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# Gender inequality and the gender-job satisfaction paradox in Europe

Cristiano Perugini [a], Marko Vladisavljević [b,*]

[a] University of Perugia, Italy, Via A. Pascoli 10, 06123, Perugia, Italy
[b] Institute of Economic Sciences, Belgrade, Serbia, Zmaj Jovina 12, Belgrade, Serbia

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ABSTRACT

Despite being paid less than men and facing worse working conditions, lower promotion opportunities and workplace discrimination, women typically report higher levels of job satisfaction. Twenty years ago Andrew Clark (Clark, 1997) suggested that this might be due to their lower expectations, driven by a number of factors related to current and past positions in the labour market. Although this hypothesis is one of the leading explanations of gender differences in job satisfaction, cross-country research on the relationship between gender inequality and the gender-job satisfaction gap is rare and only descriptive. In this paper, we use the data from EU-SILC module on subjective well-being from 2013 to analyse adjusted gender-job satisfaction gaps in 32 European countries and we relate them to country differences in gender inequalities. Our results provide extensive and robust evidence of a relationship between exposure to more gender equal settings in the early stages of life and smaller gender gaps in job satisfaction. This corroborates the hypothesis that women who grew up in contexts with higher gender equality have expectations increasingly aligned to those of their male counterparts. Our results also show that being employed in typically male occupations enables this alignment too, whereas higher levels of education do not play a similar effect.

## 1. Introduction

In parallel with extensive evidence of women having lower wages, poorer job conditions, and being worse off in terms of discrimination, job content and promotion opportunities, female workers are frequently found to have equal or higher levels of job satisfaction than men (e.g. Clark, 1997; Sousa-Poza and Sousa-Poza, 2003; Kaiser, 2007, Blanchflower et al., 1993). After numerous confirmations of this result, this empirical finding is now referred to as the *gender-job satisfaction paradox*.

The aim of this paper is to contribute to this body of literature, along the avenue traced by Clark (1997), by providing econometric evidence that the existence and the extent of the paradox can be explained by exposure to gender unequal socio-economic settings. To this end, we use micro data from the special module on well-being of the 2013 EU-SILC in combination with various gender inequality country-level indicators. The paper adds to the existing knowledge on the topic by: (i) providing extensive and recent cross-country evidence on the existence and size of the gender-job satisfaction paradox in Europe (32 countries, year 2013); (ii) estimating the paradox and its relationship with gender inequality indicators by means of econometric methods able to account for potential misspecification and comparability issues.

Our results show that, once all observables are controlled for, higher exposure to gender-unequal settings in early stages of life corresponds to women's higher levels of reported job satisfaction. Such evidence is consistent with the idea that women's lower expectations of their job positions, shaped by growing up in gender unequal settings, might be at the root of the paradox. This interpretation is in line with the body of evidence showing that culture and institutions shape social norms, preferences and beliefs (Guiso et al., 2006; Tabellini, 2010; Lippmann et al., 2016) that, especially if formed during crucial developmental age, persistently affect individual behaviour (Alesina and Fuchs-Schundeln, 2007). We also show that, independently from the effect of unequal settings in early stages of life, employment in typically male occupations reverses the paradox (i.e., in male occupations women have lower job satisfaction), supposedly (in our interpretative framework) by enabling women to revise their beliefs and align expectations to those of men. Attaining higher levels of education, by contrast, does not play any direct role in the paradox.

The remainder of the paper is organised as follows. In the next section we review and discuss the main existing studies on job satisfaction and the gender paradox. We elaborate in particular on the contributions that explain how women's expectations and preferences are shaped by gender inequality observed in the labour market and by beliefs in gender roles. In Section 3 we describe the data, the empirical methods and

* Corresponding author.
*E-mail addresses:* cristiano.perugini@unipg.it (C. Perugini), marko.vladisavljevic@ien.bg.ac.rs (M. Vladisavljević).
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the estimated levels of the job satisfaction gender gap across Europe. In Sections 4 and 5 we augment the empirical model in order to account for the effect of exposure to (current and past) gender equality/inequality settings on the job satisfaction paradox. Section 6 summarizes and concludes.

## 2. The gender-job satisfaction paradox and gender inequality

Job satisfaction can be defined as *"a pleasant or positive emotional state that is a result of the assessment of one's job or job experiences"* (Locke, 1970). One of the most prominent researchers on job satisfaction, Andrew Clark (1996, 1997), highlights two strong reasons for the need to investigate job satisfaction: 1) it represents a measure of individual well-being, the distribution of which is one of the most central topics in economics, and 2) it is one of the best predictors of job performance (quits, absenteeism, and productivity) as well as of customer satisfaction (Rogers et al., 1994). Therefore, although being a subjective concept, and as such exposed to fundamental criticism (Kahneman, 1999; Alexandrova, 2005; Sen, 1979), job satisfaction has been proven to be significant and complementary to objective welfare indicators (e.g. Stiglitz et al., 2009).

From the perspective of psychology and management, Hulin and Judge (2003) view job satisfaction as a multi-dimensional concept that includes a cognitive and an affective component. While the cognitive component requires evaluation of actual working conditions and their comparison to workers' expectations, the affective component refers to the level of happiness and positive emotions related to the job. Economists, on the other hand, use the concept of job satisfaction as one operationalization of total utility from work (Clark, 1996). In this line of research, job satisfaction is measured via one item which is, from the perspective of a multi-dimensional structure, typically cognitive.[1] Within this framework, job satisfaction is seen as a utility function, determined by wages ($y$), working hours ($h$) and a set of job ($j$) and individual ($i$) characteristics (Clark and Oswald, 1996):

$$JS = U = U(y, h, j, i) \tag{1}$$

wherein preferences for higher income and fewer working hours are assumed. Wages influence job satisfaction in accordance with the rule of diminishing marginal utility of income, which justifies the use of the natural logarithm of earnings instead of levels in specifying the utility function (Clark and Oswald, 1996). The relation between working hours and job satisfaction, once individual and household characteristics (including income) are accounted for, is also non-linear: satisfaction grows with the hours worked, but begins to decline when the number of hours becomes excessive and burdensome (Dolan et al., 2008; Meier and Stutzer, 2008). Among other job characteristics, smaller firms, permanent contracts and work in the public sector are frequently associated with higher levels of job satisfaction due to combined effects of higher job security and higher intrinsic motivation for work (Buelens and Van den Broeck, 2007; Ghinetti, 2007; Vladisavljević, 2017). Evidence on the effects of occupations is not conclusive and depends on the variables included in the estimation (e.g., Clark, 1996). Among individual characteristics, marital status is typically associated with higher levels of job satisfaction, though evidence is still inconclusive (Gazioglu and Tansel, 2006), while the correlation of job satisfaction and age is typically U-shaped (Clark et al., 1996). When controlled for other working conditions (salary, occupation, etc.), educational effects are frequently found to be negative. This has been attributed, by some authors (Clark and Oswald, 1996), to higher work expectations of more educated people.

In addition, existing literature shows that another major factor in job satisfaction is gender, which will be our focus here. In his reference paper, Clark (1997) finds that women in the UK have higher job satisfaction than men and discusses the potential reasons behind the differences. He classifies them into five groups: 1) differences in individual and job characteristics; 2) differences in work values; 3) selection bias; 4) differences in relative income distributions; and 5) differences in expectations. According to his findings, factors 1) to 4), although relevant, cannot fully explain the gender-job satisfaction gap. On the other hand, he finds that gender differences in job satisfaction are not significant for younger and highly educated workers, workers whose mothers had professional jobs, workers in professional or managerial positions and in male-dominated workplaces. His interpretation is that women in these groups have higher expectations from work than other women, because they had different role models in early childhood or were exposed to good jobs during their work-life. This suggests that the higher female job satisfaction generally observed might be due to lower job expectations[2] resulting *"from the poorer position in the labour market that women have held in the past"* (Clark, 1997, p.342). As a consequence, the gender-job satisfaction paradox is expected to be a *transitory* phenomenon. As soon as more women are exposed to better jobs, or to contexts enabling them to overcome gender roles beliefs, they will revise their expectations upwards and the gap in job satisfaction will disappear.

The evidence of a declining job satisfaction paradox over time provided by Green et al. (2018) and Sousa-Poza and Sousa-Poza (2003) confirms that this might have been the case in the UK over the periods 1991–2012 and 1991–2000, respectively. Senik (2017) also provides corroborative results: as aspirations and promotion opportunities for men and women become more equal over time, the gender-job satisfaction gap decreases, which is why it is lower for more recent generations. As a possible explanation of the generalised decline in women's happiness observed in the last decades, Stevenson and Wolfers (2009) propose the idea that their expectations rose faster than society was able to meet them. As a result, actual experienced lives drove women's subjective well-being (both absolutely and relative to men) downwards. Similarly, Graham and Chattopadhyay (2013) identify expectations and social norms as the factors able to play a mediating role in shaping gender well-being differences. In particular, they argue that changes in norms and expectations that accompany changes in gender rights and roles seem to be associated with a decline in women's well-being, at least in the short term, as it may take time for the new norms to become established or accepted. In other words, declining levels of satisfaction might be observed when equality *de jure* rises faster than equality *de facto*.

Besides expectations, the link between gender equality and gender job satisfaction paradox could work through another mechanism – differences in job values. Previous research (e.g., Sloane and Williams, 2000; Bender et al., 2005) suggested that women have higher job satisfaction in female-dominated occupations because they attach higher value to aspects of work such as flexibility, social connections, etc., even though these jobs involve lower wages and poorer working conditions. The idea of different job values for men and women has received extensive attention (Marini et al., 1996; Neil and Snizek, 1987; Dæhlen, 2007; Gooderham et al., 2004). As beliefs on gender roles are shaped, among other things, by observed and experienced gender inequalities (e.g., Hiller, 2014; Alesina et al., 2013; Giuliano, 2017; Giménez-Nadal et al., 2019), women in less equal societies are "socialized" to put a higher value on aspects of work consistent with the role they are (supposedly) assigned by society. This translates into higher job satisfaction than men's, when controlled for wages and other individual and job

[1] Usually the reason for using only one item is the fact that researchers are interested in nationally representative data sets which, due to their size, opt for a limited number of questions. Similarly, in EU-SILC, satisfaction with work is measured over a single, global cognitive item. See more details in Section 3.1.

[2] According to other authors (Bender et al., 2005), the notion of expectations can also be understood in terms of the effects of social norms on job satisfaction: as women are socialized not to anticipate high satisfaction from work, they can be surprised by their actual experiences and therefore have higher levels of job satisfaction.

characteristics. As a consequence, the idea of a link between the gender-job satisfaction paradox and gender inequality is, from the values perspective, the same as from the expectations perspective: higher levels of observed gender inequality lead to higher levels of job satisfaction gap.

Empirical research on gender gaps in job satisfaction and its relations with gender inequality is not extensive and has mainly focused on showing descriptively that progress in gender equality corresponds to a weakening of the job satisfaction paradox. This is the case of the work, mentioned above, by Sousa-Poza and Sousa-Poza (2003). Similarly, Kaiser (2007) compares gender-job satisfaction gaps in 14 EU states and finds that the paradox does not appear in more gender equal countries such as Denmark, Finland and Netherlands, nor in Portugal, where men enjoy better working positions and have higher job satisfaction. In all other 10 countries,[3] where according to the author gender equality is lower, he finds that job satisfaction is *ceteris paribus* higher for women. Sousa-Poza and Sousa-Poza (2000) analyse gender-job satisfaction gaps in 21 countries and find that women have higher levels of job satisfaction in Great Britain, United States, Hungary and New Zealand, while in the remaining countries the gap is not statistically significant. They point out that in the countries where gender-job satisfaction gap exists women have higher "work-role outputs", such as job security, feeling that their work is useful and good relations with management and colleagues. However, as the authors emphasize, these factors cannot fully account for the cross-country differences in gender-job satisfaction gaps and they propose Clark's hypothesis as one of the potential explanations.

Our attempt here is to contribute to this literature by providing direct econometric evidence, on a cross-country basis, on the link between levels of gender inequality and the existence and extent of the paradox.

## 3. Gender differences in job satisfaction across Europe

### *3.1. Data and variables*

To estimate the job satisfaction gender gap we use the 2013 EU Survey on Income and Living Conditions (EU-SILC), which includes information on 32 European countries (28 EU members plus Norway, Switzerland, Iceland and Serbia). We selected the data for 2013 since the survey for this year included an ad-hoc module on well-being, with a question on job satisfaction. The topics of ad-hoc modules in EU-SILC rotate on a five-year basis, so the next one on well-being will be available only when data with 2018 as a reference year are released. As a consequence, our analysis here is based on a cross-country sample of individuals observed in one year only. Previous research has largely confirmed good psychometric properties of the EU-SILC module on well-being (e.g., Vladisavljević and Mentus, 2018). Moreover, EU-SILC is especially suitable for this research as it contains country-comparable, detailed information on income, hours worked, individual and job characteristics, all of which are necessary to perform the analysis. Our total sample includes 359,695 persons in working age (19–64). Of them, 124,822 enter the sample for the estimation of the gender-job satisfaction gap, which includes workers in dependent employment and excludes the self-employed, agriculture workers, workers in training and persons not responding to the question on job satisfaction.[4] The share of men and women who enter the estimation sample is approximately the same (35.4% for women and 34.0% for men), due to the higher share of men among the self-employed and unemployed, which "compensates" for the lower labour market activity of women.

Job satisfaction is measured via response on an eleven-point Likert type scale (from 0 - "not at all satisfied" to 10 - "completely satisfied") to the question "How do you evaluate your current job?" (variable PW010 in the dataset). According to Eurostat (2012), when answering the question, *"the respondent should make a broad, reflective appraisal of all areas of his/her job in a particular point in time (current situation)"*. Global cognitive operationalization of job satisfaction fully corresponds to the total work utility approach proposed by Clark (1996).

Besides gender, we use a large set of control variables to account for gender differences in individual and job characteristics, which include: log monthly wages, weekly working hours (and a dummy for working more than 50 hours), age (and its square), marital status, education, occupation, sector of employment, presence of an additional job, firm size and type of contract (permanent vs. temporary), as well the country fixed effects (for more details and definitions of the variables used, see Table A1 in the Appendix).

### *3.2. Econometric methods and baseline empirical model*

The nature of the dataset and the aims of the analysis pose important specification issues related to: (i) comparability of job/individual characteristics; (ii) sample selection; and (iii) the multilevel structure of data. In order to address the first aspect, prior to model estimation, we applied the nearest neighbour matching technique (Abadie and Imbens, 2002) that restricts the sample to men and women whose individual and job characteristics are comparable. Since gender occupational and sectoral segregation have long been established in the literature, failing to account for comparability of empirical distribution of individual characteristics can cause severe misspecification problems, which have been largely documented in the impact evaluation literature. The recent acknowledgment of such issues has led to the development of several methods which incorporate the matching framework in analyses of gender wage differences (e.g., Nopo, 2008). This method ensures that there are men and women in the sample who have comparable observable characteristics, therefore providing a more robust method for comparing their job satisfaction.

We implement the nearest neighbour matching procedure by using *Stata nnmatch* command (Abadie et al., 2004). Applied to the investigation of gender differences in job satisfaction, the procedure can be described in the following way. Within each country $k$, we consider a male[5] worker $i$ ($i=1, 2, \dots$ p), with $x_{im}$ – vector of $m$ observed covariates determining his job satisfaction. Allowing for the possibility of ties, we define $d_{mij} = \|x_{im} - z_{jm}\|$ as a multidimensional distance from the covariates of a male worker $i$ to covariates of all potential matches from the pool of female workers, where $z_{jm}$ are the values of covariates for female worker $j$ ($j=1, 2, \dots$ q). Female worker $w$, with the values of covariates $z_{wm}$ is the "nearest neighbour" of male worker $i$ if condition $\forall j, \{d_{iwm} = \|x_{\mathfrak{Z}} - -z_{wm}\| \le d_{ijm}\}$ is satisfied, i.e., if the multidimensional distance from the covariates of male worker $i$ to the covariates of female worker $w$ is lower or equal than the distance from the covariates of male worker $i$ to the covariates of all other female workers from that country.

In this paper, the nearest neighbour matching procedure is implemented within each country $k$, demanding that men and women are matched exactly ($d_{iwm} = 0, \forall m$) on: wage quintile groups, working hours groups, education, occupation, sector (two groups: industry vs. ser-

[3] Austria, Belgium, France, Germany, Greece, Ireland, Italy, Luxemburg, Spain and the UK.

[4] We dropped the self-employed due to inapplicability of the some of the questions such as firm size and temporary work (similarly to Clark, 1997) and differences in job utility determinants such as income and other working conditions (Blanchflower and Oswald, 1992). Similarly, we excluded agriculture workers as their job satisfaction can be under the strong influence of weather and other unobservable working conditions, and persons in training as their job satisfaction might be confounded with training satisfaction.

[5] We define the matching procedure from the perspective of men, but the procedure and its outcomes would be the same if we take women as the reference group, as we are requesting the exact match, and not using the procedure to estimate the "treatment" effect. Instead, we are using the procedure option that keeps the results of the matching in the database; this enables us to identify, for each observation, and regardless of whether they are in the "treated" (men) or "control" group (women), the counterpart with the same characteristics.

vices), temporary/permanent contract, and age group.[6] Since we do not want to estimate the gap as a treatment effect, but rather only to restrict our sample, we demand the procedure to choose only one nearest neighbour, and allow observations to be used as nearest neighbours more than once, which makes the matching order irrelevant. Female and male respondents who do not have exact opposite sex matches are then dropped from the sample. After matching, 83,555 out of 124,822 individuals (67.0%) are kept in the analysis.[7]

The second issue is related to selection bias. The selection of women and men into the sample of dependent employment, for which we observe job satisfaction, might not be random. Since being in dependent employment could be systematically correlated with job satisfaction, estimated coefficients from the job satisfaction equation could be biased (Clark, 1997). This is particularly important when estimating gender gaps (and their interactions with other variables), as different mechanisms could be behind female and male sample selection. For example, women dissatisfied with market jobs could opt for home work more frequently than dissatisfied men (Stevenson and Wolfers, 2009). If this is the case, the observed distribution of job satisfaction between genders would be biased. Although the share of women and men in dependent employment in our data is approximately the same, these effects need to be accounted for.

Typically, selection bias is addressed by using Heckman selection correction (Heckman, 1979), and inclusion of the "omitted variable" – Inverse Mills ratio (IMR). IMR is based on the probability to be in the estimation sample (in our case dependent employment), estimated via probit regression conditional on the set of selection variables. However, as gender equality indicators that we use in the latter parts of the analysis potentially vary with current participation rates and not with shares of dependent employment (which would be modelled in the Heckman-two stage model) we need to account for a more complex structure of the selection. To this aim, we make use of Bourguignon et al. (2007) correction of the Dubin and McFadden (1984) model that allows us to simultaneously control for several (multinomial) selection effects. Within this procedure we divide working age population (19–64) into four groups, based on the self-declared labour market status and availability of the estimation variables: (1) job satisfaction gap estimation group; (2) other employed (self-employed, workers in training or education, agriculture workers, missing values for job satisfaction, dropped from the matching)[8]; (3) unemployed; and 4) inactive. The procedure is similar to the Heckman correction as it also consists of two steps. In the first step, we estimate, via multinominal logit, the selection into one of the four groups,[9] conditional on the set of personal and household characteristics: age and age squared, education, marital status, status of the household head, number of children in the household (three variables for children aged up to two years, two to six years and seven to fourteen), number of elderly and household size. Identification of the selection model relies on the distributional assumptions of the method proposed in Bourguignon et al. (2007) [10] and is further strengthened by estimating selection equations separately for each gender and country. Based on the estimated probabilities of participation in each of the four groups, we compute the inverse Mills ratios (IMRs) as the ratios of the probability density function to the cumulative distribution function (Wooldridge, 2002). Four IMRs, each derived from the probability to be in one of the four statuses, are then added to the list of covariates in the job satisfaction equation presented below.[11]

Lastly, pooling data for different countries creates a multilevel structure of the data, in which observations at the individual level are nested within the country level. Given the nature of our dataset, and relying on Bryan and Jenkins (2016) (see also, for example, Perugini et al., 2019), we deal with this multilevel structure by: (i) implementing a fixed effect (FE) estimation approach, i.e., pooled country surveys with the inclusion of distinct country intercepts; and (ii) clustering standard errors at the relevant (country/gender/age or country/gender) level.

The basic form of the job satisfaction model is given by the following equation:

$$JS_{ik} = \alpha + \beta_1 female_{ik} + X'_{ikn}\gamma_n + u_k + IMR'_{ikm}\lambda_m + \varepsilon_{ik} \tag{2}$$

where $i$ and $k$ denote individuals and countries, respectively; $u_k$ denotes country fixed effects, $X_{ikn}$is the regressor matrix, $\gamma_n$ the vector of associated coefficients, $IMR'_{ikm}\lambda_m$ is a set of variables and associated coefficients used to correct for the selection effects and $\varepsilon_{ik}$ represents the error term. The matrix $X_{ikn}$ consists of the control variables described in Section 3.1. The coefficients $\beta_1$ next to the dummy variable for gender measures the adjusted gender-job satisfaction gap which, according to Clark (1997), is a proxy for gender differences in expectations, as personal characteristics and objective working conditions are controlled for. Model (2) is applied to estimate adjusted gender-job satisfaction gap in the whole sample and separately for every country. The set of variables that we are using is extensive and allows controlling for many important aspects; however, it is not exhaustive. Therefore, caution is needed when interpreting the estimated adjusted job satisfaction gap since some important factors that may differ across genders (e.g., job tenure) cannot be fully accounted for.

Although job satisfaction is measured on a Likert type scale, which produces ordinal type variables, results from the measurement literature (e.g., Norman, 2010; Brown, 2011) suggest that ordinary least squares (OLS) estimates do not differ in results or conclusions when applied to interval and Likert scale type measures. We therefore opt to estimate

[6] Working hours groups: part-time, full-time and overtime (while part- vs. full-time distinction is self-assessed, overtime workers are identified as full time workers who work more than 50 hours per week); education groups (primary, secondary and tertiary education, see Table A1 for details); Occupation - ISCO 1 groups (see Table A1 for details on the occupation groups); age groups: 19/24, 25/34, 35/44, 45/54, 55/64.

[7] In other words, 33% of workers cannot be matched (about 37% of men and 29% of women). In most of the cases men and women cannot be matched by occupation (26% of men and 20% of women), followed by wage quintile (16% of men and 10% of women) and age group (7% of men and 6% of women), while for other used characteristics matching does not occur in less than 1% of cases. Percentages by groups do not sum up total number of cases that cannot be matched, as the characteristics that cannot be matched are not exclusive, i.e., it is possible that a person has more than one characteristic that cannot be matched.

[8] In our estimations we use the sample constructed after the matching; the full sample is instead used for robustness checks. In this first case, workers excluded via matching from the estimation sample are included in the group for the estimation of the selection effects; in the second case this group is included in the estimation sample.

[9] Given the structure of the sample we use for the control of the selection bias (two groups of employed: employees and self-employed; and two groups of out of work: unemployed and inactive), a natural concern is whether the IIA (Independence of Irrelevant Alternatives) assumption, which is an integral part of multinominal logit estimates, is too restrictive, and whether another model, e.g. multinomial normal model is more appropriate. However, according to Bourguignon et al. (2007), if the aim of the multinomial logit is the correction of the selection bias in the outcome equation (rather than the estimation of the selection process itself) multinominal logit is, even if the IIA is severely violated, a reasonable alternative to a multinomial normal model. As multinominal logit is easier to implement then the multinomial normal model, and both yield consistent estimates, we opted for the former.

[10] Unlike the original Dubin-McFadden (1984) method, which restricts the class of allowed distributions of the main equation residuals, Bourguignon et al. (2007) correction allows main equation residuals to be normally distributed, by normalizing selection equation residuals and assuming that they are related linearly to the main equation residuals (Bourguignon et al., 2007, p. 179). As a robustness check of our results we also apply original Dubin-McFadden (1984), estimation of the first part of the model. Results based on these estimates confirm the results presented in the remaining part of the paper and are available upon request.

[11] In order to estimate multinominal selection effects we use *selmlog* stata procedure by Bourguignon et al. (2007).

the model by using the OLS method, as we are then able to compare adjusted gender-job satisfaction gaps in different countries. However, as a robustness check for the analysis of the pooled data, we use ordinal probit model and recode the job satisfaction variable into three categories: low (0–5), median (6–8) and high (9 and 10).[12]

For both procedures (OLS and probit) we use clustered standard errors at the relevant level. Although the covariates presented in the Eq. (2) are individual-level variables, variables used to test the impact of the gender equality on the gender-job satisfaction gap are of higher level of aggregation (country/gender/age or country/gender). As these indicators are constant within clusters, residuals of the observations might be correlated, resulting in biased estimates of the standard errors (Moulton, 1986). In order to account for within-cluster correlation we use parametric correction for the Moulton factor suggested by Angrist and Pischke (2009).[13] We further test the robustness of our results by applying both OLS and ordered probit estimates on the total sample, without the matching restriction.

### 3.3. Baseline estimations

Results of the estimation of the baseline model are presented in Table 1.[14] In the first two columns we report OLS estimates of Eq. (2) implemented on the matched and full sample, respectively; columns 3 and 4 report corresponding estimates from ordinal probit. All estimates include the correction for (multinomial) sample selection (IMR ratios at the bottom of the table). Generally speaking, the signs of the coefficients in the selection equations (not reported here but available upon request) are as expected and indicate that selection into dependent employment is not random: the probability of being in dependent employment compared to other labour market statuses increases with age (at diminishing rate) and education levels. For women, being married and having children decreases the likelihood of being in dependent employment, when compared to inactivity, while for unemployment and other employment statuses the effects are mixed. For men, being married and having children increases the likelihood of being in dependent employment, compared to all other three conditions. Lastly, the likelihood of being in dependent employment is lower for both genders in larger households and in households with elderly household members. The effects of selection variables, as evidenced in the Table 1, are statistically significant, regardless of the model and the sample, and suggest that selection variables do have an impact on the job satisfaction in our estimates.[15]

Results in Table 1 reveal a strong stability across alternative samples and estimators, with coefficients for covariates in Eq. (2), exhibiting largely expected signs. In line with our theoretical model, job satisfaction is higher for people who receive higher wages and lower for people working longer hours (Bender et al., 2005; Linz and Semykina, 2013). Working overtime (longer than 50 hours per week) has no stable additional effect on job satisfaction, hence non-linear effects of hours worked on satisfaction do not clearly emerge from our results. Coefficients for both age and age squared are significant, indicating a well-known U-shaped relation between age and job satisfaction (Linz and Semykina, 2012; Ghinetti, 2007), while job satisfaction is higher for married individuals (Clark, 1996; Linz and Semykina, 2012). Sectoral dummies indicate that, compared to manufacturing, workers in public administration (NACE sector O), education (P), health (Q) and arts, sports and NGOs sectors (R to U), have higher levels of job satisfaction, probably due to the combination of higher intrinsic motivation for work and higher job security (Buelens and Van den Broeck, 2007; Ghinetti, 2007). Intrinsic motivation and job security are also frequently used to explain two other results from our estimates - higher job satisfaction for working in smaller firms and higher job satisfaction for working on permanent contracts (Clark, 1996). Compared to elementary occupations, all other occupations have *ceteris paribus* higher levels of job satisfaction, the effects being the strongest for Managers (ISCO group 1) and Professionals (group 2). Lastly, in line with the argument from Section 2, after controlling for all other covariates, the effects of education are negative, indicating higher work expectations of more educated workers (Clark and Oswald, 1996; Bender et al., 2005).

As all models in Table 1 yield similar results, for the following estimates we rely on OLS, estimated on the matched sample in the following empirical steps. Results with full sample and oprobit estimates are largely consistent and available upon request.

### 3.4. Adjusted gender-job satisfaction gap in Europe

We now turn to our main interest, gender differences in job satisfaction. The coefficient for *female* in model (2) represents the so-called adjusted gap in job satisfaction, i.e., the gender differences in job satisfaction once all other observable job and individual characteristics are statistically controlled for. The estimated coefficient for the gender dummy (female = 1), regardless of the sample (matched or full) and estimation procedure (OLS or oprobit), is positive and statistically significant, indicating that on average, in the sample of 32 European countries (country-fixed effects included), women have a higher level of job satisfaction then men. However, we observe large differences across countries in the size and the sign of the job satisfaction adjusted gap (Fig. 1 and Table A3 in the appendix, in which we also report the estimated unadjusted gender gap).

Our outcomes are only partially consistent with the existing evidence reviewed in Section 2. Job satisfaction is, *ceteris paribus*, higher for men in seven countries of Central-Eastern Europe: Lithuania, Slovakia, Bulgaria, Croatia, Czech Republic, Poland and Romania, as well as for Sweden, although this difference is statistically significant only for Lithuania and Slovakia. In all other countries, women have higher job satisfaction (conditional on covariates), although the difference is statistically significant only for the UK, Iceland, Malta, Cyprus, Portugal, Estonia, Hungary, Netherlands, France and Spain.

## 4. Current gender inequality and the job satisfaction paradox

According to Clark's conjecture (1997), after controlling for covariates in model (2), the gender gap in job satisfaction reflects differences in work expectations between men and women: women tend to have lower expectations regarding their jobs, and are consequently more satisfied (than men) with the same job. As a consequence, we should expect that the paradox exists in the countries with low levels of gender equality (similarly to Kaiser, 2007), since women in these countries have lower job expectations than men.

[12] In accordance with the Eurostat analysis of job satisfaction in Europe: http://ec.europa.eu/eurostat/statistics-explained/index.php/Quality_of_life_in_Europe_-_facts_and_views_-_employment.

[13] We use *moulton* command provided at Mostly Harmless Econometrics data archive website. Retrieved from http://economics.mit.edu/faculty/angrist/data1/mhe/brl on 3/12/2018. We check the robustness of our results by using country level clustering. The results, available upon request do not change the conclusions obtained from the estimates.

[14] Descriptive statistics of the variables (including the later indicators of gender equality) included in the analysis are presented in Table A2 in the Appendix.

[15] We also perform a robustness check of the results by using a simple Heckman procedure. Results (available upon request) suggest that selection into dependent employment is not random, while the effects of selection variables are similar to the ones observed in the multinominal model. Additionally, the effect of IMR from the Heckman procedure is significant, suggesting again that selection variable has an impact on job satisfaction. We additionally estimated the effects of the squared IMRs variables from the Bourguignon et al. (2007) procedure, to account for the possible non-linearity between the IMRs and job satisfaction. However, these estimates yielded very large values of the coefficients and standard errors, suggesting multicollinearity issues of such specification. Finally, we check the robustness of these results by using the original Dubin-McFadden (1984) method of multinominal selection. Results (available upon request) also suggest that selection is not random, while the effects of selection variables are similar to the ones from the Bourguignon et al. (2007) procedure.

**Table 1**
Estimation of the baseline specification (summary of results).

| | Matched sample OLS | | Full sample OLS | | Matched sample oprobit | | Full sample oprobit | |
|---|---|---|---|---|---|---|---|---|
| Female | 0.068*** | (0.025) | 0.044** | (0.022) | 0.043*** | (0.012) | 0.029*** | (0.010) |
| ln wage | 0.353*** | (0.017) | 0.318*** | (0.013) | 0.189*** | (0.010) | 0.167*** | (0.007) |
| Hours | −0.010*** | (0.001) | −0.004*** | (0.001) | −0.006*** | (0.001) | −0.003*** | (0.001) |
| Hours50 | −0.056 | (0.038) | −0.118*** | (0.025) | 0.015 | (0.024) | −0.015 | (0.016) |
| Age | −0.053*** | (0.008) | −0.058*** | (0.007) | −0.034*** | (0.005) | −0.037*** | (0.004) |
| $Age^2$ / 100 | 0.051*** | (0.010) | 0.057*** | (0.008) | 0.034*** | (0.006) | 0.037*** | (0.005) |
| Married | 0.083*** | (0.017) | 0.079*** | (0.014) | 0.028*** | (0.011) | 0.032*** | (0.009) |
| *Primary education* (omitted) | | | | | | | | |
| Secondary education | −0.289*** | (0.030) | −0.210*** | (0.022) | −0.183*** | (0.019) | −0.134*** | (0.013) |
| Tertiary education | −0.416*** | (0.039) | −0.337*** | (0.028) | −0.254*** | (0.024) | −0.203*** | (0.017) |
| Managers | 0.822*** | (0.044) | 0.726*** | (0.033) | 0.483*** | (0.028) | 0.418*** | (0.020) |
| Professionals | 0.708*** | (0.035) | 0.622*** | (0.027) | 0.401*** | (0.023) | 0.349*** | (0.017) |
| Technicians | 0.603*** | (0.033) | 0.530*** | (0.025) | 0.339*** | (0.022) | 0.295*** | (0.016) |
| Clerks | 0.492*** | (0.035) | 0.434*** | (0.027) | 0.265*** | (0.022) | 0.233*** | (0.016) |
| Service / sales workers | 0.390*** | (0.031) | 0.339*** | (0.025) | 0.210*** | (0.021) | 0.183*** | (0.016) |
| Craft / trades workers | 0.082** | (0.038) | 0.207*** | (0.028) | 0.039 | (0.024) | 0.114*** | (0.017) |
| Plant / mach. operators | 0.172*** | (0.040) | 0.221*** | (0.028) | 0.085*** | (0.025) | 0.117*** | (0.018) |
| *Elementary occupations* (omitted) | | | | | | | | |
| *Sectors B-E* (omitted) | | | | | | | | |
| Sector F | 0.107*** | (0.038) | 0.037 | (0.027) | 0.047** | (0.023) | 0.017 | (0.016) |
| Sector G | −0.060** | (0.028) | −0.020 | (0.021) | −0.044** | (0.018) | −0.021 | (0.013) |
| Sectors H - I | 0.048 | (0.031) | 0.059*** | (0.022) | 0.026 | (0.019) | 0.033** | (0.014) |
| Sectors J - K | −0.059* | (0.031) | −0.041 | (0.025) | −0.040** | (0.019) | −0.025* | (0.015) |
| Sectors L - N | −0.025 | (0.031) | −0.010 | (0.024) | −0.006 | (0.019) | 0.006 | (0.015) |
| Sector O | 0.210*** | (0.028) | 0.254*** | (0.022) | 0.122*** | (0.017) | 0.146*** | (0.013) |
| Sector P | 0.398*** | (0.029) | 0.419*** | (0.024) | 0.240*** | (0.018) | 0.251*** | (0.014) |
| Sector Q | 0.282*** | (0.029) | 0.309*** | (0.023) | 0.156*** | (0.018) | 0.170*** | (0.013) |
| Sectors R - U | 0.374*** | (0.040) | 0.441*** | (0.031) | 0.212*** | (0.025) | 0.245*** | (0.020) |
| Additional job | 0.040 | (0.030) | 0.030 | (0.025) | 0.038** | (0.019) | 0.028* | (0.015) |
| *Firm size 1–10* (omitted) | | | | | | | | |
| Firm size 11/19 | −0.103*** | (0.022) | −0.086*** | (0.018) | −0.053*** | (0.014) | −0.044*** | (0.011) |
| Firm size 20/49 | −0.102*** | (0.022) | −0.080*** | (0.018) | −0.069*** | (0.014) | −0.054*** | (0.011) |
| Firm size 50+ | −0.157*** | (0.019) | −0.131*** | (0.016) | −0.093*** | (0.012) | −0.081*** | (0.010) |
| Temporary contract | −0.318*** | (0.030) | −0.257*** | (0.020) | −0.151*** | (0.020) | −0.121*** | (0.013) |
| IMR1 (dep. employ.) | −0.114* | (0.059) | −0.064 | (0.047) | −0.075** | (0.033) | −0.038 | (0.026) |
| IMR2 (oth. employ.) | −0.438*** | (0.155) | −0.294** | (0.129) | −0.233*** | (0.079) | −0.156** | (0.063) |
| IMR3 (unemployed) | 0.768*** | (0.176) | 1.068*** | (0.143) | 0.469*** | (0.096) | 0.598*** | (0.074) |
| IMR2 (inactive) | −0.270** | (0.126) | −0.169 | (0.104) | −0.137** | (0.069) | −0.077 | (0.057) |
| Constant (cut 1) | 6.491*** | (0.301) | 6.793*** | (0.253) | −0.808*** | (0.142) | −0.961*** | (0.112) |
| (Constant cut 2) | | | | | 0.865*** | (0.142) | 0.676*** | (0.112) |
| (pseudo) r square | 0.0899 | | 0.0607 | | 0.04 | | 0.0374 | |
| Observations | 83,555 | | 124,822 | | 83,555 | | 124,822 | |

*Notes:* Country-fixed effects omitted from the table and available upon request. Standard errors clustered at country/gender level (parametric correction for Moulton factor); Estimated coefficients from the oprobit indicate the sign and significance of the coefficient, but cannot be interpreted as marginal effects. Significance levels: ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$.

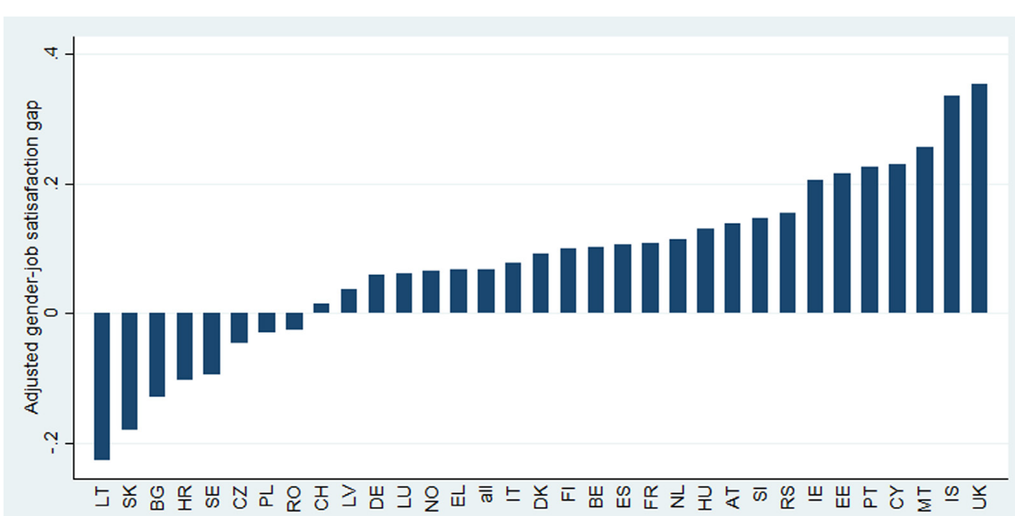


**Fig. 1.** Adjusted gender-job satisfaction gap (female to male job satisfaction) by country.
*Source:* Authors' elaboration on SILC data.
*Note:* Full tables with estimations by single countries are available from the authors upon request.

**Table 2**
Job satisfaction gender gap and current level of gender equality (OLS, matched sample).

| | 1 | 2 | 3 | 4 |
|---|---|---|---|---|
| Female | 0.068*** | 0.068*** | 0.073*** | 0.066*** |
| | (0.025) | (0.025) | (0.025) | (0.025) |
| Female * Overall Gender Equality Index | | −0.053 | | |
| | | (0.373) | | |
| Female * Economic Participation and Opportunity Score | | | −0.387 | |
| | | | (0.281) | |
| Female * Political Empowerment Score | | | | 0.059 |
| | | | | (0.122) |

*Notes:* Gender equality indicators normalized at mean (i.e., variables have the mean at zero, while preserving the original variation). Multinomial selection effects included. Full models results are presented in Table A4 in the appendix. Standard errors clustered at country/gender level (parametric correction for Moulton factor). Significance levels: ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$.

To test the effects of gender equality explicitly, in Table 2 we augment Eq. (2) with an interaction of the gender dummy variable with measures of gender equality. Indicators of current gender equality at country level are taken from the World Economic Forum's Global Gender Gap Report 2013. We consider, in the first place, the Gender Gap Index, which varies between 0 (maximum inequality) and 1 (equality); this overall index is calculated as the un-weighted average of four sub-indexes (again ranging from 0 to 1) which describe four main gender gap dimensions: economic participation and opportunity, educational attainment, health and survival, and political empowerment. The four sub-indexes are calculated as weighted averages of 14 different indicators that form them (see WEF, 2013, for details on the methodology and the base indicators).

The use of country fixed effects obviously prevents the inclusion of additional country-level predictors in the empirical model, since the country intercepts already fully encapsulate cross-country differences (Snijders and Bosker, 1999). However, additional country-level variables can be interacted with individual-level variables so as to obtain the additional effect that a country-level factor produces on the main (individual-level) effect. The augmented model therefore reads:

$$JS_{ik} = \alpha + \beta_1 fem_{ik} + \beta_{11} fem_{ik} * GenEq_k + X'_{ikn}\gamma_n + u_k + IMR'_{ikm}\lambda_m + \varepsilon_{ik} \tag{3}$$

where $\beta_{11}$ represents the impact of currently observed gender equality on gender-job satisfaction gap and all other coefficients and variables are the same as in model (2). If higher gender equality contributes to increase women's job expectations (hence lowering their job satisfaction), we should observe a negative sign for $\beta_{11}$.

Table 2 (see Table A4 in the Appendix for complete results) indicates that the current level of gender equality in the country, as measured by the WEF global gender index, is not significantly correlated with the job satisfaction gap. Replacing the global index with its two sub-components that show sufficient variability across countries[16] also indicates that the current level of gender equality has no impact on the gender job satisfaction gap.

One reason why a link between gender inequality and the job satisfaction gap does not emerge in Table 2 could be related to the fact that observed contemporaneous gender inequality is not sufficient to align expectations of women to those of men. This might occur if, as emphasised by the extensive literature reviewed in Section 2, beliefs and preferences built and internalised in early stages of life tend to be persistent over time. Should this be the case, the currently observed low levels of gender inequalities might not be sufficient to revise upwards the low expectations built in early stages of life by having experienced (and lived in context with) high gender inequality. In the following section we empirically test this possibility.

## 5. Past gender inequality and the job satisfaction paradox

### 5.1. Job satisfaction and gender inequality in early stages of life

The nature of our sample is particularly suited to our aims, since it includes European countries with very different histories of gender equality, strictly related to their political and ideological systems. As a consequence, people in a similar age and with similar characteristics, but who grew up in different countries, might have experienced very different gender inequality settings during their early stage of life. The emphasis on economic and social equality was a hallmark of the socialist ideology; before the transition to market economy started in 1989, countries of Central and Eastern Europe were actually able to maintain remarkably equal distributions of income and were often identified as the most equal countries in the world (Atkinson and Micklewright, 1992). In particular, equality of men and women was proclaimed as one of the key ideological tenets of socialism (Little, 2011), deeply rooted in the thinking of the founding fathers and emphasized as a key achievement of overcoming capitalism which, by nature, favoured women's oppression (see, for example, Friedrich Engels in his 1884 book, The Origin of the Family, Private Property and the State). Even though horizontal and vertical gender segregation still penetrated many fields of social life (Jurajda, 2003 and 2005; Pollert, 2005) and family loads were largely asymmetric (La Font, 2001; Gal and Kligman, 2000), women's participation in the labour market and their access to education, healthcare and political life were incomparably higher compared to Western Europe (Blau and Ferber, 1992; Brainerd, 2000). It is widely documented that this contributed to the development of remarkably different attitudes and beliefs about the position of women in the labour market and in society (Blanchflower and Freeman, 1997; Campa and Serafinelli, 2016; Lange, 2008; Fargher et al., 2008). The transition to market economy started in the 1990s entailed important changes in this regard too (Vecernìk, 2003), not only because the economic environment changed dramatically and forced many men and women into unemployment or out of the labour force. Central and Eastern European governments widely endorsed more conservative gender policies, emphasizing women's roles as mothers rather than workers and making labour market participation more difficult (Pascall and Manning, 2000); at the same time, the change in regime led many citizens of post-communist economies to support market justice norms and outcomes merely in contradistinction to socialist norms (Mason and Kluegel, 2000).

As a consequence, while in western European countries younger cohorts of women have been gradually exposed to more progressive and gender-neutral policies, attitudes and environments compared to their older counterparts, the opposite has happened in Central and Eastern Europe. This might explain why the estimation of model 3 does not provide evidence of an impact of gender inequality on the gender-job

[16] Education and health indexes have limited variability across the countries included in our analysis, ranging from 0.982 to 1 and from 0.964 to 0.980, respectively, and were not included in the analysis. Conversely, the participation index ranges from 0.565 to 0.836 and the political empowerment index from 0.057 to 0.754. Finally, the Gender Gap Index varies from 0.674 to 0.873.

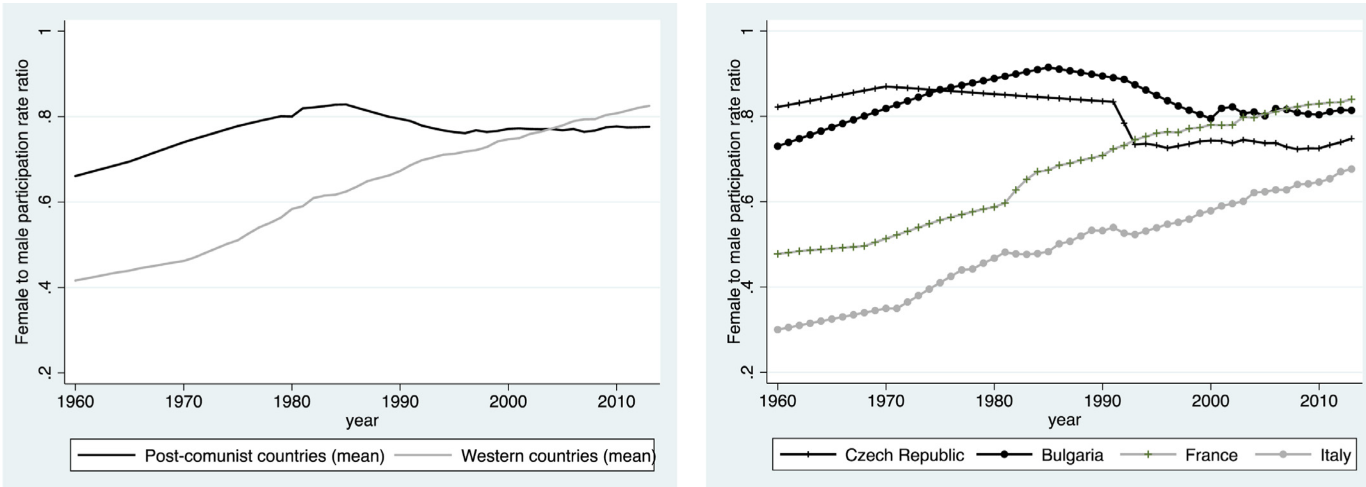


**Fig. 2.** Gender equality patterns across Europe (1960–2015).
*Source:* Our elaborations from WDI and various sources (see footnote 17).

satisfaction paradox; the effect of current equality simply does not apply to all women in the sample in the same way, and ignores the fact that gender equality has developed differently in different countries. To overcome this issue and investigate the direct link between historical changes in the gender equality and the gender-job satisfaction paradox, we need to introduce this historical dimension directly into the model.

To this aim we use the female/male participation (activity) rates ratio, one of the few available indicators of gender equality that can be constructed and traced back as close as possible to the age of birth of the oldest respondents in our sample.[17] Fig. 2 provides a snapshot for the (un-weighted) average levels of the indicator for Central Eastern and Western Europe and for some selected countries since 1960 and clearly shows how gender equality evolved quite differently in the two groups of countries.

In order to test the idea that early exposure to gender equal settings, via the proposed mechanism of values and beliefs (in this case regarding gender roles and consequent expectations), affects the current gender-job satisfaction paradox (Clark, 1997; Loscocco and Spitze, 1991; Miller, 1980) we construct an indicator of early life exposure to gender equality (ELGE) as the average of the female/male participation ratio over the first 20 years of life of each respondent in her/his country.[18] The ELGE indicator shows considerable variation as it ranges from 0.209 to 0.915, with a mean at 0.622. Although this measure does not fully capture the cultural and social setting in which the individual was raised (in particular with reference to her/his family characteristics), it provides a broad measure of the socio-economic gender environment in which work relations between the genders and gender-job expectations were constructed. This variable (ELGE), normalized at the mean in order to render the interpretation of remaining coefficients more straightforward, is then merged by country and age with the EU-SILC data and interacted with the gender dummy so as to estimate the model:

$$JS_{ik} = \alpha + \beta_1 fem_{ik} + \beta_2 ELGE_{ik} + \beta_{12} fem_{ik} * ELGE_{ik} + X'_{ikn}\gamma_n + IMR'_{ikm}\lambda_m + u_k + \varepsilon_{ik} \tag{4}$$

where $\beta_{12}$ indicates the impact of exposure to gender equality in childhood and youth on the gender-job satisfaction gap, and all other coefficients and variables are the same as in model 2.[19]

In line with our expectations $\beta_{12}$ is negative (Column 2 in Table 3, see Table A5 for full results), indicating that increase in ELGE, which signals the advancement of gender equality, is indeed related to lower levels of the gender-job satisfaction gap. Women who spent their childhood and youth in contexts with higher female-to-male participation ratios have statistically lower job satisfaction than their counterparts who were exposed to more gender unequal environments, while we find no such effects for men (coefficient $\beta_2$ is not significant). This corroborates the idea that surrounding context in early stages of life is, via development and internalisation of beliefs about gender roles and expectations, among the drivers of the gender-job satisfaction paradox.[20]

As ELGE is normalized to zero, the coefficient next to *female* in column 2 ($\beta_1$ from Eq. (4)) has the interpretation of the adjusted gender-job satisfaction gap conditional on and at the average level of ELGE (0.622). Table 2 indicates that the average gap remains positive and significant. The significant interaction term of variables female and ELGE enables us to calculate the gender-job satisfaction gap (i.e., marginal effects of coefficient next to female) at different levels of ELGE. The calculation of the marginal effects indicates that gender-job satisfaction gap becomes insignificant at the ELGE level of about 0.7. In other words, according

[17] The main data source was the World Bank World Development Indicators (WDI), providing data from 1960 onwards for a large number of countries (variable SL.TLF.CACT.FM.NE.ZS - Ratio of female to male labour force participation rate (%), national estimate). Missing data for Central and Eastern EU countries prior to 1989 were integrated using a large number of national specific information, which include: Godfrey and Richards (1997); Kinsella and Taeuber (1993), Elias (1972); Fullerton (1999); Sorrentino (1983); Statistics of the USSR (various years, in Russia); Federal Statistical Office of Yugoslavia (various years, in Serbian). The remaining few missing data (for both Central Eastern and Western countries) were reconstructed by linear interpolation.

[18] Our sample includes individuals born from 1949 to 1994. For the18,102 individuals born before 1960 (about 21 percent of the sample) the average of the gender equality indicator is calculated on a smaller number of years, as the data on gender inequality are available only up to 1960.

[19] In order to make its effects more visible we multiply the ELGE indicator by 100.

[20] To further corroborate our results we assembled a new database in which the observations represent five different ages (20, 30, 40, 50 and 60) in 32 countries (i.e. 5*32 = 160 observations). The variables in this database are: (i) the adjusted gap in gender job satisfaction, calculated as marginal effects of gender at different age using the EU-SILC database; and (ii) female-to-male participation ratios from the database on historical development of gender equality, matched again by age and country. We then performed a regression analysis in which the dependent variable is the gender-job satisfaction gap and the main regressor is the female to male participation ratio. Results are displayed in Fig. A1, and the estimated coefficient from the regression analysis is approximately equal to the one presented in Table 3 ($b = -0.436$, $SE_b = 0.123$; $p < 0.01$). Results confirm the evidence of lower gender-job satisfaction gap being associated to higher ELGE index.

**Table 3**
Job satisfaction gender gap and gender equality in early stages of life (OLS, matched sample).

| | 1 | 2 | 3 | 4 | 5 |
|---|---|---|---|---|---|
| Female | 0.068*** | 0.069*** | 0.070*** | 0.052** | 0.110*** |
| | (0.025) | (0.020) | (0.020) | (0.026) | (0.021) |
| Early Life Gender Equality (ELGE) Index | | −0.029 | −0.041 | −0.036 | −0.081 |
| | | (0.155) | (0.156) | (0.155) | (0.155) |
| Female * ELGE Index | | −0.411*** | −0.385*** | −0.407*** | −0.342*** |
| | | (0.082) | (0.089) | (0.083) | (0.083) |
| Female * Age | | | 0.001 | | |
| | | | (0.001) | | |
| Female * Tertiary Education (ED) | | | | 0.031 | |
| | | | | (0.028) | |
| Female * Male Occupation (Mocc) | | | | | −0.248*** |
| | | | | | (0.035) |

*Notes:* ELGE index and age normalized at mean (i.e. variables have mean at zero, while preserving the original variation). Multinomial selection effects included. Full models results are presented in Table A5 in the appendix. Standard errors clustered at country/gender/age level (parametric correction for Moulton factor). Significance levels: $^{***}p < 0.01$; $^{**}p < 0.05$; $^{*}p < 0.1$.

to our model, women exposed in their early childhood to female participation rates that represent 70% or more of the male participation rates will, on average, have the same job satisfaction as men, conditional on other controlled factors.

In columns 3 to 5 we test the robustness of the observed effect of ELGE by including the interaction of the gender dummy with variables describing age (normalized at mean so to preserve the interpretation of the coefficient next to female), education and male occupations. Such indicators are those identified and used by Clark to corroborate his conjecture about the role of women's expectations in shaping the job satisfaction gap; younger age, higher levels of education and employment in male-dominated occupations should help women re-align their expectations to those of men. While the age variable is simply expressed in years, the attainment of a high education level (*ED*) is a dichotomic variable that equals one if the individual completed tertiary education and zero otherwise. Male occupations are those in which the majority of the workers are male and the variable for male occupations (*Mocc*) is coded as 1 for: managers, crafts and trade workers and plant machine operators (in which men account for about 61%, 86% and 75% of employment respectively) and zero otherwise.[21]

Results indicate that the effect of ELGE is robust to the inclusion of all additional cross-terms. The non-significance of the additional interaction in column 3 indicates that ELGE already captures the effect that Clark attributes, indirectly, to age (i.e., weaker or no paradox for younger cohorts due to exposure to more gender equal contexts). However, ELGE has the advantage of including more refined and accurate information (associating to each age a country-specific level of exposure to gender inequality in early stages of life) and therefore direct evidence on the link between gender equality and the job satisfaction paradox.[22] While our results do not supply evidence of higher levels of education being associated with a smaller extent of the paradox (column 4), the effect of the gender structure of occupations is significant.[23] Results in column 5 indicate that while women in female-dominated sectors on average have higher job satisfaction than men (coefficient *female*), the opposite is true for male-dominated sectors (the negative coefficient of the interaction term *female*Mocc* overweighs the positive one of female). In an interpretative framework centred on expectations, this means that being exposed to male-dominated working environment can enable women to revise their beliefs and align expectations to those of men. The fact that both interactions with ELGE and gender structure of occupations are significant indicates that this mechanism and the one channelled by ELGE have independent effects on lowering the job satisfaction gaps.

We further test robustness of the ELGE indicator effect by including current equality indicators from the WEF global gender index (presented in Table 2). Table A6 in the Appendix indicates that inclusion of these indicators has no effect on the impact of ELGE on the job satisfaction gap; as in Table 2, indicators of current gender equality have no effect on the job satisfaction gap. Further robustness checks performed on the full sample, with oprobit estimator and with multinominal selection effects, confirm the sign and significance of the $\beta_{12}$ coefficient and are available upon request.

### 5.2. *Heterogeneous effects of early life gender equality across education groups and occupations*

As a last step of our analysis, we investigate whether the effects of ELGE are heterogeneous across education groups or gender structure of occupations. To answer this question we augment model (4) by another, triple interaction term obtained by multiplying the gender dummy, the (normalized) ELGE index and the indicators for higher levels of education and male-dominated occupations (*ED* and *Mocc)* defined above. The two empirical models, that also include all relevant double interaction terms, read:

$$\begin{aligned} JS_{ik} = {} & \alpha + \beta_1 fem_{ik} + \beta_2 ELGE_{ik} + \beta_{12} fem_{ik} * ELGE_{ik} + \beta_{13} fem_{ik} \\ & * ED_{ik} + \beta_{23} ELGE_{ik} * ED_{ik} + \beta_{123} fem_{ik} * ELGE_{ik} * ED_{ik} \\ & + X'_{ikn}\gamma_n + u_k + IMR'_{ikm}\lambda_m + \varepsilon_{ik} \end{aligned} \tag{5}$$

[21] Level terms of age, education and occupations are already included as covariates in the baseline specification. Similar results, available upon request, are obtained when we replace the dummy variable (Mocc) with country specific share of men by occupation.

[22] When ELGE variable and Female*ELGE term are left out of the specification the interaction term Female*Age becomes statistically significant ($b = 0.004$; s.e. $= 0.001$; $p < 0.01$; results available upon request), and we therefore replicate the result from Clark (1997). The fact that after the inclusion of ELGE variable and Female*ELGE the effect of Female*Age disappears further strengthens our conclusion that ELGE captures the effects improvement of the gender equality much better than age. Investigating the notion of effects of different gender equality settings via interacting the gender variable with age assumes that improvement of gender equality was linear with the age of the respondents, and that gender equality developed in the same way in all the countries. Fig. 2 illustrates that this assumption does not stand, and our results confirm that the complex and diverse histories of gender inequalities have a higher explanatory power of the job satisfaction paradox than age.

[23] When ELGE variable and Female*ELGE term are left out of the specification, we obtain similar results as in the Table 3 (results available upon request). Interaction of female and education is not statistically significant, while the interaction with the male-dominated occupations variable is significant and the coefficient is very similar ($b = -0.266$; s.e. $= 0.035$; $p < 0.01$). Therefore, the introduction of ELGE does not change the interpretation of these results.

**Table 4**
Heterogeneous effects of ELGE across education groups and occupations (OLS, matched sample) (summary).

| | 1 | | 2 | | 3 | |
|---|---|---|---|---|---|---|
| Female | 0.069*** | (0.020) | 0.076*** | (0.026) | 0.110*** | (0.021) |
| Early Life Gender Equality (ELGE) Index | −0.029 | (0.155) | −0.285* | (0.166) | −0.078 | (0.160) |
| Female * ELGE Index | −0.411*** | (0.082) | −0.577*** | (0.109) | −0.340*** | (0.093) |
| Female * Tertiary Education | | | 0.009 | (0.028) | | |
| ELGE Index * Tertiary Education | | | 0.576*** | (0.123) | | |
| Female * ELGE Index * Tertiary Education | | | 0.343** | (0.159) | | |
| Female * Male-dominated occupations | | | | | −0.245*** | (0.036) |
| ELGE Index * Male-dominated occupations | | | | | −0.012 | (0.134) |
| Female * ELGE Index * Male-dominated occupations | | | | | −0.042 | (0.211) |

*Notes:* ELGE index normalized at mean (i.e. the variable has mean at zero, while preserving the original variation). Multinomial selection effects included. Full models results are presented in table A7 in the appendix. Standard errors clustered at country/gender/age level (parametric correction for Moulton factor). Significance levels: ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$.

**Table 5**
Marginal effects of ELGE indicator for men and women at different education levels and for female and male dominated occupations.

| | | |
|---|---|---|
| *Marginal effects of ELGE for different levels of education* | | |
| Men (primary or secondary education) | −0.285* | (0.166) |
| Men (tertiary education) | 0.291 | (0.173) |
| Women (primary or secondary education) | −0.863*** | (0.175) |
| Women (tertiary education) | 0.056 | (0.171) |
| *Marginal effects of ELGE for female and male-dominated occupations* | | |
| Men (female-dominated occupations) | −0.078 | (0.160) |
| Men (male-dominated occupations) | −0.091 | (0.178) |
| Women (female-dominated occupations) | −0.418*** | (0.144) |
| Women (male-dominated occupations) | −0.472** | (0.213) |

***$p < 0.01$; **$p < 0.05$; *$p < 0.1$. Marginal effects calculated after the estimation of the model presented in Table 4 (full model in Table A7 in the appendix).

$$JS_{ik} = \alpha + \beta_1 fem_{ik} + \beta_2 ELGE_{ik} + \beta_{12} fem_{ik} * ELGE_{ik} + \beta_{13} fem_{ik} * Mocc_{ik} + \beta_{23} ELGE_{ik} * Mocc_{ik} + \beta_{123} fem_{ik} * ELGE_{ik} * Mocc_{ik} + X'_{ikn}\gamma_n + u_k + IMR'_{ikm}\lambda_m + \varepsilon_{ik} \quad (6)$$

In model (5) $\beta_{123}$ indicates whether exposure to gender equality setting in early life produces different effects on job satisfaction for women with tertiary education compared to their female counterparts with identical characteristics but lower levels of education. Similarly, in model (6), the same coefficient describes the differential effect of ELGE produced by holding a job in a male dominated occupation. Due to normalization of the ELGE variable all the coefficients next to constant terms (Female, Female*Tertiary Education, and Female*Tertiary Male-dominated occupations) should be interpreted as marginal effects at mean levels of ELGE (0.622) in our sample.

With reference to the double interaction terms, results from the estimation of models (5) and (6) (Tables 4 and A7) confirm the result presented in Table 3: while there are no differences between education levels in the size of the gender-job satisfaction gap, women in female-dominated sectors on average have higher job satisfaction than men, at the average level of ELGE indicator.

However, the significance of the triple interaction term of model (5) indicates that the impact of ELGE on the job satisfaction gap is heterogeneous across education levels. As the nature of the triple interaction specification does not allow easy interpretation of the coefficients, we compute the marginal effects of ELGE indicator for women and men at different levels of education and present them in Table 5. Results in the top panel indicate that for both men and women with lower levels of education ELGE decreases job satisfaction. However, while the effect is limited and only marginally significant for men (at 0.1 level), it is much stronger and significant for women (as indicated by Female*ELGE Index term in column 2 of Table 4). In other words, higher ELGE lowers job satisfaction for both genders, but much more for women, leading to the conclusion that higher ELGE decreases the gender-job satisfaction gap among those with lower levels of education. On the other hand, ELGE has no effect on either men or women with tertiary education, indicating that, for the pool of the highly educated the reasons behind gender differences in job satisfaction do not lie in exposure to gender inequalities in the early stages of life.

On the other hand, the triple interaction between gender, gender structure of occupations and ELGE (Table 4, column 3) is not significant, suggesting that ELGE lowers the gender-job satisfaction gap in both female and male dominated sectors, as confirmed by marginal effects reported in Table 5.

## 6. Discussion and conclusions

In 1997, Andrew Clark proposed an explanation of the *gender-job satisfaction paradox* – the paradox of women's higher job satisfaction despite lower wages and poorer working conditions – based on women's lower expectations from work. In the conclusion of his article, Clark (1997) suggested that such lower expectations are at least partially formulated early in life, under the influence of the observed position of women in the labour market. He further argued that the paradox is transitory and that advances in gender equality will diminish such gender differences in expectations. Therefore, Clark adumbrated a somewhat counterintuitive, but intriguing idea that the higher the gender equality is in a country, the lower women's "advantage" in job satisfaction is compared to men. This hypothesis was not tested explicitly by Clark in his 1997 work. Research that followed Clark's line of argumentation by comparing gender-job satisfaction gaps between countries or across time provided only descriptive evidence, since the link between gender-job satisfaction gap and gender equality was never econometrically modelled and the conclusions reached, as the authors themselves admitted, offered room for different interpretations (Kaiser, 2007; Sousa-Poza and Sousa-Poza, 2000, 2003).

In this paper, we aimed at filling this gap by providing explicit econometric evidence on the link between gender equality and the gender-job satisfaction paradox. We analysed the EU-SILC data for 32 countries for 2013 and applied the nearest neighbour matching procedure to address potential misspecification and comparability issues. Our analysis indicates that women in Europe, once all other possible drivers are controlled for, have on average higher levels of job satisfaction than men. We also show that there is a considerable variation in the gender-job satisfaction gap across Europe.

We first attempted to explain this cross-country heterogeneity by merging the EU-SILC data with the data on current levels of gender inequality, in order to explicitly test the hypothesis that the variability in gender inequality is behind countries' differences in gender-job satisfaction gaps. Our results indicate that contemporaneously observed levels of gender inequality (measured via WEF's Gender equality indices) do not have an impact on the size of the satisfaction gap. However, as Clark argued, it is the gender equality *in early stages of life*, rather than current levels of gender equality, that might have a crucial role in shaping expectations and therefore the size of the paradox.

To test this possibility, we attempted to place differences in gender inequalities across countries in their historical context by taking advantage of the geographical coverage of our sample. In order to explicitly model the idea that higher job satisfaction of women depends on exposure, in early life, to the poor position of women in the labour market, we construct an indicator of early life exposure to gender equality (ELGE). This is defined as the average female-to-male labour market participation ratio in the first 20 years of life of each respondent in her/his country. Results clearly indicate that exposure to higher gender equality in early stages of life is strongly and robustly connected to the lower levels of gender-job satisfaction gap and, therefore, to a decline of the paradox. Results also show that, independently of ELGE, employment in typically male occupations also decreases female job satisfaction, even to a level lower than male. On the other hand, education, although it has no direct effect on gender-job satisfaction paradox, plays a moderating role: significant and sizable effects of ELGE on gender-job satisfaction gap materialize only for low and medium-educated workers. For workers with tertiary education some other factors (not investigated in our paper) can be the source of the gender differences in job satisfaction.

The lowering effect of ELGE on the gender-job satisfaction paradox is consistent with the idea that gender inequality experienced in early stages of life has a persistent effect on beliefs and expectations: women who lived in more gender-equal contexts during childhood and youth might have developed expectations, and reported job satisfaction, more aligned to men's. Similarly, women who lived in less equal societies, besides developing lower expectations drawn from observing lower participation of women in the labour market, might have been "socialized" to put higher value on aspects of work such as flexibility and social connections, which translates into higher job satisfaction then men once wages and other individual and job characteristics are controlled for.

This interpretation draws attention to one of the aspects of subjective measures of well-being, such as job satisfaction, which have been exposed to criticism in previous research: their ability to deal with mechanisms of psychological adaptation (Frey and Strtzer, 2002). Despite being in very different objective conditions, two individuals could report similar levels of satisfaction due to adaptation (to) and acceptance of them by the one being worse-off. At the same time, however, Sen (1999) recognises that people's judgments may be constrained by political and social conditions, rather than psychology. As a consequence, any action able to play a constructive role in reducing obstacles for people to determine their own values or priorities is important not only for intrinsic reasons, but also because it allows people to be free to come to their own decisions (an essential point in the capability approach). This is a crucial logical step in drawing policy implications from our results, which would otherwise sound paradoxical: measures aimed at reducing gender inequality would result, years later, in lower job satisfaction (and well-being) of women, due to alignment of their expectations to those of men. Lower gender inequalities, besides having an intrinsic value, would enable women to shape expectations and preferences not downward biased by the circumstances in which they grew up, but equal to men's. This would facilitate the formulation of self-reflective and deliberate judgments and relevant decision making in all fundamental, and intertwined, domains of life (work and family in particular).

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## Conflict of Interest

The authors declare no potential conflicts of interest with respect to the research, authorship, and/or publication of this article.

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## Appendix

**Table A1**
List of variables and abbreviations used in the text and in the tables.

| | |
|---|---|
| ***Main variables*** | |
| *Job satisfaction* | Likert type response on a scale from 0 - “not at all satisfied” to 10 “I’m completely satisfied” (PW010) |
| *Job satisfaction categories* | Variable PW010 recoded to three categories: low (0–5), median (6–8) and high (9 and 10) job satisfaction. Used in oprobit robustness checks. |
| *Female* | Gender dummy (female = 1) |
| ***Covariates*** | |
| *Ln wage* | Log of monthly wage |
| *Hours* | Number of hours usually worked per week in main job (PL060) |
| *Hours50* | Dummy variable if person is working 50 h or more |
| *Married* | Marital status dummy (married = 1) |
| *Age* | Number of years |
| *Age2* | Number of years (squared) |
| *Education* | Highest level of education attained (1 = primary education; 2 = secondary education; 3 = tertiary education) |
| *Primary Education* | Primary Education dummy (ISCED levels 0–2 = 1) |
| *Secondary Education* | Secondary Education dummy (ISCED levels 3–4 = 1) |
| *Tertiary Education* | Tertiary Education dummy (ISCED levels 5–6 = 1) |
| *Occupation* | Type of Occupation (ISCO 08 classification): 1. Managers, 2. Professionals (army personal included), 3. Technicians and Associate Professionals, 4. Clerical Support Workers, 5. Services and Sales Workers, 6. Craft and Related Trade Workers, 7. Plant and Machine Operators and Assemblers, 8. Elementary Occupations |
| *Sector* | *Note: Skilled Agricultural, Forestry and Fishery Workers excluded, see Section 3.1* |
| | Sector of employment (NACE sections): 1. Industry (Sectors B-E), 2. Construction (F), 3. Trade (G), 4. Transport, Hotels and Restaurants (H-I), 5. Information and Communications; Financial and Insurance Activities (J-K), 6. Real Estate, Professional and Administrative Activities (L-N), 7. Public Administration (O), 8. Education (P), 9. Health and Social Work Activities (Q), 10. Other services (R-U) *Note: Agriculture (section A) excluded, see Section 3.1* |
| *Additional Job* | Additional job dummy (second job = 1) |
| *Firm size* | Size of the employer (coded as 1 if between 0 and 10 employees; 2 if between 11 and 49; 3 if over 50) |
| *Firm size (0–10)* | Small firm size dummy (between 0 and 10 employees = 1) |
| *Firm size (11–19)* | Medium firm size dummy (between 11 and 19 employees = 1) |
| *Firm size (20–49)* | Medium-large firm size dummy (between 20 and 49 employees = 1) |
| *Firm size (over 50)* | Large firm size dummy (over 50 employees = 1) |
| *Temporary contracts* | Employment status dummy (temporary = 1) |
| ***Gender equality indicators*** | |
| Overall GEI | *Overall Gender Equality Index (Source: World Economic Forum)* |
| EP Score | *Economic Participation and Opportunity Score (Source: World Economic Forum)* |
| PE Score | *Political Empowerment Score (Source: World Economic Forum)* |
| ELGE Index | *Early life Exposure to Gender Equality Index (WDI and various sources, see footnote 12)* |
| ***Selection variables*** | |
| IMR1 | *Inverse Mills Ratio based on the probability to be in dependent employment* |
| IMR2 | *Inverse Mills Ratio based on the probability to be in other employment* |
| IMR3 | *Inverse Mills Ratio based on the probability to be unemployed* |
| IMR4 | *Inverse Mills Ratio based on the probability to be inactive* |

**Table A2**
Descriptive statistics of the variables included in the job satisfaction model (matched sample).

| | Observations | Mean | Std. Deviation | Minimum | Maximum |
|---|---|---|---|---|---|
| Job satisfaction | 83,555 | 7.3 | 1.9 | 0 | 10 |
| Female | 83,555 | 0.558 | 0.497 | 0 | 1 |
| Ln wage | 83,555 | 7.2 | 1.1 | 0.0 | 11.6 |
| Hours | 83,555 | 38.6 | 7.2 | 1 | 99 |
| Hours50 | 83,555 | 0.042 | 0.202 | 0 | 1 |
| Age | 83,555 | 44.4 | 10.2 | 19 | 64 |
| Age2 / 100 | 83,555 | 20.7 | 8.9 | 3.6 | 41.0 |
| Married | 83,555 | 0.615 | 0.487 | 0 | 1 |
| Secondary education | 83,555 | 0.445 | 0.497 | 0 | 1 |
| Tertiary education | 83,555 | 0.460 | 0.498 | 0 | 1 |
| Senior officials and managers | 83,555 | 0.053 | 0.225 | 0 | 1 |
| Professionals | 83,555 | 0.275 | 0.447 | 0 | 1 |
| Technicians and ass. professionals | 83,555 | 0.185 | 0.388 | 0 | 1 |
| Clerks | 83,555 | 0.108 | 0.311 | 0 | 1 |
| Service and sales workers | 83,555 | 0.172 | 0.377 | 0 | 1 |
| Craft and trades workers | 83,555 | 0.082 | 0.274 | 0 | 1 |
| Plant and machine operators | 83,555 | 0.056 | 0.229 | 0 | 1 |
| Sector F | 83,555 | 0.038 | 0.192 | 0 | 1 |
| Sector G | 83,555 | 0.130 | 0.336 | 0 | 1 |
| Sectors H - I | 83,555 | 0.078 | 0.268 | 0 | 1 |
| Sectors J - K | 83,555 | 0.079 | 0.270 | 0 | 1 |
| Sectors L - N | 83,555 | 0.082 | 0.274 | 0 | 1 |
| Sector O | 83,555 | 0.119 | 0.324 | 0 | 1 |
| Sector P | 83,555 | 0.131 | 0.337 | 0 | 1 |
| Sector Q | 83,555 | 0.130 | 0.336 | 0 | 1 |
| Sectors R - U | 83,555 | 0.038 | 0.191 | 0 | 1 |
| Additional job | 83,555 | 0.051 | 0.221 | 0 | 1 |
| Firm size 11/19 | 83,555 | 0.156 | 0.363 | 0 | 1 |
| Firm size 20/49 | 83,555 | 0.171 | 0.376 | 0 | 1 |
| Firm size 50+ | 83,555 | 0.455 | 0.498 | 0 | 1 |
| Temporary contract | 83,555 | 0.058 | 0.234 | 0 | 1 |
| Overall GEI | 83,555 | 0.731 | 0.047 | 0.674 | 0.873 |
| EP Score | 83,555 | 0.699 | 0.062 | 0.566 | 0.836 |
| PE Score | 83,555 | 0.253 | 0.142 | 0.057 | 0.754 |
| ELGE Index | 83,555 | 0.622 | 0.169 | 0.209 | 0.915 |

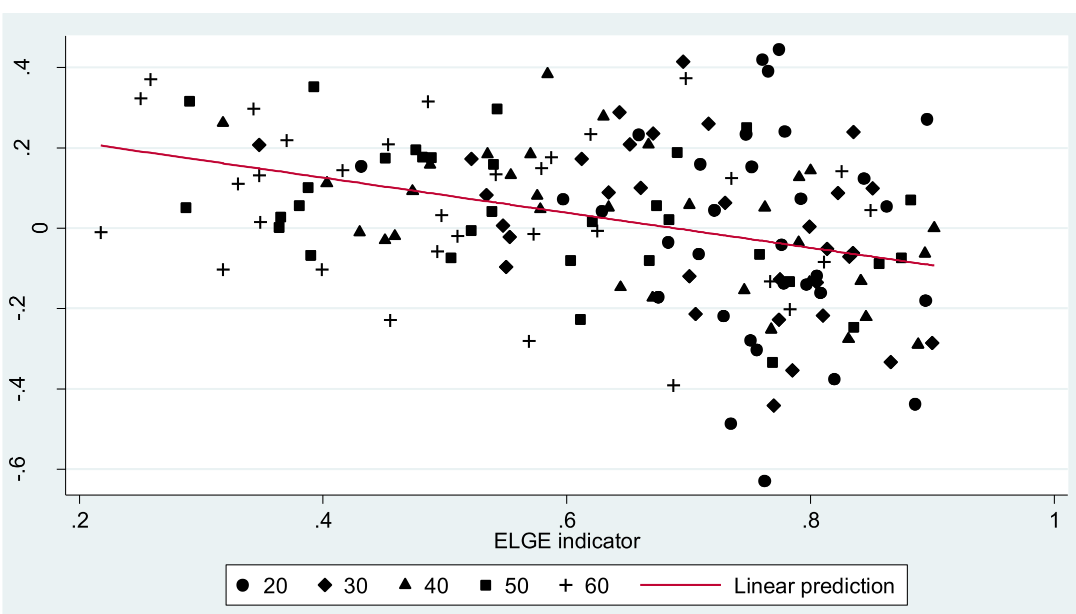


**Fig. A1.** Job satisfaction gender gap and gender equality in early stages of life, by age group.
*Notes:* Estimated based on the regression analysis in which the dependent variable is the country-level gender-job satisfaction gap at different ages (20, 30, 40, 50 and 60) and the only regressor is the female to male participation ratio. Estimated coefficient from the regression analysis is approximately equal to the one presented in Table 3 ($b = -0.436$, $SE_b = 0.123$; $p < 0.01$, full results and data for the estimation available upon request). Results confirm the evidence of lower gender-job satisfaction gap being associated to higher ELGE index.

**Table A3**
Unadjusted and adjusted gender gaps in job satisfaction by country (OLS, matched sample).

| | Unadjusted gap | | Adjusted gap | |
|---|---|---|---|---|
| all | 0.080*** | (0.013) | 0.068*** | (0.014) |
| AT | 0.178** | (0.073) | 0.138 | (0.084) |
| BE | 0.115** | (0.057) | 0.101 | (0.066) |
| BG | −0.019 | (0.090) | −0.130 | (0.095) |
| CH | 0.047 | (0.055) | 0.015 | (0.069) |
| CY | 0.240*** | (0.080) | 0.230*** | (0.082) |
| CZ | −0.122* | (0.070) | −0.046 | (0.077) |
| DE | 0.028 | (0.056) | 0.060 | (0.065) |
| DK | 0.050 | (0.093) | 0.092 | (0.106) |
| EE | 0.059 | (0.078) | 0.215** | (0.091) |
| EL | −0.129 | (0.096) | 0.067 | (0.094) |
| ES | 0.142*** | (0.051) | 0.105* | (0.055) |
| FI | 0.159*** | (0.054) | 0.100 | (0.061) |
| FR | 0.124** | (0.058) | 0.108* | (0.062) |
| HR | 0.070 | (0.141) | −0.103 | (0.150) |
| HU | 0.192*** | (0.055) | 0.130** | (0.059) |
| IE | 0.063 | (0.135) | 0.205 | (0.150) |
| IS | 0.185 | (0.146) | 0.336** | (0.162) |
| IT | 0.078 | (0.049) | 0.077 | (0.054) |
| LT | −0.148** | (0.073) | −0.228*** | (0.078) |
| LU | 0.112 | (0.095) | 0.061 | (0.103) |
| LV | 0.128* | (0.074) | 0.037 | (0.084) |
| MT | 0.230* | (0.121) | 0.257* | (0.135) |
| NL | 0.042 | (0.049) | 0.115** | (0.058) |
| NO | 0.083 | (0.068) | 0.065 | (0.079) |
| PL | 0.091* | (0.055) | −0.031 | (0.063) |
| PT | 0.202** | (0.090) | 0.225** | (0.100) |
| RO | −0.033 | (0.049) | −0.026 | (0.055) |
| RS | 0.200* | (0.118) | 0.155 | (0.110) |
| SE | −0.067 | (0.087) | −0.095 | (0.096) |
| SI | 0.213** | (0.099) | 0.146 | (0.108) |
| SK | −0.100 | (0.061) | −0.181** | (0.073) |
| UK | 0.287*** | (0.071) | 0.353*** | (0.079) |

*Notes:* Robust standard errors in parentheses. ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$. The unadjusted gap is estimated by running model (2) country by country with only the gender dummy on the right hand side.

**Table A4**
Job satisfaction gender gap and current gender equality (OLS, matched sample).

| | 1 | | 2 | | 3 | | 4 | |
|---|---|---|---|---|---|---|---|---|
| Female | 0.068*** | (0.025) | 0.068*** | (0.025) | 0.073*** | (0.025) | 0.066*** | (0.025) |
| Female*Overall GEI | | | −0.053 | (0.373) | | | | |
| Female*EP Score | | | | | −0.387 | (0.281) | | |
| Female*PE Score | | | | | | | 0.059 | (0.122) |
| Ln wage | 0.353*** | (0.017) | 0.352*** | (0.017) | 0.352*** | (0.016) | 0.353*** | (0.017) |
| Hours | −0.010*** | (0.001) | −0.010*** | (0.001) | −0.010*** | (0.001) | −0.010*** | (0.001) |
| Hours50 | −0.056 | (0.038) | −0.056 | (0.038) | −0.055 | (0.038) | −0.057 | (0.038) |
| Age | −0.053*** | (0.008) | −0.007*** | (0.001) | −0.007*** | (0.001) | −0.007*** | (0.001) |
| $Age^2$ / 100 | 0.051*** | (0.010) | 0.001*** | (0.000) | 0.001*** | (0.000) | 0.001*** | (0.000) |
| Married | 0.083*** | (0.017) | 0.083*** | (0.017) | 0.081*** | (0.017) | 0.084*** | (0.017) |
| *Primary education (omitted)* | | | | | | | | |
| Secondary education | −0.289*** | (0.030) | −0.289*** | (0.030) | −0.293*** | (0.030) | −0.289*** | (0.030) |
| Tertiary education | −0.416*** | (0.039) | −0.416*** | (0.039) | −0.422*** | (0.039) | −0.415*** | (0.039) |
| Managers | 0.822*** | (0.044) | 0.822*** | (0.044) | 0.822*** | (0.044) | 0.821*** | (0.044) |
| Professionals | 0.708*** | (0.035) | 0.708*** | (0.035) | 0.709*** | (0.035) | 0.708*** | (0.035) |
| Technicians | 0.603*** | (0.033) | 0.603*** | (0.033) | 0.604*** | (0.033) | 0.603*** | (0.033) |
| Clerks | 0.492*** | (0.035) | 0.492*** | (0.035) | 0.492*** | (0.035) | 0.491*** | (0.035) |
| Service / sales workers | 0.390*** | (0.031) | 0.390*** | (0.031) | 0.391*** | (0.031) | 0.389*** | (0.031) |
| Craft / trades workers | 0.082** | (0.038) | 0.082** | (0.038) | 0.083** | (0.038) | 0.081** | (0.038) |
| Plant / mach. operators | 0.172*** | (0.040) | 0.172*** | (0.040) | 0.173*** | (0.040) | 0.172*** | (0.040) |
| *Elementary occupations (omitted)* | | | | | | | | |
| *Sectors B-E (omitted)* | | | | | | | | |
| Sector F | 0.107*** | (0.038) | 0.107*** | (0.038) | 0.107*** | (0.037) | 0.106*** | (0.038) |
| Sector G | −0.060** | (0.028) | −0.060** | (0.028) | −0.060** | (0.028) | −0.060** | (0.028) |
| Sectors H - I | 0.048 | (0.031) | 0.048 | (0.031) | 0.048 | (0.031) | 0.048 | (0.031) |
| Sectors J - K | −0.059* | (0.031) | −0.059* | (0.031) | −0.060* | (0.031) | −0.059* | (0.031) |
| Sectors L - N | −0.025 | (0.031) | −0.025 | (0.031) | −0.026 | (0.031) | −0.025 | (0.031) |
| Sector O | 0.210*** | (0.028) | 0.210*** | (0.028) | 0.210*** | (0.028) | 0.209*** | (0.028) |
| Sector P | 0.398*** | (0.029) | 0.398*** | (0.029) | 0.398*** | (0.029) | 0.398*** | (0.029) |
| Sector Q | 0.282*** | (0.029) | 0.282*** | (0.029) | 0.283*** | (0.029) | 0.281*** | (0.029) |
| Sectors R - U | 0.374*** | (0.040) | 0.374*** | (0.040) | 0.373*** | (0.040) | 0.374*** | (0.040) |
| Additional job | 0.040 | (0.030) | 0.040 | (0.030) | 0.040 | (0.030) | 0.040 | (0.030) |
| *Firm size 1–10 (omitted)* | | | | | | | | |
| Firm size 11/19 | −0.103*** | (0.022) | −0.103*** | (0.022) | −0.103*** | (0.022) | −0.103*** | (0.022) |
| Firm size 20/49 | −0.102*** | (0.022) | −0.102*** | (0.022) | −0.102*** | (0.022) | −0.102*** | (0.022) |
| Firm size 50+ | −0.157*** | (0.019) | −0.157*** | (0.019) | −0.157*** | (0.019) | −0.157*** | (0.019) |
| Temporary contract | −0.318*** | (0.030) | −0.318*** | (0.030) | −0.318*** | (0.030) | −0.318*** | (0.030) |
| IMR1 | −0.114* | (0.059) | −0.113* | (0.059) | −0.109* | (0.059) | −0.114* | (0.059) |
| IMR2 | −0.438*** | (0.155) | −0.441*** | (0.156) | −0.457*** | (0.154) | −0.427*** | (0.157) |
| IMR3 | 0.768*** | (0.176) | 0.769*** | (0.176) | 0.801*** | (0.176) | 0.765*** | (0.176) |
| IMR4 | −0.270** | (0.126) | −0.268** | (0.127) | −0.241* | (0.127) | −0.273** | (0.126) |
| Constant | 6.491*** | (0.301) | 5.165*** | (0.165) | 5.166*** | (0.163) | 5.164*** | (0.165) |
| Adj. r square | 0.0892 | | 0.0892 | | 0.0892 | | 0.0892 | |
| Observations | 83,555 | | 83,555 | | 83,555 | | 83,555 | |

*Notes:* Country-fixed effects omitted from the table and available upon request. Gender equality indicators normalized at mean (i.e., variables have the mean at zero, while preserving the original variation). Standard errors clustered at country/gender level (parametric correction for Moulton factor). Significance levels: ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$.

**Table A5**
Job satisfaction gender gap and gender equality in early stages of life (OLS, matched sample).

| | 2 | | 3 | | 4 | | 5 | |
|---|---|---|---|---|---|---|---|---|
| Female | 0.069*** | (0.020) | 0.070*** | (0.020) | 0.052** | (0.026) | 0.110*** | (0.021) |
| ELGE index | −0.029 | (0.155) | −0.041 | (0.156) | −0.036 | (0.155) | −0.081 | (0.155) |
| Female*ELGE index | −0.411*** | (0.082) | −0.385*** | (0.089) | −0.407*** | (0.083) | −0.342*** | (0.083) |
| Female*Age | | | 0.001 | (0.001) | | | | |
| Female*Ed | | | | | 0.031 | (0.028) | | |
| Female*Mocc | | | | | | | −0.248*** | (0.035) |
| Ln wage | 0.350*** | (0.014) | 0.350*** | (0.014) | 0.350*** | (0.014) | 0.347*** | (0.014) |
| Hours | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.009*** | (0.001) |
| Hours50 | −0.064* | (0.037) | −0.064* | (0.037) | −0.063* | (0.037) | −0.067* | (0.037) |
| Age | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.009*** | (0.001) |
| $Age^2$ / 100 | 0.001*** | (0.000) | 0.001*** | (0.000) | 0.001*** | (0.000) | 0.001*** | (0.000) |
| Married | 0.093*** | (0.017) | 0.094*** | (0.017) | 0.093*** | (0.017) | 0.092*** | (0.017) |
| *Primary education (omitted)* | | | | | | | | |
| Secondary education | −0.288*** | (0.030) | −0.286*** | (0.030) | −0.287*** | (0.030) | −0.291*** | (0.030) |
| Tertiary education | −0.414*** | (0.038) | −0.411*** | (0.038) | −0.430*** | (0.041) | −0.416*** | (0.038) |
| Managers | 0.824*** | (0.043) | 0.825*** | (0.043) | 0.824*** | (0.043) | 0.940*** | (0.046) |
| Professionals | 0.712*** | (0.035) | 0.712*** | (0.035) | 0.710*** | (0.035) | 0.720*** | (0.035) |
| Technicians | 0.605*** | (0.033) | 0.606*** | (0.033) | 0.604*** | (0.033) | 0.614*** | (0.033) |
| Clerks | 0.491*** | (0.034) | 0.492*** | (0.034) | 0.490*** | (0.034) | 0.493*** | (0.034) |
| Service / sales workers | 0.392*** | (0.031) | 0.393*** | (0.031) | 0.392*** | (0.031) | 0.396*** | (0.031) |
| Craft / trades workers | 0.078** | (0.037) | 0.079** | (0.037) | 0.072* | (0.037) | 0.165*** | (0.039) |
| Plant / mach. operators | 0.169*** | (0.039) | 0.170*** | (0.039) | 0.165*** | (0.039) | 0.276*** | (0.042) |
| *Elementary occupations (omitted)* | | | | | | | | |
| *Sectors B-E (omitted)* | | | | | | | | |
| Sector F | 0.097*** | (0.037) | 0.097*** | (0.037) | 0.094** | (0.037) | 0.061 | (0.037) |
| Sector G | −0.059** | (0.028) | −0.059** | (0.028) | −0.059** | (0.028) | −0.072** | (0.028) |
| Sectors H - I | 0.048 | (0.030) | 0.048 | (0.030) | 0.047 | (0.031) | 0.037 | (0.031) |
| Sectors J - K | −0.060* | (0.031) | −0.060* | (0.031) | −0.059* | (0.031) | −0.067** | (0.031) |
| Sectors L - N | −0.028 | (0.031) | −0.028 | (0.031) | −0.029 | (0.031) | −0.036 | (0.031) |
| Sector O | 0.209*** | (0.028) | 0.209*** | (0.028) | 0.207*** | (0.028) | 0.201*** | (0.028) |
| Sector P | 0.397*** | (0.029) | 0.397*** | (0.029) | 0.395*** | (0.029) | 0.380*** | (0.029) |
| Sector Q | 0.279*** | (0.028) | 0.278*** | (0.028) | 0.277*** | (0.028) | 0.259*** | (0.028) |
| Sectors R - U | 0.371*** | (0.039) | 0.372*** | (0.039) | 0.371*** | (0.039) | 0.361*** | (0.040) |
| Additional job | 0.043 | (0.030) | 0.043 | (0.030) | 0.043 | (0.030) | 0.045 | (0.030) |
| *Firm size 1–10 (omitted)* | | | | | | | | |
| Firm size 11/19 | −0.104*** | (0.022) | −0.104*** | (0.022) | −0.105*** | (0.022) | −0.103*** | (0.022) |
| Firm size 20/49 | −0.102*** | (0.022) | −0.102*** | (0.022) | −0.102*** | (0.022) | −0.099*** | (0.022) |
| Firm size 50+ | −0.157*** | (0.019) | −0.158*** | (0.019) | −0.157*** | (0.019) | −0.152*** | (0.019) |
| Temporary contract | −0.316*** | (0.030) | −0.316*** | (0.030) | −0.316*** | (0.030) | −0.312*** | (0.030) |
| IMR1 | −0.104* | (0.055) | −0.105* | (0.055) | −0.105* | (0.055) | −0.106* | (0.055) |
| IMR2 | −0.367*** | (0.133) | −0.378*** | (0.134) | −0.356*** | (0.133) | −0.342*** | (0.133) |
| IMR3 | 0.737*** | (0.156) | 0.717*** | (0.158) | 0.738*** | (0.156) | 0.740*** | (0.156) |
| IMR4 | −0.192 | (0.122) | −0.203* | (0.123) | −0.208* | (0.123) | −0.223* | (0.122) |
| Constant | 5.170*** | (0.131) | 5.158*** | (0.132) | 5.183*** | (0.132) | 5.167*** | (0.131) |
| r square | 0.0895 | | 0.0895 | | 0.0895 | | 0.0901 | |
| Observations | 83,555 | | 83,555 | | 83,555 | | 83,555 | |

*Notes:* Country-fixed effects omitted from the table and available upon request. ELGE index and age normalized at mean (i.e., variables have the mean at zero, while preserving the original variation). Standard errors clustered at country/gender/age level (parametric correction for Moulton factor). Significance levels: ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$. Full estimates for column 1 from Table 3 already presented in table A4.

**Table A6**
Job satisfaction gender gap and gender equality in early stages of life (OLS, matched sample).

| | 1 | | 2 | | 3 | | 4 | |
|---|---|---|---|---|---|---|---|---|
| Female | 0.069*** | (0.020) | 0.070*** | (0.022) | 0.071*** | (0.022) | 0.070*** | (0.022) |
| ELGE index | −0.029 | (0.155) | −0.027 | (0.167) | −0.031 | (0.167) | −0.028 | (0.168) |
| Female*ELGE index | −0.411*** | (0.082) | −0.412*** | (0.090) | −0.400*** | (0.092) | −0.413*** | (0.091) |
| Female*Overall GEI | | | −0.096 | (0.331) | | | | |
| Female*EP Score | | | | | −0.163 | (0.257) | | |
| Female*PE Score | | | | | | | −0.012 | (0.110) |
| Ln wage | 0.350*** | (0.014) | 0.350*** | (0.015) | 0.350*** | (0.015) | 0.350*** | (0.015) |
| Hours | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.009*** | (0.001) |
| Hours50 | −0.064* | (0.037) | −0.064* | (0.037) | −0.063* | (0.037) | −0.064* | (0.037) |
| Age | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.008*** | (0.001) | −0.009*** | (0.001) |
| $Age^2$ / 100 | 0.001*** | (0.000) | 0.001*** | (0.000) | 0.001*** | (0.000) | 0.001*** | (0.000) |
| Married | 0.093*** | (0.017) | 0.093*** | (0.017) | 0.092*** | (0.017) | 0.093*** | (0.017) |
| *Primary education (omitted)* | | | | | | | | |
| Secondary education | −0.288*** | (0.030) | −0.288*** | (0.030) | −0.289*** | (0.030) | −0.288*** | (0.030) |
| Tertiary education | −0.414*** | (0.038) | −0.415*** | (0.038) | −0.417*** | (0.039) | −0.414*** | (0.038) |
| Managers | 0.824*** | (0.043) | 0.824*** | (0.044) | 0.824*** | (0.044) | 0.824*** | (0.044) |
| Professionals | 0.712*** | (0.035) | 0.712*** | (0.035) | 0.712*** | (0.035) | 0.712*** | (0.035) |
| Technicians | 0.605*** | (0.033) | 0.606*** | (0.033) | 0.606*** | (0.033) | 0.606*** | (0.033) |
| Clerks | 0.491*** | (0.034) | 0.492*** | (0.034) | 0.492*** | (0.034) | 0.491*** | (0.034) |
| Service / sales workers | 0.392*** | (0.031) | 0.393*** | (0.031) | 0.393*** | (0.031) | 0.392*** | (0.031) |
| Craft / trades workers | 0.078** | (0.037) | 0.078** | (0.038) | 0.078** | (0.038) | 0.078** | (0.038) |
| Plant / mach. operators | 0.169*** | (0.039) | 0.169*** | (0.039) | 0.169*** | (0.039) | 0.169*** | (0.039) |
| *Elementary occupations (omitted)* | | | | | | | | |
| *Sectors B-E (omitted)* | | | | | | | | |
| Sector F | 0.097*** | (0.037) | 0.097*** | (0.037) | 0.097*** | (0.037) | 0.097*** | (0.037) |
| Sector G | −0.059** | (0.028) | −0.059** | (0.028) | −0.059** | (0.028) | −0.059** | (0.028) |
| Sectors H - I | 0.048 | (0.030) | 0.049 | (0.030) | 0.048 | (0.030) | 0.049 | (0.030) |
| Sectors J - K | −0.060* | (0.031) | −0.060* | (0.031) | −0.060* | (0.031) | −0.060* | (0.031) |
| Sectors L - N | −0.028 | (0.031) | −0.028 | (0.031) | −0.029 | (0.031) | −0.028 | (0.031) |
| Sector O | 0.209*** | (0.028) | 0.209*** | (0.028) | 0.209*** | (0.028) | 0.209*** | (0.028) |
| Sector P | 0.397*** | (0.029) | 0.397*** | (0.029) | 0.397*** | (0.029) | 0.397*** | (0.029) |
| Sector Q | 0.279*** | (0.028) | 0.279*** | (0.028) | 0.279*** | (0.028) | 0.279*** | (0.028) |
| Sectors R - U | 0.371*** | (0.039) | 0.371*** | (0.040) | 0.371*** | (0.040) | 0.371*** | (0.040) |
| Additional job | 0.043 | (0.030) | 0.043 | (0.030) | 0.043 | (0.030) | 0.043 | (0.030) |
| *Firm size 1–10 (omitted)* | | | | | | | | |
| Firm size 11/19 | −0.104*** | (0.022) | −0.104*** | (0.022) | −0.104*** | (0.022) | −0.104*** | (0.022) |
| Firm size 20/49 | −0.102*** | (0.022) | −0.102*** | (0.022) | −0.102*** | (0.022) | −0.102*** | (0.022) |
| Firm size 50+ | −0.157*** | (0.019) | −0.157*** | (0.019) | −0.157*** | (0.019) | −0.157*** | (0.019) |
| Temporary contract | −0.316*** | (0.030) | −0.316*** | (0.030) | −0.317*** | (0.030) | −0.316*** | (0.030) |
| IMR1 | −0.104* | (0.055) | −0.103* | (0.057) | −0.102* | (0.057) | −0.103* | (0.057) |
| IMR2 | −0.367*** | (0.133) | −0.372*** | (0.143) | −0.376*** | (0.143) | −0.369** | (0.144) |
| IMR3 | 0.737*** | (0.156) | 0.740*** | (0.164) | 0.752*** | (0.166) | 0.738*** | (0.164) |
| IMR4 | −0.192 | (0.122) | −0.189 | (0.124) | −0.182 | (0.124) | −0.191 | (0.124) |
| Constant | 5.170*** | (0.131) | 5.171*** | (0.146) | 5.170*** | (0.146) | 5.170*** | (0.146) |
| r square | 0.0895 | | 0.0895 | | 0.0895 | | 0.0895 | |
| Observations | 83,555 | | 83,555 | | 83,555 | | 83,555 | |

*Notes:* Country-fixed effects omitted from the table and available upon request. Gender equality indicators, ELGE index and age normalized at mean (i.e., variables have the mean at zero, while preserving the original variation). Standard errors clustered at country/gender level (parametric correction for Moulton factor), accept for the column one where standard errors are clustered on the country/gender/age level. Significance levels: ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$.

**Table A7**
Heterogeneous effects of ELGE across education groups and occupations (OLS, matched sample).

| | 1 | | 2 | | 3 | |
|---|---|---|---|---|---|---|
| Female | 0.069*** | (0.020) | 0.076*** | (0.026) | 0.110*** | (0.021) |
| ELGE index | −0.029 | (0.155) | −0.285* | (0.166) | −0.078 | (0.160) |
| Female*ELGE index | −0.411*** | (0.082) | −0.577*** | (0.109) | −0.340*** | (0.093) |
| Female * Tertiary Education | | | 0.009 | (0.028) | | |
| ELGE * Tertiary Education | | | 0.576*** | (0.123) | | |
| Female * ELGE * Tertiary Education | | | 0.343** | (0.159) | | |
| Female * Male-dominated occ. | | | | | −0.245*** | (0.036) |
| ELGE * Male-dominated occ. | | | | | −0.012 | (0.134) |
| Female * ELGE * Male-dominated occ. | | | | | −0.042 | (0.211) |
| Ln wage | 0.350*** | (0.014) | 0.352*** | (0.014) | 0.348*** | (0.014) |
| Hours | −0.009*** | (0.001) | −0.009*** | (0.001) | −0.009*** | (0.001) |
| Hours50 | −0.064* | (0.037) | −0.056 | (0.037) | −0.067* | (0.037) |
| Age | −0.009*** | (0.001) | −0.008*** | (0.001) | −0.009*** | (0.001) |
| $Age^2$ / 100 | 0.001*** | (0.000) | 0.001*** | (0.000) | 0.001*** | (0.000) |
| Married | 0.093*** | (0.017) | 0.101*** | (0.017) | 0.092*** | (0.017) |
| *Primary education (omitted)* | | | | | | |
| Secondary education | −0.288*** | (0.030) | −0.206*** | (0.031) | −0.290*** | (0.030) |
| Tertiary education | −0.414*** | (0.038) | −0.324*** | (0.042) | −0.415*** | (0.038) |
| Managers | 0.824*** | (0.043) | 0.797*** | (0.044) | 0.939*** | (0.047) |
| Professionals | 0.712*** | (0.035) | 0.685*** | (0.035) | 0.720*** | (0.035) |
| Technicians | 0.605*** | (0.033) | 0.580*** | (0.033) | 0.613*** | (0.033) |
| Clerks | 0.491*** | (0.034) | 0.467*** | (0.034) | 0.493*** | (0.034) |
| Service / sales workers | 0.392*** | (0.031) | 0.383*** | (0.031) | 0.396*** | (0.031) |
| Craft / trades workers | 0.078** | (0.037) | 0.079** | (0.037) | 0.165*** | (0.039) |
| Plant / mach. operators | 0.169*** | (0.039) | 0.172*** | (0.039) | 0.277*** | (0.042) |
| *Elementary occupations (omitted)* | | | | | | |
| *Sectors B-E (omitted)* | | | | | | |
| Sector F | 0.097*** | (0.037) | 0.094** | (0.037) | 0.061 | (0.037) |
| Sector G | −0.059** | (0.028) | −0.060** | (0.028) | −0.072** | (0.028) |
| Sectors H - I | 0.048 | (0.030) | 0.044 | (0.031) | 0.036 | (0.031) |
| Sectors J - K | −0.060* | (0.031) | −0.063** | (0.031) | −0.068** | (0.031) |
| Sectors L - N | −0.028 | (0.031) | −0.035 | (0.031) | −0.036 | (0.031) |
| Sector O | 0.209*** | (0.028) | 0.202*** | (0.028) | 0.201*** | (0.028) |
| Sector P | 0.397*** | (0.029) | 0.401*** | (0.029) | 0.380*** | (0.029) |
| Sector Q | 0.279*** | (0.028) | 0.278*** | (0.028) | 0.259*** | (0.028) |
| Sectors R - U | 0.371*** | (0.039) | 0.364*** | (0.039) | 0.361*** | (0.040) |
| Additional job | 0.043 | (0.030) | 0.042 | (0.029) | 0.045 | (0.030) |
| *Firm size 1–10 (omitted)* | | | | | | |
| Firm size 11/19 | −0.104*** | (0.022) | −0.102*** | (0.022) | −0.103*** | (0.022) |
| Firm size 20/49 | −0.102*** | (0.022) | −0.102*** | (0.022) | −0.099*** | (0.022) |
| Firm size 50+ | −0.157*** | (0.019) | −0.156*** | (0.019) | −0.152*** | (0.019) |
| Temporary contract | −0.316*** | (0.030) | −0.315*** | (0.030) | −0.312*** | (0.030) |
| IMR1 | −0.104* | (0.055) | −0.082 | (0.055) | −0.106* | (0.055) |
| IMR2 | −0.367*** | (0.133) | −0.434*** | (0.134) | −0.343*** | (0.133) |
| IMR3 | 0.737*** | (0.156) | 0.525*** | (0.158) | 0.739*** | (0.156) |
| IMR4 | −0.192 | (0.122) | −0.170 | (0.123) | −0.223* | (0.122) |
| Constant (cut 1) | 5.170*** | (0.131) | 5.024*** | (0.134) | 5.166*** | (0.132) |
| (pseudo) r square | 0.0895 | | 0.0905 | | 0.0900 | |
| Observations | 83,555 | | 83,555 | | 83,555 | |

*Notes:* Country-fixed effects omitted from the table and available upon request. ELGE index and age normalized at mean (i.e., variables have the mean at zero, while preserving the original variation). Standard errors clustered at country/gender/age level (parametric correction for Moulton factor). Significance levels: ***$p < 0.01$; **$p < 0.05$; *$p < 0.1$.

## Supplementary materials

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.labeco.2019.06.006.

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