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DOES AGE EXACERBATE THE GENDER-WAGE GAP? NEW METHOD AND EVIDENCE FROM GERMANY, 1984–2014

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ABSTRACT

Given theoretical premises, the gender-wage gap adjusted for individual characteristics is likely to vary according to age. This study adapts John DiNardo, Nicole M. Fortin, and Thomas Lemieux's (1996) semi-parametric technique to disentangle year, cohort, and age effects in adjusted gender-wage gaps. The study relies on a long panel of data from the German Socio-Economic Panel (SOEP) covering 1984–2015. The results indicate that, in Germany, the gender-wage gap increases over a birth cohort's lifetime, including in the post-reproductive age for some birth cohorts. The results suggest that age and gender are overlapping handicaps in the labor market and call for a policy intervention.

KEYWORDS

Gender-wage gap, age, cohort, decomposition, non-parametric estimates, Germany

JEL Codes: J31, J71

INTRODUCTION

Although the age dimension is often absent from economic analyses of the gender-wage gap, there are reasons why it could matter for the gender disparity in earnings.[1] Policy reports – such as Organisation for Economic Co-operation and Development (OECD; 2012) – show that raw differences in earnings tend to increase as men and women grow older. With population aging, this should imply increasing raw gaps in aggregate terms due to a greater share of older employees in the population. However, the aggregate *raw* gender-wage gap remained fairly constant over time in most OECD countries, which hints at considerable changes in the *adjusted* gender-wage gaps in the life cycle.

While looking deeper into the age and cohort differences in adjusted wage gaps seems particularly relevant, this dimension is nearly absent in the literature. Perhaps the most advanced attempts are the ones linking time trends to institutions (Blau and Kahn 2003; Weichselbaumer and Winter-Ebmer 2007). Yet, the time span under these analyses combines a wide variety of changes, including declines in fertility, absolute and relative increases in women's educational attainment, and changes in occupational and industrial structure, as well as a number of other societal and lifestyle changes. With gradually closing gaps in education between young men and women, as well as near-to-par participation in the labor market for entry cohorts, one should expect considerable changes in the *adjusted* gender-wage gap for consecutive cohorts and within a cohort's life cycle.

One of the reasons for the absence of age patterns in the analyses of gender-wage disparity may be the technical challenge associated with the fact that age and cohort effects are at play jointly, together with period effects. Separating them statistically poses a difficulty, given that these variables are perfectly colinear: age equals year minus birth cohort. Our objective in this paper is to fill this gap in the literature, disentangling age and cohort effects and, thus, providing new insights into the age patterns in unexplained women's penalty in wages. We propose to address the identification problem by a novel extension of the semi-parametric decomposition developed by John DiNardo, Nicole M. Fortin, and Thomas Lemieux (1996). This decomposition allows flexibility in defining the counterfactual structure of wages. To identify the effects of aging on the adjusted gender-wage gap, we propose to utilize the most reliable counterfactual: one's past and/or future earnings. Our approach consists of providing a *double* decomposition of the changes in the gender-wage gap: *against the same birth cohort* at different points in time, and *against a different birth cohort* at the same point in time. We estimate this for both men and women; thus our estimates are, in fact, differences-in-counterfactuals.

To apply the proposed empirical strategy, one needs a relatively long panel, ideally observing the entire professional path, which is rare to find. Perhaps the one with the highest quality is the German Socio-Economic Panel (SOEP), spanning years from 1985 to 2015 for West Germany. In addition to offering high-quality data, Germany itself is a particularly relevant case. The increase in educational attainment among women, the drop in fertility rates, and the postponement in childbearing in Germany were among the most pronounced of all advanced economies (OECD 2012). At the same time, Germany remains characterized by relatively high gender inequality. It has the highest raw gender-wage gap in the European Union (OECD 2012), with estimates of 22 percent relative to men's wages at the median. Not only is the pay gap well above the OECD average (15 percent), it is also the only country where the difference has effectively grown over the last decade. Raw gaps actually understate the scope of gender inequality. Stephen Machin and Patrick A. Puhani (2003) show that the unexplained component was twice as large in Germany as it was

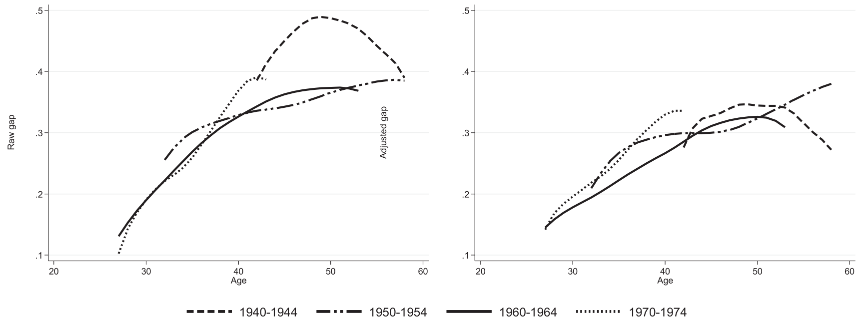

Figure 1 Gender-wage gaps in Germany, evolution for selected cohorts
Notes: Figure shows values of the raw and adjusted gap for four selected cohorts. Adjusted gaps estimated using the DiNardo, Fortin, and Lemieux (1996) decomposition. Controls include marital status, age, education level, tenure, and experience.
Source: SOEP, 1984–2015.

in the United Kingdom in 1996 (see also Arulampalam, Booth, and Bryan [2007]). Using high-quality matched employee–employer data, Thomas Hinz and Hermann Gartner (2005), as well as Anja Heinze and Elke Wolf (2006), provide estimates of the adjusted gap that amount to as much as 20 percent of women's wages.

Although our paper is not the first to study the gender-wage gap in Germany, to the best of our knowledge, it is the first to focus on wage patterns over the life cycle. Figure 1 is indicative of why this perspective is particularly relevant. The left panel displays the raw gender-wage gap for a few selected cohorts, whereas the right panel does the same for the estimates of the adjusted gender-wage gap. Both raw and adjusted gender gaps vary with age. Naturally, Figure 1 confounds year effects with age effects. For example, a hike in both raw and adjusted gap experienced by cohorts born between 1950 and 1954 is reflected in a similar hike in ten-years-older age brackets for the cohorts born between 1940 and 1944. This tentative evidence demonstrates an important role for cohort, age, and year effects in determining the adjusted gender-wage gaps.

Also, subsequent cohorts appear to display different age patterns: the oldest cohort displayed in Figure 1 is characterized by an inverted U-shaped pattern in both the raw and the adjusted gender-wage gap, whereas the subsequent cohorts display a less curved inverted U-shape for the raw and an increasing pattern for the adjusted gap. Cohort effects might stem from societal changes, such as a revision of gender roles inside the household, the surge in women's university enrollment rates in most countries, and postponed childbearing. They may also be driven by institutional changes, such as those in pension systems, maternity leave, and access to childcare facilities. Year effects may reflect business cycle effects, tax and other welfare reforms, and other sources of changes in labor demand and labor supply, such as, for example, the adoption of new technologies that reduce the effort demanded by household activities. Finally, age effects, being entirely individual, reflect changes in human capital as well as social norms related to employability and expected productivity of employees.

Our results suggest the existence of a strong age-related pattern – meaning that the adjusted gender-wage gap increases with age, even at later stages in life. The rate of growth at various stages throughout the life cycle differs across birth cohorts. Such patterns suggest that the main explanations of the adjusted gender-wage gap – that is, career interruptions, time spent on unpaid domestic work, and returns to investment – may be interacting with societal explanations, such as the doubling of age and gender penalties in the wages of women older than the prime age.

THEORY AND CONTEXT

A raw difference in wages across genders may be a misleading measure for the actual gender-wage inequality. Imagine a scenario where men and women receive, on average, the same wages, but women are better educated. In this case, the raw difference in wages suggests no inequality, yet behind this apparent equality, we observe different rewards to the human capital embodied in men and women. Consequently, in order to obtain reliable measures of inequality, empirical studies strive to adjust the gender-wage gap for observable, and thus explained, differences between men and women. In this framework, the part of the gap that is not attributed to characteristics (called the adjusted gender-wage gap) is a proxy for discrimination. Because of its residual nature, the adjusted gender-wage gap lumps together both *true* gender discrimination and unobserved gender differences in productivity. These unobserved differences are attributable to, among others, women's reproductive roles, which tend to interrupt women's careers, limit their occupational choices, and decrease their efforts and energy devoted to paid work, all of which are only poorly captured by surveys. These departures of empirical observations from theoretical concepts (in other words, mismeasurement) depend on a variety of factors that might be intertwined with other labor market processes, including occupation (and occupation sorting), industry (and industry sorting), and time (that is, due to technology adoption).

Insights from theory

Ever since the human capital theory of Gary S. Becker (1985), investment in human capital throughout the lifespan has become a central topic of economic analyses. Viewed through this lens, childbearing and childrearing induce career interruptions (Mincer and Polachek 1974; Iversen and Rosenbluth 2006), which are treated as human capital depreciation. Moreover, these expected interruptions might even discourage investment in human capital by primary caregivers, typically women (Polachek, Zhang, and Zhou 2014). Caring obligations may limit women's occupational choices to jobs that allow accommodations for nonmarket work, which are also characterized by fewer opportunities to accumulate human capital and yield lower rewards to seniority.[2] If human capital investment is cumulative in nature, the motherhood wage penalty could continue beyond reproductive age. These differences in human capital investment and accumulation should be reflected in raw wage gaps. Properly accounting for differences in human capital investment should eliminate the differences in the observed wage gaps. Yet, experience and years of formal education are highly imperfect proxies of human capital, and the more imperfect they are, the longer the labor market tenure of an individual.

In addition, time endowments available for market work differ between primary caregivers and other household members, even many years after a career interruption, due to specialization within the household. Nested in a standard framework, specialization of household members in market and nonmarket work is rational, as it allows exploiting the comparative advantage that one member has in the labor market. As household activities demand more effort than leisure does, those who engage in household labor, usually women, have less energy to devote to paid work and thus become less productive in that paid work, even if they spend the same number of hours at their place of employment (Becker 1985). Different levels of engagement in household activities could yield raw gaps increasing with age until the caring activities are completed.

The division of tasks in the household serves to perpetuate and magnify even small gender differences in the labor market. Notice that the prior existence of labor market gender discrimination is not necessary for this channel to operate; some policies to facilitate women's participation, such as maternity leave, could have similar unintended consequences. Pia S. Schober (2013) estimates that women who take longer maternity leaves tend to carry a larger share of household duties even upon their return to the labor market. While Schober interprets her findings in terms of habit formation, it could also result from the development of a comparative advantage in household tasks. In this setup, raw gender-wage gaps should widen with age. However, the implications of this mechanism are somewhat less clear for the relationship between adjusted gender-wage gaps and age. The level effect of career interruptions is typically accounted for by including years of experience in the adjustment process. Once adjusted for level effects, the adjusted gaps would widen with age if returns to experience accumulation were somehow nonlinear. Women who suspend careers to engage in domestic activities might not accumulate experience at the same pace as men do and lag behind in terms of productivity and, thus, seniority premiums. If that is the case, econometric models would yield uneven estimates of gender-wage gaps across age as a mismeasurement of the actual differences in experience rather than returns to experience.

Although theories related to the supply side – based on human capital and career interruptions – suggest that gender-wage gaps might increase with age, it remains unclear whether this increase should continue in the post-reproductive years. Indeed, childrearing eventually ends, which implies that whatever gap was accumulated up until that stage in life may remain fairly constant in subsequent years. Given that the motherhood wage penalty cannot be fully accounted for by the observed individual characteristics, its effect on the adjusted gender-wage gap could in principle go beyond reproductive age. Arguably, the gap may continue to widen if returns to earlier investment in human capital are nonlinear. They might give rise to a difference in productivity between men and women, including in the post-reproductive stage of life, which is hardly captured by commonly available proxies of human capital. By compounding, a small difference in productivity between men and women in the early stages of their careers could grow larger with time. However, stylized facts appear to be at odds with this proposition, as it would require ever-growing wages for men and women, while empirical studies suggest that hourly wages for both genders stagnate at ages above 50 years old and might even decline later on (Smyk, Tyrowicz, and Liberda 2014; Rupert and Zanella 2015; Bhuller, Mogstad, and Salvanes 2017; Hanushek et al. 2017). If wages decline at different rates for men and women, the gender-wage gap may continue to grow in age groups of 50 years and older, but the mismeasurement of human capital is not likely to be a mechanism behind these patterns.

On the labor demand side, rational employers – expecting women to give birth and subsequently carry a larger share of household chores – will account for that productivity shortfall in lower wages. In turn, rational employees of both genders will incorporate this insight in formulating the reservation wages (Arrow 1973; Spence 1973; Dahlby 1983). This explanation, sometimes referred to as statistical discrimination, provides no insights into the age pattern of the adjusted gender-wage gap. However, one may expect that with gradually declining fertility and delayed childbearing, that discount should also decline gradually.[3]

Furthermore, if taste-based discrimination and/or statistical discrimination transform the labor market into an occupationally segregated market, where women are concentrated in occupations that reward seniority less than occupations held by men, then one can expect the raw gender-wage gaps to grow with age. Naturally, after proper accounting for occupations held, the adjusted gender-wage gap need not display any specific age pattern. Especially if occupational mobility declines with age, then the differences accumulated up to the career height should continue as men and women age, but it is not obvious that they would widen any further.

Another group of demand-side reasons for why the adjusted wage gaps should *continue to increase* in the post-reproductive age relates to the institutional theories for the prevalence of gender inequality in the labor market (Ferber and Nelson 1993; Kabeer 1994; Bergmann 1996; Agarwal 1997). One of these is the "double standard of aging" hypothesis (Bergmann 1981; Sontag 1972). Sara Wilcox formulates the double standard of aging as a "differential treatment of aging, in which women lose value and see themselves more negatively with increasing age, whereas men maintain or gain value" (1997: 550).[4]

The case of Germany

Social norms tend to be strongly biased against employed mothers in Germany. In a comparative study, Judith Treas and Eric D. Widmer (2000) found that women were expected to not do market work full time between the time of giving birth and the time of school enrollment of their children. In fact, only 1 percent of the respondents indicated that they believed otherwise. Only after all children are in school does it become socially acceptable for women to re-enter the labor market full time (more than 50 percent of the respondents supported this option).

A mixture of welfare policies reinforces these social norms. The "tax splitting" system lowers the average tax rate for the household, but imposes a high shadow tax rate for the second earner.[5] In addition, childcare facilities are rare in Germany, with only 3 percent of children under 3 years old having attended these institutions in the late 1980s. Despite the subsequent increased availability, in the early 2000s, the percentage still hovered around 12 percent, which implies a high shadow price of professional work. The duration of parental leave, which can only be utilized by the mother, increased from 4 months in 1984 to up to 34 months in 2006, which provides incentives for women to specialize in home production (during the leave period there is only a limited number of hours that a woman might work for pay, albeit this has been gradually raised from 15 hours per week in 1986 to 30 hours per week in 2001). Moreover, childcare benefits to be paid to the mother during the leave are fixed and do not depend on previous earnings, while childrearing benefits are means tested, and dual-earner families are typically ineligible.[6] This legislation has been proven to adversely affect women's labor supply (Merz 2004; Schober 2012), further reinforcing the existing prejudices delineated by Treas and Widmer (2000).

On top of the gender-related concerns, participation of employees above age 55 is an important policy concern in Germany, even though the *de* *jure* minimum eligibility retirement age is set at 65. Although the use of early retirement schemes is voluntary, labor contracts can include an additional provision whereby the employees commit to retiring early. Such provisions were abolished as of 1992. A correspondence study run by Victoria Büsch, Svenn-Åge Dahl, and Dennis A. V. Dittrich (2009) shows that older applicants are 22 percent less likely to be called for an interview than younger applicants. In addition, employees over 55 years old are allowed to unilaterally reduce workload from full time to part time. Overall, the effective age of leaving the labor market and claiming pension benefits is around 60 (or even 58 if employees decide to retire and claim unemployment benefits for two years, which yields a replacement rate of roughly 60 percent).

METHODS AND DATA

In this paper, we contribute to the literature by delineating a life-cycle and a cohort perspective on the adjusted gender-wage gaps. Despite the richness of the parametric and non-parametric methods developed to estimate the gender-wage gaps – see a recent review by Nicole Fortin, Thomas Lemieux, and Sergio Firpo (2011) – no method has yet disentangled age, time, and cohort effects with respect to the gender-wage gap. However, DiNardo, Fortin, and Lemieux (1996) proposed a versatile semi-parametric approach for the decomposition of the entire wage distribution. As the estimation of the differences is based on a non-parametric kernel, it avoids the perils of specifying a functional form for wages.[7]

We propose a novel application of DiNardo, Fortin, and Lemieux (1996), utilizing one's wages from the past/future to provide a "first counterfactual" – that is, isolating the age effects. We then exploit differences between men and women to provide an estimator based on "difference-in-counterfactuals" – that is, how the gender gaps change as men and women both age. Below, we discuss the proposed methodology and subsequently, the data properties.

Methodology

The DiNardo, Fortin, and Lemieux (1996) decomposition is based on the idea that average wages represent an integral over density function of individual characteristics. To build a counterfactual distribution, one modifies the density function, adapting it to the group of interest. Thus, to obtain women's distribution of wages if they were paid as if they were men, it is sufficient to reweight the density function of men according to women's distribution of characteristics. The difference that remains between the actual and counterfactual women's wage distribution cannot be explained by differences in characteristics and thus is analogous to the traditional understanding of the adjusted gender-wage gap.

Even though it is possible to discretize the distribution of characteristics and reweight each resulting bin using the relative probabilities of belonging to each group, such an approach would be inefficient: the number of bins grows and the statistical power declines with the number of characteristics to be included. To circumvent this issue, DiNardo, Fortin, and Lemieux (1996) introduced a weighting function $\Psi(x)$, to "weight" all observations of men by the probability of being a woman, given their characteristics. The procedure consists of estimating a probit model, where the dependent variable is the membership in a group of interest (men or women, young or old, etc.) – for example, weighting function of women's wages for the year t would be given by:

$$\Psi(\mathrm{x}) = \frac{\mathrm{P}(\mathrm{s} = \mathrm{men} = \mathrm{x}, \mathrm{j} = \mathrm{t})}{\mathrm{P}(\mathrm{s} = \mathrm{women}|\mathrm{X} = \mathrm{x}, \mathrm{j} = \mathrm{t})} * \frac{\mathrm{P}(\mathrm{s} = \mathrm{women}|\mathrm{j} = \mathrm{t})}{\mathrm{P}(\mathrm{s} = \mathrm{men}|\mathrm{j} = \mathrm{t})}$$

By changing the definition of $\Psi(x)$, one defines the specific counterfactual, which allows a comparison of the distribution of wages of women employees in two periods, and even to compare men and women in different periods.

We exploit this feature of the methodology to produce a *double decomposition* across time and gender (see also Cho and Cho [2011]). Let Δ_{t} be the raw gender-wage gap in year t. It represents a sum of the explained (by differences in endowments) and the unexplained (due to differences in rewards, also known as adjusted wage gap) components:

$$\Delta_t = \underbrace{\int f_{men,j}(w|x)h(x|s = men, t = t)\,dx - \int f_{women,j}(w|x)\Psi(x)h(x|s = women, t = t)\,dx}_{Unexplained\ component}$$

$$+ \underbrace{\int f_{women,j}(w|x)\Psi(x)h(x|s = women, t = t)\,dx - \int f_{women,j}(w|x)h(x|s = women, t = t)\,dx}_{Explained\ component}$$

The interest lies in estimating the changes across time, in short $\Delta_{\mathrm{t}} - \Delta_{t-1}$. This difference consists of four components. The first one is the measure of changes in the characteristics of men and women, keeping the rewards to the characteristics constant at a given base (for example, men in period t). A positive sign indicates that the change in endowments was larger for men than for women. The second is the change in the unexplained components, keeping the characteristics fixed at the level of women from period t. A positive change indicates that the adjusted wage gap grew over the period under analysis. The remaining components correspond to interactions between changes in endowments and changes in rewards. Since these interactions lack a clear interpretation, they are grouped under the term residuals. Clearly, the equation could also be employed to study the changes over time for a particular cohort. The full derivation of the decomposition is presented in the Supplemental Online Appendix.

The decomposition proposed above offers a straightforward analysis of the changes in the wage gap over time for the entire population. For example, we can decompose the changes in the wage structure for women ages 30–34 in 1989, with respect to their situation in 1984, when they were 25–29. At this point, the advantage of using SOEP becomes evident: by following individuals through the life cycle, we may repeat the procedure to obtain the estimates of changes in adjusted gender-wage gaps as individuals age, controlling for individual- and cohort-specific effects.

Data

We work with the SOEP for West Germany for the period 1984–2015. The SOEP is a longitudinal survey conducted in annual face-to-face interviews (see Pannenberg [2000]).[8] We follow each individual for as many years as (s)he is available in the sample, with over 1,810 individuals observed over the entire span of thirty years and over 360,000 total individual-year observations.[9]

Individuals report family situation, which includes marital status and household composition. Along these household variables, the SOEP also provides valuable data on the labor market status of the individuals: it contains information on net wages, paid working hours, type of employment (whether part time or full time), experience (also split for part and full time), positions held, and tenure, as well as firm industry and size. We construct (log) hourly wages by dividing the regular gross wage by the usual paid working hours reported in each period. Wage data were deflated using SOEP-provided inflation and converted to euros at the official conversion rate.

The sample comprises all West German nationals ages 25–59 who were wage employed in at least one period.[10] The double decomposition requires that each individual be observed at least twice. Since the SOEP is a panel, this requirement is not particularly binding. In total, the SOEP contains 208,589 observations for West Germany nationals with non-missing data for relevant characteristics (age, education, household structure, and so on). Out of this sample, 201,846 observations reflect year-persons observed at least twice.

While thanks to the novel adaptation of DiNardo, Fortin, and Lemieux (1996), our decomposition unveils the role of aging in determining women's wage penalty, it has a rather nonstandard way to account for selection effects. Namely, the nature of this decomposition is to reweight the distribution of wages for a given group by the distribution of characteristics pertaining to a counterfactual group. We utilize one's past/future as the counterfactual, observing the differences between men and women. Thus, potential bias in estimating the age pattern of the adjusted gender-wage gap could arise if the event of observing only once women's wages in the sample was substantially more frequent than for men. After testing this explicitly, we found that, indeed, women have more interruptions, but the differences between men and women concern never-employees rather than single-year employees. In fact, the share of observations dropped to 4.9 percent for men and 3.8 percent for women because an individual worked for pay only once in the sample amounts. By contrast, the share of observations dropped because the share of respondents who never worked for pay in the sample is 14.9 percent for men and 25.6 percent for women.

Clearly, as our decomposition cannot account for selection into never-employed, this topic remains beyond the scope of our analysis. Childbearing is an individual characteristic utilized to obtain counterfactual distributions (changing with age in the majority of cases). Since for each ever-employee we utilize all available observations, our estimates are not troubled by the bias from potential unobserved heterogeneity in preferences for temporary inactivity (for example, preference to stay away from the labor market longer after childbearing). Naturally, the conclusions apply to the 85.1 percent of men and 74.4 percent of women who have ever been employed in the SOEP.

It may occur that an individual works for pay in one period and does not work for pay in the subsequent one. Given that employment decisions are not random, this could lead to well-known selection bias problems. Moreover, the selection bias need not be constant over the life cycle. One could expect selection bias to be particularly acute among employees close to retirement age: individuals who value careers more are potentially more likely to work for pay for the entire period between 45 and 59 years of age, hence ushering an additional reason for employment selection, which could be related to the experienced degree of labor market fairness. Simultaneously, age- and gender-specific selection may emerge if, for example, only more capable women stay in the market. From descriptive statistics reported in Supplemental Online Appendix Table B1, labor market exit is indeed early in Germany, and it is likely to be selective in a sense that women with relatively higher wages are more likely to remain active in the labor market than those with relatively lower wages, ceteris paribus. The German system provided relatively low incentives to postpone retirement, and the eligibility age for women has been low for most of the analyzed period.

In order to mitigate selection bias,[11] we assign, in separate estimations, the temporarily non-employed respondents (1) the previous non-missing wage; or (2) zero wages when they do not work in the labor market.[12] To account for the fact that assigning wages may be more probable for some employees than for others, we introduce a specification where we include employment status in the estimation of the reweighting function. Since our estimates utilize wage distribution from previous age as a counterfactual, this source of potentially unobserved heterogeneity will affect results only if it is not constant across time. If, as hinted above, selection into employment is adversely and jointly age and gender specific, the true age effects would be larger than those estimated in our model.[13]

Figure B1 and Table B1 in the Supplemental Online Appendix portray the changes in individual characteristics and gender-wage gaps over the age groups for selected periods. The data reveal a clear time pattern toward postponing getting married and raising children in favor of tertiary education and employment. Between 1984 and 2009, the proportion of women living with at least one child in their households fell considerably.[14] The incidence of marriage has also decreased, though to a lesser extent. Men experienced comparable trends toward postponing marriage and children, but at a lower rate. These changes in household composition and formation were accompanied by an almost twofold increase in the proportion of young women with tertiary degrees. As a result, toward the end of the sample, the proportion of women with tertiary education among the youngest age group exceeded that of men of the same age.

Table B2 in the Online Supplemental Appendix reports raw and adjusted gender-wage gaps for selected birth cohorts in selected years. Estimated adjusted gender-wage gaps are high, indicating a large extent of the unexplained gender inequality in the German labor market. We proceed to identification of the age pattern when controlling for cohort and year effects in the subsequent section.

RESULTS

The first stage in the DiNardo, Fortin, and Lemieux (1996) decomposition is the estimation of a probit model, where the dependent variable is gender (it takes the value of 1 if the respondent is a man). We include three sets of controls: household characteristics, human capital, and employment variables. Variables describing household characteristics are marital status of an individual (married or cohabiting) and a dummy for the presence of children younger than 5 years old in the household. The variables describing human capital are: educational attainment, tenure with the current employer, and years of experience. In order to accommodate for interruption in the career, experience is obtained as the difference between the actual experience and the average experience for employees of the same gender and of a similar age. Finally, employment is a dummy to distinguish employed respondents from non-employed respondents. We want to utilize the full extent of the available data – that is, including the data points for the non-employed periods of the individuals in the survey. Information is missing on occupation and industry at periods of non-employment. We update this information in coherence with wages: previous occupation/industry if previous wages are used or zeros in the specification where missing wages are treated as zeros.

After the estimation of probit, we recover counterfactual distributions of wages for individuals. These are obtained for each birth cohort for each available age group. To assure a sufficient sample size, we define birth cohort as individuals born within a five-year span. Birth cohorts are not overlapping. Following a similar logic, we also pull together individuals of a similar age, forming five-year age groups. Thus, when we refer to a subsequent period, we mean a period when a given individual is five years older, whereas a subsequent cohort signifies an individual born in the five-year period following the five-year period relevant for a given woman. Hence, this second step of the analysis yields a matrix of relevant counterfactual wage distributions – that is, distributions that are relevant for a given woman as she ages, relative to a man of the same birth cohort as he ages.

The final step of our analysis is the double decomposition, which identifies – in relative terms – the contribution of each age group to the pattern of the adjusted gender-wage gap.

Quality of the estimates

The semi-parametric DiNardo, Fortin, and Lemieux (1996) decomposition relies on the quality of the first-stage probit, thus it is useful to analyze the results of the first stage. Figure C1 in the Online Supplemental Appendix plots the value under the Receiver Operating Characteristic (ROC) curve for each year and age group: the number of correct predictions varies for subsequent models, but it is satisfactory.[15] Our model performs especially well in the case of older employees. Among the youth, the predictive ability of the model also varies across time: in the early years of the sample, probit models distinguish between men and women more accurately than in the last years of the sample. This suggests that the differences in observable characteristics between men and women are closing, especially after 1994.

Identifying age patterns in adjusted gender-wage gaps

Table 1 focuses on age patterns in the adjusted gender-wage gaps. Rows indicate the initial age, whereas columns display the initial period. For example, the first row in the first column indicates the change in the adjusted gender-wage gap experienced by women ages 25–29 in 1984 over the upcoming five years, when they eventually turn 30–34. The last row in the first column concerns individuals ages 50–54 as they become 55–59 (that is, our last age group of observation). To better identify the role of changing age, the sample of respondents is kept constant in each adjacent

Table 1 Changes in the adjusted gender-wage gap

Age		*Initial year*						*Average change*
t	*t* + *1*	*1984*	*1989*	*1994*	*1999*	*2004*	*2009*	*in AGWG*
Missing wages replaced with previous wage, w/o employment dummy in ψ function								
25–29	30–34	– 0.02	0.01	0.07	0.02	0.07	0.04	**0.03**
30–34	35–39	– 0.06	0.06	0.00	– 0.01	0.01	0.20	**0.03**
35–39	40–44	0.05	0.04	0.01	0.08	0.06	0.05	**0.05**
40–44	45–49	0.25	0.02	0.07	– 0.01	0.02	0.03	**0.06**
45–49	50–54	– 0.08	– 0.02	0.08	0.09	0.00	0.01	**0.01**
50–54	55–59	0.02	– 0.16	– 0.01	0.16	0.06	0.08	**0.02**
Missing wages replaced with previous wage, w/ employment dummy in ψ function								
25–29	30–34	– 0.08	0.02	0.07	0.02	0.03	0.06	**0.02**
30–34	35–39	– 0.09	0.09	– 0.01	– 0.01	0.02	0.25	**0.04**
35–39	40–44	0.06	0.05	0.03	0.08	0.07	0.07	**0.06**
40–44	45–49	0.23	0.05	0.05	0.00	0.02	0.03	**0.07**
45–49	50–54	– 0.06	0.02	0.04	0.10	0.00	0.02	**0.02**
50–54	55–59	0.14	– 0.14	0.06	0.13	0.06	0.09	**0.06**
Missing wages replaced with zeros, w/o employment dummy in ψ function								
25–29	30–34	0.12	0.16	0.00	0.19	0.05	0.25	**0.13**
30–34	35–39	– 0.11	– 0.13	0.13	0.17	0.10	0.62	**0.13**
35–39	40–44	– 0.05	0.07	0.16	– 0.02	0.06	0.09	**0.05**
40–44	45–49	0.25	0.04	0.10	– 0.06	0.05	0.03	**0.07**
45–49	50–54	0.05	– 0.24	0.03	0.18	0.03	0.10	**0.03**
50–54	55–59	0.23	0.15	0.01	0.09	0.03	0.14	**0.11**
Missing wages replaced with zeros, w/ employment dummy in ψ function								
25–29	30–34	0.11	0.14	– 0.01	0.17	0.04	0.18	**0.10**
30–34	35–39	– 0.18	– 0.10	0.15	0.20	0.07	0.61	**0.12**
35–39	40–44	– 0.06	0.10	0.14	– 0.02	0.08	0.12	**0.06**
40–44	45–49	0.22	0.03	0.07	– 0.01	0.06	0.09	**0.08**
45–49	50–54	0.10	– 0.17	0.01	0.15	– 0.01	0.13	**0.04**
50–54	55–59	0.33	0.16	0.09	0.10	0.00	0.21	**0.15**

Notes: SOEP 1984–2014. Adjusted gender-wage gaps (AGWG) estimated using double DiNardo, Fortin, and Lemieux (1996) decomposition. Zero wage specification takes effectively 0.1€/CPI – that is, the real value in 2005 euros of 10 euro cents. Changes in the adjusted gender-wage gap decomposed by age, reweighed by the distribution of men's characteristics. Controls include marital status, age, education level, tenure, years of employment experience (difference with respect to reference group mean), a dummy for employees in high-skilled occupations (ISCO code smaller than three), and industry dummies. Sample in the upper panel includes only individuals who were employed in the t and $t+1$. In the middle and bottom panels, sample includes individuals who were employed for at least one period in our sample. Estimates in the bottom panel include an employment dummy in recovering the ψ function.

two periods.[16] These estimates are additive – that is, for each birth cohort, as they age, one should follow the diagonal to observe the changes in the adjusted gender-wage gap. Positive values along diagonals signify that the adjusted gender-wage gap widens as a given cohort ages. The last column summarizes the average increase – over the birth cohorts/periods – in the adjusted gender-wage gap for given age brackets.

Estimates from Table 1 suggest that the adjusted gender-wage gap increases as women age. The age profile that emerges from Table 1 indicates a steep increase from the beginning of the career that lasts beyond the reproductive age. Estimates presented in the middle and bottom panel of Table 1 reveal the importance of controlling for selection bias. In a majority of cases, the introduction of corrections for selection results in higher estimates of changes in the adjusted gender-wage gap, which is consistent with the insights of Schober (2013). With the employment dummy in ψ function, women ages 40–44 experienced on average a 7–8 percentage point increase in the adjusted gender-wage gap over the following five years. In the subsequent age brackets, the increase of the adjusted gender-wage gap is slower and the actual size of the increment depends on the specification: if wages missing due to non-employment are anchored to the last observed wage, they tend to grow somewhat slower than if wages missing due to non-employment are replaced by zeros. Indeed, that may be related to the fact that pre-retirement or early retirement benefits could be superior outside options relative to, for example, unemployment benefits or social assistance, albeit this is not directly observed in our data. As revealed by Table B1 in the Supplemental Online Appendix, over most of the analyzed periods, women catch up with men in terms of observable characteristics. Indeed, the opposite adjustment in rewards stands behind relatively stable raw wage gaps in Germany.

Cohort and year effects also appear to play a substantial role. The increases in the adjusted gender-wage gap are non-monotonic, which points to the role of year effects, such as changes of legislation, but also the overall economic landscape. However, given how data intensive our methodology is, the SOEP sample size is too small to quantify such effects.

The size of the increase in the adjusted gender-wage gaps in post-reproductive age differs between cohorts. In general, the adjusted gaps increase in the 45–49 age group (with the exception of one or two birth cohorts, depending on specification) and continue to increase in the 50–54 age group (with the exception of one or two birth cohorts, depending on specification). Low incentives to postpone retirement should imply that higher-earning women are more likely to be observed among the employed population. Although we are not able to fully disentangle the selection effect from the age patterns in the adjusted gender-wage gaps (larger datasets, possibly administrative data would be needed), the estimated patterns increase in age on average. If adjusted gaps are lower among low earners, our estimates could be overstated. Conversely, if those with more unequal wages are more likely to leave early, our estimates would be understated.

For numerous cohorts, the unexplained component continues to increase after the 40–44 age brackets, despite adjusting for years of experience and tenure with the current employer (among other controls). This is equivalent to stating that as a man ages beyond 45 years old, his wage grows by more (or declines by less) than that of an observationally equivalent woman, as she ages beyond the post-reproductive age, despite accounting for the changing individual characteristics. One of the potential explanations for our findings may be dynamic mismeasurement of human capital. For example, some investments in human capital made at the age of 30 may be reflected in productivity and, thus, in wages, only at the age of 50. Then, women with a career gap due to childbearing and childrearing may be less likely to have made those investments. If this investment remains unobservable and if, at the same time, its relationship with experience is nonlinear, the omitted variable bias would exhibit as an increasing adjusted gender-wage gap in the post-reproductive age. While theoretically possible, the effects of these investments would need to be, indeed, large to widen the gap by an additional 7–8 percentage points twenty years later. More research would be needed on details of investment in human capital to identify if such gender differences exist and if their bearing on individual productivity could plausibly be as high as 7–8 percentage points.

An alternative explanation for the increasing adjusted gender-wage gaps after women have passed reproductive age is the "double standard of aging" hypothesis. If age and gender are both considered handicaps in the labor market (see Deutsch, Zalenski, and Clark [1986]; Wilcox [1997]; England and McClintock [2009]), the overlap of both may explain the persistence of adjusted gender-wage gaps in the post-reproductive period. A woman's age may be a more salient characteristic in labor demand than a man's age, especially in the case of leadership roles – that is, for the upper quartile in the income distribution (Nelson 1996). In some professions, for example, aging women are regarded as less valuable than young women (for the movie industry, see Lauzen and Dozier [2005]; in regards to TV anchors and reporters, see Saner [2010]). The evidence is scarcer on more common occupations, but the double standard of aging hypothesis could contribute to explaining a widening gender-wage gap with age.

CONCLUSIONS

Raw gender-wage gaps prove to be remarkably persistent over time. Referring to differences in the accumulation of human capital for both genders, some theories predict an inverted U-shaped pattern of the adjusted wage gap with respect to age, whereas some of those theories, particularly social theories, can be interpreted to suggest an ever-increasing age pattern. With aging, composition effects should imply gradually increasing aggregate estimates of the gender-wage gap. However, subsequent cohorts of women are gradually better educated than men, their fertility decreases, childbearing is delayed, and access to care facilities increases with an apparent trend toward an equalization of the division of labor within households. These trends imply smaller grounds for the statistical discrimination for the youngest age groups in subsequent birth cohorts. Theoretical implications for the cohort effects and for the age effects work in the opposite direction, blurring the analysis of changes in the aggregate inequality of men and women in the labor market.

Extending the DiNardo, Fortin, and Lemieux (1996) decomposition – a key methodological innovation of our study – helps to disentangle age and cohort effects. We construct estimates for the age patterns in gender-wage gaps as differences-in-counterfactuals: differences between men and women, as both men and women age. As the first counterfactual, we use the opposite gender, but as a second counterfactual, we use the individual's own wage at earlier/later stages of the life cycle. This method requires panel data, for which we utilize the SOEP for 1984–2015.

We find that women's unexplained penalty in wages is increasing with age and continues to increase for many cohorts, including in post-reproductive age. Some of the earlier literature has suggested gender inequality prevails among older employees. This study shows that the scope of that inequality is actually increasing with age, including among older employees. This suggests that age and gender are overlapping handicaps in the labor market, which calls for a policy intervention.

An obvious caveat of our study is that it only concerns one country – Germany. While it is an interesting case for a number of policy-related reasons, it may also prove singular, as the secular trends in education and fertility experienced in Germany were particularly strong. A second caveat concerns the data. Although the SOEP contains very high-quality data, in cohort- and age-specific analyses, like ours, the sample size proves to be a constraint. Research with the use of large administrative data could corroborate the findings with more precise estimates, also over finer defined age groups and birth cohorts. Moreover, with such a large sample size one could enrich the analysis by looking into occupational specificity, thus providing additional answers about the universality of the identified age and birth cohort patterns.

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NOTES

[1] In this paper, unless stated otherwise, the terms “gender-wage gap” and “gender differences in earnings” refer to the differential adjusted for individual characteristics. The unadjusted gap is identified by the term “raw.”

2 A similar argument applies to women's decisions on which career to follow. Anticipating interruptions, women might select the careers that impose a lower penalty for interruptions. For example, Claudia Goldin and Lawrence F. Katz (2008) show that women with more children tend to work in careers with lower wage penalties for interruptions.

3 Importantly, alternative individualist explanations for the persistent gender-wage gap even after accounting for individual characteristics – such as lower wage expectations of women (Blau and Ferber 1991; Reuben, Wiswall, and Zafar 2013) or taste-based discrimination (Becker 1971; Lang 1986; Duncan and Loretto 2004) – have no clear time/age-related patterns. These explanations are also weakly founded in feminist theories.

4 This approach finds some empirical support: Peter Kuhn and Kailing Shen (2013) analyze job ads from China and find that some job openings are subjected to strong bias toward young and attractive women. Moreover, firms tend to have a higher preference for hiring men when looking for older employees than when looking for employees in other age groups (Duncan and Loretto 2004; Lincoln and Allen 2004; Lauzen and Dozier 2005; Neumark, Burn, and Button 2015).

5 Charlotte Lauer (2000) and Elke Holst and Anne Busch (2009) explore the glass ceiling in wages by studying individuals in managerial positions, suggesting that the unexplained component within the top occupation is approximately 40 percent of men's wages, once selection bias is taken into account. David Reimer and Jette Schröder (2006) also explore the relation between the adjusted wage gap and the field of education, but they do so in a much more homogeneous population (university students). Their results indicate that the adjusted gap among former students was between 4.3 percent and 7.6 percent of women's wages. Though the value is much lower than in the previous case, it must be taken into account that the sample covers only individuals at the beginning of their careers. Doreen Triebe (2013) finds that while an increase in the salaries of women leads to reduced labor supply of men, as men's labor is subject to a higher shadow tax, which leads to lower marginal income. Women's response to wage increases by men differs. Namely, married women do reduce paid working hours because of tax splitting, but cohabiting women do not.

6 This legislation was changed in 2007. Benefits are not means tested, and they are proportional to previous earnings. Yet, the effects of this reform are to be observed only in the cohorts eligible between 2008 and 2015.

7 Though originally intended to measure the consequences of the changes in the unionization rate, it was adopted to measure the impacts of other variables as well, among them gender. Casey Warman, Frances Woolley, and Christopher Worswick (2010) use the DiNardo, Fortin, and Lemieux (1996) decomposition to measure the differences in earnings between university professors in different periods, from the early 1970s to the early 2000s. Their analysis bears some similarities with ours, as we also consider a time dimension. However, the most important difference is that, in our paper, we focus on the gender-wage gap at different ages. Eva M. Sierminska, Joachim R. Frick, and Markus M. Grabka (2010) also employ the DiNardo, Fortin, and Lemieux (1996) decomposition for Germany, using data from the SOEP to study the wealth gap, of which salary is just one of the components.

8 While data from East Germany are also available, the longitudinal dimension is substantially shorter. Moreover, the communist legacy and the process of economic transition suggest that trends in the gender-wage gap in East Germany might be driven by different factors.

[9] Data correspond to the German national sample. Immigrants are not included in the analysis.

[10] Although in Germany the minimum legal employment age is 15, in the most recent year, only 30 percent of young people entered the labor market before their 25th birthday. In 1984, however, this percentage was twice as high. Thus, analyzing the individuals under the age of 25 would have involved additional selection issues and educational choices. The employment ratio among individuals above the age of 60 remained below 10 percent during the entire sample period.

[11] The literature typically relies on cross-sectional data and hence frequently utilizes parenthood or age of children as an exclusion restriction in estimating the selection bias. However, most of the men and women in our sample eventually have at least one child. Clearly, never-parents are not directly comparable to ever-parents.

[12] Wages are taken in logarithms; hence we take 0.1€/CPI – that is, the real value in 2005 euros of 10 euro cents.

[13] Another source of bias, particularly toward the upper tail of the age distribution, stems from the fact that if motivation to work for pay is age dependent and heterogeneous across individuals and genders – the preference argument – our estimates would be biased upward because only individuals motivated enough to work for pay will be observed in the employed sample past certain age thresholds. Thus, our estimate of the age pattern in the gender-wage gap may partially confuse pure age effects and – should they indeed be heterogeneous across age and genders – preferences.

[14] The 2010 wave of the SOEP introduced a new wave of respondents to the panel. As the wave focused on families, this raised the average of couples living with children in the following years.

[15] A full results output from each of the numerous probit models is available upon request.

[16] The sample of respondents used to compare 1994 to 1989 is the same for both years; it might differ from the sample used to compare 1989 to 1984. This choice is motivated by panel attrition. If we require individuals to stay longer in the panel, up to thirty-two years, the sample size often drops below 150 observations. Estimates for those small groups are reported as a robustness check in Table D1 in the Supplemental Online Appendix.

SUPPLEMENTAL DATA

Supplemental data for this article can be accessed at https://doi.org/10.1080/13545701.2018.1503418.

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