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The gender gap in workplace authority in Sweden 1968–2000

– a family affair?

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The gender gap in workplace authority in Sweden 1968–2000 – a family affair?

by

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Abstract

We assess whether the gender gap in authority in Sweden has changed during the period 1968–2000, and investigate to what extent family factors are responsible for this gap. We find that the gap has narrowed modestly during this period, and identify the life-event of parenthood as a major cause of the gap. When men become fathers, they gain authority; when women become mothers, they do not. Our fixed effects panel estimates of the effects of family factors deviate from the cross-sectional estimates, suggesting that unobserved individual heterogeneity – routinely neglected in this line of research – matters.

Keywords: Workplace authority, gender gap, work-family balance, Sweden
JEL-codes: J12; J13; J16

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1 Introduction

The Swedish society has become well-known for its high degree of equality between men and women. Sweden consistently ranks in the top in international comparisons of the degree of gender equality (e.g. United Nations Development Programme 2005). Swedish women have a higher labor market participation rate than women in most other countries (Jaumotte 2003) and the wage gap between working men and women is low in an international perspective (Blau & Kahn 2003; Waldfogel 1998). Furthermore, Swedish women are not economically dependent on their husbands to the same extent as are their counterparts in other countries (Evertsson & Nermo 2004; Hobson 1990; Sørensen 1994).

Yet, there remain substantial inequalities between Swedish men and women, and one of the more pronounced is the gender gap in *workplace authority*. Few women are found at the top of business firms (Henrekson 2004), and female employees less often have managerial positions and subordinates (Hultin 1998; Mueller, Kuruvilla & Iverson 1994). This may not be surprising, given the fact that the gender gap in authority seems to be a universal phenomenon. Surprising, however, is the finding that the gap seems to be relatively large in Sweden compared to that in many other developed countries (Baxter & Wright 2000; Mandel & Semyonov 2006; Rosenfeld, Van Buren & Kalleberg 1998; Wright, Baxter & Birkelund 1995), even though it has been an explicit policy goal that more women should fill managerial positions (e.g., Ministry of Industry, Employment and Communications 2003).

An equal gender distribution in authority in the workplace is important not only because authority positions reward the individual with prestige, power, autonomy, and status; gender differences in authority attainment also account for a part of the gender wage gap (England et al. 1994; Halaby & Weaklim 1993; McGuire & Reskin 1993). Moreover, authority is of particular pertinence for another reason: ”(B)ecause of the real power associated with positions in authority hierarchies, gender inequalities in authority may constitute one of the key mechanisms that sustain gender inequalities in workplace outcomes. The under-representation of women in positions of authority, especially high levels of management, is not simply an *instance* of gender inequality; it is probably a significant *cause* of gender inequality” (Wright et al. 1995:407).

In other words, more women on positions of authority could contribute to gender equality also on lower levels in the work organization.[1] Another argument for increased gender equality in workplace authority, though less often mentioned, is that if women – for different reasons – choose not to strive for power and authority, or if they are systematically debarred from these positions, the society as a whole suffers; it misses the unexploited talent that these women constitute.

Studies on the gender gap in authority have exclusively been based on cross-sectional (and predominantly U.S.) data. Hence our knowledge about this phenomenon is limited. Although studies have uniformly shown that women have less authority than men do, we do not know much about whether the size of the gap has changed over time or about the mechanisms governing the process by which individuals move into and out of authority positions.

In this paper, we will contribute to research aiming at describing and explaining the gender gap in authority in the following ways. We will assess whether the gender gap in authority in Sweden has changed during the period 1968–2000, with and without control for relevant variables. We will further examine the mechanisms of allocation responsible for this gap by following individuals over time to examine what the correlates of change in authority are. We will focus on a hitherto surprisingly neglected aspect of the process behind the gap; the role of family-related responsibilities of men and women. It is a well-known fact that women in all societies take a larger responsibility for children and household chores. Relatively little is still known, however, about the impact of these responsibilities on women's chances of gaining authority in the labor market. The results of the few studies that have been conducted are inconclusive. As pointed out in a recent review by Smith (2002), additional research in this area is sorely needed in order to specify the relationship between these responsibilities and the attainment of workplace authority.

[1] Studies have shown that the more male-dominated the managerial staff, the larger the gender wage gap among subordinate men and women (Cohen & Huffman 2007; Hultin & Szulkin 2003) and the lower the chances of promotion for subordinate women (Cohen, Broschak & Haveman 1998).

2 Literature review

2.1 Workplace authority

Researchers have defined and operationalized the concept workplace authority in numerous ways, for example as the right to hire and fire, set the rate of pay, supervise the work of subordinates, the direct participation in policy decisions in the workplace, and formal hierarchical position (McGuire & Reskin 1993; Wolf & Fligstein 1979; Wright et al. 1995).[2] A common characteristic for these dimensions, however, seems to involve ”legitimated control over the work process of others” (Wolf & Fligstein 1979:236). Regardless of definition, previous studies have uniformly shown that women are less likely to exert workplace authority than men are (Smith 2002). Moreover, when women are in positions of authority they are located at lower levels of management, i.e. they exert less authority than male managers (Hopcroft 1996; Jacobs 1992; Jaffee 1989). In accordance with most former studies the empirical part of this paper is based on measures of whether the respondent has subordinates (supervisory authority) and, if so, the number of subordinates (span of control).[3]

2.2 Mechanisms behind the gender gap in workplace authority

2.2.1 Family responsibility and human capital

A large amount of evidence indicates that there are gender-specific ways of reconciling family and career-related demands (see, e.g., Blossfeld & Drobnic 2001; Hakim 2000). In Sweden, gender differences in time devoted to

[2] In his review Smith (2002:511) defines four types of workplace authority (not including *ownership* which may be viewed as “the ultimate form of authority”): i) *Sanctioning authority or span of responsibility*, i.e. “the ability to influence the pay or promotions of others”, ii) *Decision-making or managerial authority*, i.e. authority which “relates to organizational policy decisions, control over products, services, budgets or purchases”, iii) *Hierarchical authority position*, i.e. “an individual’s formal location within the structure of organizational hierarchies” and iv) *Supervisory authority*, i.e. “whether an individual ‘*supervises anyone on the job*’” and, if so, *span of control*, i.e. “the number of people under direct supervision”. It should be noted that the different types of authority are not mutually exclusive.

[3] Unless otherwise stated this type of authority is what we henceforth refer to as “authority” and “workplace authority”. When we refer to previous studies we will explicitly mention if they apply to other types of authority (only).

household tasks seem to be triggered by cohabitation and childbearing: it is relatively small for single persons, but grows substantially larger among childless cohabiting couples and further increases when they have children, even when spouses devote identical time to paid work (Flood & Gråsjö 1997). In addition, women dedicate much more time to child care (The National Social Insurance Board 2003; Statistics Sweden 2003). As the children grow older, or leave the nest, this gender difference in household work remains largely unchanged (Ahrne & Roman 1997; Hörnqvist 1997). Fathers normally work full-time irrespectively of the age of their children, whereas mothers adjust their labor force attachments to the demands of parenthood (cf. Gornick & Meyers 2003; Sundström 1997). A consequence of these differences is that Swedish women (mothers) commonly have less work experience than men (Edin & Richardson 2002; Hultin 1998; Mueller et al. 1994). Because prospects to reach positions with power, influence, and authority are conditioned by the individual's human capital, there are reasons to believe that women's chances to reach these positions deteriorate when they have children. Accordingly, human capital differences normally explain some of the gender gap in authority. Studies however consistently reveal a remaining and pronounced authority gap between women and men even after controlling for various indicators on human capital (Baxter & Wright 2000 [Australia, Sweden and USA]; Huffman & Cohen 2004 [USA]; Hultin 1998 [Sweden]; Jaffee 1989 [USA]; Kraus & Yonay 2000 [Israel]; Mitra, 2003 [USA]; Wright et al. 1995 [seven Western countries]).[4] Some studies also reveal that women receive smaller authority returns for the same formal merit as men (McGuire & Reskin 1993; also see review by Smith 2002),[5] but two Swedish studies do not find this (Hultin 1998; Mueller et al. 1994).[6]

Over and above human capital differences, family responsibilities might affect women's authority attainment also in other ways. A net wage penalty associated with motherhood is a fairly stable empirical regularity across institutional settings (see, e.g., Budig & England 2001; Waldfogel 1998 [the U.S.];

[4] In Baxter and Wright (2000) authority is measured by formal position only whereas Kraus and Yonay (2000) and Wright et al. (1995) add measures on sanctioning and decision-making authority.

[5] McGuire and Reskin (1993) use an index containing indicators on different types of authority, including supervisory authority.

[6] Also see Hopcroft (1996) and Jaffee (1989) for American studies (based on the same data set). Both studies use a variety of indicators on workplace authority, among them indicators on supervisory authority.

Joshi, Paci & Waldfogel 1999 [the U.K.]; Albrecht et al. 1999 [Sweden]). The results for authority outcomes are less unequivocal. Hultin (1998) finds that Swedish women with children in the household have lower chances of exerting workplace authority compared to childless women, but no such child penalty is found for men. Rosenfeld et al. (1998) (using a nine-country dataset) and Wolf and Fliegstein (1979) (using two U.S. samples) report equal results for women, but in their studies men with children rather benefit. In addition, a British study finds that among top managers, women are much more often childless compared to men (50 vs. 7 percent), suggesting that men can usually have a career as well as a family whereas women appear to choose between the two (Wyatt & Langridge 1996). In contrast to these results, more recent U.S. studies find no effects of children on authority, neither for men nor women (see Hopcroft 1996; Huffman & Cohen 2004; Mitra 2003).

Comparing the authority outcomes of cohabiting/married working men and women to those of single working men and women, results are even more mixed. Hultin (1998) finds that Swedish men who are cohabiting or married experience an authority premium, but that civil status is unrelated to workplace authority among Swedish women. Rosenfeld et al. (1998) find a marriage/cohabitation authority premium, but that it is given to both men and women. By contrast, Huffman and Cohen (2004) and Mitra (2003) (both using U.S. data) find that civil status is unrelated to authority for men and women alike.[7]

Controlling for human capital indicators, why would family conditions affect men's and women's labor market careers differently? According to Becker (1985), the remaining precedence for men in the labor market may be due to gender differences in work *effort*. Women's greater household work and child care effort demand a large amount of time and energy. As a consequence, women may seek less demanding jobs, to be able to meet family and work obligations, and/or put less effort into market work. In contrast, others argue that women's main responsibility for household chores and child care may rather signal (than actually cause) less work commitment among employers (see below) whereas it may point to more work commitment and stability for male employees (e.g., Bielby & Baron 1986).

[7] In contrast to these conflicting results on the role of cohabitation/marriage for the gender gap in authority, a male marriage wage premium has been documented, in Sweden (Richardson 2000) and elsewhere (Ribar 2004).

2.2.2 Gender discrimination

Most studies, then, show that the gender gap in workplace authority can not entirely be explained by differences in observed individual traits between men and women. Explanations of the remaining gap have therefore been sought on the demand side. If employers' decisions to hire or promote, for some reason, are gender biased in favor of men, they will disproportionately end up in positions of authority. Given that men and women are equally merited for these positions, gender bias in employer decisions are due to gender discrimination.

Taste discrimination refers to employers preferring men for these positions irrespectively of them being more qualified than women, which would imply that employers are willing to pay for not having women in authority positions. *Statistical discrimination* refers to a decision rule when information about potential employees' productivity is limited. Employers may then use easily observable characteristics, such as sex, as proxies for promotion suitability. For example, if employers know that women, on average, are more constrained by family responsibilities than are men, they should act on this information in their promotion decisions. High turnover rates in positions of authority can also be assumed to be especially costly for an employer, and to avoid these costs, they may avoid to hire women "at risk" of having children, or mothers of small children, for these positions.

Direct evidence on gender discrimination as a mechanism behind the gender gap in workplace authority is rare. Some scholars argue along the line of proof by residual, i.e., that any unexplained gender difference after controls for indicators on other possible predictors, must, at least partly, be due to discrimination (see, for example, Hultin 1998; Smith 2002 for discussions). The residual may, however, be due to measurement error and/or unobserved gender differences that are relevant to the exertion of authority, observed by employers but unobserved in the data used by researchers.

Setting the difficulties with measuring discrimination aside, however, according to the taste discrimination hypothesis we would expect women to exert less workplace authority than men irrespective of their family situation. This perspective gives us no reason to expect that childless women will be treated differently by employers than women with children.[8] According to the

[8] This assertion hinges on the assumption that employers do not have a taste for discriminating women with children. One could of course make an argument that because employers think a mother's place is not in authority positions, they discriminate against mothers and rather choose men, and women who are not yet mothers, for these positions. Although we would not like to

statistical discrimination hypothesis, on the other hand, we would expect employers to treat married and cohabiting women and mothers differently from the way they treat single women, women without children and men. Given the gendered division of household labor and child care, rational employers could assume married and cohabiting women and mothers to be less productive than other women and men and hence be less willing to award them with positions of authority.

2.3 The gender gap in workplace authority – development over time

We know very little about the changes in the workplace authority gender gap over time, in Sweden and internationally. This is due to the fact that previous studies are almost uniformly based on non-repeated cross-sectional data.[9] What, then, could we expect of this development? Since the 1970s gender inequality in a number of domains – education, employment rates, work hours, the gender wage gap, economic dependency within couples, household work hours, and use of parental leave – has decreased in Sweden (Bygren & Duvander 2004; Bygren, Gähler & Nermo 2004; le Grand, Szulkin & Tåhlin 2001).[10] Although major inequalities remain, Swedish women of today should meet the labor market with more assets and possibilities relative to men than women did just a couple of decades ago. Moreover, the signalling effect of household variables should have become noisier over time, i.e. statistical discrimination might not be as prevailing as it used to be. Hence we would expect the aggregate gender gap in workplace authority to have decreased over time.

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discard this hypothesis altogether, we find it hard to believe that the very same employers who discriminate against mothers would treat men and women without children equally. Of course, one could imagine both kinds of taste discrimination operating, but then again, we find the reasons for employers exerting discrimination against all women more compelling, not least because sex is more easily observed compared to parenthood.

[9] A recent study by Meyersson Milgrom and Petersen (2006) shows, however, that the gender gap in “rank”, i.e. a job’s level of difficulty and responsibility among other things, narrowed in Sweden during the period 1970–1990 (also see Jacobs 1992 and Reskin & Padavic 1994 for American studies).

[10] During the period, Sweden has implemented a number of family-related policies aiming at replacing the male-breadwinner model with that of a dual-earner model. Important elements driving this transformation were the introduction of individual taxation of married couples in 1971, the introduction of parental-leave benefits in 1974, and the expansion of public child care during the 1970s and 1980s (Duvander, Ferrarini & Thalberg 2005).

Several researchers assert, however, that the extensive parental-leave program in Sweden is a probable cause of the relatively large gender gap in workplace authority there (see, e.g., Hakim 2000; Mandel & Semyonov 2006; Rosenfeld et al. 1998). In order to have access to the parental-leave benefit women must have and retain an employment relation when they have children. In practice, though, mothers have a relatively weak connection to the labor market in the period following childbirth, due to long periods of absence and/or part-time work. As employers normally prefer to promote (at least) full-time employees for authority positions, many women with small children may effectively be sorted out of the pool of employees considered suitable for filling these kinds of positions. Absence and part-time work among mothers arguably also give rise to statistical discrimination of female employees in fertile ages (Skard & Haavio-Mannila 1984, cited in Rosenfeld et al. 1998; Mandel & Semyononv 2006). According to this view, we would expect the Swedish family policy model to have reinforced, or at least to have upheld, the gender gap in authority over time.

2.4 The contributions of the present study

As stated, data on the historical development of workplace authority is rare. More serious, however, is the fact that longitudinal panel data has not been used in this line of research. As a consequence the issue of unobserved selection effects has routinely been ignored. Selection issues have been extensively dealt with in studies of the male marriage wage premium and the motherhood penalty on wages, but in studies of workplace authority, none has hitherto used panel designs. All of what we know about the attainment of authority is based on data from cross-sections, which seems far from ideal if we want to approach a causal understanding of the authority attainment process. A well-known weakness of cross-sectional data is that the effect of unobserved individual characteristics, which is correlated with observed independent variables, may give rise to omitted variable bias in the estimated coefficients. For example, a motherhood penalty on authority in a cross-section may have arisen because employers have used traits unobserved in the data to sort women into positions of authority. If these traits are correlated with the probability of getting a child, omitted variable bias in the estimated effect of motherhood would arise. In this study we have access to an individual-level panel data set covering a period of over thirty years. Because many individuals in this data set have been observed at more than one point in time, we are in our estimations able to account for time-invariant characteristics sorting individuals into authority positions (see, e.g., Halaby 2004; Petersen 2004). In this way, we will be able to considerably extend present knowledge about the role of selection processes generating the gender gap in authority.

3 Data and analytical strategy

The basic research questions we want to answer are: How has the gross gender gap in authority changed over time during the past decades? What is the role of family-related responsibilities in explaining the gender gap in authority? To answer these questions, we use cross-sectional and panel data from the Swedish Level of Living Survey (LNU) collected in 1968, 1974, 1981, 1991 and 2000.[11] The respondents in each cross-section are a 0.1 percent random sample of the Swedish adult population aged 18–75 years. The unique strength of these data is that they include detailed and longitudinally identical measures of labor market related variables and family related variables. We will first model authority for the five cross-sections separately. Of particular interest here, besides describing the gross gender gap, is whether and in what direction the gender gap in authority has changed between 1968 and 2000, whether the correlates of authority have changed during the period of study, and to what extent it is possible to statistically explain the gender gap using our independent variables. In a second part of the empirical analysis, we will make use of the panel information, and model changes in authority as a function of changes in the independent variables.

4 Variables

To enable our empirical analysis, we excluded respondents who did not work at least one hour a week. Additionally, we excluded a small number of observations with internal non-responses on relevant questions. After these selections, the samples consisted of 2,816 respondents in 1968, 3,213 respondents

[11] The response rates in the three cross-sections range from 90.8 percent in 1968 to 76.6 percent in 2000. See Gähler (2004) and Jonsson and Mills (2001) for more thorough descriptions of the Swedish Level of Living Surveys.

in 1974, 3,438 respondents in 1981, 3,333 respondents in 1991, and 2,998 respondents in 2000 (which equals 15,798 observations in total). Due to the panel structure, the data contain information from 7,170 unique respondents, of which 63 percent participated in more than one cross-section of the survey. Below, we describe the variables used in the analyses. Descriptive statistics are presented in *Table 1*.

In each survey, all employed respondents were asked identical questions about whether they currently had any supervisory tasks. Respondents with supervisory tasks were also asked how many subordinates they had. In 1968, 1974 and 1981, the number of people supervised was collapsed into discrete groups at the time of the interview, whereas the exact reported number was used in 1991 and 2000. To achieve consistency over time, we constructed identical measures of authority across the panel cross-sections, using the following four categories: 0, 1–5, 6–10, and 11+ subordinates. For the independent variable *woman*, we assigned the value 0 for males and 1 for females. We constructed three dummy variables indicating various dimensions of the *respondent's children*. The first indicates the presence of one or more pre-school aged children (0–6 years) in the household. The second indicates the presence of one or more children between the ages 7–20 in the household.[12] Whether an individual exerts workplace authority or not is not only due to the present situation but also historical conditions, i.e. former family and household responsibilities. Given the present situation, former family responsibilities may have lasting effects on career outcomes. This has often been neglected in previous research in the field. With our third child-related variable we therefore indicate whether the respondent has grown up children (20+) that live in or outside the respondent's household.[13] We used a dichotomous measure of marital status with value 1 if respondent is *cohabiting or married*, and value 0 if not. Further, we included a measure of *household work hours per week*, i.e. the sum of the number of hours per week the respondent reports to spend on food purchasing, cooking, doing dishes,

[12] 7-18 years in the 1991 and 2000 surveys.

[13] 18+ years in the 1991 and 2000 surveys. In a previous version of this paper, we also included a fourth child variable referring to whether the respondent has any children below the age of 20 that do not live in the respondent's household, indicating, in most cases, that the other biological parent or any other adult has the main responsibility for the child. This variable uniformly turned out to be insignificantly related to the dependent variable, why we excluded it from our models.

laundering, and cleaning.[14] In the 1968, 1974 and 1981 surveys, the questions were not asked to single respondents. We therefore only use this variable for the years 1991 and 2000. To study the impact of labor market conditions we used four different indicators. We measured *education* in years of formal education. It refers to the question: "For how many years have you been in full-time school and vocational education (from first grade and onwards)?".

Table 1 Descriptive statistics for cross-sectional data (individuals working at least one hour a week)

	Men					*Women*				
Year	1968	1974	1981	1991	2000	1968	1974	1981	1991	2000
Number of subordinates, percent										
0	71.9	70.5	69.7	69.3	70.4	85.2	85.6	81.0	81.6	80.3
1-5 subordinates	11.7	13.5	13.0	13.0	13.1	9.5	8.5	11.1	10.4	11.7
6-10 subordinates	5.5	5.8	7.3	7.1	6.3	2.2	2.4	3.8	3.7	3.7
11+ subordinates	10.9	10.2	9.9	10.5	10.2	3.1	3.5	4.0	4.3	4.3
R's children, percent										
Child 0-6 yrs. in HH	24.8	25.1	19.3		19.4	14.4	20.4	20.0	20.7	19.2
Child 7-20 yrs. in HH[a]	34.4	33.6	36.3	31.7	31.5	34.3	37.5	40.4	35.0	36.5
Child 20+ yrs. in/out of HH[b]	24.3	23.3	23.2	26.8	28.0	26.6	25.7	28.5	32.2	35.9
Cohabiting/married, percent	73.6	74.3	71.9	70.7	70.9	63.1	72.2	72.6	72.4	74.4
Hours of paid work per week, mean	42.8	39.7	38.9	39.1	39.1	34.4	31.8	31.5	33.8	35.1
Education in years, mean	8.7	9.8	10.7	11.7	12.6	8.7	9.6	10.3	11.5	12.7
Work experience in years, mean	22.8	21.5	20.7	20.0	20.1	15.1	14.3	15.6	16.7	19.2
Seniority in years, mean	10.1	10.0	9.9	10.3	10.3	6.4	6.8	7.8	9.5	10.3
Household work hours per week, mean	n.a.	n.a	n.a.	7.3	7.3	n.a	n.a	n.a.	17.6	14.5
Private sector, percent	78.6	74.4	70.2	70.7	71.8	54.5	49.1	40.3	39.1	39.1
Number of observations	1750	1808	1801	1663	1531	1066	1405	1637	1670	1467
Share of cross-section	62.1	56.3	52.4	49.9	51.1	37.9	43.7	47.6	50.1	48.9

[a] 7-18 years in the 1991 and 2000 surveys.

[b] 18+ years in the 1991 and 2000 surveys.

[14] Whereas marital status and (number of) children in the household are only proxies for family and household obligations, household work hours per week is a direct measure on work effort at home and constraints to put energy on market work (alternatively preferences to use energy on family and household tasks). Only few, if any, previous studies include such a direct measure. Thus our study – unlike previous studies – can account for the fact that, e.g., women devote more time than men to household chores even when children are older or have left the household.

We measured work experience as the respondent's *total number of years in employment*. It refers to the answer to the question: ”How many years altogether have you spent in gainful employment?”. We measured *seniority* as the number of years that had passed since the respondent was first employed by the current employer. The measure of *hours of paid work per week* refers to the question “How many hours per week do you regularly work?” As can be noted, the gender difference in labor market variables has declined markedly during the period. Still, however, there is a clear gender gap, in men's favor, between fathers and mothers (not shown). Finally, we used a variable of sector of employment, *private*, with value 1 if the individual is employed by a privately owned firm, and 0 otherwise.

5 Results

5.1 Descriptives of the gross gender gap over time and age

In *Figure 1*, we display the odds of finding a woman in a particular position with (or without) authority, relative to the share of women working. A number of patterns are noteworthy here. First, consistent with all other studies in this area of research, we find that women less often than men execute supervisory tasks. Second, the gender gap in authority becomes more and more pronounced the higher up in the authority hierarchy one gets. This finding is also consistent with other studies (e.g., Hopcroft 1996; Jacobs 1992; Jaffee 1989). The odds of a woman having a relatively small number of subordinates (1-5) is, averaging over time, around .9. For those with 6-10 subordinates, the odds drops to an average level of around .6, and for those with 11+ subordinates, it drops further to an average of .5. Third, during the period, the odds of women exerting authority increased substantially on all levels between 1974 and 1981 and, to a smaller extent, between 1991 and 2000. Hence during the 1970s there was a clear change in the direction towards more of gender equality in authority outcomes, but this development levelled out in the beginning of the 1980s. The odds presented in *Figure 1* hold constant the share of women working. However, one might also ask what the probability is that a position of authority is held by a woman in a given year. From *Table 1* it is possible to calculate these proportions, and over the period it has increased from 24 percent in 1968 to 39 percent in 2000. This is a combined effect of two changes that mainly took place in the 1970s: that of women increasing their relative share of the labor force and that of a modest gender equalization of chances of getting into authority positions (displayed in *Figure 1*).

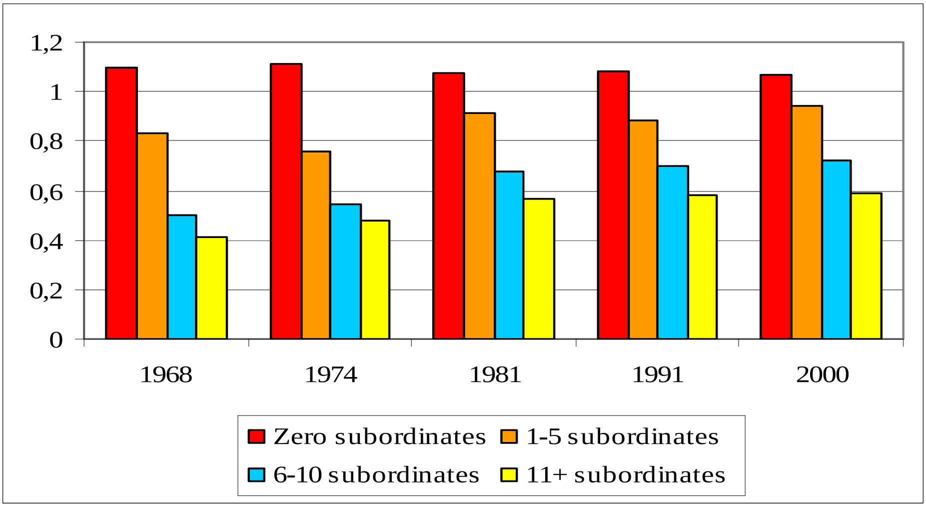

Figure 1 The odds[a] of a working woman being located in positions with different amounts of authority, 1968, 1974, 1981, 1991, and 2000

[a] Share of women in position i among all individuals in position i/share of women working >1 hour a week among all individuals working >1 hour a week

In *Figure 2*, we display the proportions having authority for different age groups, separately for men and women. One can note that men and women seem to start out at the same low level in their 20s, but that men have a much steeper “career curve”, peaking in their late 40s, declining thereafter. The women’s chances of having authority increase until they reach their 30s, but thereafter remain at a more or less flat level, significantly below that of men. We also broke down the figure by survey year. We found that, although the proportion with subordinates differs by year (for women), the age profile by gender only differs marginally. Hence men’s and women’s career development does not differ a great deal up until they reach their 30s, when they separate ways. Coincidentally, this point in life is also the age at which many men and women are in the stage of forming a family. The question that remains to be answered is whether these gender-specific career profiles are associated with family-related events. We next turn our attention to this issue.

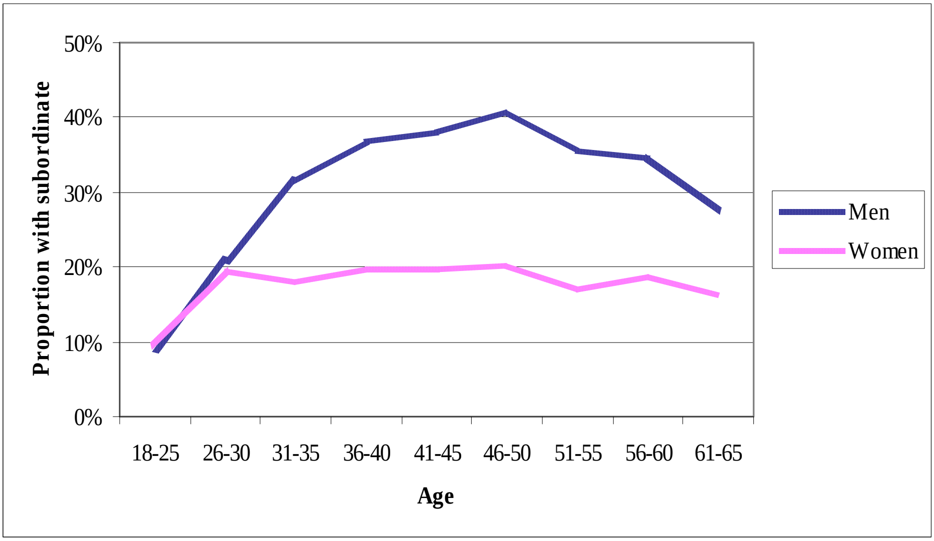

Figure 2 Age profiles of proportions having at least one subordinate, by gender (pooled 1968-2000)

5.2 Cross-sectional regressions

We conducted a number of alternative analyses, using various cutting points and categorizations of the measure of authority. Independently of authority measure and statistical method used, the results from these analyses were very similar.[15] For reasons of simplicity we therefore decided to use a dichotomized measure of workplace authority, i.e. whether the respondent has any subordinates or not, and estimated the following linear probability model using this measure, for each survey year, and for men and women separately:

$$P(A_i) = \alpha + \beta_j X_{ij} + \varepsilon_i \tag{1}$$

[15] Using our four-category measure, we have estimated multinomial logit models, ordinal logit models, and linear regression (OLS) models. For 1991 and 2000, we had a continuous measure available. Using this measure, we estimated tobit models, and linear regression models on the logged and untransformed number of subordinates. We further experimented with different cut-points for the dichotomous measure.

A_i equals 1 if individual *i* has subordinates, otherwise zero. β_j is the coefficient vector associated with X_{ij}, and ε_i is the random error term for individual *i*. If the coefficient vector β_j is multiplied by 100, we get the percentage point change associated with a unit increase in the independent variable. The linear probability model has two well-known weaknesses: it may give probability predictions outside of the interval 0 to 1, and the error term violates the linear regression assumption of homoscedasticity. As our main purpose with the regressions is to estimate the mean effects of the family-related variables, we do not consider the first weakness a serious one. The second weakness has potentially more serious implications, as standard errors thereby get biased and tests of significance will be invalid. To avoid this problem, we use Huber/White sandwich estimators of variance to obtain robust standard errors. These are valid in the presence of any kind of heteroscedasticity, and yield valid tests of significance. To have meaningful intercepts, we centered all continuous variables on their grand mean in the survey year. For each survey year, we introduce the family-related variables in a first model. In a second model we add education and our labor market variables. For some cross-sections, we have access to information on household work hours, and we then add this variable to the second model.

Table 2 Linear probability model estimates of authority on independent variables, and gender gap decomposition, by survey year[a]

	1968				1974				1981			
	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women	Men	Women
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Child in HH 0-6	0.002	-0.025	0.011	0.003	-0.014	-0.043	-0.017	-0.021	-0.010	0.010	0.001	0.006
	(0.030)	(0.032)	(0.029)	(0.030)	(0.029)	(0.025)	(0.028)	(0.024)	(0.031)	(0.026)	(0.029)	(0.024)
Child in HH 7-20	0.069**	-0.011	0.024	0.015	0.131**	-0.021	0.082**	0.019	0.094**	-0.007	0.035	0.001
	(0.026)	(0.024)	(0.026)	(0.024)	(0.026)	(0.020)	(0.026)	(0.021)	(0.026)	(0.020)	(0.026)	(0.021)
Child 20+	0.035	-0.055*	0.016	-0.020	0.082**	-0.037	0.075*	-0.002	0.051	0.002	-0.001	0.019
	(0.029)	(0.024)	(0.033)	(0.024)	(0.030)	(0.023)	(0.034)	(0.025)	(0.029)	(0.023)	(0.035)	(0.025)
Cohabiting/married	0.164**	-0.013	0.094**	0.007	0.114**	0.018	0.055*	0.036	0.128**	-0.045	0.055*	-0.037
	(0.027)	(0.025)	(0.027)	(0.025)	(0.028)	(0.024)	(0.027)	(0.023)	(0.027)	(0.024)	(0.025)	(0.023)
Education yrs			0.049**	0.032**			0.040**	0.028**			0.037**	0.030**
			(0.004)	(0.005)			(0.004)	(0.004)			(0.004)	(0.004)
Hrs paid work/wk (/10)			0.056*	0.047**			0.071**	0.043**			0.086**	0.035**
			(0.023)	(0.010)			(0.020)	(0.008)			(0.016)	(0.009)
Seniority yrs (/10)			0.061**	0.041*			0.049**	0.034*			0.088**	0.071**
			(0.011)	(0.019)			(0.011)	(0.015)			(0.014)	(0.018)
Work experience yrs			0.016**†	0.012**†			0.021**†	0.004†			0.017**†	0.007
			(0.003)	(0.003)			(0.003)	(0.003)			(0.003)	(0.003)
Work experience squared (/100)			-0.027**†	-0.020**†			-0.038**†	-0.004†			-0.025**†	-0.012
			(0.006)	(0.007)			(0.006)	(0.007)			(0.006)	(0.008)
Private sector			-0.030	-0.044*			-0.064*	-0.056**			-0.034	-0.074**
			(0.028)	(0.021)			(0.025)	(0.018)			(0.024)	(0.018)
Intercept	0.127**	0.178**	0.236**	0.257**	0.150**	0.157**	0.280**	0.189**	0.166**	0.223**	0.258**	0.288**
	(0.016)	(0.020)	(0.033)	(0.029)	(0.017)	(0.019)	(0.030)	(0.028)	(0.017)	(0.021)	(0.029)	(0.030)
Adj. R squared	0.04	0.00	0.16	0.12	0.05	0.00	0.17	0.10	0.04	0.00	0.16	0.10
Number of observations	1750	1066	1750	1066	1808	1405	1808	1405	1801	1637	1801	1637
Decomposition of the gross gender gap[b]												
Endowments (E)	-0.003		0.070**		0.000		0.054**		0.000		0.031*	
	(0.004)		(0.015)		(0.002)		(0.012)		(0.002)		(0.012)	
Coefficients/Unexplained (C)	0.116**		0.030		0.156**		0.064		0.120**		0.006	
	(0.016)		(0.026)		(0.014)		(0.024)		(0.015)		(0.022)	
Interaction EC	0.019		0.031		-0.005		0.033		-0.008		0.075**	
	(0.007)		(0.027)		(0.004)		(0.023)		(0.004)		(0.021)	

** p<0.01, * p<0.05 that coefficient is equal to zero (t-test), † p<0.05 that both terms in the polynomial are equal to zero (F-test).
a Robust standard errors in parentheses. All continuous variables are centered on their grand mean in the survey year.
b The variances/standard errors of the components in the decomposition of the gap are computed according to the method detailed in Jann (2005).

Table 2 Continued [a]

	1991				2000			
	Men	Women	Men	Women	Men	Women	Men	Women
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Child in HH 0-6	-0.005	-0.028	0.027	-0.004	0.069*	-0.001	0.082*	0.016
	(0.032)	(0.024)	(0.031)	(0.025)	(0.034)	(0.029)	(0.034)	(0.030)
Child in HH 7-18	0.127**	0.003	0.039	-0.037	0.065*	0.045	0.013	0.060*
	(0.028)	(0.021)	(0.028)	(0.023)	(0.028)	(0.023)	(0.029)	(0.026)
Child 18+	0.073*	0.017	-0.001	0.005	0.102**	0.004	0.058	0.025
	(0.029)	(0.023)	(0.036)	(0.027)	(0.030)	(0.024)	(0.038)	(0.027)
Cohabiting/married	0.140**	0.027	0.060*	0.013	0.058*	0.014	0.007	0.014
	(0.028)	(0.022)	(0.029)	(0.022)	(0.029)	(0.025)	(0.029)	(0.025)
Education yrs			0.038**	0.033**			0.037**	0.023**
			(0.004)	(0.004)			(0.004)	(0.004)
Hrs paid work/wk (/10)			0.097**	0.029**			0.132**	0.052**
			(0.016)	(0.010)			(0.019)	(0.011)
Seniority yrs (/10)			0.062**	0.039**			0.017	0.006
			(0.014)	(0.013)			(0.015)	(0.014)
Work experience yrs			0.017**†	0.013**†			0.009**†	0.003
			(0.003)	(0.003)			(0.003)	(0.003)
Work experience squared (/100)			-0.023**†	-0.022**†			-0.010†	-0.002
			(0.008)	(0.008)			(0.007)	(0.008)
Private sector			0.054*	-0.047**			0.047	0.060**
			(0.024)	(0.018)			(0.027)	(0.021)
Household work hrs/wk (/10)			-0.045*	-0.002			-0.042	-0.028*
			(0.020)	(0.009)			(0.024)	(0.013)
Intercept	0.149**	0.164**	0.180**	0.253**	0.193**	0.169**	0.209**	0.155**
	(0.016)	(0.019)	(0.030)	(0.029)	(0.019)	(0.022)	(0.038)	(0.035)
Adj. R squared	0.06	0.00	0.17	0.09	0.02	0.00	0.11	0.05
Number of observations	1663	1670	1663	1670	1531	1467	1531	1467
Decomposition of the gross gender gap[b]								
Endowments (E)	-0.001		0.017		-0.003		0.053**	
	(0.002)		(0.014)		(0.003)		(0.013)	
Coefficients/Unexplained (C)	0.133**		-0.017		0.112**		0.010	
	(0.015)		(0.028)		(0.016)		(0.025)	
Interaction EC	-0.009*		0.122**		-0.010*		0.035	
	(0.004)		(0.029)		(0.004)		(0.024)	

** $p<0.01$, * $p<0.05$ that coefficient is equal to zero (t-test), † $p<0.05$ that both terms in the polynomial are equal to zero (F-test).

[a] Robust standard errors in parentheses. All continuous variables are centered on their grand mean in the survey year.

[b] The variances/standard errors of the components in the decomposition of the gap are computed according to the method detailed in Jann (2005).

To analyze closer how much of the gender gap in authority that our variables are responsible for, we also conducted Blinder-Oaxaca decompositions for each cross-section. This procedure decomposes the gap into an "explained" part, due to average differences in endowments between men and women, and an "unexplained" part, due to differences in returns (coefficients, including the constants) to these endowments between men and women, and an interaction between the "explained" and "unexplained" part.[16]

In *Table 2,* column 1 and 2 for each year, we report coefficients for the family-related variables only, for men and women separately. We can see that there are large gender-specific effects of the family situation of individuals.[17] Throughout the period, having (had) children is associated with an authority premium for men (columns 1). This "fatherhood premium" varies somewhat between survey year, but is present in all cross-sections. For women, the corresponding effect of having (had) children is, almost exclusively, insignificantly different from zero (columns 2). We can further see that there is a large cohabitation/marriage premium for men, but that this premium has grown smaller over time; the net difference in having authority between cohabiting men and women has shrunk from 17.7 percentage points in 1968, to 4.4 percentage points in 2000. For women, marital status is generally not significantly associated with workplace authority in any year. It is also interesting to note that the intercepts, which in these models indicate the probability that noncohabiting childless individuals have subordinates, show that in this group the gender gap is very small, and even reversed in some cross-sections.

The decompositions of the gender gap, reported in the lower rows of *Table 2*, indicate that the gap in these models almost entirely can be attributed to the fact that men have higher "returns" to their family situation. Taken

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[16] The gross authority gap $\overline{A}^M - \overline{A}^F$ (superscript M indicating males, F females) can, after some algebra, be decomposed into the following parts,

$$\sum_{j=1}^{J} \beta_j^F (\overline{X}_j^M - \overline{X}_j^F) + (\alpha^M - \alpha^F) + \sum_{j=1}^{J} \overline{X}_j^F (\beta_j^M - \beta_j^F) + \sum_{j=1}^{J} (\beta_j^M - \beta_j^F)(\overline{X}_j^M - \overline{X}_j^F),$$

where the first term is the "explained" part of the gap, the second and third term is the "unexplained" part, and the fourth term is the part of the gap that comes from the interaction between the "explained" and "unexplained" components. The Blinder-Oaxaca decomposition technique is attributed to Blinder (1973) and Oaxaca (1973). See Jones and Kelley (1984) for a graphical description of the equation.

[17] We use the word "effect" for language simplicity reasons. We are of course aware that it is not possible to determine the order of causality between the dependent and independent variables in this setting.

together, the estimated effects indicate large, and to some extent, persistent gender-related effects of family situation on the attainment of authority, and that the gender gap is statistically explained, or even reversed, once family-related variables and gender-specific effects thereof are accounted for.

When we add controls for labor market conditions, and education (columns 3 and 4), the tendency is that the raw fatherhood premiums are attenuated towards zero. Additional analyses (not shown) indicate that differences in education, work experience, and, to some extent, seniority, are responsible for this premium. That is, part of the reason why fathers, and primarily fathers to older children, more often exert workplace authority is that they score higher on education and work attachment.[18] Moreover, the cohabitation/marriage premium for men is also attenuated once the controls are introduced, indicating that more work attachment and higher education among married and cohabiting men seems to be a mechanism partly generating this effect.

The decompositions of the gender gaps in the full models essentially tell the same story. Once the controls have been included, the part of the gender gap that is unexplained/generated by differences in coefficients shrinks considerably in all cross-sections. Instead, the parts of the gap that can be attributed to endowments, and the interactions between endowments and coefficients, increase in size. This indicates that the gap, apparently generated by male family premiums, in fact is generated by married or cohabiting fathers having higher levels of labor market and education endowments, and also reaping higher returns to these endowments.

When we add household work to the last models for the 1991 and 2000 surveys, we find that it has the expected negative sign. However, adding household work does not significantly alter the other estimated coefficients of family-related variables, suggesting that these effects are not mediated by gender-specific differences in household work (results not shown).

More generally, a clear impression of the indicators of model fit is that it has become increasingly more difficult to predict authority with observables. More associations turn insignificant and the R square values get lower over

[18] The overall gender differences in these variables have decreased over time (see *Table 1*). However when mothers and fathers are compared, the gender differences in these variables tend to be higher. For example, fathers have more work experience than mothers. This is true in 1968 (27 years compared to 17 years) as well as in 2000 (25 years compared to 22 years).

time, implying that factors other than those observed in the dataset have become more important for the attainment of authority over time.

To sum up, we find that the gender gap in authority is mainly a gap existing among men and women who have (had) children and/or are cohabiting/married. Whereas the gap between men and women with children is stable over time, the gender gap in authority within the group of married/cohabitants has declined. A substantial part of the gender gap in authority among parents, and the once large gender gap in authority among married/cohabitants is mainly due to differences in education and labor market variables, as well as differences in the returns to these characteristics. Throughout the period, the gap was absent or even reversed among single men and women.

5.3 Panel analysis

In *Table 3* we present results from regressions using data aggregated over the whole observation period 1968–2000. As a benchmark, we estimate the following "total" models pooled over time, for men and women separately.

$$P(A_{it}) = \alpha + \beta_j X_{ijt} + \gamma_t + \varepsilon_{it} \tag{2}$$

This model is identical to the models estimated on the cross-sections, but the data here may contain the same individuals *i* across *t*, and we have added a period dummy γ_t for each survey year. For the fixed effects models, we add to these models a dummy δ_i for each individual *i*.

$$P(A_{it}) = \alpha + \beta_j X_{ijt} + \gamma_t + \delta_i + \varepsilon_{it} \tag{3}$$

In effect, we thereby capture all observed and unobserved individual attributes 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 the probability to be transferred into a position of authority (because the models are estimated by sex). Our point of estimating this model is to show what happens to the dependent variable when individuals experience a family-related change, for men and women separately.[19] What one needs to be aware of is that the fixed effects model uses only intra-individual variation to estimate effects, which means that these are only representative for the subselection(s) of individuals exhibiting such variation. Out of the sample of 7,170 individuals, only 902 men and 593 women switch between having and not having authority during the period. A minority of the sample thus contributes to the estimation of these effects. This "selection criterion" also has the unintended consequence of making the group age with the panel. Its mean age is 32 in 1968 and 48 in 2000. This is not a serious problem as our main purpose is to estimate unbiased effects of family-related changes. Note that the age effects are picked up by the time dummy variables γ_t in the fixed effects regressions.

The total pooled linear effects models reported in columns 1–2 and 5–6 in *Table 3* can be seen as a summary of what has been shown in the previous analysis.[20] Hence estimates from this model show that women who have (had) children/are cohabiting/are married do not significantly differ in terms of authority from women who do not belong to these categories. Men, on the other hand, differ a great deal in their degree of authority depending on their family situation. Cohabiting and married men and fathers with children aged 7+ are much more likely to exert authority than are single and childless men. The sizes of the differences are non-trivial, with cohabiting men having a 12 percentage point "authority premium" compared with noncohabiting men, and fathers of children above the age of six having a 7–10 percentage point "authority premium" compared to childless men. As for the cross-sectional regressions, a substantial part of these premiums can be statistically explained with education and labor market related controls.

Results from the fixed effects model, reported in columns 3–4 for men and 7–8 for women, partly underscore the findings from the total models but they also deviate from these findings in important ways, which suggests that the total model effects are biased due to unobserved individual heterogeneity.

[19] It should be underscored that the fixed effects model does not eliminate the problem with the order of causality. The reason is that independent variables might not be exogenous. Hence we can not rule out a reverse causality. For example, individuals may change their family situation as a consequence of their labor market position (authority or no authority) rather than the other way around.

[20] Because individuals can be observed repeatedly, the standard errors are adjusted for intra-individual correlation, using Stata's cluster option. Still, the standard errors are reduced in size due to the increased number of individuals contributing to the estimation of effects.

Similar to what was found in the total model, it seems that fathers gain authority compared to non-fathers. The estimated gains of fatherhood are even higher compared to the estimates from the total model, and the effect appears already when children are small (0-6 years old). Moreover, contrary to what was found in the total model, motherhood is accompanied by an increase of authority, but only of a lagged kind: it kicks in when the children are above the age of 20. When we add the controls the motherhood effect turns insignificant, and the fatherhood effects tend to decrease substantially, but stay significantly above zero; implying that much of the fatherhood effect, and all of the motherhood effect, can be attributed to an increase in labor market related inputs.

We find that in the fixed effects model the cohabitation effect for women is positive and very close to that of men. There is, in other words, no clear male cohabitation/marriage premium visible in the fixed effects model, just a gender-neutral positive effect of cohabitation. For the cohabiting/married coefficients, we can note that the effect for men is more than halved in the fixed effects model without controls. Once individual time-constant heterogeneity is accounted for, the "pure" positive effect on authority for men by entering into cohabitation is much smaller. This pattern suggests that there is a positive selection of men into cohabitation and marriage. All of the male (and female) cohabitation premium is captured by the controls, which suggests that cohabitation for men *and* women is accompanied by an increase in investments relevant to the exertion of authority. Because men, but not women, are positively selected into cohabitation, we only observe a male cohabitation/marriage premium in the (early) cross-sections.

Table 3 Linear probability model estimates of authority on independent variables. Individual panel data 1968–2000[a]

	Men, total estimators		Men, fixed effects estimators		Women, total estimators		Women, fixed effects estimators	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Child in HH 0-6	0.006	0.015	0.039*	0.047**	-0.015	-0.001	-0.014	0.001
	(0.014)	(0.014)	(0.017)	(0.016)	(0.012)	(0.012)	(0.016)	(0.016)
Child in HH 7-20[b]	0.098**	0.037**	0.120**	0.039*	0.002	0.005	0.014	-0.006
	(0.012)	(0.013)	(0.014)	(0.016)	(0.010)	(0.011)	(0.013)	(0.015)
Child 20+[c]	0.068**	0.031	0.093**	-0.007	-0.007	0.001	0.051**	-0.006
	(0.014)	(0.017)	(0.018)	(0.022)	(0.011)	(0.013)	(0.017)	(0.021)
Cohabiting/married	0.121**	0.062**	0.054**	-0.024	0.001	0.007	0.045*	0.025
	(0.014)	(0.013)	(0.020)	(0.020)	(0.012)	(0.012)	(0.018)	(0.019)
Education yrs		0.039**		0.015**		0.030**		0.019**
		(0.002)		(0.005)		(0.002)		(0.006)
Hrs paid work/wk (/10)		0.086**		0.076**		0.041**		0.039**
		(0.009)		(0.012)		(0.004)		(0.007)
Work experience yrs		0.016**†		0.013**†		0.008**†		0.007*†
		(0.001)		(0.003)		(0.001)		(0.003)
Work exp. squared (/100)		-0.025**†		-0.027**†		-0.012**†		-0.014**†
		(0.003)		(0.004)		(0.003)		(0.005)
Seniority yrs (/10)		0.056**		0.016		0.032**		0.013
		(0.007)		(0.010)		(0.007)		(0.011)
Private sector		-0.003		0.014		-0.032**		0.012
		(0.014)		(0.025)		(0.010)		(0.019)
Year 1974		0.006		0.085**		-0.020		0.015
		(0.012)		(0.019)		(0.011)		(0.018)
Year 1981		-0.010		0.134**		-0.007		0.059*
		(0.014)		(0.032)		(0.013)		(0.027)
Year 1991		-0.049**		0.186**		-0.067**		0.041
		(0.015)		(0.052)		(0.015)		(0.041)
Year 2000		-0.095**		0.196**		-0.106**		0.022
		(0.017)		(0.071)		(0.017)		(0.055)
Intercept	0.157**	-0.735**	0.185**	-0.416**	0.179**	-0.336**	0.125**	-0.271**
	(0.009)	(0.045)	(0.014)	(0.074)	(0.010)	(0.028)	(0.014)	(0.066)
R squared/R squared within	0.04	0.15	0.03	0.08	-0.00	0.09	0.01	0.03
Number of observations	8553		8553		7245		7245	
Number of individuals	3844		3844		3326		3326	
Number of auth. switchers			902				593	
Number of switches: up			727				457	
Number of switches: down			444				314	

** p<0.01, * p<0.05 that coefficient is equal to zero (t-test), † p<0.05 that both terms in the polynomial are equal to zero (F-test).

[a] Robust standard errors in parentheses. The standard errors are adjusted for intra-individual correlation in columns (1)-(2) and (5)-(6).

[b] 7-18 years in the 1991 and 2000 surveys.

[c] 18+ years in the 1991 and 2000 surveys.

With the caveat that the estimated effects pertain to a selected group of men and women (those who exhibit variation in the variables of interest), the results indicate that when men and women become parents, men gain authority, whereas women's chances stay constant. Further, men, unlike women, who are married or cohabiting are more likely to exert authority, but this seems to be an effect of a selection of authority-prone men into marriage/cohabitation. Taking individual selection and controls into account, the net effect of marriage/ cohabitation is zero for men, as well as for women.

6 Concluding discussion

Research on workplace authority has uniformly shown that women are under-represented in such positions. In this paper we raise two questions on this gender gap in workplace authority. First, how has it changed over time? Second, to what extent are family conditions, i.e. marital status and children, associated with the gender gap in workplace authority? We extend the second question by asking how the impact of family conditions has changed over time and whether family conditions are *causally* linked to workplace authority. Since almost all previous studies in the field are based on non-repeated cross-sectional data, the present knowledge on these issues is limited.

What, then, has happened to the gender gap in workplace authority over time? To answer this question we use Swedish Level of Living Survey data, covering the period 1968–2000. We find that in 1968, the proportion of women in authority positions, irrespective of the number of subordinates, was 15 percent. This proportion had increased to just below 20 percent in 2000. For men the corresponding proportion has remained about or just below 30 percent throughout the period. Hence the relative chance for women to reach a position of authority in the labor market improved during the period. The most rapid change appeared in the 1970s, whereas the equalization has moved at a slower pace since the early 1980s. Moreover the levelling has been stronger on higher hierarchical levels (6+ subordinates), i.e. at levels where the gender gap was, and still is, most pronounced.[21]

[21] It is interesting to note that the development for the gender gap in workplace authority shows similarities to the gender wage gap trend. Analyses based on the same data set as here show that the gender difference in wages decreased substantially between 1968 and 1981 but has remained stable thereafter, despite the fact that women have increased their human capital, i.e. education and work experience, relative to men (Edin & Richardson 2002; le Grand et al. 2001).

The descriptive analyses also showed that for men workplace authority is clearly linked to age. As men get older their chance to hold an authority position increases sharply and peaks in the age span 40–50 years, to gradually decrease thereafter. Women exhibit a different pattern. In their twenties they exert authority to the same extent as men do. At older ages, however, their promotion chances level off and remain at a substantially lower level than for men of the same age. Coincidentally, the gender dividing point appears in the early thirties, when many men and women have formed, or are about to form, a family.

This finding leads us to our second main question: how are family conditions related to the gender gap in workplace authority? In fact, our results point out family formation as a watershed generating this gap. As long as male and female employees are single and childless there is no male advantage in workplace authority. In this group women are sometimes even overrepresented in positions of authority. This result contradicts the hypothesis on taste discrimination. If employers were systematically involved in such behavior, we would expect all women, regardless of their family situation, to exert less workplace authority than men.

After men and women have become parents, a gender gap arises and lasts throughout the career (see Meyersson Milgrom & Petersen 2006 for a similar conclusion). Results from cross-sectional analyses reveal a "male marriage premium", as well as a "fatherhood premium".[22] However, no marriage premium can be observed in the panel analysis with individual fixed effects. This finding suggests that the premium was primarily generated by the fact that authority-prone men were selected into marriage and cohabitation. This selection effect seems, however, to have become much weaker over time. The

[22] To some extent the male premiums can be explained by the fact that married/cohabiting fathers have higher education and have more work experience and seniority than mothers and childless/single men. We further found the correlates of authority to have changed during the period of study. It has become increasingly difficult to predict authority with observables, which indicates that factors other than those observed in the data set have become more important for the attainment of authority over time. A parallel to this result can be found in Swedish studies of wage differentials, where the variance explained by traditional Mincer equations has diminished somewhat over time (Edin & Richardsson 2002; le Grand et al. 2001). For one reason or the other, employers have begun to let individual differences other than those that are easily observed affect the decision to promote. Whether this is a long-term trend or just random fluctuation is too early to tell. More (longitudinal) studies of the predictors of authority are needed to know whether our result indicates a more general phenomenon, rather than being idiosyncratic to the Swedish labor market during this particular period of time.

reason for this change is not clear. Authority-prone men, for some reason, may have become less interested in marriage/cohabitation, or women to a lesser extent in the 1990s than in the late 1960s and 70s have preferences to mate with authority-prone men.[23] Employers may also decreasingly choose to promote married and cohabiting men. Like the cross-sectional analysis, the fixed effects panel analysis suggests that having children does not affect the chance for women to have authority, but increases the chance for men to have authority. In other words, the gender gap in workplace authority, that appears only after parenthood, is due to the fathers' gain in chances for promotion compared to when they were childless.

The differences in findings and conclusions between analyses based on cross-sectional and panel data respectively show that the association between family conditions and the gender gap in workplace authority is more complex than could be grasped in cross-sectional data, and the evidence we have had access to, so far, has relied on such data. Hence we would like to underscore the importance of longitudinal data for future studies in this field of research. Previous estimates of gender-specific effects of family factors on workplace authority are likely to be biased, because men and women are not randomly selected into cohabitation, marriage, and parenthood.

A main finding in this paper, then, is that men face better promotion chances when they become fathers. We outline two, possibly complementary, explanations of this empirical fact. First, it may be generated by a typically male response to the life-event of becoming a father. Because of socialization/norms/biology men react to this event in a gender-stereotypical way. They take the traditional household provider role, and entering into an authority position in their workplace may be one way of doing this. Women, instead, take the traditional role of caring for their children, which only with difficulty can be combined with an active career involving authority and the like. A substantial part of the gender gap in authority among parents can be explained by fathers' higher levels of, and higher returns to, education, work experience, and seniority. Becoming a father therefore seems to make men more prone to devote energy to market work, and human capital accumulation relevant to the exertion of authority.

[23] One way this change may have come about is that women with increasing labor participation (and divorce) rates have become accustomed not to rely on their men to take responsibility for their economic well-being.

Second, employers may begin to treat men differently once they (are about to) become fathers. Because of norms and expectations around parenthood, employers may interpret the event of a man becoming a father as a signal that he is now prepared to take on more responsibilities at work, an interpretation that is unlikely to be made when a woman becomes a mother (see Bielby & Baron 1986). On the contrary, employers may (often correctly) anticipate that when women become parents, they take the main responsibility for the children. Given the gender bias in child and household responsibilities, statistical discrimination against (potential) mothers, by not promoting them, may appear as a rational strategy for employers.

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