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Gender Wage Gap in European Emerging Markets: A Meta-Analytic Perspective[*]

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Abstract: In this paper, we report the results of a meta-analysis of 670 estimates extracted from 53 previous research works to estimate the gender wage gap in European emerging markets. A meta-synthesis of collected estimates exhibits that the gender differences have a statistically significant and economically meaningful impact on wage levels. Synthesis results also reveal that the gender wage gap in countries with EU membership is lower than that in non-EU member states and, nevertheless, the wage gap between men and women has a tendency to diminish over time in the region as a whole. The meta-regression analysis of literature heterogeneity and test for publication selection bias back up the findings obtained from the meta-synthesis.

JEL classification numbers: D31, I26, J31, P23, P36

Keywords: gender wage gap, research synthesis, meta-regression analysis, publication selection bias, European emerging markets

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1 Introduction

After the fall of the Berlin Wall in November 1989, the countries of Central and Eastern Europe and the former Soviet Union rebuilt their socioeconomic systems based on the two pillars of democratization and the market economy. In addition, some of these countries set a goal of returning to Europe, that is, joining the European Union (EU). Such a major shift in national regimes and policy goals will inevitably have a major impact on the presence of female workers in these European emerging markets[1] and the distribution of wages between men and women. As explained later, in these countries, the female participation rate was relatively high during the socialist era. However, this does not mean that there was gender equality in wages. Thus, it was expected that the transformation to a democratic market-oriented state and accession to the EU might have a certain impact on the gender wage gap in the former socialist economies. Unsurprisingly, this issue has attracted the attention of many researchers.

In the Cold War era, the European socialist countries centered around the Soviet Union moved toward achieving legal equality of the sexes and working rights for women based on the ideology of socialism that denies gender distinction. For example, the constitution of the German Democratic Republic (East Germany) guaranteed gender equality, and its labor law stipulated the principle of equal pay for equal work. Thus, East Germany established legal gender equality in the field of labor (Satogami, 2011). At the same time, these socialist countries struggled with "the economics of shortage" (Kornai, 1980). They continuously lacked not only goods and services but also an adequate labor force. Under such circumstances, governments of these countries regarded women as a valuable economic resource, and their policies promoted their mobilization as a labor force. The fact that the development of childcare facilities and the institutionalization of maternity and parental leave have been realized in these European socialist countries at a rate no less than that of the Western developed countries with which they are in conflict is emblematic of their achievements (Fodor, 2005).

As a result, European socialist countries all boasted high female labor participation rates. Based upon this fact, these countries strongly promoted to the public that the

[1] In this paper, we adopt a broad definition of European emerging markets as economies in the post-communist Central and Eastern European countries that have experienced transition from the Soviet-type planned system to a market economy from the end of 1980s or the early 1990s and are in the process of catching up to the advanced industrial states of North America and Western Europe. It should be noted, in this regard, that some European emerging markets are now classified as high-income countries in the World Bank definition.

gender inequality in the labor field had been resolved in the Communist Bloc. Indeed, throughout the socialist era, the advancement of women's participation in the workplace has progressed considerably. There has also been a remarkable improvement in the level of women's education, and the opportunities for women who have completed higher education to play active roles as highly skilled personnel have greatly increased. However, at the same time, these countries, which were originally underdeveloped countries on the frontiers of Europe, still had a strong traditional social aspect in terms of gender division of labor; as a manifestation of this, the double burden of productive and reproductive labor fell heavily on women. In fact, according to Fodor and Balogh (2010), although some of the domestic work and childcare have been borne by society due to the development of nursery schools and after-school facilities, women are assumed to have primary responsibility for housework and childcare in labor legislation, and they were rather discriminated against as "workers different from men," since the parental leave system is exclusively for female workers.

This tendency was further accelerated when governments, confronting a decline in the birth rate against the background of long-term stagnation of the planned economy system that became apparent from around the 1980s, established generous systems[2] for female workers in order to balance childbirth and labor.

In other words, in European socialist countries, in spite of the outward signs of "gender equality," there were major gender problems not only in the labor market but also in the society as a whole. As will be described later, the gender gap in terms of wages was also quite serious in the European socialist bloc in this social situation (Brainerd, 2000).

Since around 1990, these European socialist countries have been promoting democratization and market economies. As mentioned earlier, gender equality during the socialist era was promoted to compensate for labor shortages caused by the socialist economic system and to embody socialist ideology. However, such goals were eliminated with regime change. The transition to a market economy wiped out the

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[2] For example, in East Germany, support for women workers was expanded in the 1970s, including better support for single mothers, one year's paid parental leave, and a guarantee of a return to work after parental leave. In 1977, the Labor Code included a section requiring all companies and factories to draw up affirmative action plans to enhance training for women, promote their appointment to leadership positions, and improve their working conditions. The expansion of support for women during this period is said to have been a response to the decline in the birth rate (Bauernschster and Rainer, 2012).

economy of shortage, and democratization brought freedom of speech and exposure to feminism, which had been excluded during the socialist era. In addition, countries that specifically aim to join the EU as a return to Europe must embrace EU principles in all aspects, including the acceptance of diversity; labor gender equality is a goal that must be achieved. Thus, the factors of democratization, the transition to a market economy, and the return to Europe are all factors that can be expected to work in the direction of increasing the status of women workers. In this way, regime change has also had a profound impact on the presence of female workers in the region and the distribution of wages between men and women, as it has led to a major shift in the economic system and policy goals related to female labor.

In reality, however, a number of issues needed to be considered. For example, in the Czech Republic, the idea of a division of labor between gender roles remained strong during the socialist period. In addition, generous maternity leave schemes, such as the extension of maternity leave to 28 weeks in 1987, meant that women were regarded as handicapped workers who had to combine work with housework and child rearing (Fultz et al., 2003). It has also been noted that the Czech Republic has a significant maternity and childcare penalty (Razzu, 2018), which suggests that these conditions worked against women in the tough job search during the transition recession. On the other hand, in Hungary, the reduction in the supply of childcare services and the system of childcare allowances that were not linked to employment, which had been abundant in the socialist era, occurred at an early point in the transition. At that time, there was no change in the values of the gender role division of labor that had continued even during the socialist era. This led women away from gainful employment (Fultz et al., 2003). Pollert (2005) also noted that, in CEE countries, achieving the major goal of economic reform was given top priority, and attention to gender equality took a backseat. In addition, there are non-negligible differences among emerging market countries in terms of minimum wage systems and wage determination systems, and the degree of gender wage inequality resulting from these differences in institutional arrangements varies widely (Brainerd, 2000).

A country's relationship with the EU could also have a significant impact on the distribution of wages between men and women in European emerging market countries. As is well known, countries wishing to join the EU are required to develop national laws that conform to the *acquis communautaire*, the general system of EU law. This includes legislation that guarantees gender equality in the field of labor. These changes in the legal system must be effective, and the creation of mechanisms and institutions to ensure that gender equality actually works was an important part of the EU accession negotiations (Sedelmeier, 2009). While the accession process would help close the gender gap in terms of wages in the countries involved in the negotiations, it was never clear in advance how much impact the advancement of accession process would actually have on the candidate countries as compared with other emerging market countries.

To settle these numerous issues, a number of empirical studies have attempted to measure the gender wage gap in European emerging market countries. A Russian study by Reilly (1999) and a study of eight Central and Eastern European countries by Newell and Reilly (2001) are the pioneers, and a Russian study by Rudakov and Prakhov (2020) and a Polish study by Magda and Sałach (2021) are the latest works in this field. Furthermore, even in literature that does not focus on the gender wage gap, the estimation results almost always reported that gender is the third most important variable in the wage function after education level and work experience. As a result, we now have gained some insight into the gender wage gap in almost all European emerging market countries, due to the research efforts over the past two decades.

However, Horie and Iwasaki (2022), in their study about the rate of return to education in European emerging markets, pointed out that the problem remains extremely serious in this research field as well. In other words, no previous studies have investigated the long-term trend of the gender wage gap throughout the transition period, and the reality remains unknown even more than 30 years after the collapse of the socialist bloc. In addition, wage research on European emerging markets has also lacked significant cross-national empirical results; therefore, cross-national deviations and regional trends in terms of the gender wage gap are almost completely unclear.

This paper conducts a meta-analysis covering a wide range of previous studies in order to solve the above-mentioned problems that single studies have struggled to address by showing a big picture of gender wage gap in European emerging markets. Specifically, it considers the lack of knowledge of the long-term trend of the gender wage gap and the inadequacy of cross-national comparisons.

Meta-analysis can provide certain answers to problems that standard empirical studies cannot or that are very difficult to verify due to technical reasons such as data limitations (Stanley and Doucouliagos, 2012). This approach is also useful in emerging market wage studies, as is proved by Horie and Iwasaki (2022), meta-analyses of Chinese studies by Iwasaki and Ma (2020, 2021), and Ma and Iwasaki (2021). In this paper, we propose our own hypothesis on the gender wage gap in European emerging markets, and then empirically investigate it through meta-integration, meta-regression analysis, and tests for publication bias to prove the validity of the research synthesis.

2 Gender Wage Gap in European Emerging Markets: Hypothesis Development

As mentioned at the beginning of the paper, European socialist countries have touted the ideology of gender equality and boasted a high rate of female labor participation. However, this does not mean that these countries have been free of the gender wage gap. In fact, Ferge (1997), Brainerd (2000), and many other studies have repeatedly pointed out the existence of a significant gender wage gap in the European socialist bloc. The reason for this is the difference in labor patterns between men and women in terms of industrial fields, occupations, and job classifications, which researchers mention in unison. In fact, the share of male workers in construction and heavy industry is clearly large, while the percentage of female workers in public service fields such as health and education is remarkably high, which is a conspicuous feature common to the entire European socialist bloc (Brainerd 2000; Satogami 2011). In the capitalist market economy, it is quite natural that there is some disparity in wage levels among industries; however, in the socialist planned economy, the national economic growth strategy tended to set higher wages in the heavy and chemical industries, which were considered important for national strategy, and lower wages in the service sector. This fact is well known among scholars of Central and Eastern Europe and the Soviet Union (Brainerd, 2000).

In addition to the differences in which industries they are mainly engaged and their wage levels, there is also discrimination between men and women in the matter of occupations and job classifications, with women being at a distinct disadvantage compared to men in terms of the types of jobs they can actually take, the frequency with which they can become managers, and the speed with which they can be promoted. This discrimination is also thought to have contributed greatly to the gender wage gap in the European socialist bloc (Brainerd, 2000). It is indisputable that occupations and job classifications strongly depend on the human capital of individual workers. However, as suggested by the case of East Germany, where the percentage of women enrolled in higher education institutions in the 1980s[3] surpassed that of men (Statistisches Amt der DDR, 1988), it is difficult to fully explain the low-wage occupations and lower job

[3] In fact, in East Germany, 70.6% and 50.2% of the students enrolled in vocational schools and universities, respectively, were women in 1987 (Statistisches Amt der DDR, 1988).

classifications of female workers in European socialist countries in terms of human capital. According to Ingham et al. (2001), many occupations in these countries were clearly differentiated into "male" and "female" occupations, with the latter concentrated in low-paying jobs with low potential for future development. Such institutional gender barriers should be considered as the main culprit. In addition, it has been noted that the gender norm that women were the ones responsible for housework and childcare remained strong, positioning women as handicapped workers and making it difficult for them to take up leadership positions (Fultz et al., 2003).

As mentioned above, there is no doubt that policy and institutional gender barriers existed in the European socialist countries in terms of labor. Consequently, it is highly possible that a significant gender wage gap has emerged in these countries. In fact, as will be shown in detail later, a series of empirical studies—such as that of Newell and Reilly (2001), whose observation period was in the 1980s—strongly suggest the possible existence of an extremely large gender wage gap at the end of the socialist era. As argued earlier by Stark (1996) and supported by a systematic review by Mizobata and Horie (2019), the process of market economization in European emerging market countries is strongly determined by initial conditions and is highly path dependent. On that account, it is unlikely that the "negative legacy" of socialism in the labor market has been immediately eliminated in these countries during the period from the 1990s to today. Therefore, we propose the following hypothesis about the gender wage gap in European emerging market.

Hypothesis 1: *An economically meaningful gender wage gap exists in European emerging markets.*

In the mid-1990s, some European emerging market countries began negotiating for EU accession in order to realize an early return to Europe while promoting transition to a capitalist market economy. In the so-called "Eastern Enlargement of the EU" process, the EU Commission not only assessed whether the negotiating parties had met the conditions for accession but also provided various technical and financial support to ensure that the conditions were met (Iwasaki and Suganuma, 2009). At this time, the implementation of a unified policy for the prohibition of discrimination, including gender discrimination, became possible within the EU, based on the Treaty of

Amsterdam (Sedelmeier, 2009).[4] Among the conditions for accession proposed by the EU side was the incorporation of EU directives into national laws. The directive on combating sex discrimination in the field of labor—which consists of the principle of equal treatment of men and women in terms of wages, employment, vocational training, promotion, and social security; the parental leave system; and the principle of burden of proof in the law of sex discrimination proceedings, etc.—was disseminated into the national laws of the candidate countries with the active support of the EU Commission.

As Křížková and Hašková (2008) have argued—that gender equality would not have been taken seriously in the Czech policy agenda without the external pressure of the EU—the effect of the negotiations with the EU in promoting the formation of institutions to eliminate the gender gap in the candidate countries has been remarkable.

In the area of gender, transnational policy propagation has been reported to be highly effective (True and Mintrom, 2001). In the socialist era, the propagation of feminist ideas from the Western world to the Eastern world was blocked; therefore, as mentioned above, women were taken for granted when it came to family and childcare responsibilities, and they were forced to balance family and work (Razzu, 2018). In this way, the understanding in European socialist countries regarding gender equality fell into a kind of self-righteousness, and the recognition of gender issues in the true sense was actually very weak.

It can be said that external pressure from the EU has strongly encouraged candidate countries to take a new approach to the problem of gender inequality in the labor field. In fact, in these countries, the replacement of EU directives with national laws, as well as the establishment of labor discrimination supervisory bodies, has been realized progressively. At the same time, the awareness of companies and citizens regarding the social roles of men and women was changing significantly.[5] The impact of conditionality on EU accession negotiations has been strong, not only during the negotiation phase but also after accession (Sedelmeier, 2012). These series of policy

[4] Since Article 119 of the Treaty of Rome of the European Community (EC), the predecessor of the EU, referred to gender equality in the field of labor (Bego, 2015), efforts toward gender equality in Europe have gradually gained momentum from the late 1970s to the 1990s.

[5] For example, in a poll conducted in Poland in 1992, 60% of women said that women should stay at home to take care of the children, and even women were positive about the traditional gender role division of labor. However, in a similar survey conducted in 1996, 70.2% of women said that earned labor was the most important thing in their lives, far more than the 2.3% who said housework was the most important. In just three years, there has been a remarkable shift in citizens' values regarding work and housework (Fodor and Balogh, 2012).

measures and social changes could more significantly reduce the gender gap in member countries as compared to nonmember countries. Therefore, we predict the effect of EU member state status on the gender wage gap as follows.

Hypothesis 2: *EU membership and candidate status tend to restrain the gender wage gap in European emerging markets.*

The factors that strongly influence the level of wages between men and women in European emerging markets are not limited to EU accession. Rather, the great wave of democratization and transition to market economy has brought about significant changes in public awareness and business activities in these countries, beyond the differences between EU member and nonmember countries. It is noted that the aforementioned transnational policy propagation, catalyzed by the increased press activities and speech associated with democratization, has had a remarkable effect on public awareness of gender equality (Fodor and Balogh, 2012); however, it is highly likely that the continuing transition to a market economy has also acted to reduce the gender gap in wage levels. This is because intensified market competition may have the effect of forcing companies in the country to pay wages commensurate with workers' productivity. It must also eliminate the unreasonable gender gap in wages (Iwasaki and Ma, 2020). Furthermore, special factors in the European emerging markets enhanced the gender wage equalization effect of the transition to a market economy: the presence of Western multinationals and the integration of local companies into global supply chains.

After the collapse of socialism, companies whose home countries were in the developed world competed to enter the European emerging markets. Through the activities of their local subsidiaries, these multinational corporations broke the rigid wage system of the socialist era and acted as a powerful driving force to create a closer relationship between human capital and wage levels. A number of previous studies have pointed out that these foreign companies paid highly educated and skilled workers higher wages than domestic companies paid, resulting in a widening of the wage gap due to differences in education and skills (Horie and Iwasaki, 2022).

However, foreign companies' emphasis on education and skills can be expected to have the effect of correcting the economically irrational wage gap at the same time. As mentioned above, in European emerging market countries, the educational level of women has improved dramatically throughout the socialist era, and the gender gap in educational investment has not been so large. Moreover, it is also important to note that since the 1990s, the rate of women's enrollment in higher education has increased further in most European emerging market countries, and the gap in educational attainment between women and men has been reduced significantly (European Commission, 2020). In this sense, it is highly likely that the initial conditions favorable to women and the narrowing of the gap in investment in education between men and women during the transition period have gradually reduced the gender gap in wages in European emerging market countries.

In European emerging market countries, the inclusion of local companies in the global supply chain also may have helped to reduce unreasonable sexual discrimination in wages. This is because a series of factors, such as the penetration of Western corporate culture into local companies through their participation in global supply chains, international standardization of work evaluation and wage systems, disclosure of information on management activities, and active CSR activities, are all sure to strongly encourage the equalization of employee treatment in companies of European emerging markets.[6] Based on the above arguments, the authors make the following prediction about the chronological trend of the gender wage gap in European emerging markets.

Hypothesis 3: *There is a decreasing trend of gender wage gap in European emerging markets.*

In the following sections, we examine the above three hypotheses by performing a meta-analysis of the extant literature.

3 Literature Selection Procedure and Overview of Selected Studies

This section describes how we searched for and identified papers to be included in the meta-analysis in this paper. It also provides an overview of the selected literature and the estimates collected from it.[7]

As the first step in searching studies in which estimates of the gender wage gap in European emerging markets are available, we utilized the electronic literature databases of EconLit and Web of Science and accessed the websites of major academic publishers

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[6] According to the mathematical model of Costinot et al. (2021), a global supply chain can have opposite effects on wage standardization between its upper- and lower-tier companies. The fact that these problems are becoming a reality in European emerging markets between multinational corporations of a developed country as manufacturers and their local suppliers has led to the fostering of social dissatisfaction and the rise of populism in the European emerging markets. Although it is beyond the scope of this paper, this problem should not be overlooked.

[7] The literature selection and meta-analysis in this paper were carried out in general conformity with the guidelines described in Havránek et al. (2020).

and international organizations to identify relevant research works. The search covered a period from 1990 to 2021.[8] We conducted an AND search using the term "*wage*" in combination with one of the terms "*emerging markets*," "*transition economies*," "*Central Europe*," "*Eastern Europe*," "*former Soviet Union*," and the name of one of the European emerging markets,[9] obtaining 453 articles. We then inspected each of the collected works and narrowed the literature to those studies that provide estimates of the gender wage gap in European emerging markets as outcomes from regression estimation of a Mincer-type wage function. More specifically, we sought out research works that typically estimated the wage function formulated in the following equation (1) using household-level data and, as a result, reported all information necessary to carry out the meta-analysis in this paper with respect to the coefficient ϑ of the gender dummy variable and its estimation conditions:

$$wage_i = \mu + \gamma \cdot working\ year_i + \delta \cdot working\ year_i^2 + \theta \cdot education\ year_i + \vartheta \cdot gender_i + \sum_{n=1}^{N} \varphi_n \cdot x_n + \varepsilon_{i,} \quad (1)$$

where $wage_i$, $working\ year_i$, $education\ year_i$ and $gender_i$ are wage level (log-transformed in most cases), years of working experience, years of schooling, and gender of the i-th worker, respectively. x_n is the other n-th wage determinant. ε_i is the error term. η is the constant term. γ, δ, θ, ϑ, and φ are parameters to be estimated.

In this sense, the gender wage gap addressed in this paper means a purely female-discriminatory difference in wages detected econometrically after controlling for various factors that may significantly affect wage levels. These factors include working knowledge, education level, job type, occupation, and a series of attributions of the employing firm. Therefore, it should be noted that, in this paper, the gender wage gap never means the observable difference in wages between men and women taken up by the mass media, governments, or international organizations.

As a result of the literature search mentioned above, we selected a total of 53

[8] They include Emerald Insight, Oxford University Press, Sage Journals, ScienceDirect, Springer Link, Taylor & Francis Online, Wiley Online Library, EU, IMF, OECD, and World Bank. The final literature search was conducted in June 2022.

[9] This includes Albania, Belarus, Bosnia and Herzegovina, Bulgaria, Croatia, the Czech Republic/Czechia, Estonia, Hungary, Kosovo, Latvia, Lithuania, (North) Macedonia, Moldova, Poland, Romania, Russia, Serbia, Slovakia, Slovenia, and Ukraine.

research works that target one or more European emerging markets.[10] These 53 selected papers that met the above criteria are useful for testing all three hypotheses described in the previous section. In fact, the selected literature deals with 18 economies in Central and Eastern Europe, therefore, enabling us to estimate the effect size of gender on wage levels across the region to examine Hypothesis 1 and to compare countries with different EU membership statuses from this perspective for testing Hypothesis 2. The selected works also allow us to grasp the long-term time trend of the gender wage gap in European emerging markets by covering the 34-year period from 1984 to 2017 as a whole. Moreover, only eight of 53 studies used panel data, meaning that the overwhelming majority of estimates reported in the selected papers are empirical results of the wage gap between male and female workers in a particular year. These conditions are quite favorable for testing Hypothesis 3.

We extracted a total of 670 estimates from the 53 selected research works. The mean (median) of the number of collected estimates per study is 12.6 (6). All of these represent single-term estimates of gender dummy variables, of which 275 are male dummy variables, while the remaining 395 are female dummy variables.[11] Following the meta-analysis of the gender wage gap in China by Iwasaki and Ma (2020), we use the reversed values of the estimates of male dummy variables together with the estimates of female dummy variables in order to focus on discrimination against women in terms of wage level. In other words, the meta-analysis in this paper investigates how much lower female wages are than male wages in European emerging markets, *ceteris paribus*.

We transformed all 670 collected estimates to partial correlation coefficients (PCCs) in order to adjust the difference in the units of estimation results and with or without logarithmic transformation of the wage variable. The PCC is a unitless statistic that measures the association of a dependent variable and the independent variable in question when other variables are held constant. It ranges between -1.0 and 1.0. When t_k and df_k denote the t value and the degree of freedom of the k-th estimate ($k = 1, ..., K$), respectively, the PCC (r_k) is calculated with the following equation:

[10] **Appendix Table A1** lists these 53 selected studies. Bibliographic information of the papers is provided in the **Supplement**. Although a series of empirical studies, including those of Adamchik and Bedi (2003), Perugini and Selezneva (2015), Goraus et al. (2017), Tyrowicz and Smyk (2019), and Magda et al. (2021), are very useful for understanding the gender wage gap in Central and Eastern Europe, these works could not be included in our meta-analysis due to their reported contents and other technical reasons.

[11] Estimates of interaction terms of a gender dummy variable and other independent variables are not included in the meta-analysis in this paper.

$$r_k = \frac{t_k}{\sqrt{t_k^2 + df_k}}. \quad (2)$$

The standard error (SE_k) of r_k is computed by $\sqrt{(1 - r_k^2)/df_k}$.

As the evaluation criteria for PCCs, a guideline for microeconomic labor research proposed by Doucouliagos (2011) sets 0.048, 0.112, and 0.234 as the lowest thresholds of small, medium, and large effects, respectively (ibid., Table 3, p. 11). In this paper, we evaluate the scale of the gender wage gap in accordance with these criteria.

Table 1 shows the descriptive statistics of the PCCs of the collected estimates, as well as the results of a t-test of means and the Shapiro–Wilk normality test. **Figure 1** draws from the kernel density estimation results. To examine Hypotheses 2 and 3, we computed descriptive statistics and estimated the kernel density by dividing the collected estimates into four different EU membership statuses and three different time periods in addition to those for all studies to test Hypothesis 1.

According to **Table 1**, the mean and median of the estimates extracted from all studies are -0.173 and -0.165, respectively. In addition, Panel (a) of **Figure 1** illustrates their highly skewed distribution toward the negative side. These findings imply that the selected literature as a whole reports a sufficiently large wage gap between men and women in European emerging markets through the transition period in line with Hypothesis 1. Furthermore, the descriptive statistics by period and Panel (c) of **Figure 1** strongly indicate that the gender wage gap tends to diminish as the time period approaches the present, as expected with Hypothesis 3, while those by EU membership status and Panel (b) of **Figure 1** do not show the predicted relationship between different EU membership statuses and the scale of the gender wage gap in accordance with Hypothesis 2.

However, we must interpret the above findings with caution because the simple aggregation of their reported results and an illustration of their distribution may lead us to a false conclusion. In other words, we should synthesize and compare the collected estimates, taking into account their precision and heterogeneity, as well as the possible influence of publication selection bias. The next section reports the results of meta-analysis that properly deals with these issues using advanced techniques.

4 Results

A meta-analysis conventionally consists of three steps: (a) meta-synthesis of collected estimates, (b) meta-regression analysis (MRA) of heterogeneity across studies, and (c) testing for publication selection bias (Stanley and Doucouliagos, 2012; Iwasaki, 2020a). This paper follows this standard procedure.[12] In this section, we report the results obtained from a meta-analysis conducted in accordance with the procedure step by step.

4.1 Meta-Synthesis

As the first step of the meta-analysis, **Table 2** presents the meta-synthesis results. As in **Table 1** and **Figure 1**, **Table 2** shows the results by EU membership status and by period in addition to the synthesis result using all 670 collected estimates.

In Column (b) of **Table 2**, Cochran's Q test of homogeneity consistently rejects the null hypothesis at the 1% significance level, and the I^2 and H^2 statistics also suggest the presence of heterogeneity across studies in all eight cases. Therefore, the synthesized effect sizes of the random-effects model in Column (a) are preferred to those of the fixed-effect model. With respect to the results of the UWA and WAAP estimations in Column (c), in all eight cases, a considerable number of estimates whose statistical power exceeds the threshold of 0.80 are secured. Accordingly, we adopt the WAAP synthesis values, which are more reliable than those of the UWA and the random-effects model. **Figure 2** presents a visual comparison of the adopted synthesized effect sizes.

As shown in Column (c) of **Table 2** and **Figure 2**, the synthesized effect size for all studies is statistically significant at the 1% level and takes a value of -0.098 with the WAAP approach in terms of PCC. According to the Doucouliagos criteria, the gender wage gap in all 18 European emerging markets through the period of 1984–2017 is well above the "small" effect threshold and approaching the "medium" threshold, suggesting that the gender wage discrimination in the region is of economically meaningful severity, as Hypothesis 1 predicts.

A comparison of the synthesis values by EU membership status in 2020 reveals that the gender wage gap in EU member countries reaches almost the same level regardless of the year of accession, while that in non-EU member countries significantly exceeds it. In fact, the wage effect of gender in EU member countries can be rated as small. In contrast, the gender differences have a medium impact on wage levels in countries without EU membership. These results well correspond with Hypothesis 2.

[12] The **Appendix** provides a methodological note regarding the meta-analysis applied in this paper. The methodological description of this meta-analysis is kept to a minimum due to space limitations. For more details, see Borenstein et al. (2009), Stanley and Doucouliagos (2012), and Iwasaki (2020b, Chapter 1).

Further, the synthesized effect sizes computed by period clearly support Hypothesis 3. Actually, the reference value for the period of 1995 or earlier is -0.206, while that for 1996–2005 is -0.171 and for 2006 or later is -0.042. In other words, European emerging markets have experienced a decline in the gender wage gap from a medium scale to an economically insignificant level with the advancing systemic transformation to a market-oriented economy. To examine the reliability of these synthesis results, in **Figure 3**, we look at changes over time in the scale of the gender wage gap through a more detailed subdivision of collected estimates. The slope of the approximate line in the figure is estimated to be positive with statistical significance at the 1% level. Its coefficient implies that, as the average estimation period approaches the present time year by year, the gender wage gap decreases by 0.0066 in terms of the PCC. As demonstrated in **Figure 3**, Hypothesis 3 is verified, even when the estimation period is divided into single-year units.

4.2 Meta-Regression Analysis

The meta-synthesis presented in the previous subsection enables explicit hypothesis testing by providing point estimates as synthesized effect sizes. Nevertheless, it fails to sufficiently consider the influence of heterogeneity across the selected studies on their reported estimates. This subsection, therefore, examines the reliability of synthesis results by estimating a multivariate meta-regression model in which diversity in study conditions and research quality is simultaneously controlled for.

With meta-independent variable x_{kn}, in addition to variables designed to capture differences in EU membership statuses and estimation periods that are keys to hypothesis testing, we also employed a series of moderators in respect to target region, target cohort, target firm ownership, wage payment period, data type, survey data used, estimator, presence of control for selection bias, selection of control variables with potentially significant impact on estimates, and the quality level of the study. The names, definitions, and descriptive statistics of these variables are provided in **Table 3**. As expounded in the **Appendix**, the meta-independent variables are estimated along with the standard errors of the PCCs using five different estimators.[13]

[13] To avoid multicollinearity that may arise from the simultaneous estimation of a large number of independent variables, we have inspected the correlation matrix and variance inflation factor (VIF) of all of the coded variables and, as a result, narrowed down the variables to the 28 listed

Estimation results of Eq. (3) with all moderators from urban region to quality level are reported in **Table 4**. With regard to the key variables for hypothesis testing, the variables of membership candidacy, nonmembership, and average estimation year show robust estimates, while the variable of 2007 and 2013 membership is estimated to be insignificant. These results imply that, first, in comparison with the studies of countries that joined the EU in 2004, those of EU candidate states and non-EU countries tend to find larger gender wage gaps. Actually, the wage effect of gender in EU candidates/non-EU countries is reported to be higher in range of 0.0498–0.0922/0.0614–0.0750 than that in countries with 2004 membership in terms of PCC. Second, if other study conditions are held constant, there is no statistically significant difference in empirical results of gender wage discrimination between the leading and the trailing EU member states. Third, the significant and positive coefficients of average estimation year denote that the gender wage gap in European emerging markets tends to decrease by 0.0031–0.0065 per year through the observation period from 1984 to 2017.

The above findings are reproduced in **Table 5**, which displays estimation results of a meta-regression model with moderators selected through a BMA analysis and an OLS frequentist check.[14] From this table, we confirm that not only the MRA using full moderators but also the MRA that takes model uncertainty into account produced estimates of the key variables that are highly consistent with the meta-synthesis results reported in **Table 2**.

Furthermore, from the estimates of selected moderators in **Table 5**, we found that, first, in European emerging markets, young workers face greater gender wage discrimination than do their older counterparts, *ceteris paribus*. Second, there exists a notable difference in the scale of the gender wage gap between state and private sectors in the sense that state enterprises unreasonably pay male employees more than female ones as compared with the private sector. Third, non-OLS estimation of a Mincer-type wage function with control for location fixed-effects tends to generate a smaller estimate of the wage effect of gender. In addition to the impacts of EU membership and estimation period on the gender wage gap, these findings also serve as insights for deeper understanding of the literature.

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in **Table 3** that fully met the criteria of a correlation coefficient of less than 0.7 and a VIF of less than 10.

[14] See **Appendix Table A2** for the procedure of selecting moderators.

4.3 Test for Publication Selection Bias

As the final step of meta-analysis, this subsection tests for publication selection bias and the presence of genuine evidence in the literature.

Figure 3 illustrates a funnel plot of all 670 collected estimates. The figure shows that the estimates reported in the selected works do not form an ideal distribution from the viewpoint of statistical theory that the shape of the plot must look like an inverted funnel in the absence of publication selection bias. If the true effect is assumed to be zero, as the dotted line in **Figure 3** depicts, the ratio of positive to negative estimates is 32:638; therefore, the null hypothesis that the number of positive estimates equals that of negative ones is strongly rejected by a goodness-of-fit test (z = -23.411, p = 0.000). If the WAAP synthesis value reported in **Table 2** is assumed to be the approximate value of the true effect, as drawn by the solid line in **Figure 3**, the estimates have a ratio of 480:190, with a value of -0.098 being the threshold; therefore, the null hypothesis that the ratio of estimates below the WAAP value versus those over it is 50:50 is again rejected strongly (z = -11.203, p = 0.000). In sum, both the funnel plot and the goodness-of-fit test suggest that there is a high risk of publication selection bias in this research field.

The FAT–PET–PEESE procedure endorses the above findings. In fact, as Panel (a) of **Table 6** shows, the null hypothesis that the intercept γ_0 is zero is rejected by the FAT in four of five models, suggesting a high likelihood of publication selection bias. Even when funnel symmetry is not present, however, the selected studies may contain genuine evidence. Actually, the PET rejects the null hypothesis that the coefficient of the inverse of the standard errors (γ_1) is zero in all five models, meaning that the collected estimates do contain evidence of a nonzero true wage effect of gender. Furthermore, the PEESE approach in Panel (b) of **Table 6** shows that the coefficient (γ_1) is statistically significantly different from zero in four models, indicating that the real scale of the gender wage gap should be in a range from -0.0970 to -0.0513 in terms of PCC.

As pointed out in the **Appendix**, the FAT–PET–PEESE method implicitly assumes a linear relationship between the standard error and publication selection bias, which may not be realistic in the case of this study. For a robustness check, therefore, we performed alternative estimations of publication selection bias–corrected effect size. **Table 7** shows the results. Although the synthesis value varies depending on the applied method, all of the estimates demonstrate the existence of a statistically significant and economically meaningful wage gap between different genders and, accordingly, support

Hypothesis 1.

We also carried out the FAT–PET–PEESE procedure by EU membership and by period. The test results are summarized in **Table 8**, along with those for all of the studies mentioned above. As shown in this table, the publication selection bias was detected by the FAT in the studies of countries with 2004 membership and EU candidate countries as well as the estimates with an average estimation year of 2006 or later. Nevertheless, the PET rejected the null hypothesis of the nonexistence of genuine evidence in all seven cases, and the PEESE approach generated a nonzero publication selection–adjusted effect size in each of these seven cases that is highly compatible with the corresponding WAAP synthesis value reported in **Table 2**. Hence, we judge that the test results of publication selection bias strongly back up both Hypotheses 2 and 3 as the meta-synthesis and the MRA did in the previous subsections.

5 Conclusions

In this paper, we attempted to estimate the scale of the gender wage gap in European emerging markets and compare countries with different EU membership statuses and time periods from this perspective by applying advanced meta-analytic techniques. The meta-synthesis of 670 estimates extracted from 53 selected works exhibited that gender has a statistically significant and economically meaningful impact on wage levels in European emerging markets through the transition period. The synthesis results also revealed that the gender wage gap in countries with EU membership is much lower than that in non-EU member states. Despite of the presence of such regional disparities, we found that the wage gap between men and women has a tendency to diminish over time in the region as a whole. The MRA of literature heterogeneity and the test for publication selection bias strongly support the findings obtained from the meta-synthesis.

As mentioned in the introduction, in European emerging market countries, there were a variety of factors that could strongly influence the level of the gender wage gap. However, the results of the meta-analysis in this paper, summarized above, clearly show that the drastic historical changes that have characterized the period from the 1990s to the present—democratization, marketization, and the return to Europe—have strongly determined the level of and change in the gender wage gap across the region. Furthermore, our results also demonstrate that transition to a democratic capitalist market economy and new membership in the European Union have brought a blessing to the former socialist economies in the sense that they have remarkably eliminated gender inequality in terms of wages. The reform efforts of governments and citizens over the past three decades have certainly paid off.

In lieu of final remarks, we emphasize that, as stated in Section 3, this study dealt with the econometrically detected female-discriminatory difference in wages, not with the difference in average labor incomes between men and women that is often considered a problem in the real world. Hence, the above findings do not necessarily imply a trend toward the elimination of the observed wage gap between male and female workers in European emerging markets. In fact, while the average hourly wage difference between men and women in the EU27 is 13.0 percent in 2020 according to Eurostat, some countries in this region—such as the Czech Republic, Hungary, Estonia, and Latvia—largely exceed that figure (16.4%, 17.2%, 21.1%, and 22.3%, respectively).[15] As repeatedly pointed out in Razzu (2018), this is likely due to various factors of gender differences created by social values and other factors, as some countries in this region have large differences in industrial sectors and job classifications and statuses between men and women. The gender wage gap discussed in this paper is the difference that remains after controlling for these factors, but we do not condone such gender differences. It is necessary to implement policies to eliminate gender discrimination in society as a whole so that the income gap between men and women can be eliminated from all angles.

Appendix Methodology of Meta-Analysis: A Short Guide

In this appendix, we will provide a brief description of the methodology of meta-analysis performed in this paper.

To synthesize PCCs, we use the meta fixed-effect model and the meta random-effects model. According to Cochran's Q test of homogeneity and I^2 and H^2 heterogeneity measures, we adopt the synthesized effect size of one of these two models. In addition to this traditional synthesis method, we also utilize the unrestricted weighted least squares average (UWA) approach proposed by Stanley and Doucouliagos (2017) and Stanley et al. (2017) as a new synthesis method. The UWA is less subject to

[15] Eurostat, "Gender pay gap in unadjusted form": https://ec.europa.eu/eurostat/databrowser/view/tesem180/default/table?lang=en (accessed on December 17, 2022). On the other hand, it is also true that the gender wage gap in Romania, Slovenia, and Poland in this sense is considerably lower than the EU27 average (2.4%, 3.1%, and 4.5%, respectively).

influence from excess heterogeneity than the fixed-effect model. The UWA method regards as the synthesized effect size a point estimate obtained from the regression that takes the standardized effect size as the dependent variable and the estimation precision as the independent variable. Specifically, we estimate Eq. (A1), in which there is no intercept term, and the coefficient, α, is utilized as the synthesized value of the PCCs:

$$t_k = \alpha(1/SE_k) + \varepsilon_k, \quad \text{(A1)}$$

where SE is the standard error of the PCC of the k-th estimate, and ε_k is a residual term. In theory, α in Eq. (A1) is consistent with the estimate of the meta fixed-effect model.

Further, Stanley et al. (2017) proposed conducting a UWA of estimates, the statistical power of which exceeds the threshold of 0.80, and called this estimation method the weighted average of the adequately powered (WAAP). They stated that WAAP synthesis has less publication selection bias than the traditional random-effects model. Accordingly, we adopt the WAAP estimate as the best synthesis value whenever available. Otherwise, the traditional synthesized effect size is used as the second-best reference value.

Following the synthesis of collected estimates, we conduct an MRA to explore the factors causing heterogeneity between selected studies. More concretely, we estimate a meta-regression model:

$$y_k = \beta_0 + \sum_{n=1}^{N-1} \beta_n x_{kn} + \beta_N SE_k + e_k, \quad \text{(A2)}$$

where y_k is the PCC (i.e., r_k) of the k-th estimate, β_0 is the constant, x_{kn} denotes a meta-independent variable (also known as a moderator) that captures the relevant characteristics of an empirical study and explains its systematic variation from other empirical results in the literature, β_n denotes the meta-regression coefficient to be estimated, and e_k is the meta-regression disturbance term.

As pointed out in Iwasaki et al. (2020, 2022), there is no clear consensus among meta-analysts about the best model for estimating Eq. (A2). Hence, to check the statistical robustness of coefficient β_n, we perform an MRA using the following five estimators: (1) the cluster-robust ordinary least squares (OLS) estimator, which clusters the collected estimates by study and computes robust standard errors; (2) the cluster-robust weighted least squares (WLS) estimator, which uses the inverse of the standard error ($1/SE$) as an analytical weight; (3) the multi-level mixed-effects restricted maximum likelihood (RLM) estimator; (4) the cluster-robust random-effects panel generalized least squares (GLS) estimator; and (5) the cluster-robust fixed-effects panel least squares dummy variable (LSDV) estimator. In this paper, we interpret the results based on the assumption that estimates that not only are statistically significant but also have the same sign in at least three of five models constitute statistically robust estimates.

It is argued that MRA involves the issue of model uncertainty, in the sense that the true model cannot be identified in advance. Thus, in addition to a meta-regression model that takes all available moderators on the right-hand side, we also estimate a model with selected moderators following the approach of Bayesian meta-analysts (Bajzik et al., 2020; Havranek and Sokolova, 2020). More concretely, we first estimate the posterior inclusion probability (PIP) of each meta-independent variable other than the variables needed for hypothesis testing and standard error of PCCs, using the Bayesian model averaging (BMA) method that utilizes conventional noninformative priors on the focus parameters and a multivariate Gaussian prior on the auxiliary parameters. Then, we conduct an OLS frequentist check of variables with PIPs of 0.50 or more, adopting a policy of employing variables for which the estimates are statistically significant at a level of 10% or above as selected moderators in Eq. (A2).

As the final stage of meta-analysis, we examine publication selection bias using a funnel plot, by conducting a goodness-of-fit test of proportional distribution, and by performing an MRA test procedure consisting of a funnel-asymmetry test (FAT), a precision-effect test (PET), and a precision-effect estimate with standard error (PEESE), which were proposed by Stanley and Doucouliagos (2012) and have been used widely in previous meta-studies.

A funnel plot is a scatter plot with the effect size (in the case of this paper, the PCC) on the horizontal axis and the precision of the estimate ($1/SE$) on the vertical axis. In the absence of publication selection bias, effect sizes reported by independent studies vary randomly and symmetrically around the true effect size. Moreover, according to the statistical theory, the dispersion of effect sizes is negatively correlated with the precision of the estimate. Therefore, the shape of the plot must look like an inverted funnel. In other words, if the funnel plot is not bilaterally symmetrical but is deflected to one side, then an arbitrary manipulation of the study area in question is suspected, in the sense that estimates in favor of a specific conclusion (i.e., estimates with an expected sign) are more frequently published.

The goodness-of-fit test examines the proportional distribution of the reported estimates. The test is performed based on either the assumption that the true effect size is zero or the assumption that the selected meta-synthesis value approximates the true effect. By conducting this univariate test, we inspect whether the estimates in question are distributed evenly around the true effect size.

The FAT–PET–PEESE procedure has been developed to test publication selection bias and the presence of genuine evidence in a more rigid manner: FAT can be performed by regressing the *t* value of the k-th estimate on $1/SE$ using Eq. (A3), thereby testing the null hypothesis that the intercept term γ_0 is equal to zero:

$$t_k = \gamma_0 + \gamma_1(1/SE_k) + v_k, \quad \text{(A3)}$$

where v_k is the error term. When the intercept term γ_0 is statistically significantly different from zero, we can interpret that the distribution of the effect sizes is asymmetric.

Even if there is publication selection bias, a genuine effect may exist in the available empirical evidence. Stanley and Doucouliagos (2012) proposed examining this possibility by testing the null hypothesis that the coefficient γ_1 is equal to zero in Eq. (A3). The rejection of the null hypothesis implies the presence of genuine empirical evidence. γ_1 is the coefficient of precision; therefore, it is called a PET.

Furthermore, Stanley and Doucouliagos (2012) also stated that an estimate of the publication selection bias–adjusted effect size can be obtained by estimating the following equation (A4), which has no intercept. If the null hypothesis of $\gamma_1 = 0$ is rejected, then the nonzero true effect does actually exist in the literature, and the coefficient γ_1 can be regarded as its estimate.

$$t_k = \gamma_0 SE_k + \gamma_1(1/SE_k) + v_k \quad \text{(A4)}$$

This is the PEESE approach.

To test the robustness of the regression coefficients obtained from the above FAT–PET–PEESE procedure, we estimate Eqs. (A3) and (A4) using not only the unrestricted WLS estimator, but also the WLS estimator with bootstrapped standard errors, the cluster-robust WLS estimator, and the unbalanced panel estimator for a robustness check. In addition to these four models, we also run an instrumental variable (IV) estimation with the inverse of the square root of the number of observations used as an instrument of the standard error, because "the standard error can be endogenous if some method choices affect both the estimate and the standard error. Moreover, the standard error is estimated, which causes attenuation bias in meta-analysis" (Cazachevici et al., 2020, p. 5).

The above FAT–PET–PEESE approach implicitly relies on the assumption that publication selection bias is linearly proportional to the size of the standard error, which might not be practical in some cases (Zigraiova et al., 2021). To deal with the possible nonlinear relationship between the two, some advanced techniques have been developed recently. They include the "Top 10" approach, proposed by Stanley et al. (2010), who discovered that discarding 90% of the published findings greatly reduces publication selection bias and is often more efficient than conventional summary statistics; the selection model, developed by Andrews and Kasy (2019), which tests for publication selection bias using the conditional probability of publication as a function of a study's results; the endogenous kinked model, innovated by Bom and Rachinger (2019), which presents a piecewise linear meta-regression of estimates of their standard errors, with a kink at the cutoff value of the standard error below which publication selection bias is unlikely; and the *p*-uniform method, introduced by van Aert and van Assen (2012), which is grounded on the statistical theory that the distribution of *p*-values is uniform conditional on the population effect size. In this paper, we apply these four techniques to provide alternative estimates of the publication selection bias–corrected effect size and compare them with the selected synthesized values and PEESE estimates for a robustness check.

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Table 1. Descriptive statistics of partial correlation coefficients, *t*-test, and Shapiro–Wilk normality test of collected estimates

Study type	Number of estimates (K)	Mean	Median	S.D.	Max.	Min.	Kurtosis	Skewness	t- test [a]	Shapiro–Wilk normality test (z) [b]
All studies	670	-0.173	-0.165	0.115	0.055	-0.570	2.708	-0.332	-38.893 ***	4.371 †††
By EU membership status										
2004 member	265	-0.186	-0.179	0.117	0.007	-0.445	2.076	-0.193	-25.879 ***	4.185 †††
2007 and 2013 member	85	-0.147	-0.132	0.122	0.024	-0.390	1.893	-0.245	-11.153 ***	2.846 †††
Membership candidate	19	-0.009	0.033	0.086	0.054	-0.183	2.882	-1.278	-0.450	4.108 †††
Nonmember	301	-0.179	-0.165	0.104	0.055	-0.570	3.902	-0.745	-29.884 ***	4.646 †††
By period										
1995 or before	163	-0.233	-0.227	0.117	0.024	-0.445	2.197	0.240	-25.284 ***	2.562 †††
1996–2005	300	-0.181	-0.176	0.103	0.055	-0.570	3.448	-0.295	-30.586 ***	2.666 †††
2006 or later	207	-0.114	-0.102	0.102	0.054	-0.517	4.478	-0.989	-16.009 ***	4.961 †††

Notes: This table shows the number of estimates of the gender dummy variable in the Mincer-type wage function estimated in 53 research works listed in Online Appendix Table A1 and the Supplement. Descriptive statistics of the estimates and the results of t-test and Shapiro-Wilk normality test are also reported for all studies and by EU membership status and estimation period.

[a] ***: Null hypothesis that the mean is zero is rejected at the 1% level

[b] †††: Null hypothesis of normal distribution is rejected at the 1% level.

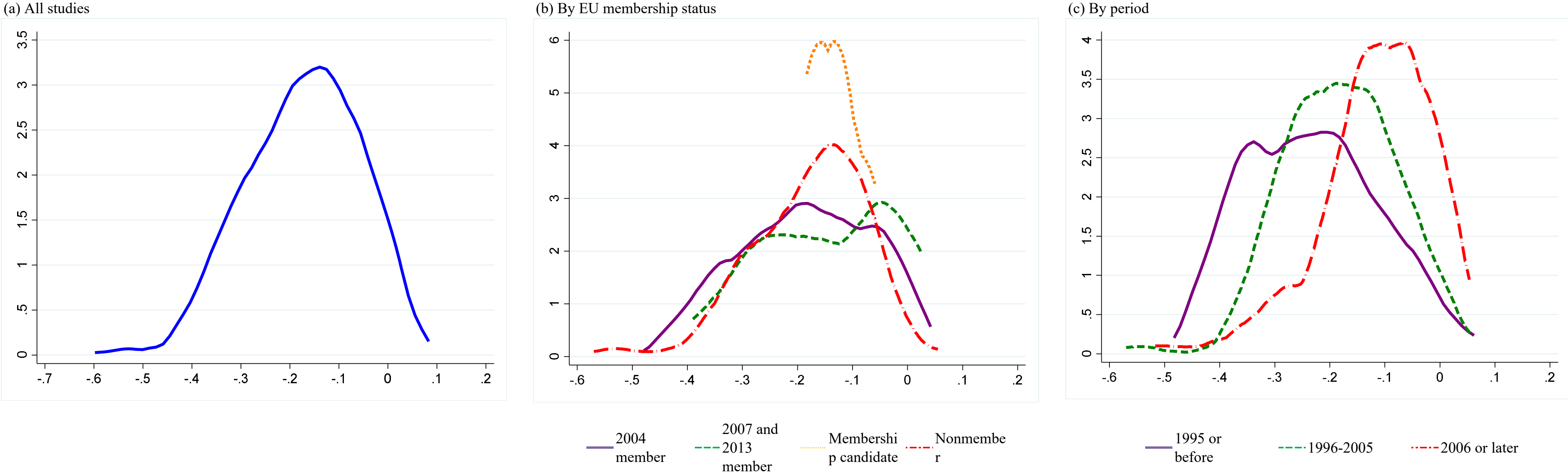

Figure 1. Kernel density estimation of collected estimates

Notes: This figure illustrates Kernel density estimation of 670 estimates of the gender dummy variable included in the meta-analysis. The Epanechnikov kernel function was used for estimation. The vertical axis is the kernel density. The horizontal axis is the partial correlation coefficient of collected estimates. Kernel density estimation was conducted by EU membership status and estimation period, and the results are shown in Panels (b) and (c), respectively, in addition to that for all studies in Panel (a). See Table 1 for the descriptive statistics of estimates.

Table 2. Synthesis of estimates

Study type	Number of estimates (K)	(a) Traditional synthesis: Fixed-effect model (z value)[a]	(a) Traditional synthesis: Random-effects model (z value)[a]	(b) Heterogeneity test and measures: Cochran's Q test of homogeneity (p value)[b]	(b) I^2 statistic[c]	(b) H^2 statistic[d]	(c) Unrestricted weighted least squares average (UWA): UWA of all estimates (t value)[a,e]	(c) Number of the adequately powered estimates[f]	(c) WAAP (weighted average of the adequately powered estimates) (t value)[a]	(c) Median S.E. of estimates	(c) Median statistical power
All studies	670	-0.098 ***	-0.173 ***	540000 ***	99.9	1134.8	-0.098 ***	620	-0.098 ***	0.018	1.000
		(-765.4)	(-38.9)	(0.00)			(-26.9)		(-25.9)		
By EU membership status											
2004 member	265	-0.097 ***	-0.186 ***	510000 ***	100.0	2903.9	-0.097 ***	245	-0.097 ***	0.008	1.000
		(-741.1)	(-25.9)	(0.00)			(-16.9)		(-16.2)		
2007 and 2013 member	85	-0.104 ***	-0.145 ***	11589 ***	99.3	147.5	-0.104 ***	81	-0.104 ***	0.013	1.000
		(-97.9)	(-11.0)	(0.00)			(-8.3)		(-8.1)		
Membership candidate	19	-0.062 ***	-0.009	276 ***	89.7	9.7	-0.062 **	3	-0.176 ***	0.033	0.469
		(-9.7)	(-0.4)	(0.00)			(-2.5)		(-36.4)		
Nonmember	301	-0.190 ***	-0.180 ***	13346 ***	97.1	34.3	-0.190 ***	289	-0.190 ***	0.020	1.000
		(-188.2)	(-29.8)	(0.00)			(-28.2)		(-27.7)		
By period											
1995 or before	163	-0.206 ***	-0.233 ***	14967 ***	99.2	117.5	-0.206 ***	157	-0.206 ***	0.021	1.000
		(-252.9)	(-25.5)	(0.00)			(-26.3)		(-25.8)		
1996–2005	300	-0.171 ***	-0.182 ***	180000 ***	99.9	804.3	-0.171 ***	290	-0.171 ***	0.018	1.000
		(-844.9)	(-30.5)	(0.00)			(-34.9)		(-34.3)		
2006 or later	207	-0.042 ***	-0.113 ***	97341 ***	99.9	1555.3	-0.042 ***	89	-0.042 ***	0.018	0.648
		(-250.2)	(-16.1)	(0.00)			(-11.5)		(-7.6)		

Notes: This table reports meta-synthesis results of estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement. See Table 1 for the descriptive statistics of estimates. The appendix describes the methodology of meta-synthesis applied in this table. *** and ** denote statistical significance at the 1% and 5% levels, respectively.

[a] Null hypothesis: The synthesized effect size is zero.

[b] Null hypothesis: Effect sizes are homogeneous.

[c] Ranges between 0 and 100% with larger scores indicating heterogeneity

[d] Takes zero in the case of homogeneity

[e] Synthesis method advocated by Stanley and Doucouliagos (2017) and Stanley et al. (2017)

[f] Denotes number of estimates with statistical power of 0.80 or more, which is computed by referring to the UWA of all collected estimates

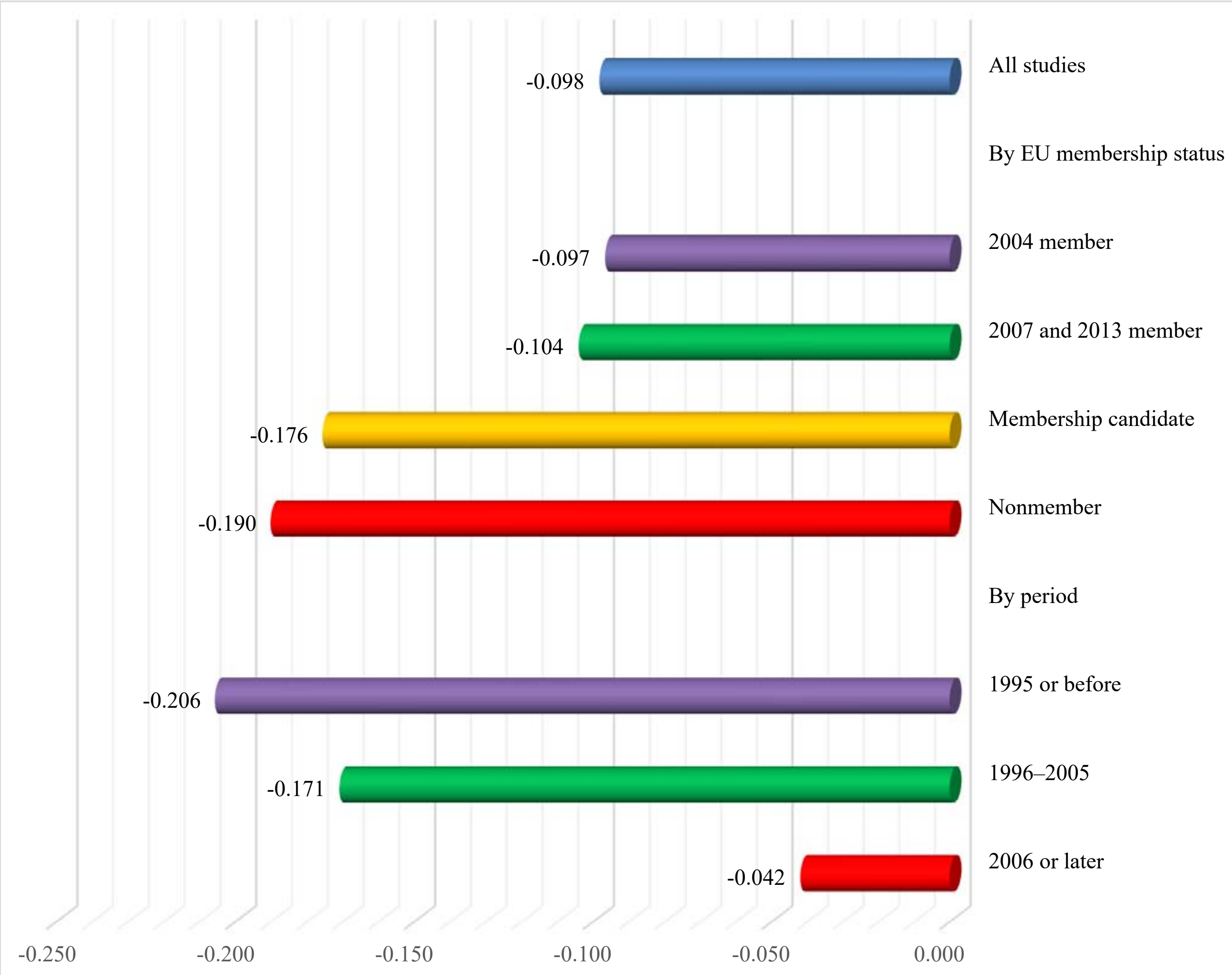

Figure 2. Illustrated comparison of synthesis results

Notes: This figure illustrates the synthesized effect size by WAAP estimation reported in Table 2 using estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement. See Table 1 for the descriptive statistics of estimates.

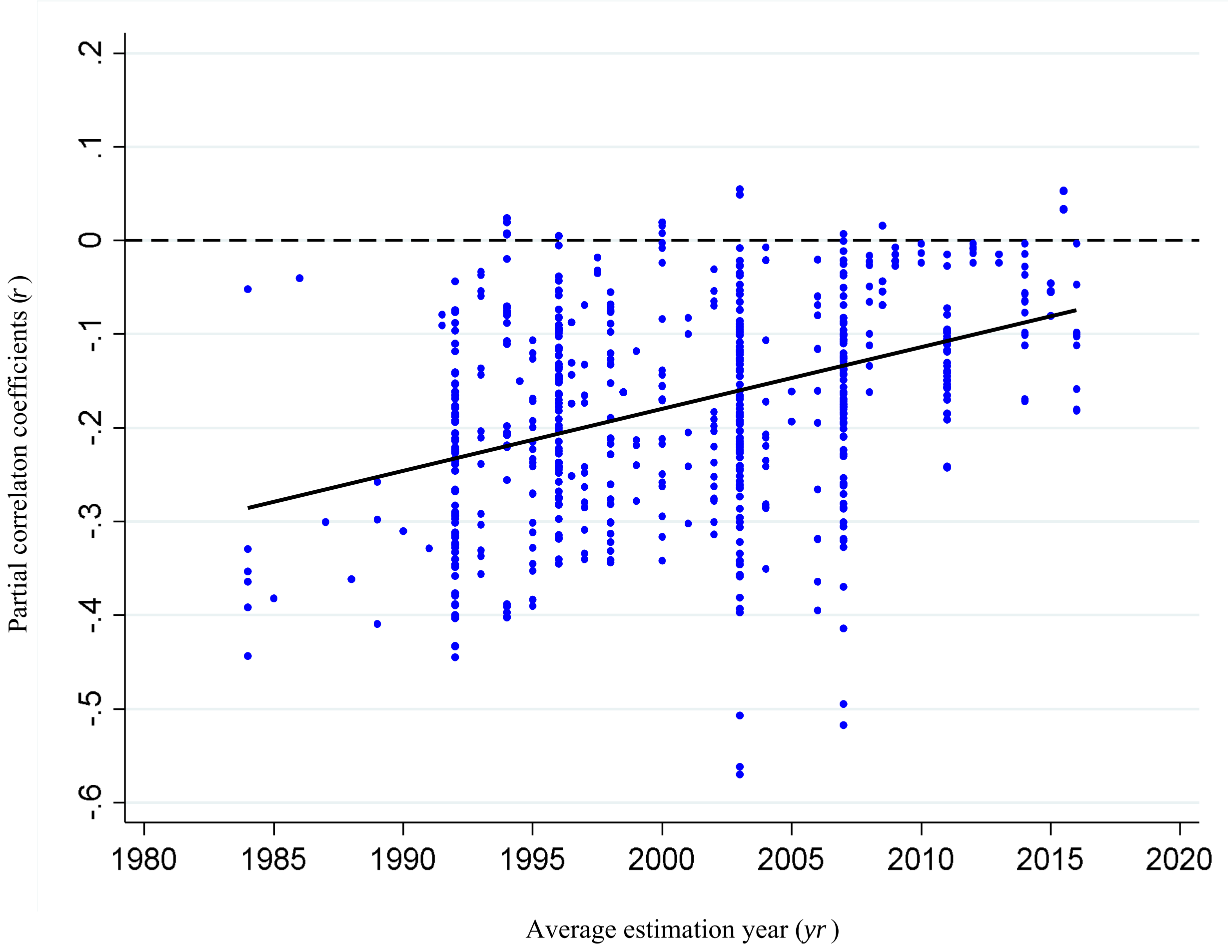

Figure 3. Chronological order of partial correlation coefficients

Notes: This figure shows the chronological order of partial correlation coefficients of estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement. See Table 1 for the descriptive statistics of estimates. The solid black line in the figure displays the fit line. The equation of the fit line is described below. The values in parentheses below the coefficients in the equation are robustness standard errors. *** denotes statistical significance at the 1% level.

$$r = \underset{(1.1291)}{-13.3680^{***}} + \underset{(0.0006)}{0.0066^{***}}\,yr$$

Adj. $R^2 = 0.1685$ $F = 136.55^{***}$

Table 3. Name, definition, and descriptive statistics of meta-independent variables

Variable name	Definition	Descriptive statistics		
		Mean	Median	S.D.
2007 and 2013 member	1 = if the target country joined the EU either in 2007 or 2013, 0 = otherwise	0.127	0	0.333
Membership candidate	1 = if the target country is a candidate for EU membership, 0 = otherwise	0.028	0	0.166
Nonmember	1 = if the target country is a non-EU member state, 0 = otherwise	0.449	0	0.498
Average estimation year	Average estimation year	2001	2002	7.193
Urban region	1 = if the target region is urban, 0 = otherwise	0.110	0	0.314
Rural region	1 = if the target region is rural, 0 = otherwise	0.001	0	0.039
Young age	1 = if the sample is limited to young workers, 0 = otherwise	0.009	0	0.094
Middle age	1 = if the sample is limited to middle-aged workers, 0 = otherwise	0.012	0	0.109
Older age	1 = if the sample is limited to older workers, 0 = otherwise	0.009	0	0.094
State enterprise	1 = if the target firm is a state enterprise, 0 = otherwise	0.049	0	0.217
Private firm	1 = if the target firm is a private firm, 0 = otherwise	0.091	0	0.288
Annual	1 = if annual wage is employed for empirical analysis, 0 = otherwise	0.082	0	0.275
Weekly	1 = if weekly wage is employed for empirical analysis, 0 = otherwise	0.039	0	0.193
Hourly	1 = if hourly wage is employed for empirical analysis, 0 = otherwise	0.467	0	0.499
Panel data	1 = if panel data is employed for empirical analysis, 0 = otherwise	0.067	0	0.250
Original household survey	1 = if the results of an original household survey are used as the data source, 0 = otherwise	0.481	0	0.500
Non-OLS	1 = if an estimator rather than OLS is used for estimation, 0 = otherwise	0.118	0	0.323
IV/2SLS/3SLS	1 = if IV, 2SLS, or 3SLS estimator is used for estimation, 0 = otherwise	0.001	0	0.039
Control for selection bias	1 = if the sample selection bias of employment is controlled for, 0 = otherwise	0.031	0	0.174
Occupation	1 = if the estimation simultaneously controls for occupation, 0 = otherwise	0.351	0	0.478
Health	1 = if the estimation simultaneously controls for the health of workers, 0 = otherwise	0.012	0	0.109
Firm size	1 = if the estimation simultaneously controls for the size of firms to which workers belong, 0 = otherwise	0.354	0	0.478
Trade union	1 = if the estimation simultaneously controls for the presence of trade unions, 0 = otherwise	0.045	0	0.207
Location fixed effects	1 = if the estimation simultaneously controls for location fixed effects, 0 = otherwise	0.427	0	0.495
Industry fixed effects	1 = if the estimation simultaneously controls for industry fixed effects, 0 = otherwise	0.373	0	0.484
Time fixed effects	1 = if the estimation simultaneously controls for time fixed effects, 0 = otherwise	0.051	0	0.220
Quality level	20-point scale of the quality level of the study	16.233	18	3.839
SE	Standard error of partial correlation coefficient	0.020	0.0183	0.017

Note: This table lists meta-independent variables used in regression estimation in Tables 4 and 5 and thier descriptive statistics. We evaluated the quality levels of studies included in our meta-analysis as follows: As the basic source of information for evaluating the quality levels of journal articles, we used the rankings of economic journals published on February 1, 2018 by IDEAS (http://ideas.repec.org/), a bibliographic database dedicated to economics and available freely on the Internet. IDEAS provided the most comprehensive rankings of 2159 economic journals as of February 2018. We conducted a cluster analysis by using the comprehensive evaluation scores provided by IDEAS to divide the 2159 journals into 20 clusters and then graded (weight) each journal on a scale from 20 (a group of journals belonging to the highest cluster) to 1 (a group of journals belonging to the lowest cluster). For journals not indexed by IDEAS, we referred to impact factors (Thomson Reuters) and other journal rankings that allowed us to compare them against the journals indexed by IDEAS, and then graded them according to the scores given to the equivalent journals listed in IDEAS. For academic books or book chapters, we gave an initial score of 1 and upgraded them to a median score of 10 when any of the following conditions were satisfied: (1) it is clearly specified that the book or article in question has been peer reviewed; (2) the book or article in question has been published by a major academic publisher that is assessed by outside experts; and, (3) the quality level of the research in question is clearly high.

Table 4. Meta-regression analysis of literature heterogeneity: Estimation using all moderators

Estimator	Cluster-robust OLS	Cluster-robust WLS [1/*SE*]	Multi-level mixed-effects RML	Cluster-robust random-effects panel GLS	Cluster-robust fixed-effects panel LSDV
Meta-independent variable (default category)/model	[1]	[2]	[3]	[4]	[5]
EU membership status (2004 member)					
2007 and 2013 member	0.0392	0.0729	0.0045	0.0047	-0.0036
	(0.037)	(0.050)	(0.013)	(0.014)	(0.011)
Membership candidate	-0.0498 **	-0.0748 *	0.0408	0.0416	-0.0922 *
	(0.023)	(0.046)	(0.026)	(0.027)	(0.054)
Nonmember	-0.0054	-0.0242	-0.0617 ***	-0.0614 ***	-0.0750 ***
	(0.038)	(0.041)	(0.018)	(0.018)	(0.004)
Estimation period					
Average estimation year	0.0065 ***	0.0057 ***	0.0031 **	0.0032 **	0.0012
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Taget region (region unspecified)					
Urban region	-0.0642	-0.1651 ***	-0.0577 *	-0.0580	-0.0216 ***
	(0.044)	(0.061)	(0.035)	(0.036)	(0.001)
Rural region	0.0402	-0.0387	-0.0508	-0.0503	dropped
	(0.067)	(0.046)	(0.043)	(0.044)	
Target cohort (cohort unspecified)					
Young age	-0.1368 **	-0.0620	0.0094	0.0081	0.0646 ***
	(0.064)	(0.074)	(0.020)	(0.021)	(0.011)
Middle age	-0.0425	0.0507	0.1033 ***	0.1020 ***	0.1585 ***
	(0.064)	(0.074)	(0.020)	(0.021)	(0.011)
Older age	0.0810	0.1740 **	0.2266 ***	0.2254 ***	0.2818 ***
	(0.064)	(0.074)	(0.020)	(0.021)	(0.011)
Target firm ownership (ownership unspecified)					
State enterprise	-0.0571 ***	-0.0175	-0.0375 **	-0.0375 **	-0.0375 **
	(0.021)	(0.027)	(0.018)	(0.018)	(0.017)
Private firm	0.0383 *	0.0150	0.0391 *	0.0391 *	0.0389
	(0.020)	(0.024)	(0.022)	(0.023)	(0.024)
Payment period (monthly)					
Annual	0.0432	0.0394	0.0262	0.0265	dropped
	(0.040)	(0.041)	(0.041)	(0.042)	
Weekly	0.0494 *	-0.0046	-0.0247	-0.0242	dropped
	(0.029)	(0.029)	(0.029)	(0.029)	
Hourly	0.0258	-0.0306	0.0789 ***	0.0787 ***	0.0874 ***
	(0.023)	(0.029)	(0.009)	(0.009)	(0.009)
Data type (cross-sectional data)					
Panel data	0.0360	0.0743	0.0121	0.0125	dropped
	(0.051)	(0.060)	(0.040)	(0.041)	
Survey data used (government statistics)					
Original household survey	0.0018	0.0809 **	0.0177	0.0176	0.0188 ***
	(0.030)	(0.037)	(0.015)	(0.016)	(0.004)
Estimator					
Non-OLS (OLS)	0.0289	0.0078	0.0254 **	0.0255 **	0.0200
	(0.027)	(0.016)	(0.012)	(0.013)	(0.015)
IV/2SLS/3SLS	-0.0324	-0.0890	0.0438	0.0434	dropped
	(0.042)	(0.056)	(0.045)	(0.046)	
Control for selection bias					
Control for selection bias	0.0501	0.0926 ***	0.0670 **	0.0670 **	0.0617
	(0.045)	(0.034)	(0.033)	(0.034)	(0.053)
Selection of control variable					
Occupation	0.0132	0.0135	-0.0271	-0.0269	-0.0363
	(0.026)	(0.025)	(0.021)	(0.022)	(0.023)
Health	-0.0204	0.0022	0.0034	0.0032	dropped
	(0.078)	(0.099)	(0.049)	(0.051)	
Firm size	-0.0374	0.0087	-0.0092	-0.0092	-0.0155
	(0.036)	(0.033)	(0.018)	(0.018)	(0.022)
Trade union	0.0085	0.0239	0.0998	0.0995	0.1121 *
	(0.059)	(0.040)	(0.062)	(0.063)	(0.067)
Location fixed effects	0.0489 **	0.0227	0.0261	0.0261	0.0396 *
	(0.023)	(0.025)	(0.017)	(0.017)	(0.020)
Industry fixed effects	0.0015	-0.0195	0.0322 *	0.0321	0.0370
	(0.039)	(0.037)	(0.019)	(0.020)	(0.023)
Time fixed effects	-0.0051	-0.0566	0.0604 **	0.0598 **	0.0881 ***
	(0.046)	(0.056)	(0.025)	(0.026)	(0.020)
Quality level					
Quality level	-0.0001	-0.0057 **	-0.0031	-0.0031	dropped
	(0.003)	(0.003)	(0.003)	(0.003)	
SE	-0.1479	-1.9418 **	0.5074 *	0.5040 *	0.6273 **
	(0.425)	(0.819)	(0.265)	(0.271)	(0.248)
Intercept	-13.1792 ***	-11.4567 ***	-6.4962 **	-6.5623 **	-2.7206
	(2.227)	(1.871)	(2.730)	(2.778)	(3.304)
K	670	670	670	670	670
R^2	0.325	0.556	-	0.123	0.041

Notes: This table shows estimation results of meta-regression analysis using estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement as the dependent variable. See Table 1 for the descriptive statistics of the estimates. See Table 3 for the definitions and descriptive statistics of meta-independent variables. Figures in parentheses beneath the regression coefficients are robust standard errors. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

Table 5. Meta-regression analysis of literature heterogeneity: Estimation with selected moderators

Estimator	Cluster-robust OLS	Cluster-robust WLS [1/*SE*]	Multi-level mixed-effects RML	Cluster-robust random-effects panel GLS	Cluster-robust fixed-effects panel LSDV
Meta-independent variable (default category)/model	[1]	[2]	[3]	[4]	[5]
EU membership status (2004 member)					
2007 and 2013 member	0.0336	0.0491	0.0273	0.0276	0.0265
	(0.034)	(0.047)	(0.020)	(0.020)	(0.022)
Membership candidate	-0.0149 ***	-0.0133 *	0.0590	0.0672	-0.0205 **
	(0.006)	(0.007)	(0.038)	(0.043)	(0.009)
Nonmember	-0.0109	-0.0095	-0.0940 ***	-0.0872 ***	-0.1284 ***
	(0.035)	(0.038)	(0.028)	(0.032)	(0.004)
Estimation period					
Average estimation year	0.0067 ***	0.0072 ***	0.0052 ***	0.0054 ***	0.0037 *
	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Selected moderators					
Urban region	-0.0494	-0.0597	-0.0480	-0.0482	-0.0206 ***
	(0.036)	(0.036)	(0.033)	(0.033)	(0.002)
Young age	-0.1238 ***	-0.1317 ***	-0.1225 ***	-0.1230 ***	-0.1202 ***
	(0.033)	(0.026)	(0.002)	(0.003)	(0.000)
State enterprise	-0.0534 **	0.0096	-0.0362 *	-0.0359 *	-0.0387 **
	(0.022)	(0.041)	(0.019)	(0.020)	(0.017)
Private firm	0.0384 **	0.0281	0.0409 **	0.0416 **	0.0371
	(0.019)	(0.032)	(0.020)	(0.020)	(0.024)
Non-OLS	0.0380	0.0150	0.0457 ***	0.0468 ***	0.0375 ***
	(0.025)	(0.016)	(0.012)	(0.012)	(0.014)
Firm size	-0.0339 *	-0.0398	0.0025	-0.0001	0.0134
	(0.019)	(0.029)	(0.013)	(0.013)	(0.016)
Location fixed effects	0.0578 ***	0.0326	0.0270 **	0.0295 **	0.0161
	(0.019)	(0.030)	(0.013)	(0.014)	(0.015)
SE	-0.0281	-1.4516	0.3226	0.2912	0.4593
	(0.399)	(0.979)	(0.344)	(0.353)	(0.324)
Intercept	-13.7018 ***	-14.5798 ***	-10.6369 ***	-11.0617 ***	-7.6550 *
	(1.758)	(2.796)	(2.705)	(2.556)	(4.273)
K	670	670	670	670	670
R^2	0.299	0.433	-	0.163	0.050

Notes: This table shows estimation results of meta-regression analysis using estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement as the dependent variable. See Table 1 for the descriptive statistics of the estimates. See Table 3 for the definitions and descriptive statistics of meta-independent variables. Selected moderators denote the meta-independent variables having a PIP of 0.50 or more in the Bayesian model–averaging estimation and a p-value of 0.10 or less in the OLS frequentist check reported in Online Appendix Table A2. Figures in parentheses beneath the regression coefficients are robust standard errors. ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

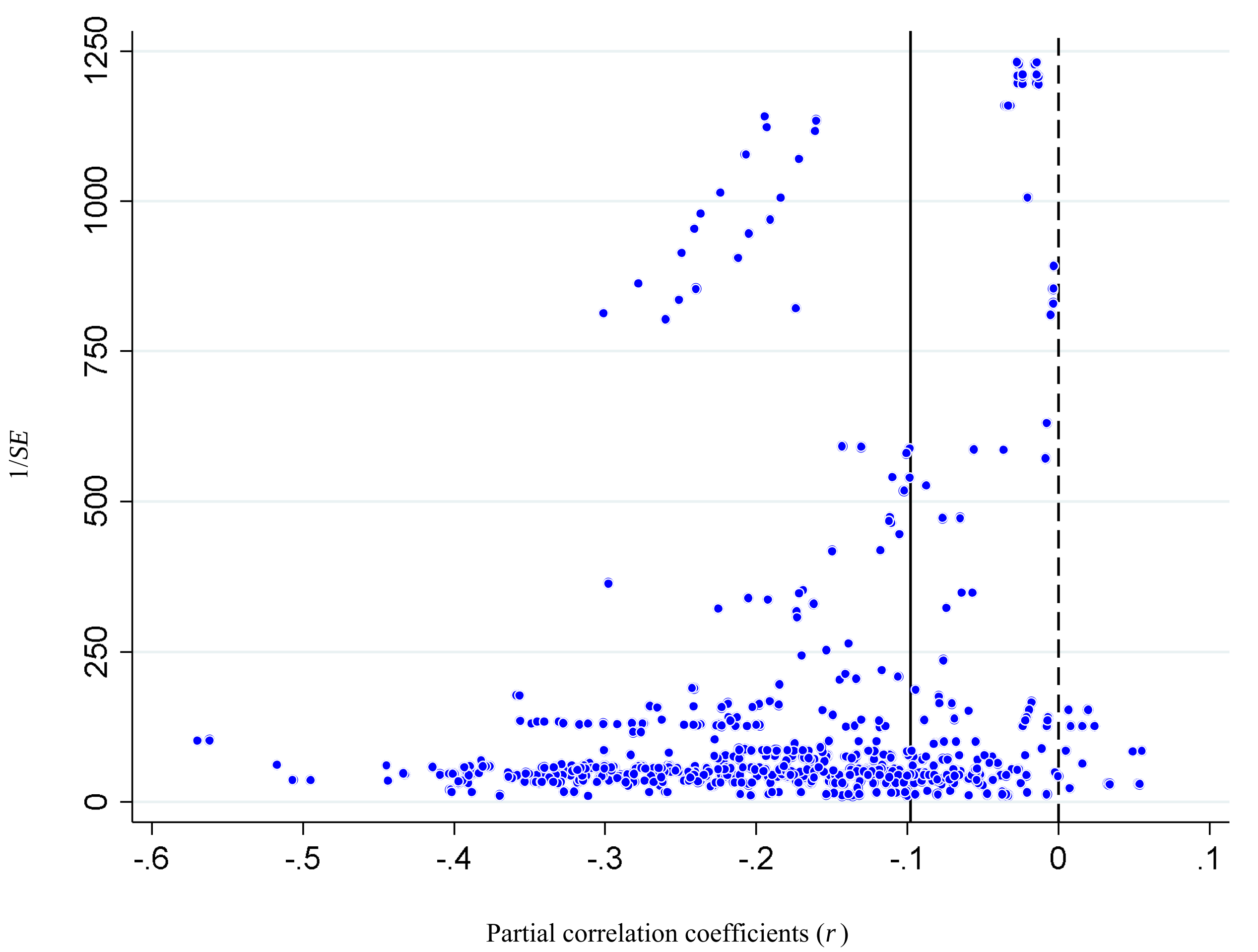

Figure 4. Funnel plot of partial correlation coefficients

Note: This figure displays a funnel plot of partial correlation coefficients of estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement as the dependent variable. See Table 1 for the descriptive statistics of the estimates. The solid line indicates the synthesized effect size by WAAP estimation as reported in Table 2.

Table 6. Meta-regression analysis of publication selection bias: All studies

(a) FAT–PET test (Equation: $t = \gamma_0 + \gamma_1(1/SE) + v$)

Estimator	Unrestricted WLS	WLS with bootstrapped standard errors	Cluster-robust WLS	Cluster-robust random-effects panel GLS	IV
Model	[1]	[2]	[3]	[4][a]	[5]
Intercept (FAT: H_0: $\gamma_0 = 0$)	-6.4514 ***	-6.4514 ***	-6.4514 *	-9.4834	-3.4118 **
	(1.038)	(0.996)	(3.831)	(7.115)	(1.734)
$1/SE$ (PET: H_0: $\gamma_1 = 0$)	-0.0875 ***	-0.0875 ***	-0.0875 *	-0.0569 **	-0.1076 ***
	(0.013)	(0.013)	(0.049)	(0.027)	(0.009)
K	670	670	670	670	670
R^2	0.400	0.399	0.400	0.400	0.379

(b) PEESE approach (Equation: $t = \gamma_0 SE + \gamma_1(1/SE) + v$)

Estimator	Unrestricted WLS	WLS with bootstrapped standard errors	Cluster-robust WLS	Random-effects panel ML	IV
Model	[6]	[7]	[8]	[9]	[10]
SE	-111.5465 ***	-111.5465 ***	-111.5465	165.6966 ***	388.5471
	(18.967)	(16.454)	(68.935)	(31.308)	(494.315)
$1/SE$ (H_0: $\gamma_1 = 0$)	-0.0970 ***	-0.0970 ***	-0.0970 **	-0.0513 ***	-0.4378
	(0.012)	(0.011)	(0.044)	(0.006)	(0.320)
K	670	670	670	670	670
R^2	0.524	0.523	0.524	-	-

Notes: This table reports the test results for publication selection bias in estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement using the FAT–PET–PEESE procudure. See Table 1 for the descriptive statistics of the estimates. The appendix describes the methodology applied in this table. Figures in parentheses beneath the regression coefficients are standard errors. Models [3], [4], and [8] report standard errors clustered by study. Models [5] and [10] use the inverse of the square root of the number of observations used as an instrument of the standard error. *** and ** denote statistical significance at the 1% and 5% levels, respectively.

[a] Hausman test: $\chi^2 = 0.55$, $p = 0.4571$

Table 7. Alternative estimates of publication selection bias–corrected effect size

Method	Top 10[a]	Selection model[b]	Endogeneous kinked model[c]	*p* -uniform[d]
Model	[1]	[2]	[3]	[4]
Publication selection bias–corrected effect size	-0.1034 ***	-0.1490 ***	-0.0875 ***	-0.0968 ***
	(0.011)	(0.010)	(0.004)	(0.001)
K	67	670	670	670

Notes: This table reports the alternative estimation results of the publication selection bias–corrected effect size using estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement as the dependent variable using the FAT–PET–PEESE procudure. See Table 1 for the descriptive statistics of the estimates. The appendix describes the methodology applied in this table. Figures in parentheses are standard errors. *** denotes that the coefficient is statistically significantly different from zero at the 1% level.

[a] Arithmetic average of the top 10% most precise estimates (Stanley et al., 2010)

[b] Test for publication selection bias using the conditional probability of publication as a function of a study's results (Andrews and Kasy, 2019)

[c] Piecewise linear meta-regression of estimates on their standard errors with a kink at the cutoff value of the standard error below which publication selection bias is unlikely (Bom and Rachinger, 2019)

[d] Method based on the statistical theory that the distribution of *p* -values is uniform conditional on the population effect size (van Aert and van Assen, 2021)

Table 8. Summary of publication selection bias test

Study type	Number of estimates (K)	Test results[a]		
		Funnel asymmetry test (FAT) (H_0: $\gamma_0 = 0$)	Precision-effect test (PET) (H_0: $\gamma_1 = 0$)	Precision-effect estimate with standard error (PEESE) (H_0: $\gamma_1 = 0$) [b]
All studies	670	**Rejected**	**Rejected**	**Rejected (-0.0970/-0.0513)**
By EU membership status				
2004 member	265	**Rejected**	**Rejected**	**Rejected (-0.0958/-0.0421)**
2007 and 2013 member	85	Not rejected	**Rejected**	**Rejected (-0.1126/-0.0955)**
Membership candidate	19	**Rejected**	**Rejected**	**Rejected (-0.1597/-0.1217)**
Nonmember	301	Not rejected	**Rejected**	**Rejected (-0.1974/-0.1972)**
By period				
1995 or before	163	Not rejected	**Rejected**	**Rejected (-0.2013/-0.1918)**
1996–2005	300	Not rejected	**Rejected**	**Rejected (-0.1705/-0.0965)**
2006 or later	207	**Rejected**	**Rejected**	**Rejected (-0.0419/-0.0357)**

Notes: This table summarizes the test results for publication selection bias in estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement using the FAT–PET–PEESE procudure by EU membership status and estimation period in addition to the results for all studies reported in Table 6. The appendix describes the methodology applied in this table.

[a] The null hypothesis is rejected when more than three of five models show a statistically significant estimate. Otherwise not rejected.

[b] Figures in parentheses are PSB-adjusted estimates. If two estimates are reported, the left and right figures denote the minimum and maximum estimates, respectively.

Appendix Table A1. List of selected studies on the gender wage gap in European emerging markets for meta-analysis

Author(s) (publication year)	Target country	Estimation period	Number of collected estimates
Flanagan (1995)	Czech Republic	1994 - 1994	11
Rutkowski (1996)	Poland	1987 - 1993	5
Grogan (1997)	Russia	1994 - 1994	1
Bedi (1998)	Poland	1996 - 1996	2
Noorkôiv et al. (1998)	Estonia	1989 - 1995	5
Reilly (1999)	Russia	1992 - 1996	24
Gustafsson et al. (2001)	Russia	1989 - 1989	1
Lehmann and Wadsworth (2001)	Russia	1994 - 1998	6
Newell and Reilly (2001)	CEE8	1984 - 1996	136
Jolliffe (2002)	Bulgaria	1995 - 1995	1
Puhani (2002)	Poland	1994 - 1998	6
Guariglia and Kim (2003)	Russia	1994 - 1998	1
Delteil et al. (2004)	Hungary	1989 - 1998	10
Andrén et al. (2005)	Romania	1990 - 2000	16
Co et al. (2005)	Hungary	1993 - 1993	10
Goh and Javorcik (2005)	Poland	1994 - 2001	1
Gorodnichenko and Sabirianova (2005)	Ukraine, Russia	1985 - 2002	20
World Bank (2005)	Ukraine	2003 - 2004	24
Brown et al. (2006)	Ukraine	1997 - 2003	5
Ogloblin and Brock (2006)	Russia	2002 - 2004	1
Pastore and Verashchagina (2006)	Belarus	1996 - 2001	4
Earle and Telegdy (2007)	Hungary	1992 - 2003	3
Kazakova (2007)	Russia	1996 - 2002	20
Myck et al. (2007)	Poland	1996 - 1996	1
Cattaneo (2008)	Albania	2002 - 2002	1
Csengödi et al. (2008)	Hungary	1992 - 2001	5
Dohmen et al. (2008)	Russia	1997 - 2002	10
Yamaguchi (2008)	Poland	1995 - 2002	4
Jackson and Mach (2009)	Poland	1988 - 1998	3
Nestić (2010)	Croatia	1998 - 2008	12
Bouton et al. (2011)	Moldova	2006 - 2006	5
Eriksson and Pytlikova (2011)	Czech Republic	2006 - 2006	6
Hölscher et al. (2011)	CEE10	2007 - 2007	28
Kovacheva (2011)	Bulgaria	1995 - 2003	10
Andrén (2012)	Romania	1994 - 2000	16
Pignatti (2012)	Ukraine	2003 - 2007	100
Voinea and Mihaescu (2012)	Romania	2004 - 2009	8
Eriksson et al. (2013)	Czech Republic	1998 - 2006	18
Vodopivec (2014)	Slovenia	1992 - 2001	2
Gustafsson et al. (2015)	Russia	2003 - 2003	22
Tiwari et al. (2015)	Russia	2006 - 2010	1
Balcar and Gottvald (2016)	Czech Republic	2008 - 2014	14
Grotkowska et al. (2018)	Poland	2012 - 2012	4
Vilerts (2018)	Latvia	2015 - 2015	8
de Silva and Kupets (2019)	Serbia	2015 - 2016	14
Vahter and Masso (2019)	Estonia	2011 - 2011	28
Vasilescu and Begu (2019)	Romania	2016 - 2016	1
Grabowski and Korczak (2020)	Poland	2010 - 2016	4
Lehmann et al. (2020)	Latvia	2007 - 2015	3
Rudakov and Prakhov (2020)	Russia	2015 - 2017	4
Tovar-Garcia (2020)	Russia	2000 - 2017	10
Krstić (2021)	Croatia, Slovenia, Serbia	2016 - 2016	3
Madga and Salach (2021)	Poland	2014 - 2014	12

Note: See **Online Supplement** for bibliographic information on the listed articles.

Appendix Table A2. Meta-regression analysis of model uncertainty for selection of moderators

Estimator	Bayesian model averaging				OLS frequentist check			
Meta-independent variables/model	[1]				[2]			
	Coef.	S.E.	*t*	PIP	Coef.	S.E.	*t*	*p*
Focus regressors								
2007 and 2013 member	0.0267	0.0137	1.95	1.00	0.0336	0.0128	2.62	0.01
Membership candidate	0.1378	0.0303	4.55	1.00	0.1491	0.0275	5.42	0.00
Nonmember	-0.0097	0.0113	-0.86	1.00	0.0109	0.0095	1.15	0.25
Average estimation year	0.0067	0.0008	8.78	1.00	0.0067	0.0006	10.83	0.00
SE	-0.0621	0.2639	-0.24	1.00	-0.0281	0.2499	-0.11	0.91
Auxiliary regressors								
Urban region	-0.0491	0.0263	-1.86	0.86	-0.0494	0.0148	-3.35	0.00
Rural region	0.0014	0.0210	0.07	0.04				
Young age	-0.0905	0.0700	-1.29	0.70	-0.1238	0.0420	-2.95	0.00
Middle age	-0.0033	0.0170	-0.19	0.07				
Older age	0.0518	0.0633	0.82	0.47				
State enterprise	-0.0416	0.0302	-1.38	0.73	-0.0534	0.0182	-2.93	0.00
Private firm	0.0299	0.0248	1.21	0.67	0.0384	0.0145	2.64	0.01
Annual	0.0031	0.0106	0.29	0.12				
Weekly	0.0045	0.0141	0.32	0.13				
Hourly	0.0035	0.0080	0.44	0.21				
Panel data	0.0020	0.0098	0.20	0.08				
Original household survey	0.0011	0.0052	0.22	0.08				
Non-OLS	0.0268	0.0211	1.27	0.69	0.0380	0.0122	3.12	0.00
IV/2SLS/3SLS	-0.0002	0.0196	-0.01	0.04				
Control for selection bias	0.0168	0.0318	0.53	0.27				
Occupation	0.0003	0.0029	0.10	0.05				
Health	-0.0010	0.0090	-0.11	0.05				
Firm size	-0.0271	0.0163	-1.67	0.81	-0.0339	0.0097	-3.48	0.00
Trade union	-0.0005	0.0051	-0.10	0.05				
Location fixed effects	0.0537	0.0112	4.81	1.00	0.0578	0.0093	6.20	0.00
Industry fixed effects	0.0002	0.0042	0.04	0.06				
Time fixed effects	0.0002	0.0043	0.05	0.04				
Quality level	-0.0001	0.0004	-0.18	0.06				
K	670				670			

Notes: This table reports the results of Bayesian model averaging and OLS frequentist check using estimates of the gender dummy variable reported in 53 research works listed in Online Appendix Table A1 and the Supplement as the dependent variable for the selection of moderators to deal with the issue of model uncertainty in meta-regression analysis. See Table 1 for the descriptive statistics of the estimates. See Table 3 for the definitions and descriptive statistics of meta-independent variables. The appendix describes the methodology applied in this table. Estimate of the intercept is omitted. S.E. and PIP denote standard errors and posterior inclusion probability, respectively. In Model [1], the variables of target countries, average estimation year, and standard errors of partial correlation coefficients are included in the estimation as focus regressors. Therefore, the PIP of these key variables is 1.00.

Supplement

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