The Constant Gap: Parenthood Premiums in Sweden, 1968–2010

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We know that parenthood has different consequences for men’s and women’s careers. Still, the research remains inconclusive on the question of whether this is mainly a consequence of a fatherhood premium, a motherhood penalty, or both. A common assumption is that women fall behind in terms of pay when they become mothers.

Based on longitudinal data from the Swedish Level of Living Survey (LNU), and individual fixed-effects models, we examine the support for this assumption by mapping the size of parenthood effects on wages during the years 1968–2010. During this period, Swedish women’s labor supply increased dramatically, dual-earner family policies were institutionalized, and society’s norms on the gendered division of labor changed. We describe the development of parenthood effects on wages during this transformative period.

Our results indicate that both genders benefit from a gross parenthood premium, both at the beginning of the period and in recent years, but the size of this premium is larger for men. Individual fixed-effects models indicate that the wage premium is mainly the result of parents’ increased labor market investments. Controlling for these, women suffer from a small motherhood penalty early in the period under study whereas parenthood is unrelated to women’s wages in later years and to men’s wages throughout the period. Neither for men nor for women do we find a statistically significant period change in the parenthood effects. Instead, patterns are remarkably stable over time given the radical changes in family policies and norms that took place during the period examined.

The authors contributed equally to this article and are listed alphabetically. Financial support from the Swedish Research Council for Health, Working Life and Welfare (Forte) (2016–00661) is gratefully acknowledged. A previous version of this paper was presented at the Network Meeting for Research on Social Policy and Welfare, Stockholm, October 2016, the Swedish Institute for Social Research, Stockholm, March 2017 and the European Consortium for Sociological Research conference in Milan, September 2017, where we benefited from comments and suggestions from the participants. Corresponding author: Charlotta Magnusson, Swedish Institute for Social Research (SOFI) Stockholm University, SE-106 91 Stockholm, Tel: +46-8-162607; E-mail: Charlotta.magnusson@sofi.su.se

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Social Forces 100(1) 137–168, September 2021
doi:10.1093/sf/soa087
Advance Access publication on 21 October 2020
Introduction

We know that men, in general, earn more than women. We also know that this gender gap is triggered or accentuated at parenthood (Angelov et al. 2013; Blau and Kahn 2017; Boye et al. 2017). The fatherhood premium, that is, that men who become fathers fare better on the labor market than men who remain childless (Koslowski 2011; Bygren and Gähler 2012), and/or the motherhood penalty, that is, that women who become mothers fare worse on the labor market than women who remain childless (Gash 2009; Budig et al. 2016), are both potential sources of this inequality. Some comparative research has found none or only small motherhood wage and earnings penalties in the Nordic countries and has applauded the generous family policies, mainly introduced in the 1970s–1990s, for this phenomenon (Gash 2009; Budig et al. 2016) while other researchers (Hakim 2000; Mandel and Semyonov 2006; Datta Gupta et al. 2008; Mandel 2012) assert that such policies rather tend to increase gender inequality.

Since the 1960s, there has been a shift from male breadwinners to dual-earner families in many Western countries. The gradual enrolment of women into paid work following the housewife era has had consequences for the composition of the workforce—not least when it comes to mothers’ labor market participation. These changes should have influenced the gendered association between parenthood and labor market outcomes. However, with a few exceptions (Bygren and Gähler 2012; Petersen et al. 2014; Pal and Waldfogel 2016; Glauber 2018; Jee et al. 2019), there are very few studies describing the development of parenthood effects on the careers of men and women during this transformative period. Thus, we largely lack knowledge of whether the changes in the labor market and in family policies over recent decades have increased, decreased or been neutral in relation to parenthood penalties and premiums in terms of wages.

We contribute to filling this research void by mapping out the size of parenthood effects on wages, and determinants of these effects, during the period 1968–2010, using data from six nationally representative waves of the Swedish Level of Living Survey (with the Swedish acronym LNU, see https://www.sofi.su.se/english/2.17851/research/three-research-units/lnu-level-of-living). We thus cover a longer period, and are able to show more recent results for a Nordic country, than the previous study from Norway (1979–1996) by Petersen et al. (2014). In addition to describing developments over time, the panel data structure, that is, that the same respondents are observed on several occasions, gives us an opportunity to analyze wage changes associated with family formation. Since we are then able to control for individual time-invariant determinants of wages, we can judge whether parenthood effects are generated by parenthood as such, or selection by such determinants.

Background

Mechanisms Behind Gendered Career Effects of Parenthood

Explanations for gendered career effects of parenthood may analytically be divided into (1) “treatment” effects, implying that there is a causal effect of
parenthood on career outcomes, and (2) selection effects, implying that there is no causal effect of parenthood on career outcomes, but nonrandom selection processes which generate spurious correlations between parenthood and career outcomes (see Petersen et al. 2014). Actors involved in generating the treatment effects may be subdivided into those located on (1.1) the employee side or (1.2) the employer side of the labor market.

**Treatment Effects Part 1: The Employee Side**

Human capital theory explains women’s lower wages by reference to gender differences in individual characteristics associated with productivity, such as level of education, labor market experience, on-the-job training and other aspects of productivity that affect earnings (Becker 1991). During the 1960s, 1970s and 1980s there were quite large gender differences in human capital. Since then, women have improved their human capital and today their level of education is equal to or even exceeds that of men (Blau and Kahn 2017 [United States]; Bygren and Gähler 2012 [Sweden]; Goldin 2006 [United States]). Still, there is a large amount of evidence indicating strongly gender-specific ways of reconciling family and career-related demands that affect gender labor supply differences among parents. One consequence of the gender division of labor is that women (mothers) tend to work part-time to a larger extent than men, which results in less work experience for women. Because the prospects of reaching well-paid positions in the labor market are conditioned by the individual’s human capital, such differences usually explain some of the gender gap in wages. However, studies consistently reveal a remaining and pronounced wage gap between women and men even after controlling for various indicators of human capital, across countries (Arulampalam et al. 2007; Blau and Kahn 2017; Boye et al. 2017). One mechanism behind this residual gap might be gender differences in more hard-to-observe employer-specific human capital investments, caused by career interruptions because of childrearing.

The mother-friendly job hypothesis postulates that women, and mothers in particular, choose jobs that are easy to reconcile with family responsibilities, that is, jobs with flexible work arrangements. The idea rests on a compensating principle: female-dominated occupations are assumed to have more flexible working arrangements that facilitate work-family balance, but are not well paid. Women (or mothers) will choose to work in female-dominated occupations to facilitate everyday family life (Becker 1991). The empirical support for this hypothesis seems to be rather weak however (England 2005; Magnusson 2019).

As dual-earner couples have become increasingly common, the traditional male breadwinner model has lost in significance; men and women increasingly share their unpaid household work and some men reduce the time they devote to market work following childbirth (see Bygren et al. 2004). Nevertheless, the male “good-provider model” prevails in the sense that men most commonly continue working full time or even increase their market work hours when they become fathers, whereas women often decrease their labor supply when they become mothers (Kennerberg 2007). The gender difference in time devoted to household
tasks seems to be triggered by marriage/cohabitation and childbearing: It is relatively small for single persons, but grows substantially among childless couples and further increases when couples have children (Boye 2014). There are also indications of gender differences in work commitment, with women’s labor market attachment and commitment decreasing following family formation (Becker 1991; Evertsson 2013; cf. Hakim 2000).

**Treatment Effects Part 2: The Employer Side**

Women’s disadvantages might also be due to structural constraints in the labor market, such as discrimination against women on the employer side (Glauber 2007). According to many theories on the demand side, motherhood is devalued on the labor market while fatherhood is valued (Ridgeway and Correll 2004). Family responsibilities are regarded as intrinsic to motherhood but for fathers, care responsibilities, at least from the perspective of employers and co-workers, are more often understood as optional (Kugelberg 2006). If employers assume or know that women, on average, are more constrained by family responsibilities than men, and/or leave the labor market for longer periods, they may statistically discriminate using this information in their employment and promotion decisions. Accordingly, employers may be less likely to invest in firm-specific on-the-job training for women (Polavieja 2008) and women may be excluded from occupations requiring such training (Estévez-Abe 2005). There are some indications that employers treat employees differently depending on whether they are parents. Parental status has been shown to have a polarizing effect on judgments of individual men and women in evaluations of job competence and decisions on employment and promotion, to the disadvantage of mothers and the advantage of fathers (Correll et al. 2007; Oesch et al. 2017).

Assumptions about fathers/mothers and men/women are also central to the gendered organizations perspective, which asserts that cultural assumptions about men and women tend to influence organizational rewards. Organizations are not bureaucratic gender-neutral institutions; all actions entail a distinction between masculine and feminine. The ideal worker, Acker (1990) stresses, is a disembodied worker, totally committed to work without any obligations outside work, a figure that is thus not consistent with assumptions about motherhood. For instance, long work days, which are hard to reconcile with parenthood, have been stressed as an important factor behind the wage gap between mothers and fathers (Weeden et al. 2016; Magnusson and Nermo 2017). Goldin (2014) asserts that a large part of the gender wage gap is due to firms rewarding individuals who work long days and who are available most of the time.

Gendered cultural assumptions could also affect differences in labor market outcomes between fathers and childless men. Fatherhood signals reliability and commitment—characteristics that are coveted on the labor market and consistent with hegemonic masculinity (Glauber 2008; cf. Collinson and Hearn 2005).
Selection effects

While treatment hypotheses imply that parental premiums or penalties are caused by behavioral changes on the employee and/or employer side, the selection perspective instead asserts that the wage premium for parenthood is caused by stable differences in productivity and other wage-relevant personal characteristics between individuals who become parents and those who do not (Becker 1991; Ludwig and Brüderl 2011; Petersen et al. 2014). This perspective implies that parents (commonly fathers in previous research) have unobserved characteristics that are coveted both in the labor market and in the marriage (or reproduction) market. The hypothesis then is that men who eventually end up having children have permanent traits that set them apart from men who never have children, and that these permanent traits increase their earning power. If true, the apparent effect of fatherhood on wages is then a spurious effect of these permanent traits. Likewise, this perspective explains lower wages for mothers compared with non-mothers by reference to a negative selection of women into parenthood (Lundberg and Rose 2000). The selection effect may also change over time. For instance, prior research shows a shift in the association between marriage and level of education, whereby highly educated women, who were less likely to get married in earlier cohorts, are now the most likely to get married (Isen and Stevenson 2010). Taking selection effects into account, which we attempt to do here, is crucial when analyzing the association between parenthood and wages. To summarize, there are several potential explanations, both on the employer and employee side, and in terms of selection, as to why parenthood should impact differently on women’s and men’s careers.

Previous Findings Regarding Parenthood Penalties and Premiums

For men, most studies, both American and European, show that fatherhood is positively associated with wages and earnings (Angelov et al. 2013 [Sweden]; Boschini et al. 2011 [Sweden]; Cooke 2014 [Australia, Great Britain, and United States]; Glauber 2008 [United States]; Hodges and Budig 2010 [United States]; Magnusson and Nermo 2017 [Sweden]; Petersen et al. 2014 [Norway]; Simonsen and Skipper 2012 [Denmark]; Koslowski 2011 [14 countries]). Wage and earnings penalties associated with parenthood, net of, for example, human capital, work effort, family structure, and job characteristics, have been found among women (Budig and Hodges 2010 [United States]; Budig et al. 2016 [22 countries]; Gangl and Ziefle 2009 [Germany, UK, United States]; Gash 2009 [6 countries]; Gough and Noonan 2013 [United States]; Harkness and Waldfogel 2003 [7 countries]; Joshi et al. 1999 [UK]; Sige-Rushton and Waldfogel 2007 [8 countries]; Simonsen and Skipper 2012 [Denmark]). In Sweden and the other Nordic countries, which have policies that support maternal employment, for example, publicly funded childcare, paid parental (and paternal) leave, and individual taxation, motherhood wage and earnings penalties are small compared to many other countries (Datta Gupta and Smith 2002; Harkness and Waldfogel...
Some studies even show a parenthood wage (Cukrowska-Torzewska 2017) and earnings premium among Swedish women (Sigle-Rushton and Waldfogel 2007; Boschini et al. 2011).

A small number of previous studies, mostly from the United States, have investigated how parenthood effects on labor market outcomes have evolved over time. Pal and Waldfogel (2016) analyzed wage differences between mothers and non-mothers in the United States over five decades and found a large decline in the motherhood wage penalty. In 1967, the overall gap between women with and without children was about 6 percent. In 2013, the gap was down to 1 percent. In an American study, Jee et al. (2019) instead found a rather stable pattern in the motherhood wage penalty during the period 1986–2014 (see also Avellar and Smock 2003). Petersen et al. (2014) used Norwegian register data to analyze how marriage and parenthood were related to men’s and women’s wages during the period 1979–1996. They found that the gender wage gap was primarily due to a motherhood penalty in the 1970s but that this penalty became much smaller in the 1990s. At the same time, however, a fatherhood premium evolved during this period.

**Development Over Time: What to Expect for the Swedish Case**

Sweden consistently ranks at the top in international comparisons of the degree of gender equality (United Nations Development Programme 2016). Swedish women have a higher labor market participation rate than women in most other countries (ibid.). In particular, the enrolment of Swedish mothers in paid labor, in comparison with other countries, is very high (Thévenon 2011). Since the 1970s, gender equality has increased in many domains: employment rates, work hours, wages, occupational skill level, education, economic dependency within couples, household work hours, and the use of parental leave (Bygren et al. 2004; Boye 2014). During this period, Sweden has implemented a number of family-related policies aimed at replacing the male breadwinner model with that of a dual-earner model. Important elements driving this transformation were the introduction of the individual taxation of married couples in 1971, the introduction of parental leave benefits in 1974, the expansion of subsidized public childcare during the 1970s and 1980s, and the introduction of “daddy-quotas” for parental leave in the 1990s (Ferrarini and Duvander 2010).

Based on these policy changes, two conflicting hypotheses can be discerned in the literature. The *attenuation hypothesis* states that with regard to parenthood “treatment” effects, the policy changes referred to above should have produced a push towards gender convergence in supply side behaviors following parenthood, making any motherhood penalties decrease (or vanish) over time. Swedish family policies make motherhood easier to reconcile with job demands, at least compared to other countries, and there is a culture of supporting working mothers in Swedish workplaces (Collins 2019). Although major inequalities remain, Swedish women of today, and mothers in particular, should meet the labor market with more assets and possibilities, relative to men, than was the
case for women and mothers before the implementation of these family policies. Since fathers, partly as a consequence of these policies, have increased their engagement in family responsibilities over time, any fatherhood wage premiums are also expected to be decreasing over time. For instance, the fathers’ share of used parental leave days increased from approximately 0.5 percent of all days in 1974 to approximately 23 percent in 2010 (Duvander and Johansson 2012). According to this hypothesis, then, we would expect possible parenthood effects among men and women to converge over time.

However, the shift to dual-earner families, in which both partners may face problems in combining paid and unpaid work, may lead to new challenges for women (and men) in the labor market. The extensive parental leave program in Sweden has been identified as a potential cause of gender inequalities in the labor market (see, Wright et al. 1995; Rosenfeld et al. 1998; Hakim 2000; Mandel and Semyonov 2006). Although parental leave benefits can be used by any parent, in reality it is the mothers who make the most extensive use of them (Duvander and Viklund 2014). Moreover, mothers with small children, 3–5 years, are much more likely than fathers to work part-time (40 vs. 8 percent) (Statistics Sweden 2018). Even though the long work day culture is less established in Sweden than in the United States, Swedish women are less likely to work overtime, go on business trips, or to be required to be constantly available (Magnusson and Nermo 2017, see also Jacobs and Gornick 2002). In practice, then, a large proportion of Swedish mothers have a relatively weak connection to the labor market in the period following childbirth as a result of rather long periods of absence and/or part-time work. This may give rise to statistical discrimination against female employees of fertile age. The exacerbation hypothesis states that the development of Swedish family policies has led to increased motherhood penalties, especially among women in higher positions (Mandel and Semyonov 2006). Family-friendly policies involving long parental leave and the possibility to reduce working hours and to stay home with sick children are expected to exacerbate gender differences in labor supply choices as mothers make use of these possibilities to a larger extent than fathers.

Finally, structural labor market changes witnessed over recent decades may have had consequences for the composition of the workforce, and in particular the female workforce. The rise in female labor force participation suggests that the size of the group of working women with relatively weak work orientation has increased. Thus, the positive selection of women into paid labor during the 1960s and 1970s, when the female participation rate was low, and when it was less common for mothers to engage in paid work, has most probably weakened over time, which all else being equal would lead to lower relative wages for mothers over time. On the other hand, in more recent decades the labor market has also undergone a shift from low-skill to high-skill jobs (Tåhlin 2014). On the supply side, the average educational level has increased quite substantially. Swedish women have increased their enrolment in higher education much more than men, and from the late 1970s women have outnumbered men in tertiary education (Universitetskanslersämbetet 2016). So, even if the positive selection into paid labor among women was higher at the beginning of the period, the
overall skill level, both at the occupational level and at the individual level, has increased substantially over time, leading to higher relative wages for women in general, but not necessarily for mothers in comparison to non-mothers.

In comparison with Petersen et al.’s (2014) study on Norway, another Scandinavian welfare state, we may assume the development of parenthood premiums and penalties to be similar in Sweden. However, there are differences that might impact gender inequality in the labor market (Datta Gupta et al. 2008; Grönlund et al. 2017). One important difference is that the expansion of childcare facilities took place earlier and was more rapid in Sweden (Ellingsæter and Jensen 2019). Another is that the share of part-time workers is larger in Norway. Thus, the entry and establishment of mothers in paid labor took place later in Norway (Ellingsæter 2018), implying that the timing of any changes in parenthood effects on wages plausibly took place at an earlier point in time in Sweden compared to Norway.

Before we move on, it is important to bear in mind that a one-country study such as this cannot resolve the role of the dual-earner model for the parenthood effect on wages. Although our data cover a period before and after the implementation of these policies, we cannot analytically separate the effects of different dual-earner policies (for instance, the public provision of childcare may promote gender equality, whereas the right to take long parental leave may not, see Budig et al. 2016; Halldén et al. 2016), and we cannot distinguish between the effects of these policies and those of other causes that vary over time and are unobserved in our data (e.g., other policy changes, changes in norms that are exogenous to family policies). We turn next to what we instead consider to be our main contribution to research.

The Contribution of the Present Study

In the present study, we have access to an individual-level panel data set covering a period of over 40 years. Because many individuals in these data have been observed at more than one point in time, our estimates are able to account for time-invariant characteristics that sort individuals into jobs with a certain wage level. Our basic analytical strategy is to estimate fixed-effects (FE) regression models accounting for individual observed and unobserved heterogeneity, and to compare these to standard regression models which account for observed heterogeneity but ignore individual unobserved heterogeneity on stable characteristics. We adopt this strategy to test the relative weight of causal treatment effects of parenthood versus selection effects. In terms of treatment effects, we will test whether these are mediated through labor market related inputs. In addition, we study whether, and how, these associations have changed over time, dividing our study period into three periods, roughly corresponding to “before”, “during”, and “after” the gradual implementation of the dual-earner family policy model.

First, we estimate the following “total” model, for all cross sections and pooled over time, for men and women separately: 

$$ Y_{it} = \alpha + \beta_j X_{ijt} + \gamma_t + \varepsilon_{it}, $$

where $Y_{it}$ equals individual $i$’s wage rank at time point $t$, $\beta_j$ is the coefficient
vector associated with $X_{ijt}$, and $\varepsilon_{it}$ is the random error term for individual $i$ at time $t$. As the data contain individuals $i$ across $t$, we have added a period dummy $\gamma_t$ for each survey year. Second, we add a dummy $\delta_i$ for each individual $i$: $Y_{it} = \alpha + \beta_j X_{ijt} + \gamma_t + \delta_i + \varepsilon_{it}$. We thereby capture wage effects of all observed and unobserved individual characteristics that do not change over the observation period, as well as all observed and unobserved gender-specific factors that for a particular survey year affect an individual’s wage rank (because the models are estimated by gender). This (FE) model shows what happens to the dependent variable when individuals experience a change in parental status, for men and women separately. Because the FE model uses only intra-individual variation to estimate effects, this model is only representative of the subselection(s) of individuals exhibiting such variation. Our fixed-effects modeling approach over time parallels that of Datta Gupta and Smith (2002). Petersen et al.’s (2014), approach differs from ours as our fixed-effects are located at the observation-unit individual level while theirs are located at the establishment level, occupation level, and the establishment-occupation levels combined.

**Data and Variables**

The data used here were retrieved from the Swedish Level of Living Surveys (LNU) collected in 1968, 1974, 1981, 1991, 2000, and 2010. In 1968, a random, nationally representative sample of approximately 0.1 percent of the Swedish adult population (then aged 15–75 years, in later waves aged 18–75) was selected for the survey. Individuals who still fulfilled the age criteria were invited to take part in each subsequent survey wave. Thus, LNU is based on a panel structure in which individuals participate repeatedly. Individuals who pass age 75 leave the sample frame but are replaced by individuals who were too young to take part in the previous survey. Moreover, a proportional sample of immigrants, who moved to Sweden between survey waves, is added to the sample. Thus, each survey wave can also be used as a nationally representative sample for cross-sectional purposes. A unique feature of these data is that they include detailed and identical measures of labor market and family-related variables over time.

Respondents were excluded during years in which they did not work at least 1 hour per week. We also excluded a small number of observations with internal non-response on relevant questions, and respondents past the age of 65 (the common retirement age in Sweden). The samples then consisted of 2,757 respondents in 1968, 3,012 respondents in 1974, 3,301 respondents in 1981, 3,317 respondents in 1991, 2,956 respondents in 2000, and 2,291 respondents in 2010 (which equals 17,634 observations in total). The data contain information from 7,637 unique respondents and 84 percent of these contribute with more than one cross-sectional observation.

Below we describe the variables used in the analyses. To net out time-varying effects of inflation and wage dispersion, we used an individual’s within-year percentile *hourly wage rank* as the dependent variable for this study. This measure is comparable over time and reflects an individual’s wage (defined
as the sum of the respondent’s pre-tax monthly pay, divided by hours worked per month) relative to other employees at the time. Coefficients thus refer to percentile positional differences (the total model) or changes (the fixed-effects model) in the total wage distribution for a one unit difference (the total model), or change (the fixed-effects model) in the independent variable. For the independent variable woman, we assigned the value 0 for males and 1 for females. Age is measured in years since birth. We constructed a dummy variable indicating whether the respondent is a parent of nonadult child(ren) in the household (child aged 0–20 years in the 1968, 1974, and 1981 surveys and 0–18 years in the 1991, 2000 and 2010 surveys). We measured education in years of formal education (“For how many years have you been in full-time school and vocational education (from first grade and onwards)?”). To study the mediating impact of labor market inputs, we used five different indicators. We measured work experience as the respondent’s total number of years in employment (“How many years altogether have you spent in gainful employment?”). We measured seniority as the number of years since the respondent was first employed by the current employer. This variable is most probably an overestimation of actual seniority, and more so for women, as it may include periods of temporary leave. The measure of part-time work has the value 1 if the respondent worked more than 0 but less than 30 hours per week, and 0 if the respondent worked 30 hours or more per week. We measure workplace authority in terms of number of subordinates, with three dummy variables (0 (ref.); 1–5; 6–10; 11+) and we used a variable for sector of employment, private, with the value 1 if the individual is employed by a privately owned firm and 0 otherwise. Finally, becoming and being a parent is correlated with conjugal union status. Parents are more prone to be married or cohabiting with a partner than are non-parents. Research shows that cohabitation and marriage are positively related to both men’s and women’s wages (Budig and England 2001; Killewald and Gough 2013; Budig and Lim 2016), although the premium is larger for men than for women (Petersen et al. 2014). Given this association, and given the correlation between parenthood and union status, we added a dummy variable for the respondent being married/cohabiting (value 1) or not (0) to our models. To account for potential nonlinear effects, squared terms of the seniority, work experience, and age variables were included in the empirical models.

Descriptive statistics for the variables included in the analyses are presented in Figure 1 and Table 1. As noted, only women and men working at least 1 hour per week are included here and changes in distributions over time are in line with expectations. One of the most startling changes during the observation period is the gender equalization of labor force participation. In 1968, less than every second woman aged 18–65 was gainfully employed. Forty years later, the share had increased to around 70 percent, in parity with the share for men (Figure 1). There has also been an equalization between men and women regarding their work experience and seniority, and whereas women still work part-time to a higher extent than men, the difference is substantially smaller in 2010 than it was four decades earlier (Table 1). Men and women are still found in different sectors of the labor market, however, and there is no sign of integration over
time. The primary focus in this paper, however, is directed at how the wage rank of women and men developed during the period examined. From the data, it is clear that men on average are advantaged in relation to women but also that the gender gap declined during the period. In 1968, it was equal to 24 percentile points (62–38). In 2010, the gap had decreased to 12 percentile points, half of its level 40 years earlier.

**Findings**

We will now turn to the questions of whether parenthood is associated with wage rank, whether this association varies by gender, and whether there is change in these associations over time. To investigate these patterns more closely, we model how parenthood is associated with women’s and men’s wage rank, with and without controls. We repeat this approach for all survey years, which gives us an indication of whether the impact of these conditions is stable or varies over time. In Table 2, Model 1, we report the coefficient for our primary focus, parenthood, for men in 1968. This year it seems that being a father to child(ren) aged 0–18 had a strongly positive association with men’s wage rank. On average, fathers were located around 12 percentiles higher up the wage distribution than non-fathers. This is a substantial difference. Model 3 shows, however, that this difference could largely be explained by differences in human capital, labor market position, and civil status between fathers and non-fathers, to the advantage of fathers. However, the association between fatherhood and wage
### Table 1. Means for Cross-Sectional Data Variables (individuals working at least 1 hour a week)

| Variable                        | 1968 | 1974 | 1981 | 1991 | 2000 | 2010 | 1968 | 1974 | 1981 | 1991 | 2000 | 2010 |
|---------------------------------|------|------|------|------|------|------|------|------|------|------|------|------|
| Wage rank                       | 62.43| 60.21| 58.97| 59.71| 58.47| 57.21| 38.14| 38.68| 39.24| 39.39| 41.68| 45.07|
| Wage (ln), 2010 prices          | 4.56 | 4.70 | 4.69 | 4.74 | 4.94 | 5.15 | 4.23 | 4.44 | 4.50 | 4.54 | 4.77 | 5.01 |
| 1–5 subordinates                | 0.12 | 0.14 | 0.13 | 0.13 | 0.13 | 0.14 | 0.09 | 0.09 | 0.11 | 0.10 | 0.12 | 0.08 |
| 6–11 subordinates               | 0.05 | 0.06 | 0.08 | 0.07 | 0.06 | 0.06 | 0.02 | 0.02 | 0.04 | 0.04 | 0.04 | 0.04 |
| 11+ subordinates                | 0.11 | 0.11 | 0.10 | 0.10 | 0.10 | 0.10 | 0.03 | 0.04 | 0.04 | 0.04 | 0.04 | 0.07 |
| Married/cohabiting              | 0.74 | 0.75 | 0.72 | 0.71 | 0.71 | 0.72 | 0.63 | 0.72 | 0.72 | 0.72 | 0.75 | 0.73 |
| Hours of paid work per week (/10)| 42.77| 39.73| 38.95| 39.07| 39.11| 39.08| 34.36| 31.54| 31.41| 33.82| 35.30| 35.88|
| Education in years              | 8.76 | 9.85 | 10.73| 11.72| 12.62| 13.59| 8.76 | 9.70 | 10.36| 11.52| 12.77| 13.99|
| Work experience in years        | 22.86| 21.82| 20.76| 20.08| 20.17| 21.71| 15.12| 14.66| 15.64| 16.78| 19.28| 21.25|
| Seniority in years              | 10.07| 10.17| 9.97 | 10.41| 10.28| 8.22 | 6.40 | 6.95 | 7.88 | 9.53 | 10.40| 8.22 |
| Private sector                  | 0.78 | 0.74 | 0.70 | 0.70 | 0.71 | 0.74 | 0.55 | 0.47 | 0.39 | 0.39 | 0.39 | 0.43 |
| Parent of nonadult child(ren)   | 0.66 | 0.68 | 0.65 | 0.65 | 0.67 | 0.68 | 0.60 | 0.70 | 0.71 | 0.71 | 0.75 | 0.77 |
| Age                             | 39.80| 39.55| 39.25| 39.70| 40.79| 42.80| 38.65| 38.56| 38.79| 39.61| 41.84| 44.07|
| Proportion of cross section     | 0.62 | 0.57 | 0.52 | 0.50 | 0.51 | 0.52 | 0.38 | 0.43 | 0.48 | 0.50 | 0.49 | 0.48 |
| N                               | 1700 | 1709 | 1727 | 1659 | 1508 | 1198 | 1057 | 1303 | 1574 | 1658 | 1448 | 1093 |
Table 2. OLS Estimates of Wage Rank on Independent Variables. Cross-Sectional Data 1968 and 1974

| Variable/model                  | 1968      | 1974      |
|--------------------------------|-----------|-----------|
|                                | (1) Men   | (2) Women | (3) Men   | (4) Women | (5) Men   | (6) Women | (7) Men   | (8) Women |
| Parent of nonadult child(ren)   | 12.431**  | 5.027**   | 2.019     | 1.269     | 11.034**  | 5.738**   | 1.056     | 3.490*    |
|                                | (1.161)   | (1.596)   | (1.246)   | (1.453)   | (1.215)   | (1.459)   | (1.367)   | (1.494)   |
| Education years                 | 2.688**   | 5.000**   | 2.394**   | 3.915**   |
|                                | (0.219)   | (0.240)   | (0.209)   | (0.219)   |
| Part-time work                  | 4.706     | −0.719    | −0.286    | −0.044    |
|                                | (4.438)   | (1.457)   | (3.394)   | (1.335)   |
| Seniority years (/10)           | 3.374*    | 7.915**   | 5.433**   | 10.742**  |
|                                | (1.405)   | (2.393)   | (1.572)   | (2.110)   |
| Seniority squared (/100)        | −0.368    | −0.976    | −0.618    | −1.953**  |
|                                | (0.354)   | (0.808)   | (0.382)   | (0.523)   |
| Work experience years           | 0.484     | 0.894**   | 0.613     | 0.381     |
|                                | (0.332)   | (0.237)   | (0.326)   | (0.256)   |
| Work experience squared (/100)  | −1.115*   | −1.315*   | −0.590    | −0.056    |
|                                | (0.561)   | (0.515)   | (0.560)   | (0.555)   |
| Private sector                  | −1.872    | −5.270**  | 0.180     | −4.313**  |
|                                | (1.254)   | (1.213)   | (1.258)   | (1.231)   |
| 1–5 subordinates                | 9.109**   | 9.324**   | 8.087**   | 8.329**   |
|                                | (1.553)   | (2.053)   | (1.579)   | (2.150)   |
| 6–10 subordinates               | 10.602**  | 7.882*    | 11.799**  | 2.025     |

(Continued)
| Variable/model                        | 1968         | 1974         |
|--------------------------------------|--------------|--------------|
|                                      | (1) Men      | (2) Women    | (3) Men      | (4) Women    |
|                                      | (2.177)      | (3.920)      | (2.248)      | (3.897)      |
| 10+ subordinates                     | 16.220**     | 15.644**     | 16.005**     | 15.211**     |
|                                      | (1.658)      | (3.428)      | (1.822)      | (3.218)      |
| Age                                  | 1.930**      | 2.348**      | 2.381**      | 2.642**      |
|                                      | (0.576)      | (0.441)      | (0.588)      | (0.458)      |
| Age squared (/100)                   | −2.103**     | −2.879**     | −3.106**     | −3.129**     |
|                                      | (0.660)      | (0.441)      | (0.672)      | (0.553)      |
| Cohabiting/married                   | 7.834**      | 1.870        | 6.576**      | −0.376       |
|                                      | (1.437)      | (1.416)      | (1.514)      | (1.459)      |
| Constant                             | 56.210**     | 35.945**     | −15.542      | −61.494**    |
|                                      | (0.821)      | (1.055)      | (8.697)      | (7.537)      |
| Observations                         | 1,700        | 1,057        | 1,700        | 1,057        |
| Adjusted R-squared                   | 0.063        | 0.008        | 0.367        | 0.481        |
|                                      |             |             |             |             |
|                                      |             |             |             |             |
|                                      | (Continued)  | (Continued)  | (Continued)  | (Continued)  |

**OLS Estimates of Wage Rank on Independent Variables. Cross-Sectional Data 1981 and 1991**

| Variable/model                        | 1981         | 1991         |
|--------------------------------------|--------------|--------------|
|                                      | (9) Men      | (10) Women   | (11) Men     | (12) Women   |
|                                      | (13) Men     | (14) Women   | (15) Men     | (16) Women   |
| Parent of nonadult child(ren)        | 11.527**     | 2.351        | 0.229        | −2.589       |
|                                      | (1.317)      | (1.341)      | (1.503)      | (1.448)      |
| Education years                      | 2.500**      | 3.387**      |             |             |
|                                      | (0.233)      | (0.212)      |             |             |
### Table 2. Continued

| Variable/model                        | 1981          | 1991          |
|---------------------------------------|--------------|--------------|
|                                       | (9) Men      | (10) Women   | (11) Men | (12) Women | (13) Men | (14) Women | (15) Men | (16) Women |
| Part-time work                        | -0.184       | 4.707**      | -9.422** | 4.042**    |
|                                       | (2.690)      | (1.244)      | (2.691)  | (1.310)    |
| Seniority years (/10)                 | 9.049**      | 11.533**     | 6.477** | 2.429      |
|                                       | (2.049)      | (2.409)      | (1.877)  | (1.940)    |
| Seniority squared (/100)              | -1.249*      | -2.182**     | -1.130* | 0.452      |
|                                       | (0.589)      | (0.792)      | (0.523)  | (0.591)    |
| Work experience years                 | 0.722        | 0.998**      | 1.932** | 1.427**    |
|                                       | (0.375)      | (0.297)      | (0.381)  | (0.314)    |
| Work experience squared (/100)        | -1.285*      | -1.436*      | -1.997**| -2.204**   |
|                                       | (0.610)      | (0.608)      | (0.690)  | (0.651)    |
| Private sector                        | 2.015        | -2.449**     | 7.856** | 4.476**    |
|                                       | (1.284)      | (1.202)      | (1.281)  | (1.138)    |
| 1–5 Subordinates                      | 6.079**      | 4.562*       | 6.249** | 6.411**    |
|                                       | (1.741)      | (1.827)      | (1.719)  | (1.803)    |
| 6–10 Subordinates                     | 8.280**      | 8.516**      | 12.106**| 11.105**   |
|                                       | (2.217)      | (2.962)      | (2.243)  | (2.887)    |
| 10+ Subordinates                      | 14.645**     | 12.372**     | 17.619**| 17.343**   |
|                                       | (2.024)      | (2.969)      | (1.954)  | (2.756)    |
| Age                                   | 1.723**      | 2.074**      | 0.406   | 0.838      |

(Continued)
### OLS Estimates of Wage Rank on Independent Variables. Cross-sectional Data 2000 and 2010

| Variable/model                              | 2000          | 2010          |
|---------------------------------------------|---------------|---------------|
|                                             | (17) Men      | (18) Women    |
|                                             | (19) Men      | (20) Women    |
| Parent of nonadult child(ren)               | 8.425**       | 1.833         |
|                                             | (1.433)       | (1.429)       |
| Education years                             | 3.609**       | 3.873**       |
|                                             | (0.257)       | (0.227)       |
| Part-time work                              | −11.315**     | −2.235        |
|                                             | (2.987)       | (1.589)       |
| Seniority years (/10)                       | 4.398*        | 5.306*        |
|                                             | (1.977)       | (2.092)       |
| Variable/model                      | 2000               | 2010               |
|------------------------------------|--------------------|--------------------|
|                                    | (17) Men          | (18) Women        | (19) Men          | (20) Women        | (21) Men          | (22) Women        | (23) Men          | (24) Women        |
| Seniority squared (/100)           | -0.529 (0.568)    | -0.915 (0.632)    | 0.837 (0.671)     | 0.816 (0.742)     |
| Work experience yrs                | 2.083** (0.356)   | 1.296** (0.334)   | 0.614 (0.412)     | 0.878* (0.352)    |
| Work experience squared (/100)     | -2.360** (0.643)  | -1.244 (0.660)    | 0.175 (0.763)     | -0.043 (0.689)    |
| Private sector                     | 11.427** (1.376)  | 9.890** (1.275)   | 8.530** (1.646)   | 7.786** (1.390)   |
| 1–5 Subordinates                   | 8.962** (1.784)   | 8.163** (1.859)   | 9.472** (1.997)   | 6.707** (2.436)   |
| 6–10 Subordinates                  | 15.566** (2.423)  | 12.296** (3.099)  | 13.044** (2.878)  | 11.644** (3.558)  |
| 10+ Subordinates                   | 19.394** (2.001)  | 23.510** (2.921)  | 17.872** (2.312)  | 22.693** (2.635)  |
| Age                                | -0.374 (0.689)    | 2.358** (0.634)   | 2.408** (0.756)   | 2.001* (0.682)    |
| Age squared (/100)                 | -0.164 (0.797)    | -3.120** (0.715)  | -2.863** (0.866)  | -2.570** (0.776)  |
| Cohabiting/married                 | 5.597** (1.505)   | 2.291 (1.403)     | 7.382** (1.737)   | 2.627 (1.578)     |
| Constant                           | 54.770** (0.949)  | 40.784** (0.998)  | -15.854 (11.821)  | -75.470** (10.869) |
|                                    | 52.049** (1.098)  | 41.926** (1.138)  | -66.967** (13.090) | -70.434** (11.995) |
| Observations                       | 1,508 1,508       | 1,448 1,448       | 1,508 1,448       | 1,448 1,448       |
| Adjusted R-squared                 | 0.022 0.352       | 0.000 0.329       | 0.038 0.013       | 0.322 0.340       |

**p < 0.01, *p < 0.05 that coefficient is equal to zero (t-test, two-tailed).
rank remains positive (although not statistically significant). Given previous research in the field, this finding may come as no surprise. Less expected, perhaps, is that we also find a parenthood wage premium for women (see Model 2). After controls, the premium remains but loses statistical significance, just as for men (Model 4).

What about trends over time? Men experience a gross fatherhood premium throughout the period. Controls for human capital, labor market position, and civil status explain most of this premium, however, and net effects hover around zero. For women, gross parenthood coefficients are substantially smaller than for men but they never turn negative. These coefficients again approach zero once controls are added to the models, and even turn significantly negative in 1 year (2000). The net gender wage gap for parents is also fairly small throughout the period under study. These trends are graphically displayed in Figure 2. It should be noted that apart from the fact that men gain more than women, we do not see any clear time trend with regard to this pattern, regardless of whether we look at gross or net effects.

So far, we have only made use of the cross-sectional character of our data. We have compared wage ranks between individuals with different traits, for example, parents versus non-parents, and we have identified important wage differences between them. A weakness with this procedure, however, is that we ignore many potential selection effects. Unobserved individual characteristics, which are correlated with observed independent variables, may give rise to omitted variable bias in the estimated coefficients. For example, a fatherhood premium for wages in a cross section may have arisen because employers use traits that are unobserved in the data to raise wages. If these traits are correlated with the probability of being a father, omitted variable bias would arise in the estimated effect of fatherhood. To limit this problem, we make use of the longitudinal character of our data and estimate fixed-effects models (FE).

In Table 3, we compare results from OLS and FE regressions, using data pooled over the whole observation period (1968–2010).

Again, we introduce the parenthood variable in a first model. In a second model, we add the human capital variables, labor market position, age, civil status, and the period dummy variables. The total pooled models reported in columns 1 and 2 (men) and 5 and 6 (women) in Table 3 summarize what we have already observed in the year-specific analyses in Table 2. For women and men alike, wage rank increases with education, workplace authority, seniority, work experience, and age (although with a decreasing effect for the latter three), whereas work in the private sector and conjugal unions appear to be beneficial for men but less so for women. Part-time work seems to be negative for men but not for women. We also note that parenthood is associated with higher wages for both genders but that the associations are attenuated once we add controls. That is, parents (mothers and fathers) earn more than non-parents because they outrank them with regard to labor market inputs and labor market position. They also more often live in conjugal unions, thus enjoying marriage premiums.
### Table 3. OLS Estimates of Wage Rank on Independent Variables. Individual Panel Data 1968–2010

| Variable/model                  | (1) Men total | (2) Men total | (3) Men FE | (4) Men FE | (5) Women total | (6) Women total | (7) Women FE | (8) Women FE |
|---------------------------------|---------------|---------------|------------|------------|----------------|----------------|--------------|--------------|
| Parent of nonadult child(ren)   | 11.310**      | 0.697         | 9.144**    | 0.215      | 3.982**        | −0.408         | 4.637**      | −1.798*      |
|                                 | (0.579)       | (0.633)       | (0.547)    | (0.675)    | (0.624)        | (0.632)        | (0.591)      | (0.717)      |
| Education years                 | 2.936**       | 0.600**       | 3.851**    | 1.440**    | 1.440**        | 5.742**        |             |             |
|                                 | (0.120)       | (0.184)       | (0.115)    | (0.234)    | (0.234)        | (0.234)        |             |             |
| Part-time work                  | −4.329*       | 0.928         | 1.628*     | 5.742**    | 5.742**        | 5.742**        |             |             |
|                                 | (1.700)       | (1.431)       | (0.681)    | (0.724)    | (0.724)        | (0.724)        |             |             |
| Seniority years (/10)           | 4.942**       | 3.085**       | 5.495**    | 2.976**    | 2.976**        | 2.976**        |             |             |
|                                 | (0.755)       | (0.817)       | (0.886)    | (0.957)    | (0.957)        | (0.957)        |             |             |
| Seniority squared (/100)        | −0.623**      | −0.512*       | −0.757**   | −0.901**   | −0.901**       | −0.901**       |             |             |
|                                 | (0.192)       | (0.219)       | (0.255)    | (0.272)    | (0.272)        | (0.272)        |             |             |
| Work experience years           | 1.348**       | 0.457*        | 0.941**    | 0.799**    | 0.799**        | 0.799**        |             |             |
|                                 | (0.162)       | (0.187)       | (0.125)    | (0.157)    | (0.157)        | (0.157)        |             |             |
| Work experience squared (/100)  | −1.843**      | −0.742*       | −0.999**   | −1.106**   | −1.106**       | −1.106**       |             |             |
|                                 | (0.262)       | (0.293)       | (0.255)    | (0.287)    | (0.287)        | (0.287)        |             |             |
| Private sector                  | 4.312**       | 4.571**       | 1.822**    | 1.999*     | 1.999*         | 1.999*         |             |             |
|                                 | (0.634)       | (0.960)       | (0.625)    | (0.861)    | (0.861)        | (0.861)        |             |             |

(Continued)
### Table 3. Continued

| Variable/model          | (1) Men total | (2) Men total | (3) Men FE | (4) Men FE | (5) Women total | (6) Women total | (7) Women FE | (8) Women FE |
|-------------------------|---------------|---------------|------------|------------|----------------|----------------|-------------|-------------|
| 1–5 Subordinates        | 8.369**       | 4.014**       | 7.774**    | 5.067**    |                |                |             |             |
|                         | (0.762)       | (0.805)       | (0.923)    | (0.958)    |                |                |             |             |
| 6–10 Subordinates      | 12.184**      | 4.478**       | 10.926**   | 7.117**    |                |                |             |             |
|                         | (0.935)       | (1.061)       | (1.373)    | (1.560)    |                |                |             |             |
| 10+ Subordinates       | 16.989**      | 7.872**       | 19.349**   | 12.577**   |                |                |             |             |
|                         | (0.737)       | (0.991)       | (1.178)    | (1.434)    |                |                |             |             |
| Age                    | 0.916**       | 2.703**       | 1.961**    | 1.340*     |                |                |             |             |
|                         | (0.283)       | (0.634)       | (0.213)    | (0.566)    |                |                |             |             |
| Age squared (/100)     | −1.258**      | −2.900**      | −2.391**   | −2.475**   |                |                |             |             |
|                         | (0.311)       | (0.329)       | (0.252)    | (0.293)    |                |                |             |             |
| Cohabiting/married     | 6.118**       | 3.988**       | 1.271*     | 1.543      |                |                |             |             |
|                         | (0.720)       | (0.797)       | (0.639)    | (0.791)    |                |                |             |             |
| Year dummies           | No            | Yes           | No         | Yes        | No             | Yes            | No          | Yes         |
| Constant               | 54.613**      | −10.658*      | 55.621**   | −13.901    | 38.093**       | −48.594**      | 37.774**    | −11.984     |
|                         | (0.465)       | (4.477)       | (0.320)    | (13.430)   | (0.486)        | (3.548)        | (0.360)     | (11.476)    |
| Observations           | 9,501         | 9,501         | 9,501      | 9,501      | 8,133          | 8,133          | 8,133       | 8,133       |
| R-squared              | 0.043         | 0.324         | 0.726      | 0.767      | 0.006          | 0.326          | 0.695       | 0.751       |

**p < 0.01, *p < 0.05** that coefficient is equal to zero (t-test, two-tailed).

*aThe total estimator standard errors are adjusted for intra-individual correlation.
Figure 2. Gross and net parenthood effects on wage rank by gender and survey year, with 95 percent confidence intervals
For women and men alike, the coefficients for parenthood from the FE models, reported in Models 3 and 4 (men), and 7 and 8 (women), are parallel to the findings from the total models (1 and 2, and 5 and 6, respectively), which suggests that the total model effects are generally not biased as a result of unobserved individual heterogeneity. For both genders, the positive gross parenthood effect approaches zero for men and even turns slightly negative for women when we add controls in the FE models (4 and 8), implying that most of the positive parenthood effect can be attributed to its impact on time-varying labor market related factors, for example, increases in human capital, promotions and switches to jobs in the private sector, and, for men, the initiation of a conjugal union. Thus, the results indicate that when men and women become parents, they make wage gains mostly because of increased labor market related inputs and a marriage premium (men). Without these inputs and changes in civil status, the fathers’ wage rank would equal that of non-fathers (Model 4), whereas the mothers’ wage rank would be slightly lower than that of non-mothers (Model 8).

Changes Over Time

As has been noted, we find no consistent pattern in the development of the parenthood premium over time. In order to investigate this more closely, we partitioned our data into three periods, 1968–1974, 1981–1991 and 2000–2010. We have two reasons for this specific partitioning. First, female labor force participation was significantly lower in the late 1960s and the 1970s, but has since the 1980s been on a par with the labor force participation of men (see Figure 1). Second, a number of influential policy reforms were introduced at the beginning of the 1970s, were then developed during the 1980s and the 1990s, and came into “full force” (e.g., full availability of subsidized childcare) during the final period, 2000–2010. This approach helps us to compare the association between parenthood and wages before, during and after the introduction of important family policies. Results from this analysis are reported in Figure 3 (the period-specific effects are based on coefficients reported in Table 4). Here we find a gross parenthood premium for men at all points in time, and for women in all periods except the first. These effects tend to be higher for men. Premiums are entirely statistically explained by labor market related factors and civil status. The figure reveals that the significant motherhood penalty that we observe in the FE model with controls, when we pool all survey years 1968–2010 (see also Table 3, Model 8), can be attributed to the period 1968–1974. For the other periods, 1981–1991 and 2000–2010, we find no such (significant) negative effect. A formal test (see Table 4) shows that there are no statistically significant changes in the fixed-effects coefficients between the first period (1968–1974) and the last (2000–2010) under study, neither for women nor for men, and regardless of whether models with or without controls are compared. Thus, we cannot reject the null of no period change in these “treatment” parenthood effects. Instead, the picture from Table 2 and Figure 2 of rather little change over time is confirmed.
### Table 4. OLS Estimates of Wage Rank on Independent Variables, Across Periods. Individual Panel Data 1968–2010. Control variables (suppressed) are identical to those reported in Table 3

| Period          | Men total | Men total + controls | Men FE | Men FE + controls | Women total | Women total + controls | Women FE | Women FE + controls |
|-----------------|-----------|----------------------|--------|-------------------|-------------|------------------------|----------|---------------------|
| 1968–1974       |           |                      |        |                   |             |                        |          |                     |
| Parent of nonadult child(ren) | 11.738**  | 1.501                | 5.898**| −1.623            | 5.331**     | 2.349*                 | 1.243    | −4.035*             |
|                 | (0.910)   | (0.927)              | (1.197)| (1.391)           | (1.197)     | (1.098)                | (1.480)  | (1.709)             |
| 1981–1991       |           |                      |        |                   |             |                        |          |                     |
| Parent of nonadult child(ren) | 11.842**  | 0.433                | 7.353**| 0.336             | 3.011**     | −1.291††               | 5.678**† | 1.289†              |
|                 | (0.973)   | (1.085)              | (1.278)| (1.516)           | (0.956)     | (0.977)                | (1.337)  | (1.654)             |
| 2000–2010       |           |                      |        |                   |             |                        |          |                     |
| Parent of nonadult child(ren) | 9.673**   | 0.466                | 6.583**| −1.320            | 3.801**     | −1.678††               | 3.508*   | −1.461              |
|                 | (1.090)   | (1.230)              | (1.333)| (1.570)           | (1.124)     | (1.163)                | (1.404)  | (1.603)             |

Robust standard errors in parentheses, significance levels related to coefficients equal to zero: ** \( p < 0.01 \), * \( p < 0.05 \) (t-test, two-tailed). Significance levels in tests of zero coefficient differences between current and previous period: †† \( p < 0.01 \), † \( p < 0.05 \); significance levels in tests of zero differences in coefficients between the 2000–2010 period and the 1968–1974 period: ††† † † † † † † † † (Cohen et al. 2003).
**Sensitivity Analyses**

First, in accordance with some previous research, we replicated our analyses using the logarithm of wage instead of wage rank as the dependent variable. These analyses yielded some dissimilarities between the models (see Appendix Figures S1 and S7 and Tables S1 and S2). In the analyses where the dependent variable is the logarithm of wage i) the 1968 net parenthood premium for men is statistically significant, ii) the 1974 net parenthood premium for women is statistically non-significant, and iii) the 1981 net parenthood penalty for women is statistically significant. These differences do not, however, suggest that we should alter any of the substantial conclusions drawn from our main analysis. Second, we re-estimated the models reported in Table 3 for men and women with at least median educational level at a cross section. These analyses indicate that highly educated men and women receive higher gross parenthood premiums, but net effects of parenthood are similar to those for the entire sample (see Appendix Table S3). Third, we re-estimated the models for the age span 25–55. The estimated total gross effects of parenthood are uniformly closer to zero for both men and women in this age span, but parenthood net effects do not differ much from those reported for the whole sample (see Appendix Table S4). Fourth, we re-estimated the models with dummies indicating the age of the children in the household instead of the simple yes/no dummy used in the analyses reported above. In the conditional fixed-effects models, we find that having younger children (aged 0–6 years) is associated with a wage premium for both genders,
whereas having older children (7–20 years) is associated with a wage penalty for both genders, but it is larger for women (see Appendix Table S5). Thus, patterns are similar for mothers and fathers but premiums are smaller for women when children are young and penalties are larger when children are older. Fifth, we re-estimated the models with dummies indicating the presence of 1, 2, or 3+ children in the household. For both genders, coefficients in the fixed-effects gross models tend to increase with the number of children (see Appendix Table S6), but an important distinction seems to be that between childless households and households with children.

Discussion

The present paper evaluates the size of parenthood effects on wages during the period 1968–2010 using data from six nationally representative waves of the Swedish Level of Living Survey. The length of the period under study is one of the merits of the present analysis, as we can show how associations between parenthood and wages evolved during a transformative period with an inflow of women to the labor market, during which Sweden changed from a male-breadwinner to a dual-earner society. We suggest a number of reasons why this development should alter the association between parenthood and wages.

According to the attenuation hypothesis, the policies introduced in the 1970s and 1980s to stimulate dual-earner families, for example, individual taxation, parental leave benefits, subsidized public childcare, earmarked parental leave for fathers, and men’s greater involvement in childcare and household work (Ferrarini and Duvander 2010; Boye 2014), should have led to convergence in men’s and women’s career opportunities, resulting in attenuated fatherhood wage premiums and motherhood wage penalties over time. In contrast, the exacerbation hypothesis suggests that because mothers still assume the main responsibility for childcare, increased periods of parental leave, part-time work, and staying home with sick children, as an effect of changed policies, would harm mother’s careers and result in increased motherhood wage penalties over time. In addition, the large and rapid increase in female labor force participation, from the 1960s onwards, may have resulted in an inflow of less work-oriented women, and mothers in particular, and the previous positive selection of women into paid labor may have weakened, resulting in a less positive (or more negative) association between parenthood and wages for women over time.

Our results however demonstrate a remarkable stability over time in the estimated wage effects of parenthood. In fact, in 2010 the associations between parenthood and wage rank are very similar to those in 1968; regardless of gender and regardless of whether fixed-effects models with or without controls are compared, there is no statistically significant change between the start and the end of the time period under study. Thus, the hypotheses presented above, both of which predict change over time, receive no support from the analyses. We are not aware of any theory postulating stability in parenthood premiums/penalties over time, particularly not during a period of such radical
institutional transformation as the one we study here. We might speculate that there may have been countervailing influences from the policy changes implemented during the period: increasing incentives for dual-earner households from individual taxation and subsidized childcare but prolonged periods of work absence from paid parental leave and part-time work may have led to an inflow of less work-committed mothers to the labor market, while at the same time drawing more career-oriented mothers from the labor market, at least temporarily. These processes counteract one another and result in a steady status quo in the association between parenthood and wage rank.

The stability in parenthood effects for men may be surprising given fathers’ increasing involvement in childcare and household work. However, it is consistent with American (Glauber 2018) and Norwegian (Petersen et al. 2014) studies on this development over time. These studies also, however, find a strong decrease in the motherhood penalty during the periods 1979–1996 (Norway) and 1980–2014 (United States). Petersen et al. see much of this change in a period during which family policies were expanded (1988–1993). We see no such development in our data. Instead, the motherhood penalty (in fixed-effects models with controls) was rather modest already in 1968–1974 and statistically non-significant in 2000–2010.

The different patterns found for Swedish and Norwegian (Petersen et al. 2014) mothers might be due to timing differences in the inflow of women to the labor market. In the 1960s, the share of women in paid labor was much lower in Norway than in other Nordic countries. Moreover, studies indicate that gender attitudes in Norway were relatively traditional during the 1970s (Ellingsæter 2018). These country differences may account for some of the differences in parenthood effects among mothers in Norway and Sweden.

In our fixed-effects models without controls, both men and women receive parenthood premiums (although the premium is larger for men). The higher wages among parents in our analyses could largely be attributed to the fact that mothers and fathers on average outrank nonparents with regard to labor market related investments and achieved positions. Fixed-effects models with controls indicate that the positive parenthood effect can be explained by its impact on time-varying labor market related inputs, for example, increases in human capital and promotions. Fathers also gain from the initiation of a conjugal union, which often appears in conjunction with becoming a parent. Such a “marriage premium” has repeatedly been found in a number of studies (Budig and England 2001; Killiewald and Gough 2013; Budig and Lim 2016), including Scandinavian and Swedish research (Dribe and Nystedt 2013; Petersen et al. 2014).

The gender difference in wage rank is rather substantial during the entire observation period (see Table 1). This could be seen as support for the criticism that has been raised against family-friendly policies. Dual-earner policies may pave the way for statistical discrimination among employers, which implies that all women in gender egalitarian countries, and not only mothers, get punished on the labor market. Our results do not support this conjecture, however, as the association between motherhood and wages is rather stable over time. We find the same effect of parenthood among women at the end of the period, when
the dual-earner model is well established and labor force participation among mothers is high. Family-friendly policies are also assumed to have particularly negative consequences for high-skilled women. Again, our sensitivity analyses for highly educated men and women dispute this assumption, as their positive gross parenthood effects are higher than those for mothers with less education.

The difference in wage rank between fathers and mothers could partially be caused by men and women working in different occupations. However, part of the occupational gender segregation is accounted for when we control for sector of employment. Gender differences in working conditions could be another explanation for the gender wage gap. Fathers tend to have time-consuming working conditions that are difficult to combine with family life but associated with higher wages (Magnusson and Nermo 2017; cf. Goldin 2014). Women may choose mother-friendly jobs when they become parents. The fact that we find no net motherhood penalty during later time periods speaks against the assumption of mothers choosing flexibility ahead of wages. However, the mother-friendly job hypothesis could still serve as an explanation for wage differences between mothers and fathers. Moreover, there is a gender difference among parents in the tendency to attain a supervisory position, which is a major cause of gender differences in wages. While men’s chances of holding a supervisory position increase when they become parents, women’s chances are unaffected (Bygren and Gähler 2012). When we control for the number of subordinates, this contributes to removing the larger parenthood premium for men. Finally, employers may perceive fathers as more stable and committed workers, as compared to mothers, which makes employers more likely to invest in fathers (Acker 1990; Budig and Hodges 2010). Our data give us no possibility of observing employers’ behavior, but a recent Swedish study based on a correspondence audit with gender and parenthood randomized to job applications, does not support the notion that Swedish employers treat mothers and fathers differently in hiring situations, at least not in 2013 to 2015 (Bygren et al. 2017).

Overall, we show a rather stable picture of the association between parenthood and wage rank in the Swedish labor market. This study has made important contributions to the field by showing the development of this association over several decades, covering a transformative period for Swedish family policy and labor market structure and composition. Moreover, we have made use of longitudinal data and fixed-effects models to remove any impact of time-invariant selection for wage outcomes, thus approaching a causal interpretation of the association between parenthood and wages that is not present in most previous studies in the field. Further study is warranted to chisel out the precise mechanisms underlying the remarkable stability in the (gendered) parenthood wage premium during this period of radical institutional and normative change.

Supplementary Material

Supplementary material is available at Social Forces online.
About the Authors

Magnus Bygren is a professor of sociology at the department of sociology, Stockholm University and affiliated to the Institute for Future Studies, Stockholm. His research centers on understanding ethnic, gender, and class inequality processes. His work has appeared in European Sociological Review, Journal of Marriage and Family, Social Forces, and Sociology of Education.

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