The Reversal of the Gender Gap in Education: What Does it Mean for Gender Differences in the Relationship between Education and Health

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Abstract

Recent studies have indicated that the relationship between education and health differs between men and women. In line with the resource substitution theory, these studies demonstrate that women receive greater health returns to an additional year of schooling than men do. In this article, we argue that this female advantage in the effect of education can change across different socio-historical contexts. In particular, we investigate the impact of the reversal of the gender gap in education on gender differences in the relationship between education and health. Using data from the European Social Survey (eight waves: 2002–2016) for individuals of age 25–75 years ($N = 265,299$) in 31 European countries, we find evidence for our central hypothesis: it appears that the female advantage in the effect of education on health decreases as the gender gap reverses. According to our results, however, this situation could be seen as a cost that does not outweigh the benefits provided by the unprecedented gains that female education has brought about in terms of population health.

Introduction

The existence of an educational gradient in health is well documented. People with higher levels of education generally experience better health, less psychological distress, and fewer depressive symptoms than do those with less education (Ross and Wu, 1996; Bracke, Van de Straat, and Missine, 2014). These associations are partly due to selection processes, whereby social background or health in childhood influence both educational attainment and health in adulthood (Montez and Friedman, 2015; Lynch and von Hippel, 2016). However, studies using innovative econometric and meta-analytic techniques that adjust for these selection mechanisms (Lleras-

Muney, 2005; Silles, 2009; Ljungdahl and Bremberg, 2015) reveal enduring correlations, indicating that at least part of the association between education and health is causal. In this article, we take this causal ordering as starting point to build our theoretical framework.

Education protects against ill health through several paths. First, advancing education provides people with effective skills, a sense of personal control and acquired competences that can benefit health (i.e. *the socialization function of education*) (Ross and Mirowsky, 1999). Second, education leads to favourable positions in the labour and marriage markets, which help to improve an individual's standard of living (i.e. *the allocation function*

of education) (Mirowsky and Ross, 2005). On the labour market, people with higher levels of education are generally at lower risk for unemployment, and they are more likely to have access to high-status and well-paid work and secure employment that provides more intrinsic and extrinsic rewards (Allmendinger, 1989). All of these benefits are well-known contributors to better health (Mirowsky and Ross, 2005). On the marriage market, people with higher qualifications are more likely to marry and stay married, thereby experiencing intimate support (Chiappori, Iyigun, and Weiss, 2009). In this regard, marriage constitutes a buffer against ill health (Schoenborn, 2004). In summary, the socialization and allocation functions of education reinforce each other, resulting in an accumulation of educational benefits with regard to health (Mirowsky and Ross, 2005).

In recent decades, a growing body of studies has provided evidence that the effect of education on health differs between men and women (Marmot *et al.* 1997; Ross and Mirowsky, 2006, 2010). Social-epidemiological researchers have generally found evidence for the ‘resource substitution theory’, which implies that the health returns to education are greater for women than they are for men. As described by Ross and Mirowsky (2010: p. 1), the general idea behind this theory is that ‘when resources substitute for each other, the presence of one makes the absence of another less harmful. Conversely, the less there is of one resource, the more important another will be’. Since women generally have fewer alternative resources (such as power, authority and earnings) compared to men, women depend more heavily on education to achieve good health. We nevertheless argue that such studies are not complete without considering the socio-historical context in which the gender differences in the relationship between education and health unfold. One fundamental historical shift appears to be highly relevant in this regard: the reversal of the gender gap in tertiary education.

Since the 1960s, participation in higher education has increased considerably in the Western world and beyond (Schofer and Meyer, 2005). However, institutions of tertiary education have witnessed a far greater increase in the number of female students than in male students, such that men now generally lag behind women in terms of several key educational benchmarks (Vincent-Lancrin, 2008; DiPrete and Buchmann, 2013). Drawing on data from eight cross-sectional waves (2002–2016) of the European Social Survey (ESS), we investigate the implications of this reversed gender gap in education for gender differences in the effect of education on health.

The corresponding hypothesis is that the female advantage in the effect of education decreases as the gender

gap reverses. This expectation is inspired by theories from economic and social stratification literature, which indicate that the absolute value of education has decreased in the context of educational expansion. With the massive growth in tertiary education in European countries, nominal levels of education have gradually devalued from one birth cohort to another (Ortiz and Rodríguez-Menés, 2016). Given the steeper rise in female education, this process of credential devaluation may be more pronounced among women than among men, which in turn may pose limits to education as a compensatory instrument for women to achieve good health. Although, we are convinced that this scenario is likely to have occurred, the background section also considers arguments that operate against the central hypothesis.

Background

Arguments in Support of the Central Hypothesis

Both the notions of ‘education as a positional good’ (Bol, 2015) and ‘credentialism’ (Collins, 1979) argue that educational expansion has considerably decreased the absolute value of education. As a concrete example, they suggest that, the 1950s, an individual who had completed 16 years of education could reap greater benefits from the degree than would be possible for an individual who had completed 16 years of education in the 1990s.

Sociologists following the positional good perspective stress that an additional year of education adds no additional occupational value to the process unless it confers a better position in the ‘educational queue’ relative to others (Bol, 2015; Ortiz and Rodríguez-Menés, 2016). It is argued that employers reward the highest credentials because they regard them as a signal of productivity (Arrow, 1973). When hiring, employers have only imperfect information about the knowledge and skills that jobseekers bring with them to the labour market. For this reason, they screen applicants based on crude signals that indicate low training costs. In that sense, individuals are constantly in competition to stay ahead of the queue (Bol, 2015). A corollary of this positional competition is that in the context of educational expansion, over-education becomes a persistent phenomenon in the labour market: people generally invest in levels of education that are not required for their future jobs. Such mismatches between the labour market and education have previously been shown to decrease returns to absolute levels of education (Gesthuizen and Wolbers, 2010; Bracke, Van de Straat and Missine, 2014).

Collins' (1979) theory on credentialism differs from the 'positional good model' in that it considers the link between educational degrees and productivity irrelevant (Di Stasio, Bol, and Van de Werfhorst 2015). One of Collin's (1979: 19) central claims is that 'what is learned in school has much more to do with conventional standards of sociability and propriety than with instrumental and cognitive skills'. In line with this view, social scientists argue that the expansion of education has brought people to value educational degrees, regardless of the skills and knowledge acquired, as the institutionalization of education is a system of legitimation (Meyer, 1977). Many scholars thus now speak of education as an 'artificial good', representing a symbolic value rather than being of functional use (Ross and Mirowsky, 1999).

Whatever the differences there might be between the positional good perspective and the credentialism perspective, both theories indicate that the historical increase in advanced schooling has weakened the absolute value of education. People born after the onset of expansion in educational systems may therefore realize lower returns, both economic (e.g. in terms of occupational prestige or wages) (Bol, 2015) and non-economic (e.g. in terms of relevant health knowledge) (Lynch, 2006) from similar levels of education than would have been the case for people born in the first half of the 20th century (Delaruelle, Buffel, and Bracke, 2015). However, this pattern is even more applicable for women, as they have experienced the steepest increase in educational level (Sandefur and Park, 2007; Blossfeld *et al.* 2015). In other words, given that the process of educational expansion has favoured women over men, women in particular may be exposed to educational devaluation, which in turn may have posed limits to the theory of resource substitution.

Arguments Counterintuitive to the Central Hypothesis

Although the above described scenario is likely to have occurred, the literature outlines two consequences of the gender reversal for demographic behaviours that are counterintuitive to the central hypothesis. First, evidence seems to indicate that the dramatic increase in female education (Grow, Schnor and Van Bavel, 2017) and, more broadly, the rising levels of gender egalitarianism (Kalmijn, 2013) are combining to make college-educated women more attractive to men than has been the case in the past. Several decades ago, higher qualifications were of little value to women in the marriage market. This finding should be interpreted within the framework of gender

specialization theory (Becker, 1981), which stereotypes men as 'breadwinners' and women as 'primary homemakers and caregivers'. In contexts characterized by such traditional concepts, highly educated men should be particularly likely to 'marry down' in terms of education, as a means of ensuring that their mates would be willing to take care of the children and household responsibilities (Kalmijn, 2013; De Hauw, Grow and Van Bavel, 2017). In response to the substantial increase in female participation in tertiary education, however, gender specialization has gradually faded into the background, and education has gained trading value on the marriage market for women. As the gender gap reversed, men became less likely to favour wives with less education and more likely to marry women who could bring resources to the union (Goldscheider, Bernhardt and Lappegård, 2015). Consequently, in addition to gaining improved access to health-improving positions in the marriage market, women with higher levels of education have experienced substantial increases in household income (Klesment and van Bavel, 2017).

The improvements in the position of college-educated women on the marriage market are even more remarkable in light of changes in the likelihood of staying married (Schwartz and Han, 2014). It is important to consider this aspect when examining health, as the experience of divorce is known to be negatively related to mental and physical well-being (Prigerson, Maciejewski and Rosenheck, 1999). In the past, women with higher levels of education were more likely to divorce. At present, however, marriage dissolution is more common amongst those with lower levels of education. Grow and colleagues (2017) explain this trend by emphasizing the newly emerging opportunities for meeting on the marriage market. The reversal of the gender gap in education has altered the number of alternative mates from which highly educated people can choose for purposes of re-partnering. Over time, the pool of potential marriage partners with more education that are available to college-educated women has decreased, while the pool of similar options for highly educated men has increased. This shift might have reduced divorce rates amongst women with higher levels of education, while increasing marital dissolution amongst people with less education.

A second consideration has to do with an increase in opportunities that highly educated women have for participating in the labour market. Whereas their male counterparts have always been active in the labour force, it is only relatively recently that women with high levels of education have gained access to the labour market

(Brynin and Francesconi, 2004; England, Gornick, and Shafer, 2012; Klesment and van Bavel, 2017). One possible explanation for this trend is related to the reversal of the gender gap in education. More specifically, as female education has increased, the institution of marriage has shifted from one based on gender specialization (with wives being responsible for childcare, regardless of their level of education) to one based on collaboration (with the division of labour being negotiated according to the principles of opportunity-cost theory) (Brynin and Francesconi, 2004; Schwartz, 2010). In fact, as women have attained higher credentials, they have become more likely to strive for equal employment opportunities. In line with the ‘opportunity-cost’ argument, women with higher levels of education are now more career-oriented, as they stand to gain more from being employed than they would from staying at home. According to a recent study by Klesment and van Bavel (2017), the ‘motherhood penalty’ no longer constitutes an important opportunity-cost in the decision-making processes of highly educated women with regard to participation in the labour market.

In summary, the gendered expansion of education has yielded substantial improvements in the marriage outcomes and employment opportunities of well-educated women relative to well-educated men. This in turn may mask evidence for our central hypothesis.

Data and Methods

Data

For our analysis, we use data from eight waves (2002–2016) of the biennial ESS. The data of the ESS are obtained through face-to-face interviews and are representative of the non-institutionalized populations of between 21 (Wave 7) and 31 (Wave 4) countries. National response rates vary considerably by countries and waves, ranging from 31.4 per cent for Germany in 2014 to 81.43 per cent for Bulgaria in 2010. We restrict the present sample to the population between the ages of 25 and 75 years, excluding respondents from the countries that were observed in a single period (i.e. Albania, Latvia, Romania, and Kosovo). In addition, Estonia is removed from the sample, as this country has a substantial influence on the results. A complete overview of the countries in the sample and the data periods considered is provided in Appendix A. Respondents with missing data on self-rated health, education, gender, employment, marital, and cohabitation status were omitted from the sample (4.8 per cent, $N = 13,307$). In all, the pooled data set contains information from respondents 265,299.

Measurements

Dependent variable

Health is assessed by the question, ‘How good is your health in general?’ Respondents could choose from the following answer categories: very good, good, fair, bad, or very bad. Earlier studies have underscored the validity of self-reported assessments (Bracke, Van de Straat and Missine, 2014). Answers are coded from 0 to 4, with higher scores indicating better health.

Independent variables—individual level

For *gender*, men form the reference category. *Education* is assessed continuously to enhance the comparability of educational systems between countries (Berigan and Irwin, 2011). The impact of outliers is minimized by equalizing all scores higher than the country-year-specific mean plus 3 standard deviations. In addition, the upper limit is set at 25 years. Employment status, economic hardship, marital status, cohabitation, and the number of children living in the household are regarded as individual-level mediators, as they have previously been shown to be correlated to both education and health (Lynch, 2006; Zajacova, Montez and Herd, 2014). *Employment status* contrasts employed people (reference category) with unemployed people, who are further categorized as: student; unemployed, looking for work; unemployed, not looking for work; permanently sick or disabled; retired; housework; or a remaining category for community or military services and ‘other’. *Household income* is used as an indicator of economic hardship. The net household is weighted based on the modified Organisation for Economic Co-operation and Development (OECD) scale (OECD, 2013), which assigns a weight of 1 to the first adult in the household, 0.5 to all other adults (>14 years old) and 0.3 to children (<14 years of age). To enhance the comparability of the weighted household incomes across Europe’s high-income and middle-income countries, we create five household income categories relative to each country’s median income: lowest income (<50 per cent of the median income), modest income (>50 per cent and <80 per cent of the median), high income (>80 per cent and <120 per cent of the median [reference category]), and missing values. *Marital status* indicates whether a respondent was married or in a civil partnership (reference category), divorced or separated, widowed, or single. We also include a dummy for *cohabitation with a partner* regardless of marital status (no = reference category). *Number of children living in the household* is split into

two variables, one representing children younger than 12 years and the other representing children between the ages of 12 and 21 years. Three control variables are added: age, cohort, and parental education. Age is measured in years. *Birth cohort* is entered as a categorical fixed effect, clustered into 5-year groups to ensure that the micro-level data derived from the ESS can be linked to the macro-level data derived from the Barro–Lee database. Age and cohort are both included in the models to control for mutual confounding (Bell and Jones, 2014)¹. In line with previous research (Triventi, 2011), we adjust for *parental education* as a general indicator of social background. It indicates the highest level of education completed by the respondent's father or mother, and it is operationalized according to six categories: primary education (reference category), lower secondary education, upper secondary education, post-secondary non-tertiary education, tertiary education, with an additional category for missing values.

Independent variables—macro level

To assess the impact of the *reversed gender gap in education*, we define an indicator at the country-cohort level to measure the ratio of female-to-male educational attainment. The raw data for measuring this macro-level variable are derived from the Barro–Lee database (see: <http://www.barrolee.com/>), which provides country-specific averages of years of total schooling at 5-year intervals from 1950 to 2010 disaggregated by 5-year age groups and by gender. The indicator is created in three steps. First, we calculate the corresponding birth cohort for each age group in each time period to make it possible to attach the information to the exact matching country-cohort in the ESS. Second, we average the country-cohort information over the year 2005–2010, since these years fall within the period range covered by the ESS. Third, within each country-cohort, the female-specific average is divided by the male-specific average. Consequently, values greater than 1 indicate a gender gap in education favouring women. Appendix B depicts the frequency distribution of this measure.

Analysis

To test our hypothesis, we draw on a two-level multilevel design in which 265,299 respondents are nested within 339 country-cohorts (estimated in MLwiN version 2.35),² combined with country-level and cohort-level fixed effects. A similar design is used by Bol (2015) and Scheeren and colleagues (2018). The multilevel clustering at the country-cohort level ensures that standard errors are correctly estimated by recognizing that

individuals observed in the same cohort in the same country have more in common than do individuals observed within the same cohort but in a different country (Schmidt-Catran and Fairbrother, 2015). Another advantage of this strategy is that it allows the exploitation of contextual information to explain country-cohort variation. In addition, since we are neither interested in cross-national variation nor in general variation across cohorts, we are able to include country-level and cohort-level fixed effects. The incorporation of country dummies in the fixed part of the equation protects against bias due to unobserved cross-country heterogeneity (e.g. institutional differences in educational and health systems, or differences in family-oriented policies). The inclusion of a cohort-fixed effect excludes the possibility of spurious results due to general cohort trends that may have a similar impact on population health across all of the European countries observed. In sum, the analyses thus exploit the within-country variation across cohorts to model the impact of the female-to-male educational attainment ratio. The model is represented by the following equation:

$$y_{ij} = \beta_0 + \beta_1 X_{ij} + \beta_2 Z_j + \beta_3 X_{ij} Z_j + \beta_4 \zeta + \beta_5 \eta + \mu_j + e_i, \\ e_i \sim N(0, \sigma_e^2), \quad \mu_j \sim N(0, \sigma_\mu^2),$$

where β_0 is the average health score across individuals and country-cohorts when all predictor values equal 0. The term X_{ij} represents a vector of individual-level variables (including education), and Z_j represents the variable at the country-cohort level (i.e. the indicator that captures the reversed gender trend in education). Cross-level interaction terms are exemplified by $X_{ij} Z_j$. Furthermore, ζ represents country-fixed effects, and η represents cohort-fixed effects. Finally, the random effects (μ_j and e_i) are distributed independently and normally with zero means and constant variances. To test whether our results are sensitive to the modelling approach that we employ, we re-estimate the models using a full multilevel model in which individuals are nested within country-cohorts which are in turn cross-classified within countries and cohorts. As can be seen in Appendix C, similar results are obtained.

The analyses are constructed stepwise, starting with the estimation of a null model, containing only the intercept. In the first model, we add the parameters of gender, education, and the female-to-male educational attainment ratio. The second model includes the overall interaction between education and gender, while the third model examines the gender differences in relation to the reversed gender gap in education, as it contains

the three-way interaction term for education, gender, and health. In the final model, we add potential mediators. All of the models are controlled for parental education, age, and fixed effects for country and cohort. To facilitate the interpretation of the unstandardized coefficients, continuous variables are divided by 10 and centred on the grand mean. Finally, we apply design weights to account for differences in sampling probability between respondents, but we do not include population weights.

Results

An overview of the descriptive statistics is provided in Appendix D. The mean score for self-rated health is 2.8 (standard deviation = 0.9) (range 0–4). Slightly more than half (53.7 per cent) of the respondents in the sample are female, and the mean years of full-time education completed is 12.5 (standard deviation = 4.0). Furthermore, Figure 1 displays the female-to-male educational attainment ratio across cohorts for each country separately and for the pooled data. As evident in these graphs, most of the countries have experienced a reversal in the gender gap in education. The overall reversal in the average number of years of education appears to have taken place within the cohort range 1966–1970.

The results of the multilevel analyses are displayed in Table 1. Before turning to the actual testing of the hypothesis, we discuss the variance decomposition of the null model. As can be seen, 25.0 per cent of the variance in self-rated health is at the country-cohort level (i.e. $\rho = [\sigma^2_{\text{country*cohort}}]/[\sigma^2_{\text{country*cohort}} + \sigma^2_{\text{individual}}]$). This is relatively large, which supports the relevance of using country-cohorts as a contextual level. Next, we report the main results of the different models. All models are controlled for country-fixed dummies, cohort-fixed dummies (with younger cohorts reporting better health), age (with younger respondents reporting better health), and parental education (with respondents whose parents had completed primary education being particularly likely to report bad health).

The results for Model 1 clearly indicate that, on average, men report better health than women do ($b = -0.072$, standard deviation = 0.005). These results also reveal a positive relationship between educational attainment and self-rated health. Other things being equal, people with higher levels of education tend to have better health ($b = 0.328$, standard deviation = 0.009). The significant positive association of the female-to-male educational attainment ratio with health implies that population health generally increases as the gender

imbalance in education reverses ($b = 0.402$, standard deviation = 0.150).

In line with previous research, the results of Model 2 indicate that the association between education and health is stronger for women than it is for men. Whereas the health of men increases by 0.028 point scores ($b = 0.280$ —divided by 10, standard deviation = 0.011) with each additional year of education, women's health increases by 0.037 ($b = 0.280 + 0.090$ —divided by 10, standard deviation = 0.012). However, Model 3—which includes the three-way interaction term—shows that this pattern is less marked in country-cohort combinations in which the gender gap in education has reversed in favour of women. This can be seen in Figure 2, which depicts the coefficients of the gender*education interaction term by values of the female-to-male educational attainment ratio. It appears that the stronger the reversal in the gender gap in education, the smaller the female advantage in the effect of education on health.

In Model 5, we add the mediating variables. All of these effects are significant and in the expected direction. Higher self-rated health scores are found for employed people, married people, people living with a partner, and people with higher household income. Moreover, each additional child living in the household is associated with an increase in self-rated health. Controlling for these predictors does not change the negative direction of the interaction between gender, education, and the ratio. One surprising outcome, however, is that the coefficient for this three-way interaction in Model 5 is greater than the coefficient in Model 4. This indicates the presence of a suppression effect,³ as visualized in Figure 3. More precisely, Figure 3 plots gender differences in self-rated health ($SRH_{\text{women}} - SRH_{\text{men}}$) across different levels of education, calculated, respectively, from Model 4 (without mediators) and Model 5 (including mediators). Panel A represents a context in which men are more highly educated on average (defined as the lowest 25th percentile in the female-to-male odds ratio), while Panel B represents a context in which women are more highly educated on average (defined as the highest 25th percentile in the female-to-male odds ratio).

Several observations can be derived from this figure. First, it clearly reveals, once more, that the female advantage in the effect of education on health is stronger in country-cohort combinations in which men have more education on average than is the case in country-cohort combinations in which the gap has reversed in favour of women (i.e. differences in self-rated health between well-educated women and well-educated men are

Figure 1. The female-to-male educational attainment ratio across cohorts for each country separately and for the pooled data

Figure 1. Continued

larger in Panel A than in Panel B). Second, the clearest evidence of a suppression effect can be seen at the higher levels of education (in both Panel A and Panel B). In this case, controlling for the mediating pathways widens the health gap between women and men. We re-introduce the mediating variables stepwise to identify factors that contribute significantly to this suppression effect. The addition of household income and number of children

living in the household apparently has little impact. In contrast, the inclusion of marital status increases the effect size by approximately 0.036 point scores, whereas the inclusion of employment status increases the effect size by approximately 0.091 point scores. These results suggest that the opportunities of highly educated women on the marriage and labour markets improve more as the gender gap reverses than is the case for their male

Table 1. Self-rated health regressed on gender, education, and the female-to-male educational attainment ratio Total population of age 25–75 ($N_{cc} = 339$; $N_{individuals} = 265,299$, weighted sample)

	Null-model <i>b</i> (SE)	Model 1 <i>b</i> (SE)	Model 2 <i>b</i> (SE)	Model 3 <i>b</i> (SE)	Model 4 <i>b</i> (SE)
Intercept	2.728*** (0.025)	2.957*** (0.060)	2.960*** (0.060)	2.924*** (0.060)	3.145*** (0.059)
Parental education (primary = reference)					
Lower secondary		0.034*** (0.006)	0.034*** (0.006)	0.034*** (0.006)	0.017** (0.006)
Upper secondary		0.071*** (0.007)	0.071*** (0.007)	0.071*** (0.007)	0.039*** (0.006)
Post-secondary non-tertiary		0.072*** (0.010)	0.072*** (0.010)	0.073*** (0.010)	0.035*** (0.009)
Tertiary		0.073*** (0.009)	0.073*** (0.009)	0.074*** (0.009)	0.040*** (0.008)
Missing values		−0.050*** (0.013)	−0.050*** (0.013)	−0.050*** (0.013)	−0.045*** (0.012)
Age (/10)		−0.122*** (0.007)	−0.122*** (0.007)	−0.122*** (0.007)	−0.065*** (0.007)
Gender (men = reference)		−0.072*** (0.005)	−0.072*** (0.005)	−0.064*** (0.005)	−0.044*** (0.005)
Education (/10)		0.328*** (0.009)	0.280*** (0.011)	0.284*** (0.011)	0.163*** (0.009)
Ratio		0.402** (0.150)	0.397** (0.149)	0.260 (0.154)	0.343* (0.143)
Education × gender (/10)			0.090*** (0.012)	0.083*** (0.011)	0.094*** (0.010)
Education × ratio (/10)				0.030 (0.100)	−0.031 (0.088)
Gender × Ratio				−0.140* (0.067)	−0.102 (0.065)
Education × gender × ratio (/10)				−0.623*** (0.099)	−0.750*** (0.090)
Employment status (employed = reference)					
Student					0.021 (0.017)
Unemployed, looking for work					−0.127*** (0.010)
Unemployed, not looking for work					−0.232*** (0.016)
Permanently sick or disabled					−1.237*** (0.021)
Retired					−0.272*** (0.012)
Housework					−0.114*** (0.008)
Others					−0.212*** (0.022)
Marital status (married = reference)					
Divorced or separated					−0.040*** (0.009)
Widowed					−0.033** (0.010)
Single—never married/civil partnership					−0.005 (0.008)

(continued)

Table 1. (Continued)

	Null-model <i>b</i> (SE)	Model 1 <i>b</i> (SE)	Model 2 <i>b</i> (SE)	Model 3 <i>b</i> (SE)	Model 4 <i>b</i> (SE)
Household income (>120 per cent of median income = reference)					
<50 per cent of median income					−0.233*** (0.008)
50–80 per cent of median income					−0.156*** (0.006)
80–120 per cent of median income					−0.083*** (0.005)
Missing values					−0.056*** (0.007)
Cohabitation (no = reference)					0.023** (0.008)
Number of children younger than 12 years					0.040*** (0.003)
Number of children between the ages of 12 and 21 years					0.027*** (0.003)
Country dummies	No	Yes	Yes	Yes	Yes
Cohort dummies	No	Yes	Yes	Yes	Yes
Variance					
Country-cohort	0.212 (0.015)	0.014 (0.001)	0.014 (0.001)	0.015 (0.001)	0.012 (0.001)
Individual	0.635 (0.008)	0.615 (0.008)	0.615 (0.008)	0.615 (0.008)	0.565 (0.006)
−2*Log-likelihood	659,411.2	650,315.3	650,184.3	650,093.0	627,516.9

* $P < 0.05$; ** $P < 0.01$; *** $P < 0.001$ (two-tailed test).

Note: All continuous variables are centred on their grand mean.

counterparts. This tendency partially cancels out the decreasing pattern of the female advantage in the effect of education on health. In other words, if education had not favoured women's health-relevant life chances more in societies where the gender gap has reversed, the three-way interaction term in Model 4 would have been even more negative. Last, the figure reveals an interesting side effect that has not been considered theoretically. It appears that gender differences in health at the lower levels of education are smaller in country-cohort combinations in which women are more highly educated on average (Panel B) than is the case in country-cohort combinations in which men have more education on average (Panel A), which suggests that women at the lower end of the educational distribution benefit from the reversal in the gender imbalance in education.

Sensitivity Checks

We performed several sensitivity checks on the robustness of the findings in Model 5. The results are reported in Appendix E. First, we assessed Model 5 according to different operationalizations of education. In Column 1,

the maximum of education is set at 20, while in Column 2, education is measured categorically. Second, we also used different operationalizations of the female-to-male educational attainment ratio to check whether the results are sensitive to the chosen measure of the macro-level indicator. Column 3 presents the results of the analysis in which the impact of the reversed gender gap is measured by aggregating and dividing micro-level information of the ESS itself on years of full-time education completed. In addition, Column 4 reports the results of the analysis in which extreme values on the female-to-male educational attainment ratio are removed (extreme values fall above or below 2 standard deviations from the mean). Third, Model 5 was re-estimated for Western (Column 4) and Central/Eastern European countries separately (Column 5). The rationale is that gender differences in health have been traditionally larger in the latter group of countries than in the former group (Van de Velde, Bracke, and Levecque, 2010; Dahlin and Härkönen, 2013), which in turn may affect the potential of education as a compensatory mechanism in these countries. Fourth, we attempted to reduce the impact of mortality selection by cutting the sample at 65 years of

Figure 2. The coefficient of the interaction term between gender and education by values of the female-to-male educational attainment ratio

age. Column 6 reports the results of this check for Western European countries, while Column 7 presents the results for Central/Eastern European countries. Finally, we used a hierarchically ordered logit model to control for ‘ceiling-effect bias’, which is likely to occur, given the ordinal nature of the dependent variable. All of these additional analyses confirmed our main findings, with the exception of the results presented in Column 7. When excluding people of age 65 and older, the three-way interaction term in CEE countries is no longer significant at the 0.05 level. This is most likely due to a lack of variation in the female-to-male educational attainment ratio. As can be seen in Figure 1, most of the CEE countries (especially ex-Soviet countries) have experienced a reversal in the gender gap in education within the oldest birth cohorts. When these cohorts are removed from the sample (which is the case, as age is strongly correlated to cohort in the current study), the analysis may not be able to detect trends that are found in analyses that can exploit more variation in the core variable (such as the analysis for the pooled data or the analysis for CEE countries that includes the oldest cohorts).

In addition to the sensitivity checks, we also wanted to take account of the problem of unobserved factors. Although we focus on the reversal of the gender gap in education as the main contextual explanation, other related macro-mechanisms can drive the results (such as increases in labour force participation among women or gender-ideological transformations). Unfortunately, we are not able to control for these factors in the present analysis, as historical information is lacking or incomplete. To the best of our knowledge, there is no database available that contains macro-level measures that can be linked to birth cohorts dating back to 1930 for so many

European countries. In an effort to deal with this problem, we conducted a sensitivity analysis that uses a full fixed approach. In other words, the models include country dummies, cohort dummies, and the interactions between both. In this way, we avoid the omitted variable bias through controlling for country, cohort, and country-cohort heterogeneity. A drawback of this full fixed-effect approach is that no main effects of country-cohort variables can be included, since all the variances at this level are already explained (i.e. this approach does not allow to assess the overall effect of the female-to-male educational attainment ratio). Nonetheless, it is possible to include a cross-level interaction term (which is the main focus of our article), since this coefficient focuses only on potential non-linearities in the impact of individual-level variables (Bol *et al.*, 2014; Bennett and Moehring, 2015). As can be seen in Appendix F, the coefficients are very similar to those presented in the main analysis, indicating that unobserved factors are not confounding the main finding.

Discussion

Past studies have provided evidence to support the resource substitution theory, which argues that the beneficial effect of education on health is greater for women than it is for men (Marmot *et al.*, 1997; Ross and Mirowsky, 2006, 2010). As argued in this article, however, such studies overlook the changing socio-historical context in which the complex interaction between gender, education, and health unfolds. To address this gap, our investigation focuses on the implications of the reversed gender gap in education for the gender differences in the effect of education on health. Our central hypothesis predicts that the pattern of resource substitution becomes weaker as the gender gap in education reverses.

Similar to previous studies, the overall results are in line with the resource substitution theory (Ross and Mirowsky, 2006, 2010). Abstracting from the context, women experience greater health returns to education than men do. More importantly, however, the strength of this female advantage depends heavily on the gender distribution within the educational realm. In countries and cohorts where women on average have more education than men, the female advantage in the effect of education on health appears to be smaller. As indicated by the suppression effect, this occurs despite the fact that the reversal of the gender gap in education has yielded substantial improvements in the marriage outcomes and employment opportunities of well-educated women relative to well-educated men.

Figure 3. Gender differences in self-rated health ($SRH_{\text{women}} - SRH_{\text{men}}$) across different levels of education—calculated, respectively, from Model 4 and Model 5 in (A) country-cohort combinations in which men are more highly educated on average, and (B) country-cohort combinations in which women are more highly educated on average.

The main observed result is consistent with our expectation that the steeper rise in female education has placed women at greater risk for the effects predicted by the theories of ‘positional education’ and ‘credential inflation’. This is obviously only one explanation. However, as this study does not test the mechanisms that account for the three-way interaction effect, alternative explanations are likely to apply. One issue that will need to be addressed in future studies is whether the reversal in the gender gap affects the resource substitution pattern through changes in horizontal gender segregation in the field of education and vertical gender segregation in the occupational sector (Smyth 2005; Ortiz and Rodríguez-Menés, 2016). However, perhaps even more importantly, is to consider the observed finding from a health selection perspective. The current article builds on the observation that education has a causal effect on health, and therefore complies with the social causation theory that predominates in social epidemiological research. However, theories of social selection—which propose that dispositions and traits of individuals influence both educational attainment and health—have been offered as a serious challenge to the

presumption that educational attainment and health are causally related (Montez and Friedman, 2015; Lynch and von Hippel, 2016). From such a selection perspective, it could be that the gendered expansion in education has increased the heterogeneity more among well-educated women than among well-educated men (Montez *et al.*, 2011). In other words, as the gender gap in education reversed, the group of highly educated women may have become less selected on characteristics related to health. As a consequence then, the relationship between education and health may have weakened more among the female population mainly due to compositional changes.

Although it is obviously of interest, the fact that the reversed gender gap in education has decreased the advantage that well-educated women have over well-educated men in terms of health is but one small part of a larger picture. Restricting our focus to this aspect while ignoring the broader picture could lead to the erroneous impression that the massive advances in female education have done more harm than good. This is not the case. As indicated by the stepwise regression analysis in this study, the reversal in the gender imbalance in

education has contributed to the broader goal of population health gains. These improvements are highly likely a by-product of disproportional advances in the health of women with low levels of education (as suggested by Figure 3). Until further studies are conducted, the reason for this trend will remain speculative. It is nevertheless plausible that spillover effects are at play. It could be that the unprecedented growth in female tertiary education has generally motivated women at the lower end of the educational spectrum at least to complete secondary education. In addition, as documented in previous studies, the transition from primary to secondary education is more important for improving health than is the transition from the initial stages of tertiary education to more advanced stages (Groot and van den Brink, 2007; Bracke, Pattyn, and Von dem Knesebeck, 2013). In conclusion, our results imply that the reversal of the gender gap has been accompanied by positive externalities for population health, albeit to a certain extent at the cost of women at the higher end of the educational distribution. We therefore encourage policymakers to make further investments in female education while guarding against side payments from winners to losers.

Although the results reported here appear robust, several limitations should be acknowledged. First, measuring health according to self-reports could potentially result in the over-estimation or under-estimation of the actual prevalence of poor health (Subramanian, Huijts and Avendano, 2010). Future studies should address this concern by incorporating objective health indicators, such as chronic conditions or functional limitations. In addition, they should focus on mortality as an outcome, since previous research (Montez *et al.*, 2009) has shown that the association between education and mortality is stronger for men than it is for women, unlike the association between health and education. Therefore, another line of reasoning may be applicable with regard to this outcome variable. Second, as mentioned above, our study implicitly assumes a causal effect of education on health. However, this assumption cannot be granted, given that the cross-sectional nature of the ESS restricts our capacity to rule out health selection processes (as well as any changes that occur within them). We have attempted to reduce such potential selection effects by controlling for parental education. Although this is an important confounder, selection may operate on other covariates, including childhood health environment and intelligence (Hayward, Hummer and Sasson, 2015). Future studies could provide more certainty by examining the expansion of compulsory education as a social experiment with which to test the direction of the relationship (Oreopoulos, Page and Stevens, 2003). A third

limitation of this study has to do with another form of potential selection bias stemming from the structure of our data set: mortality selection. Given the short time span of our measurements, low values on the female-to-male educational attainment ratio partially correspond to the population of older respondents, who are vulnerable to selective survivorship. Even if this bias was at play, however, it would have led only to the underestimation of the positive impact of the rise in female education on population health. Finally, a minor yet noteworthy shortcoming of the present study is that the ESS data do not allow any fine-grained distinctions between formal marriage and civil partnership (the two statuses are grouped into the category of ‘married’). Considering these statuses as separate categories would increase the accuracy of the test, given that they signify different things in different countries (Van Bavel, 2012).

Even with the limitations noted, this study makes a significant contribution to existing literature concerning the contextualization of gender differences in the relationship between education and health. Our findings contribute to the further development of a macro-sociology of education and health by demonstrating that education is not exclusively an acquired social characteristic of individuals. Education is a societal institution, and its expansion has the potential to transform societies, permeate various life domains, and thus have profound effect on population health. All macro-sociological theories aim to explain transformations in societies. With regard to the social positions of women and men, the gendered expansion of education is probably one of the most pervasive transformations in recent history, having consequences for the health-relevant life chances of all people, regardless of gender, for decades to come.

Notes

- 1 Because we control for cohort and age effects in the fixed part of the equation, we cannot include dummies for the survey year to account for period differences in self-rated health due to identification problems (Bell and Jones, 2014). Results from additional analyses nevertheless support the assumption of no period effects (see also: Delaruelle, Buffel, and Bracke, 2015).
- 2 In a recent article, Schmidt-Catran and Fairbrother (2015) advocate including all observation levels in the nesting structure, to avoid misleading inferences. Given that almost every country was observed in two different periods, this would imply a need to control for country-period variation in the random

part of the equation. We ran these analyses, but they made no difference to the more parsimonious analyses presented.

- 3 To provide evidence of this suppression effect, the mediators must be related to both education and the three-way interaction term. Additional analyses show that this is the case.

Supplementary Data

Supplementary data are available at ESR online.

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