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Labour Market Institutions, Crisis and Gender Earnings Gap in Eastern Europe

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Abstract

This paper studies gender earnings inequality in ten Central and Eastern EU countries before (2007) and during the ongoing crisis (2009), using quantile regression methods. The analysis reveals remarkable cross-country diversity in levels and patterns of the gender gap along the earning distribution. We address then the role played by country-specific labour market institutions in forming this variety. Labour market deregulation increases gender inequality, particularly reinforcing the glass-ceiling effect. Higher union density and wage coordination reduce the pay gap, with stronger equalizing effects again in the better-paid jobs. Lastly, the crisis seems to further weaken the already poor role of institutions in the low-pay sector.

JEL Classification: J16, J31, P50
Keywords: earnings gender gap, institutions, quantile regression

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1 Introduction

The emphasis on economic and social equality was a hallmark of the socialist ideology. Countries of Central and Eastern Europe (CEE) and the Soviet Union were actually able to maintain remarkably equal distributions of income under central planning and were often identified as the most equal countries in the world (Atkinson and Micklewright, 1992). Yet, remarkable forms of disparities in living standards – not associated to monetary flows or property rights, and thus invisible to statistics – certainly existed and often reflected the position of individuals in the political sphere (Milanovic, 1998). Inequality between genders was flourishing despite the equality of men and women proclaimed as one of the key ideological tenets of socialism, deeply rooted in the thinking of the founding fathers and emphasized as a key achievement of overcoming capitalism which, by nature, favoured women's oppression (see, for example, Friedrich Engels in his 1884 book, *The Origin of the Family, Private Property and the State*). While the relatively low employment and wage gaps were a fruit of labour market participation policies on one side and generalised wage compression on the other, horizontal and vertical gender segmentation penetrated all fields of social life (Jurajda, 2003 and 2005). The market-oriented reforms undertaken over the last two decades allowed existing visible and hidden inequalities to develop, and new ones, associated to restructuring and vast structural change, to unfold. In the 90s, distributional patterns in Central and Eastern Europe and former-Soviet Union were evolving at a quite different pace, with inequalities reaching (and in some cases stabilising at) diversified levels after twenty years of transition (Aristei and Perugini, 2012). Particularly, visible gender disparities and their evolution played a crucial role in the process.

This paper aims at: (i) providing comparative analyses of gender earnings inequalities – over the actual pay distributions – in Central and Eastern European countries in the context of the ongoing crisis; (ii) assessing the role labour market institutions play in shaping the variety of the gender gaps. To these aims, we first provide a review of the major relevant theoretical and empirical literature (section 2), which among all shows that the most recent evidence on the countries of interest mainly dates back to the mid-2000s and that comparative studies are scanty. We fill the research gap with an analysis covering all ten CEE members of the EU for 2007 and 2009. We explore the latest available information (from EU-Silc) also in view of the effects produced by the outburst of the global crisis and the policy responses emerged (Glassner and Keune, 2012). In section 3 we present the datasets used and some descriptive evidence. Section 4 explains the empirical methods, centred upon quantile regressions, which allow to investigate the unexplained gender gap over the earnings distribution, and to signal for the presence, which is of the particular interest, of sticky floors and/or glass ceiling effects. Additionally, we examine explicitly the effects of three labour market institutional aspects (deregulation, union density and wage coordination) on earnings gap and we investigate if these effects are diversified for low, middle or top income recipients. Up to our knowledge, it is the first time in the literature, when the quintile regression approach is applied to a large sample of Eastern European countries in order to do so. A further distinctive feature of our analysis is the consideration of self-employed earnings in addition to wages. Despite the empirical difficulties that this entails, we believe that the inclusion of this segment of employment into analysis provides important information for the contexts, like those under scrutiny, in which self-employment (particularly in subsistence agriculture and small trade sectors) traditionally represented the only alternative to unemployment and to employment with very low wages (Earle et al., 1994). As the literature recommends, we control for potential self-selection bias in all model specifications. In section 5 we discuss results, which show: (a) a remarkable cross-country heterogeneity in gender gaps ranges and in their changes in response to the crisis, (b) a strong variability of the gap size across the distributions, and (c) a significantly different impact of the three institutional settings on the gender gaps across the earnings range. Section 6 provides a summary of outcomes and concludes.

2 A review of the relevant literature

Gender wage inequality studies in Central and Eastern Europe trace back to the end of the communist regime times, when equality of men and women was proclaimed at governmental level. Women received access to education, healthcare and political life, but in return bared a triple burden of paid employment, unpaid housework, and social/political activities (called "pseudo emancipation" in La Font, 2001). Men dominated in top occupations and gender earnings gap existed in all communist countries, although its size was restricted due to the low wage dispersion (Jurajda, 2003 and 2005).

The development of gender labour income differences (often measured in terms of gender wage gap – GWG) in the first decade of transition was not homogeneous – although it was generally decreasing until mid-1990s – across the CEE countries (Newell and Reilly, 2001; Brainerd, 2000). Regardless the improvement of female-male ratio of (monthly) earnings, feminization of poverty was observed both due to segregation of women into low-paid female dominated occupations and discrimination in local hiring practices (Jurajda and Harmgart, 2007; Orazem and Vodopivec, 2000). During the communist times legally established social entitlements, such as parental leave (gender neutral only since recently) and employment return guarantee (see La Fonte, 2001, for Lithuania and Bulgaria), augmented the costs associated with female labour force compared to men. GWG decomposition studies of that period were attributing more than one third of the gap to the gender segregation by occupation and industry in majority of transitional countries[1]; differences in the quality of labour of the two genders, although not negligible, explained only a relatively smaller portion of the gap (Jurajda, 2003 and 2005; Myslikova, 2012).

During the second decennium of transition, a wide range of GWG levels – from 0.067 in Lithuania to 0.313 in Slovakia in 2002 (Simon, 2012) – was documented, as well as both improvement (for example in Hungary and Poland) and reversal (in Czech Republic and Slovak Republic) of the gender wage equalization process. The harmonization of anti-discriminatory legislation across the countries candidates for the EU-admission brought little change, as the process was not reinforced by practice (see

[1] Make note that the case of the Eastern Germany differed in many ways from other transitional countries, for example, in terms of restructuring process and higher wages in predominantly female occupations in the beginning of transition (see among others, Jurajda, 2005; Hunt, 2002; Jurajda and Harmgart, 2007).

Jurajda, 2005, on Slovak Republic). However, the first signs of the improvement – in terms of significant returns on women's individual characteristics, such as education and age – started to appear (see Simon, 2012, on Latvia and Lithuania in 2002).

Studies of the gender wage inequality in the countries of Central and Eastern Europe rarely go beyond mid-2000s, with rare exceptions being, for example, Myslikova (2012), Rigler and Vanicsek (2008) and Christofides et al. (2010). The first study, using EU-silc data for 2008, emphasizes that at the outset of the financial crisis a remarkable heterogeneity across the old and new EU countries persisted, both in terms of gender wage and employment gaps. While the observed GWG accounted for 22.6% in Czech Republic, 18.4% in Slovakia, 8.9% in Hungary and 8.6% in Poland (Myslikova, 2012), the official European figures for EU27 reported 17.6% as the EU average unadjusted gap (Ponzellini et al., 2010). The second study, focused on Hungary in 2006 and 2007, estimates the gender pay gap to be as high as 17.7%; higher education and longer job experience are found to strengthen women's disadvantage in the labour market. Christofides et al. (2010), on a sample of 24 European Union member states in 2007, concluded that the glass ceiling effect might be observed in majority of the states (including Estonia, Hungary and Poland), while the sticky floor effect can be found only in some of them (for example Slovenia). No evidence of either of the effects is found in Latvia and Lithuania.

As regards the country level drivers of gender earning differences, of interest here, recent comparative studies have underlined the role of institutions in shaping gender inequality/GWG, although the evidence is not conclusive (see, Blau and Khan, 2003; Pastore and Verashchagina, 2011; Heinze and Wolf, 2010). The impact of alternative institutional settings may differ depending on the labour market outcome (Checchi and García-Peñalosa, 2008; Koninger et al., 2007) or on the employment scheme (full-time vs. part-time) considered (McGuinness et al., 2011), as well as on the position of individuals along the wage distribution. The effect might be also both direct – through negotiations over pay – or indirect, – through negotiations over working conditions (Ponzellini et al., 2010).

Since the pioneering contribution by Freeman and Katz (1995), the main debate is mainly focused on GWG in the presence of stronger wage-setting institutions[2] which are expected to compress the wage structure, reduce differences within and across sectors and firms and inhibit discriminatory practices. Card (1992) and Freeman (1993) argued for example that the de-unionisation process was responsible for a remarkable part of the increase of wage inequality in the 80s in the USA. Dustmann et al. (2009) provided similar evidence for Germany. However, (de)unionization dynamics can produce heterogeneous effects across the distribution (Di Nardo et al., 1996). As Firpo et al. (2010) interpret their empirical evidence, stronger unions may correspond to between-group (unionized/non-unionized workers) effects prevailing over within-group ones at the low-paid segment, leading to an increase in inequality. The opposite may hold at the top of the distribution, where the within-group effect of unions dominates. These effects may render the impact of unions on gender wage gap ambiguous, due to different unionization rates across male/female dominated sectors. In the framework of insider/outsiders models, unions can even reinforce the existence of dual segments in the labour market (incumbent/new-hire or temporary/permanent workers), in which the gender distribution is not random. Additionally, the empirical literature shows that the union wage effect explains a substantial proportion of the observed wage gap between union and non-union workers for men but not for women (Cai and Liu, 2008). This may be also due to the fact that men and women tend to pursue different goals in bargaining, with the latter more likely to struggle for an improvement of working conditions rather than for the pay. Under such circumstances, GWG paradoxically increases in case of high proportion of female union members (Heinze and Wolf, 2010).

The literature has also provided relevant evidence on gender gap reducing effects of minimum wage provisions and collective agreements (see Lee, 1999; Blau and Kahn, 2003; McGuinness et al., 2011). In particular, the decline in minimum wage was blamed for the increase in inequality at the bottom part of the distribution (Autor et al., 2010), especially by pushing downward remunerations of the weakest segments of workers, that is low educated women and youth (Di Nardo et al., 1996), thus also increasing the gender gap. On the side of collective agreements, stronger wage-setting centralisation and coordination, which tend

[2] The role of family support schemes, such as parental leave and child-care provisions (Arulampalam et al., 2007; Christofides et al., 2010), income policies and specific anti-discrimination legislative provisions (see Zabalza and Tzannatos, 1985) are out of the focus of this study and are not discussed in detail.

to reduce inter-firm and inter-industry wage variation, may indirectly reduce the gender pay gap if it is associated to inter-firm or inter-industry wage differences. However, the weaker tradition of collective bargaining practices in female-dominated private sectors, such as trade (Rubery et al., 2005), compared to male-dominated sectors, may paradoxically result in an increase of the gender gap (see Ponzellini et al., 2010, for Hungary).

Lastly, labour market deregulation patterns also contribute to shaping gender differences. On the side of quantity, more stringent employment protection legislation (EPL) has been found to impact negatively especially on female employment levels (Kahn, 2007; Bertola et al., 2007). Generally speaking, on the side of wages the impact of EPL depends on the bargaining strength of workers vis-à-vis employers, which is related on the position held by a worker in the labour market, her/his characteristics and the aggregate labour market conditions shaping the outside options (Leonardi and Pica, 2012). If the distribution of workers by gender in the groups with different bargaining strength is not random, a change in EPL may contribute to re-shaping of the gender wage gap, especially in the presence of asymmetries in EPL for different segments of workers that may favour new dual labour market structures (Boeri and Garibaldi, 2007; Belot et al., 2007). For example, in case of substantial firing and hiring costs for permanent contracts and low protection for term positions, firms will prefer placing new entrants into temporary jobs: since new entrants often include a significant share of women, deregulation of temporary work may lead to a higher incidence of temporary employment among women (Kahn, 2007), to an expansion of the gender experience/informal skills gap and, ultimately, to an increase of the wage gap (Ponzellini et al., 2010).

Studies of the gender pay inequality/institutions covering Central and Eastern European countries are very limited. Among them, Simon (2012) finds that “the gender wage gap is not significantly correlated across countries either with the minimum wage or with the collective bargaining coverage rate” (p. 1996). Rigler and Vanicsek (2008), for Hungary in 2006/2007, argue for the presence of discrimination in case of collective bargaining agreement, which is more usual for the firms/organizations dominated by men; trade union membership seems to have little impact on the wage differential. Christofides et al. (2010), using quantile regression approaches on a EU-27 sample for 2007, conclude that “unionism appears to be associated with reductions in the wage gap at the center of wage distribution”.

3 Raw gender earnings gap and distributions in Eastern Europe before and during the crisis

3.1 Data

The datasets used for the empirical analysis are the 2008 and 2010 releases of the EU-Silc (European Union Statistics on Income and Living Conditions) cross-section samples, containing data for reference years 2007 and 2009, respectively. 2009 is the most recent reference year available at the beginning of the study and the use of 2007 enable us to analyse inequality levels, and its drivers, before and after the outburst of the global crisis for all ten central and eastern EU members. We focus our attention on individuals aged between 16 and 65, not in education and not retired. The two samples are composed of 102,960 (2007) and 116,907 (2009) individuals. Of them, 73,354 and 70,608, respectively, are employed; the remaining ones, not in employment, are used in the estimates to account and correct for sample selection bias.

We include in the analysis incomes for both employees (permanent and temporary) and self-employed. The measure of employees income (variable PY010G) is defined as the gross total (yearly) remuneration, in cash or in kind, payable by an employer to an employee in return for work done in the reference period. It includes wages and salaries paid in cash, holiday payments, thirteenth month payment, overtime payment, profit sharing, bonuses and productivity premia, allowances paid for transport or for working in remote locations, as well as the social contributions and income taxes payable by employees. The use of gross wages is common in the literature which considers within-countries wage and earnings inequality (Antonczyk et al., 2010). Brandolini et al. (2011) explain in detail why the use of gross wages is, in fact, the only alternative when EU-Silc data are concerned. Earnings from self-employment (PY050G-Gross Cash Benefits and Losses from Self-Employment and PY070G-Value of goods produced for own consumption) are defined as the income received in the reference period, as a result of current or former involvement in jobs where the remuneration is directly dependent upon the profits derived from the goods and services produced.

To avoid dis-homogeneities in cross-individuals earnings comparison due to different hours of work, we computed all earning measures on a hourly basis. This is done using the information on the number of hours usually worked per week in the main job and the number of months spent at full-time and part-time work. The only assumption needed is that all employed individuals work four weeks per month. Within the group of those in employment, we have trimmed 1% of lower and top hourly wage and self-employment earnings at country level. All monetary variables are expressed in 2005 Euro PPPs. The use of hourly earnings implies studying only one component of annual earnings inequality, the "price effects", while setting aside the "quantity effect" (hours worked per year). A variance decomposition of annual gross earning into the two effects, based on the methodology by Blau and Kahn (2009)[3], reveals that the price effect plays the lion's share on earnings variability, accounting for over 70% of it in the whole sample in both 2007 and 2009. The comparison between annual and hourly earnings inequality is consistent with previous evidence (OECD, 2011) and shows that the former tends to be higher in most of the countries both in 2007 and 2009 due to the presence of people working part-time or part-year and/or to the positive slope of labour supply (Sila, 2012).

EU-Silc data allow considering in the analysis of gender earnings gap and of participation into employment a large set of information referred to households and to individuals. It includes the number of household members, their age profile, the household localization (urban/non-urban region); individual age, marital status, level of education (primary, secondary and tertiary), self-reported health status (on a *1–very good* to *5–very bad* scale), employment status (temporary/permanent employee, self-employed), presence of second job, type of occupation, sector and size of the firm in which the individual is employed[4].

As for the analysis of the impact of labour market institutional variables, we consider the following three indicators: (i) *Hiring regulations and minimum wage* from the Fraser Institute's Economic Freedom World (EFW) database which is essentially a summary indicator of labour market deregulation, particularly on the side of temporary jobs; (ii) *Union density* from the Visser Institute for Advanced Labour Studies database (ver-

[3] The decomposition simply reads: Var(ln AE) = Var (ln hw) + Var (ln ah) + 2 Cov(ln hw, ln ah), where AE, hw and ah stand for annual earnings, hourly wage and annual hours, respectively.

[4] The three levels of education correspond to ISCED classification levels 0–2, 3–4, and 5–6, respectively. Occupations are classified into six categories: 1. Managers & Senior Officials, 2. Professional & Technicians, 3. Clerks, 4. Skilled agricultural & Craft workers, 5. Machine Operators, 6. Elementary Occupations. Industry breakdown has been limited to 8 sectors: 1. Agriculture, 2. Industry, 3. Construction, 4. Trade, 5. Transports, 6. Hotels & Restaurants, 7. Business services, 8. Other services. Lastly, we consider three firm size classes: 0–10, 11–49, 50 and over employees.

sion 2, January 2009); (iii) the level of *Coordination of wage bargaining* from the Visser Institute[5]. We do not use the widely employed OECD labour market institutions indicators, namely the EPL-Employment Protection Legislation for temporary jobs index, as it would imply dropping various countries due to data unavailability (Bulgaria, Romania, Latvia and Lithuania). However, using EPL on the subsample of countries for which it is available provides results strongly consistent with the indicator (i). This is consistent with the evidence in Perugini and Pompei (2012) on western European countries.

Despite fragmentary data availability, the evolution of labour market institutions in Central and Easter Europe has attracted much scholars' attention (see, for example, Cazes and Nesporova, 2003). The literature and the evidence provided by a variety of data sources unanimously show a fairly modest level of institutional rigidities in the labour market and a general trend towards liberalization since the mid-1990s in the whole transition region (see Lehmann and Muravyev, 2012). However, there are also important differences across countries that deserve attention (see Cazes and Nesporova, 2007; Fialova and Schneider, 2009). As for the aspects and the countries considered here, the data show a very complex picture, depicted in table A1 in the appendix in which the average values for western EU countries are also reported as a reference. Data confirm how various countries pursued a flexible model of labour market, without completely giving up an active role of collective organisations (particularly Bulgaria, Slovak Republic and Hungary). In other countries, such as the Czech Republic, the process of deregulation was more limited, while the Baltic states offer a combination of low age coordination and unionisation and intermediate (Estonia and Latvia) or low (Lithuania) labour market deregulation. Slovenia shows limited deregulation on the labour market, high levels of bargaining coordination and relatively high union density; this evidence confirms the convergence of Slovenian capitalism towards the neo-corporatitivist model

[5] The 5.b.i indicator of the Fraser database is in fact based on the World Bank's Doing Business Difficulty of Hiring Index, which is described as follows: "The difficulty of hiring index measures (i) whether fixed-term contracts are prohibited for permanent tasks; (ii) the maximum cumulative duration of fixed-term contracts; and (iii) the ratio of the minimum wage for a trainee or first-time employee to the average value added per worker". In the dataset countries with higher difficulty of hiring are given lower ratings. See: http://www.doingbusiness.org/ and http://www.fraserinstitute.org (particularly the appendix to Gwatney et al., 2010). The Union density rate is calculated as the net union membership as a proportion wage and salary earners in employment. Wage coordination ranges from 5 (economy-wide bargaining, based on a-enforceable agreements between the central organisation of unions and employers affecting the entire economy or entire private sector, or on b-government imposition of a wage schedule, freeze, or ceiling) to 1 none of the above, fragmented bargaining, mostly at company level).

(Bohole and Greksovtis, 2007). Of course, the very short distance between the two years of reference does not allow for any variations over time of institutional indicators, the only exception being a generalised decline in unionisation.

As suggested by macro-economic models of employment and wage determination (Carlin and Soskice, 1990), when running pooled regressions for the whole Eastern EU region we also included macroeconomic (namely unemployment rate) and other institutional controls at country level (namely, a Product Market Deregulation index, again by the Fraser Institute), as well as country dummies to account for unobserved residual country-level heterogeneity.

3.2 Preliminary descriptive evidence on raw earnings gaps

Table 1 reports basic descriptive statistics on hourly earnings by gender in the two years considered.

The raw gender gap, calculated as the female/male earnings ratio, varies remarkably between countries, ranging from 0.70 in Estonia to 0.95 in Slovenia in 2007. The remaining countries tend to polarize into two groups, around 0.90 (Poland and Hungary) and 0.80 (the remaining ones). After the outburst of the crisis, the gender gap declines in all countries except Hungary, while their ranking remains unchanged. Inequality of earnings within the two gender groups does not appear as significantly different, with the only exceptions of Romania in which dispersion of female labour incomes is three Gini points higher than for man in both years. The small differences in dispersion of male and female earnings and over time, as measured by the coefficient of variation, suggest that asymmetric evolutions of the shape of the distribution – particularly increasing inequality within men only – cannot be considered as major sources of gender inequality, as otherwise reported (Gregory, 1999).

The generalised gender earnings gap decline in 2009, compared to 2007, is clearly the result of convergence between the median males' and the lowest females' earnings. The same happens within the three employment status groups (see Figure 1). Permanent contracts (about 78% of the sample against 9% of temporary and 13% self employed) are associated with the highest average earnings for both genders; temporary and self-employed earnings approximately coincide only for males. For women, self-employment leads to significantly lower returns when compared to temporary positions. The kernel density plots in Figure 2 provide a snapshot of the gender distribution differences across Eastern Europe. First, the position of the density curves reveals the remarkable differences in the modal values of hourly earnings across the ten countries, with Romania, Bulgaria and to a lesser extent Slovak Republic located at the extremely low values area. On the opposite side, the density curves for Slovenia show the most rightward position and less concentration around a single modal value. In general, all plots shift back to lower values in 2009 compared to 2007, with increasing density associated to the modal value. The distributions for Estonia, and to a lesser extent for the Czech Republic, confirm their specificity in terms of remarkably high gender gap, providing evidence of poorly overlapping plots for men and women.

Table 1 Male/Female Hourly earnings in Eastern EU member countries, 2007 and 2009 (2005 Euro PPP)

	2007								
	Obs		*Median*			*Coef. Var.*		*Gini*	
	M	F	M	F	F/M	M	F	M	F
BG	2411	1983	3.08	2.40	0.78	0.64	0.65	0.31	0.30
CZ	6048	4713	6.55	5.15	0.79	0.45	0.44	0.24	0.23
EE	2536	2474	5.48	3.83	0.70	0.56	0.57	0.30	0.28
HU	3977	3416	4.35	4.06	0.93	0.61	0.58	0.30	0.29
LT	2243	2258	5.33	4.18	0.78	0.60	0.63	0.31	0.32
LV	2301	2354	4.86	3.93	0.81	0.67	0.69	0.35	0.35
PL	7004	5494	5.02	4.50	0.90	0.66	0.70	0.34	0.36
RO	3773	2600	2.85	2.20	0.77	0.60	0.65	0.32	0.35
SI	5926	4931	8.92	8.51	0.95	0.54	0.54	0.28	0.28
SK	3739	3173	4.30	3.48	0.81	0.43	0.42	0.23	0.22
Total	39958	33396	5.19	4.36	0.84	0.67	0.72	0.30	0.30

Table 1 (continued)

	2009								
	Obs		*Median*			*Coef. Var.*		*Gini*	
	M	F	M	F	F/M	M	F	M	F
BG	3256	2806	2.75	2.38	0.87	0.58	0.58	0.29	0.29
CZ	4656	3654	5.83	4.86	0.83	0.48	0.47	0.24	0.24
EE	2048	2213	4.61	3.50	0.76	0.59	0.60	0.31	0.30
HU	4315	4007	3.47	3.18	0.92	0.57	0.55	0.29	0.27
LT	2091	2383	3.49	3.20	0.92	0.76	0.77	0.36	0.38
LV	2077	2420	3.19	2.93	0.92	0.72	0.69	0.37	0.35
PL	6253	4702	4.24	4.04	0.95	0.69	0.69	0.34	0.34
RO	3555	2591	1.91	1.71	0.89	0.58	0.63	0.31	0.34
SI	5942	4952	8.45	8.20	0.97	0.62	0.57	0.31	0.30
SK	3537	3150	4.88	4.30	0.88	0.44	0.39	0.23	0.21
Total	37730	32878	4.38	3.85	0.88	0.77	0.78	0.31	0.30

Figure 1 Male / Female hourly earnings in Eastern EU member countries by employment status, 2007 and 2009 (2005 Euro PPP)

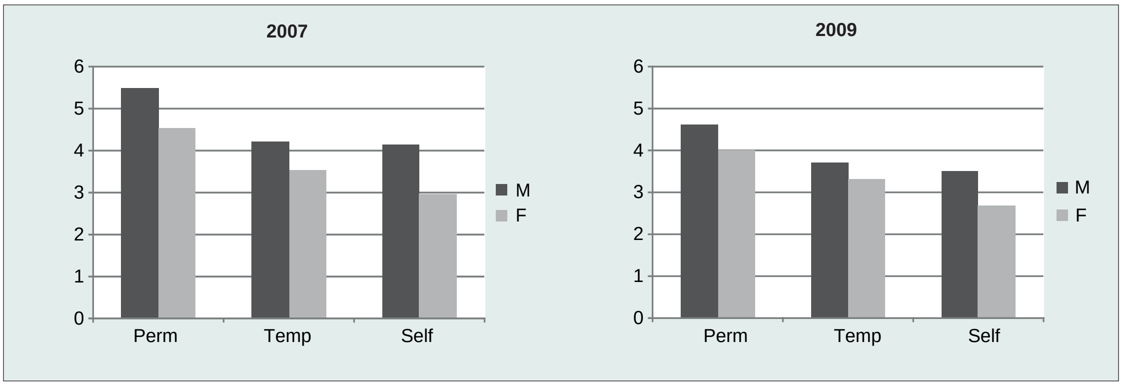
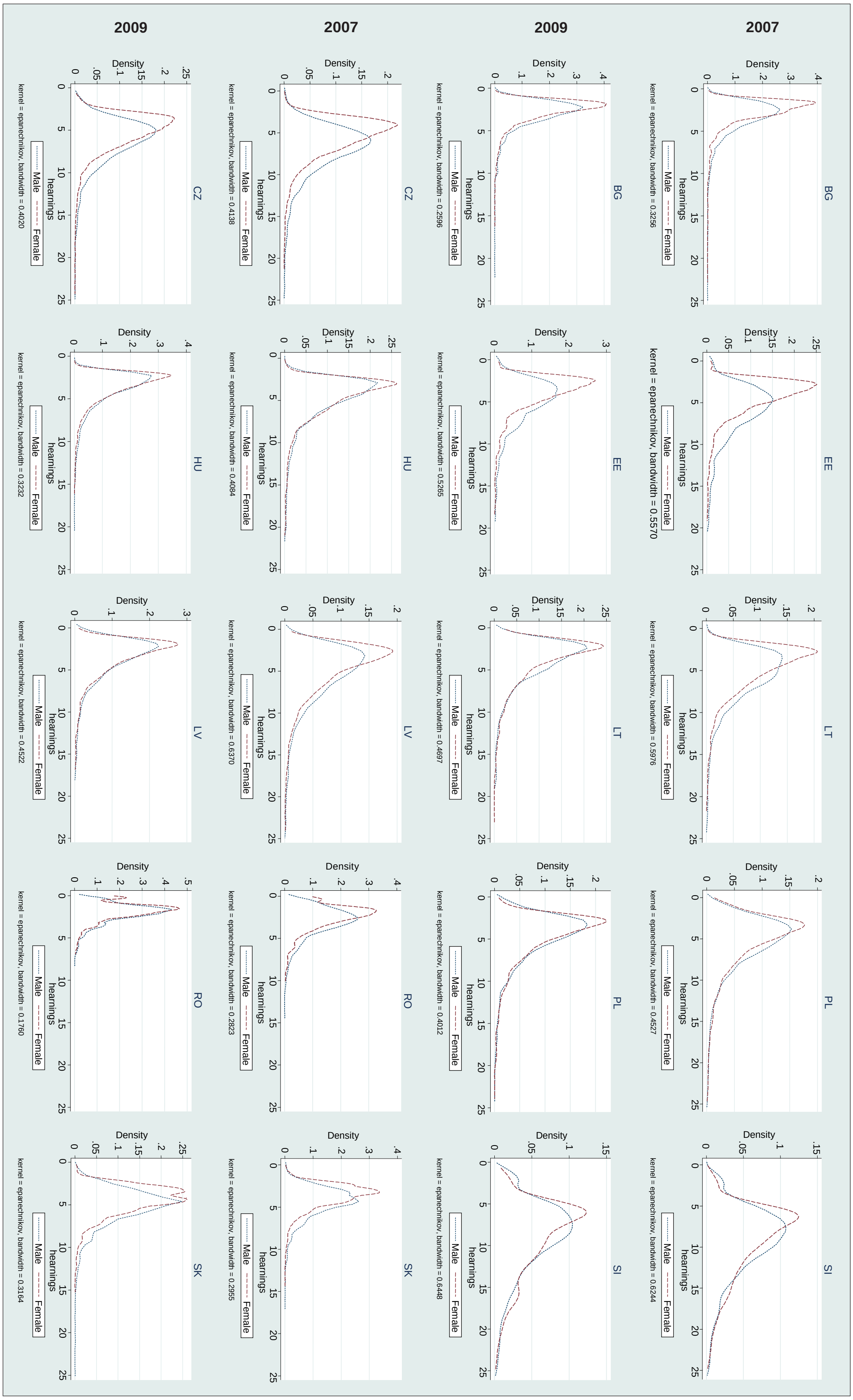


Figure 2 Kernel density distributions of Male/Female hourly earnings in Eastern EU member countries, 2007 and 2009 (2005 Euro PPP)



4 Econometric methods

The raw (unadjusted) gender earning gap presented in Table 1 does not account for many characteristics relevant in shaping male and female earnings, such as the level of education, experience, skills, employment status, occupation, sector of employment. Comparison of raw earnings gaps would therefore not compare like with like. To allow for the role of observable characteristics we estimate a log hourly earnings (*lhean*) equation in which the coefficient of the gender dummy (*male* = 1) provides, *coeteris paribus*, the estimate of the percent *residual gender earnings gap* (Newell and Reilly, 2001). The estimation of one equation on a pooled sample of two sexes implies assuming that men and women are paid the same rewards for their labour characteristics, which – as the literature suggests – might not be the case (Albrecht et al., 2003). The application of decomposition methods allow identifying the extent to which the gender gap can be explained by differences in characteristics versus differences in labor market rewards to those characteristics. Given the aims of the present study we limit the presentation and discussion of empirical evidence to regression analyses only; the outcomes of the decomposition of earning gaps at different percentiles of the distribution, according to the Machado and Mata (2005) approach, are available upon request.

As far as the drivers of individual earnings are concerned, we rely on the human capital model as the theoretical basis for the earnings function (Becker, 1964; Mincer, 1958). We therefore assume that labour income increases first of all with measures of accumulated formal (education) and informal (experience) skills. Education is measured by the highest level of education achieved (primary, secondary, tertiary). We approximate work experience with age, as the work experience measure (PL 200 – number of years spent in paid work) in EU-Silc is not available for all the relevant countries and presents many missing values[6]. We further control for other explanatory variables available in the dataset, namely marital staus (*married*), health status (*health*); urban/non-urban region of residence (*urb*), employment status (*temp, self*); part-time job position (*part*); presence of a second job (*sjob*); sector of employment (*sec*); occupation (*occ*); size of the firm (*size*). All these variables, but especially those referring to the sector,

[6] The correlation between age and experience, where available, is above 0.7 in both the 2007 and 2009 samples (significant at 1%).

employment status and occupation play a crucial role in explaining gender earning differences (see, for example, Manning and Petrongolo, 2008), particularly related to vertical segregation. Information about occupations is also aimed at internalizing in the analysis the distinction between routine/not-routine task and their role in shaping relative earnings (Autor et al., 2003). This wide range of information allows for interpreting the gender dummy variable (*male*) coefficient as a reliable proxy of revealed unexplained gender earnings gap, that is, gender pay discrimination. In this manner we estimate *k* country specific empirical models (with k = 10 Central and Eastern EU members) for 2007 and 2009.

In order to investigate the impact of labour market institutional factors on gender earning gap we then estimate a pooled model (without Romania, for which the information on occupation is not available). The role of the three institutional settings (*INST*: (i) Hiring regulations and minimum wage; (ii) Union density; (iii) Collective bargaining coordination) on the gender gap is identified by means of interaction terms ((vector) *INST x male*), which specify their additional (gap augmenting or diminishing) effect on gender differences, with respect to coefficient of the gender dummy.

Our baseline pooled empirical model takes therefore the following form:

$$\begin{aligned} lheran_{i,k} &= c_i + \alpha_1 age_{i,k} + \alpha_2 age^2_{i,k} + \alpha_3 married_{i,k} + \alpha_4 health_{i,k} + \alpha_5 urb_{i,k} + \alpha_6 male_{i,k} + \omega\, male_{i,k} \cdot INST_k + \\ &\beta_1 temp + \beta_2 self + \beta_3 part + \beta_4 sjob + \sum_{s=1}^{2} \beta_s size_s + \sum_{n=1}^{7} \beta_n \sec_n + \sum_{r=1}^{5} \beta_r occ_r + \gamma_1 INST_k + \gamma_2 PMD_k + \gamma_3 UR_k + \delta_k + \varepsilon_{i,k} \end{aligned} \tag{1}$$

where subscripts *i* and *k* stand for individuals and countries, respectively; δ_k represents unobservable country-specific effects and ε_{ik} the usual error term. PMD and UR are country level controls for the level of product market deregulation and unemployment rate, respectively. As customary in the literature about institutions (Bassanini et al., 2009; Bourlès et al., 2012), institutional variables are lagged one period in order to alleviate endogeneity issues and to account for the fact that *de jure* institutional reforms take time to become effective. The country specific estimates of the gender earnings gap are obtained as the coefficient of *male* simply restricting the model in equation 1 to single countries, consequently dropping the country specific variables *INST*, *PMD*, *UR* and the country fixed effects.

The empirical models following from equation 1 allow estimating only average effects of explanatory variables on log earnings and, in particular, average gender earning gaps. However, such approach does not allow understanding how gender differences unfold across the earning distribution and whether women encounter a glass ceiling. This would be revealed by women's earnings falling behind those of men at a higher extent at the top of the wage distribution than at the middle or bottom. Analogously, a higher or an increasing gap at the bottom of the distribution would reveal a sticky floor effect. In order to investigate whether these effects exist a wider view of the gender gap over the entire distribution is needed, which is rendered possible by quantile regression (QR) approaches. The estimated gender dummy coefficients in these regressions thus indicate the extent to which the gender gap remains unexplained at the various quantiles when we control for individual differences of characteristics.

Following Koenker and Basset (1978), the model of QR can be simply described in terms of conditional θ^{th} quantile (instead of conditional mean as in the standard regression) distribution of y_i conditional on a vector of covariates x_i under the assumption of linear specification:

$$Q_\theta(y_i \mid x_i) = x_i\beta_\theta \qquad (2)$$

implying $y_i = x_i\,\beta_\theta + \varepsilon_{\theta,i}$. The semiparametric nature of the approach lies in the fact that the distribution of the error term $\varepsilon_{\theta,i}$, $F_{\varepsilon,\theta}(\cdot)$, is left unspecified, and $\varepsilon_{\theta,i}$ satisfies $Q_\theta(\varepsilon_{\theta,i} \mid x_i) = 0$.

The θ^{th} QR estimator $\hat{\beta}_\theta$ minimizes over β_θ the following objective function:

$$Q(\beta_\theta) = \sum_{i:y_i \geq x_i\beta}^{n} \theta\left|y_i - x_i\beta_\theta\right| + \sum_{i:y_i < x_i\beta}^{n} (1-\theta)\left|y_i - x_i\beta_\theta\right| \qquad (3)$$

The estimated vectors of QR coefficients $\hat{\beta}_\theta$ measures the marginal change in the conditional quantile θ due to a marginal change in the corresponding element of the vector of coefficients on x, obtained using the optimization techniques described for example in Cameron and Trividi (2009), as the usual gradient optimization method cannot be applied since equation (3) is not differentiable.

QR estimations are run using the simultaneous quantile regression (sqreg) routine available in STATA, which allows specifying simultaneously different values of θ and testing whether regression coefficients of interest for various θ do differ (by means of a Wald test). This option provides bootstrap standard errors, which are robust and assume independence over *i* but do not require errors to be identically distributed.

Table 2 Gender labour market differences in Eastern European countries, 2007 and 2009

	Employment rate, %		*Part-time workers, %*		*Temporary contracts, %*		*Tertiary education, %*	
Men	***2007***	***2009***	***2007***	***2009***	***2007***	***2009***	***2007***	***2009***
BG: Bulgaria	73.4	73.8	1.3	2.0	5.0	5.2	14.6	15.0
CZ: Czech Republic	81.5	80.2	2.3	2.8	7.3	7.0	12.4	13.7
EE: Estonia	81.4	71.0	4.3	7.0	2.7	3.0	20.8	21.4
LV: Latvia	80.1	67.4	4.9	7.5	5.5	4.7	14.8	15.6
LT: Lithuania	76.5	66.9	7.0	7.0	4.9	2.9	20.2	20.7
HU: Hungary	70.2	67.0	2.8	3.9	7.7	9.0	13.8	14.7
PL: Poland	70.2	72.6	6.6	5.8	28.4	26.3	13.1	14.8
RO: Romania	71.0	70.7	9.2	9.1	1.7	1.1	10.0	10.9
SI: Slovenia	77.5	75.6	7.7	8.4	16.5	15.1	15.5	15.8
SK: Slovakia	76.0	74.6	1.1	2.7	4.9	4.6	11.5	12.5
Women	***2007***	***2009***	***2007***	***2009***	***2007***	***2009***	***2007***	***2009***
BG: Bulgaria	63.5	64.0	2.1	2.7	5.5	4.2	22.4	22.7
CZ: Czech Republic	62.4	61.4	8.5	9.2	10.2	10.2	10.9	13.1
EE: Estonia	72.5	68.8	12.1	13.8	1.6	2.0	33.2	37.6
LV: Latvia	70.7	66.8	8.0	10.2	2.9	2.9	22.6	27.4
LT: Lithuania	69.5	67.5	10.2	9.5	2.3	1.6	27.8	30.0
HU: Hungary	55.5	54.4	5.8	7.5	6.8	7.8	16.9	19.0
PL: Poland	55.5	57.6	12.5	11.6	27.9	26.6	18.1	21.3
RO: Romania	57.9	56.3	10.4	10.6	1.5	1.0	9.8	11.6
SI: Slovenia	67.1	67.9	11.3	13.2	20.8	17.8	21.6	23.6
SK: Slovakia	58.7	58.2	4.5	4.7	5.3	4.1	12.3	14.4

Notes: education (15 – 64 y.o.); employment rate (20 – 64 y.o.); part-time workers and temporarily contracts – in % of all employed.

Source: Eurostat

A last crucial aspect we need to address refers to a possible bias originated by sample selection. If selection of individuals into employment is nonrandom, the direction in which it may affect the level of earnings is a concern. In the field of gender studies, a growing literature has for example recognized that employed women tend to have –

more often – characteristics normally associated to high wages (De la Rica et al., 2008; Heckman, 1979; Buchinsky, 1998). As a consequence, low female employment rates may become consistent with low gender wage gaps simply because the low-wage women would not feature in the observed wage distribution. Differences in participation in employment may result from a number of factors, especially at cross-country level (Albrecht et al., 2009). They include differences in labor supply behavior related to household structure or social norms, and in institutional settings such as unionization or minimum wages (Olivetti and Pietrolongo, 2008). The countries considered in this study show impressive gender differences in terms of remarkably low female employment rates, higher incidence of part-time, temporary contracts and higher education levels (see table 2).

Heckman in 1974 and 1979 proposed a parametric estimator to estimate covariates with selection bias; Powell (1987) and Newey (1991) developed a semi-parametric estimator for the sample selection model. More recently Das et al. (2003) introduced a nonparametric estimator. Buchinsky (1998 and 2001) was the first to apply the semiparametric sample selection model for quantile regression. We follow here the approach by Buchinsky (1998), explained in more detail in Albrecht et al. (2009) and Nicodemo (2009). As the recent literature shows that also men do not randomly select into employment (Christofides and Vrachimis, 2007), we control for sample selection for both genders. We therefore estimate the quantile regression of individuals employed (for which we observe the log earning rate) as:

$$Q_\theta(y \mid x) = x\beta_\theta + h_\theta(z\lambda) \tag{4}$$

where z is the set of observable characteristics that influence the probability that an individual is employed which must also contain, for the identification, at least one variable that is not included in x. In our case, in addition to the individual characteristics associated to coefficients (α) in equation 1 and the country level institutional and macroeconomic controls in the case of pooled sample, we add variables related to household structure, namely: number of household components, number of children (less than 3, 4–6 and 7–15 years old), number of elderly (65–74 and over 75 years old). The term $h_\theta(z\lambda)$ corrects for selection at the θ^{th} quantile, playing the role that the Mills ratio plays in Heckman (1979) procedure, but it is quantile-specific and more general so as not to assume normality (Albrescht et al., 2009). Following the Buchinky's method, the $h_\theta(z\lambda)$ can be approximated by a power series whose coefficients has to be estimated and should define a function which is larger when the impact of unobservable is larger (Bosio, 2009). This function is the inverse Mill's ratio, being small for those with an high probability of being temporary and increasing monotonically as the probability of being temporary reduces. Following Arumpalan et al. (2006) we therefore control for the selectivity bias in QR earning equation expanding $h_\theta(z\lambda)$ as a power series in the inverse Mill's ratio, derived from a participation equation dependent on the vector of explanatory variables *z*. The latter is estimated by both a standard probit model and a single index model (Ichimura, 1993), by means of the semiparametric ML estimator of Klein and Spady (1993). In the second stage, QR are augmented by the derived inverse Mill's ratio and its square[7].

[7] Results of the QR are substantially invariant to the use of Mill's ratio calculated with the two methods and we report here the results obtained with the probit estimation. A model specification with a series of power 3 (of the Mill's ratio) is also tested, but the additional term generally did not turn out statistically. Results of the QR are also substantially invariant to the use of Mill's ratio calculated with the two methods and we report here those obtained with the probit estimation.

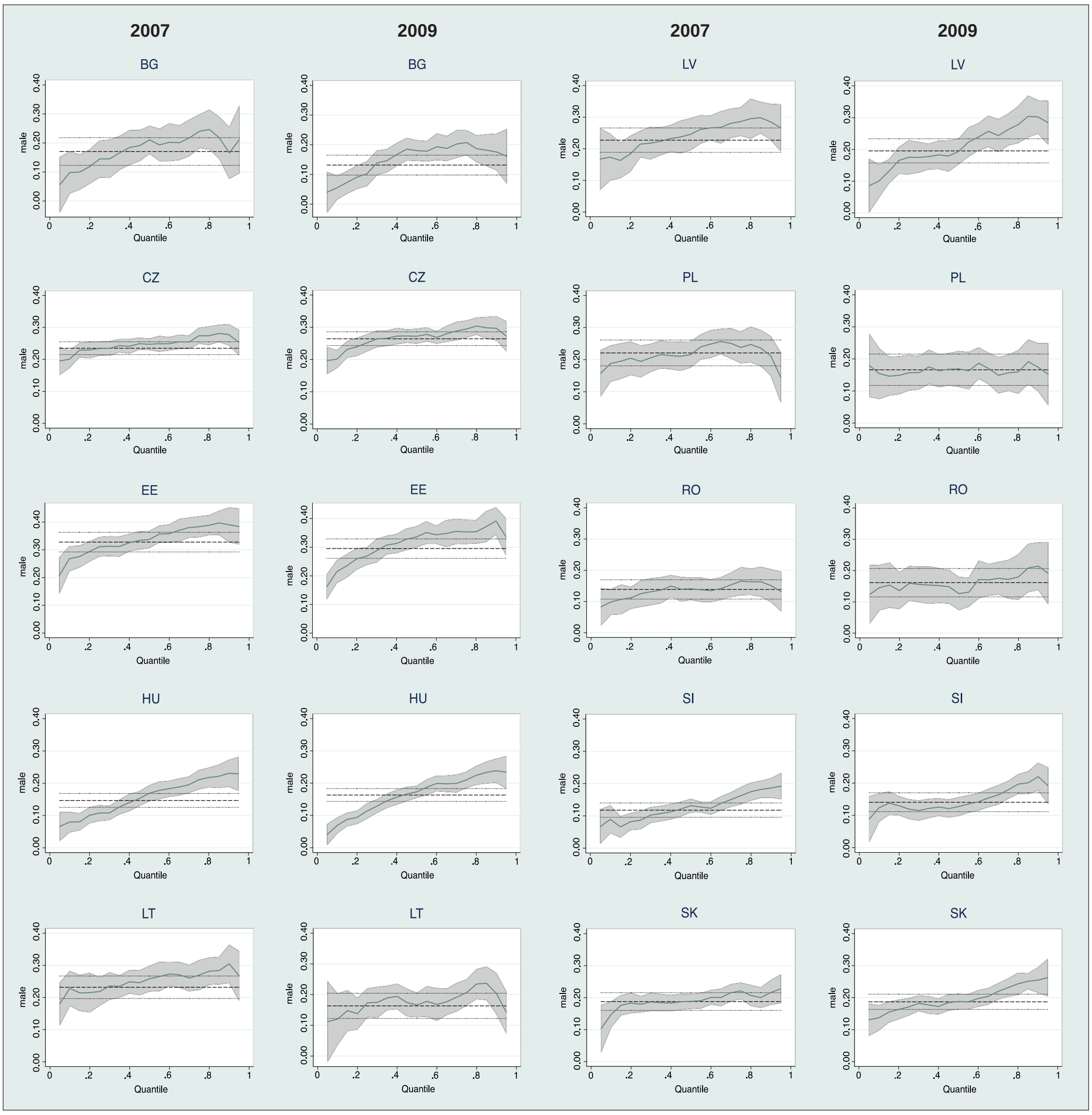
5 Econometric analysis: Adjusted gender earnings gap and the role of institutions

5.1 Adjusted gender earnings gap across the distribution in East EU countries

Using the quintile regression method we show how the (adjusted) unexplained gap changes at different points (quintiles) of the hourly earnings distribution. The log-earnings equation (equation 1) is estimated at various percentiles (from 0.05 to 0.95, with a 0.05 interval) of the earnings distribution, country by country, using bootstrap standard errors (obtained with 400 replications) and controlling for self-selection of both men and women into employment. The estimation results country-by-country, available upon request, show that the explanatory variables of log hourly earnings play the role expected ex-ante. We summarize graphically the outcomes of interest in Figure 3, which plots the OLS and the quantile regressions coefficients (and 95% confidence intervals) of the male dummy for each of the 10 East EU members, for 2007 and 2009. Note that the estimates for Romania are not directly comparable to the results for other countries, as the occupation variable was missing for this country and the gender gap estimate does not control for this information.

Our estimates reflect the improvement in the relative female wage position in Czech Republic and Estonia (in 2009 compared to 2007); in other countries the evidence is not clear. It is widely confirmed that in case of developed countries (old-EU members), "the gap typically widened toward the top of the wage distribution (the "glass ceiling" effect), especially in northern and central European countries (Dolado and Llorens, 2004); in a few cases it also widened at the bottom (the "sticky floor" effect)" (Arulampalam et al. 2006). Like Newell and Reilly (2001), who found a steady rise in the adjusted gender wage gap across the wage distribution for majority of the transition countries (in 1992–1996), we find this tendency for most of the countries of our sample. Hence, we can conclude on lower gender inequality (gender discrimination) at the bottom of the conditional earnings distribution for Bulgaria, Estonia, Hungary, Latvia, Slovenia and Slovakia (more important "glass ceiling" effect). In some of the cases the estimated gender effect is rather flat over the different percentile of the earnings distribution: Czech Republic, Poland, Lithuania (in 2009) and Romania (2009).

Figure 3 Adjusted gender earnings gap in 2007 and 2009, by quantiles



Note: For Romania the role of occupations is not controlled for due to missing information.

In the literature, a negative relationship between gender employment gap and GWG was found (Olivetti and Petrongolo, 2008), which is in general confirmed in our data in terms of earnings. This evidence is consistent with the fundamental Becker's work on discrimination (Becker, 1971): if there are discriminatory employers, as more women enter the labour market they will have to more frequently deal with discriminatory employers and this will lower their relative wage. In our data the high gender earnings gaps observed for the Baltic States correspond to the lowest levels of employment gap (below 10%); the opposite holds for Slovenia, Hungary, Romania, where the employment gap exceeds 10% and approaches 15%. The Baltic countries, however, also have a higher share of men (than women) with temporary contracts (in the literature, higher share of temporary workers, especially among women, is associated with a higher pay gap). We can also note virtually no (or very low) difference between the OLS estimates of the gap and the quantile regression estimate at the median for majority of the countries. Bulgaria is the only case in which a remarkable difference between the mean gap and gap at the median emerges in both years.

5.2 Institutions and gender gap: Econometric evidence

The average characteristics of the pooled sample used in the econometric analysis, which does not include Romania due to missing data on occupations, are summarized in Table A2. Female workers are, on average, older than males, report better health status and have remarkably higher incidence of tertiary education. As for employment characteristics, they hold more part-time and more permanent positions; the incidence of temporary jobs is not remarkably different, whereas self-employment incidence among men is twice as high as among women. The sector breakdown of employment by gender confirms the usual vertical segmentation pattern, with male workers employed relatively more often in agriculture, industry, transports and constructions and women into services sectors. As for occupations, the share of females is higher among professionals and technicians, clerks and elementary occupations. Along with the evidence of a lower share of women in the top positions (managers and senior officials), this outcome provides another representation of the glass ceiling /sticky floor effects. Overall, the information supplied by average characteristics of the sample confirms what is common knowledge. However, if the sample is analyzed at different parts of the earnings distribution, useful additional aspects emerge and they will be used in the presentation and discussion of the econometric results.

The baseline quantile regression estimates are reported in full in Table 3 (for θ =0.10, 0.50, 0.90); the OLS specification and the quantile regressions with interaction effects are reported in Tables A3–A6 of the appendix; in the main text we limit their presentation to a summary of the effects of the labour market institutions on the earnings gap, both in tabular and graphical form (Table 4 and Figure 4, respectively).

Before focusing on the gender variable (male) and on the interaction terms, we should briefly note that the estimated models of earnings provide a remarkably solid – and consistent with ex-ante expectations – picture (across all estimations) of earnings determinants. The baseline regressions (Table 3 and A3) show that age, which is also a proxy for experience, has the expected positive (and non linear) effect; better health status is positively associated to earnings, as well as being married, residing in urban areas and achieving higher levels of education. Holding a temporary position or being self-employed is in general associated to lower hourly returns compared to permanent jobs; part-time employment is associated to higher earnings at the median and at the top of the distribution, but negatively at the bottom tail. This may reflect the nature of part-time job choices, driven by pull factors in the first case and by push factors in the second. Being employed in larger firms also guarantees higher earnings. The sector and occupation controls provide expected hierarchies of coefficients (not reported for the sake of brevity, but available upon request). As for the institutional variables, a discussion of their effects on earnings is beyond the scope of this paper and the object of an extensive and controversial literature (see Batcherman, 2013, for a survey). Here, we only briefly emphasise that in our results labour market deregulation is associated, on the median, to lower earnings. This also holds, in 2009, for the incomes at the top and at the bottom of the distribution. As for 2007, the impact of deregulation is positive for the bottom end and non statistically significant for the upper tail; this complexity probably drives the positive coefficient obtained on the OLS regressions and confirms the inadequacy of this average approach to represent the complex effects of institutional settings. The presence of unions is found to impact positively on earnings, in every segment of the distribution, but with stronger effects for those at the bottom. On the contrary, stronger wage coordination tends to decrease hourly earnings.

Table 3 Quantile regression estimates, pooled model (2007 and 2009)

	2007			*2009*		
	θ =.10	θ = 50	θ = 90	θ =.10	θ = 50	θ = 90
male	0.151*** (0.008)	0.212*** (0.005)	0.236*** (0.009)	0.117*** (0.008)	0.193*** (0.005)	0.245*** (0.009)
married	0.019*** (0.007)	0.035*** (0.004)	0.034*** (0.007)	0.024*** (0.006)	0.031*** (0.005)	0.026*** (0.007)
age	0.021*** (0.005)	0.022*** (0.003)	0.020*** (0.006)	0.016*** (0.004)	0.025*** (0.003)	0.038*** (0.005)
age^2	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)
Health	–0.045*** (0.009)	–0.049*** (0.005)	–0.052*** (0.009)	–0.037*** (0.007)	–0.040*** (0.005)	–0.052*** (0.008)
Temp	–0.159*** (0.010)	–0.138*** (0.007)	–0.086*** (0.012)	–0.152*** (0.010)	–0.126*** (0.007)	–0.090*** (0.014)
Self	–0.630*** (0.025)	–0.134*** (0.010)	0.076*** (0.013)	–0.698*** (0.023)	–0.147*** (0.009)	0.088*** (0.014)
Secondary	0.045*** (0.014)	0.086*** (0.008)	0.095*** (0.016)	0.022*** (0.011)	0.098*** (0.009)	0.154*** (0.014)
Tertiary	0.244*** (0.020)	0.345*** (0.012)	0.395*** (0.023)	0.179*** (0.018)	0.350*** (0.013)	0.491*** (0.021)
Part	–0.093*** (0.017)	0.044*** (0.012)	0.209*** (0.020)	–0.050*** (0.015)	0.064*** (0.013)	0.214*** (0.019)
sjob	0.167*** (0.008)	0.101*** (0.005)	0.088*** (0.007)	0.750*** (0.045)	0.763*** (0.048)	0.917*** (0.062)
Size (11–49)	0.140*** (0.008)	0.098*** (0.005)	0.042*** (0.008)	0.154*** (0.008)	0.114*** (0.005)	0.057*** (0.008)
Size (over 50)	0.240*** (0.009)	0.193*** (0.005)	0.125*** (0.009)	0.263*** (0.008)	0.211*** (0.006)	0.147*** (0.009)
urb	0.060*** (0.007)	0.080*** (0.005)	0.101*** (0.008)	0.054*** (0.006)	0.077*** (0.005)	0.097*** (0.008)
LM_dereg	0.013*** (0.003)	–0.005* (0.002)	–0.002 (0.004)	–0.102*** (0.003)	–0.101*** (0.002)	–0.098*** (0.003)
PMD	0.056*** (0.008)	0.090*** (0.005)	0.127*** (0.008)	1.328*** (0.074)	0.750*** (0.054)	0.384*** (0.067)
UR	–0.105*** (0.002)	–0.095*** (0.002)	–0.090*** (0.002)	–0.373*** (0.017)	–0.225*** (0.012)	–0.122*** (0.015)
Union Density	0.115*** (0.002)	0.094*** (0.001)	0.071*** (0.002)	0.353*** (0.020)	0.188*** (0.014)	0.085*** (0.018)
W_coord	–0.516*** (0.008)	–0.404*** (0.006)	–0.277*** (0.008)	–2.843*** (0.159)	–1.518*** (0.114)	–0.708*** (0.143)
_cons	–0.085 (0.107)	0.295*** (0.069)	0.739*** (0.112)	–3.519*** (0.290)	–1.346*** (0.210)	0.041 (0.274)
Sector/occup/country dummies	yes	yes	yes	yes	yes	yes
Sample-selection correction	yes	yes	yes	yes	yes	yes
Test F [q10=q50=q90]: male	37.30*** [2, 66944]			66.30*** [2, 64425]		
Obs	66981	66981	66981	64462	64462	64462
Adj. R-Sq	0.296	0.307	0.286	0.315	0.343	0.333

Notes: Robust standard errors in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. W_coord, LM_dereg, UD, PMD: lagged one year.

Table 4 Quantile regression estimates, pooled model (2007 and 2009): Summary of the effects of labour market institutions on the gender earnings gap

	2007			*2009*		
	θ =.10	θ = 50	θ = 90	θ =.10	θ = 50	θ = 90
Labour market deregulation						
Male	0.164***	0.172***	0.161***	0.137***	0.142***	0.202***
	(0.020)	(0.011)	(0.017)	(0.021)	(0.012)	(0.018)
LM_dereg	0.014***	–0.007***	–0.009**	–0.101***	–0.104***	–0.101***
	(0.004)	(0.002)	(0.004)	(0.003)	(0.002)	(0.003)
LM_dereg_male	–0.002	0.006***	0.012***	–0.003	0.007***	0.006***
	(0.003)	(0.001)	(0.002)	(0.002)	(0.001)	(0.002)
Test F [q10=q50=q90]: LM_dereg_male	7.79*** [2, 66943]			8.00*** [2, 64424]		
Union Density						
Male	0.286***	0.429***	0.500***	0.136***	0.316***	0.375
	(0.021)	(0.014)	(0.023)	(0.021)	(0.014)	(0.020)
UD	0.117***	0.100***	0.079***	0.356***	0.188***	0.084
	(0.002)	(0.001)	(0.002)	(0.021)	(0.014)	(0.020)
UD_male	–0.007***	–0.011***	–0.013***	–0.001	–0.007***	–0.007
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Test F [q10=q50=q90]: UD_male	11.76*** [2, 66943]			12.66*** [2, 64424]		
Wage bargaining coordination						
Male	0.210***	0.291***	0.338***	0.111***	0.259***	0.313
	(0.014)	(0.010)	(0.015)	(0.015)	(0.010)	(0.014)
W_coord	–0.503***	–0.388***	–0.247***	–2.851***	–1.486***	–0.659
	(0.009)	(0.007)	(0.010)	(0.154)	(0.107)	(0.151)
W_coord_male	–0.028***	–0.036***	–0.045***	0.003	–0.032***	–0.032***
	(0.005)	(0.003)	(0.005)	(0.007)	(0.004)	(0.005)
Test F [q10=q50=q90]: W_coord_male	2.85* [2, 66943]			13.40*** [2, 64424]		

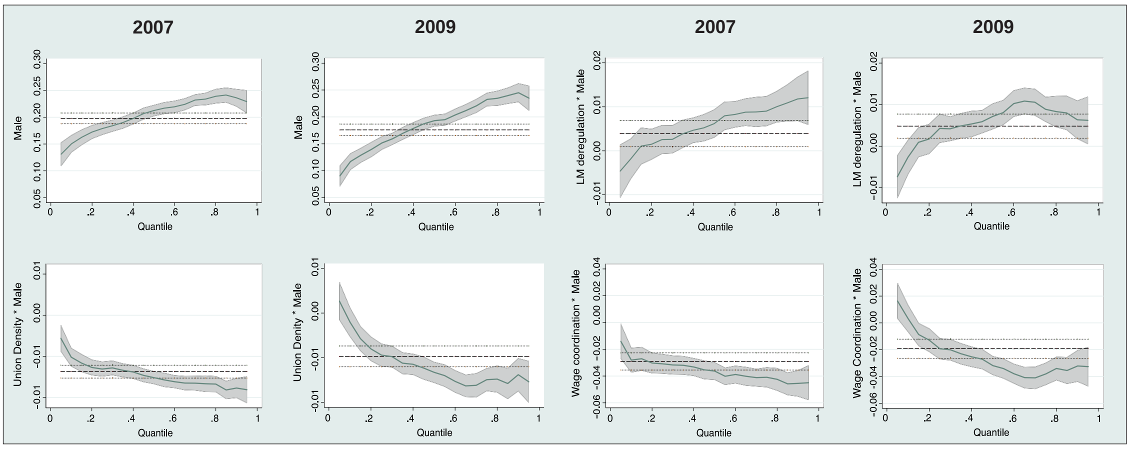
Notes: Robust standard errors in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. Complete estimations results are available in the appendix (Tables A4–A6). All estimations include all workers and employment characteristics (see table A1), plus country level controls for unemployment rate (UR), Product Market deregulation (PMD) and country dummies. W_coord, LM_dereg, UD, PMD: lagged one year.

The coefficients of the gender dummy variable measure the earnings gap in different parts of the earnings distribution, once all observable workers' and job characteristics are controlled for. Table 3 shows that the size of discrimination varies across the distribution, getting higher as earnings grow. This confirms the presence of a clear glass-ceiling effect. The top-left panel of Figure 4 displays the size of the coefficients at every 5th percentile of the distribution. Although, consistent with the evidence already provided, the average and median gap decreased in 2009 compared to 2007, it is noteworthy that the slope of the plot increased, which means a reduction of the gap for the bottom earnings, but also an increase at the top. In other words, behind the average/median decrease of the wage gap there is a deepening of the glass ceiling effect. Therefore, as already emphasized, the reduction of average earnings gap is not good news: the gap decreases at the bottom tail due to a downwards convergence of male earnings to the lowest levels of female earning, but discrimination increases at the top.

So the crisis renders labour income between genders more similar only because men are pushed towards the lowest, incompressible female levels: that is, where there is not any room for discrimination left. As a matter of facts, the pool of lowest earning workers (bottom 10%) in 2007 contained 46% of man; in 2009 this share grows to 49%. This evidence is the starting point to interpret the outcomes on the role of labour market institutions in shaping the gender gap (Table 4 and top-right/bottom panels of Figure 4).

Stronger unions and wage coordination (UD_male and W_coord_male) reduce the average and median gender gap, but their effects are weaker at the bottom of the distribution and are not statistically significant in 2009. This suggests that these labour market institutions have a weak role in addressing discrimination for the most disadvantaged segments of the labour market, especially as the crisis starts to produce its effects. This may depend on the fact that employment in this segment of the labour market is concentrated in industries typically less affected by unions or collective bargaining (agriculture or certain services), or by higher shares of self-employment induced by push forces (Falter, 2007), as it was typically the case in transition countries (Earle and Zakova, 2000). In addition, as emphasized in the literature (Gregory, 1999), in the presence of very low pay for male workers, especially for those employed in typically female occupations, the room for reducing discrimination is smaller and the role of institutions declines accordingly. The evidence on the role of labour market deregulation (LM_dereg_male), which in our case includes information on temporary job regulations and on minimum wage provisions, corroborates this interpretation. Our findings show that deregulation tends to favour gender discrimination: this is certainly due to the fact that a significant part of female's flows into employment has been associated, in the last decades, to the deregulation of temporary jobs, often with low levels of pay (OECD, 2012). However, our evidence also shows that the effect increases as earnings grow, indicating that the opportunities of high incomes offered by more flexible labour markets are only (or relatively more) grasped by men. As emphasized by, for example, Leonardi and Pica (2012), this could be the effect of mechanisms related to differences in bargaining power between genders. Women have normally higher turnover rates and higher exposition to discontinuities in labour force participation, related to the extra-work factors (Zabalza and Tzannatos, 1985), particularly in contexts where the asymmetry in the load of family care is remarkable or the welfare state does not provide appropriate support, as it is the case in Eastern Europe (Viitanen, 2007; European Commission, 2009). As a consequence, women experience on the one side lower levels of specific training; on the other side, they are less able to supply the flexibility and adaptability over time and space needed to grasp the best opportunities associated to high risk/returns that temporary positions may provide. In addition, the dismissal of minimum wage provisions favours the opening up of the gender pay gap, by removing the lower bound for low-pay workers for which the minimum wage was binding, and in which women are typically overrepresented (Di Nardo et al., 1996).

Figure 4 Gender earning gap in Eastern Europe and additional effects of labour market institutions (pooled sample, 2007 and 2009)



However, again, the effect of labour market deregulation on the gender gap is not statistically different from zero at the bottom of the distribution. So, weaker minimum wage regulations and more deregulation on the side of temporary jobs do not produce any effect on gender pay differences among the most disadvantaged workers: they seem to be already trapped into a situation where the regulatory framework is not effective or is not able to produce any re-equilibrating effect between genders. As already highlight-ed, the bottom tail of the distribution is almost completely balanced between man and women; the dismissal of the minimum wage then hits equally males and females especially if, as we will see, their distribution into sectors and positions in which the regulation is binding is not remarkably asymmetric.

All these elements seem to basically provide additional aspects to a picture of a duality in the labour market of Eastern European countries, already discussed in Hölscher at al. (2011), which is not only based on differences in the productive attributes of workers or structural factors (Doeringer and and Piore, 1971), but also increasingly depends on the employment status of workers, as a result of on-going or incomplete labour market reforms which generate asymmetries between employment positions (Boeri and Garibaldi, 2007; Belot et al., 2007). Our findings support the view that the disadvantaged pool of workers is so disadvantaged that there is little or no room for gender discrimination and in which labour market institutions, effective in other parts of the distribution, are not able to play a role anymore. And the crisis seems to reinforce this picture. To corroborate this interpretation, we can use some descriptive evidence drawn by our samples, focusing on the differences of workers and job characteristics at the various parts of the distribution, particularly at the lower and top 10% of hourly earnings (for the sake of brevity, figures refer to 2007 only, since the 2009 sample provides very similar evidence). Low-income workers held more temporary positions (12%) compared to the median (9%) and to the top earners (4%); and were almost four and three times more self-employed (30%) than the median (8%) and the high income ones (11%), respectively. This means that in the lowest 10% of earners there is a much higher incidence of workers out of the scope of protective labour market institutions, both due to holding temporary positions in which job protection is normally lower (see Venn, 2009) and to resorting to some sort of subsistence self-employment. As a matter of facts, only 9% of self-employed in the bottom 10% are professionals or technicians, compared to 55% of the top pool; 20% of them is employed in the primary sector, compared to 2% of the top earners. Over 70% of low-income workers lives in rural areas, compared to less than 25% of the high-paid pool; 45% of them works in small firms (less than 10 employees), in which the presence of unions is normally scanty or ineffective, compared to 54% of top earners employed in firms with over 50 employees. Employment in low union and collective bargaining density sectors is remarkably higher for the low earners: agriculture is ten times higher than for the top 10% (21% versus 2.3%); trade services are twice as important as for top earners (18% versus 9%); the share of employed in hotels and restaurants five times bigger (5% and 1%, respectively). On the contrary, their employment in the most regulated sector, industry, is only around 18% compared to over 23%. Similar corroborative evidence is provided by the breakdown of workers into occupations, with low-income workers disproportionately more employed in farming-related tasks. If we look at the composition by gender of employment within sectors in the low-income pool, we also notice that the usually reported vertical segregation tends to fade, especially in the most unions intensive sectors/segments. The share of women employed in industry, within the bottom 10% of earners, is balanced with the share of men (44% and 56%, respectively); similarly, the share of women in permanent jobs is around 55%. This means that, in the low wage area, in the sectors/segments in which unions might play a role in reducing gender discrimination, there is not room for action since the intensive presence of men earning low salaries constraints anti-discriminatory actions based on achieving comparability on a gender basis (Gregory, 1999). This is consistent with our empirical result of unions and collective bargaining being not significant in reducing the gap at the bottom of the distribution.

6 Summary and Final Remarks

In this paper we provide a picture of earnings gender gap in Central and Eastern European countries, shortly before and during the ongoing economic crisis. In particular, we use the quantile regression approach to: (i) highlight the variability of unexplained gender earnings gap and hence of gender discrimination across the earnings distribution within each country; (ii) investigate the effects of labour market institutional settings on the gender gap in different segments of the distribution. All estimations account for potential sample selection bias. Our samples include the 10 Eastern European EU members and refer to 2007 and 2009. We found remarkable differences across countries in terms of: (i) average and median gender (hourly earnings) gap, with Bulgaria, Romania, Slovenia and the Slovak Republic presenting relatively low rates of discrimination and Latvia and Estonia the highest; (ii) patterns of the gender gap along the earnings distribution, with Latvia, Estonia, Hungary and Slovenia facing the strongest glass-ceiling effect; (iii) dynamics of the gap as a reaction to the crisis. Poland, Bulgaria, Latvia and Lithuania demonstrated a decrease of the gender gap; Czech Republic, Hungary, Romania and Slovenia a slight increase. In Estonia and the Slovak Republic the gap remained stable, although the glass-ceiling effect increased.

For the majority of countries and in the pooled sample, the crisis produced a decrease of the gender gap at the bottom end of the distribution. However, this is not good news since our descriptive evidence clearly shows that this was the result of a convergence of male earnings towards the lowest levels of the female workers. Our analysis suggests that deregulation of labour markets, in the form of less regulated temporary jobs and weakened minimum wage provisions, increases the gender earnings gap. On the contrary, higher presence of unions and stronger bargaining coordination reduce the gap. The quantile regressions allow demonstrating that the impact of these institutions is stronger in the upper part of the earnings distribution, that is, for high-pay workers. On the contrary, institutions are less effective in reducing the gap for the low-paid segment of workers, especially during the crisis. The complementary descriptive evidence corroborates the idea that the low-pay workers belong to an area of the labour market that the institutions either cannot reach, or in which they are unable to address the issue of gender discrimination any longer. This contributes to the description of new, complex dualities in the labour market, this time related to the effectiveness or not of institutional settings.

Our findings are not necessarily in contrast with the evidence of pro-inequality effects of declining labour market institutions, especially those unfolding at the lower end of the distribution (see, for example, Lemieux, 2011). All institutional indicators available undoubtedly indicate weakening of labour market institutions over the last two decades, characterized by the evolution of labour market regulations and functioning, especially at the European level (Checchi and Lucifora 2002), towards greater flexibility and more liberalistic models (ILO, 2012). Our results suggest that, at the end of the years 2000, this long-lasting process may have already pushed out of the reach of institutions an important segment of the labour market, at least as regards their capacity to reduce gender disparities. This means, on the policy side, that further waves of liberalisations might not only lead to an increase of the gender gap, but also favour a further polarization of workers in the labour market and expand the area in which the capacity of institutions to address discrimination issues is dampened.

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Appendix

Table A1 Labour market institutional settings in Central and Eastern European countries

	2006			*2008*		
	LM deregulation	***Union Density***	***Wage Coordination***	***LM deregulation***	***Union Density***	***Wage Coordination***
Bulgaria	8.3	21.15	4	8.3	20.09	2
Czech Republic	6.7	18.70	2	6.7	17.42	2
Estonia	6.7	8.20	1	6.7	7.27	1
Hungary	10.0	17.04	2	10.0	16.81	2
Latvia	5.0	17.60	1	5.0	14.80	1
Lithuania	6.7	10.46	1	6.7	8.47	1
Poland	8.9	16.77	1	8.9	15.60	1
Slovak Republic	8.3	20.56	2	8.3	17.17	2
Slovenia	2.2	29.66[a]	4	2.2	29.66	4
Western EU	6.9	36.46	3.6	6.8	34.91	3.3

[a] data refer to 2008, due to missing data for the reference year

Table A2 Sample average characteristics and gender differences (pooled sample, 2007 and 2009)

	2007			*2009*		
	Female	***Male***	***Diff.***	***Female***	***Male***	***Diff.***
Ln(hourly earnings)	1.536	1.672	–0.136***	1.427	1.527	–0.100***
Age	43.73	42.61	1.126***	42.51	41.55	0.964***
Health status	2.332	2.286	0.046***	2.294	2.261	0.033***
Urban	0.477	0.439	0.038***	0.488	0.458	0.029***
Partime	0.058	0.026	0.032***	0.058	0.024	0.034***
Second job	0.020	0.025	–0.058***	0.040	0.070	–0.003***
Status						
Permanent workers	0.827	0.748	0.079***	0.836	0.742	0.093***
Temporary workers	0.091	0.086	0.004*	0.081	0.081	0.001
Self-employed	0.082	0.166	–0.084***	0.083	0.177	–0.094***
Education						
Primary	0.093	0.115	–0.022***	0.084	0.105	–0.022***
Secondary	0.636	0.720	–0.084***	0.598	0.700	–0.102***
Tertiary	0.271	0.165	0.106***	0.318	0.195	0.123***
Firm size						
0 – 10	0.321	0.332	–0.011**	0.295	0.312	–0.017***
11 – 49	0.317	0.320	–0.003	0.342	0.337	0.005
50 and over	0.362	0.347	0.014***	0.363	0.351	0.012**
Sector						
Agriculture	0.046	0.0776	–0.032***	0.0355	0.0795	–0.044***
Industry	0.219	0.320	–0.101***	0.188	0.301	–0.114***
Construction	0.017	0.165	–0.148***	0.0161	0.149	–0.133***
Trade	0.163	0.105	0.058***	0.168	0.108	0.06***
Transport	0.049	0.113	–0.063***	0.051	0.117	–0.067***
Hotels and restaurants	0.046	0.019	0.026***	0.046	0.0220	0.024***
Business services	0.086	0.063	0.024***	0.095	0.0721	0.023***
Other services	0.373	0.137	0.236***	0.401	0.151	0.250***
Occupation						
Managers and Senior officials	0.047	0.068	–0.020***	0.053	0.074	–0.021***
Professional and Technicians	0.365	0.195	0.170***	0.395	0.213	0.182***
Clerks	0.303	0.120	0.184***	0.304	0.129	0.175***
Skilled agricultural and craft workers	0.097	0.339	–0.242***	0.075	0.314	–0.238***
Machine operators	0.068	0.189	–0.121***	0.056	0.182	–0.126***
Elementary occupations	0.119	0.090	0.029***	0.117	0.088	0.029***
Observations	30,796	36,185		30,287	34,175	

Notes: Significance level for the t-test (H0: diff.=0): * p<0.05, ** p<0.01, *** p<0.001

Table A3 OLS estimates, pooled model (2007 and 2009)

	2007				2009			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
male	0.198***	0.171***	0.387***	0.258***	0.176***	0.143***	0.263***	0.214***
	(0.005)	(0.011)	(0.014)	(0.009)	(0.006)	(0.012)	(0.013)	(0.010)
married	0.038***	0.038***	0.034***	0.035***	0.036***	0.036***	0.034***	0.035***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)
age	0.021***	0.021***	0.021***	0.021***	0.025***	0.024***	0.025***	0.025***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)
age^2	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
health	–0.050***	–0.050***	–0.050***	–0.050***	–0.038***	–0.038***	–0.039***	–0.039***
	(0.006)	(0.006)	(0.006)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)
Temp	–0.133***	–0.133***	–0.133***	–0.133***	–0.125***	–0.125***	–0.125***	–0.125***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Self	–0.202***	–0.201***	–0.195***	–0.198***	–0.208***	–0.207***	–0.204***	–0.206***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Secondary	0.067	0.067***	0.067***	0.067***	0.080***	0.080***	0.081***	0.081***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Tertiary	0.311***	0.311***	0.311***	0.311***	0.324***	0.323***	0.324***	0.324***
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
Part	0.039***	0.039***	0.039***	0.040***	0.070***	0.070***	0.071***	0.071***
	(0.012)	(0.012)	(0.012)	(0.012)	(0.011)	(0.011)	(0.011)	(0.011)
sjob	0.157***	0.157***	0.159***	0.158***	0.836***	0.836***	0.835***	0.835***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.033)	(0.033)	(0.033)	(0.033)
Size (11–49)	0.101***	0.101***	0.101***	0.101***	0.127***	0.127***	0.127***	0.127***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Size (over 50)	0.198***	0.197***	0.197***	0.197***	0.233***	0.233***	0.233***	0.233***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
urb	0.086***	0.086***	0.086***	0.086***	0.077***	0.077***	0.077***	0.077***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
LM_dereg	0.010***	0.008***	0.010***	0.010***	–0.094***	–0.096***	–0.094***	–0.094***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
PMD	0.080***	0.080***	0.080***	0.080***	0.793***	0.792***	0.789***	0.791***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.044)	(0.044)	(0.044)	(0.044)
UR	–0.101***	–0.101***	–0.100***	–0.100***	–0.235***	–0.235***	–0.234***	–0.234***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.010)	(0.010)	(0.010)	(0.010)
UD	0.094***	0.094***	0.099***	0.094***	0.203***	0.203***	0.204***	0.203***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.012)	(0.012)	(0.012)	(0.012)
W_coord	–0.393***	–0.393***	–0.392***	–0.376***	–1.641***	–1.639***	–1.630***	–1.626***
	(0.005)	(0.005)	(0.005)	(0.006)	(0.094)	(0.094)	(0.093)	(0.093)
LM_dereg_male	–	0.004***	–	–	–	0.005***	–	–
		(0.001)				(0.001)		
UD_male	–	–	–0.010***	–	–	–	–0.005***	–
			(0.001)				(0.001)	
W_coord_male	–	–	–	–0.029***	–	–	–	–0.019***
				(0.003)				(0.004)
_cons	–0.337***	–0.323***	–0.432***	–0.366***	–1.125***	–1.104***	–1.160***	–1.143***
	(0.069)	(0.069)	(0.069)	(0.069)	(0.174)	(0.174)	(0.174)	(0.174)
Sector/occ/country dummies	yes	yes	yes	yes	yes	yes	yes	yes
Sample-selection corr.	yes	yes	yes	yes	yes	yes	yes	yes
Obs	66981	66981	66981	66981	64462	64462	64462	64462
Adj. R-Sq	0.452	0.452	0.454	0.453	0.508	0.508	0.508	0.508

Table A4 Quantile regression estimates, pooled model (2007 and 2009): Labour market deregulation and earnings gap

	2007			*2009*		
	θ =.10	θ = 50	θ = 90	θ =.10	θ = 50	θ = 90
male	0.164***	0.172***	0.161***	0.137***	0.142***	0.202***
	(0.020)	(0.011)	(0.017)	(0.021)	(0.012)	(0.018)
married	0.020***	0.035***	0.031***	0.025***	0.031***	0.027***
	(0.007)	(0.004)	(0.007)	(0.007)	(0.005)	(0.007)
age	0.021***	0.023***	0.021***	0.016***	0.024***	0.039***
	(0.005)	(0.004)	(0.006)	(0.005)	(0.003)	(0.005)
age^2	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
health	–0.045***	–0.049***	–0.053***	–0.038***	–0.040***	–0.053***
	(0.009)	(0.006)	(0.009)	(0.006)	(0.005)	(0.007)
Temp	–0.159***	–0.138***	–0.088***	–0.152***	–0.126***	–0.089***
	(0.010)	(0.007)	(0.012)	(0.011)	(0.007)	(0.013)
Self	–0.629***	–0.129***	0.083***	–0.696***	–0.144***	0.087***
	(0.026)	(0.010)	(0.013)	(0.023)	(0.008)	(0.013)
Secondary	0.045***	0.086***	0.091***	0.025**	0.096***	0.157***
	(0.013)	(0.008)	(0.014)	(0.010)	(0.009)	(0.014)
Tertiary	0.245***	0.347***	0.394***	0.182***	0.349***	0.495***
	(0.020)	(0.013)	(0.021)	(0.018)	(0.014)	(0.021)
Part	–0.093***	0.043***	0.209***	–0.049***	0.060***	0.211***
	(0.017)	(0.012)	(0.019)	(0.016)	(0.012)	(0.019)
sjob	0.167***	0.102***	0.084***	0.744***	0.766***	0.926***
	(0.008)	(0.005)	(0.007)	(0.045)	(0.049)	(0.060)
Size (11–49)	0.140***	0.097***	0.041***	0.155***	0.113***	0.057***
	(0.008)	(0.005)	(0.008)	(0.007)	(0.005)	(0.008)
Size (over 50)	0.240***	0.192***	0.127***	0.264***	0.210***	0.145***
	(0.008)	(0.005)	(0.009)	(0.008)	(0.006)	(0.008)
urban	0.060***	0.082***	0.103***	0.054***	0.077***	0.099***
	(0.007)	(0.005)	(0.008)	(0.006)	(0.004)	(0.008)
LM_dereg	0.014***	–0.007***	–0.009**	–0.101***	–0.104***	–0.101***
	(0.004)	(0.002)	(0.004)	(0.003)	(0.002)	(0.003)
PMD	0.056***	0.091***	0.126***	1.329***	0.747***	0.376***
	(0.007)	(0.005)	(0.008)	(0.071)	(0.053)	(0.064)
UR	–0.105***	–0.096***	–0.089***	–0.374***	–0.225***	–0.120***
	(0.002)	(0.002)	(0.002)	(0.016)	(0.012)	(0.014)
UD	0.115***	0.094***	0.071***	0.353***	0.187***	0.083***
	(0.002)	(0.001)	(0.002)	(0.019)	(0.014)	(0.017)
W_coord	–0.517***	–0.405***	–0.278***	–2.847***	–1.513***	–0.694***
	(0.009)	(0.006)	(0.010)	(0.151)	(0.112)	(0.136)
LM_dereg_male	–0.002	0.006***	0.012***	–0.003	0.007***	0.006***
	(0.003)	(0.001)	(0.002)	(0.002)	(0.001)	(0.002)
_cons	–0.510***	–0.227***	0.209*	–3.543***	–1.293***	0.087
	(0.097)	(0.071)	(0.109)	(0.275)	(0.211)	(0.255)
Sector/occup/country dummies	yes	yes	yes	yes	yes	yes
Sample-selection correction	yes	yes	yes	yes	yes	yes
Test F [q10=q50=q90]: LMD_male	7.79*** [2, 66943]			8.00*** [2, 64424]		
Obs	66981	66981	66981	64462	64462	64462
Adj. R-Sq	0.297	0.307	0.286	0.315	0.343	0.333

Notes: Robust standard errors in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. W: coord, LM_dereg, UD, PMD: lagged one year.

Table A5 Quantile regression estimates, pooled model (2007 and 2009): Union density and earnings gap

	2007			*2009*		
	θ =.10	θ = 50	θ = 90	θ =.10	θ = 50	θ = 90
male	0.286*** (0.021)	0.429*** (0.014)	0.500*** (0.023)	0.136*** (0.021)	0.316*** (0.014)	0.375*** (0.020)
married	0.019*** (0.007)	0.030*** (0.005)	0.026*** (0.008)	0.024*** (0.006)	0.028*** (0.004)	0.025*** (0.007)
age	0.019*** (0.005)	0.022*** (0.003)	0.018*** (0.006)	0.016*** (0.005)	0.026*** (0.003)	0.040*** (0.005)
age^2	–0.000*** (0.000)	–0.000*** (0.000)	–0.000** (0.000)	–0.000** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)
health	–0.043*** (0.008)	–0.048*** (0.005)	–0.048*** (0.009)	–0.037*** (0.007)	–0.041*** (0.005)	–0.052*** (0.007)
Temp	–0.161*** (0.010)	–0.139*** (0.007)	–0.087*** (0.012)	–0.152*** (0.010)	–0.127*** (0.007)	–0.091*** (0.013)
Self	–0.624*** (0.025)	–0.120*** (0.010)	0.091*** (0.011)	–0.697*** (0.022)	–0.142*** (0.009)	0.088*** (0.013)
Secondary	0.050*** (0.012)	0.087*** (0.007)	0.090*** (0.013)	0.022* (0.011)	0.098*** (0.008)	0.151*** (0.013)
Tertiary	0.246*** (0.018)	0.348*** (0.011)	0.393*** (0.021)	0.179*** (0.019)	0.352*** (0.013)	0.494*** (0.020)
Part	–0.086*** (0.015)	0.048*** (0.011)	0.208*** (0.020)	–0.047*** (0.014)	0.057*** (0.011)	0.219*** (0.020)
sjob	0.168*** (0.008)	0.102*** (0.005)	0.083*** (0.007)	0.746*** (0.048)	0.763*** (0.053)	0.914*** (0.056)
Size (11–49)	0.139*** (0.008)	0.097*** (0.005)	0.041*** (0.008)	0.153*** (0.008)	0.112*** (0.005)	0.058*** (0.008)
Size (over 50)	0.238*** (0.008)	0.193*** (0.005)	0.126*** (0.008)	0.262*** (0.008)	0.209*** (0.006)	0.143*** (0.008)
urb	0.059*** (0.007)	0.081*** (0.005)	0.099*** (0.008)	0.054*** (0.006)	0.075*** (0.005)	0.098*** (0.008)
LM_dereg	0.010*** (0.003)	–0.004* (0.002)	–0.000 (0.004)	–0.103*** (0.003)	–0.101*** (0.002)	–0.097*** (0.003)
PMD	0.059*** (0.008)	0.092*** (0.004)	0.125*** (0.007)	1.339*** (0.079)	0.740*** (0.052)	0.361*** (0.077)
UR	–0.103*** (0.002)	–0.096*** (0.002)	–0.090*** (0.002)	–0.376*** (0.018)	–0.223*** (0.012)	–0.117*** (0.018)
UD	0.117*** (0.002)	0.100*** (0.001)	0.079*** (0.002)	0.356*** (0.021)	0.188*** (0.014)	0.084*** (0.020)
W_coord	–0.511*** (0.008)	–0.404*** (0.006)	–0.277*** (0.008)	–2.866*** (0.169)	–1.494*** (0.113)	–0.667*** (0.165)
UD_male	–0.007*** (0.001)	–0.011*** (0.001)	–0.013*** (0.001)	–0.001 (0.001)	–0.007*** (0.001)	–0.007*** (0.001)
_cons	–0.133 (0.096)	0.171** (0.068)	0.645*** (0.108)	–3.558*** (0.303)	–1.377*** (0.215)	0.033 (0.304)
Sector/occ dummies	yes	yes	yes	yes	yes	yes
Sample-selection correction	yes	yes	yes	yes	yes	yes
Test F [q10=q50=q90]: UD_male	11.76*** [2, 66943]			12.66*** [2, 64424]		
Obs	66981	66981	66981	64462	64462	64462
Adj. R-Sq	0.298	0.309	0.288	0.315	0.344	0.344

Notes: Robust standard errors in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. W: coord, LM_dereg, UD, PMD: lagged one year.

Table A6 Quantile regression estimates, pooled model (2007 and 2009): Wage coordination and earnings gap

	2007			*2009*		
	θ =.10	θ = 50	θ = 90	θ =.10	θ = 50	θ = 90
male	0.210*** (0.014)	0.291*** (0.010)	0.338*** (0.015)	0.111*** (0.015)	0.259*** (0.010)	0.313*** (0.014)
married	0.020*** (0.007)	0.032*** (0.005)	0.027*** (0.008)	0.024*** (0.007)	0.028*** (0.004)	0.025*** (0.007)
age	0.021*** (0.005)	0.022*** (0.004)	0.019*** (0.006)	0.016*** (0.004)	0.025*** (0.003)	0.037*** (0.005)
age^2	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)	–0.000*** (0.000)
health	–0.047*** (0.009)	–0.049*** (0.006)	–0.050*** (0.009)	–0.037*** (0.006)	–0.040*** (0.005)	–0.050*** (0.008)
Temp	–0.162*** (0.010)	–0.139*** (0.007)	–0.092*** (0.011)	–0.152*** (0.011)	–0.125*** (0.007)	–0.093 (0.013)
Self	–0.620*** (0.025)	–0.124*** (0.010)	0.084*** (0.012)	–0.698*** (0.021)	–0.145*** (0.008)	0.088*** (0.012)
Secondary	0.049*** (0.013)	0.087*** (0.008)	0.092*** (0.013)	0.023** (0.010)	0.098*** (0.009)	0.150*** (0.014)
Tertiary	0.244 (0.019)	0.348 (0.013)	0.393*** (0.021)	0.181*** (0.017)	0.351*** (0.013)	0.490*** (0.021)
Part	–0.087*** (0.016)	0.049*** (0.011)	0.210*** (0.021)	–0.049*** (0.015)	0.058*** (0.012)	0.219*** (0.018)
sjob	0.168*** (0.008)	0.102*** (0.005)	0.088*** (0.007)	0.749*** (0.045)	0.766*** (0.053)	0.931*** (0.063)
Size (11–49)	0.141*** (0.008)	0.097*** (0.005)	0.039*** (0.008)	0.156*** (0.008)	0.113*** (0.005)	0.056*** (0.008)
Size (over 50)	0.239 (0.009)	0.192*** (0.006)	0.125*** (0.009)	0.264*** (0.008)	0.211*** (0.006)	0.141*** (0.008)
urban	0.062 (0.007)	0.081*** (0.005)	0.098*** (0.008)	0.054*** (0.006)	0.077*** (0.005)	0.098*** (0.008)
LM_dereg	0.012*** (0.003)	–0.005** (0.002)	–0.002 (0.004)	–0.102*** (0.003)	–0.101*** (0.002)	–0.097*** (0.003)
PMD	0.060*** (0.007)	0.090*** (0.005)	0.123*** (0.007)	1.331*** (0.072)	0.742*** (0.050)	0.364*** (0.071)
UR	–0.105*** (0.002)	–0.095*** (0.001)	–0.089*** (0.002)	–0.374*** (0.016)	–0.223*** (0.011)	–0.117 (0.016)
UD	0.115*** (0.002)	0.094*** (0.001)	0.070*** (0.002)	0.354*** (0.019)	0.186*** (0.013)	0.081*** (0.019)
W_coord	–0.503*** (0.009)	–0.388*** (0.007)	–0.247*** (0.010)	–2.851*** (0.154)	–1.486*** (0.107)	–0.659*** (0.151)
W_coord_male	–0.028*** (0.005)	–0.036*** (0.003)	–0.045*** (0.005)	0.003 (0.007)	–0.032*** (0.004)	–0.032*** (0.005)
_cons	–0.146 (0.097)	0.260*** (0.074)	0.732*** (0.117)	–3.532*** (0.272)	–1.353*** (0.199)	0.104 (0.291)
Sector/occ dummies	yes	yes	yes	yes	yes	yes
Sample-selection correction	yes	yes	yes	yes	yes	yes
Test F [q10=q50=q90]: W_coord_male	2.85* [2, 66943]			13.40*** [2, 64424]		
Obs	66981	66981	66981	64462	64462	64462
Adj. R-Sq	0.297	0.308	0.287	0.315	0.344	0.333

Notes: Robust standard errors in parentheses. ***, ** and * denote significance at the 1, 5 and 10 percent level, respectively. W: coord, LM_dereg, UD, PMD: lagged one year.