplement.

Although the inclusion of additional variables improves model fit and reduces the size of the gender earnings gap, often quite substantially (see Appendix Table 2 for estimated gaps), the additional variables do not account for the observed relationship between paid parental leave over 6 months and marginal effect of gender (female) on annual earnings percentile. Overall, the inclusion of additional variables *increases* the negative association between long paid leave and earnings percentile among non-tertiary educated mothers (panel B). In the final model controlling for full-time status and occupation the mean gender earnings gap for this group is reduced by two-fifths (not shown), but the negative association between paid leave and earnings increases by half from -.21 to -.33. Differences between women and men in full-time status and occupation were partially suppressing the negative association between long leave and earnings for this group. Results remain largely null for non-tertiary-educated childless (panel D).

Among tertiary-educated parents (panel A) and the tertiary-educated childless (panel C), some of the association is explained by control variables with the largest declines generated by full-time status and occupation. For example, in the final model controlling for both full-time status and occupation the association is reduced by about 20 percent for tertiary-educated parents (from -.29 to -.23) and one-third for the tertiary-educated childless (from -.26 to -.17).

Interestingly, among the tertiary-educated work hours do not explain any of the association between leave and the gender earnings gap, whereas full-time status does. This suggests that differences in contract type may be more important for understanding this relationship than work hours. The importance of contract type is consistent with research finding that leave expansion in India led to downgrading in contract type among women (Bose and Chatterjee 2024).

In the final model, the negative association between paid leave and the gender earnings gap is higher among non-tertiary-educated parents than the tertiary-educated, unlike the results

presented in Table 4. Educational differences in the association between leave and the gender earnings gap among parents are sensitive to control variables and thus are inconclusive.

[FIGURE 2 ABOUT HERE]

As noted above, we performed a variety of specification checks including to our EDV approach, measurement of parental leave, and using number of children to examine the motherhood penalty. We also tested the robustness of our results to outliers with the omission of each country, and to using logged earnings as the dependent variable. Lastly, we tested paternity leave weeks with null results. All test results are in the replication package under Tables.

DISCUSSION

We begin our discussion by returning to our two questions: (1) Do work-family policies explain variation in the size of the gender earnings gap and/or the motherhood penalty and does this vary by educational attainment? Contrary to our expectations, measured work-family policies do not explain variation in the size of the motherhood penalty. This is consistent with recent findings by Brady, Blome, and Kmec (2020), but inconsistent with several studies finding that ECEC and moderate amounts of paid leave narrow the motherhood wage gap, whereas long leave widens it (Budig, Misra and Boeckmann 2016, Cukrowska-Torzewska 2017, Halldén, Levanon and Kricheli-Katz 2015, Misra, Budig and Boeckmann 2011). Our methods are most similar to the former, so we conducted several sensitivity checks to interrogate our findings (e.g., per child penalties among women, log earnings instead of earnings percentile). We were unable to find support for policy effects on the motherhood penalty. One advancement in our design is that we not only consider variation between countries, but within countries over time, which contributes 20 percent of the variation in our model of the motherhood penalty. Additionally, our data spans 1999-2019, covering a more recent period than previous analyses.

We do find that longer paid leave (> 6 months) is associated with larger gender earnings gaps among people of childbearing age (25–44), indicating that extended leave policies may exacerbate inequality for all women in this age group. This is partly mechanistic because most employees ages 25–44 are parents, so the gender earnings gap is larger than the motherhood penalty, including penalties for mothers and/or bonuses for fathers. Yet, there is a gender earnings gap among the childless that is larger than the motherhood penalty and roughly half the size, on average, of the gap among parents regardless of education level. And we find that long paid leave affects the tertiary-educated childless as well. Our results align with prior research showing substantial gender wage gaps among childless workers (Cha, Weeden and Schnabel 2023 in the US, Cooke, Hägglund and Icardi 2022 in Finland, Germany, and the UK), and that the length of paid leave is associated with larger gender wage gaps (Christofides, Polycarpou and Vrachimis 2013), even among the childless (Cukrowska-Torzewska and Lovasz 2020). By distinguishing between the motherhood penalty and the broader gender earnings gap, our analyses further underscore that long paid leave does not exacerbate the motherhood penalty because its effect is dispersed among women, not just mothers.

Regarding differences by skill-level, our findings suggest a more complex picture than the welfare state paradox (Mandel and Semyonov 2005, Mandel and Semyonov 2006, Mandel 2012). We do not find support for larger policy-linked gender earnings gaps among higher skilled parents. Longer paid leave is consistently associated with larger gender earnings gaps between mothers and fathers of *both* educational levels. We find education-based gaps in policy effects among parents are sensitive to the selection of model controls. The effect of longer paid leave on the gender gap is larger for tertiary-educated mothers than non-tertiary educated mothers in some models, but this is not robust to the inclusion of controls for characteristics of

workers and work. This is contrary to the hypothesis that larger policy effects for the higher educated would persist after controlling for micro-economic mechanisms. This may explain inconsistencies in educational gaps across previous studies (as noted in Hook and Li 2020) and has implications for future research.

We do, however, find support for larger policy-linked gender earnings gaps among the higher skilled childless, consistent with aspects of the welfare state paradox. Longer paid leave is consistently associated with larger gender earnings gaps between tertiary-educated childless women and men, but not among the non-tertiary educated childless. Why would we find support for the welfare state paradox among the childless, but not among parents? We posit that unlike tertiary-educated women, women working in lower-skilled jobs are seen as more replaceable, thus the possibility of taking leave sometime in the future may be seen as less of a “threat” to employers.

We find no evidence that work-family policies that promote women's inclusion into the labor market -- ECEC and short paid leave (< 6 months) -- have negative or positive associations with earnings for any group of women. This is contrary to our hypotheses, but consistent with some previous research (Brady et al., 2020). These null findings are important given the growing evidence that ECEC supports mothers' employment, particularly among lower educated mothers (Hook and Paek 2020, Scherer and Pavolini 2023). That is, ECEC includes lower educated mothers into employment, without exacerbating their earnings disadvantage. Another potential explanation for null findings for ECEC is our measure of spending, which combines various factors that may affect usage and ultimately earnings. For example, opening hours of childcare facilities per week – as opposed to spending or enrollment - may be particularly consequential for earnings allowing mothers to compete on more equal footing with men and childless women.

As new policy data becomes available, it opens important opportunities for further research.

Another limitation, shared with other studies using cross-sectional labor market or income surveys, is that we do not have fertility histories, so we limit our analysis to ages 25 to 44 to identify parents by presence of household children. If earnings disadvantages associated with motherhood accumulate or attenuate over the life course, our findings may not capture later-life dynamics. Furthermore, our results provide an average effect of motherhood by education. Although results are robust to using the number of children, detailed data on birth histories would allow us to consider differences in timing and parity (see Doren 2019 for differences in the motherhood penalty by timing and parity in the US).

(2) Do commonly hypothesized micro-level mechanisms explain observed relationships between work-family policy and the size of the gender earnings gap? Although human capital focused micro-level mechanisms help explain the gender earnings gap, the effect of longer paid leave on the gender earnings gap is not well explained by full-time status, work hours, experience, occupation, or sector. The negative association between long leave and the gender earnings gap actually increases for non-tertiary educated mothers with the inclusion of these characteristics. This is inconsistent with the argument that discrimination is greatest for women with high human capital (Mandel 2012), and overlaps with Cooke and colleagues' (2022) argument that lower-skilled women are likely to be seen by employers as the least reliable absent work-family policy supports. Our finding, however, suggests that long leave functions as a work-family policy that *heightens* employers' concerns about the lower-skilled mothers' reliability. For the tertiary-educated, these characteristics explain about one-fifth of the gender earnings gap for parents and one-third for the childless. This is consistent with the hypothesis that micro-economic explanations may be more helpful for explaining policy effects among the highly

educated given their greater returns to human capital. Yet, associations remain substantively large and statistically significant.

We were surprised, however, that controlling for occupation did not account for the negative effect of long paid leave on mothers' earnings percentile. Recent research finds that long paid leave increases occupational sex segregation, especially for mothers (Hook et al 2023). One limitation of our study is that we are only able to measure occupation at the 1-digit level, which may obscure some of the effects of occupational segregation on earnings. Resolving this will require high-quality data on earnings and detailed occupation codes, currently unavailable. We acknowledge that measurement issues likely downwardly bias the explanatory power of these characteristics, yet we note that these characteristics do more to reduce the size of the gender gap than to reduce the association of long leave with the gender gap.

It is also important to note that we do not know if mothers take leave, only that the policy exists. Effects for childless women and persistent effects when controlling for selection and the characteristics of workers and work hours suggests that normative expectations set by leave policy (Gangl and Ziefle 2015) may be as or more important for gender earnings gaps than actual uptake by individual mothers. Yet, recent research suggests that at the individual-level, use of part-time, versus full-time, leave can lessen the motherhood penalty for individual women (Domínguez-Folgueras, González and Lapuerta 2022). Although a recent field experiment in Sweden found that fictitious applicants signaling intentions to take part-time parental leave compared to no mention of leave) were half as likely to receive job interview invitations (Ahmed et al. 2025). It is unclear how flexibility of leave may affect discrimination mechanisms; this is a fruitful area for research. Another avenue for research concerns the length of job guarantees. Job-guarantee rights increase tenure with pre-leave employers (Mari and Cutuli 2021), and paid

leave that extends beyond job-guarantee rights negatively impacts mothers' employment and earnings (Schönberg and Ludsteck 2014). As Mari and Cutuli (2021) suggest, more attention should be paid to the “nuts and bolts” of parental leave schemes that minimize gender inequality.

We pay careful attention to modeling the length of paid leave, improving upon previous work that has relied on a quadratic model. Although we use splines to identify the point at which leave becomes “too long,” there is nothing magical about 6 months, with our result indicating substantively similar results with a knot at 9 months. With a standard deviation of over 10 months, +/- 3 months of leave is a relatively small increment. Thus, any negative effects of moderate length leave in the 6 to 12-month range are estimated to be quite small. We believe the spline is largely distinguishing shorter well-paid periods of leave (including maternity leave) from longer less well-paid periods of care leave. Across 28 European countries in the 2010s, median well-paid leave (≥ 60 percent replacement rates) was 4.5 months with a mean of 7.5 months (Cukrowska-Torzewska, 2017). Given our findings, we anticipate that the length of leave that employers expect mothers to take, not the replacement rate, is particularly relevant.

In sum, although some scholars hypothesize that work-family policies exacerbate the gender earnings gap, especially for mothers and/or tertiary-educated women, we find that long paid leave specifically, as opposed to work-family policies more generally, cleave the labor market outcomes of women from men. Furthermore, the negative relationship between leave length and the gender earnings gap extends to tertiary-educated childless women and is not well-explained by hypothesized individual-level mechanisms measuring the characteristics of workers and work. Thus, we argue that findings are consistent with generalized discrimination against mothers and tertiary-educated women of child-bearing age more broadly as suggested by theory and recent empirical findings (Glass and Fodor 2011, Hutt Cabello 2023). We conclude by

emphasizing that inclusive policies that support mothers' employment, here ECEC and moderate length paid leave, do so without exacerbating the motherhood penalty or the gender earnings gap.

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Table 1. Individual-level Descriptives, LIS Data, Ages 25-44

	Women				Men			
	Mean	SD	Min	Max	Mean	SD	Min	Max
<i>Employee Sample (N = 2,957,461)</i>								
Earnings percentile	42.6	(27.8)	34.6	48.8	57.3	(27.8)	51.4	64.8
Tertiary degree (%)	46.8		13.0	76.6	38.9		9.8	68.5
Parent (living with children, %)	60.2		39.2	81.2	51.2		35.5	71.3
Partnered (%)	54.5		34.8	80.0	55.4		27.0	80.2
Age	34.8	(5.8)	32.4	36.2	34.6	(5.7)	32.2	36.2
Tertiary x Parent								
Non-Tertiary Parent	35.4		17.3	67.6	32.2		16.8	62.0
Non-Tertiary Childless	17.8		6.1	39.2	28.8		13.6	52.8
Tertiary Parent	24.8		7.0	47.1	18.9		3.4	38.1
Tertiary Childless	22.0		4.8	38.4	20.0		3.9	34.9
<i>Selection Equation (N = 4,213,021)</i>								
Employed (%)	71.7		52.0	85.8	86.4		67.3	96.1
Child <= 5 years (%)	32.6		19.0	49.0	28.3		15.5	40.4
Transfer income (log)	6.3	(3.8)	2.4	12.7	5.9	(4.0)	2.9	11.7
Other household income (log)	2.5	(3.8)	0.6	4.4	1.0	(2.3)	-0.5	2.1
Education (%)								
Low	13.0		3.3	43.8	14.7		3.0	50.2
Medium	45.7		17.7	80.5	49.4		19.4	81.7
High	41.2		11.2	69.0	36.0		11.6	62.9
	Women		Men		Sample N			
<i>Additional Variables & Samples</i>	Mean	SD	Mean	SD	Country	Survey	Employees	
Full-time (%)	68.1		86.7		23	236	2,070,692	
Work hours	35.9	(10.9)	42.4	(9.2)	21	237	1,442,025	
Work experience	12.0	(6.5)	13.2	(6.8)	12	238	395,425	
Work in public sector (%)	23.5		15.7		15	239	1,868,820	
Occupation (10 categories, %)					22	240	1,741,226	
Managers	9.1		11.0					
Professionals	24.8		18.1					
Technicians/assoc professionals	19.0		13.7					
Clerical support workers	16.9		5.9					
Services & sales workers	18.1		10.2					
Agricultural, forestry, & fishing	0.3		1.1					
Craft & related trades workers	2.3		19.0					
Plant and machine operators	2.8		10.8					
Elementary occupations	6.5		8.8					
Armed forces occupations	0.2		1.4					

Notes: Min and Max refer to the minimum and maximum survey-level means (N = 280).

Table 2. Survey-level Descriptives, N= 280

	Unadjusted Mean	Adjusted (Heckman Correction)			
		Mean	SD	Min	Max
Estimates as Dependent Variables					
Gender Earnings Gap	-19.1	-17.3	4.2	-27.7	-7.0
Motherhood Penalty	-11.0	-7.3	5.4	-20.6	6.1
<i>Gender Earnings Gap by Group</i>					
Tertiary-Educated Parents	-23.1	-19.3	5.9	-36.0	-4.8
Non-Tertiary-Educated Parents	-25.5	-22.6	5.9	-36.4	-6.2
Tertiary-Educated Childless	-9.3	-8.8	4.1	-18.1	4.8
Non-Tertiary-Educated Childless	-11.6	-11.6	3.8	-21.9	0.0
<i>Motherhood Penalty by Group</i>					
Tertiary-Educated Women	-9.3	-5.7	5.9	-21.4	11.9
Non-Tertiary-Educated Women	-11.9	-8.1	5.7	-22.1	4.9
<i>Fatherhood Premium by Group</i>					
Tertiary-Educated Men	4.6	4.9	3.6	-6.5	15.5
Non-Tertiary-Educated Men	2.0	2.9	2.9	-5.8	10.6
Independent Variables					
ECEC Expenditure (\$1,000s)		4.3	2.4	0.1	11.7
Months Paid Leave		12.5	11.3	0.0	49.4
# Months ≤ 6		5.0	1.8	0.0	6.0
# Months > 6		7.5	10.5	0.0	43.4
Gender Egalitarianism (scale)		2.6	0.3	2.0	3.4
Women's Part-Time (%)		32.0	15.9	3.8	75.4
Public Sector Size (%)		17.3	5.7	2.4	30.3
Year		2009.6	5.5	1999	2019

Note: Gender Earnings Gap refers to the marginal effect of the variable “female” on annual earnings percentile as estimated from 280 OLS (unadjusted) and 280 Heckman selection (adjusted) models using LIS data on the full sample. Motherhood Penalty refers to the marginal effect of “parent” when gender equals female. The unadjusted model includes female, parent, and their interaction with controls for tertiary degree, partnered interacted with female, age and its square. The adjusted model adds the Heckman selection equation described in text.

Table 3. EDV Regression Predicting Country-Year Motherhood and Gender Earnings Gaps (Annual Earnings Percentile)

	<u>Motherhood Penalty</u>		<u>Gender Earnings Gap</u>						
	(1a)	(1b)	(2a)	(2b)	(3)	(4)	(5)	(6)	(7)
	Full Sample OLS	Heckman	Full Sample OLS	Heckman	Full-Time Sample	Work Hours Sample	Occupation Sample	Public Sector Sample	Experience Sample
ECEC Expenditure (in \$1,000s)	.30 (.74)	.36 (.65)	.32 (.29)	.36 (.25)	.23 (.24)	.23 (.25)	.31 (.30)	.74 ** (.22)	.11 (.50)
Months Paid Leave ≤ 6 Months	-.65 (.56)	-.57 (.49)	-.04 (.36)	-.18 (.32)	-.06 (.29)	.01 (.32)	-.15 (.38)	-.42 (.35)	.59 (.63)
Months Paid Leave > 6 Months	-.13 (.09)	-.06 (.08)	-.18 * (.07)	-.16 * (.06)	-.23 *** (.05)	-.24 *** (.05)	-.17 * (.07)	-.15 * (.06)	-.28 ** (.07)
Gender Egalitarianism (scale)	2.95 (4.28)	1.81 (3.64)	1.56 (2.48)	.38 (2.22)	-.32 (1.79)	.11 (1.75)	1.40 (2.66)	1.03 (2.99)	-.24 (2.48)
Women's Part-Time (as % of Total)	-.34 ** (.10)	-.26 *** (.07)	-.23 *** (.04)	-.22 *** (.04)	-.23 *** (.03)	-.23 *** (.03)	-.22 *** (.04)	-.21 * (.08)	-.23 ** (.06)
Public Sector Size (as % of Total)	.01 (.22)	.01 (.19)	-.18 (.10)	-.16 (.09)	-.16 (.08)	-.17 * (.08)	-.17 (.13)	-.28 * (.13)	-.19 (.09)
Year	.01 (.10)	.05 (.09)	.19 * (.08)	.20 * (.07)	.32 *** (.05)	.30 *** (.06)	.19 (.11)	.16 (.13)	.30 ** (.09)
Constant	-11.08 *** (1.76)	-7.82 *** (1.56)	-21.03 *** (.73)	-19.40 *** (.70)	-20.56 *** (.54)	-20.51 *** (.68)	-19.30 *** (1.13)	-18.94 *** (1.32)	-20.83 *** (.87)
N - country-year	280	280	280	280	236	217	220	159	113
N - country	26	26	26	26	23	21	22	15	12
R ²	0.52	0.47	0.53	0.55	0.66	0.67	0.53	0.62	0.60

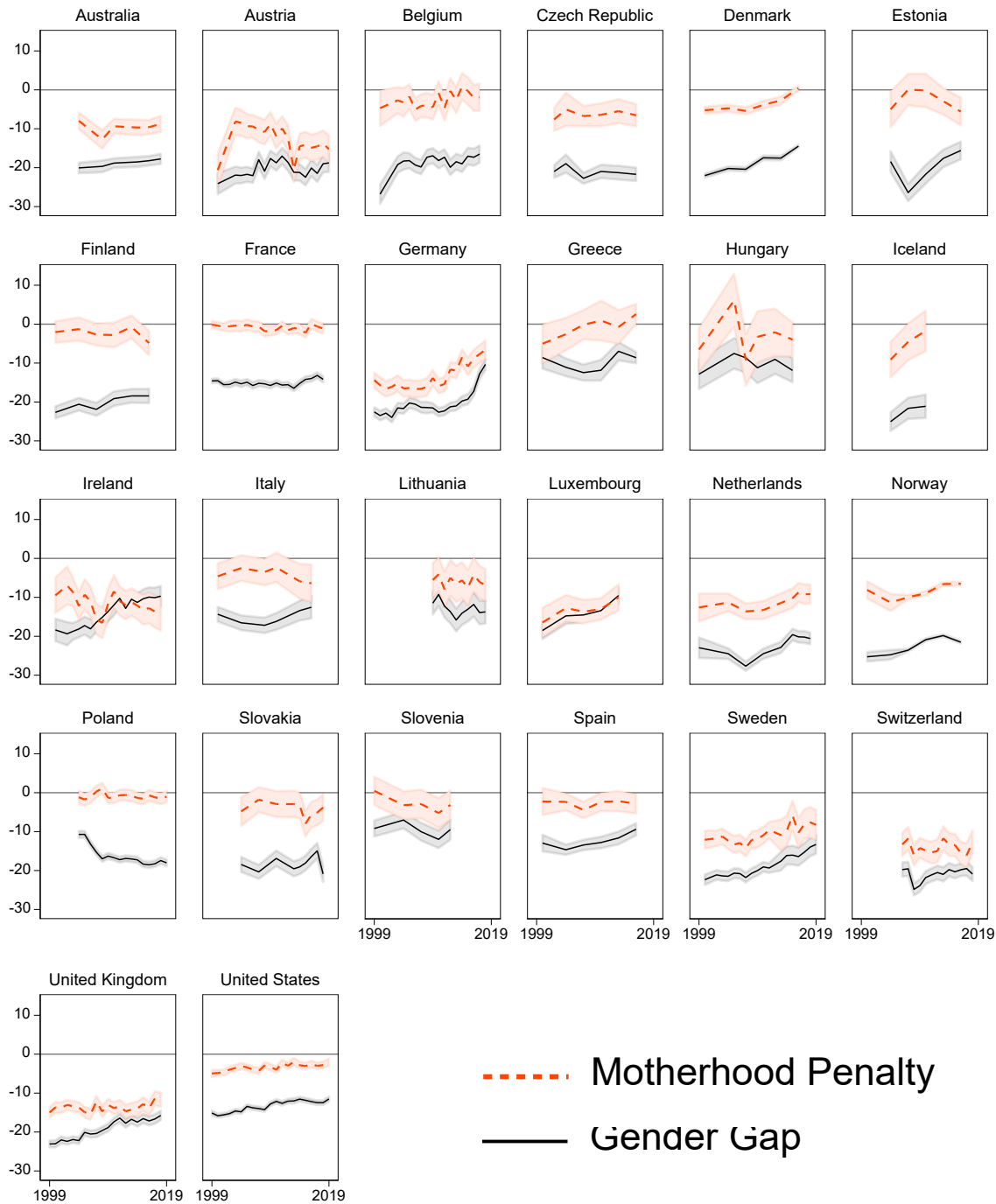
Notes: *P < .05; **P < .01; ***P < .001 (two-tailed tests). Models 3-7 are Heckman corrected. EDV = estimated dependent variable.

Table 4. EDV Regression Predicting Country-Year Motherhood and Gender Earnings Gaps (Annual Earnings Percentile, Heckman Adjusted), by Education and Parenthood

	<u>Motherhood Penalty</u>		<u>Gender Earnings Gap</u>			
	Tertiary Women	Non-Tertiary Women	Tertiary Parents	Non-Tertiary Parents	Tertiary Childless	Non-Tertiary Childless
ECEC Expenditure (in \$1,000s)	.45 (.77)	.32 (.60)	.23 (.40)	.69 * (.27)	-.37 (.45)	.21 (.38)
Months Paid Leave ≤ 6 Months	-.13 (.62)	-.82 (.47)	-.08 (.26)	-.44 (.40)	.31 (.45)	.03 (.51)
Months Paid Leave > 6 Months	-.03 (.08)	-.06 (.08)	-.26 *** (.06)	-.16 * (.06)	-.18 * (.08)	-.06 (.08)
Gender Egalitarianism (scale)	1.30 (4.23)	1.18 (3.41)	2.30 (2.27)	-.92 (2.18)	3.08 (3.91)	-1.31 (3.23)
Women's Part-Time (as % of Total)	-.25 ** (.09)	-.27 *** (.06)	-.34 *** (.05)	-.35 *** (.04)	-.05 (.05)	-.01 (.04)
Public Sector Size (as % of Total)	-.24 (.25)	.16 (.16)	-.11 (.12)	-.17 (.12)	-.03 (.14)	-.23 (.13)
Year	-.01 (.11)	.06 (.09)	.16 (.10)	.14 (.10)	.11 (.09)	.10 (.08)
Constant	-5.50 ** (1.85)	-8.71 *** (1.52)	-21.03 *** (1.20)	-24.06 *** (.97)	-9.97 *** (1.11)	-12.64 *** (1.00)
N - country-year	280	280	280	280	280	280
N - country	26	26	26	26	26	26
R ²	0.29	0.56	0.61	0.61	0.21	0.17

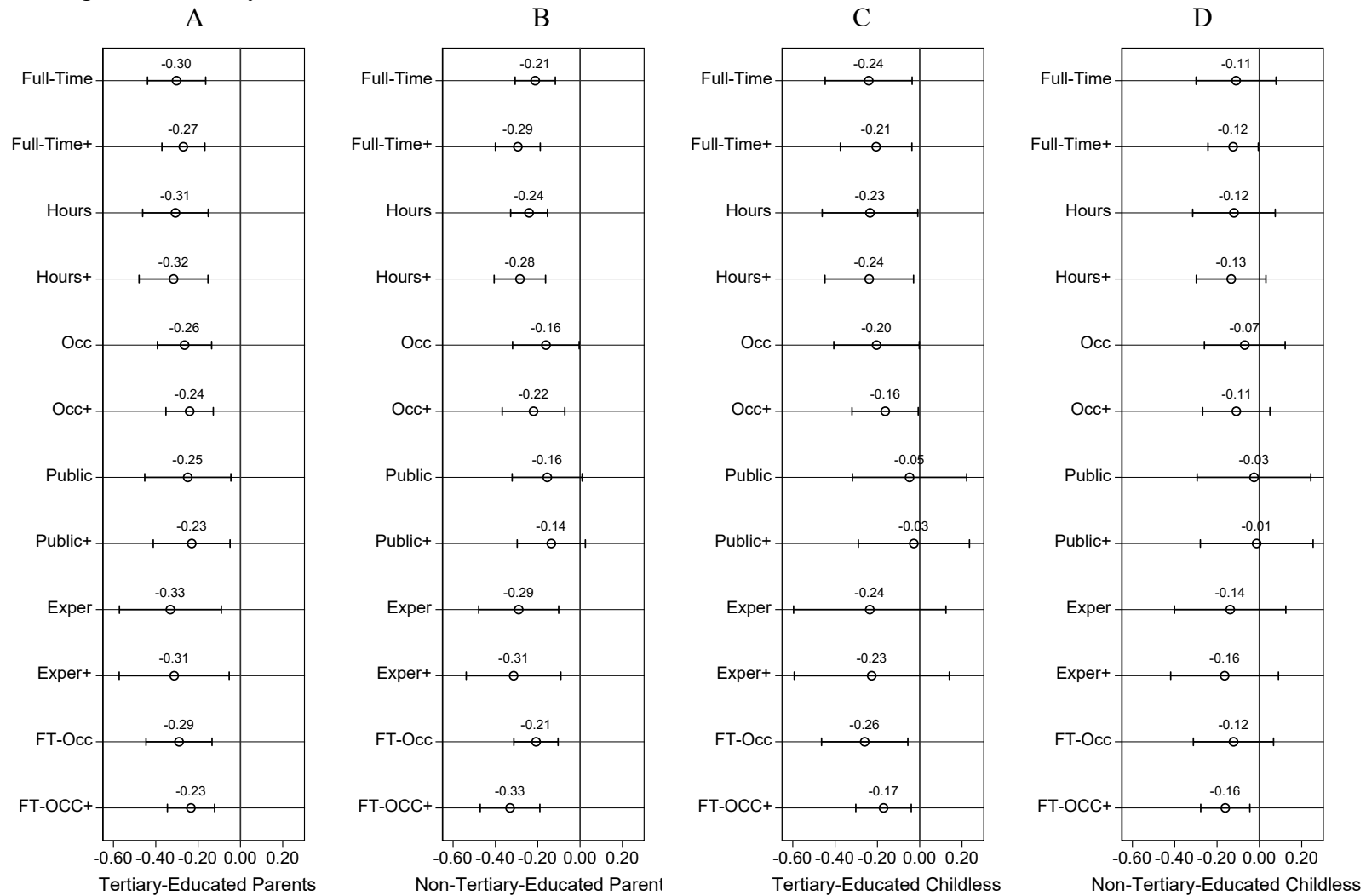
Notes: *P < .05; **P < .01; ***P < .001 (two-tailed tests). Full sample from Table 3 (model 1b and 2b). EDV = estimated dependent variable.

Figure 1. Trends in Earnings Percentile Gaps by Country, 1999-2019, Heckman-Adjusted



Notes: Heckman corrected gaps as shown in Table 2. 95% confidence intervals calculated from the standard errors of the marginal effects. Gender Earnings Gap refers to the marginal effect of the variable “female” on annual earnings percentile as estimated from 280 Heckman selection models using LIS data on the full sample. Motherhood Penalty refers to the marginal effect of “parent” when gender is set to female. The correlation between the Gender Earnings Gap and the Motherhood Penalty is .49.

Figure 2. Coefficient Plots: Effect of Months of Paid Leave > 6 Months on the Marginal Effect of Gender (Female) on Annual Earnings Percentile, by Education and Parental Status



Notes: Model name indicates sub-sample. “+” indicates the addition of the micro-level control variable for which the sub-sample is named. Sample sizes are listed in Table 1. FT-OCC is a sub-sample containing both full-time status and occupation (N= 20/197). See Online Supplement Tables 3-6 for full results.

Appendix Table 1. Country-level Descriptives

	Year		Obs. (#)	N		Min N Group	# Children		ECEC		Paid Leave		GenderNorms		Part-Time		PublicSector	
	First	Last		Mean	Min		Min	Max	Min	Max	Min	Max	Min	Max	Min	Max	Min	Max
Australia	2004	2018	6	6,522	4,623	379	1.9	2.0	2.5	4.9	0.0	4.2	2.9	3.0	44.5	46.8	13.3	13.7
Austria	2000	2019	18	2,449	1,324	47	1.7	1.9	1.6	4.9	13.8	31.8	2.0	2.5	32.4	47.7	10.2	17.2
Belgium	2000	2017	16	2,538	1,431	69	1.8	1.9	3.0	6.0	6.5	7.5	2.5	3.0	37.4	43.6	16.9	18.8
Czechia	2002	2016	6	3,738	1,948	53	1.7	1.8	1.4	2.5	25.4	49.4	2.5	2.7	8.5	10.4	12.9	16.4
Denmark	2000	2016	6	54,136	31,085	1877	1.9	1.9	8.7	11.3	11.5	12.5	3.1	3.4	32.7	37.5	23.9	29.4
Estonia	2004	2016	5	2,330	1,901	59	1.7	1.8	0.7	3.0	38.3	38.3	2.3	2.9	11.5	15.3	14.9	23.7
Finland	2000	2016	6	3,568	2,986	167	2.0	2.0	4.5	6.6	36.7	37.2	2.6	3.1	16.8	20.1	21.1	25.0
France	1999	2018	20	21,635	15,221	900	1.8	1.9	6.2	8.7	3.7	9.7	2.4	3.0	29.5	31.7	21.7	22.6
Germany	1999	2018	20	6,233	4,030	276	1.7	1.8	2.5	6.5	13.4	25.2	2.1	2.8	36.4	47.4	10.4	11.7
Greece	2000	2016	6	3,356	1,417	65	1.7	1.8	0.1	1.7	3.7	9.9	2.0	2.1	7.7	13.2	2.4	17.7
Hungary	1999	2015	6	904	724	24	1.6	1.8	1.6	3.0	36.9	36.9	2.2	2.5	5.5	9.1	19.0	20.4
Iceland	2004	2010	3	1,579	1,395	52	2.0	2.0	5.3	9.0	6.0	6.0	2.9	3.0	37.0	37.5	25.0	25.0
Ireland	2000	2018	18	1,741	1,179	56	1.8	2.1	1.4	3.9	3.2	6.0	2.6	2.9	30.6	35.7	14.6	15.8
Italy	2000	2016	6	2,716	1,551	45	1.6	1.7	3.4	4.2	11.0	11.0	2.0	2.3	15.6	32.5	13.7	14.5
Lithuania	2009	2018	10	1,516	1,369	64	1.6	1.7	2.6	3.5	14.3	14.3	2.2	2.4	8.7	11.3	22.2	24.8
Luxembourg	2000	2013	5	2,494	1,698	81	1.6	1.8	5.3	10.9	9.7	9.7	2.1	2.5	24.6	36.3	10.1	11.7
Netherlands	1999	2018	9	4,261	2,421	133	1.8	1.9	1.9	6.4	3.7	9.7	2.5	3.1	67.2	75.4	12.0	13.1
Norway	2000	2016	6	97,029	6,810	365	1.9	2.0	5.3	11.7	21.0	32.8	3.0	3.4	37.4	44.7	28.2	30.1
Poland	2004	2019	16	17,401	14,809	526	1.7	1.8	0.6	2.4	3.7	12.0	2.1	2.5	10.2	14.2	10.7	16.5
Slovakia	2004	2018	9	3,056	2,591	134	1.8	1.9	1.4	2.7	37.8	39.2	2.5	2.8	3.8	8.5	9.4	18.9
Slovenia	1999	2012	5	2,298	1,994	88	1.7	1.8	2.6	3.0	12.0	12.0	2.5	2.9	7.5	13.3	13.1	16.2
Spain	2000	2016	6	5,435	2,635	179	1.6	1.8	2.0	3.2	3.7	3.7	2.6	3.0	16.8	25.2	12.1	15.2
Sweden	2000	2019	18	6,353	2,257	85	1.9	2.0	5.4	9.9	12.9	13.8	2.8	3.4	29.8	36.4	28.3	30.3
Switzerland	2006	2018	13	2,875	2,503	187	1.7	1.9	3.8	7.2	1.8	3.2	2.1	2.8	56.5	59.4	7.1	10.1
UK	1999	2018	20	10,380	7,165	436	1.7	1.9	2.9	6.0	4.2	9.0	2.5	2.9	41.3	44.8	16.0	19.6
US	1999	2019	21	37,367	24,924	1429	2.0	2.0	2.0	2.7	0.0	0.0	2.5	2.8	23.9	26.6	7.7	16.7

Notes: Calculated on the employee sample. # Obs indicates the number of surveys included for each country; Min N and Group N list the minimum number of observations ever observed for each country and for a group with each country (e.g., tertiary-educated mothers), respectively. # Children reports the minimum and maximum mean number of children observed in each country among mothers. ECEC and Paid Leave report the minimum and maximum ECEC spending in \$1,000s and months of paid parental leave.

Appendix Table 2. Gender Earnings Gap in Sub-Samples, Below and After Inclusion of Micro-level Control

Model	N	Before	After	% Change
Full-Time	236	-17.3	-11.2	35.5
Work Hours	217	-17.5	-10.5	40.0
Occupation	220	-17.1	-16.3	4.3
Public Sector	159	-16.9	-16.8	0.9
Experience	113	-18.0	-17.0	5.4
Full-Time & Occupation	197	-17.4	-11.3	35.1

Notes: Model name indicates sub-sample. “After” indicates the addition of the micro-level control variable for which the sub-sample is named. Gender Earnings Gap refers to the marginal effect of the variable “female” on annual earnings percentile as estimated from Heckman selection (adjusted) models using LIS data.