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GENDER GAP CLOSURE AND REVERSAL: THE CAUSES AND CONSEQUENCES OF THE RISE OF WOMEN IN HIGHER EDUCATION IN EUROPEAN COUNTRIES

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Institute of Sociology, Czech Academy of Sciences, Prague

Call for Papers for a thematic issue on 'Revisiting the Sociology of Literature: Towards Epistemological Symmetry in Literature-and-Society Research'

Guest editors: Jan Váňa and Hernán Maltz

(Institute of Czech literature of the Czech Academy of Sciences)

Deadline for submission of abstracts: 31 March 2025

This proposal builds on our long-term effort to facilitate dialogue between literary theoretical and sociologically oriented approaches in examining the complex relationship between the social and the literary. Since the inception of the social sciences, sociology and literary studies have profoundly influenced one another. However, due to institutional and disciplinary divides, sociology has historically placed greater emphasis on the 'social' aspects of literary communication. We can observe this trend even today. Contemporary sociology of literature continues to emphasize literary production, often relegating literary works to the status of passive objects shaped by social interactions and institutions—with Pierre Bourdieu's work serving as a prominent example. Even sociologists who acknowledge the intrinsic value of literature frequently treat literary texts as resources to be translated into sociological discourse.

At the Literature & Society Laboratory we advocate for both sociologists and literary scholars to consider extra-textual and intra-textual elements to gain a more comprehensive and enriching understanding of social-literary interaction. Successive synthetic accounts show that the sociology of literature remains a fragmented and dispersed field, offering a diverse array of approaches, with Bourdieusian field theory being just one among many. We argue that the existing epistemological asymmetry in this area of study can be addressed through a dialogue with those whose professional calling is to understand literature—with literary theorists and critics. Therefore, the proposed issue aims to bring together sociologists, cultural scholars, literary theorists, historians, and critics to explore various facets of the literary-social link, providing rich and compelling analyses through their collaborative effort.

We welcome analyses that engage with texts and contexts, fictions and discourses, theories and practices, themes and motifs, traditions and novelties, spaces and trajectories, and more. We are particularly interested in submissions that critically examine the category of 'sociology of literature' itself and explore its connections with various fields of knowledge, especially in areas where sociology and literary studies intersect or compete. Topics of interest include, but are not limited to:

- Case studies examining the creation and/or impact of literary works across different cultural and historical contexts.
- Research on literary production (such as authorship, publishing institutions, and editorial practices), readership (including reader interviews, statistical surveys, and discourse analyses of reception), and literary fields that combine sociological perspective with close reading of relevant literary texts.
- Methodological contributions that propose new ways of conceptualizing the relationship between literature and society or that draw on the analysis of specific literary works within their social contexts.
- Theoretical contributions that discuss models, paradigms, and schools of thought, or propose new theoretical approaches.
- Intellectual history exploring how different schools and paradigms have approached—and are approaching—the study of the literature-society relationship, and how these approaches have evolved over time.

Instructions for authors

Please send a working title and an abstract of 250–500 words, outlining your central research question, theoretical framework, and general argument or expected findings to casopis@soc.cas.cz, vana@ucl.cas.cz, and maltz@ucl.cas.cz.

Authors whose abstracts are accepted will be invited to submit full papers, which will undergo a fully anonymized peer review process. Please note that an invitation to submit a full paper does not guarantee publication in the issue, as acceptance is subject to the outcome of the peer review process.

Deadline for submission of abstracts: **31 March 2025**

Decisions on abstracts communicated to accepted authors: Late April 2025

Submission of full papers: **31 October 2025**

For any additional information or questions regarding the issue, please contact the guest editors vana@ucl.cas.cz and maltz@ucl.cas.cz.

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Thematic Issue

Gender Gap Closure and Reversal: The Causes and Consequences of the Rise of Women in Higher Education in European Countries

Guest Editors

Tomáš Katrňák and Tomáš Doseděl
Faculty of Social Studies, Masaryk University, Brno

2024

Call for Papers for a thematic issue on 'Building a Responsible Future: Civic and Environmental Education in Central Europe'

Guest editors: Petr Bláha (Jan Evangelista Purkyne University), **Roman Kroufek** (Jan Evangelista Purkyne University) & **Vladimír Naxera** (University of West Bohemia)

Deadline for submission of abstracts: 16 February 2025

Note: We accept manuscripts in both Czech and English for the special issue

The Central European region is currently facing a number of challenges of a complex nature, including ongoing climate change, problems of social inequality and weakening democratic citizenship. This calls for innovative and integrative approaches in education that enable learners to better understand the links between society and the environment. It seems that the integration of civic and environmental education can play a key role in preparing the younger generation for the global challenges of the future.

The synthesis of civic education, reflecting active citizenship, critical thinking and a deeper understanding of democratic processes, with education for a sustainable future, reflecting environmental issues, creates a potentially comprehensive educational framework for exploring the complex relationships between human activity and nature.

Western Europe, especially the Scandinavian countries, already work routinely with the integration of citizen and environmental education. Their education systems place a strong emphasis on interdisciplinarity and the practical application of knowledge, leading to greater civic engagement and environmental responsibility among their citizens. Students are generally encouraged to participate actively in community projects and to develop critical thinking in the context of sustainability.

Education systems in the CEE region generally keep these areas separate. As a result, students are not prepared to solve future and current problems and do not acquire the necessary skills to do so. This is usually due to a lack of resources, limited state support and a preference for traditional approaches to teaching.

In view of the above, the aim of the forthcoming special issue is primarily to map the state of current research on civic and/or environmental education in the region and, above all, to promote dialogue between the two fields.

We welcome contributions that focus on (but need not be limited to) the following general topics:

- Case studies of successful models for integrating civic and environmental education.
- Review articles describing the current understanding of the topic.
- Comparative analyses of approaches between Central Europe and other regions.
- Theoretical papers offering new concepts and frameworks for linking these educational areas.
- Innovations, including the development of new teaching methods, tools and materials.

PIf you are interested in being published in the forthcoming monothematic issue, please send an abstract (250–500 words) to the editorial office at casopis@soc.cas.cz and to the guest editors at petr.blahaml@ujep.cz, roman.kroufek@ujep.cz, and vnaxera@ff.zcu.cz by **16 February 2025**. Please mention 'civic education' in the subject line. Notification of acceptance or non-acceptance of abstracts will be sent by 28 February 2025. The deadline for submission of manuscripts for peer review is **31 August 2025**. Publication of papers will depend solely on the outcome of the standard peer review process. We also welcome book reviews and reports relevant to the topic of the monothematic issue. The monothematic issue is scheduled for publication in 2026.

If you have any questions or concerns, please feel free to contact the guest editors.

Gender Gap Closure and Reversal

Since at least the middle of the 20th century, sociology has paid significant attention to inequalities in access to higher education and the search for ways to reduce them. A wide range of work has addressed the impact of parental education, parental employment, family wealth, or class background, while others have looked at the impact of ethnicity and living in an excluded area. Other texts have evaluated the impact on inequalities of increasing school capacity or removing barriers in the form of tuition fees. Despite these in-depth analyses, there has been an almost unnoticed change, the reasons for which we have not yet been able to uncover and the consequences of which we can only cautiously estimate. That change is the reversal of the gender gap in tertiary education.

The gender gap reversal is a new and unique situation in the history of education systems. Women have been a severely marginalised group throughout the existence of universities. Until the end of the 20th century, the proportion of university-educated women was still low. It was only after the Bologna Declaration (in 2000) that the situation began to change rapidly. The massive increase in the capacity of universities, coupled with labour market changes and continued women's emancipation, has led to a rapid increase in the proportion of women in higher education. More female students began to enrol in university than male students, which, a few years later, led to a gradual change in the structure of graduates in favour of women. We are now in a situation where we can identify an overrepresentation of women among college students as well as among university graduates. The gender gap reversal has become a global phenomenon and one of the highlights of women's emancipation.

The higher proportion of women with tertiary education compared to men has causes that have not yet been reliably identified by sociologists. It is generally said that women are more ambitious, more conscientious during their school career, and have higher educational aspirations, and that discrimination based on gender has disappeared. However, this phenomenon has also specific consequences for the labour market, assortative mating, partnership and family life, and public and political engagement. In all these areas, we can expect changes in the traditional roles of men and women. This special issue of the *Sociologický časopis/Czech Sociological Review* offers four interesting texts that try to reflect at least partially on the gender gap reversal (GGR).

The first text by Tomáš Katrňák discusses the impact of the GGR on inequality of educational opportunity (IEO) by educational origin in the transition to tertiary education. Using data from the European Social Survey (ESS), the author shows that the post-2000 educational expansion has indeed slightly reduced inequality of educational opportunity, but it has also strengthened the gender effect. IEO decreased more for women than for men in the tertiary education transition.

This gender-specific effect is especially identified in higher educational origins, which indicates that family origin and gender equalisation at expanding tertiary education are competing processes. The decrease of IEO from bottom of educational structure means an increase in gender inequality, albeit reversed in the top of educational structure.

Petr Kužel's text shifts the focus from the educational system to the labour market. Specifically, he focuses on issues relating to the adjusted gender pay gap (AGPG). In his study, he shows the main gaps in gender pay inequality. Kužel demonstrates that the currently high GPG rates in European countries are at least partly due to inappropriate calculation methodologies because age instead of total job experience is the category used in the calculations. His study should thus contribute to the more accurate measurement of this phenomenon. We can certainly expect that income inequalities between men and women will become the target of several analytical and political activities, since, at a time when women are on average more highly educated than men, differential pay for the same work is not justifiable.

The third text was prepared by a group of authors – Magdalena Adamus, Denisa Fedáková, and Vladimíra Čavojová. They take another look at the GGR, this time focusing on the cohabitation of tertiary-educated partners in a shared household. By conducting the survey among HEED and STEM university students, who differ significantly in gender structure, the authors aimed to find out what expectations students have about the future division of unpaid work in their households. The results of the research have shown that even the significant increase in the number of women in tertiary education has not yet led to a shift in traditional gender roles in household care. Both male and female students in the survey stated that women should do more of the unpaid work in their future households than their male counterparts.

The last text, by Mike Smith, deals with social mobility in Central European countries from a gender perspective. It follows up on the first text by Tomáš Katrňák and demonstrates women's higher mobility chances compared to men's. This gender gap in favour of women is increasing in time. The causes for this increase are seen in the changes occurring in the occupational structures in the countries analysed in this text. This suggests that women are more effectively taking advantage of the changes in the labour market (a continuous upgrade of occupational structure that goes hand in hand with technological changes) than men.

Let's hope that this issue of the *Sociologický časopis/Czech Sociological Review* is the first stepping stone on which further research on individual aspects of the educational gender gap reversal will be built. We wish you pleasant and inspiring reading.

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The Gender-gap Reversal in Tertiary Education and Its Implications for Inequality of Educational Opportunity in European Countries*

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Abstract: Tertiary education has expanded in European countries since 2000. One consequence of this expansion is the growth of the gender-gap reversal (GGR), in which proportion of women in tertiary education is increasing faster than that of men. This article deals with the historically new gender arrangement of tertiary education. It answers the question of whether GGR, as part of educational expansion, means different gender trends in inequality of educational opportunity (IEO) by educational origin in the tertiary education transition. The author analyzed European Social Survey (ESS) data on the 25–34 age group from 20 European countries over five rounds (2002, 2006, 2010, 2014 and 2018). A three-level (random) binary logistic regression model was used to cover individual variables by period by country. The results show that the recent educational expansion has slightly weakened the IEO in tertiary education transition and that it is significantly different for men and women. Gender is important in IEO in a time of GGR. The author discusses what the empirical results mean for the theory of maximally maintained inequality (MMI), which is used in social stratification research as a general explanation for persistent inequality in a time of educational expansion.

Keywords: gender-gap reversal, educational expansion, inequality of educational opportunity, European countries

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Introduction

The educational structures of European populations have undergone a significant change over the past 20 years. In 2000, the Bologna process was launched with the aim of increasing the proportion of people with tertiary education in EU countries. New educational institutions arose, new study fields emerged and the process of enrolling more people in university studies was initiated (Kogan, 2012). The result was that from 2000 to 2020, the average proportion of young people (aged 25–34) in EU countries who had attained tertiary education increased from 25% to 41% (Eurostat Statistics, 2022). This meant a slight decrease in the inequality of educational opportunity (IEO) by educational origin in this age group (cf. Katrňák & Hubatková, 2022). A concomitant process of this educational expansion was the change in the gender ratio in favour of women in some countries and the growth of gender-gap reversal (GGR) in other countries. Women's odds of transitioning to tertiary education have increased faster than men's, as has their probability of completing tertiary education. The numbers of women with completed tertiary education in the cohorts who have left the education system exceed those of men in all European countries.[1]

Some sociologists have formulated the consequences of GGR for events such as assortative mating (De Hauw et al., 2017; Esteve et al., 2012), marriage stability and divorce risks (Van Bavel et al., 2018), and the division of paid and unpaid work inside marriages (Van Bavel, 2012). Others have addressed the consequences of GGR for labour market outcomes, especially for gender gaps in income (Klesment & Van Bavel, 2017), labour market positions (Goldin et al., 2006), and living standards (DiPrete & Buchmann, 2006). This article analyzes the consequences of GGR for IEO by educational origin. The key questions of this article are as follows: Could GGR, as part of the recent expansion of tertiary education, also mean changes in IEO by gender in the transition to tertiary education? Does the change in the proportion of men and women in tertiary education in European countries lead to differences in trends in IEO by gender in access to this education?

The analysis of IEO is aimed at the tertiary education level in European countries for two reasons: first, because of its massive expansion in the last two decades (2000–2020) and second, because more than half of the contemporary IEO in completed education is generated at this educational level (Katrňák & Hubatková, 2022). Moreover, GGR at the tertiary education level is a completely new situation in the history of European education systems, creating conditions for a natural experiment (Dunning, 2012) to measure IEO. Due to the different expansion intensities for the two gender groups, it is possible to test sociological theories about the effects of educational expansion on the IEO for men and women separately ('persistent inequality' versus 'non-persistent inequality'). From the view-

[1] The same trend can be seen in many other countries (cf. Schofer & Meyer, 2005; Vincent-Lancrin, 2008). DiPrete and Buchmann (2013) wrote about a global GGR phenomenon.

point of 'persistent inequality' (Shavit & Blossfeld, 1993), which does not assume that expansion weakens IEO, it is expected that IEO would not change differently by gender because IEO should not change at all. From the viewpoint of 'non-persistent inequality' (Breen et al., 2009), according to which educational expansion weakens IEO, there is room for differential weakening of IEO by gender when there are differences in the pace of expansion by gender (Blossfeld et al., 2017).

In this article, I target a young age group (25–34 years) and analyze the data from five European Social Survey (ESS) rounds (2002, 2006, 2010, 2014 and 2018) for 20 European countries. I use a three-level (random effects) logit model in which the dependent variable is a binary for transition to tertiary education (1 'success', 0 'failure') and the key explanatory variables are gender and educational origin (controlling for parental educational heterogamy and parents' labour market position; level-1 of analysis). The tertiary education transition occurs in the context of periods (level-2 of analysis), for which I use the variable of educational expansion, and in the context of countries (level-3 of analysis). The analytical strategy is based on three two-way and one three-way model interactions between the three key explanatory variables. The two-way interactions are (1) educational origin and gender, which indicate gender differences in IEO; (2) educational expansion and gender, which indicate trends in GGR; and (3) educational expansion and educational origin, which indicate trends in IEO. The three-way interaction is among these three variables (origin effect by expansion by gender), which indicates whether expansion is associated with IEO changes differently for men and women.

The results show that gender and trends in GGR matter. Recent educational expansion has slightly weakened the IEO in the transition to tertiary education by educational origin in European countries. The probability of transition to tertiary education increases faster for lower educational origins than for higher ones. This probability is significantly differentiated by gender over educational expansion levels. These results are not consistent with 'persistent inequality' results (Shavit & Blossfeld, 1993) or with the theory of maximally maintained inequality (MMI; Raftery & Hout, 1993) that is used as a general explanation of persistent inequality.

Educational expansion and gender-gap reversal in European countries after 2000

While European systems of higher education have been expanding since at least the 1960s (Boliver, 2011; Haim & Shavit, 2013; Schofer & Meyer, 2005),[2] the expansion after 2000 is closely linked to structural changes brought by the Bologna

[2] In the late 1960s and the 1970s, the number of university students more than doubled in many Western European countries, which can be marked as the first educational expansion period (Trow, 1973).

process (Bologna Declaration, 1999; Keeling, 2006). The main objective of this process was to establish a single European higher education area within which national higher education systems would be mutually comparable and compatible. This end-state would be achieved through the adoption of comparable degrees and credit systems, as well as the adoption of three main study cycles: a widely accessible bachelor's cycle, a more selective master's cycle and a scientific doctoral cycle (i.e., ISCED97 levels 5 and 6; Bologna Declaration, 1999). Later additions included an emphasis on employability, lifelong learning, and – crucially – the reduction of IEO in access to higher education. All of these changes were in response to the need for highly educated people in European labour markets (Kogan, 2012). Tertiary education has progressively become a necessary entry-level qualification for obtaining a well-paid job, so more individuals are striving to attain it (Marginson, 2016).

In 2020, the average proportion of women with a tertiary education in the 20 analyzed countries was 50.6%; the average proportion of men was 37.6%. The number of women with tertiary education had increased (on average) by 1.9 times (about 24.5 percentage points) since 2000; men had increased by 1.7 times (about 15.3 percentage points). The fastest changes in the gender ratio in favour of women (from 2000 to 2020) occurred in Switzerland, the Czech Republic, Slovakia, Austria, Great Britain, and Germany – in countries where a GGR did not exist or was very low in 2000. No change or a slightly negative trend in the gender ratio in favour of women was seen in France, Finland, Estonia, Portugal and Slovenia – in countries where a GGR already existed in 2000 (cf. Table 1). In all countries, the change in the gender ratio was not due to stagnation or even a decline in men's tertiary education attainment. It was due to a faster increase in the educational attainment of women.

Many factors have contributed to this shift in the gender ratio. From a historical point of view, these factors include changes in the labour market, marriage and family lives (DiPrete & Buchmann, 2013). These changes produce greater opportunities and incentives for women to obtain more education. The economic and social rewards for higher education increase more greatly for women than for men, so women are motivated to achieve higher education. Women's participation in the labour force increases because labour markets demand specific occupations requiring higher education. Higher education for women means greater insurance against poverty, a higher standard of living and the probability of a better match in family life. Parental incentives directed at children by gender have been equalized inside families, and because girls perform better than boys at university (higher level of preparedness, differences in school-related attitudes, and lower dropout rate), gender differences occur not only in enrolment in higher education but also in the likelihood of completion (DiPrete & Buchmann, 2013). The combination of all these factors led first to closing the gender gap and then to GGR in higher education in advanced countries. The educational expansion in European countries since 2000 has accelerated this process and made it more obvious.

Table 1. The ratio of women to men with tertiary education (ISCED11 5–8 levels) for 25–34 age group by years and countries

Country	Code	Year 2000	2004	2008	2012	2016	2020	Difference in ratios 2000–2020
Switzerland	CH	0.51	0.62	0.84	0.97	1.00	1.09	*2.15*
Czechia	CZ	0.97	1.03	1.25	1.41	1.47	1.55	*1.59*
Slovakia	SK	1.08	1.15	1.33	1.53	1.61	1.68	*1.55*
Austria	AT	0.90	0.96	1.08	1.14	1.19	1.23	*1.35*
Great Britain	GB	0.92	1.03	1.09	1.11	1.10	1.15	*1.26*
Germany	DE	0.87	0.97	1.08	1.16	1.06	1.07	*1.24*
Denmark	DK	1.19	1.20	1.23	1.60	1.36	1.42	*1.20*
Italy	IT	1.29	1.45	1.58	1.57	1.63	1.53	*1.18*
Netherlands	NL	1.02	1.09	1.15	1.19	1.25	1.20	*1.18*
Belgium	BE	1.18	1.25	1.34	1.40	1.34	1.38	*1.17*
Norway	NO	1.29	1.35	1.48	1.46	1.40	1.45	*1.13*
Spain	ES	1.20	1.26	1.30	1.32	1.34	1.30	*1.08*
Poland	PL	1.52	1.43	1.51	1.55	1.58	1.61	*1.06*
Hungary	HU	1.38	1.38	1.43	1.48	1.50	1.44	*1.04*
Ireland	IE	1.11	1.22	1.34	1.29	1.28	1.14	*1.03*
France	FR	1.17	1.18	1.24	1.23	1.24	1.15	*0.98*
Finland	FI	1.52	1.62	1.67	1.60	1.47	1.45	*0.95*
Estonia	EE	1.83	1.64	1.40	1.74	1.68	1.73	*0.95*
Portugal	PT	1.57	1.70	1.77	1.46	1.59	1.42	*0.90*
Slovenia	SI	1.93	1.94	1.71	1.80	1.72	1.58	*0.82*
mean		*1.17*	*1.23*	*1.31*	*1.37*	*1.36*	*1.35*	*1.15*

Source: Eurostat statistics (2022).

Note: The ratio is computed as the proportion of women vs. proportion of men with tertiary education. Countries are sorted from the highest to the lowest ratio difference between years 2020 and 2000 (last column in the table).

Should we expect a change in IEO by gender during a time of GGR?

The decline of the gender gap in educational attainment in the 20th century has been documented in many social stratification studies (cf. Breen, 2004; Breen & Müller, 2020; Erikson & Goldthorpe, 1992; Shavit & Blossfeld, 1993). These studies referred to various periods of the 20th century in different countries. They did not relate to specific educational expansions and they did not measure the direct effect of particular educational expansions on IEO by family background (social or educational origin). They assumed that educational systems expanded evolutionarily during the 20th century (the shift of an increasingly large population towards higher levels of education), and they analyzed the changes in IEO against the background of this expansion.

The findings of these studies can be summarized in two general conclusions. According to the first and prevalent one (Breen & Jonsson, 2005), the expansion of the educational system in most countries does little to help economically and socially disadvantaged groups to gain degrees and qualifications. The IEO remains more or less stable regardless of educational expansion (cf. Alon, 2009; Blossfeld et al., 2015; Boliver, 2011; Haim & Shavit, 2013; Hannum & Buchmann, 2005; Raftery & Hout, 1993; Shavit, 2011; Shavit & Blossfeld, 1993; Triventi, 2013). According to the second conclusion, educational expansion widens the proportion of people with higher levels of education from different family backgrounds. There is a negative relationship between expansion and IEO, although this does not necessarily occur with the same intensity in all countries (cf. Ballarino et al., 2009; Bernardi & Ballarino, 2016; Blossfeld et al., 2017; Breen et al., 2009; Brown, 1995; Katrňák & Hubatková, 2022; Liu et al., 2016; Thélot & Vallet, 2000).[3]

The first conclusion about 'persistent inequality' (Shavit & Blossfeld, 1993) is explained by the theory of MMI (Raftery & Hout, 1993). MMI was originally formulated for Irish society. It acquired wider support through an over-cohort comparative analysis of trends in IEO in 13 countries (cf. Shavit & Blossfeld, 1993). According to this theory, educational opportunities are primarily used by higher social classes rather than by lower social classes. This is because the demand for higher education is more widespread among higher social classes. Even if access to higher education increases, class inequalities in education (measured as the odds ratios of educational transitions) are maintained because students from lower-class backgrounds lack the appropriate knowledge and skills to navigate the educational system towards successful transitions to higher education levels.

[3] These two conclusions (no-relationship and negative relationship) can be joined by a third conclusion about a positive relationship between educational expansion and IEO (cf. Haim & Shavit, 2013; Halsey et al., 1980). This can happen when there is a different 'reaction' to educational expansion according to family origin (higher social classes 'react' faster than lower classes and take most of the advantages connected with expanded educational stages). Because this conclusion is not widely empirically supported in social stratification research (cf. Erikson & Jonsson, 1996), I do not address it further in this text.

When the demand for higher education is 'saturated' for the upper classes, then the educational opportunity window also opens to the lower classes and IEO begins to decrease.

The second conclusion about the negative relationship between IEO and educational expansion can be called 'non-persistent inequality' (Breen et al., 2009). It lacks a consistent explanatory theory such as MMI in the case of the first conclusion. Some factors have been identified as contributing to declining IEO during times of educational expansion. These include structural factors, such as vertical and horizontal differentiations of higher education in degrees, tracks and study fields (Reimer & Jacob, 2011); the increase of homogeneity on unobserved variables in population education transitioning (Katrňák & Hubatková, 2022); a change in meritocratic selection and in competition at expanded educational levels (Alon, 2009, 2014; Treiman et al., 2003); a change in the perception of educational failure risks (Ballarino et al., 2009; Ballarino & Schadee, 2010; Erikson & Jonsson, 1996); and the equalization of living conditions and levelling of educational barriers – especially the costs of education (Erikson & Jonsson, 1996; Jonsson, 1993).

When studies about 'persistent' and 'non-persistent' inequality have measured trends in IEO for men and women separately, they have not shown that these trends lead to a systematically different change in IEO by gender. In these studies, class inequalities in IEO have been identified as more important and surpassing gender differences. From the standpoint of 'persistent inequality', this conclusion is logical, because this theory does not assume a relationship between educational expansion and IEO changes. Even though the proportion of women with completed higher education increases faster than that of men (which means that women's opportunities for enrolment and study at higher education levels are increasing), this does not mean that a faster decrease in IEO for women than for men should be expected on the basis of this theory. From the standpoint of 'non-persistent inequality', this conclusion is inconsistent with the result that educational expansion weakens the IEO. If educational opportunities increase for women faster than for men, there should also be a faster decrease in IEO for them.[4]

Studies in the Netherlands (De Graaf & Ganzeboom, 1993), Sweden (Jonsson, 1993), Great Britain, Germany (Jonsson et al., 1996), and the United States (Buchmann & DiPrete, 2006) did not find systematic different decreases of IEO by gender across birth cohorts in the 20th century. Gender differences in IEO were only identified in earlier cohorts, but across cohorts, they converged. This 'earlier' gender equalization is understood in terms of women's changing attitudes and ambitions in regard to education, and not by changes in educational systems themselves (cf. Bukodi & Goldthorpe, 2019). Breen et al. (2010) analyzed seven

[4] Using PIAAC data from 22 countries, Blossfeld et al. (2017) showed the relationship between the pace of educational expansion and the decrease of IEO. More rapid educational expansion leads to greater decrease of IEO and vice versa.

European countries over cohorts born between 1900 and 1965 and confirmed these findings. Gender differences in IEO were identified in earlier cohorts in some countries, but they were not systematic or consistent. The trend towards a general decline of IEO rather than no change ('persistent inequality') is much more regular than differences in this trend over gender, according to Breen et al. (2010).

Based on these studies, I assume that regardless of whether IEO remains stable or declines during educational expansion, it does not happen significantly differently for women and men, although there is GGR at the expanded tertiary education level. Thus, although the assumption about the relationship between educational expansion and IEO is twofold ('persistent' and 'non-persistent' trends), this relationship should not differ by gender. The hypothesis asserts that IEO is 'persistent' or 'non-persistent' regardless of gender at the time of recent tertiary education expansion. This hypothesis was tested in the analysis.

Data and analyzed variables

To answer the question of whether the GGR trend in tertiary education in European countries also means IEO gender differences in the transition to this education, I used ESS data from 2002, 2006, 2010, 2014 and 2018. I analyzed 20 European countries, aiming for respondents aged 25–34 years. These young respondents are the most strongly influenced by education expansion. A comparison of the same age group over the educational expansion period is much more reliable than a comparison of all survey respondents. The numbers of respondents in the ESS data by country and year are listed in Table A1 in the Appendix.[5]

The respondent's *completed tertiary education* is a key variable. In the ESS data, the highest level of a respondent's education is measured by the harmonized categorical variable ISCED (cf. Schneider, 2013). I collapsed this variable into four categories (primary, lower secondary, upper secondary and tertiary education). These categories retrospectively delimit transitions from one educational level to the next (Mare, 1980, 1981). For the population exposed to tertiary education transition (upper secondary and tertiary education attained), I created a dummy variable: 0 – did not transition, and 1 – transitioned to tertiary education (56.4% of respondents did not transition and 43.6% transitioned on average over countries and ages 25–34 in ESS data).[6]

[5] Appendix is available online at https://doi.org/10.13060/csr.2024.008. For 2002, I used the Integrated file edition 6.6 round 1; for 2006, the Integrated file edition 3.7 round 3; for 2010, the Integrated file edition 3.4 round 5; for 2014, the Integrated file edition 2.2 round 7; and for 2018, the Integrated file edition 3.0 round 9. As not all countries took part in all ESS rounds, the closest available year was used to replace any missing year. This applied to four countries: the Czech Republic, Italy, Estonia, and Slovakia.

[6] Respondents who had attained only primary and lower secondary education were excluded from the analysis, because they are structurally excluded from 'the risk' of tertiary education transition.

There are four individual-level explanatory variables. The first is *gender* (a dummy variable: 0 – man, 1 – woman). The second variable is *educational origin*, defined as the highest level of parental education. Like a respondent's education, the father's and mother's highest education is measured in the ESS data by the categorical variable ISCED. I transformed this variable into one continuous variable, ISLED (international standard level of education; cf. Schröder & Ganzeboom, 2014), for the father and mother separately and then indicated educational origin by the higher of these two values.[7] I used continuous ISLED for educational origin because the variability of this explanatory variable in level-1 can be more reliably analyzed by a contextual variable (levels of educational expansion). I worked with a standardized ISLED variable (z-scores: mean = 0, SD = 1, ages 25–34, standardization around the grand mean of this variable in the data).

Besides these two key explanatory variables, in the transition to tertiary education, I controlled for *parental educational heterogamy* and *parents' higher social origin.* To construct parental educational heterogamy, I used a harmonized categorical of the father's and mother's highest levels of education (in ESS data: primary, lower secondary, upper secondary, and tertiary education). Because I indicate educational origin by the more highly educated parent, I was able to create a heterogamy variable by combining these four categories of both parents: 0 – the same education of parents (homogamy); 1 – father's lower education; and 2 – mother's lower education (for more about combining parental variables into one, cf. Thaning & Hällsten, 2020, and especially with respect to the gender of parents, cf. Ballarino et al., 2021). To identify parents' social origins, I used variables capturing the father's and mother's occupations when the respondent was 14 years old. I created three categories differentiating three basic types of occupations in European labour markets: (0) semi-routine and routine manual and service occupations; (1) technical, craft occupations and farmers; and (2) professional, clerical and intermediate occupations, managers and administrators. Because the analysis compares countries at different time points, I used ESS post-stratification weights including design weights (variable *pspwght* in ESS data).

The data has a hierarchical structure: on level-1 there are individuals; these individuals are nested in years at level-2, then nested in countries at level-3 (together, 100 contexts are created by five years in 20 countries). Because of this country–year structure (Schmidt-Catran & Fairbrother, 2016),[8] I analyzed the data with multilevel (random regression) models (cf. Kreft & Leeuw, 1998; Gelman & Hill, 2006).

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[7] The ISLED variable is absent from the original ESS dataset. However, this variable may be easily imputed using a prepared syntax (cf. Harry Ganzeboom website: http://www.harryganzeboom.nl/ISLED/isled_56.txt).

[8] These data can be labelled as 'comparative longitudinal survey data' (CLSD). The data are drawn from multiple waves of comparative surveys inside individual countries, meaning that they are comparative over countries as well as longitudinally inside countries (cf. Fairbrother, 2014).

Table 2. Descriptive statistics

Variable	Mean	Std. Dev.	Min.	Max.	Categories
Individual variables (level-1)					
Tertiary education transition	0.44	0.50	0	1	0) no; 1) yes
Gender	0.53	0.50	0	1	0) man; 1) woman
Educational origin	0.00	1.00	−1.69	2.16	highest parents' ISLED standardized
Parents' educational heterogamy	0.58	0.82	0	2	0) the same education (homogamy); 1) father's lower education; 2) mother's lower education
Occupational origin	1.05	0.84	0	2	0) semi-routine and routine manual and service occupations; 1) technical, craft occupations and farmers; 2) professionals, clerical and intermediate occupations, managers and administrators
Contextual variables (level-2)					
Educational expansion	0.00	9.55	−16.70	21.35	grand mean centering
Contextual levels					
Period (level-2)	–	–	1	5	
Country (level-3)	–	–	1	20	

Recent *educational expansion* is a macro time-varying variable. It is measured by the proportions of people aged 25–34 with tertiary education (indicated by ISCED11, levels 5 to 8) in each period and country (together, 100 numbers given by 5 periods in 20 countries). Respondents in the 25–34 age group in the 2002 ESS data were born in 1968–1977 and passed into tertiary education in 1987–1996.[9] Respondents aged 25–34 in the 2006 ESS data passed into tertiary education in

[9] The time windows defining ESS respondents' passing to tertiary education were computed as follows: ESS round year minus age 25–34 plus age of 6, which I consider the general age for beginning primary education, and the most common duration of each educational level: primary education 6 years, lower secondary 3 years, and upper secondary 4 years (cf. UIS, 2012).

1991–2000, respondents in the 2010 ESS data in 1995–2004, respondents in the 2014 ESS data in 1999–2008, and respondents in the 2018 ESS data in 2003–2012. For each of the years' windows of tertiary transition, I computed the average proportion of people who completed their tertiary education according to the Eurostat statistics (2022). I used the differences in these average proportions of completed tertiary education to measure the trends in educational expansion (how much tertiary education was available) in the analyzed periods and countries (level-2 of the analysis). Table A2 in the Appendix presents these average proportions. I standardized these values by grand mean centring in the analysis (cf. Gelman & Hill, 2006; Kreft et al., 1995; Raudenbush & Bryk, 2002).

Table 2 presents the descriptive statistics of all analyzed variables and the description of their categories.

Methods and statistical analysis

With regard to the hierarchical structure of the data and the binary dependent variable (failure or success in tertiary transition), I used three-level (random) logistic regression models (Gelman & Hill, 2006; Rabe-Hesketh & Skrondal, 2012). Using the random component in these models, I was able to analyze the direct effects of educational expansion (identified at period levels) on the variability of educational origin and gender on tertiary education transition over periods in countries. The specification of the three-level random-intercept logit model for passing tertiary educational transitions is as follows:

$$E(y_{ijk} = 1 \mid \mathbf{x}_{ijk}, \zeta^{(2)}_{jk}, \zeta^{(3)}_{k}) = \hat{p}_{ijk} = \Lambda(\beta_0 + \beta_1 x_{1ijk} + \ldots + \beta_6 x_{6ijk} + \beta_7 x_{7,jk} + \zeta^{(2)}_{jk}, \zeta^{(3)}_{k}) \quad (1)$$

where $\Lambda(\cdot)$ denotes the cumulative logistic distribution function $\Lambda(\cdot) = \frac{e^{(\cdot)}}{1+e^{(\cdot)}}$ and y_{ijk} is the expected value of 1 (success in transition) with respect to observed variables. This is the conditional probability $\hat{p}_{ijk}$ that respondent i nested in period j nested in country k passes tertiary transition if they are exposed to this transition. $\mathbf{x}_{ijk} = (x_{1ijk}, \ldots, x_{6ijk}, x_{7,jk})'$ is a vector containing all level-1 covariates (x_{1ijk} = gender – woman; x_{2ijk} = educational origin; x_{3ijk} = parental educational heterogamy – father's lower education; x_{4ijk} = parental educational heterogamy – mother's lower education; x_{5ijk} = occupational origin – technical and craft occupations; x_{6ijk} = occupational origin – professional and administrative occupations, and level-2 covariate ($x_{7,jk}$ = educational expansion) with estimated intercept β_0 and parameters $\beta_0 \cdots \beta_7$. $\zeta^{(2)}_{jk}$ is a random intercept varying over periods (level-2) and $\zeta^{(3)}_{k}$ is a random intercept varying over countries (level-3).

Table 3 presents the estimated random logistic regression models.[10] I began with the null random intercept model M0, in which no covariates were consid-

[10] All models are estimated using the Stata command *melogit*. I will provide the do-file with data on request.

Table 3. Random logistic regression models for tertiary education transition in age group 25–34 – first part

	Variable	Levels	M0	M1	M2	M3	M4
Individual level	Gender	*man*		*ref.*	*ref.*	*ref.*	*ref.*
		woman		0.446***	0.446***	0.455***	0.446***
				(0.030)	(0.030)	(0.041)	(0.040)
	Edu origin		0.572***	0.569***	0.606***	0.579***	
			(0.019)	(0.020)	(0.032)	(0.035)	
	Parents' heterogamy	*homogamy*		*ref.*	*ref.*	*ref.*	*ref.*
		father's education lower		−0.190***	−0.191***	−0.213***	−0.218***
				(0.044)	(0.044)	(0.045)	(0.045)
		mother's education lower		−0.255***	−0.251***	−0.264***	−0.270***
				(0.040)	(0.040)	(0.041)	(0.041)
	Occupation origin	*manual/service occup.*		*ref.*	*ref.*	*ref.*	*ref.*
		technical/craft occup.		0.265***	0.263***	0.261***	0.259***
				(0.039)	(0.039)	(0.040)	(0.040)
		professionals/ admin.		0.687***	0.686***	0.662***	0.657***
				(0.042)	(0.042)	(0.042)	(0.042)
	Constant		−0.205	−0.677***	−0.693***	−0.683***	−0.670***
			(0.132)	(0.145)	(0.117)	(0.119)	(0.119)

Table 3. Random logistic regression models for tertiary education transition in age group 25–34 – second part

	Variable	Levels	M0	M1	M2	M3	M4
Contextual level	Educational expansion			0.028***	0.022***	0.022***	
					(0.006)	(0.007)	(0.007)
Interactions	Gender*Edu origin					0.074**	
							(0.033)
	Gender*Edu expansion					0.009**	
							(0.004)
	Edu origin*Edu expansion					−0.008**	
							(0.003)
	Gender*Edu origin*Edu Expansion					0.003	
							(0.003)

Table 3. Random logistic regression models for tertiary education transition in age group 25–34 – third part

	Variable	Levels	M0	M1	M2	M3	M4
Random effects parameters	Country variation constant	0.318***	0.367***	0.228***	0.239***	0.241***	
	Period variation constant	0.134***	0.117***	0.100***	0.0742*	0,072	
	Period variation Gender				0.074***	0.064***	
	Period variation Edu origin				0.057***	0.053***	
	ICC Country		0,084	0,097	0,063		
	ICC Period/ Country		0,121	0,128	0,091		
LL Model			−14,263.9	−13,043.9	−13,034.2	−12,972.0	−12,963.5
BIC			28,558	26,178	26,168	26,094	26,016
AIC			28,534	26,106	26,089	25,974	25,965
N (respondents)			21,270	21,270	21,270	21,270	21,270

Note: Number of contexts 100 (5 periods by 20 countries). standard errors in parentheses. ***$p < 0.01$. **$p < 0.05$. *$p < 0.1$

ered ($x_{1ijk},...,x_{7,jk} = 0$ in Equation 1). The intercept indicates the average odds (in periods and over countries) for respondents to pass a tertiary transition. It was 0.82(exp(−0.205)). The intercept country variance is higher than the intercept period variance, which indicates that the odds of tertiary transition differ more between countries than between periods inside the countries.

In model M1, the individual-level factors are added to the null model M0 ($x_{1ijk},...,x_{6ijk}$ in Equation 1). The LL statistics improved considerably, and the likelihood-ratio (LR) test (comparing models M0 and M1) indicated that individual variables significantly increased the model fit to the data. Women had higher odds than men of tertiary transition: 1.56 times greater (exp(0.446)), which means about 56% (100*[exp(0.446)−1]). Educational origin positively influenced the transition to tertiary education. Better educated parents use their cultural, economic and social advantages to help their offspring attain the highest educational level possible (e.g. Blau & Duncan, 1967; Breen, 2004; Breen & Müller, 2020; Erikson & Goldthorpe, 1992). Because educational origin is a standardized variable (z-scores), it must be interpreted in changes given by standard deviation (SD). For instance, if educational origin increases by 1SD, the odds of tertiary transition increase by 1.77 times (exp(0.572)).

Tertiary education transition is also influenced by parental educational heterogamy. A lower education level of one parent generally weakens the odds of transition compared to parental homogamy. Nevertheless, the gender of the parents must be considered. In our case, the father's lower education weakened the tertiary transition by about 17% (100*[exp(−0.190)−1]) and the mother's lower education by about 23% (100*[exp(−0.255)−1]) in comparison with parental homogamy.

Higher categories of social origin strengthen the odds of tertiary transition. Technical and craft occupations increase the odds by about 30% (100*[exp(0.265)−1]), and professionals, intermediate occupations, managers and administrators increase the odds by about 99% (100*[exp(0.687)−1]) compared to semi-routine, routine manual and service occupations. Because the effects of all these individual variables do not change significantly across the other (more complex) models (M2, M3 and M4), we consider them as invariant across all estimated models.

Model M2 adds a level-2 variable to model M1, which is the educational expansion ($x_{7,jk}$ in Equation 1). The effect of this variable is positive, significant and almost invariant in the other estimated models (M3 and M4). Educational expansion – the quantitative increase of tertiary education – increases the odds of passing to this level of education. LL statistics improved after including this variable in the analysis, and the LR test of the nested model M1 in M2 indicates that educational expansion significantly increased the model fit to the data.

Models M1 and M2 assume that the effects of key explanatory variables (educational origin and gender) are constant over periods (levels of educational expansion). I tested the hypothesis about 'persistent' and 'non-persistent' trends in IEO by gender and educational expansion via model interactions between educational origin, gender and educational expansion. To do this, it is necessary to allow educational origin and gender to vary randomly across periods. Model M3 is therefore the random effect model that contains the period variability of gender and educational origin (the $\zeta_{jk}^{\prime(2)}x_{1ijk} + \zeta_{k}^{\prime(2)}x_{2ijk}$ term is added to Equation 1). This variability could then be modelled by a contextual variable: educational expansion. In model M4, I added four interactions: (1) the two-way interaction between gender and educational origin (the term $\beta_8 x_{21ijk}x_{2ijk}$ is added to Equation 1); (2) the two-way cross-level interaction between gender and educational expansion (the $\beta_9 x_{1ijk}x_{7,jk}$ term is added to Equation 1); (3) the two-way cross-level interaction between educational origin and educational expansion (the term $\beta_{10}x_{2ijk}x_{7,jk}$ is added to Equation 1); and (4) the three-way cross-level interaction between gender, educational origin, and educational expansion (the term $\beta_{11}x_{2ijk}x_{7,jk}$ is added to Equation 1). The last interaction indicates the differences in the educational origin effect on tertiary transition over expansion levels for men and women (for the complete equation for model M4, see the Appendix).[11]

The interaction between gender and educational origin is positive (0.074) and significant ($p<.5$). There are two possible interpretations of this interaction. The first is that gender gaps differ by educational origin. This means that educational origin influences disparities between men and women in tertiary education transition differently (different gendered parental investments by educational origin; cf. Breen et al., 2010). The second interpretation is that the effect of educational origin varies by gender. This means that tertiary transition depends on parental origin, but this relationship is differentiated by gender (higher educational origin advantages are more frequently utilized by women than men compared to those with lower educational origins).[12]

The interaction parameter between gender and educational expansion is positive (0.009) and significant ($p<.5$). Recent educational expansion in European countries has been gendered. It influences women more than men in transitioning to tertiary education. Because women have higher chances for this transition (positive model parameter for gender), educational expansion does not mean gender equalization at this education level (cf. Shavit & Blossfeld, 1996) but disequalization or strengthening of GGR. Women are no longer approaching men; they are moving away from them.

Even though the expansion of tertiary education increases the probability of transition to tertiary education (positive and significant parameter 0.022), the interaction between educational expansion and educational origin was negative (−0.008) and significant ($p<.05$). In other words, IEO decreased over expansion levels, which does not correspond with the thesis about 'persistent inequality' in

[11] In a hierarchical approach, all these interactions are necessary when a three-way interaction is part of the statistical model (e.g. Powers & Xie, 2012).

[12] All two-way interactions can be interpreted from the perspective of both interacting variables, because formally it is one parameter in a statistical model. Nevertheless, both interpretations are not meaningful in all cases.

times of educational expansion (Shavit & Blossfeld, 1993). This does not mean, however, that the relationship between educational origin and tertiary education transition has ceased to exist. It is still relatively strong, and the last educational expansion weakened it only partially (cf. Katrňák & Hubatková, 2022).

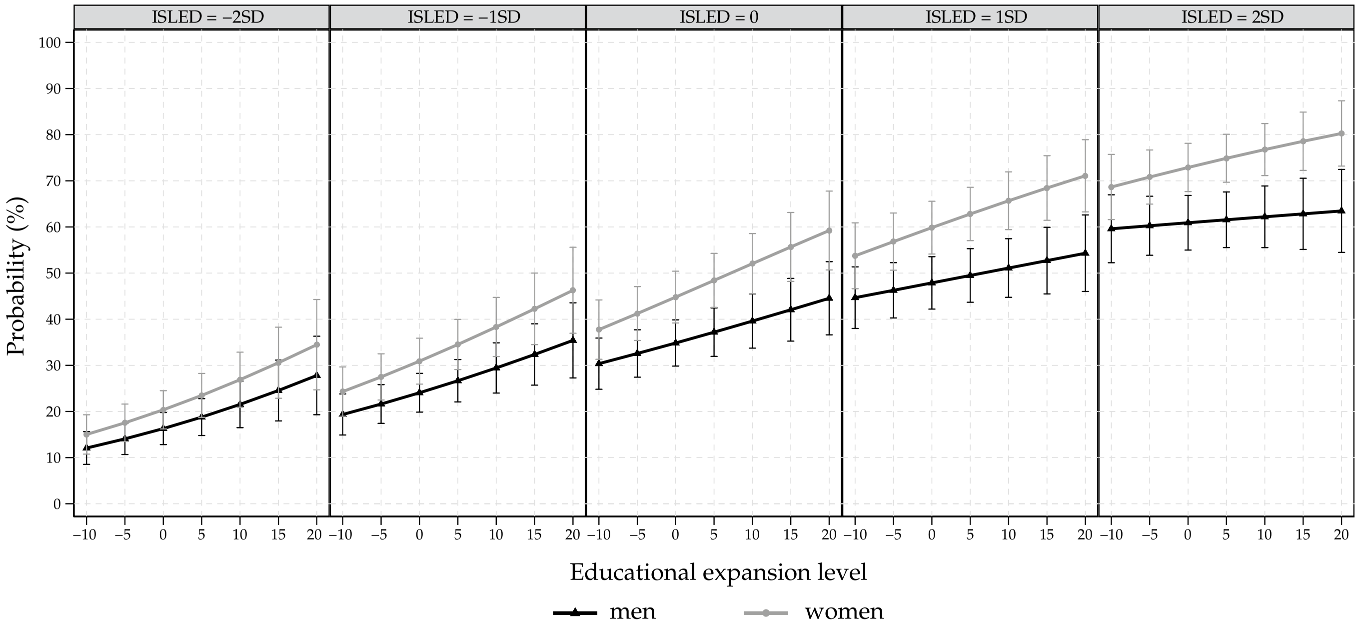
The final three-way interaction answers the key research question. This interaction was positive (0.003) but not significant. However, the no-significance model of this interaction is not appropriate for making inferences or conclusions about the differences between men and women by educational origin and educational expansion (Mitchell, 2021). Both these findings (positivity and non-significance) must be examined in more detail.

The positivity of the three-way interaction indicates that recent educational expansion decreased IEO differently by gender. Not only did more women complete a tertiary transition because of this expansion (GGR), but IEO decreased for them with higher intensity than for men. To better illustrate this three-way interaction, I estimated the conditional probability of tertiary education transition by educational origin levels (SD of average ISLED) and over levels of educational expansion for men and women separately (cf. Equation 1).[13] Figure 1 shows that this probability increased from the lower to the higher educational origin and that educational expansion raised it. In higher educational origins (1SD and 2SD of ISLED), this increase was not as steep as it was for lower origins (−2SD and −1SD of ISLED). The distance between men and women in this trend increased over levels of educational expansion, and in higher educational origins, significantly. For instance, for average educational origin (ISLED = 0), the probability of passing tertiary transition for women increased from 40% to 65% over expansion levels, which corresponds to the probability of men from the highest educational origin (ISLED = 2SD) in average educational expansion (0). In this case, the probability connected with women increased so that one educational origin category (indicated as 1SD) was overcome.

The confidence intervals presented in Figure 1 do not concern the differences between men and women in the probability of passing tertiary transition; they relate to the estimated parameters of model M4 and their difference from 0. I tested these differences using the statistical procedure *contrast* (Mitchell, 2021) for all combinations of categories of educational origin and expansion levels shown in Figure 1. The lowest gender differences were in low levels of educational expansion combined with a low level of educational origin. Together with the growth of educational expansion and increasing educational origins, gender differences also grew. All these differences are statistically significant (cf. Table A3 in the Ap-

[13] These are margins estimated from model M4, which are specified for *educational origin* (ISLED) from −2 SD to 2 SD (standard deviations), levels of *educational expansion* from −10% to 20%, *parental homogamy* and *manual and service occupations*. Stata command: margin *gender*, at(*edu_origin* = (−2(1)2) *expansion* = (−10 −5 0 5 10 15 20) *heterogamy* = 0 *occup_origin* = 0).

Figure 1. Probability of passing tertiary education transition by educational origin (ISLED) and levels of educational expansion for men and women



Note: Expansion level 0 is average expansion in %. ISLED 0 is standardized average ISLED. SD is standard deviation. Whiskers denote 95% confidence interval.

pendix). Reversed gender inequality destroys educational origin inequality with the help of educational expansion. Based on this, I reject both the assumption about the 'persistent inequality' of IEO in times of educational expansion (Shavit & Blossfeld, 1993) and the assumption about no gender-specific effect of educational expansion in IEO (Breen et al., 2010).

Consequences of empirical results for MMI theory

What do these results add to explanations of the relationship between educational expansion and IEO? In the theoretical part of this text, I introduced two conclusions about this relationship: a non-relationship or 'persistent' trend in IEO and a negative relationship or 'non-persistent' trend in IEO. I have assumed that these two conclusions are valid regardless of gender. The persistent trend in IEO has been explained by the theory of MMI (Raftery & Hout, 1993). If MMI is a universal theory, then it should be applied equally to both men and women. This means that the probabilities (or odds ratios) of transitioning to tertiary education do not change differently by gender (net of educational origin), even though enrolment is higher for women than for men. The results do not support this conclusion and call the universality of MMI theory (from the perspective of gender) into question.

By leaving out what is missing from MMI theory and has been elaborated elsewhere (cf. Breen & Jonsson, 2000; Lucas, 2001),[14] a weak point emerges in the concept of 'saturation'. According to MMI, the effect of educational expansion reaches the lower classes (and IEO measured as odds ratios starts to decline) after the saturation of higher social classes in transitions to educational stages. However, the theory does not indicate what the saturation level is, and one can only make assumptions about when saturation is reached. If it is 100%, it is empirically rare, especially at the tertiary level of education, that the demand for this level of education can be completely saturated for higher social classes. Therefore, there is only hypothetical space for the decrease in IEO. MMI should be confirmed in all cases in which IEO does not decline. Paradoxically, a vague definition of the saturation point makes it possible to confirm MMI even in situations when IEO declines. If the decrease in IEO is empirically measured, the MMI makes it possible to say that saturation has already occurred, which does not mean rejecting it.

[14] MMI is criticized for aiming only at quantitative (vertical) differences in education – whether there is a transition to a certain level of education or not. It means that MMI does not consider qualitative (horizontal) dimensions in IEO. The point is that the decline of IEO in one (vertical) dimension can lead to the increase of IEO in the other (horizontal) dimension (cf. a special issue of *American Behavioral Scientist* about Effectively Maintained Inequality [EMI], Lucas & Byrne, 2017).

Because of the vagueness of the saturation concept, it is difficult to evaluate and reject the MMI theory empirically. Lucas (2009) suggested aiming at the internal logic of MMI and testing it from the point of coherence. Lucas's formal analysis shows that MMI is tautological or internally inconsistent, depending on the scenario. In terms of tautology, the visible pattern in the empirical data is not explained by the mechanism that generates it. MMI only states that transition rates and IEO (as measured by odds ratios) remain constant unless forced to change by increasing enrolments. This is a different way of expressing what is seen in the data and includes both 'no change' and 'decline' in IEO. In terms of inconsistency, the MMI assumes that 'a margin-free measure of association is not margin-free' (Lucas, 2009, p. 471), because it claims that the relative measures (odds ratios) should begin to change due to the change of absolute measures (in marginal distributions).

The data analysis presented here adds an alternative approach to this formal analysis, aimed at conditions in which MMI should be valid. Even though MMI does not say anything about gender, as a universal theory, it should be valid regardless of whether it considers gender or any other social group (e.g. defined by ethnicity, age, geography or time criterion). Because the data analysis does not confirm this expectation, this theory should not be considered a universal theory for explaining the relationship between IEO and educational expansion. However, it is possible that MMI is still valid as an idiosyncratic theory. In this case, it could be applicable to specific social groups, regions or times, or it could be valid with some specific deviations or corrections (for more on this, cf. Shavit et al., 2007).

Conclusion

Several studies have shown that the recent educational expansion (since 2000) in European countries is not gender equal. Women benefit from it much more than men. Their odds of educational transitions increase, which leads to gender-gap reversal (GGR) in tertiary education. In this text, I started with this fact and analyzed whether the increasing proportion of women in tertiary education compared to men leads to differences in transitional IEO by educational origin. I formulated two assumptions. In the first, IEO does not change; in the second, IEO weakens in times of educational expansion. I further hypothesized that both these assumptions would hold equally for men and women, although their proportions increased differently in tertiary education after 2000 (a hypothesis that I tested empirically). The analysis yielded two basic results. First, IEO was 'non-persistent': the probability of transitioning to tertiary education increased faster for those from lower educational origins than from higher ones. The effect of educational origin has not ceased to exist in European countries. It is still alive; the last educational expansion weakened it only slightly. Second, IEO decreased more for women than for men; women's probability of transitioning to tertiary education increased significantly faster than men's probability. This gender-specific effect was especially valid for those with higher educational origins.

This text brings gender and GGR to the long-term analyzed relationship between IEO and educational expansion in social stratification research (Shavit et al., 2007). According to Shavit and Blossfeld (1996), class and gender equalizations in expanding tertiary education levels are competing processes. These authors stated that when women approach men at this educational level, gender and family origin interact; through this interaction, the education level may not accommodate the increasing demand of people from lower social classes. My results show that slight educational origin equalization occurs during times of increasing GGR, meaning when gender inequality (albeit reversed) increases. In the estimated statistical models, educational origin equalization was controlled for in gender, and gender differences were controlled for in the educational origin effect. It is possible that the gender effect would block educational origin equalization in a situation in which women were closer to men and educational systems had expanded 'evolutionarily' since the mid-20th century. However, the analyzed data are beyond this situation. The recent educational expansion in European countries is 'revolutionarily' fast, and women have moved away from men. The positive interaction between gender and educational expansion must therefore be interpreted so that the GGR (dis-equalization) is a part of educational origin equalization (lowering of IEO). Decreasing origin inequality is compensated for by increasing reversed gender inequality.

Empirical analyses of IEO over time and across countries are important for the economic, political and social development of each society. A low IEO by family origin means strong economic progress, political stability and societal fairness (Bukodi & Goldthorpe, 2019; Van de Werhorst, 2014). Educational expansion is then understood as a social policy instrument by which IEO can be changed. The theory of maximally maintained inequality (MMI) nevertheless implies the opposite. From its perspective, educational reforms that are carried out with the aim of decreasing IEO by increasing the proportion of students in higher education levels seem pointless. Because the analyzed data allowed it, the analytical approach and results have been framed as an empirical test of conditions under which the MMI theory should apply. The results indicate that MMI should not be considered a universal theory. It is not valid when both educational origin and gender are considered over levels of educational expansion. In addition to the theoretical ambiguity associated with MMI's key concept of 'saturation', and after a formal test showing that MMI asserts either a tautology or a contradiction (Lucas, 2009), this analysis is the next challenge to MMI theory as a prominent explanation of 'persistent' IEO in times of educational expansion.

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Is There Really Unequal Pay for Equal Work Between Men and Women in the Czech Republic? Problems with the Decomposition of Wage Determinants*

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Abstract: This study focuses on issues of the adjusted gender pay gap (AGPG) and problematises existing approaches to calculating this indicator, especially Eurostat's methodology. It analyses the different factors and variables with which Eurostat and other authors work, noting flaws in their measurement methods. The unadjusted gender pay gap (GPG) is typically divided into explained and unexplained parts, with the latter interpreted as the effect of unequal pay for equal work. This study demonstrates why the unexplained part might be considerably smaller than reported by existing studies (typically at 14%–15% and 17% in the case of Eurostat). What is key to determining the size of the explained part of the GPG is what productive characteristics and how many of them are included in statistical model. Existing analyses have artificially increased the adjusted part of the GPG due to simplifications in their application. For example, as this study shows, substituting the category of total job experience with the category of age has a significant impact, along with several minor shifts in the statistical analysis. When combined, these shifts are responsible for the substantial overestimation of the adjusted GPG. This study aims to eliminate these flaws and provide a theoretical and descriptive account of the reasons behind the overestimation.

Keywords: unadjusted gender pay gap, adjusted gender pay gap, Blinder–Oaxaca decomposition, equal pay, discrimination, Czechia

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Introduction

The issue of gender-based pay inequality has been resonating in European societies at an accelerated rate. As such, it has increasingly been the focus of academic studies, books, media appearances, popularisation articles, political statements and concrete national policies, action plans and European directives.[1]

Discussions of gender-based pay inequality are centred on the concept of the unadjusted gender pay gap (GPG), which is defined as the difference between men's and women's mean gross wages divided by men's mean gross wage and expressed as a percentage. The GPG tells us how many per cent lower an average working woman's income is than an average working man's income.[2] Therefore, *by definition*, the difference obtained does not reflect different productive characteristics of the two groups, such as representation in various (differently remunerated) occupations, total years of job experience, level of economic activity or the fact that men with the same contracted working time as women work in average three hours per week longer than women. Irrespective of the number of hours written in people's contracts, the gender gap in 'hours actually worked per week' has reached as many as 4.2 hours (Czech Statistical Office [CZSO], 2022a, p. 218).[3] Another neglected factor is that men are statistically significantly more likely to work overtime or at night[4] (Czech law mandates a bonus of 'at least 25%' for overtime hours and '10% of average earnings' for night work [Labour Code, Sections 114 and 116]) and to pursue more at-risk or physically/mentally more demanding occupations (CZSO, 2021; European Working Conditions Survey, 2023). Many more items could be added to the list of differences in average productive characteristics between both groups, but doing so would not be purposeful here. The logical inference is that if each group exhibits different productive characteristics that influence their wage levels, then their average wages will also

[1] See, e.g., the Action Plan for Equal Remuneration of Women and Men 2023–2026 adopted by the Czech government in December 2022; the Gender Equality Strategy for 2021–2030 elaborated by the Office of the Government of the Czech Republic in 2021; the project '22% towards Equality' launched by the Ministry of Labour and Social Affairs; and at the European level, the Pay Transparency Directive (European Commission, 2021).

[2] Some authors work with median wages, others convert mean monthly wages into mean hourly wages, etc.

[3] This aspect requires some clarification. Major analyses of the unadjusted (GPG) and adjusted gender pay gap (AGPG) are calculated on the basis of hourly earnings, so it would appear that the difference in hours worked is indeed included in the GPG analysis. As we will show, the reality is much more complicated, and this gap in actual hours worked is not (or only partially) involved. Moreover, the conversion to hourly earnings is not unproblematic, and it introduces quite important biases into the analysis. See below sections "Hours actually worked" and "The Relationship Between Hours Worked and Wages is not a Linear Function".

[4] There are only 70 women for every 100 men who 'sometimes' work nights. The ratio is the same for women and men who 'usually' work at night. (CZSO, 2022a, p. 227, data for 2020)

Table 1. Average gross (unadjusted) gender pay gap in EU countries

Territory, country	2010	2014	2015	2016	2017	2018	2019	2020
EU27	*15.8*	*15.7*	*15.5*	*15.1*	*14.6*	*14.4*	*14.1*	*13.0*
Belgium	10.2	6.6	6.4	6.0	5.8	5.8	5.8	5.3
Bulgaria	13.0	14.2	15.5	14.6	14.3	13.9	14.1	12.7
Czechia	*21.6*	*22.5*	*22.5*	*21.5*	*21.1*	*20.1*	*18.9*	*16.4*
Denmark	17.1	16.0	15.1	15.1	14.8	14.6	14.0	13.9
Estonia	27.7	28.1	26.7	24.8	24.9	21.8	21.7	21.1
Finland	20.3	18.4	17.5	17.5	17.1	16.9	16.6	16.7
France	15.6	15.5	15.6	15.9	16.3	16.7	16.5	15.8
Croatia	5.7	8.7	.	11.6	12.3	11.4	11.5	11.2
Ireland	13.9	13.9	13.9	14.2	14.4	11.3	.	.
Italy	5.3	6.1	5.5	5.3	5.0	5.5	4.7	4.2
Cyprus	16.8	14.2	13.2	12.3	11.2	10.4	10.1	9.0
Lithuania	11.9	13.3	14.2	14.4	15.2	14.0	13.3	13.0
Latvia	15.5	17.3	18.4	19.7	19.8	19.6	21.2	22.3
Luxembourg	8.7	5.4	4.7	3.9	2.6	1.4	1.3	0.7
Hungary	17.6	15.1	14.0	14.0	15.9	14.2	18.2	17.2
Malta	7.2	10.6	10.7	11.6	13.2	13.0	11.6	10.0
Germany	22.3	22.3	21.8	21.1	20.4	20.1	19.2	18.3
Netherlands	17.8	16.2	16.1	15.6	15.1	14.7	14.6	14.2
Poland	4.5	7.7	7.3	7.1	7.0	8.5	8.5	4.5
Portugal	12.8	14.9	16.0	13.9	10.8	8.9	10.6	11.4
Austria	24.0	22.2	21.8	20.8	20.7	20.4	19.9	18.9
Romania	8.8	4.5	5.6	4.8	2.9	2.2	3.3	2.4
Greece	15.0	12.5	.	.	.	10.4	.	.
Slovakia	19.6	19.7	19.7	19.2	20.1	19.8	18.4	15.8
Slovenia	0.9	7.0	8.2	8.1	8.4	9.3	7.9	3.1
Spain	16.2	14.9	14.1	14.8	13.5	11.9	11.9	9.4
Sweden	15.4	13.8	14.0	13.3	12.5	12.1	11.8	11.2

Source: Eurostat as at 8 June 2022, CZSO (2022a). Focus on Women and Men–2022, chapter 4.37, p. 240

Note: Includes only businesses with 10+ employees. Includes economic activities under sections B through S, except for section O of NACE Rev.2.

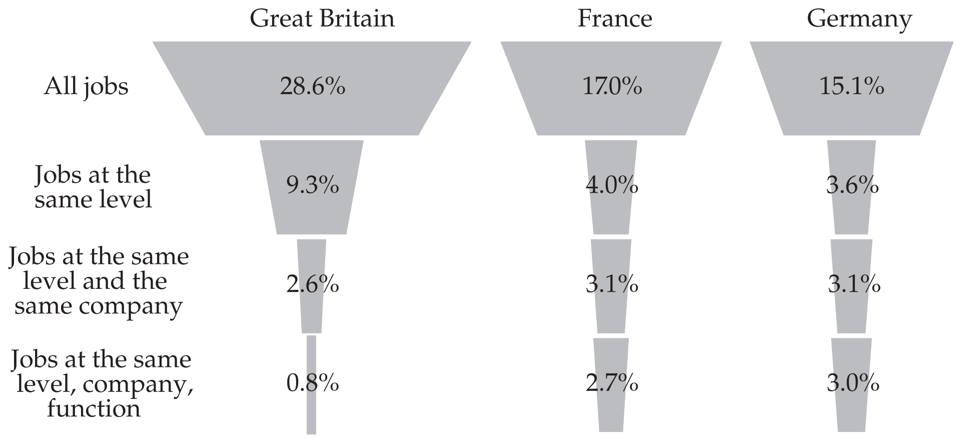
differ. In this vein, it cannot be surprising that men have different average pay than women. Clearly, given the different characteristics of both groups, there was no statistical or logical reason for them to have the same pay. In other words, unadjusted GPG levels do not help us determine whether this difference exists due to pay inequality or not; whether or not women are being discriminated against by receiving less money than men for equal work. Unadjusted GPG could serve merely as a starting point for a more detailed analysis. Table 1 shows the development of the unadjusted gender pay gap (i.e. the difference between men's and women's average earnings in the Czech Republic and other countries). The Czech GPG, based on Eurostat, decreased from 20.1% in 2018 to 15.2% in 2021 (CZSO, 2022a, p. 240).[5] However, if we want to determine the extent of the gender pay gap for the same work (i. e. AGPG), we need to compare men and women with the same productive characteristics. In other words, we need to eliminate the effect of women and men having, on average, different productive characteristics. The pay gap between women and men with the same productive characteristics is expressed by the adjusted gender pay gap category (AGPG). In contrast to the GPG, the adjusted gender pay gap (AGPG) serves as an approximate indicator of unequal pay for equal work and should be strictly distinguished between both categories.

Adjusted gender pay gap

Most empirical studies on gender wage discrimination calculate the AGPG using a formal statistical technique designed by Oaxaca (1973) based on Becker's (1957) theory of labour market discrimination. Given that Blinder (1973) designed a similar method, the technique has been named the 'Blinder–Oaxaca decomposition'. This approach defines discrimination as the difference between an observed gender pay gap and one that would exist if women and men were remunerated on the basis of the same 'productive characteristics' (e.g. qualifications, job experience, hours actually worked, etc.) and exhibited equal levels of those characteristics. In contrast, wage differences based on the 'non-productive characteristics' of workers, including their gender (but also sympathy, etc.), are viewed as discriminatory (Grimshaw & Rubery, 2002). The practical uses of the Blinder–Oaxaca approach

[5] Eurostat calculates the unadjusted GPG by including occupational sectors from 'NACE sections B to S, without O' (European Commission, Eurostat, 2018, p. 6). Therefore, the excluded sections are agriculture, forestry and fishing (A); public administration, defence and compulsory social security (O); activities of households as employers (T); and activities of extraterritorial organisations and bodies (U). It is reasonable to assume that, in particular, the decision to exclude the sector of agriculture, which is dominated by men with low wages, and the exclusion of sector O, where the GPG is significantly lower than in the other sectors, increases even the unadjusted GPG, from which the adjusted GPG is subsequently calculated.

Figure 1. International comparison of the unadjusted and adjusted gender pay gap, pay gap between women and men, 2016, percentage of men's wages (full-time pay)

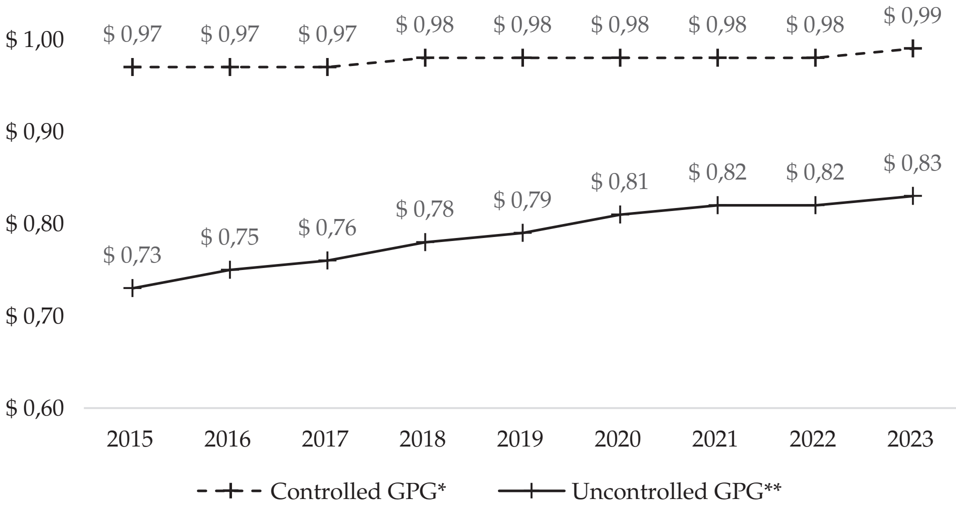


Source: Korn Ferry database, cited from The Economist (2017, 1 August).

include placing decomposition into individual productive factors that each explain a portion of the unadjusted GPG. To simplify, the sum of such productive factors equals the 'explained' part of the GPG, and the portion that cannot be explained by the different productive characteristics of men and women is referred to as the 'adjusted' or 'unexplained' gender pay gap. This adjusted gap (or unexplained part of the GPG) is interpreted as a 'gap in pay for equal work and work of equal value' (Ministry of Labour and Social Affairs, 2022, p. 7) or similarly as 'approximation of potential discrimination, i.e. how the labour market rewards men and women in an unequal way for the same work.' (European Commission, Eurostat, 2018, p. 17; Weichselbaumer & Winter-Ebmer, 2005).

AGPG values, as an approximation of the unequal remuneration of men and women for equal work, are shown in Figure 1–2 and Table 2. Based on highly representative data from the Korn Ferry database of the wages of 12.3 million workers at 14,284 companies in 53 countries, the journal stated that 'when all job differences are accounted for, the [gender] pay gap almost disappears' (*The Economist*, 2017). For the Czech Republic, it indicates an even higher difference between men's and women's average wages (i.e. *unadjusted* GPG) than Eurostat or the CZSO. However, a comparison of workers in the same position, occupation and company indicates an AGPG of as little as 3.8% in the Czech Republic (Korn Ferry, 2016, p. 7). *The Economist* (2017) also presents comparable results for other countries (see Figure 1) as well as Chamberlain, Zhao, Stansell (2019) (see Table 2). The Payscale (2023) database (Figure 2) indicates a 1–2% gender pay gap

Figure 2. Unadjusted and adjusted gender pay gap in the USA, 2015–2023



Source: Payscale 2023

Note: * Uncontrolled Gender Pay Gap (Opportunity Pay Gap): Measures median salary for all men and all women regardless of job type, seniority, location, industry, years of experience etc.

** Controlled Gender Pay Gap (Equal Pay for Equal Work): Measures pay for men and women with the same job and qualifications

for ‘equal work’ in the US. (Notably, according to the Organisation for Economic Co-operation and Development, 2023, the unadjusted GPG in the US is higher than in the Czech Republic.) In general, the AGPG—in contrast to the GPG—is in the range of lower units of percentage according to these data (see Figure 1–2 and Table 2).

This is a much lower gap than that indicated by Křížková and Pospíšilová (2023), among others. They argued that the ‘equal remuneration of women and men for equal work is not safeguarded in the Czech Republic. If men and women work in the same positions,[6] women earn 9% lower hourly wages, on average’ (Křížková & Pospíšilová, 2023, p. 54). Křížková et al. (2018) stated the following regarding the Czech Republic:

[6] ‘Same position’ is defined by the authors as a combination of the same workplace and the same occupational category according to the four-digit ISCO classification (International Standard Classification of Occupations). In their analysis, the term ‘same position’ does not mean the same position within the company’s hierarchy. Neither does it say anything about the worker’s level of responsibility, decision-making autonomy, length of job experience, being a subordinate/manager, etc.

Table 2. The unadjusted and adjusted gender pay gap by country

	Unadjusted base gender pay gap		Adjusted base gender pay gap	
	Average cents/ pence earned by women per dollar/pound/ euro of male earnings	Percentage male pay advantage	Average cents/ pence earned by women per dollar/pound/ euro of male earnings	Percentage male pay advantage
Australia	0.85	15.1%	0.97	3.1%
France	0.88	11.6%	0.96	3.7%
Canada	0.84	16.1%	0.96	4.0%
United States	0.79	21.4%	0.95	4.9%
United Kingdom	0.82	17.9%	0.95	5.0%
Singapore	0.87	12.8%	0.95	5.2%
Germany	0.78	22.3%	0.94	6.4%
Netherlands	0.81	18.9%	0.93	6.6%

Source: Glassdoor Economic Research, cited from Chamberlain et al. (2019, p. 4)

> [I]n 2016, the gender pay gap was around 26%, of which 11 percentage points (43% of the total GPG) were explained by the factors included in the model, while the remaining 15 percentage points (57% of the total GPG) remained unexplained. (p. 95)

Similarly, in another study, Křížková et al. (2020, p. 25) claimed that the unexplained part of the GPG was 14%. Furthermore, Eurostat indicated even higher levels of the AGPG (i.e. the unexplained part of the GPG)—as high as 17% (Leythienne & Pérez-Julian, 2021, p. 20).

The conclusions of the studies mentioned above were subsequently incorporated into government documents (Office of the Government of the Czech Republic, 2021; Ministry of Labour and Social Affairs, 2022).[7]

The Eurostat data provide the key practical basis for EU and individual member states' policies. Eurostat elaborated on its own methodology for determining the AGPG, published in the European Commission document named 'Adjusted Gender Pay Gap' (European Commission, Eurostat, 2018). In the following section, we critique Eurostat's methodology and its approach to the decomposition of the factors influencing the AGPG. However, as our critique im-

[7] For example, the Czech government's website states, 'We know that women receive lower pay than men for equal work with the same employer' (Government of the Czech Republic, 2022), referring directly to Křížková et al. (2018).

plicitly affects other approaches to AGPG measurement, we also review the work of Křížková and Pospíšilová (2023) (who, however, do not use the Blinder-Oaxaca decomposition in their study, but use a different methodological approach) and Křížková et al. (2018).

Eurostat's approach to calculating the adjusted gender pay gap and the decomposition of factors

As stated above, Eurostat's methodology standardly divides the unadjusted (gross) gender pay gap into an explained part (explained by the different characteristics of both groups) and an unexplained part, i.e. AGPG (European Commission, Eurostat, 2018, p. 17). As shown in Table A1 in online Appendix[8], the explained part comprises as few as three percentage points for the Czech Republic (Gender Pay Gap Statistics—Analytical Tables, Table 2, 2022; in this paper Table A1). When reviewing the weights attached to the individual factors included in the explained 3 percentage points, we found low values representing different characteristics of both groups *prima facie* implausible. 'Economic activity' is the only factor in excess of 1 percentage point: According to Eurostat, it explains 3.2 percentage points of the unadjusted gender pay gap in the Czech Republic (Gender Pay Gap Statistic—Analytical Tables, Table 2, 2022; in this paper Table A1). Importantly, the CZSO (2022b) data indicate that men's level of economic activity (83.8%) exceeds that of women by 13.8 percentage points.[9] The gap is not due to women's lower diligence; instead, it can be almost fully explained by the fact that it is mostly women who take care of disabled, sick or otherwise unable family members; women are more likely than men to study at university; personal health problems are a third important reason for inactivity or low economic activity. Other causes are relatively small (CZSO, 2022a, p. 196).

As shown in Table A3, men's economic activity reaches high levels as early as the 25–29 age category (93.9%, 2021 data) and continues to increase with age, being consistently above 90% until the age of 60. As for women, only the 45–55 age group exhibits economic activity levels comparable to men's (above 90%). Outside that decade, women exhibit much lower levels of economic activity. This significantly impacts the different remunerations of the two groups, and that impact is certainly greater (see next two sections) than the 3.2 percentage points reported by Eurostat (see Table A1).

[8] Available at https://doi.org/10.13060/csr.2024.032.

[9] The higher economic activity rate of men is also found in comparison to childless women, even in the EU average. 'In the EU in 2016, the employment rate for women without children was 65%, while it was 73% for men' (Czech Statistical Office, Eurostat, 2017a). Thus, the higher economic activity of men cannot be explained solely by the fact that women generally take more care of the family.

Finally, all other factors considered by Eurostat are in the range of several tenths of a percentage point, and when collectively included in the model, they decreased the explained part of the GPG by two-tenths of a percentage point (from 3.2 to 3). Next, we take a closer look at each of these factors.

The age factor as a spurious variable causing a biased adjusted gender pay gap

Age is the first factor on Eurostat's list. We find this category problematic for several reasons. The likely rationale behind its inclusion is that as wages grow with age, there is a relationship between age and wage levels. According to CZSO (2022c, A3) data, this is the case (until a certain age is reached): A 30-year-old earns 37% more than a 20-year-old, and a 40-year-old earns 43.9% more than a 20-year-old. However, this is not caused by greater *age* but by having greater *job experience*. Age represents a proxy variable here, and job experience (not age) is the actual determinant of wage level.[10]

However, Eurostat, Křížková and Pospíšilová (2023) and Křížková et al. (2018) have compared the *age* of both groups instead of their years of job experience. This substitution leads to highly biased AGPG levels. It has been shown that men exhibit substantially longer working lives (39.2 years) than women (32.7 years), with a mean gender gap of 6.5 years (CZSO, 2022a, p. 220). This gap is due to several factors, one of which is that women obtain their first jobs later than men and retire earlier.[11] This job experience gap is further increased by the fact that women predominantly take parental leave (at a mean age of 28.2 years in the EU). Note that in Czech public-sector jobs, parental leave is, by law, counted towards an adjusted length of job experience. This causes a further bias in the statistic: Men and women with equal *adjusted* years of experience have different amounts (i.e. years) of *actual* job experience. Thus, when comparing men and women *with equal adjusted (or formal) length of job experience*, a gap caused by different lengths of *actual job experience* may appear as discrimination against women. At the same time, the inclusion of parental leave in adjusted years of job experience is likely the main reason why the GPG in the public sector is lower than in

[10] According to the CZSO (2022c, A3), an average worker earns 32,279 CZK at the age of 20–25 years, 44,278 CZK 10 years later and 46,462 CZK another 10 years later (at the age of 40–44 years). Thus, the average worker's wage grows at an average pace of 2.2% for each year worked in the first 20 years. This later slows down, and workers older than 50 years of age even record negative growth. Despite the slowdown that occurs later in life, on average, each year of experience has a significant impact on wages.

[11] In the EU as a whole, men begin their first employment at a mean age of 22 years, one year earlier than women, and women retire 1.1 years earlier, on average (Czech Statistical Office, Eurostat, 2017b, Chapter 1.1). Thus, at the very time the average EU woman enters the labour market, the average man already has one year of job experience.

the private sector, where organisations have a stronger tendency to pay workers on the basis of *actual* length of job experience and their resulting qualifications and competences (rather than only *formal* or *adjusted*).

The above-mentioned level of economic activity is probably the best measurement of the total length of job experience of women and men as separate groups. This indicates the extent to which women and men obtain job experience by working. For example, a 50% level of economic activity among men in a given year means that the length of job experience for men as a whole has increased by half a year. Table A2 illustrates the difference in years of job experience between men and women in the Czech Republic throughout their careers. While the gender gap in the length of job experience varies from one year of a career to another, the mean values for the entire productive life based on CZSO (2022a, p. 194) data indicate that *the average Czech woman has four fewer years of job experience than the average man*. Given a wage increment of 1.1% for each additional year of job experience (CZSO, 2017),[12] this factor is one of the *key* differences in the productive characteristics of women and men, explaining approximately 4.4 percentage points of the total GPG. From this perspective, we find it striking the existing Czech analyses of the AGPG (Křížková & Pospíšilová, 2023; Křížková et al., 2018) have not included the factor of total years of job experience, which has likely caused an overestimation of the AGPG. Furthermore, official government documents were directly based on overestimated AGPG results mentioned above. This is the case with the *Gender Equality Strategy for 2021–2030* (Office of the Government, 2021) and the *Action Plan for Equal Remuneration of Women and Men 2023–2026* (Ministry of Labour and Social Affairs, 2022).

According to O'Neill (2003), narrowing the gender gap in years of job experience was 'the key factor underlying the decline in gender earnings disparities between 1979–2001' in the US. Claudia Goldin (2023) came to the same conclusion when she examined the evolution of wages for male and female lawyers in the US. The reason why women lawyers earn less, on average, she argued, is not the 'discriminatory practice of promotion and mentorship' or 'gender bias'[13] but rather the fact that 15 years after graduation, 'men accumulated more legal expe-

[12] Overall (from ages 19–64 years), wages grow, on average, 0.8% for each year worked (Czech Statistical Office, 2017, own calculations). When looking directly at the relationship between wages and length of employment, there is an annual wage growth of 1.4% during the first 30 years and 1% during the first 40 years (CZSO, 2022c). There are several ways of estimating the effect of each additional year of job experience on wage level. However, available data indicate an average increment of 0.8%–1.4% per year of job experience, with 1.1% being the central value. This data will be used in the next sections of this study. Pytlíková (2015, pp. 2, 19) reached a similar result based on EU–SILC (European Union – Statistics on Income and Living Conditions) data for 2012, indicating a wage decrease of 1.1% for each year on parental leave (which means one less year of job experience).

[13] 'Bias exists in many law firms, but it is not the primary cause of gender differences in promotion and earnings' (Goldin, 2023, p. 182).

rience' (because men work, on average, more years and more hours per month than women). She concluded that the 'underlying cause of the gap is not their genders' (Goldin, 2023, p. 180). Rather, job experience and real working hours play decisive roles.

With this in mind, let us return to the problem of the AGPG calculation method. Arguably, if a factor (e.g. total length of job experience) is not included, the analysis implicitly assumes *equal* levels of the parameter for both groups under comparison. Since this assumption regarding the total length of experience of men and women does not correspond to the facts, analyses that neglect this factor necessarily lead to a highly biased and overestimated AGPG.

However, even more biased results are obtained when total years of job experience are *substituted* with the category of age, as in the case of Křížková and Pospíšilová (2023) and Křížková et al. (2018), among others; substituted with 'job experience in current enterprise'; or even both (as in Eurostat's methodology). The next section demonstrates why this is so.

Including age factor causes a spurious increase in the adjusted gender pay gap

By neglecting the factor of total years of job experience, analyses artificially increase the unexplained part of the gender pay gap (i.e. unequal pay for equal work). A significant factor that explains a large part of the GPG is simply ignored. Nevertheless, including age instead of total experience causes a *further increase* in the AGPG and its deviation from reality. According to Křížková and Pospíšilová (2023), age can be used as an approximation for or *in lieu* of the total years of job experience category. The authors explicitly stated in their analysis that 'age partially, though of course not perfectly, *substitutes* the variable of length of experience, which it is not appropriate to include in the model *simultaneously* due to its *high correlation* with age' (Křížková & Pospíšilová, 2023, p. 55, emphasis is ours).

We agree with Křížková and Pospíšilová (2023) that instead of including *both* age and length of job experience in the model, only one of the categories should be included. Even though age is somewhat correlated with length of job experience, the problem is that the correlation is much stronger among men than women, and both groups are clearly distinguished *precisely* by this correlation gap. Women exhibit a weaker correlation due to their interrupted careers, which means that age does not automatically translate into length of job experience. At the same time, this difference is one of the main reasons behind the pay gap. However, this difference is not only *nullified* by including only the age category in the statistical model (and neglecting the length of job experience category); the involvement of the age category in the statistical analyses even *increases* the unexplained part of the GPG. The age factor represents a negative number, so it does not contribute at all to the explication of the GPG but rather reduces its explained part. Therefore, understandably, the ensuing analysis indicates a much higher level of wage discrimination (AGPG) than exists in reality.[14] Let us examine this more closely.

Including the age category in the statistical model practically means comparing the age of *working* women with that of *working* men while implicitly assuming that wages grow with age. Because young women taking maternity and parental leave do not receive wages, they are excluded from the statistical comparison of working women and men. This is once again apparent in the data on women's economic activity. There are ~20% fewer economically active women than men aged 20–24 years, and almost 30% fewer in the 25–29 age category; therefore, it comes as little surprise (given the inactive women's exclusion from wage statistics) that working women are, on average, *older* than working men. The CZSO (2022a, p. 194) data indicate a mean age of 43.6 years for working women and 42.3 years for men in the year 2021,[15] a gender gap of 1.3 years.[16] For the same reason, Eurostat (see Table A1) found a negative contribution of age to explaining the unadjusted gender pay gap.[17] In other words, as mentioned above, the age factor not only fails to explain any part of the GPG but even increases its unexplained part (by reducing the explained part). In contrast, total length of job experience is a highly significant explanatory factor for a substantial part of the GPG. Given the gender gap of ~4 years in total length of job experience and the

[14] The reason why Křížková and Pospíšilová (2023) – and other authors, including Eurostat – do not include length of job experience in their statistical models is most likely that the available data on employees lack this information. However, this does not change the fact that the absence of the length of job experience variable still makes the claims about a high AGPG very problematic. If this variable is not included (or even approximated by the age variable), the statistical model must logically always show a high AGPG (considering the fact that the difference in length of job experience between men and women is, in reality, one of the biggest differences between their productive characteristics). Therefore, it is nearly impossible to reach a result other than a high AGPG. If we encounter an absence of data on the length of job experience, it may be more appropriate to approximate this variable on the basis of the economic activity rates of the two groups throughout their careers rather than simply substituting it with age. Data on the economic activity rate will illustrate the disparity in overall job experience between genders at each life stage and the average difference in job experience length between the two groups (see Table A2).

[15] According to the CZSO (2022c), the mean age of workers was 43.4 years for men and 44.8 for women in 2022.

[16] The data indicate that working women are, on average, older than working men, despite the fact that men retire later and, even in retirement, it is mostly men who continue to work. When comparing the age of working women and men in the category up to 55 years old, there will be more significant age differences, where women will be older than men.

[17] See also Křížková et al. 2018. In addition to age as a factor that contributes negatively to the explained part of the GPG, the authors also mention 'job experience' (with 'job experience' meaning 'job experience in the current firm', p. 74). Job experience in current firm reduces the value of the explained part of the GPG by 0.52 percentage points, and education reduces the explained part by 0.67 percentage points. Křížková et al. 2018, p. 77.

1.1% increment in wage levels for each additional year of job experience, this factor *explains* ~4.4 percentage points of the GPG.

The age category, which is supposed to 'substitute' total job experience, although 'not perfectly', according to Křížková and Pospíšilová (2023), not only falls short of explaining any part of the GPG but even, as we said, increases the amount of the adjusted GPG by ~4.4 percentage points. Thus, including age instead of total job experience is not only far beyond 'not perfect' but it causes a substantial bias and spurious increase in the unexplained part of the GPG.

Another notable fact in the context of the age category is that data indicate that, for the reasons described above, working men are, on average, younger than working women (by 1.3 years). On the other hand, if we compare only employees in managerial positions, it is the women who are younger (in the private sector in the Czech Republic for the years 2020 and 2021, they are younger by about 7 months).[18] Thus, it appears that women reach managerial positions[19] at a younger age, despite having on average less total experience and working fewer hours per month than men. This seems to challenge the popular notion of the glass ceiling. As Farrell (2005) notes, despite the fact that women are less represented in managerial positions, they actually reach managerial positions after a shorter period of experience than men, who need a longer period of work experience to reach a managerial position.

We should add that male managers can be expected to retire later than female managers, which increases the average age of male managers. This could hypothetically partly explain why male managers are on average as a group older than female managers. However, even taking this into account, women still appear to reach managerial positions earlier than men. For the United States, Farrell (2005, p. 86) gives the following figures:

> Prior to age 40, women are 15 times more likely than their male counterparts to become top executives at major corporations. (Of top female executives at major companies, 21.4% are under 40, while only 1.4% of the male executives are under 40.) In a study of the top five executives at almost 3,000 of the country's largest firms, the women's average age was 48; the men's, 53.

As we said male managers can be expected to retire later than female managers, which increases the mean age of male managers. Nevertheless, women seem to reach managerial positions earlier than men on average.

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[18] We would like to thank Petr Soukup for providing us with this data from the Trexima company.

[19] Concerning the GPG among top managers, Bugeja, Matolcsy, and Spiropoulos (2012) analysed the question of the GPG among top managers (CEOs). Their results 'indicate that there is no difference in total pay, salary or bonus for female CEOs.' (p. 859)

Job experience in current enterprise

The category of 'job experience in current enterprise' plays a similar role as age in artificially increasing the AGPG. Table A1 shows that in Eurostat's analysis, 0.2 percentage points of the total GPG is explained by 'job experience'. At first sight, this suggests that Eurostat included the total length of job experience in its analysis. However, in a note attached to the above-mentioned publication on the Adjusted Gender Pay Gap (European Commission, Eurostat, 2018), it is explained that instead of total length of job experience, 'job experience' stands for 'job experience in current enterprise'. This causes additional bias in the resulting AGPG.

The fact that men change jobs more frequently than women decreases their level of 'job experience in current enterprise',[20] while the fact that men have more total years of job experience travels in the opposite direction.[21] The resulting adjusted gender gap in 'job experience in current enterprise' amounts to 0.2% in favour of men. By only including the length of 'job experience in current enterprise' in its statistical model (instead of total job experience), Eurostat effectively assumes that a worker's job experience is nullified by each change of employer. We find this assumption absurd: When a person gains experience in one job, they seek better opportunities. When changing employers, they do not lose their job experience; they 'take it with them'. The employee's prior work experience is considered in determining his or her salary. Moreover, from a career path perspective, the highest wage increases are achieved by changing employers. Employees typically accept a new job when they are offered higher earnings than in their existing job. Thus, although men have, on average, four more total years of job experience,[22] the variable used by Eurostat to account for this fact exhibits almost the same levels for both genders (by only accounting for 'job experience in current enterprise'). In Eurostat's analysis for 2014, 'job experience in current enterprise' contributed zero to the explained part of the GPG (European Commission, Eurostat, 2018, p. 12).[23]

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[20] See, for example, US Bureau of Labor Statistics, US Department of Labor (2023, pp. 2–3).

[21] If men changed jobs as frequently as women, their length of job experience in the same firm would logically be higher (given that their overall length of work experience is, on average, higher). However, when hypothetically assuming an equal job fluctuation rate between men and women, the total length of job experience increases men's length of experience in one firm (compared to women). Conversely, changing jobs more frequently decreases their length of experience in one firm. These two factors go against each other and tend to nullify their contribution to 'job experience in current enterprise'. Thus, the contribution to the explained part of the unadjusted gender pay gap by the category of 'length of experience in current enterprise' is, according to Eurostat, either exactly zero (see European Commission, Eurostat, 2018, p. 12) or very close to zero (0.2 percentage points in 2018) (see Table A1).

[22] According to the CZSO (2022a, p. 220) data for 2021, men's working lives are as many as 6.5 years longer than women's — a 16.6% difference.

[23] Křížková and Pospíšilová (2023) also mentioned the weak effect of the 'job experience

Occupational choice

Another factor significantly influencing the mean unadjusted gender pay gap is men's and women's occupational choices—they are not equally represented in various industries, and those industries exhibit different levels of mean wages.[24]

In contrast, different gender proportions in different industries have a remarkably strong effect on GPG. For example, the GPG was reduced by ~50% when men and women working in the same occupational category were compared (Křížková & Pospíšilová 2023, p. 45). In other words, the fact that women opt for different occupations than men explains about half of the unadjusted GPG.[25]

Notably, even Eurostat indicates a negative number for this factor ('occupation'; see Table A1). According to Eurostat, different occupational choices do not explain any part of the GPG and instead increase the unexplained part by 0.3 percentage points. This would mean that women are more often working in better-paid occupations than men—a result that is far from reality. On the contrary, women are more likely to be employed in lower-paid sectors. Eurostat's analysis indicates such a weak effect of occupational choice because of the method used to differentiate occupations. Eurostat uses the two-digit ISCO-08 and refrains from differentiating occupations within each two-digit category, thus treating them as *equal* or *same* for statistical purposes. However, the two-digit classification is *highly coarse-grained*. At the same time, the finer-grained a classification, the larger part of the GPG that is explained by the occupational choice category (and thus the higher the explained part of the GPG as such).

An examination of the various two-digit categories reveals the kinds of occupations that are, sometimes surprisingly, treated as equal. In other words, very different occupations belong to the same two-digit occupational category and are treated as the same in terms of statistical analysis. For instance, sub-major Group 12 contains finance managers, personnel managers and cleaning services managers. In Group 13, there is a mix of occupations, such as fishing vessel skippers (coastal waters), mine managers, warehouse managers, internet service providers, childcare centre managers, directors of nursing, housing services managers, deans (university), head teachers, finance managers, archive manag-

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in current enterprise' category. For this reason, they did not include it in their analysis: 'Neither was the length of job experience in current enterprise included because a previous analysis of GPG decomposition indicated that the variable explains just under 1% of GPG only [Křížková et al., 2018]' (Křížková and Pospíšilová, 2023, p. 55).

[24] Another issue is that men are more likely to hold managerial positions than women. Yet, this fact has a negligible effect on the total GPG. According to Křížková et al. (2018, p. 77), women's weaker representation in management positions explains as little as 0.62 percentage points of the GPG.

[25] Goldin (2023, p. 4) reported that, for the US, different occupational choices of men and women explain about one-third of the GPG.

ers, library managers and prison governors. And in Group 26 are lawyers, chief justices, judges and notaries mixed together with archivists, art gallery curators, librarians, economic analysts, philosophers, priests and poets (International Labour Office, 2012, pp. 92–104, 158–168). The consequences of this method are obvious. By considering jobs in each group as the *same* from a statistical analysis perspective and by mixing well-paid and poorly paid or qualified and unqualified occupations in the same category, the analysis nullifies any differences between the occupations predominantly chosen by men and women. Of course, this resulted in zero explanatory power for the GPG. Additionally, by valuing this factor as a negative value, Eurostat (once again) increased the unexplained part of the GPG.

Completely different results were obtained when finer-grained differentiation using the four-digit ISCO was used. This commendable choice was made by Křížková and Pospíšilová (2023, pp. 45–46), who concluded that based on this finer-grained classification, the occupational choice factor can explain about half of the GPG.

Nevertheless, even this differentiation is merely an approximation and obviously does not provide a complete picture of reality. With this differentiation, we arrive at the level of pay gaps between female and male doctors at a given hospital, female and male teachers at a given school, female and male cooks at a given restaurant, female and male lawyers at a given law firm, etc.[26] As shown by Farrell (2005) with US data, a large part of a GPG found within the same occupation at the same workplace can be explained by different *specialisations* and other factors, such as length of job experience, hours actually worked, level of risk involved (even at the same position), responsibility level (even at the same position), standby duties vs. the ability to mentally detach from work, willingness to work longer hours or return to work in case of unforeseen situations[27] (again, all at the same position[28]), amount of further education required for one's specialisation, etc. Including these factors would further reduce the level of the AGPG.

Farrell (2005) illustrated this with a 20% gender pay gap in the US medical profession. The gap was reduced to 2% when doctors with the *same speciality* were compared (Farrell, 2005, p. 75). Teachers in the US provided a similar example. Their annual earnings were 46,000 USD among men and 42,000 USD among women—a sizeable GPG. However, a closer look revealed that male teachers worked, on average, two hours a week more than female teachers, had a 25% larger share of workers with 20+ years of job experience and exhibited, on average, a 10% longer job experience with their current employer (Farrell, 2005, p. 77).

[26] The authors defined men and women working at the same positions as 'working at the same workplace and in the same four-digit category of the CZ-ISCO occupational classification' (Křížková & Pospíšilová, 2023, p. 41).
[27] See Bolotnyy and Emanuel (2022).
[28] In line with Křížková and Pospíšilová (2023), 'same position' is understood here as a combination of the same workplace and the same occupational category.

The relationship between hours worked and wages is not a linear function

Converting total wages to hourly earnings can introduce significant bias in the AGPG calculation. This is due to the non-linear relationship between working time and wages, especially in skilled occupations, which have a higher gap than unskilled jobs. Goldin (2023) provided an example of this phenomenon in the legal profession:

> The average lawyer who works sixty hours a week earns more than two and a half times what the lawyer working thirty hours a week earns. That jump in earnings with time occurs without regard to gender.[29] Both male and female lawyers earn significantly more per hour when their overall hours increase (…) When a lawyer's hours increase from thirty to sixty per week, the average hourly rate increases by almost a quarter. The more hours per week that lawyers work, the more *each of their hours* spent working is worth. If we hold hours worked constant for men and women, there is no gender component to the discrepancy. We know the difference between what the genders earn is significant. But … the underlying cause of the gap is not their gender (Goldin 2023, pp. 180–181).

Farrell (2005, pp. 78–79) demonstrated this principle of non-linear dependence using data from the US Bureau of Labour Statistics (2003), but he extended it beyond the legal profession. The data show that an individual working 45 hours per week earns 44% more than someone working 40 hours per week (Farrell, 2005, p. 78). This means that a 13% increase in working hours results in a 44% increase in total wages. Importantly, this pay difference is not based on gender but rather on the number of hours worked. However, according to the US Bureau of Labor Statistics (2003), men in the US actually work an average of 45 hours per week, while women work an average of 42 hours per week. What does that three-hour difference amount to in pay? 'The average person [regardless of gender] who works 45 hours per week earns 14% more than the 42-hour per week worker' (Farrell, 2005, pp. 78–79). Thus, there was an approximately 6% difference in hourly wages between the two groups.

It is likely that even in the Czech Republic the relationship between wage level and hours worked is not linear. The fact that men work longer per week (36.5 hours versus 32.3 hours; for the same hours [i.e. full-time job], the difference is 37.2 hours versus 34.2 hours [CZSO, 2022a, Chapters 4.7, p. 218]), means that not only are their total wages higher, but their *hourly* wages are also significantly higher. However, this is not because they are men but because they work longer hours per week and the relationship between real hours and wages is non-linearly increasing. Since the adjusted pay gap analyses, which compare the

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[29] For example, if one lawyer earns $1,000 for 30 hours and another lawyer earns $2,500 (not $2,000) for 60 hours, there is a 20% difference in their hourly earnings without any form of discrimination (gender or otherwise). Note: PK.

hourly earnings of men and women, proceed by simply dividing total wages by the number of hours worked, an implicit linear relationship between wages and hours worked is assumed. The logical result of this operation (the assumption of a linear relationship between wages and real hours worked) is that it follows that an hour of male work is worth more than an hour of female work, and this is then misinterpreted as wage discrimination against women, even though, to recall Goldin's (2023) statement, 'there is *no gender component to the discrepancy*' and the 'jump in earnings with time occurs *without regard to gender*' (Goldin 2023, pp. 180–181, emphasis ours).

Farrell (2005) concluded that 'it is possible that up to 70% of the [unadjusted] pay gap between men and women could be accounted for by differences in hours worked' (p. 79).

It would, of course, have to be rigorously calculated (on the basis of Czech data) exactly how much the erroneous implicit assumption of a linear relationship between the length of working time and the wage rate contributes to the overall virtual increase in the AGPG in the Czech Republic. However, given the US data, it is reasonable to assume that it is certainly not a small part.

A further bias in the AGPG level arises from the fact that the conversion to hourly wages is effectively made on the basis of the number of hours stated in the employment contract rather than the *actual* hours worked. This issue will be addressed in the next subsection.

Hours actually worked

Differences in the monthly number of hours worked are of utmost importance in explaining why women and men have different average earnings. At the same time, its erroneous use in the calculation of the AGPG causes the highest amounts of bias. Let us briefly look back at the category of *unadjusted* GPG (the difference between the average pay of women and men). While the CZSO quantified a gap in *monthly* pay, Eurostat quantified the gender gap in 'hourly earnings'. However, both result in almost the same amount of unadjusted GPG. The difference in monthly wages between men and women and in hourly wages is almost the same in both cases. More specifically, in its publication *Focus on Women and Men–2022*, Chapter 4.34, the CZSO (2022a, p. 240) defined the GPG as the difference in 'average gross monthly wages' and indicated values of 19.1% for 2019, 16.2% for 2020 and 15.2% for 2021. For the same years, Eurostat (2023) indicated practically equal numbers 'in *hourly* earnings': 19.2% for 2019, 16.4% for 2020 and 15.0% for 2021. One key fact emerges from these data: The difference between the monthly and hourly wage gaps was virtually non-existent (in the lower deciles of percentage points) in these analyses. In other words, the difference in that men work 8.6% longer than women on the same contract (CZSO, 2022a, Chapter 4.7, p. 218) was not involved in the calculation of the AGPG by Eurostat (see Table A1), Křížková and Pospíšilová (2023), the Office of the Government (2021) and the Ministry of Labour and Social Affairs (2022).[30] This may not be immediately apparent, as these analyses rely on the number of hours worked by men and women and then convert them into hourly earnings. However, the reality is more complicated than this suggests. The data set is also affected by a significant limitation, which presents a significant obstacle to achieving a clear conclusion. Let us now provide a more detailed explanation.

Based on the Labour Force Sample Survey (LFSS), the CZSO (2022a, Chapter 4.7, p. 218) stated that men with the same contracted working time (full-time) work 3.2 hours per week more than women. A similar difference (~three hours a week) was also shown in the SILC data.[31] Thus, LFSS showed that women working full-time hours work 8.6% fewer hours than men working full-time hours (CZSO, 2022a, Chapter 4.7, p. 218).

Given an unadjusted gender gap in *monthly* pay of 15.2% in 2021 (CZSO, 2022a, p. 218) and the fact that women work 8.6% fewer hours than men with equal contracted working time, it is impossible to obtain an *hourly* GPG only 0.2% below the monthly gap (or to obtain even higher hourly GPGs for the previous years). Let us explain this discrepancy. Eurostat, Křížková and Pospíšilová (2023), the Ministry of Labour and Social Affairs (2022) and the Office of the Government (2021) all used data sourced from *employers*, whereas few Czech workplaces currently use punch clocks or maintain any records of hours actually worked. Therefore, Czech employers account for the hours worked by their employees based on the hours written in their *employment contracts* and not their *hours actually worked*. Therefore, their regular reports are based on the hours written in employment contracts and *recorded* leave from work or overtime hours (but again, most employers do not keep overtime records). This logically results in entirely negligible differences in hours worked by men and women with the same contracted working time: merely 1.3 hours per *month* or 0.75% (according to employer reports, Czech men worked 173.6 hours and women worked 172.3 hours per month in 2022; CZSO, 2022c). For the above reasons, these reports neglected the fact that men systematically work longer hours than women with the same contracted working time. Consequently, this statistic is highly biased in a situation when, according to the European Working Conditions Survey (2023), ~58% of men and ~40% of women are obliged to work in their free time. The following answers were obtained to the question, 'How often have you worked in your free time to meet work demands?': daily, 2% of men and 1% of women; several times a week, 7% of men and 4% of women; several times a month, 20% of men and 10%

[30] More precisely, these analyses used data on working hours, which show that the difference between men's and women's actual working hours is negligible. According to data used by analyses mentioned above, the difference is only 1.3 hours per month. According to the CZSOs, however, this difference is significantly higher. See below for more details.

[31] According to the EU–SILC 2013, Czech men worked, on average, 43.3 hours per week and women worked, on average, 39.9 hours per week.

of women; less often, 29% of men and 25% of women; and never, 42% of men and 60% of women (European Working Conditions Survey, 2023).

When the calculation of *hourly* AGPG is not based on hours *actually* worked but rather on employment contracts (i.e. monthly pay is practically divided by contracted working hours), the resulting AGPG will, of course, be significantly higher than what exists in reality.

The obvious questions remain about how men and women 'really' differ in their total hours worked and whether the calculation of the AGPG should be based on employer data (Trexima and SILC) or employee data (LFSS). There are problems with both approaches. Although subjective bias may affect employee data, since most employers do not keep records of hours actually worked, employee reports (CZSO, 2022a, p. 218) seem to provide a better picture of the hours actually worked. According to CZSO (2022a, p. 218), as stated above, women working full-time hours work 8.6% fewer hours than men with the same contracted working time (their actual working time is only 91.4% of men's).

Therefore, in order to answer the question of the level of the wage gap between men and women for the same work, it would be necessary to adjust the figure of 9% reported by Křížková and Pospíšilová (2023), supposedly for 'the same work', by firstly including the factor of the length of the difference in accumulated job experience between men and women; the difference in actual working hours between men and women with the same contractual working hours (i.e. using data that actually included this difference); and taking into account the non-linear relationship between wages and working hours.

Conclusion

Let us conduct a thought experiment. We will assume that a) two groups of workers (women and men) have the same productive characteristics, and b) there is perfect wage equality between them (i.e. they are paid only on the basis of their productive characteristics and not on any other factors, such as gender, sympathy, etc.), so there is no discrimination. Assuming that women and men have the same productive characteristics and that there is perfect pay equality between them, they should receive the same hourly wage. If both groups work for the same amount of time, they receive the same monthly wage. However, if one group works 8.6% longer than the other, they should receive 8.6% more pay than the other group, assuming equal pay. This is almost exactly the pay gap that exists between the two groups to which Křížková and Pospíšilová (2023) are referring (if we disregard the difference in actual hours worked between men and women, because the data from the employers, as we have seen, do not include the difference in actual hours worked).[32]

[32] Křížková and Pospíšilová (2023) suggest that there is a 9% wage gap, which they interpret as the difference in remuneration between women and men *for the same work* (Kříž-

It should be noted that the men's group has, on average, four more years of total job experience. Additionally, an extra year of experience increases the wage rate by 1.1%. Consequently, the group with one more year of experience should, on average, earn 1.1% more than the other group. If the more experienced group has a total of four more years of experience, it should earn 4.4% more than the other group (assuming that both groups have the same other productive characteristics and that there is no gender discrimination).

As we have shown, the fact of a non-linear function between wage level and hours worked also has a very significant effect, explaining a significant part of the GPG (and thus reducing the unexplained part of GPG).

Considering these factors (i.e. the difference in real working time, the difference in length of experience between both groups and non-linear function between hours actually worked and salary), it seems *highly unlikely* that there is (in reality) any remaining space for 'unequal pay for equal work' between the two groups. Naturally, this hypothesis must be empirically proven and calculated (which will be done in the next paper; this paper deliberately focuses only on *critical* aspects of existing measurements of the AGPG). However, if the gender pay gap for equal work persists, it is likely to be an extremely small part of the unadjusted gender pay gap. This point has been well expressed by recent Nobel Laureate Claudia Goldin (2023):

> Are women actually receiving lower pay for equal work? By and large, not so much anymore. Pay discrimination in terms of unequal earnings for the same work accounts for a small fraction of the total earnings gap. Today, the problem is different. (Goldin 2023, p. 4)

We have addressed the calculation of the explained and unexplained parts of the gender pay gap in the Czech Republic by analysing the different factors influencing the size of both parts. Existing studies indicate that one-third[33] or at most one-half of the unadjusted GPG can be explained by the different productive characteristics of men and women (Křížková & Pospíšilová, 2023, p. 45). The

ková and Pospíšilová, 2023, p. 54). We respectfully disagree with this interpretation, as we believe it may be premature and potentially misleading. It would be inaccurate to say that their analysis showed that women and men receive different pay for the same work. Rather, it demonstrated a gender pay gap for workers of the same occupational category, age, workplace, and type of contracted working time (full-time hours). It is important to note that significant gender differences were not considered in their analysis. These include total length of job experience, hours *actually* worked, position within the company, and specialization within the occupational category. For this reason, we believe that their conclusion that there is unequal pay for equal work may not be fully supported and requires further analysis.

[33] '... there is still a so-called "unexplained" gender pay gap, which accounts for two-thirds of the gender pay gap in the EU Member States' (European Commission, 2021, p. 3).

other unexplained half represents (as standardly interpreted) discrimination in pay for equal work or for work of equal value. The AGPG value of 9%–11% for the Czech Republic indicated by these studies was incorporated into the work of the Czech Government[34] and is understood therein as unequal pay for equal work. The assumption of wage discrimination also occurred in a European Commission document aimed at addressing 'pay discrimination and bias in pay structures' (European Commission, 2021, p. 3).

As previously demonstrated, conclusions about high levels of unequal pay for women for the same work or work of equal value are presented as undeniable facts of current social reality. However, they are based on several problematic assumptions that artificially increase the adjusted gender pay gap in statistical analyses. It seems evident that the three main factors that determine wage levels are occupational choice, total length of experience and number of hours worked. These productive characteristics significantly differentiate working men and women. If Eurostat effectively nullifies all three of these factors or even considers their contribution to the explained part of the gender pay gap negative (see above), it is not surprising that the result of such an analysis is that the size of the explained part of the gender pay gap is negligible, while the size of the unexplained part (i.e. AGPG) is extremely high. Given that the conclusions drawn in this manner form the basis of European and national policies aimed at addressing gender pay inequality, it is desirable that the resulting conclusions of the analyses are as close to reality as possible. Therefore, the aims of this article were to highlight these problematic aspects of the AGPG calculation and to provide a theoretical and conceptual criticism of them. This critique is not an end in itself. It is a preparation for more detailed empirical modelling, which should show what the actual AGPG is, whether the relevant parameters were considered and whether the biases described in this paper were eliminated. The theoretical analysis carried out here will therefore be followed by a systematic empirical and analytical analysis that should more accurately and rigorously quantify the true AGPG and the actual impact of each variable.

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[34] Government of the Czech Republic (2022).

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The Role of Educational Choices in Support of Gender Equality in Unpaid Domestic Work: A Case Study of Psychology and STEM Students in Slovakia*

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Abstract: The present paper investigates female and male students' perceptions of descriptive and prescriptive gender norms in Slovakia and their expectations of and preferences for unpaid domestic work in the students' future family lives. We explore the Slovak subset of the 'Understanding Communal Orientation in Men' (UCOM) project, which aims to better understand the social-psychological factors associated with students' interest in taking care-oriented roles and occupations. Data were collected at several universities across Slovakia in 2018 at psychology (as an example of the HEED field of study) and STEM departments traditionally dominated by female and male students, respectively. Our final sample consisted of 129 psychology students (106 females) and 124 STEM students (39 females). The results point to differences between female and male students' perceptions of gender norms regarding unpaid domestic work, with male students reporting both descriptive and prescriptive norms as being more equal than what female students note. Interestingly, the men and women agreed in their preferences and expectations of how unpaid work should and will be distributed in their future family lives. Regardless of biological sex, the students wanted—and expected—women to take on more of unpaid work, indicating that the decision to study does not foster the desire for gender equality in either female or male students. Next, we explored the associations between the decision to study traditionally gender-incongruent majors and our dependent variables. Because of the limitations of the sample and country-specific conditions, we can only cautiously suggest that the gender-incongruent major choices may become a meaningful indicator of the changing dynamics in how gendered roles and norms are understood in society.

Keywords: Educational gender gap, educational choices, gender norms, unpaid domestic work, gender equality
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Introduction

Equal access to education, including tertiary education, is believed to be a key factor in promoting gender equality and changing gender norms and stereotypes. The United Nations hopes that higher education can empower women both politically and economically, make women's voices heard across all social spheres and contribute to the more equal division of family roles, including the ability to sustain dual careers and share family responsibilities. In a recent report, UNESCO acknowledged that, globally, the gender gap in higher education has reversed, with 54% of students now being women (Bothwell, 2022). According to OECD data, the proportion of Slovak men and women aged 25–34 years old with tertiary education was nearly equal in 1998, reaching approximately 10% (Encinas-Martíni & Cherian, 2023). In line with global trends, however, after the year 2000, women started to outperform men. Currently, 51.2% of women in this age cohort have higher education compared with only 28.2% of men, and in the general population, 24% of women and 19% of men aged 15 and older have higher education (compared with 27% and 26%, respectively, in the EU).

Nevertheless, the data indicate that the progress towards political and economic gender equality and the balance between paid and unpaid domestic work is slower than the pace at which the educational gender gap has closed. Moreover, inequalities persist within the educational systems themselves, resulting in a humanities bias or gendered segregation in fields of study (Bothwell, 2022). Globally, only 30% of students graduating from STEM are women, and the field tends to be stereotyped as a domain in which women can hardly excel (Makarova et al., 2019). In contrast, in the EU, 43% of women intend to graduate from education; health and welfare; humanities; and arts, whereas 21% of men intend to do so. In Slovakia, the humanities bias is even more pronounced: 51% of female students, compared with 26% of male students, undertake education; health and welfare; humanities; and arts as their majors (EIGE, 2022). Therefore, despite the gender gap being reversed, gendered educational patterns, along with stereotypical views of gender roles, may linger. By sorting students into future occupations, the biases could reach beyond academia, contributing to labour market segregation and a gender wage gap (Block et al., 2019) or gender inequality at home (Saxler et al., 2024). These aversive consequences may arise because the HEED and STEM occupations are associated with different rewards in the labour market, with STEM occupations being systematically better paid. Hence, the humanities bias may be associated with the petrification of the stereotypical perception of gender roles because it is believed that partners with less-demanding and lower-paid jobs should take on more childcare and household chores (Akerlof & Kranton, 2010; Becker, 1985). This calls into question whether the mere reversal of the educational gender gap could serve as an indicator of attenuating the stereotypical perceptions of gender roles, leading to an increasing level of gender equality. Thus, we investigated female and male students' perceptions of descriptive and prescriptive gender norms in Slovakia and their expectations of and preferences for the division of unpaid work in students' future family life. In addition, we explored whether gender-incongruent choices could be indicative of a changing understanding of gender roles and norms.

Literature review

Gendered division of household chores and childcare

Compared with men, women tend to spend more time at home with their newborn children and perform more tasks related to cleaning, cooking and caring (Klímová Chaloupková, 2018; Miller & Borgida, 2016; Olsson et al., 2023; Rudman et al., 2012; Rudman & Mescher, 2013). However, contrary to the standard economic expectations regarding women's comparative advantage in caregiving, this unequal division persists even among couples in which women earn more or are the sole earners. Generally, the gap between household chores and childcare is wider than the gap in earnings (Akerlof & Kranton, 2010; Becker, 1985). In addition to the economic consequences, including limited women's career opportunities and the gender wage gap, men's underrepresentation in childcare also results in lower well-being of children and of the men themselves (Meeussen et al., 2020; Olsson et al., 2023). Because the earnings data do not corroborate the comparative advantage hypothesis, the research has focused on other factors that could explain the unequal division of unpaid domestic work in families. The results have unambiguously pointed to gender norms about paid and unpaid work, showing how these are key factors preventing men from greater engagement in family life (Olsson et al., 2023).

Gender norms may have two nonmutually exclusive forms—descriptive and injunctive—that describe the social status quo and normative expectations. Both types of norms also reflect the important aspects of experiencing and appraising the social reality in which young people plan their futures and make decisions about their education trajectories and family lives. Inconsistencies between one's own preferences and socially constructed demands may have far-reaching consequences for the life choices of individuals living in a given society. Gender role theory claims that the internalisation of norms starts in early childhood (Eagly, 1987). As a result of the gendered nature of those norms, boys are socialised to be masculine (instrumental or agentic) and develop traits such as aggression, independence, ambition and rationality. On the other hand, girls are encouraged to be feminine (expressive or communal), warm, caring, emotional and socially oriented (Bem, 1974; Eagly, 1987; Eagly & Karau, 2002; Eagly & Wood, 2012). Descriptive norms reflect the frequency with which one gender is involved in some type of behaviour. In this case, the frequency itself is a consequence of socialisation and social expectations and may lead people to believe that relevant skills and abilities are directly related to biological sex and are immutable (Eagly et al., 2000, 2020). As a result of gendered socialisation, women *take on* more unpaid work and, thus, are believed to be more competent caregivers—a role that requires strong communal traits—and *are expected* to take on more tasks associated with unpaid domestic work. Men—who are less often seen as primary caregivers—are believed to be incapable of caring activities and are expected to refrain from becoming excessively involved in family life and prioritise what they are best at—their paid careers (Cukrowska-Torzewska & Lovasz, 2020). The perception of descriptive norms about unpaid work may reflect sensitivity to social cues as well as beliefs about the competence of men and women in childcare and household chores. Injunctive norms, in turn, reflect the normative expectations of society; these norms can be further divided into two subtypes and indicate either what behaviour is expected from men and women (prescriptive norms) or discouraged (proscriptive norms). The perception of injunctive norms is associated with beliefs about what one *should* and *must not* do. For example, those men who—despite the gendered expectations—endorse gender equality tend to show greater intentions to take parental leave (Olsson et al., 2023). However, this does not come without a cost: A recent study showed that men may face backlash for being family oriented. Specifically, those men who take more days of leave related to caring for family members or sickness face a heightened risk of being dismissed compared with women leaving to care for family members (Adamus & Ballová Mikušková, 2024). If behaviour that violates gender norms is associated with social punishment, those norms could interfere with an individual's preferences and expectations of future family life and career.

The distinction between expectations and preferences is vital for understanding the gendered division of caring and household chores. Despite living in a society with predominantly conservative views on gender roles, such as Slovakia, it is likely that young people would *prefer* a different division of paid and unpaid work than the one they observe in their everyday lives. However, if they assume that their partners are socialised to endorse the gender norms prevalent in this society and/or are afraid of the backlash of non-normative choices, they may expect that young people will not achieve what they would prefer. Hence, they may adjust their own expectations to the circumstances in which they live. Therefore, the present study investigated the differences between female and male students' perceptions of descriptive and prescriptive gender norms as well as their preferences and expectations for the future division of unpaid domestic work. We explored whether perceptions of descriptive and prescriptive gendered norms in society are associated with students' expectations and preferences for unpaid work (household chores and childcare) in their own family life in the context of gender and educational choices (majors). If closing the gender gap in higher education was associated with more progressive views about gender roles, then we would observe the desire for and expectation of equality in students' future family lives (Objective 1).

Gendered educational choices

Although the data and literature show that education *per se* is an important factor in shaping gender norms and promoting gender equality, the decision to study is not the only driver of views about gender roles and norms associated with the division of unpaid work (Šprocha et al., 2020). Gender role theory posits (Eagly, 1987; Eagly & Karau, 2002) that educational choices themselves are gendered. This theory claims that society values and expects different behaviours from men and women. Therefore, if different behaviours are considered acceptable and desirable for men and women, occupations and fields of study—such as STEM or psychology as examples of HEED fields of study—are not gender neutral either; some fields would then be considered appropriate for women and others consider for men. Indeed, in addition to quantitative data, the literature shows that STEM, particularly math and physics, is a field that is stereotypically viewed as associated with masculinity and manliness (Archer & Freedman, 1989; Cheryan et al., 2015; Dicke et al., 2019; Makarova et al., 2019). Even children as young as kindergarten age imagine and draw scientists as men do (Chambers, 1983), and both students and teachers assign more masculine than feminine traits to a person who is a scientist (Archer et al., 2010; Hand et al., 2017).

Because of persistent gender norms, individuals may refrain from choosing a career that is socially viewed as unsuitable for them for fear of failure or because of potential backlash (Rudman & Glick, 1999, 2001; Rudman & Phelan, 2008). Generally, individuals try to align their behaviour with the norms applicable to their biological sex, and those who internalise the gendered notions associated with some fields are likely to feel discouraged from studying these fields. A recent study by Reisel and Seehuus (2023) suggested that not only are majors gendered but female-type majors are also valued less, discouraging men—who are socially required to prioritise paid careers—from choosing them. Concurrently, girls who hold more traditional beliefs about gender roles in adolescence are more reluctant to choose STEM as their major later, have lower educational attainment and earn less (Dicke et al., 2019). If we understand gender norms as a complex set of beliefs about the roles that men and women play in society, then perhaps not the size of the educational gender gap but rather the ratio of incongruent major choices could serve as an indicator of decreasing support for traditional gender norms, here assuming the unequal division of paid and unpaid work. In other words, those who have already internalised more counterstereotypical beliefs in general may more easily choose incongruent fields of study. Hence, we delved into our data to identify the potential differences between female and male students who chose congruent and incongruent majors (Objective 2).

Methods

Procedure and sample

The data were collected as a part of the 'Understanding Communal Orientation in Men' (UCOM) project. UCOM was launched in 2017 and covered 94 sites in 59 countries, aiming to better understand the social-psychological factors that predict students' interest in taking care-oriented roles and occupations. In accordance with the UCOM project criteria, the sample should be either all students from an undergraduate/bachelor (or similar) psychology major or all students from a balance of different majors, with at least 80 students from a STEM-related major (e.g., engineering, math, chemistry, computer science, physics) and 80 participants from a HEED-related major (e.g., psychology, social work, sociology, nursing, education/teaching, counselling, health professions). The selection criteria set by the international UCOM project coordinators have a twofold justification. First, the HEED and STEM major choices are gendered and, thus, may show associations with the gender norms prevalent in society and/or endorsed by the students. Second, major choices are an important factor for labour market segregation and the gender pay gap. When more women choose to study majors associated with lower-paid HEED jobs, this may contribute to the skewed division of paid and unpaid work in families. This is because partners who earn less often take on more caring and household chores on themselves and more easily resign from or reduce their paid work engagement. Finally, psychology—as an example of a HEED major—was chosen for pragmatic reasons. In many countries, psychology is a widely popular major choice; thus, the potential pool of participants may be sufficient. Additionally, among the HEED majors, psychology has a considerable number of male students, facilitating the ability to obtain responses from males for comparison purposes. The present study was approved by the Ethics Committee of Slovak Academy of Sciences. We report here only the measures that were analysed in the current study (for more information, see variable overview: https://osf.io/jw8fh/wiki/home/; or project overview: https://ucom2017.wordpress.com/).

The data were collected at universities across Slovakia in 2018. The participants completed an online survey in the Slovak language that took approximately 45 minutes to complete. In accordance with the UCOM registered participation criteria (see OSF: https://osf.io/7psh5/?view_only=a6ef288322884140b788042819d926c9), those participants who completed fewer than 80% of the questions or did not pass at least one of two attention check questions were excluded. We recorded 837 attempts to start the questionnaire in Qualtrics. After quality and completeness checks, 262 studies were included in the final sample. After selection for the major, the final Slovak subsample consisted of 129 psychology students (106 females) and 135 STEM students (39 females). The small deviations from the sample size in some of the analyses were because of missing values deemed acceptable by the UCOM protocol. The STEM participants were of the following composition of majors: business, mathematics/statistics, computer science and engineering. Both psychology and STEM students were recruited from various universities across Slovakia with the aim of collecting responses from both male and female students.

Measures

Perceptions of descriptive norms about unpaid work

Perceptions of descriptive norms were assessed using two items with an instruction: 'We are interested in your perceptions about how parents in your society handle the division of labour at home. For the following questions, please imagine a family of four living together under one roof. The family consists of a father, a mother, and two children aged 2 and 6 years old. Make ratings of how the family would most likely divide tasks in Slovakia'. 'How much of the childcare (taking care of children, spending time with them and fulfilling their physical and psychological needs) do mothers and/or fathers do, respectively?' 'How much of the unpaid domestic work do mothers and/or fathers do, respectively?' The items were rated on a scale from 0 (*father does all*) to 100 (*mother does all*); subsequently, they were centred; 0 indicates equal division, negative values indicate greater shares of fathers, and positive values indicate greater shares of mothers. The scores were then summed to create composite score of unpaid domestic work (including childcare and household).

Perceptions of prescriptive norms about unpaid work

Perceptions of prescriptive norms were again assessed by two items with an instruction: 'We are interested in your beliefs about others' perceptions of the division of labour at home. For the following questions, please imagine a family of four living together under one roof. The family consists of a father, a mother, and two children aged 2 and 6 years old. Make ratings of how others in Slovakia believe these tasks should be distributed'. 'How much of the childcare (taking care of children, spending time with them and fulfilling their physical and psychological needs) do mothers and/or fathers should do, respectively?' 'How much of the unpaid domestic work do mothers and/or fathers should do, respectively?' The items were rated on a scale from 0 (*father should do all*) to 100 (*mother should do all*); subsequently, they were centred; 0 indicates equal division, negative values indicate greater shares of fathers, and positive values indicate greater shares of mothers. The scores were then summed to create composite score of unpaid domestic work (including childcare and household).

Preferences for the ideal division of unpaid work

Preferences for the ideal division of childcare and household work were assessed by two questions with an instruction: 'We are interested in your own beliefs. For the following questions, please imagine a family of four living together under one roof. The family consists of a father, a mother, and two children aged 6 and 2 years old. Make ratings of how you think these tasks should be distributed'. 'How much of the unpaid domestic work within the household do you think mothers and/or fathers should do, respectively?' 'How much of the childcare (taking care of children, spending time with them and fulfilling their physical and psychological needs) do you think mothers and/or fathers should do, respectively?' The items were rated on a scale from 0 (*fathers should do all*) to 100 (*mothers should do all*); subsequently, they were centred; 0 indicates equal division, negative values indicate greater shares of fathers, and positive values indicate greater shares of mothers. The scores were then summed to create composite score of unpaid domestic work (including childcare and household).

Expectations about the division of unpaid work in future family life

Expectations concerning students' future family life and the division of childcare and household work were assessed by two questions with an instruction: 'We are interested in your own expectations about the division of labour at home in your own future family. For the following questions, please imagine that you have a life partner and child(ren). Make ratings of how you believe these tasks will be distributed as accurately as you can'. 'How much of the unpaid domestic work within the household do you expect you and your partner will do, respectively?' 'How much of the childcare (taking care of children, spending time with them and fulfilling their physical and psychological needs) do you expect you and your partner will do, respectively?' The items were rated on a scale from 0 (*I would do all*) to 100 (*my partner would do all*); subsequently, they were centred; 0 indicates equal division, negative values indicate greater participant's own share, and positive values indicate greater expected partner's share. The scores were then summed to create composite score of unpaid domestic work (including childcare and household).

Educational choices (majors)

Educational choices were coded as a binary variable for either psychology or STEM. Following the UCOM project assumptions, *psychology* is considered a gender-incongruent major choice for men and *STEM* for women.

Results

First, we examined female and male students' perceptions of descriptive and prescriptive gender norms in Slovakia and their expectations of and preferences for unpaid domestic work in students' future family life. All measured variables (DN, PN, PREF, EXP) are centred—variables are calculated by subtracting 50 (mean = totally equal division); thus, positive values indicate greater shares of women/mothers whereas negative values indicate greater shares of men/fathers. We then added scores for household and childcare to obtain a score for *unpaid work* used for further analyses. The descriptive statistics for all the variables are presented in Table 1, and the results of the correlational analysis are presented in Table 2.

As shown in Table 2, being female positively correlated with the perceptions that women are expected to do and actually do more unpaid work. In other words, the participants tended to report that both descriptive and prescriptive

Table 1. Minimum, maximum, mean and standard deviations for all measured and created variables

	N	Minimum	Maximum	M	SD
Descriptive norms_ childcare	262	−30	50	21.98	14.41
Decriptive norms_ household	262	−30	50	20.76	14.38
Descriptive norms_ together	262	−41	100	42.74	24.37
Prescriptive norms_ household	262	−50	50	17.79	17.71
Prescriptive norms_ childcare	262	−45	50	19.66	16.40
Prescriptive norms_ together	262	−68	100	37.46	30.10
PREF_household	261	−40	43	7.22	11.21
PREF_childcare	261	−18	50	6.40	9.08
PREF_spolu	261	−58	80	13.62	17.71
Expected share_ household	254	−50	50	−2.41	15.26
Expected share_ childcare	253	−50	50	0.25	13.51
Expected share_ together	253	−100	100	−2.16	26.14

Table 2. Pearson's correlation for all measured variables

	1.	2.	3.	4.	5.
1. gender					
2. major	*−.513**** (N = 253)				
3. DN unpaid work	*.238**** (N = 253)	*−.200*** (N = 262)			
4. PN unpaid work	*.251**** (N = 253)	*−.239**** (N = 262)	*.175**** (N = 262)		
5. PREF unpaid work	.031 (N = 252)	.018 (N = 261)	*.181**** (N = 261)	*.167**** (N = 261)	
6. EXP unpaid work	*−.381**** (N = 244)	*.350**** (N = 253)	−.094 (N = 253)	−.119 (N = 253)	*−.151** (N = 253)

Note: $^{*}p < .05$, $^{**}p < .01$, $^{***}p < .001$. Gender is coded 1 = male, 2 = female. Major is coded 1 = psychology, 3 = STEM. DN – descriptive norms. PN – prescriptive norms. PREF – Preferred share. EXP – expected share.

norms are biased against women. In addition, perceived descriptive norms, prescriptive norms and the preferred distribution of unpaid work were positively correlated. However, the perception of this unequal distribution was also related to the greater preference for this type of unequal distribution.

Correlational analysis was performed on the whole sample; therefore, we report the main analyses comparing men's and women's perceptions of descriptive and prescriptive norms as well as the preferred and expected distributions of unpaid work (childcare and household chores) in Table 3.

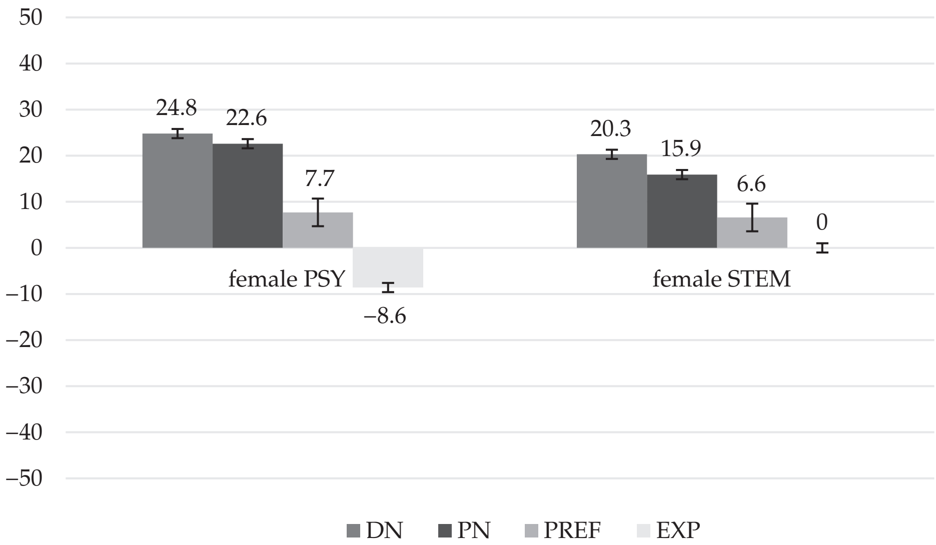
Men and women significantly differed in their perceptions of descriptive and prescriptive norms regarding the division of unpaid work. Although participants reported that women actually do (descriptive norms) and should (prescriptive norms) take on more unpaid work, men tended to perceive this inequality as smaller. Moreover, both men and women indicated that they would prefer the distribution of unpaid work to remain unequal (women are preferred to do more unpaid work). Nevertheless, the 'ideal' division in their view should be much closer to equality than how they perceive descriptive and prescriptive norms. Interestingly, the preferences of both the male and female participants were very similar. Finally, even though both men and women preferred that mothers do slightly more, they expected a more equal division of unpaid work than they reported. Although the numbers may suggest otherwise, men's and women's expectations about the division of unpaid work in their future family lives were strikingly aligned. Notably, negative values, in this case, indicated that the

Table 3. Comparison of men and women

	gender	N	M	SD	*t*	*p*	*d*	*[95% CI]*
DN_childcare	male	108	18.78	15.5	−3.08611	0.002	−0.39226	[−0.646, −0.137]
	female	145	24.37	13.21				
DN_household	male	108	17.45	14.3	−3.50181	< .001	−0.44510	[−0.700, −0.188]
	female	145	23.55	13.27				
DN_unpaid work	male	108	36.23	25.4	−3.88603	< .001	−0.49394	[−0.751, −0.235]
	female	145	47.92	22.29				
PN_childcare	male	108	14.96	14.2	−4.31217	< .001	−0.54810	[−0.807, −0.287]
	female	145	23.68	17.2				
PN_household	male	108	14.49	16.3	−2.90846	0.004	−0.36968	[−0.623, −0.115]
	female	145	20.81	17.67				
PN_unpaid work	male	108	29.45	25.9	−4.10271	< .001	−0.52148	[−0.779, −0.262]
	female	145	44.48	30.78				
PREF_childcare	male	107	6.38	10.2	−0.00847	0.993	−0.00108	[−0.251, 0.249]
	female	145	6.39	8.40				
PREF_household	male	107	6.62	12.8	−0.77458	0.439	−0.09872	[−0.349, 0.152]
	female	145	7.72	9.85				
PREF_unpaid work	male	107	13.00	20.8	−0.48854	0.626	−0.06226	[−0.312, 0.188]
	female	145	14.12	15.49				
EXP_childcare	male	103	6.43	11.7	6.47595	< .001	0.83940	[0.559, 1.117]
	female	141	−4.18	13.29				
EXP_household	male	103	3.19	12.8	5.13960	< .001	0.66520	[0.394, 0.933]
	female	142	−6.23	15.11				
EXP_unpaid work	male	103	9.62	21.9	6.41766	< .001	0.83185	[0.552, 1.109]
	female	141	−10.45	25.66				

Note: DN – descriptive norms, PN – prescriptive norms, PREF – preferred share, EXP – expected share.

Figure 1. Unpaid domestic work distribution differences between females in psychology majors and females in STEM majors with confidence interval values



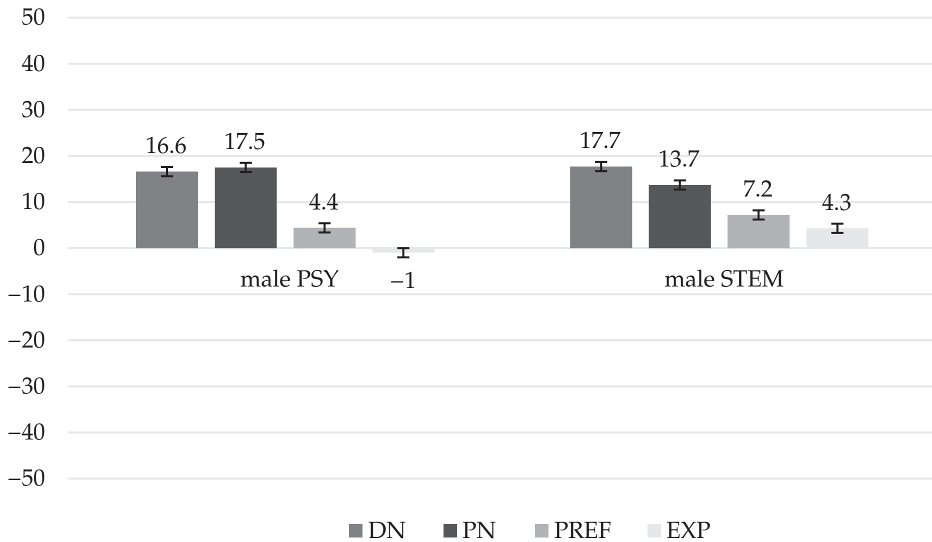
Note: DN – descriptive norm; PN – prescriptive norm; PREF – own preferences; EXP – expected share.
For DN, PN and PREF, 0 indicates equal division, negative values indicate greater shares of fathers whereas positive values indicate greater shares of mothers and summed.
For EXP: 0 indicates equal division, negative values indicate greater shares of participants whereas positive values indicate greater shares of expected partners.
Females in psychology CI values: DN [22.3–27.3]; PN [19.4–25.8]; PREF [5.8–9.6]; EXP [−11.3–−5.9]. The STEM CI values of the females were as follows: DN [16.2–24.4]; PN [10.0–21.8]; PREF [4.7–10.7]; and EXP [−4.9–4.9].

participants expected to do more of the unpaid work themselves whereas positive values indicated that they expected their partners to do more. Because the questionnaire asked about heterosexual families, we can see that women expected to do more themselves (negative values), whereas men expected their (female) partners to do more (positive values) in nearly equal proportions (−9.59 vs. 9.49).

As a secondary aim, we explored whether gender-incongruent choices could be indicative of a changing understanding of gender roles and norms.

The findings regarding female perception of the distribution of unpaid domestic work (see Fig. 1) show the decreasing participation of women in direction from descriptive norms to the expected share in one's own household. The Mann-Whitney U test was applied to analyse the differences between female psychology students and female STEM students. The mean rank differences were signifi-

Figure 2. Differences in the unpaid domestic work distribution between males in psychology majors and males in STEM majors with confidence interval values



Note: DN – descriptive norm; PN – prescriptive norm; PREF – own preferences; EXP – expected share.
For DN, PN and PREF, 0 indicates equal division, negative values indicate greater shares of fathers whereas positive values indicate greater shares of mothers and summed.
For EXP: 0 indicates equal division, negative values indicate greater shares of participants whereas positive values indicate greater shares of expected partners.
Males in psychology CI values: DN [12.1–21.1]; PN [12.1–22.9]; PREF [1.2–7.6]; EXP [1.2–7.6]. The STEM CI values for males were as follows: DN [14.5–20.9]; PN [10.1–17.3]; PREF [4.3–10.1]; and EXP [1.7–7.1].

cant for DN (U = 1576.5; p = .029), PN (U = 1565.5; p = .025) and EXP (U = 1483; p = .019). The expected share of unpaid work was around the equal gender distribution but (still) keeping the higher share of women's shoulders.

A similar 'decreasing' pattern as that for the female sample was observed for the male sample. The findings regarding male perception of the distribution of unpaid domestic work (see Fig. 2) show the decreasing participation of women in direction from descriptive norms to the expected share in one's own household. The Mann-Whitney U test was applied to analyse the differences between male psychology students and male STEM students. The mean rank differences were not significant. The expected share of unpaid work was close to the equal gender distribution, but the higher share of work was still expected to be done by the partner with whom the male respondent would be living.

The findings regarding male perceptions of the distribution of unpaid domestic work show the decreasing participation of women (increasing the participation of men) in the direction of descriptive norms to the expected share in one's own household. The differences in the expected share (EXP) between males with different majors indicated a desire for greater gender equality in domestic work distribution among male psychology students than among male STEM students (see Fig. 2). However, the differences were not statistically significant.

Discussion

The primary aim of the present study was to investigate female and male students' perceptions of descriptive and prescriptive norms regarding the division of unpaid domestic work in Slovak families as well as the students' preferences and expectations about this division in their future family lives. The secondary aim was to explore whether gender-incongruent major choices could be indicative of a changing understanding of gender roles and norms. Although our results need to be interpreted with the utmost caution because of the small sample sizes—which, however, reflect the real-life gender distributions in the two majors—being underpowered if placed in more advanced analyses, the present study signals interesting patterns that deserve further investigation. The findings highlight the differences between female and male students' perceptions of gender norms and between students who chose congruent and incongruent majors (psychology for men and STEM for women). The results indicate that it may not be the mere decision to study—which is relatively common now—but rather that an increasing number of gender-incongruent major choices may be indicators of the desire for greater equality in the division of unpaid domestic work between male and female partners.

Gendered perceptions of prescriptive and descriptive norms

Our results indicate that, regardless of major choices, there is a systematic difference in how the female and male students perceived prescriptive and descriptive gender norms. Male students believed that mothers are expected to do more unpaid domestic work. Nevertheless, compared with female students, mothers' share of unpaid work was lower. The data provided by the Gender Equality Index and Eurobarometer (Cukrowska-Torzewska & Lovasz, 2020) suggest that female students' beliefs—that women perform more household chores and caring activities—may actually be more realistic. Although our dataset does not allow us to explain this pattern, we can speculate that male students' perceptions of more engaged fathers could be associated with the students' early experiences. It is possible that fathers engage or feel more competent in rearing sons. Even if male students' memories and perceptions are biased, there is a constant need to emphasise that fathers' presence at early stages of life is important for both their daughters' and sons' well-beings and educations. Additionally, it seems important to reliably inform the extent of inequalities in the division of unpaid domestic work to stimulate greater understanding of and desire for equality in the future. If young men consistently underestimate the scope of inequality and the burden of unpaid work left on women's shoulders, it would be difficult to expect them to become vocal about the need for equality.

Expectations about the future division of household work and childcare

Our data show that, although female and male students are aware of the existence of gender descriptive and prescriptive norms, they strive to make their future family lives fairer and try to obtain a more equal division of unpaid domestic work than they observed in society. Regardless of their gender and major, the students generally voiced an expectation that their own family lives could be more balanced than those of their parents and people in contemporary society. Although the results show that current students do not wish to follow and reproduce prescriptive gender norms regulating the division of unpaid work in their own family life, they are far from encouraging. Primarily, it seems that there is still a persistent expectation that women should perform more caring and household chores. The results, however, need to be interpreted within the social context in which the data were collected. According to Eurobarometer data (Cukrowska-Torzewska & Lovasz, 2020), Slovakia is among the most conservative EU countries in terms of gender equality and the social perception of gender roles. Fifty-one percent of Slovaks believe that men are less competent than women in household tasks, and 48% support the view that a father must put his career ahead of looking after young children. Unsurprisingly, in the gender equality subindex time, which measures engagement in household chores and cooking and caring activities, Slovak men score far below the EU average: Slovakia scores 25th among all 27 countries and is ahead of only Greece and Bulgaria. Hence, we cannot forget that both preferences and expectations do not emerge in a void; they are shaped by and renegotiated in opposition to the descriptive and prescriptive norms prevalent in society. Thus, the results concerning the preferred and expected division of unpaid work—compared with the perception of gender norms—may seem to reflect generational changes rather than the effect of tertiary education. Although male students generally prefer their future female partners to take on a greater share of unpaid work, they also seem to understand and come to terms with the fact that gender norms have evolved and that they are expected to participate in unpaid work more. Finally, if we take a closer look at major choices, we can see an interesting pattern. Although the differences in preferences tend to be small, students who choose gender-incongruent majors (psychology for men and STEM for women) expect a more equal division of unpaid work in the future than their congruent counterparts. In the current study, the sample of students who had chosen a gender-incongruent major was too small to draw far-reaching conclusions. Nevertheless, the findings suggest that gender incongruence in major choices may be associated with the desire for and support of a more gender-balanced division of unpaid work. Even in the absence of clear causal patterns, this may indicate that the share of young women and men choosing gender-incongruent majors could reflect an increasing climate of gender equality.

Limitations

Despite our best efforts, the present study is not free from limitations associated with the data availability and quality, which were beyond our control. First, the data were cross-sectional, hence not allowing for observing possible evolution over time. Moreover, we are unable to draw far-reaching conclusions about the actual choices participants are likely to make in the future. Second, in the present analysis, we used the Slovak subset of the UCOM datasets, which focused only on psychology and STEM students; thus, the picture we have obtained is necessarily limited. Additionally, despite the attempts to reach as many male and female students in both fields, the sample did not meet the boundary conditions; thus, the resulting sample was underpowered for more advanced analysis. The inherent limitations of the capacity of the educational system in Slovakia forced us to use convenience sampling and snowball techniques to collect as many questionnaires as possible. Therefore, we need to remember that the sample composition may not be fully representative and, to some extent, biased by self-selection. To compensate for these limitations, future research could focus on longitudinal studies that compare students' attitudes, beliefs and expectations at least twice during their academic lives—at the beginning of their studies and around graduation—and across various faculties and fields of study.

Conclusions

Our results contradict previous findings that higher education could be the sole driver of gender equality. In contrast, the current study indicates that gender-incongruent major choices could become the indicators, if not drivers, of the desire for gender equality in unpaid domestic work to a greater extent than the mere decision to study. Although, regardless of their biological sex or major, all students showed preferences differing from descriptive and prescriptive norms, the students who chose majors that are traditionally seen as gender-incongruent strived for more equality in their future family lives. Moreover, the results concerning preferences and expectations about future family life show that, without policy interventions and support, gender norms and acceptance of men being more involved in family life are unlikely to change or change quickly enough to avoid perpetuating the unequal distribution of unpaid work. This inequality may be a disappointment for both men and women, who strive for greater work-life balance and general well-being.

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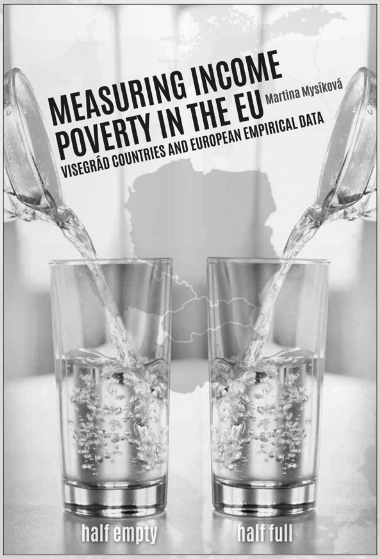
Introducing a new publication from the Institute of Sociology of the Czech Academy of Sciences

MEASURING INCOME POVERTY IN THE EU: VISEGRÁD COUNTRIES AND EUROPEAN EMPIRICAL DATA

Martina Mysíková

This book recapitulates the methodology of income poverty measurement applied in the EU and provides statistics and outlines the characteristics of the poor in Visegrád countries, and includes appendices with results for EU countries. After introducing the data, which is drawn from EU-SILC 2005-2018 and HBS 2010, the main analytical chapter focuses on methodological issues connected to measuring income poverty in a European context, with a focus on the suitability of the currently applied equivalence scales. The sensitivity of the at-risk-of-poverty rate to the OECD-type equivalence scale differs across countries. If the equivalence scale applied does not fit national conditions well, resulting income poverty rates may fail to accurately inform social policies, especially in countries with high sensitivity. Two sets of county-specific equivalence scales are estimated in this work: an expenditure-based scale using HBS data and a subjective equivalence scale based on subjective poverty lines and EU-SILC data. The book discusses the impacts of the estimated scales on income poverty rates and provides alternative subjective income poverty measures, which can usefully supplement objective income poverty data.



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Gender Differences in Intergenerational Occupational Persistence and Mobility in Central Europe*

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Abstract: This article investigates intergenerational occupational persistence and mobility across Central Europe (Austria, Czech Republic, Hungary, Poland and Slovakia) based on EU-SILC survey data from 2005, 2011 and 2019. Social Stratification in Eastern Europe survey data from 1993 is also used as a historical comparison. These surveys are uniquely suited for the analysis of occupational mobility because of their large sample sizes and the inclusion of detailed parental occupation data. I report gender differences in total and net mobility rates based on the analysis of 7×7 occupational mobility tables as well as predicted probabilities (derived from log odds from multinomial regression) of attaining specific occupational destinations based on parental occupational origins. The reproduction of occupational status is particularly strong in professional occupations (for both men and women), trade and crafts (for men) and sales/clerical occupations (for women), which seem to be in dynamic equilibrium. Compared with men, women's increases in social fluidity (and higher rates of upward mobility) are shaped much more strongly by changes in occupational structure, although this has weakened in both the Czech Republic and Slovakia. Finally, I find that women have much greater chances than men of upward mobility in attaining professional occupations from lower family origins, and this trend seems to have been strengthening in recent years.

Keywords: occupational persistence, occupational mobility, social mobility, social reproduction

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Introduction

The Visegrad states of Central Europe (Czech Republic, Hungary, Poland and Slovakia) have been subject to many political transformations in recent decades, including the collapse of communism, large-scale privatisation and market reforms, liberal (and illiberal) constitutional changes and accession to the European Union. Although it is easy to presume that the expansion of economic opportunities that have accompanied these changes would lead to greater social mobility, this is not necessarily the case. If social elites benefit the most from the changing opportunity structure, then political and economic transformation does not entail changing social mobility at all, as many sociologists have found in other and past contexts (Erikson & Goldthorpe, 1992); rather, this entails the persistence of the same elites in new clothes (Eyal et al., 1998). One reason for this seemingly counterintuitive finding is the way social mobility and other dimensions of relative inequality are measured: in terms of odds ratios, which are insensitive to the marginal distribution of social origins and destinations, such as changes in social class structure or any other analysed societal outcome. In fact, the current consensus among sociologists of the region is that the transition from communism to democracy did not entail any substantial change in social mobility at all (Bukodi & Goldthorpe, 2010; Džambazovič & Gerbery, 2018; Jackson & Evans, 2017).

The present article revisits the important question of whether and how social mobility has changed in Central Europe through the use of new data, a detailed look at change between specific strata and special attention to the role of gender. The article examines both occupational mobility and persistence, which are two sides of the same coin. *Occupational persistence*, also referred to as occupational reproduction, is indicated by the strength of the association between parents' occupational status and that of their children (Hout, 2018), such as the odds that the children of lawyers will also attain the same or similar professional occupations. The opposite of occupational persistence is *occupational mobility*: Individuals who hold occupational titles different from their parents exhibit intergenerational occupational mobility. That mobility may be *upward* or *downward*, depending on the assumptions we make about the rank order of occupations. An open society has a lot of intergenerational occupational mobility, while a closed society has a lot of persistence.

Early studies of intergenerational social mobility focused exclusively on men (e.g., Blau & Duncan, 1967), particularly mobility between fathers and sons, in part because of data limitations on women's occupations stemming from lower workforce participation. Although women's participation in the labour market substantially increased across postwar generations, social mobility researchers have continued to neglect gender differences in social mobility (Luke, 2019) and in fact theorise and generalise trends in international social mobility based on men's data (Bukodi & Goldthorpe, 2022). Even when data on women's mobility are available, it is not given the same analytical treatment as men's mobility, and direct comparisons of men's and women's relative mobility rates are often not made (e.g., Bukodi & Paskov, 2020). Even though social mobility research has a deep analytical tradition, there are major opportunities for better understanding international social mobility through the lens of gender.

That being said, the present study of intergenerational occupational persistence and mobility makes use of well-established methods in social stratification research (DiPrete & Grusky, 1990; Duncan & Hodge, 1963; Erikson & Goldthorpe, 1992; Goodman, 1979; Hout, 1988). However, there have been relatively few studies of intergenerational occupational mobility in Central Europe over the past several decades (the most notable exceptions are Gugushvili, 2017; Jackson & Evans, 2017; Mach, 2004; Róbert & Bukodi, 2004; Želinský et al., 2016), especially those comparing the countries under investigation here: the Visegrad states of the Czech Republic, Hungary, Poland and Slovakia, along with their most similar neighbour that was spared (thanks to the Austrian State Treaty of 1955) the experience communist authoritarianism. Even if the methods used in the present article are not new, the data and results presented here provide a fresh look into the question of the openness of Central European societies in recent years.

In the present article, I will first provide an overview of a cross-section of the key research on intergenerational social mobility in Europe, highlighting some of the important findings about postcommunist societies that inform the hypotheses enumerated below. In the data section, I introduce the 1993 Social Stratification in Eastern Europe survey and the 2005, 2011 and 2019 EU-SILC (European Union—Statistics on Income and Living Conditions) surveys, which are very large datasets with strong cross-national comparability and detailed information on the variables of interest, particularly parental occupation. In the results section, I present findings on intergenerational occupational persistence as well as absolute and relative intergenerational mobility. *Absolute mobility* refers to the extent to which individuals end up in different occupations from their parents, while *net mobility* or *social fluidity* refers to the strength of the association between occupational origins and destinations. The weaker the association between origins and destinations, the greater the social fluidity or 'openness' of the society in this key dimension of social stratification (Breen & Jonsson, 2007; Breen & Luijkx, 2004; Featherman et al., 1975). Net mobility is a summary statistic reflecting the *relative mobility rates* between the origins and destinations for specific occupational trajectories (49 such data points for a 7×7 mobility table for each social group of interest). I depart from the more technical social mobility literature by emphasising substantive differences between these five countries by gender and with respect to mobility between specific occupational categories.

Changes in social mobility in Central Europe

Although occupations are more detailed and less complex than the categories of social class, they are analysed in the social stratification literature in similar ways. That is, in the analysis of a social mobility table, the same analytical approach can be used regardless of whether occupational or social class origins and destinations are being examined. Depending on issues of data availability, many social mobility researchers have focused on occupational mobility (e.g., Connelly et al., 2016; Simkus, 1995). The use of occupational categories also has the advantage of being easy to understand what they refer to in the real world, are not subject to theoretical assumptions of different class schemas and do not require the additional variables required for constructing class schemas. For these reasons, the present article focuses on occupational mobility. However, because the literature on occupational and social class mobility is in fact the same (as they deal with the exact same issues and use the same methods), I will bring these strands of research together. Therefore, in this section, I will refer to the more general concept of social mobility or indicate whether occupational or class mobility was examined, when relevant.

In the late Cold War period, analyses of intergenerational social mobility in Europe showed divergent trends. Thélot (1983) found increasing relative mobility in France, as did Erikson et al. (1983) for Sweden and Ganzeboom et al. (1989) for the Netherlands. On the other hand, Erikson et al. (1983) did not find evidence of more fluidity in either France or England. These studies contested the FJH hypothesis (Featherman et al., 1975) that industrialised nations with capitalist economies and nuclear families would have roughly similar patterns of social mobility. These divergent results on trends in social mobility in the second half of the twentieth century were systematically revisited in Erikson and Goldthorpe's (1992) *A Constant Flux*, the culmination of the Comparative Analysis of Social Mobility in Industrial Nations (CASMIN) project, which involved the analysis of both absolute and relative social mobility in 12 industrialised nations, including Hungary and Poland, where relative mobility rates were stable during the communist period. The key finding of their breathtaking work is that, over time, intergenerational social mobility does not systematically increase or decrease, nor are countries converging towards a similar mobility regime; instead, relative social mobility resembles a constant flux across countries.

In a similar seminal work, Breen's (2004) *Social Mobility in Europe* used the same statistical techniques as Erikson and Goldthorpe (1992) but with more recent data covering the 1990s. Breen found constant flux in the mobility rates in Germany, Great Britain and Sweden, for example—for both men and women—but also found very modest increases in social fluidity (i.e., more relative mobility) in Hungary and Poland during the late communist period. However, those results were based on only several surveys, and the increase in fluidity was very modest. The most recent surveys covered by Breen in both countries indicate a flattening out of the relative mobility trend.

Our understanding of social mobility in European communist regimes is very much influenced by Hungarian and Polish data: These political regimes were the most reformist in the Eastern bloc in the 1980s, enabling sociologists to access large demographic datasets before it was possible in the 'normalisation' era of the more hardline Czechoslovak regime, for example. Mach (2004) found that Polish socialist industrialisation in the 1940s and 1950s led to high rates of upward and absolute mobility, which also endured longer for women than for men. The modest trend of increasing social fluidity in the 1970s and 1980s in Hungarian and Polish data (Breen, 2004) have suggested that socialist countries were somewhat successful in the forceful reallocation of occupations according to the needs of the command economy. This is also confirmed by Grusky and Hauser's (1984) finding that country-level exogenous interventions impacted relative social mobility. The same authors (Hauser & Grusky, 1998) later found 27% more exchanges in the mobility table between manual and nonmanual sectors in socialist countries compared with nonsocialist ones. Most of this change was because of government-dictated changes in the occupational structure rather than to the changing associations between occupations (Simkus, 1985; Zagorski, 1976).

The impact of the command economy on occupational structure is referred to as 'counter-selection' (Jackson & Evans, 2017) because the ideological objective of these regimes was to counter the effects of family background in determining occupational and class positions. Matějů (1993) pointed out that some occupations in communist Czechoslovakia were only available to people from working-class families. However, these counter-selective policies did not make communist regimes any more open than Western countries (Erikson & Goldthorpe, 1992). Sociologists' reliance on survey data of a limited quality and size means that it is difficult to reach robust conclusions about social mobility during the communist period.

As Andorka and Zagorski (1980) pointed out, the collectivisation of agriculture led to increased absolute mobility in socialist countries. However, this also accounted for an important divergence in the Polish and Hungarian data: Because Poland was the only socialist country that failed in the collectivisation of agriculture—thus maintaining a large sector of family farms—there was more intergenerational persistence in agriculture in Poland compared with Hungary. Erikson and Goldthorpe (1992) also found strong persistence in the upper service class, which could be linked to a '*nomenklatura*' or 'political capital' effect, meaning that the children of politically privileged socialist managers, for example, were able to attain similar class positions. Toth and Szelenyi (2019) and Toth (2019) similarly observed social closure within the Hungarian upper-middle class today, pointing to the 'Great Gatsby Curve' linking inequality and immobility. Along this line of thought, higher intergenerational occupational persistence in professional occupations in Central Europe indicates social closure, but it is not yet clear whether this social closure has increased, decreased or remained stable during the postcommunist transformation.

It should be noted that much of what we know about relative social mobility in Hungary and Poland is based on simple 3×3 mobility tables that were used but also criticised by Erikson and Goldthorpe (1992). Breen (2004) was also hesitant to make bold conclusions based on these data. A part of the problem of revisiting these old datasets is that the occupational categories used at the time were highly influenced by the socialist system and do not correspond to the ISCO categories developed in Western nations (Connelly et al., 2016). Analysing them, therefore, requires folding smaller occupations that lack comparability across time and space into much larger occupational categories that still make sense today, but doing so comes at the cost of losing a great deal of variance in the occupational categories.

This is also the case with the Czechoslovak datasets. For instance, despite the groundbreaking and important sociological achievements of Machonin's 1969 book *Czechoslovak society* (Machonin, 1969; 1970; Havelka & Machonin, 1997), the difficulty of collecting data on parental occupation, as opposed to characteristics of respondents, meant a much greater emphasis of Czech social stratification researchers on questions of social structure and differentiation than intergenerational mobility. Similar to Polish and Hungarian scholars, data limitations and the political salience of the 'class struggle' (including the need to censor variables that might call into question the victoriousness of the proletariat) entailed that Czech social stratification researchers were able to examine largely 2×2 intergenerational mobility tables, contrasting manual and nonmanual male workers (e.g. Charvát, 1978). These limitations continued in e.g. Machonin et al.'s (2000) important study on the development of the social structure in Czech Society from 1988 to 1999, analysing *intragenerational* mobility (change in a respondent's occupation from 1988 to 1999) in only five occupational categories, and did not analyse intergenerational occupational or class mobility (analysing educational mobility instead, and without respect to gender). In fact, a more detailed and sophisticated account of intergenerational social mobility using well-established contemporary methods did not emerge until Katrňák's and Fučík's (2010) exceptional study.

If command economies forced social mobility, it is likely that the transition to market economies in the early 1990s led to a reversal of direction. After the collapse of communism, countries in Central and Eastern Europe engaged in comprehensive market reforms: the liberalisation of markets, the stabilisation of public finances and the weaning away of enterprises from state subsidies and the comprehensive privatisation of enterprises, housing, land and other aspects of the property market (Gerber & Hout, 1998). Although the specific policies implemented in Poland, Hungary, Slovakia and the Czech Republic were different, by the early 2000s, all countries emerged from deep recessions because of these comprehensive reform packages and had functional, growing markets that enabled them to enter the European Union in 2004.

Therefore, the 1990s witnessed a modest strengthening of the association between family origin and occupational destinations or a 'return to social origin', as Katrňák and Fučík (2010) aptly expressed it. Much of Czech social stratification research in the 1990s conceived of occupations as hierarchically structured, thus entering into models of status attainment as an Index of Socio-economic Status (ISEI) or as a component of a composite variable of family SES. This reflects the lasting attraction of the Blau–Duncan model of status attainment and other indices of occupational status compared with the conceptualisation of occupations or social classes as discrete categories to be analysed via contingency tables. Thus, Katrňák and Fučík (2010) were the first Czech sociologists to estimate absolute and relative social mobility according to mobility tables, finding that the 1990s in the Czech Republic was marked by a decrease in social fluidity. This finding coincides with Czech research on educational inequality: Although educational expansion increased opportunities to study tertiary education, this did not lead to an increase in social fluidity (Simonová & Katrňák, 2016), largely because of the offsetting role of family origin. In other words, a modest decrease in Czech social fluidity may be linked to a modest increase in Czech educational inequality, akin to what the Great Gatsby Curve might predict (Jerrim & Macmillan, 2015).

These results on declining social fluidity are surprising because there has been a growing consensus that, in the 1990s, social fluidity increased in many developed countries (Breen & Jonsson, 2005; Breen & Luijkx 2004; Jonsson et al., 2011). One reason for this increase in social fluidity could be the role of educational expansion (Pfeffer & Hertel, 2015), which may provide students of different occupational origins with the qualifications needed for upward mobility. Nonetheless, the different trajectory of the trend in social mobility in Central Europe in the 1990s could be because of 'marketisation' (Jackson & Evans, 2017) or that Central Europe was 'catching up' to Western patterns. In the case of educational fluidity, Jackson and Evans (2017), Katrňák and Fučík (2010) and Simonová and Katrňák (2016) saw a levelling out or reversal of this divergence in the early 2000s.

In terms of the role of these structural changes on occupational persistence and mobility, we can note that the 1990s witnessed not only educational expansion but also an increase in income inequality (Večerník, 1996; Večerník & Matějů, 1999). In addition, educational expansion cannot keep pace with the demand for skills in certain occupations, leading to a rapid increase in wage returns to education across the region. It is likely that these increasing returns on education could strengthen—not weaken—social fluidity, depending on the strength of the association between family origin and education. Some research on educational inequality in the 1990s indicated a small decrease in the bond between parental educational attainment and that of their offspring (Simonová, 2003), while others indicated persistent inequality (Smith et al., 2016). Therefore, we can conclude that the empirical findings on social mobility from both educational and occupational perspectives are far from definitive, and ultimately, more insights can be gained by new analyses and methodological innovations in approaches.

However, what about the more recent period after the EU accession of Central European countries in 2004? By that time, Visegrad states had more or less completed their expansion of tertiary education, and the rising returns to education in these countries also began to level off. For these reasons, we can expect that social fluidity could have also remained stable, showing no clear direction in one way or another. In fact, Jackson and Evans (2017) found decreases in social fluidity across postcommunist countries, with a 9% increase in the chances of occupational persistence from the early 1990s to mid-2000s. However, their research suffers from low cell counts in some cross-tabulations, which may impact the robustness of their results. Therefore, we should be cautious about making bold proclamations about declines in social fluidity in postcommunist Central Europe before more analyses with different data sources are conducted. For the period under study (particularly 2005–2019), there is very little published research on occupational persistence and mobility beyond Jackson and Evans' study. Although our baseline assumption is that of stability (relative flux in occupational mobility) or modest increases in occupational persistence, the fact is that we do not know the impact of the Great Recession and other recent economic factors on these trends.

The limited research on occupational persistence and mobility in Central Europe has also suggested that little is known about gender differences in persistence and mobility. The very first studies of women's relative mobility rates found that 'there are no major differences in the patterns for males and females… Generalizations about occupational mobility which have been made for males apply to females' (DeJong et al., 1971, p. 1040). More recent studies on social mobility have continued to find similar mobility patterns for men and women in Europe (Breen, 2004; Bukodi & Paskov, 2020), especially in terms of relative mobility rates. In their new theoretical summary of social mobility research, Bukodi and Goldthorpe 'focus on those findings that are largely common across gender, and the theory we subsequently advance is intended to be gender neutral' (2020, p. 273), thus dismissing any theoretical value to gender. Although this limited literature has suggested that we should not expect major gender differences in mobility rates, this should not lead us to conclude that gender should be ignored in social mobility research or that gender differences cannot be discovered with closer analytical scrutiny, especially in the newest available data.

Hypotheses

Based on the above literature, it is possible to anticipate only modest, if any, changes in social mobility in the postcommunist period. We can differentiate hypotheses into two groups: those relating to absolute and net mobility and those relating to gender differences between specific occupational groups across surveys and countries.

First, major economic change—such as the transition to market economies in Central Europe in the 1990s or the rapid economic growth of the 2000s and the subsequent Great Recession—likely increased *absolute mobility*, which is also called the total mobility in a society. This can be because of changes in the occupational structure (*structural mobility*), which can affect men and women differently. The difference between total mobility and mobility because of changes in occupational structure is referred to as *net mobility*. Charvát (1978) provided a detailed explanation of these concepts of mobility regarding a dichotomous mobility table (a 2×2 table examining fathers and sons attaining manual vs. non-manual occupations).

Similarly, the rapid expansion of educational credentials in the region in the 1990s and 2000s, which contributed to a gender gap in educational attainments favouring women (Katrňák 2024), could have impacted the gendered distribution of qualified employees who can compete for higher-paying technical and professional occupations. Assuming that total mobility between men and women is more or less constant over time, we can hypothesise that women have experienced higher degrees of structural mobility (because of changes in educational and occupational structure) and that, therefore, men's net mobility rates would be higher than women's across the region (H1). However, we do not know whether men's higher net mobility translates into more upward mobility; on the contrary, the increasing pool of highly qualified women in the service sector might imply higher rates of upward mobility for women than for men (H2).

In terms of relative mobility rates, the literature above indicates that we should not observe any substantial gender differences (H3), despite quite significant changes in the labour market during the postcommunist transformation. This means that we should not expect gender differences in the odds ratios (or their derived predicted probabilities) of mobility between different occupational categories. Even though there has been an expansion of service sector jobs across the region, we do not have reasons to speculate gender differences in upward and downward mobility by gender, unless proven otherwise.

Because the four countries have undergone somewhat similar changes in the economic structure, we do not anticipate any major differences in their social mobility regimes (H4), thus anticipating common patterns and/or convergence, including with Austrian data.

Data and methods

The analysis of social mobility tables requires exceptionally large datasets because there must be a sufficient number of cases in all the cells of a contingency table of categories of parental and respondent occupation across both sexes and for each survey year. Representative samples of 5000+ respondents who provide detailed information on parental occupations are typically required. Because very few surveys in Central Europe have achieved this threshold, many scholars pooled together a diverse range of surveys: Katrňák and Fučík (2010), for example, used 28 surveys to study Czech intergenerational social mobility between 1990 and 2009.

To avoid issues in survey comparability and weights, I use only a few but large and high-quality datasets. Our first dataset is from the 1993 Social Stratification in Eastern Europe After 1989 (SSEE) survey, which was organised by Ivan Szelenyi and Donald J. Treiman. Surprisingly, the survey has not been widely used in research on social mobility—despite having very detailed occupational and job history data—with the notable exception of Donanski (1997), who used it to demonstrate that there was no increase in social fluidity in postcommunist countries from the 1980s to 1993. We use the survey to establish a 'baseline' level of social mobility in the early 1990s for all four Central European countries.

In addition to the SSEE data, our way out of the problem of data quality comes by virtue of the Visegrad states' accession to the European Union in 2004. All EU countries implement the annual and cross-nationally comparable EU-SILC survey (Survey on Income and Living Conditions), which also contains a number of rotating modules on different topics. In the first year of Czech, Polish, Hungarian and Slovak participation in EU-SILC in 2005, a rotating module on the intergenerational transfer of poverty was implemented (and repeated in 2011 and in 2019), which also has information on the occupational and educational attainments of the respondents' parents. EU-SILC surveys are also large and typically implemented by government or statistical agencies, which ensures response rates and data quality of a very high standard. Because of the large sample sizes, it is possible to analyse social mobility tables from these surveys without having to merge them with other data sources.

There are, of course, limitations to SSEE and EU-SILC data, such as the lack of details about parents' work, such as whether they supervised others. Depending on the country, missing information on mother's occupation can reach 20–35% of respondents, which is sufficiently large of a problem that it is not possible to analyse mother's occupation separately and have sufficient sample size, especially when combined with other covariates. However, sociological research suggests that a mother's occupational status can have a major role in the life outcomes of offspring (Hout, 2018). I address this issue by integrating fathers' and mothers' occupations into the highest parental occupation and use this variable as the metric for occupational origin. Because some respondents can recall their mothers' occupations but not their fathers', this solution also modestly increases the sample size compared with using only the father's occupation.

Occupational categories have to be comparable across surveys and between parents (fathers and/or mothers) and their children (respondents). Although 2005 EU-SILC data contain information on parents' occupation at ISCO level 2, 2011 and 2019 EU-SILC data have information only at ISCO level 1, which are basically 10 broad categories of occupational status. One of the occupational categories—'armed forces' occupations—is problematic because of the small number of women in the military (not only of respondents but also their mothers). Therefore, I had two options: either omit respondents with these occupations (or if their fathers had these occupations) or merge them with another occupational category. Because many armed force occupations involve hands-on technical work, such as the operation and maintenance of technical equipment, I subsumed these occupations into the occupational category of 'machine operators'. I did not find that this analytical choice impacted the results in any substantive way, but it does reduce the low cell counts in less common origin–destination combinations.

Clerical support workers and sales workers were also merged into a single category because both sets of work are considered in the sociological literature as low-status service occupations, and thus, it is not meaningful to examine change between them. These are also occupations that are predominantly filled by women, especially among the generations of respondents' parents, so merging these categories also ensured sufficient cell counts in origin–destination combinations for men.

Another challenge involves the category of 'skilled agricultural, forestry and fishery workers'. The problem here was not occupational gender segregation but rather the low cell counts for respondents in this category compared with their parents because the number of workers in these occupations declined significantly after the collapse of communism. The choice was to either omit these respondents or merge them with another occupational category. I decided to merge these occupations with the neighbouring category of 'crafts and related trades workers', thus bringing together forestry workers and electrical workers, for example. The workers in this larger occupational category also share the characteristic that many are self-employed or work in small enterprises. Merging these occupations also ensures that the category of 'crafts and related trades workers' has sufficient sample size across generations, genders and other categories of analysis. The analysed occupations are outlined in Table 1 and are identical for both the respondents and their parents.

The other variables used in the present article are very straightforward: 'Female' is a binary variable for whether the respondent is female (=1) or male (=0). I take age and age-squared into consideration as continuous variables (there would be low cell counts in marginal situations if we recategorise age into birth cohorts). I initially planned to include educational variables, but their inclusion proved problematic (i.e., low cell counts), and the important role of education on social mobility deserves separate treatment in its own article.

Besides parental and respondent occupation, gender and age—the only other variables included in the analysis—are variables for survey year and country. Because relative mobility may vary significantly by gender, we need to allow for interaction effects. Therefore, I include two-way interactions of gender × year, gender × parental occupation, year × parental occupation and the three-way interaction of gender × year × parental occupation.

Although researchers often relied on log-linear models, for this analysis, I follow DiPrete (1990) and Wu and Treiman (2007) in their recommendations for social mobility researchers to use multinomial logistic regression, which is useful for taking into account covariates, such as gender, education or age. Multinomial regression (MNR) also simplifies the effort to include interaction effects. To interpret the results, because the odds ratios of the main effects cannot be interpreted

Table 1. Categories of parental and respondent occupations

ISCO Major Groups (Codes)	Recoding for this analysis	ISCO Occupational Group	Examples
1	1	Managers	CEOs, elected officials, managers in public and private sectors
2	2	Professionals	Scientists, doctors, teachers, lawyers and other professional occupations
3	3	Technicians and Associate Professionals	IT technicians, legal assistants, research assistants, nurses, and other associate professionals
4	4	Clerical Support Workers	Secretaries, bank tellers, and other office assistants
5	4	Services and Sales Workers	Shop clerks, salespersons, cashiers, child care workers
6	5	Skilled Agricultural, Forestry and Fishery Workers	Fishermen, skilled farmers, chicken growers, lumbermen, mixed crop growers
7	5	Craft and Related Trades Workers	Carpenters, construction workers, painters, welders, electricians, mechanics
8	6	Plant and Machine Operators And Assemblers	Factory workers at e.g. car plants, assembly line and equipment operators, truck drivers
9	7	Elementary Occupations	Cleaners and refuse operators, unskilled workers in factories, mines, food preparation, farming
0	6	Armed Forces Occupations	Officers and non-officers in the armed forces

separately from the interactions, I use the predicted probabilities derived from the odds ratios as computed for each respondent and then summarise these predicted probabilities for different occupational destinations by parental occupation, gender, country and survey year. The results are easy to interpret and can be converted to odds ratios if needed; they are also more complete than the common practice of reporting only the odds of movement between adjacent categories of the mobility table (Džambazovič & Gerbery, 2018).

Finally, any origin–destination change between occupational categories (on either side of the diagonal of the social mobility table) is considered occupational mobility, whether upward or downward. I do not make theoretically questionable assumptions that movement only into the highest or lowest occupational categories should be considered upward or downward mobility. In addition, it should be noted that the analysis of relative mobility rates using MNR computes the odds ratios for each possible origin–destination combination and, therefore, is agnostic to the theoretical question of what should be considered upward or downward mobility.

Results

Occupational persistence and total mobility

Tables 2–6 provide the 7×7 occupational classification tables for all five countries. The results in this section can be replicated from that data. In Table 2 (Austria), the top value in each cell refers to the 2005 data, the middle value 2011 data and the bottom value 2019 data. This is the same for Tables 3–6 (Czech Republic, Hungary, Poland and Slovakia), except that the top cell refers to 1993 data, with 2005, 2011 and 2019 data below it, for each cell.

Occupational persistence is measured as the percentage of respondents (sons or daughters) who report the same occupation as they reported for their father or mother, whoever had the higher status occupation. These percentages can be computed by dividing cases along the main diagonal of the mobility table by the total number of cases. We can observe modest cross-national variation in occupational persistence: It is particularly high among Polish men, 40% of whom attained the same occupational status as their parents in the 2005 data. Occupational persistence is generally lower in Austria and Slovakia compared with the other three countries. Occupational persistence among women is substantially lower than among men in Hungary and Poland, while the differences are much smaller in the other three countries (though, as a guiding principle, the social reproduction of occupational status is a bit higher among men than women). Contrary to Jackson and Evans' (2017) finding of decreasing fluidity, occupational persistence seems to indicate constant fluidity for both men and women in all countries between the observed time periods.

Total mobility is the opposite of occupational persistence; all respondents are categorised in one or the other (Table 7). Total mobility can be further differentiated into upward mobility (whether respondents have a 'higher' occupational status than their parents, as represented by the bottom left side of the main diagonal of the mobility table) and downward mobility. In all countries, the rates of upward mobility seem to reflect dynamic equilibrium (i.e., stability) for both men and women. However, women are substantially more upwardly mobile than men in all countries and years and experience less downward mobility. The ratio of upward to downward mobility is quite commonly twice as large for women as it is for men; this gender gap favouring women does not seem to change significantly over time, hence confirming hypothesis H2. This may be because of long-term changes in gender occupational segregation: Large numbers of women

Table 2. Occupational mobility table of parents to sons and daughters in Austria

	Sons							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	36	13	10	21	17	4	8	109
	31	33	31	13	19	4	6	137
	31	47	29	22	11	7	2	149
2 Professionals	15	32	15	8	2	5	3	80
	26	76	49	23	16	3	5	198
	20	104	20	30	12	9	4	199
3 Technicians	25	28	76	38	23	15	7	212
	32	44	67	28	29	12	16	228
	26	59	56	35	41	11	4	232
4 Clerical /	25	29	70	167	80	18	31	420
Sales	110	134	169	142	160	62	55	832
	75	170	199	173	161	67	44	889
5 Craftsmen	47	22	127	143	309	82	62	792
and traders	95	79	187	148	357	141	145	1,152
	42	65	133	90	243	73	48	694
6 Machine	5	4	24	15	29	24	7	108
operators	11	12	31	30	36	19	8	147
	7	12	20	13	22	13	5	92
7 Basic occu-	14	13	41	47	81	38	69	303
pations	9	4	23	22	71	28	32	189
	6	4	20	31	35	26	16	138
total	167	141	363	439	541	186	187	2,024
	314	382	557	406	688	269	267	2,883
	207	461	477	394	525	206	123	2,393

	Daughters							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	11	24	15	49	9	2	7	117
	11	44	42	39	3	0	11	150
	12	58	21	42	5	1	4	143
2 Professionals	6	32	11	27	0	0	1	77
	12	74	38	35	2	2	6	169
	9	120	37	35	3	0	5	209
3 Technicians	16	21	38	105	6	2	22	210
	18	57	52	78	6	0	16	227
	8	78	53	65	10	1	9	224
4 Clerical /	13	35	41	278	24	5	42	438
Sales	43	154	226	444	31	8	84	990
	45	269	207	394	33	14	73	1,035
5 Craftsmen	12	25	47	397	83	19	124	697
and traders	48	81	159	506	89	31	238	1,152
	27	130	98	325	82	26	125	813
6 Machine	2	7	9	77	9	5	22	131
operators	6	12	27	87	6	2	28	168
	4	17	17	46	4	3	14	105
7 Basic	9	5	22	127	24	6	101	294
occupations	4	7	21	95	14	11	65	217
	1	8	10	71	11	6	41	148
total	69	149	183	1,060	155	29	319	1,964
	142	429	565	1,284	151	54	448	3,073
	106	680	443	978	148	51	271	2,677

Note: Data in each cell refer to distributions for 2005, 2011 and 2019, from top to bottom.

Table 3. Occupational mobility table of parents to sons and daughters in the Czech Republic

	Sons							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	34	26	22	8	38	19	6	153
	18	13	24	11	18	13	5	102
	25	25	28	14	18	12	2	124
	35	46	42	22	20	20	3	188
2 Professionals	29	49	31	7	24	10	3	153
	30	46	49	20	15	9	2	171
	38	89	80	25	48	34	7	321
	53	183	118	78	73	38	18	561
3 Technicians	34	46	47	14	57	28	4	230
	42	59	136	66	125	46	8	482
	32	65	141	46	112	50	8	454
	42	103	134	67	103	72	10	531
4 Clerical /	55	41	66	39	150	60	19	430
Sales	20	41	76	75	195	80	17	504
	40	44	140	87	244	139	34	728
	68	104	179	146	316	224	48	1,085
5 Craftsmen	55	47	87	53	371	185	65	863
and traders	37	29	93	77	389	164	46	835
	33	31	96	78	343	165	42	788
	27	40	82	88	300	193	47	777
6 Machine	26	10	34	16	112	61	22	281
operators	6	6	20	16	60	37	9	154
	9	13	20	27	94	64	20	247
	4	14	28	22	94	96	22	280
7 Basic	17	7	24	17	95	47	32	239
occupations	4	3	21	7	50	31	15	131
	2	2	12	8	37	14	12	87
	1	8	4	15	50	36	22	136
total	250	226	311	154	847	410	151	2,349
	157	197	419	272	852	380	102	2,379
	179	269	517	285	896	478	125	2,749
	230	498	587	438	956	679	170	3,558

	Daughters							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	22	42	52	60	9	8	25	218
	9	26	42	41	11	2	4	135
	9	29	68	42	9	7	6	170
	14	68	35	65	6	5	10	203
2 Professionals	18	67	50	41	11	1	5	218
	9	62	62	35	5	5	4	182
	17	116	155	96	10	9	9	412
	18	230	93	172	20	14	17	564
3 Technicians	19	35	113	79	17	10	24	297
	14	76	197	125	27	12	22	473
	23	102	245	193	29	25	17	634
	19	161	116	209	20	21	24	570
4 Clerical /	24	63	114	163	44	29	62	499
Sales	15	56	129	230	48	30	30	538
	32	85	269	365	74	60	68	953
	32	200	178	507	57	78	90	1,142
5 Craftsmen	46	54	155	253	186	140	191	1,025
and traders	21	35	164	278	182	71	122	873
	26	40	182	331	134	128	130	971
	15	92	96	316	76	115	122	832
6 Machine	12	12	42	99	38	63	69	335
operators	3	11	35	54	27	31	38	199
	6	14	58	127	43	56	51	355
	4	33	34	124	24	59	65	343
7 Basic	5	10	32	60	42	42	90	281
occupations	2	5	17	33	15	10	44	126
	2	3	17	34	18	14	33	121
	3	9	14	48	9	12	36	131
total	146	283	558	755	347	293	466	2,848
	73	271	646	796	315	161	264	2,526
	115	389	994	1,188	317	299	314	3,616
	105	793	566	1,441	212	304	364	3,785

Note: Data in each cell refer to distributions for 2005, 2011 and 2019, from top to bottom.

Table 4. Occupational mobility table of parents to sons and daughters in Hungary

	Sons							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	15	10	9	11	32	9	4	90
	54	38	26	39	45	29	9	240
	38	84	38	35	53	31	8	287
	8	15	13	8	13	10	2	69
2 Professionals	16	35	12	21	25	10	2	121
	61	85	27	30	45	19	6	273
	36	205	76	46	64	25	24	476
	12	59	27	27	21	21	6	173
3 Technicians	8	19	15	16	30	8	9	105
	40	34	30	42	88	39	9	282
	40	101	75	80	130	72	34	532
	8	33	22	33	36	27	5	164
4 Clerical /	23	24	25	52	126	42	25	317
Sales	69	40	52	93	204	120	35	613
	72	116	125	200	379	254	102	1,248
	11	41	46	75	128	97	27	425
5 Craftsmen	43	37	60	74	428	137	130	909
and traders	113	57	67	161	666	276	176	1,516
	51	86	117	191	842	409	278	1,974
	10	24	28	81	273	125	71	612
6 Machine	3	6	10	11	68	35	17	150
operators	29	4	19	39	140	72	30	333
	19	31	37	63	301	208	131	790
	1	5	11	33	94	81	37	262
7 Basic	14	11	22	29	145	65	93	379
occupations	14	2	6	30	125	48	75	300
	11	12	25	45	217	112	176	598
	3	5	6	17	59	28	56	174
total	122	142	153	214	854	306	280	2,071
	380	260	227	434	1,313	603	340	3,557
	267	635	493	660	1,986	1,111	753	5,905
	53	182	153	274	624	389	204	1,879

	Daughters							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	9	28	14	32	4	0	2	89
	35	75	52	61	8	3	14	259
	29	114	66	88	10	3	17	327
	7	34	16	15	3	3	6	84
2 Professionals	3	44	21	21	6	0	3	98
	35	108	44	56	9	2	5	259
	28	282	97	125	12	15	15	574
	4	83	47	33	5	5	7	184
3 Technicians	4	27	29	56	18	2	6	142
	29	57	80	76	21	2	10	275
	36	153	129	212	21	20	34	605
	4	63	58	71	9	11	16	232
4 Clerical /	19	48	54	113	45	6	35	320
Sales	56	94	140	245	75	26	55	691
	60	255	241	573	92	90	117	1,428
	16	105	114	240	28	45	62	610
5 Craftsmen	26	72	103	286	234	50	256	1,027
and traders	88	87	216	478	303	99	366	1,637
	53	217	246	638	249	293	417	2,113
	8	58	117	226	89	131	152	781
6 Machine	2	13	23	48	38	20	38	182
operators	13	18	63	102	62	26	76	360
	30	79	93	281	81	154	178	896
	6	13	46	116	36	62	89	368
7 Basic	6	14	30	80	103	29	183	445
occupations	13	12	29	86	89	27	105	361
	12	44	61	150	90	114	232	703
	0	9	24	35	27	36	103	234
total	69	246	274	636	448	107	523	2,303
	269	451	624	1,104	567	185	631	3,831
	248	1,144	933	2,067	555	689	1,010	6,646
	45	365	422	736	197	293	435	2,493

Note: Data in each cell (from top to bottom) refer to distributions for 1993, 2005, 2011 and 2019.

Table 5. Occupational mobility table of parents to sons and daughters in Poland

	Sons							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	58	40	28	25	39	13	10	213
	66	68	64	45	61	51	12	367
	61	56	43	46	44	29	13	292
	62	117	63	43	57	59	8	409
2 Professionals	29	45	14	18	17	8	3	134
	62	206	80	77	90	55	16	586
	50	148	70	47	60	47	13	292
	74	234	106	92	112	85	16	719
3 Technicians	40	19	19	33	34	9	3	157
	67	117	124	98	180	102	23	711
	49	74	91	56	114	72	21	477
	73	138	109	87	143	104	25	679
4 Clerical /	51	19	33	59	82	32	18	294
Sales	74	109	141	182	476	235	76	1,293
	68	92	125	132	341	194	61	1,013
	69	148	129	194	341	247	69	1,197
5 Craftsmen	134	45	57	103	480	152	65	1,036
and traders	217	160	278	388	2,956	897	393	5,289
	151	115	173	244	1,834	578	269	3,364
	85	136	184	249	1,642	573	211	3,080
6 Machine	16	7	5	17	55	32	8	140
operators	33	33	60	87	282	178	76	749
	32	27	48	59	257	147	46	616
	27	35	43	62	181	156	44	548
7 Basic	13	5	6	23	76	25	23	171
occupations	19	8	23	45	224	96	79	494
	16	9	14	21	156	72	53	341
	10	13	16	39	129	63	55	325
total	341	180	162	278	783	271	130	2,145
	538	701	770	922	4,269	1,614	675	9,489
	427	521	564	605	2,806	1,139	476	6,538
	400	821	650	766	2,605	1,287	428	6,957

	Daughters							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	26	48	50	60	16	3	12	215
	48	125	69	114	20	5	23	404
	47	146	56	101	12	5	11	378
	71	183	68	108	25	5	13	473
2 Professionals	12	58	25	27	2	1	2	127
	32	368	100	121	19	9	20	669
	43	231	80	96	13	4	13	480
	56	439	127	196	32	10	22	882
3 Technicians	23	35	50	42	7	2	7	166
	39	239	187	223	62	15	48	813
	36	173	125	177	35	14	36	596
	49	280	165	245	46	24	51	860
4 Clerical /	23	54	77	117	21	6	29	327
Sales	52	277	202	504	166	45	124	1,370
	60	235	160	417	128	51	118	1,169
	83	291	182	606	160	49	150	1,521
5 Craftsmen	57	80	168	309	345	55	140	1,154
and traders	143	589	598	1,447	1,964	228	825	5,794
	99	383	376	931	1,085	167	514	3,555
	99	430	287	1,005	994	173	459	3,447
6 Machine	7	8	29	60	21	14	22	161
operators	15	90	98	261	145	46	108	763
	20	88	76	246	75	48	105	658
	24	97	66	243	122	45	139	736
7 Basic	7	13	17	62	29	13	50	191
occupations	14	48	56	142	106	26	162	554
	11	24	28	108	88	19	99	377
	8	29	31	157	70	19	111	425
total	155	296	416	677	441	94	262	2,341
	343	1,736	1,310	2,812	2,482	374	1,310	10,367
	316	1,280	901	2,076	1,436	308	896	7,213
	390	1,749	926	2,560	1,449	325	945	8,344

Note: Data in each cell (from top to bottom) refer to distributions for 1993, 2005, 2011 and 2019.

Table 6. Occupational mobility table of parents to sons and daughters in Slovakia

	Sons							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	24	15	18	3	25	13	4	102
	50	40	61	30	63	27	10	281
	37	30	41	20	26	15	5	174
	17	29	22	20	15	14	2	119
2 Professionals	23	33	19	3	27	9	3	117
	52	77	83	37	62	24	15	350
	39	86	83	49	54	29	6	346
	32	72	75	64	43	54	20	360
3 Technicians	25	24	28	17	44	18	5	161
	36	45	77	46	96	71	13	384
	34	55	118	56	98	90	13	464
	14	66	74	80	68	54	17	373
4 Clerical / Sales	31	20	36	29	124	62	13	315
	48	49	98	74	169	121	37	596
	35	62	141	130	224	168	45	805
	42	69	127	168	205	154	67	832
5 Craftsmen and traders	69	47	75	81	395	165	69	901
	46	43	90	65	316	222	77	859
	32	35	77	82	305	157	37	725
	14	44	56	91	210	124	59	598
6 Machine operators	21	14	16	22	116	69	18	276
	33	15	27	43	154	122	36	430
	15	12	33	36	108	90	33	327
	10	12	26	54	80	71	34	287
7 Basic occupations	21	12	14	17	108	67	35	274
	46	26	37	39	180	116	79	523
	6	10	30	20	129	65	38	298
	7	11	16	41	72	53	47	247
total	214	165	206	172	839	403	147	2,146
	311	295	473	334	1,040	703	267	3,423
	198	290	523	393	944	614	177	3,139
	136	303	396	518	693	524	246	2,816

	Daughters							
Parents	(1)	(2)	(3)	(4)	(5)	(6)	(7)	total
1 Managers	12	22	33	33	9	3	10	122
	26	110	79	81	11	6	9	322
	8	51	74	37	4	2	3	179
	9	37	27	37	8	4	5	127
2 Professionals	5	51	37	21	3	1	3	121
	17	148	87	75	13	8	14	362
	23	115	111	78	7	4	6	344
	23	148	66	139	11	11	13	411
3 Technicians	2	26	47	38	15	9	9	146
	30	110	117	142	23	15	19	456
	26	122	185	186	16	17	15	567
	17	97	98	166	22	15	23	438
4 Clerical /	13	41	67	122	36	26	27	332
Sales	33	106	142	241	35	32	44	633
	37	102	254	309	51	43	42	838
	37	97	149	402	54	75	63	919
5 Craftsmen	34	58	148	250	213	101	137	941
and traders	37	120	159	293	134	83	111	937
	30	74	186	278	70	63	107	808
	13	60	76	224	75	82	75	605
6 Machine	9	22	49	90	44	38	34	286
operators	20	52	91	169	57	54	56	499
	7	29	88	131	27	33	45	360
	5	15	39	128	36	45	33	301
7 Basic occu-	15	13	23	70	49	27	70	267
pations	24	53	90	168	71	55	139	600
	7	19	73	106	39	33	62	339
	5	5	23	77	31	36	54	231
Total	90	233	404	624	369	205	290	2,215
	187	699	765	1,169	344	253	392	3,809
	138	512	971	1,125	214	195	280	3,435
	109	501	478	1,173	237	268	266	3,032

Note: Data in each cell (from top to bottom) refer to distributions for 1993, 2005, 2011 and 2019.

with parents in blue-collar occupations—which were commonly assigned to both men and women in the communist period in somewhat equal numbers—now pursue careers in the service sector, such as sales or administrative occupations, which would be coded as upwardly mobile. Men, in contrast, are more likely to reproduce the occupational destinations of their parents. The data do not allow us to infer about the causes of this gender gap, however.

Net mobility and its gender gap

The fundamental problem of analysing total mobility is that it confounds changes in the occupational structure experienced by two different generations with coefficients of association between occupations. Women can have more mobility than men simply because of gender segregation in the labour market, which should not be confused with social fluidity. Recessions, government interventions, secular trends in economic development away from reliance on heavy industry towards a service-based economy and other factors can all impact the occupational structure and, thus, absolute mobility, without implying that social fluidity has changed.

Table 8 decomposes total mobility into the share that can be attributed to changes in occupational structure between generations. Structural mobility can be interpreted as a kind of 'forced' mobility: Sons and daughters have different occupations than their parents simply because the kinds of jobs open to them in the labour market have changed. Structural mobility can be computed directly from the cell counts at the margins of the mobility table, such as by subtracting from the total number of cases the marginal cell counts in the occupational categories that are smaller (either for parents or for children) and dividing that number by the total number of cases.

What is particularly striking is that structural mobility is substantially higher for women than for men in all countries and years. Given that a great deal of economic transformation took place between the survey years (when the respondent reports their occupation) and the time reference of parents' main occupation when the respondent was young, we would expect much larger shares of structural mobility. Instead, the lion's share of men's mobility is, in fact, because of social fluidity (net mobility) between origins and destinations.

In contrast, a large share of women's total mobility is structural in origin—women pursue different occupations than their parents, for example, because of gender differences in the labour market or differences in family–work preferences. There does not seem to be any cross-national or temporal patterns in this. When we subtract out these structural effects, we can observe that women's social fluidity is systematically lower than men's, hence confirming Hypothesis 1, with the women in postcommunist countries enjoying over 10% less social fluidity in Poland and Hungary, for example. What is very important to note, however, is that men's advantage in net mobility seems to be declining in Austria, the Czech

Table 7. Occupational persistence and total mobility in Central Europe, men | women, %

	Persistence	Total mobility	Upward	Downward	Ratio up/ down
			Austria		
2005	35.2 | 27.9	64.8 | 72.1	41.6 | 46.3	23.2 | 25.8	1.80 | 1.80
2011	25.1 | 24.0	74.9 | 76.0	45.1 | 51.9	29.8 | 24.1	1.52 | 2.15
2019	26.6 | 26.3	73.4 | 73.7	44.9 | 52.0	28.5 | 21.7	1.58 | 2.39
			Czech Republic		
1993	26.9 | 24.7	73.1 | 75.3	39.1 | 41.3	34.0 | 34.0	1.15 | 1.21
2005	30.1 | 29.9	69.9 | 70.3	30.6 | 39.9	39.3 | 30.2	0.78 | 1.32
2011	27.7 | 26.5	72.3 | 73.5	30.4 | 39.9	41.9 | 33.6	0.72 | 1.19
2019	25.7 | 22.8	74.3 | 77.2	29.9 | 40.5	44.4 | 36.7	0.67 | 1.10
			Hungary		
1993	32.5 | 27.4	67.5 | 72.6	34.4 | 44.6	33.1 | 27.9	1.04 | 1.60
2005	30.2 | 23.5	69.8 | 76.5	32.3 | 46.8	37.4 | 29.6	0.86 | 1.58
2011	29.5 | 24.8	70.5 | 75.2	30.6 | 44.6	39.8 | 30.6	0.77 | 1.45
2019	30.5 | 25.8	69.5 | 74.2	29.6 | 42.6	39.9 | 31.6	0.74 | 1.35
			Poland		
1993	33.4 | 28.2	66.6 | 71.8	36.3 | 47.2	30.3 | 24.6	1.20 | 1.91
2005	40.0 | 31.6	60.0 | 68.4	26.6 | 44.6	33.5 | 23.8	0.79 | 1.87
2011	37.7 | 28.4	62.3 | 71.6	28.3 | 45.5	34.0 | 26.1	0.83 | 1.74
2019	35.2 | 29.1	64.8 | 70.9	27.4 | 43.5	37.4 | 27.4	0.73 | 1.59
			Slovakia		
1993	28.6 | 25.0	71.4 | 75.0	40.2 | 47.6	31.4 | 27.4	1.28 | 1.74
2005	23.2 | 22.6	76.8 | 77.4	37.6 | 49.8	39.1 | 27.6	0.96 | 1.80
2011	25.6 | 22.8	74.4 | 77.2	33.6 | 49.2	40.7 | 28.0	0.83 | 1.76
2019	23.4 | 27.4	76.6 | 72.6	33.3 | 40.7	43.3 | 31.9	0.77 | 1.28

Republic and Slovakia. Czech women even achieved a modest advantage in relative mobility in the most recent survey year, the first time this happened in the observed data. Women's advantage in graduation rates from institutions of tertiary education—a gender gap that has been expanding in recent decades—may be impacting the gender gap in relative mobility, but more research is needed to confirm such causal linkages.

Gender differences in occupational persistence

As noted earlier, the social reproduction of occupational status is generally higher among men than among women across the region. However, there are significant gender differences in reproduction across different occupational groups. The indicators of occupational persistence are reported in Tables 9–15, which report the predicted probabilities of attaining different occupational categories (1–7), here depending on parents' occupation and gender. For instance, in Table 9, the predicted probabilities of occupational persistence are reported in the column for 'managers', that is, the probability that respondents with parents who are managers also attain managerial status. In Table 10, occupational persistence is reported in the 'professionals' column, that is, the probability of being a professional for children of parents who attained professional occupations. In each cell, I report the probabilities for men (left) and women (right).

Please note that the results from Tables 9–15 are based on a multinomial regression model (NOMREG command in SPSS) with respondent occupation (seven categories) on the left-hand side of the equation, as predicted by dummy variables for highest parental occupation (managerial occupations as the reference category), gender (men as the reference category), age and age-squared, survey year (2005 as the reference category) and two-way and three-way interactions between gender, survey year and parental occupation. The analysis was conducted separately for each country. Model fit statistics are very high (Nagelkerke r-square measures are all above 0.3), reflecting high degrees of occupational persistence.

Occupational persistence is particularly high in professional occupations, but it is even higher for women than for men in all countries and survey years. In the Czech Republic, women who have at least one parent with a professional occupation have a 40% probability of also attaining a professional job, compared with 32% for men. The gender gap in the occupational reproduction of professionals seems to be particularly large in Poland and Slovakia, favouring women. Although we do not examine here the attained income of men and women in these occupations (a topic for another paper), these are generally some of the highest earning jobs in the economy and require high educational qualifications.

There are even larger gender differences in occupational persistence in other occupations. Men hold a large advantage over women in the reproduction

Table 8. Gender gap in net occupational mobility in %, sons | daughters

	Total mobility	Structural mobility	Net mobility	Gender gap (%)
		Austria		
2005	64.8 | 72.1	18.1 | 36.6	46.7 | 35.5	11.2
2011	74.9 | 76.0	30.9 | 36.5	44.0 | 39.5	4.5
2019	73.4 | 73.7	28.4 | 30.4	45.0 | 43.3	1.7
		Czech Republic		
1993	73.1 | 75.3	16.2 | 26.9	56.9 | 48.4	8.5
2005	69.9 | 70.3	13.6 | 26.0	56.3 | 44.3	12.0
2011	72.3 | 73.5	18.0 | 21.8	54.3 | 51.7	2.6
2019	74.3 | 77.2	20.0 | 20.1	54.3 | 57.1	–2.8
		Hungary		
1993	67.5 | 72.6	12.4 | 29.3	55.1 | 43.3	11.8
2005	69.8 | 76.5	12.7 | 32.2	57.1 | 44.3	12.8
2011	70.5 | 75.2	11.0 | 27.7	59.5 | 47.5	12.0
2019	69.5 | 74.2	9.5 | 28.0	60.5 | 46.2	14.3
		Poland		
1993	66.6 | 71.8	14.5 | 35.9	52.1 | 35.9	16.2
2005	60.0 | 68.4	14.7 | 35.3	45.3 | 33.1	12.2
2011	62.3 | 71.6	17.0 | 35.1	45.3 | 36.5	8.8
2019	64.8 | 70.9	13.6 | 29.9	51.2 | 41.0	10.2
		Slovakia		
1993	71.4 | 75.0			
2005	76.8 | 77.4	16.7 | 31.0	59.6 | 46.6	13.0
2011	74.4 | 77.2	18.8 | 25.0	55.6 | 52.2	3.4
2019	76.6 | 72.6	13.2 | 13.8	63.4 | 58.8	4.6

Table 9. Predicted probabilities of attaining managerial occupations (ISCO 1) by parental occupation, sons | daughters

	Managers	Profes-sionals	Technici-ans	Clerical / Sales	Traders / Craftsmen	Machine operators	Basic occupations
				Austria			
2005	.330 | .094	.188 | .080	.118 | .076	.060 | .030	.059 | .017	.046 | .015	.046 | .031
2011	.212 | .073	.116 | .071	.114 | .071	.125 | .035	.065 | .022	.075 | .030	.048 | .009
2019	.210 | .085	.095 | .038	.114 | .036	.085 | .043	.061 | .033	.078 | .038	.044 | .007
				Czech Republic			
2005	.177 | .067	.175 | .050	.087 | .030	.040 | .028	.044 | .024	.039 | .015	.031 | .016
2011	.194 | .035	.110 | .042	.077 | .038	.048 | .027	.046 | .027	.036 | .014	.023 | .017
2019	.187 | .068	.094 | .032	.080 | .033	.063 | .028	.035 | .018	.014 | .012	.007 | .023
				Hungary			
2005	.225 | .141	.223 | .135	.142 | .106	.113 | .081	.075 | .054	.087 | .036	.047 | .036
2011	.119 | .080	.076 | .044	.055 | .055	.043 | .038	.023 | .019	.023 | .028	.015 | .013
2019	.116 | .083	.069 | .022	.048 | .017	.026 | .026	.016 | .010	.004 | .016	.017 | .000
				Poland			
2005	.180 | .119	.106 | .048	.094 | .048	.057 | .038	.041 | .025	.044 | .020	.057 | .025
2011	.216 | .111	.117 | .084	.101 | .062	.066 | .045	.043 | .027	.046 | .027	.044 | .032
2019	.152 | .150	.103 | .063	.107 | .057	.058 | .054	.028 | .029	.049 | .033	.031 | .019
				Slovakia			
2005	.178 | .081	.149 | .047	.094 | .066	.081 | .052	.054 | .040	.077 | .040	.088 | .040
2011	.184 | .045	.101 | .064	.069 | .042	.042 | .041	.043 | .031	.037 | .022	.020 | .018
2019	.141 | .070	.089 | .055	.037 | .039	.050 | .040	.023 | .021	.035 | .017	.028 | .021

Table 10. Predicted probabilities of attaining professional occupations (ISCO 2) by parental occupation, sons | daughters

	Managers	Profes-sionals	Technici-ans	Clerical / Sales	Traders / Craftsmen	Machine operators	Basic occupations
				Austria			
2005	.119 | .205	.400 | .416	.132 | .100	.069 | .080	.028 | .036	.037 | .053	.043 | .017
2011	.336 | .380	.440 | .521	.228 | .340	.171 | .205	.077 | .094	.116 | .077	.037 | .051
2019	.309 | .408	.526 | .583	.255 | .356	.192 | .261	.093 | .160	.134 | .165	.029 | .055
				Czech Republic			
2005	.128 | .193	.269 | .341	.122 | .161	.081 | .104	.035 | .040	.039 | .055	.030 | .040
2011	.226 | .212	.299 | .362	.143 | .180	.076 | .121	.042 | .066	.049 | .076	.058 | .050
2019	.248 | .336	.329 | .409	.193 | .283	.096 | .174	.052 | .111	.050 | .096	.059 | .070
				Hungary			
2005	.158 | .302	.311 | .417	.121 | .207	.065 | .136	.038 | .053	.012 | .050	.007 | .033
2011	.300 | .358	.443 | .498	.194 | .217	.093 | .173	.044 | .088	.038 | .071	.020 | .051
2019	.217 | .405	.342 | .451	.201 | .272	.097 | .172	.039 | .074	.019 | .035	.029 | .039
				Poland			
2005	.185 | .309	.352 | .550	.165 | .294	.084 | .202	.030 | .102	.044 | .118	.016 | .087
2011	.195 | .397	.345 | .498	.160 | .292	.099 | .208	.035 | .108	.047 | .134	.026 | .064
2019	.286 | .387	.326 | .498	.203 | .326	.124 | .191	.044 | .125	.064 | .132	.040 | .068
				Slovakia			
2005	.142 | .342	.220 | .409	.117 | .241	.082 | .168	.050 | .128	.035 | .104	.050 | .088
2011	.172 | .302	.240 | .340	.121 | .210	.073 | .117	.048 | .089	.037 | .092	.030 | .074
2019	.245 | .294	.200 | .360	.177 | .221	.083 | .151	.074 | .099	.042 | .050	.045 | .022

Table 11. Predicted probabilities of attaining technical occupations (ISCO 3) by parental occupation, sons | daughters

	Managers	Profes-sionals	Technici-ans	Clerical / Sales	Traders / Craftsmen	Machine operators	Basic occupations
				Austria			
2005	.092 | .128	.188 | .143	.359 | .181	.167 | .094	.160 | .067	.222 | .069	.135 | .075
2011	.146 | .160	.227 | .142	.259 | .198	.204 | .212	.170 | .148	.204 | .191	.106 | .115
2019	.199 | .148	.097 | .175	.246 | .237	.224 | .201	.191 | .122	.211 | .165	.146 | .068
				Czech Republic			
2005	.235 | .311	.287 | .341	.282 | .417	.151 | .240	.111 | .188	.130 | .176	.160 | .135
2011	.194 | .312	.212 | .235	.282 | .301	.172 | .192	.105 | .129	.085 | .104	.115 | .083
2019	.221 | .172	.208 | .165	.250 | .205	.165 | .155	.106 | .116	.101 | .099	.030 | .100
				Hungary			
2005	.108 | .210	.099 | .170	.106 | .291	.085 | .203	.044 | .132	.057 | .175	.020 | .083
2011	.139 | .196	.145 | .178	.139 | .258	.108 | .184	.063 | .063	.049 | .049	.047 | .047
2019	.188 | .191	.156 | .256	.134 | .250	.108 | .187	.046 | .150	.042 | .125	.035 | .103
				Poland			
2005	.174 | .171	.137 | .150	.174 | .230	.109 | .147	.053 | .103	.080 | .128	.467 | .101
2011	.123 | .140	.145 | .152	.179 | .213	.106 | .128	.051 | .103	.073 | .107	.035 | .061
2019	.154 | .144	.147 | .144	.161 | .192	.108 | .120	.060 | .083	.079 | .090	.049 | .073
				Slovakia			
2005	.217 | .245	.237 | .240	.201 | .257	.164 | .224	.105 | .170	.063 | .182	.071 | .150
2011	.276 | .345	.254 | .299	.280 | .333	.174 | .282	.121 | .228	.107 | .217	.110 | .207
2019	.183 | .213	.209 | .161	.199 | .224	.153 | .162	.094 | .126	.091 | .130	.065 | .100

Table 12. Predicted probabilities of attaining clerical / sales occupations (ISCO 4) by parental occupation, sons | daughters

	Managers	Profes-sionals	Technici-ans	Clerical / Sales	Traders / Craftsmen	Machine operators	Basic occupations
				Austria			
2005	.193 | .419	.100 | .351	.179 | .500	.398 | .635	.181 | .570	.139 | .588	.155 | .432
2011	.117 | .300	.111 | .213	.158 | .295	.189 | .425	.140 | .425	.184 | .482	.159 | .419
2019	.151 | .296	.154 | .167	.149 | .279	.193 | .381	.130 | .397	.145 | .428	.226 | .483
				Czech Republic			
2005	.108 | .304	.117 | .192	.137 | .264	.149 | .428	.092 | .318	.104 | .271	.053 | .262
2011	.113 | .306	.106 | .296	.132 | .368	.146 | .444	.117 | .384	.117 | .380	.081 | .314
2019	.113 | .321	.139 | .306	.127 | .366	.135 | .445	.114 | .380	.079 | .362	.111 | .371
				Hungary			
2005	.163 | .246	.110 | .216	.149 | .276	.152 | .355	.106 | .292	.117 | 283	.100 | .238
2011	.143 | .278	.111 | .211	.179 | .352	.185 | .403	.115 | .305	.092 | .318	.095 | .221
2019	.116 | .179	.157 | .180	.202 | .306	.177 | .394	.132 | .290	.126 | .316	.098 | .151
				Poland			
2005	.123 | .282	.131 | .181	.138 | .274	.141 | .368	.073 | .250	.116 | .342	.091 | .256
2011	.178 | .278	.117 | .200	.134 | .299	.154 | .368	.087 | .268	.110 | .389	.082 | .294
2019	.105 | .229	.128 | .223	.128 | .286	.162 | .399	.081 | .293	.114 | .331	.121 | .372
				Slovakia			
2005	.107 | .252	.106 | .207	.120 | .311	.124 | .381	.076 | .313	.100 | .339	.075 | .280
2011	.138 | .251	.156 | .241	.125 | .342	.176 | .399	.119 | .365	.135 | .392	.084 | .327
2019	.169 | .292	.178 | .339	.215 | .380	.202 | .438	.153 | .372	.189 | .426	.167 | .338

Table 13. Predicted probabilities of attaining trade / craftsmen occupations (ISCO 5) by parental occupation, sons | daughters

<table>
<tr><th></th><th>Managers</th><th>Profes-sionals</th><th>Technici-ans</th><th>Clerical / Sales</th><th>Traders / Craftsmen</th><th>Machine operators</th><th>Basic occupations</th></tr>
<tr><td colspan="8">Austria</td></tr>
<tr><td>2005</td><td>.156 | .077</td><td>.025 | .000</td><td>.109 | .029</td><td>.191 | .055</td><td>.390 | .119</td><td>.269 | .069</td><td>.267 | .082</td></tr>
<tr><td>2011</td><td>.131 | .027</td><td>.071 | .019</td><td>.127 | .031</td><td>.188 | .036</td><td>.347 | .109</td><td>.250 | .060</td><td>.360 | .069</td></tr>
<tr><td>2019</td><td>.069 | .029</td><td>.062 | .015</td><td>.171 | .046</td><td>.178 | .031</td><td>.351 | .102</td><td>.234 | .039</td><td>.248 | .075</td></tr>
<tr><td colspan="8">Czech Republic</td></tr>
<tr><td>2005</td><td>.177 | .082</td><td>.088 | .028</td><td>.259 | .057</td><td>.387 | .089</td><td>.466 | .209</td><td>.390 | .136</td><td>.382 | .119</td></tr>
<tr><td>2011</td><td>.145 | .053</td><td>.150 | .027</td><td>.251 | .041</td><td>.321 | .078</td><td>.440 | .140</td><td>.360 | .113</td><td>.391 | .141</td></tr>
<tr><td>2019</td><td>.108 | .030</td><td>.131 | .034</td><td>.195 | .035</td><td>.291 | .050</td><td>.385 | .091</td><td>.335 | .070</td><td>.370 | .069</td></tr>
<tr><td colspan="8">Hungary</td></tr>
<tr><td>2005</td><td>.188 | .032</td><td>.165 | .035</td><td>.312 | .076</td><td>.333 | .109</td><td>.439 | .185</td><td>.420 | .172</td><td>.417 | .247</td></tr>
<tr><td>2011</td><td>.185 | .031</td><td>.137 | .016</td><td>.243 | .035</td><td>.302 | .064</td><td>.425 | .115</td><td>.376 | .088</td><td>.365 | .117</td></tr>
<tr><td>2019</td><td>.188 | .036</td><td>.121 | .027</td><td>.219 | .039</td><td>.301 | .046</td><td>.446 | .114</td><td>.359 | .098</td><td>.339 | .115</td></tr>
<tr><td colspan="8">Poland</td></tr>
<tr><td>2005</td><td>.166 | .050</td><td>.154 | .028</td><td>.253 | .076</td><td>.368 | .121</td><td>.559 | .339</td><td>.377 | .190</td><td>.453 | 191</td></tr>
<tr><td>2011</td><td>.144 | .032</td><td>.138 | .027</td><td>.221 | .057</td><td>.325 | .111</td><td>.529 | .307</td><td>.360 | .119</td><td>.443 | .233</td></tr>
<tr><td>2019</td><td>.139 | .053</td><td>.156 | .036</td><td>.211 | .053</td><td>.285 | .105</td><td>.533 | .288</td><td>.330 | .166</td><td>.397 | .164</td></tr>
<tr><td colspan="8">Slovakia</td></tr>
<tr><td>2005</td><td>.224 | .034</td><td>.177 | .036</td><td>.250 | .050</td><td>.284 | .055</td><td>.368 | .143</td><td>.358 | .114</td><td>.344 | .118</td></tr>
<tr><td>2011</td><td>.149 | .022</td><td>.156 | .022</td><td>.218 | .023</td><td>.285 | .065</td><td>.429 | .089</td><td>.349 | .081</td><td>.433 | .118</td></tr>
<tr><td>2019</td><td>.126 | .063</td><td>.119 | .027</td><td>.182 | .050</td><td>.246 | .059</td><td>.351 | .124</td><td>.279 | .119</td><td>.292 | .119</td></tr>
</table>

Table 14. Predicted probabilities of attaining machine operator occupations (ISCO 6) by parental occupation, sons | daughters

	Managers	Professionals	Technici-ans	Clerical / Sales	Traders / Craftsmen	Machine operators	Basic occupations
				Austria			
2005	.037 | .017	.063 | .000	.071 | .010	.043 | .011	.104 | .013	.222 | .038	.125 | .020
2011	.029 | .000	.015 | .006	.053 | .000	.076 | .010	.117 | .030	.122 | .012	.153 | .051
2019	.048 | .007	.046 | .000	.048 | .005	.076 | .014	.106 | .032	.144 | .030	.190 | .041
				Czech Republic			
2005	.128 | .015	.053 | .028	.095 | .025	.159 | .056	.196 | .081	.240 | .156	.237 | .080
2011	.113 | .041	.112 | .017	.097 | .044	.199 | .067	.194 | .126	.283 | .172	.184 | .141
2019	.108 | .025	.067 | .025	.135 | .037	.206 | .069	.249 | .138	.345 | .172	.259 | .092
				Hungary			
2005	.121 | .012	.070 | .008	.138 | .007	.196 | .038	.182 | .061	.216 | .072	.160 | .075
2011	.105 | .009	.050 | .026	.134 | .031	.197 | .063	.206 | .139	.267 | .173	.184 | .172
2019	.145 | .036	.121 | .027	.165 | .048	.228 | .074	.204 | .167	.309 | .168	.161 | .154
				Poland			
2005	.139 | .012	.094 | .014	.144 | .019	.182 | .033	.170 | .039	.238 | .060	.194 | .047
2011	.110 | .013	.110 | .013	.170 | .020	.200 | .042	.182 | .047	.294 | .067	.223 | .048
2019	.144 | .011	.118 | .011	.153 | .028	.206 | .032	.186 | .050	.285 | .061	.194 | .045
				Slovakia			
2005	.096 | .019	.069 | .022	.185 | .033	.203 | .051	.258 | .089	.284 | .108	.222 | .092
2011	.075 | .017	.075 | .012	.160 | .027	.196 | .044	.195 | .062	.248 | .075	.198 | .080
2019	.117 | .032	.150 | .027	.145 | .034	.185 | .082	.207 | .135	.247 | .149	.215 | .156

Table 15. Predicted probabilities of attaining basic occupations (ISCO 7) by parental occupation, sons | daughters

	Managers	Professionals	Technici-ans	Clerical / Sales	Traders / Craftsmen	Machine operators	Basic occupations
				Austria			
2005	.073 | .060	.038 | .013	.033 | .105	.074 | .096	.078 | .178	.065 | .168	.228 | .344
2011	.029 | .060	.020 | .036	.061 | .066	.048 | .076	.083 | .174	.054 | .149	.138 | .286
2019	.014 | .028	.020 | .024	.018 | .041	.050 | .070	.068 | .154	.055 | .136	.117 | .272
				Czech Republic			
2005	.049 | .030	.012 | .022	.017 | .047	.034 | .056	.055 | .140	.058 | .191	.115 | .349
2011	.016 | .041	.013 | .022	.018 | .028	.039 | .068	.056 | .129	.069 | .141	.149 | .256
2019	.016 | .049	.032 | .030	.019 | .042	.044 | .079	.061 | .146	.076 | .189	.163 | .275
				Hungary			
2005	.038 | .057	.022 | .019	.032 | .036	.057 | .080	.116 | .224	.090 | .211	.250 | .291
2011	.011 | .049	.079 | .028	.058 | .053	.072 | .076	.124 | .196	.154 | .192	.274 | .326
2019	.029 | .071	.035 | .038	.030 | .069	.064 | .102	.116 | .194	.141 | .242	.322 | .439
				Poland			
2005	.033 | .057	.027 | .030	.032 | .059	.059 | .091	.074 | .142	.106 | .142	.160 | .292
2011	.034 | .029	.028 | .027	.036 | .057	.050 | .098	.072 | .140	.070 | .157	.147 | .268
2019	.020 | .027	.022 | .025	.037 | .059	.058 | .098	.068 | .133	.080 | .189	.169 | .259
				Slovakia			
2005	.036 | .028	.043 | .039	.034 | .042	.062 | .070	.090 | .119	.084 | .112	.151 | .232
2011	.006 | .017	.017 | .020	.028 | .023	.053 | .053	.046 | .136	.089 | .122	.124 | .177
2019	.017 | .039	.055 | .031	.045 | .052	.080 | .068	.098 | .123	.118 | .109	.189 | .230

of managerial positions, especially in Austria, the Czech Republic and Slovakia. Women are much more likely than men to have clerical/sales or basic occupations, if this was the highest attained occupation of their parents, while the opposite is true in blue-collar jobs (traders/craftsmen and machine operators), hence reflecting systemic occupational differences by gender. Over time, although gender differences in occupational persistence can be large in some occupational categories, there does not seem to be substantial change over time.

Relative mobility rates

The statistics on 'net mobility' or social fluidity in Table 8 are, in fact, summary statistics for relative mobility for each country and year, separate for men and women. The results of MNR break down those mean mobility rates into the probability of attaining each occupational destination by each occupational origin, by gender, country and year.

As Table 9 indicates, social reproduction of managerial occupations is low (typically less than 20% for men), and upward mobility of men of parents with professional occupations hovers around 10% in the Visegrad states and even lower for women. In fact, the probability of attaining a managerial position with family origins lower than technical occupations rarely exceeds 5% among men. These probabilities are higher in Austria compared with the Visegrad states, reflecting a stronger origin–destination link for the most prestigious jobs in that country, even though this association declined from 2005 to 2019.

The results for professional occupations (Table 10) deserve special attention because these occupations are numerous and often strongly linked to educational qualifications and good pay. Across all countries, women have higher probabilities of upward mobility into professional occupations than men (refuting Hypothesis 3), which is consistent over time and across all five lower occupational groups. It is also true that women have greater odds of downward mobility from managerial parents, although this impacts a much smaller share of people. Despite the high degree of social reproduction of professionals, this occupational group also exhibits a large degree of social fluidity in its inflow mobility. The evidence shows that women in particular have greater odds of obtaining professional occupations, even if coming from families with lower occupational status—likely because of investments in education and the expansion of tertiary education in Central Europe in the 1990s and 2000s, which deserves special attention in a separate paper. However, my results do not indicate large changes in these odds over the observed time period.

Moving to factory and warehouse occupations ('machine operators'), a substantial difference exists between the Visegrad states and Austria in the social fluidity of these occupations. In Austria, there seems to be rigidity between blue-collar and service occupations, with children of parents with service-based occupations being substantially less likely to become machine operators. In contrast, the probability that men whose parents had clerical/sales positions become machine operators is comparable to the probabilities among craftsmen and basic occupational origins. This is also true of parents with technical or lower-professional occupations, such as IT specialists and nurses. In other words, it is probably a lasting legacy of the communist period that machine operator positions exhibit qualities of higher social fluidity in the Visegrad states. There does not seem to be any change in this over the observed time period.

Basic occupations that require little or no skill have higher occupational reproduction among women, who are also more likely than men to face downward mobility from skilled blue-collar occupations and lower service sector positions. In 2019, in Hungary, 43% of women whose parents held basic occupations are likely to hold those same or similar occupations and are much higher than the probability of upward mobility into the next ranking occupations (16%). Both men and women whose parents held occupations in the service economy rarely fall into these rudimentary jobs.

Conclusion

In the present article, I have presented evidence for very high rates of women's social mobility, especially in the upward direction, even though much of this mobility is structural in nature. After subtracting out the effects of changes in occupational structure, men continue to have more social fluidity, except for the Czech Republic, in the 2019 data. The Czech trend—that Czech women now exhibit higher relative mobility than Czech men—is in fact quite remarkable because the trend seems to be systemic in nature and other countries in the region seem to be following Czech social trends in this same direction. To the best of my knowledge, this is the first time that it has ever been shown that Czech women exhibit more social fluidity than men, hence representing an inflection point in the labour market not so different than the reversal of educational gender gaps in recent decades.

Of course, this leads to a certain sociological curiosity as to whether there is an empirical link between these educational and labour market trends. We should note that, although the social mobility literature has shown that education can increase absolute mobility, it has not been shown to increase net or relative mobility. What we do know is that different educational pathways in the Czech Republic—which are themselves unevenly distributed by gender—impact social class destinations differently (Smith, 2019), especially regarding the highest social classes. At the same time, we can observe in the social mobility tables that women exhibit more upward mobility into professional occupations compared with men. These trends raise the possibility of an educational effect on Czech net mobility rates by gender, a topic that deserves special attention in future studies.

The above results also provide evidence of dynamic equilibrium over time. With a few exceptions, we can find very little temporal change in the overall social reproduction of occupations or in mobility between them. Patterns of mobility that are specific to occupations are present (e.g., recruitment into machine operator positions from a particularly wide range of family origins in the Visegrad states), but these patterns do not seem to markedly increase or decrease. Women experience much higher rates of upward mobility into desirable professional occupations—which we would hypothesise is because of the intervening role of educational expansion, but in this case, we would expect that these probabilities would increase over time, but they do not. The social mobility regimes of these five countries seem to be quite similar (confirming hypothesis 4), with a modest trend towards a reduced or nugatory gender gap in net mobility. A follow-up study with a larger set of European countries would be needed to determine whether this degree of similarity or convergence is distinctive to the region.

These results point to the importance of further research on social mobility in Central Europe. The mechanisms, if any, between the gender gap in educational attainments and gender differences in social mobility remain largely unknown. Similarly, research can also be conducted on the income distribution of occupations among workers who experience occupational persistence versus upward and downward mobility. Finally, regionally specific differences in social mobility between occupational categories (e.g., a 'postcommunist' effect) would be more visible with cross-national comparisons across Europe than with only Austrian data points. With the advent of larger and higher-quality datasets in Central Europe, new frontiers in research on social mobility are still on the horizon.

Finally, the results in the present article are also subject to caveats. The current research is based on high-quality EU-SILC data, but the confirmation of empirical trends should also be apparent in other sources, such as pooled European Social Survey data. My results also only speak to occupational mobility, which may or may not reflect changes and continuity in social class mobility or income mobility. To encourage more students of social stratification to examine questions of social mobility in Central Europe, I include Tables 2–6, from which my results in structural and net mobility can be replicated.

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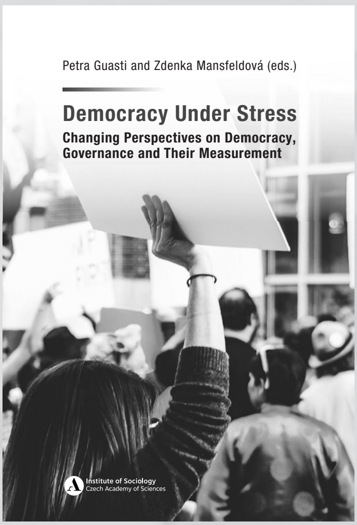
Democracy Under Stress. Changing Perspectives on Democracy, Governance and Their Measurement

Petra Guasti and Zdenka Mansfeldová (eds)

Democracy, defined as liberal pluralism, is under stress worldwide. Pluralistic democratic institutions such as: a free press, civil society, and the rule of law all seem to be under attack. Democracies are being hollowed out from within while preserving the fundamental facade of elections.



The strength of this book is in providing a range of perspectives on the study of democracy under stress. The authors, renowned scholars of democratic theory and democracy in Central and Eastern Europe, highlight the potential of different approaches – from comparative meta-assessment using indices and survey data, to case studies focused on understanding context and causal processes – for obtaining a better grasp of the loci of the stress.

Together, we offer the reader the opportunity to assess different conceptual frameworks and approaches, to reflect on their strengths and weaknesses, and to advance the study of democracy in the future. This volume is also an invitation for scholars to redirect their attention to Central and Eastern Europe, which offers an opportunity to deepen our understanding of democracy.

We see democracy in Central and Eastern Europe under stress but avoid general labels such as the crisis of democracy and deconsolidation. Instead, we argue that to understand the contemporary situation in the CEE region, we need to move beyond assessing institutional frameworks and to include citizens in our understanding and measurement of democracy.

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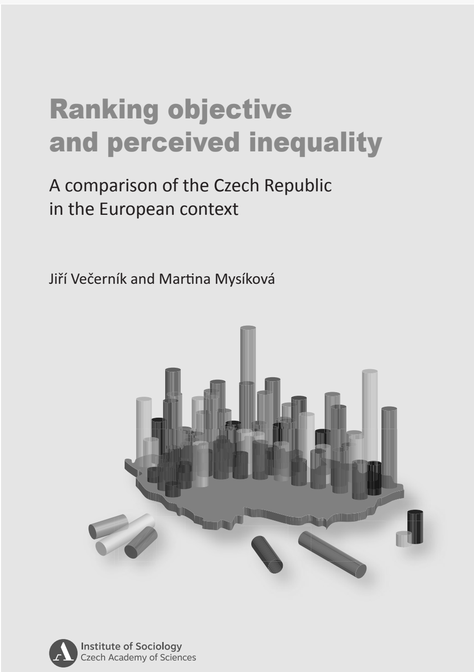
Ranking objective and perceived inequality. A comparison of the Czech Republic in the European context

Jiří Večerník and Martina Mysíková



In Czech public and professional discourse there is strong rhetoric about the rooted egalitarianism of Czech society and its extremely low socio-economic inequality. This study thus traces various objective and subjective dimensions of inequality in an attempt to examine the validity of this rhetoric. The study uses various sources of data on the levels and trends in earnings, household income, and living conditions in the Czech Republic and compares them to other European countries. It appears that although the country ranks among societies with a low level of social inequality, Czechs are not particularly 'exceptional' when it comes to objective economic equality, nor are they remarkably egalitarian in their attitudes.

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