



Policies, Skills and Earnings: How Educational Inequality Affects Earnings Inequality

Journal:	<i>Socio-Economic Review</i>
Manuscript ID	SER-2015-0140.R3
Manuscript Type:	Article
SER Keywords:	education, inequality, human capital, income distribution
 JEL Classification :	I24, I28

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Policies, Skills and Earnings: How Educational Inequality Affects Earnings Inequality

Final version, October 2016

Abstract

We study the impact of dispersions in education (both in student test scores and final educational attainment) on earnings inequality, in a country-cohort design. Neo-classical economic theory would predict a positive association between skill inequality (as measured in student test scores) and earnings inequality, while educational attainment inequality adds little on top of skills inequality. A sociological theory of social closure, however, argues that inequality in educational attainment is more important than skills inequality in the prediction of earnings inequality. Using educational policies as instruments, we find causal effects of skills inequality and educational attainment inequality, suggesting that a simple human capital model is insufficient to explain rising earnings inequalities. Nevertheless, skills inequality appeared a more important predictor of earnings inequality than educational attainment inequality. Some educational policy reforms (like public preschool provision or introducing standardised tests) led to reduced educational dispersions, and thereby reduced earnings inequality in adulthood.

Keywords: education, inequality, human capital, income distribution

Jel codes: I24 Education and Inequality - I28 Government Policy

9961 words (including text, tables, figures, references, appendices)

1. Introduction

The relationship between educational distributions and labour market distributions has been at the core of economic and sociological research. Society gets increasingly stratified on the basis of education. Economic theories and structural-functionalist allies in sociology have explained the increased importance of education through increased complexity work and the resulting higher demand for higher-level skills. Theories of social closure argue that education is increasingly rewarded for other reasons than simply rising demand for skills; education has become a core mechanism to create and maintain boundaries between social groups.

Despite the huge attention for the school-work relationship, the literature has been primarily concerned with differences in *average* educational attainment and achievement and its relationship to the *average* probability to find a higher-class occupation, or earn higher salaries. Much less attention has been paid to *dispersions* in educational distributions, and how these relate to *dispersions* in labour market outcomes such as earnings. This means that dispersions in education and in the labour market have been analyzed mostly in isolation. This is unfortunate for both scientific and political reasons. From a scientific perspective, the divorced studies of educational and earnings inequalities fail to acknowledge that one potential explanation of earnings inequalities may be found in the unequal distribution of education and skills. While many studies have related the rising earnings inequalities to rising gaps among educational categories (e.g. Goldin and Katz 2009; Hout 2012), little research has directly examined the dispersion in earnings in relation to the dispersion in skills and qualification levels. Moreover, while most studies that examined the relationship between education and labour market outcomes focused on educational attainment (e.g. in terms of qualification levels or years of schooling attained), recent scholarship argues that the quality of education in terms of skill assessments is of greater importance for economic growth than the quantity of education in terms of average years of schooling completed (Hanushek and Woessmann 2008). It is important to examine the relevance of quality and quantity of education also from an inequality perspective. Different theoretical perspectives emphasize different aspects of education to be relevant for social stratification. Some theories emphasize the productivity-enhancing skills that schooling provides, while others put more weight on the formal qualifications that provide access to desirable positions. To further scrutinize these different theoretical perspectives, it is important to apply them to inequality. It may help us not only to understand who gets ahead in the stratification order, but also how societal inequality is affected by inequalities in the various elements of education.

From this perspective it is important to know what is more important for the determination of earnings inequality: inequality in skills (student test scores) or inequality in educational

1 attainment? We derive hypotheses from standard neo-classical economics and sociological closure
2 theory, which differ in the importance that may be attached to either skills or educational
3 attainment. Thus, examining both skills and attainment helps us to understand more about how
4 mechanisms about returns to education have contributed to the rising inequality in earnings in many
5 societies.
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10 From a political perspective it is important to understand the relationships between
11 educational and income inequalities because policies that are mainly oriented towards the
12 manipulation of educational distributions can have wider repercussions on the social stratification in
13 adult life. Furthermore, for political debates it is important to study the effectiveness of policies in
14 affecting the dispersions in achievement, and/or in educational attainment.
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18 We contribute to a better understanding of the associations between educational policies,
19 educational distributions and earnings distributions by combining various datasets. We match
20 comparative mathematics achievement data (as indicator of the 'quality' of education) with
21 educational attainments (as a proxy for the 'quantity' of education), labour force participation and
22 earnings in the adult population. The match exploits combinations of cohort, country and gender,
23 and for each combination we have calculated educational and earnings inequality measures. We
24 show that educational inequality (in terms of dispersion) is correlated with earnings inequality thirty
25 years later, with 'quality' being more relevant than 'quantity' in shaping earnings inequality.
26 Furthermore, to investigate the importance of educational policies, we have matched these
27 population data (combining cohorts and countries) with data on educational policies, including
28 policies on compulsory education, school and/or teachers' autonomy, and tracking age (from Braga
29 et al. 2013). We employ instrumental variable models to validate that inequality in education
30 (measured in terms of both quality and quantity) affects earnings inequality. In addition, since
31 educational inequality responds to educational reforms, we are able to identify some educational
32 policies (like later entry into compulsory education or the introduction of standardised tests) capable
33 to reduce earnings inequalities once the same cohort is in the labour market.
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45 As a replication exercise, we match data from a survey conducted in 2000 on 15-year old
46 youngsters math test score with the corresponding cohort in the same country 12 years later. For
47 both groups of individuals we show that inequality in skills accounts for a significant portion of
48 observed earnings inequality.
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53 **2. Theoretical background: Educational distributions and earnings distributions**

54 *Mechanisms linking education and labour market outcomes*

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Education and labour market outcomes are well-known to be strongly correlated: those with more schooling and more advanced skill levels have higher earnings, higher occupational status, and preferable employment relations than those with less education. Education is therefore generally considered as one of the main determinants of the position people take in the social structure.

Studies often relate the rising earnings inequalities witnessed in many societies to rising returns to educational qualifications (although also within-group inequality has increased). According to functionalist sociological theory of modernization, and their allies from neo-classical economics, labour market outcomes are increasingly stratified on the basis of education because the labour market becomes more complex through technological developments, and rising demand for skills increases the price of labour (Bell 1974; Davis & Moore 1945; Goldin & Katz 2009). Although this model assumes that rising educational gradients correspond to a rising return on skills (due to increased productivity created by skill-biased technological change), empirical research typically examines earnings differences between qualification levels. But returns to qualification levels could also correspond to other mechanisms why education is rewarded. In particular, the sociological theory of social closure argues that qualifications are rewarded for reasons other than productivity-enhancing skills (Van de Werfhorst 2011; Di Stasio et al. 2016). Following Solga (2014) we therefore consider it relevant to study the relationship between educational distributions and income inequality from both the neo-classical/functionalist and closure perspectives.

The neo-classical/functionalist model assumes education is rewarded because of the skills that are associated to it. Within neo-classical economics, there are two arguments that call for a very strong relationship between skills and educational attainment. First, skills could be produced in the educational system, as standard human capital theory argues. Second, education sorts on skills that students already have before they go into schooling (Weiss 1995). Both perspectives hold that reward on the basis of education is efficient as education indicates the skills that people hold. The neo-classical model holds that sorting in education on the basis of ability is perfect, where further investments in education depend on expected returns and the ability students have to complete higher levels of schooling. So observing a person's skills should be highly informative of the human capital s/he commands.

The baseline argument of the closure perspective, by contrast, is that formal or informal rules are governing the placement of people in the stratification order. This perspective originates from a Weberian theory on stratification, which holds that advantaged groups search for ways to secure their advantage by norms and regulations (Brown 1995). *Informal norms* can, for instance, correspond to the cultural background of people of different qualification levels fitting different sorts of occupations. High-level occupations are allocated to graduates from tertiary education not only because they have skills to carry out the tasks, but also because they display affinity with the

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2 dominant cultural codes in society that is required to perform well in those functions (Bourdieu
3 1998; Rivera 2012). *Formal regulations* correspond to closure practices that govern the formal
4 accessibility of occupations, like licensing and certification (Bol 2014; Weeden 2002). The supply
5 of workers can then be controlled, creating rents in the form of earnings above the market value of
6 skills. As a summary of the closure perspective, Brown (2001: 20) writes, “credential requirements
7 for jobs are less concerned with concrete work skills than with demanding that recruits hold similar,
8 school-taught cultural dispositions to incumbents of positions”. Importantly, the closure perspective
9 holds, just like the functionalist/neo-classical perspective, that education is increasingly rewarded in
10 the labour market, but for different reasons.

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12 In line with closure theory, neo-institutionalists furthermore argue that the societal change in
13 relation to the ‘schooled society’ involves more than a simple school-to-work allocation process
14 (Baker 2014; Meyer and Ramirez 2000). The institutionalization of education is partly based on a
15 ‘myth’ concerning the relationship between educational attainment and individual marginal
16 productivity (Schofer and Meyer 2005). Because institutionalization also implies a social
17 construction of the value of education, an empirically demonstrable value of education can emerge
18 because we all believe in the myth and organize society on this basis. The neo-classical economic
19 model can be criticised from this perspective because educational earnings returns are both the
20 outcome of study and the (faulty) mechanism of marginal productivity explaining it.

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22 In order to differentiate between neo-classical and closure models of stratification, it is
23 essential to incorporate separate measures of skills (as indicator of the quality of education) and
24 educational attainment (as indicator of the quantity of education). Neo-classical economists show
25 that skills explain a large fraction of why education is rewarded on the labour market (Blau and
26 Kahn 2005; Leuven et al. 2004). Critics of the neo-classical model, however, have demonstrated
27 that educational gradients in earnings are only marginally reduced when skills are taken into
28 account (Bowles and Gintis 2002).

29 30 ***Education inequality and earnings inequality***

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32 The positive relationships between cognitive skills, educational attainment and earnings imply that
33 earnings inequality can be partly attributed to dispersions in educational quality and quantity. Blau
34 and Kahn (2005) claim that the greater dispersion of cognitive test scores in the United States plays
35 a part in explaining higher U.S. wage inequality. Competences also provide an autonomous
36 contribution to observed inequality, conditional on identical school attainment (Green and Riddell
37 2003). Freeman and Schettkat (2001) show that the U.S. are characterised by greater inequality in
38 adult competences than Germany, which is reflected in greater inequality in earnings. Using a wider
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1 country coverage and the more recent PIAAC data, Solga (2014) confirmed a positive relation
2 between inequality in adult competences and income inequality.
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5 Some studies have linked student competences assessed during high school with income
6 inequality. Bedard and Ferrall (2003) show a positive association between inequality in
7 mathematics performance during high school and wage inequality in adulthood of the same birth
8 cohorts. Moreover, the association was hardly affected by including trade union density as an
9 indicator of (wage-compressing) wage centralization. Busemeyer (2015) examined skill inequality
10 in terms of the slope by parental background, rather than in terms of dispersions. No statistical
11 association was found between inequality of educational opportunity in PISA test scores among 15-
12 year-old students and income inequality taken from OECD data.
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18 The two theoretical approaches summarized above have a different perspective on the
19 relevance of competences and educational attainment. According to the neo-classical/functionalist
20 model, educational attainment strongly overlaps with measured skills, so that inequality in skills
21 should be more strongly predictive of the level of earnings inequality in adulthood than inequality
22 in attained years of schooling (*hypothesis 1*). According to the closure theory, by contrast, the
23 possession of skills is not sufficient to achieve high incomes. Skills are only useful to the extent that
24 they have been ‘translated’ in attainment levels that are used to govern the allocation of workers to
25 the occupational structure. Controlling for inequality in educational attainment, skills inequality
26 should therefore be of minor importance for the explanation of earnings inequality in adulthood
27 (*hypothesis 2*).
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35 The two hypotheses follow from very strong versions of the neo-classical/functionalist and
36 closure perspectives. In the strict version of the neo-classical/functionalist theory a person’s human
37 capital is assumed to perfectly overlap with his/her measured cognitive skills (due to reduced cost
38 and/or increased productivity for abler individuals), and the strict version of the closure theory
39 assumes that the only reason that qualification levels are rewarded (conditional on skills) is that
40 informal and formal rules are used. Although these are strong assumptions, they help us to interpret
41 the relative weight of skills inequality and educational attainment inequality. So, if skills inequality
42 would be highly predictive of earnings inequality in adulthood, even after including attainment
43 inequality in the empirical model, it is unlikely that closure theory is fully able to explain rising
44 earnings inequalities in Western societies. Likewise, if attainment inequality is predominant in the
45 explanation of earnings inequality, the least we can say is that the ‘cognitive component of
46 schooling’ does not fully explain how educational inequality is associated to earnings inequality
47 (see also Bowles and Gintis 2002 for a similar argument).
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57 There is another justification for our rather stringent interpretation of these broad theoretical
58 approaches. Rising earnings inequalities could be explained from a neo-classical educational
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1 perspective by either an increase of the price for the skills people have, or a change in the labour
2 supply at different education levels. Both could work in the same, or in opposite directions (Goldin
3 and Katz 2009). For instance, the price for skills may go up due to technological innovation
4 (increasing education gaps), while the supply of higher levels of education also increases
5 (counterbalancing this increase). However, it is plausible that people adjust their educational
6 attainment in response to changing demand more easily than their level of skills. If technology
7 grows, this may lead more easily to behavioural adjustments leading to higher participation rates in
8 (tertiary) education than to higher skill levels. This could imply that the effect of skill inequality on
9 earnings inequality more directly captures skill prices (largely unaffected by the conflation of
10 supply), while the effect of educational attainment inequality also captures other, non-human capital
11 returns to education. Educational attainment inequality may even reduce (because the dispersion in
12 educational attainment reduces with expansion, Meschi and Scervini 2014) while the skill price
13 (and earnings inequality) increases, leading to stronger effects of skill inequality on earnings
14 inequality in the neo-classical model.
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25 26 27 *Education policies and inequalities*

28 While a positive association is expected between educational and income distributions, it is
29 worthwhile to be more explicit on how educational distributions are affected by education policies.
30 Some of the cited studies have looked at education policies. Busemeyer (2015) for instance, showed
31 that the size of the vocational education and training sector and public spending on higher education
32 were negatively associated to the level of income inequality in a country (see also Busemeyer and
33 Iversen 2012). However, that study has not examined the influence of policies on educational
34 distributions, as one avenue through which income inequality might be affected.
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40 Unlike previous studies, we approach educational policies as instruments of educational
41 distributions, so we can estimate the causal effects of educational distributions on earnings
42 inequality. Using educational policies as instruments is desirable because different policies can
43 create different skill equilibria and earnings distributions (Iversen and Stephens 2008). As recent
44 contributions to the Varieties of Capitalism field emphasize, educational policies can differ with
45 regard to the stratification and de-commodification they create in the educational system, which
46 results in varying levels of educational inequality (Busemeyer 2015). From a methodological
47 perspective, instruments are desirable as there is potential endogeneity, since more talented
48 individuals may possess higher level of competences (as well as achieve higher educational
49 attainments) and obtain higher earnings. In the absence of credible instruments, it is hard to accept a
50 causal interpretation of previous results.
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The impact of policies on *educational* distributions have been widely studied by now (Brunello and Checchi 2007; Hanushek and Woessmann 2011; Rindermann and Ceci 2009). Examples of policies that are associated to inequalities are the provision of pre-school education (Rindermann and Ceci 2009; Schütz et al. 2008; Heckman 2006), standardised testing (Horn 2009), tracking age (Bol et al. 2014), school accountability (Hanushek and Raymond 2004; Jennings and Sohn 2014) and the length of compulsory education (Leuven et al. 2010; Meschi and Scervini 2014). We focus on these same policies, which also enables us to study the indirect impact of these policies on *earnings* inequality.

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Because we treat policies as instruments for educational distributions, we do not pay attention to other (labour market, welfare) policies that affect earnings inequality. Such policies are less important for the production of educational inequalities, unless one sees this from an intergenerational perspective (because welfare policies affect the economic resources of parents). Educational policies can also be endogenous to educational inequalities, since government can be induced to change the organisation of the educational system by observing increasing inequality in achievements and/or attainments. Braga et al. (2013) tested the exogeneity of educational reform measures by regressing current values of the reforms against past values of the inequality in schooling, finding no statistical association. They also showed that educational reforms correlate with the left-right inclination of governments, and provided IV estimation where variations in educational reforms (induced by electoral changes) caused changes in educational inequality. It should, however, be acknowledged that education policies can also be part of larger constellations of agreements between the state and employer and employee organizations, including labour market policies (Busemeyer 2015; Iversen and Soskice 2001). For our analysis this poses less of a problem, as we are not aiming to endogenize education policies, but educational distributions.

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3. Empirical strategy

In the absence of cross-nationally comparative individual-level longitudinal data with school-age competences, schooling attainment and further labour market careers we pursue an alternative strategy of country/cohort analysis, matching aggregate inequality measures of competences, schooling and earnings based on the birth year of the relevant cohort. We use data for multiple cohorts born since 1950, observed at various ages.

By indicating with $I(x)$ a generic inequality indicator, we study the how the inequality in y (earnings) will depend on the inequality in both quality q (ability tests) and quantity h (educational attainment).

$$I(y_{jt}) = \delta_j + \delta_t + \alpha I(h_{jt}) + \beta I(q_{jt}) + \omega_{jt} \quad (1)$$

where δ_j and δ_t are country/cohort/survey fixed effects capturing any other sort of systemic earnings inequality variation beyond education, while α and β measure the association between (the distribution of) various dimensions of human capital (quantity and quality) and earnings inequality. If h and/or q are measured well in advance with respect to y (in our case h is measured at the end of schooling by the maximal educational attainment, q is measured at the age of 14 or 15, while y is measured alternatively at the ages of 28, 44 and 59), one is tempted to provide a causal interpretation of statements like “a reduction in inequality in test scores reduces income inequality by a magnitude β ”. However, unobservable components at country level (like competitiveness, solidarity, ethnic fractionalisation and so on) may drive both dimension of inequality, leading to biased estimates of the relevant coefficients. Accounting for this possibility, one can resort to an instrumental variable strategy to estimate equation (1) leading to

$$\begin{cases} I(h_{jt}) = a_j + a_t + \mathbf{b}'_j \mathbf{Z}_{jt} + e_{jt} \\ I(q_{jt}) = c_j + c_t + \mathbf{d}'_j \mathbf{Z}_{jt} + g_{jt} \\ I(y_{jt}) = \delta_j + \delta_t + \alpha \hat{I}(h_{jt}) + \beta \hat{I}(q_{jt}) + \omega_{jt} \end{cases} \quad (2)$$

where the educational inequality measures in equation (1) are replaced by their projections obtained from a vector of (supposedly) exogenous variables pertaining reforms in the educational sectors affecting the relevant birth cohorts.¹ We thus exploit both geographical and temporal variations in educational reforms by government to obtain unbiased estimates of the causal impact of educational inequality on earnings inequality.

4. Dataset 1: IMS-ECHP-SILC and educational reforms

Data on students' competences are obtained from three early surveys on mathematical competences of (around) 14-year-old students conducted in past decades (FIMS 1964 was the *First* International Mathematics Study testing mathematical competences of students born in or around 1950; the *Second* (SIMS 1980-82) tested students born in or around 1966 and the *Third* (TIMSS 1995) tested students born in or around 1981, all assumed to be nationally representative).² We thus start with a population composed by three birth cohorts, born in 1950, 1966 and 1981 respectively, in countries that participated to the student assessment surveys. Data on schooling and labour market outcomes for the same cohorts can be obtained from representative samples of the corresponding population at later stages. However, if observed at the same point in time, we would be confounding cohorts and age effects (namely, older cohorts may be characterised by higher (possibly lower, due to

1 natural decay) level of competences and earnings inequalities). For this reason, we use two
2 available datasets existing at European level and reporting data on earnings and incomes: the
3 European Community Household Panel (ECHP), which started in 1994³, and the European Union
4 Statistics on Income and Living Conditions (EU-SILC), which started in 2004 and is updated
5 annually.⁴ We selected the 1994 ECHP wave because it was the earliest available, while we resorted
6 to the 2009 SILC wave because it was the first survey reporting consistent information on gross
7 incomes for all participating countries. In table 1 we show the matching rule we followed to
8 construct our dataset. Take for example the students born in 1950 and tested in 1964 among others
9 in six European countries (Belgium⁵, Finland, France, Germany, Netherlands and United Kingdom).
10 The same birth cohort can be observed in the labour market around the age of 44 in 1994, thanks to
11 ECHP (Finnish data not available, since Finland entered in 1996), and again at age 59, using the
12 SILC data. The same logic applies to the other cohorts.

21 [Table 1 here]

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27 Labour earnings data are obtained summing earnings from dependent employment and earnings
28 from self-employment, and excluding cases of negative values. In the ECHP all values (except
29 France) are net of taxes, while in the SILC we use gross values (this is one reason to add a survey
30 fixed effect). Negative earnings values exist in the case of self-employed, they have been converted
31 into zeros in order to allow the calculation of inequality. Individuals without earnings (housewives,
32 unemployed, out of labour force) report zero values, but have been retained in our sample since
33 labour market participation is potentially endogenous and definitely correlated with educational
34 attainment and possession of competences.

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40 Years of education have been computed from maximum educational attainment according to
41 the ISCED classification converted into years by using legal duration.⁶ Mathematics test scores, the
42 only domain that has been tested since the 1964 FIMS, are calculated for the FIMS, SIMS and
43 TIMSS data based on the number of correct answers to multiple-choice items.⁷ In order to partially
44 account for gender composition and increase the degrees of freedom, we have computed these
45 inequality measures separately for males and females.

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50 The country inequality measures (averaged over genders) are reported in table A1. Overall
51 we possess an unbalanced panel covering 20 countries with 82 observations (41 country/cohort \times 2
52 genders). Figure 1 shows the plot of the relevant data for the 20 countries, while table 2 reports the
53 descriptive statistics for the relevant variables. We can observe that there is a modest positive
54 correlation (0.18) between inequality in quantity and inequality in quality of education for the
55 country/gender/cohort cell available (north-west panel).⁸ Inequality in the quality of education has a
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2 modest (but insignificant) correlation with earnings inequality (computed over employees – north-
3 east panel), while there is no bivariate correlation between inequality in educational attainment and
4 earnings inequality. In the last south-east panel we contrast earnings inequality for dependent
5 employees and for total employment. The correlation among the two inequality measures is
6 considerable (0.43).
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11 [Table 2 here]
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14 Potentially confounding factors are controlled by means of corresponding dummies (gender, birth
15 year, age, country and survey). For gender this refers to the share in the data of men in the country-
16 cohort-age group.
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21 [Figure 1 here]
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25 We do not have *a priori* about which is the most appropriate economic inequality measure to be
26 used in this analysis, since each index captures different dimensions of the underlying distributions.
27 We have decided to focus on the Gini index as our relevant measure of inequality, since it is the
28 most commonly used measure for income inequality. Table 3 shows results with fixed effects, still
29 without examining education policies. The table shows positive effects of inequalities in quantity
30 and quality of human capital (see columns (1) and (2) of table 3). Clustering errors weakens the
31 statistical significance of inequality in test scores (columns (3) and (4) of table 3).
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37 If we restrict the analysis to individuals with positive earnings values, part of the effect of
38 educational inequalities is absorbed by the sample selection (especially among women and older
39 birth cohorts). As we see from columns (5) and (6) of table 3, both inequality in test scores and
40 inequality in educational attainment (measured across individuals with positive earnings) lose
41 magnitude. This is an indication that whenever we measure earnings inequality we face a huge
42 problem of self-selection into the labour market, which we are unable to properly model in this
43 context. For this reason, we stick to the solution of including all the relevant population, imputing
44 zero earnings when they are out of the labour market.⁹
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52 [Table 3 here]
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57 Overall, the magnitude of the coefficient on inequality in test scores oscillates between 1 and 1.7,
58 while the coefficient on inequality in years of education remains in the range of half of it, between
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2 0.5 and 0.8.¹⁰ In terms of elasticities (measured at sample means), earnings inequality measured by
3 Gini concentration indices would exhibit an elasticity of 0.61-0.69 with respect to inequality in test
4 scores and 0.21-0.22 with respect to inequality in years of education (computed from columns (3)
5 and (4) of table 3).
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8 Table 4 presents the instrumental variable estimation. We use the measures of educational
9 reforms constructed by Braga et al. (2013) covering various stages of schooling. If we exclude age
10 of begin/end of compulsory education and age of first tracking, the majority of these reform
11 measures are scale-free, because they count the number (and direction) of policy intervention.¹¹
12 Table A2 in the Appendix presents the areas for which information is available, while figure A1 in
13 the same Appendix plots the cross-country averages of these educational reform measures for one
14 area of reform.
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20 For ease of comparison in columns (1) and (2) of table 4 we have reproduced columns (3)
21 and (4) of previous table 3. In columns (3) and (4) we present the corresponding IV estimations
22 using a 2SLS estimator, while in columns (5) and (6) a GMM estimator is proposed (which allows
23 for heteroskedasticity of the errors and is more convenient when the number of instruments largely
24 exceed the number of supposedly endogenous variables). The bottom part of the table reports the
25 first stage coefficients of the regression of the endogenous variables onto the instruments
26 represented by measured reforms. Starting with first stage coefficients signs and significance, we
27 notice that inequality in years of education is reduced in countries that expanded pre-primary
28 education or postponed the beginning age for compulsory education, while the school leaving age
29 seems to have a counterintuitive positive correlation. Postponing the age at which students have to
30 choose the secondary school track (wherever the educational system is stratified, like in Austria,
31 Germany, Italy and the Netherlands) seem to increase both inequalities in schooling and test
32 scores.¹² On the contrary, strengthening the standardisation of national educational systems through
33 the introduction of student testing is associated to a reduction of inequality. Finally, consistent with
34 the results of Braga et al. (2013), increasing schools/teachers and universities autonomy reinforce
35 their potential competitiveness, at the expenses of increased educational inequality. Similar patterns
36 are observed in the case of competence inequality. The statistical significance of these effects relies
37 on the method utilised to estimate the variance-covariance matrix: while educational reforms remain
38 significant for test score inequality under both IV estimators, they tend to lose significance in the
39 case of attainment inequality when passing to GMM method of estimation.¹³ Using the predicted
40 inequalities in quantity and quality of education as regressors for earnings inequality, we observe
41 that their coefficients lose statistical significance in comparison with OLS, while rising in
42 magnitude in the case of schooling inequality. As long as educational reforms are reasonable
43 exogenous instruments (which is formally tested as robustness check in Braga et al. 2013), these
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1 values can be interpreted as causal impacts.¹⁴ Looking at column (6), we may say that OLS
2 estimates for test score inequality were upward biased, while the opposite situation occurs for
3 schooling inequality, which now becomes more relevant. In terms of elasticities, the two
4 dimensions of educational inequalities get closer (0.57 for test score inequality and 0.36 for
5 schooling inequality, both computed from column (6) of table 4 at sample means).
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11 [Table 4 here]
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16 Table 5 summarises the policy implications which can be inferred by the use of instrumental
17 variables. This is done in two ways, by directly estimating a reduced form (where $I(y)$ is directly
18 regressed onto the reform variables vector Z) and by computing the overall impact of educational
19 reforms onto earnings inequalities.¹⁵ Most of these effects are consistent with the previous
20 literature: reinforcing early (pre)schooling, delaying the beginning age for compulsory education,
21 reinforcing educational standardisation by introducing standardised test scores, all reforms yield a
22 reduction in earnings inequalities observed many years later in the labour market. On the contrary,
23 increasing teachers' autonomy (in the selection of teaching contents), reinforcing school
24 accountability and/or boosting university autonomy widen earnings differentials. According to the
25 reduced form estimation, two additional reforms (increasing the years of education and delaying the
26 tracking) come out statistically insignificant with respect to earnings inequalities.
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36 [Table 5 here]
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40 In order to visualise the magnitude of these impacts, in figure 2 we have plotted the reduced form
41 coefficients (column 1 of table 5) multiplied by one standard deviation change of the reform
42 variables. The units of the horizontal axis of figure 2 are to be interpreted as Gini points. Results are
43 strongly consistent with the findings of Braga et al. (2013): some reforms are effective in reducing
44 earnings inequality, in particular the expansion of public pre-primary education. Braga et al. (2013)
45 call these "inclusive" policies, while Busemeyer (2015) uses a slightly different conceptualization
46 of de-commodification. Other reforms are effective in raising earnings inequality, via increased
47 variability in both dimensions of human capital in the relevant population. These policies are
48 "selective" because they increase the degree of competition among educational institutions (Braga
49 et al. 2013); while in Busemeyer's (2015) approach these policies would be a mixture of
50 commodified (marketised) and stratified policies. Results are almost identical when considering
51 total earnings.
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[Figure 2 here]

Overall, our results do not reject our hypothesis that inequality in education affects inequality in earnings along two dimensions, quality and quantity. In addition, we have also shown that inequality in quality (as measured by student test scores) and inequality in quantity (as measured by years of schooling) respond to educational policies.

5. Dataset 2: PISA-PIAAC

One may suspect that previous results are driven by confounding factors, not properly accounted by country/year fixed effects, since we are considering quite different birth cohorts. In order to check the consistency of previous results with respect more homogenous populations, we have performed a different exercise on more recent surveys. We consider an adult population to which we impute competences measured when they were young, showing first that their current earnings are (weakly) correlated to past (and current) competences, and second that inequality in current earnings is correlated to either past or current inequality in competences, as in the previous exercise. In practice, we have taken the first PISA survey conducted in 2000 on 15-year-old students.¹⁶ These students are aged 27 in 2012 when another survey on the adult population (aged between 16 and 65) was conducted.¹⁷ The latter survey contains information on earnings, schooling and level of competences measured at the time of the survey. Thanks to the background information available in both surveys (gender, foreign born, highest parental education, books at home),¹⁸ we can identify 48 “types” in each of the 20 countries taking part to both surveys (numbers of individual in each surveys are reported in table A3). When looking at the distribution of the background information in the two surveys (see table 6) we notice that the foreign born population share has increased between the two surveys, while parental education and books at home tend to be similarly distributed.

[Table 6 here]

We impute the PISA test results by type to the individual-level PIAAC data. Imputed mean competences are obviously correlated with currently measured proficiency (0.46), which in turn is also correlated with schooling and labour market participation.¹⁹ In table 7 we present two groups of estimates according to the dependent variable which is available: for a subset of 15 countries the hourly wage is available as a continuous variable; for the largest set of 20 countries the earnings is only available in decile. As a consequence, columns 1 to 3 of table 7 presents OLS estimates, while

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2 columns 4 to 6 utilises an order probit estimator: the coefficients are therefore only comparable in
3 terms of sign and significance, but not of size.
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7 [Table 7 here]
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10 The first model considers quantity (years of schooling)²⁰ and contemporaneous quality (proxied by
11 numeracy test score, computed as mean of ten plausible values) of education: both are positively
12 correlated to earnings, but the former dominates in terms of magnitude. In the restricted dataset of
13 15 countries, we find that the average PISA score of a person's type is not significantly correlated
14 to average earnings in adulthood. Models 4-6, covering 20 countries, show that adult competences,
15 high-school competences and years of education all are directly and significantly associated to
16 earnings.
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23 [Figure 3 here]
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26 The various inequality measures are presented in figure 3. The strongest correlation is between
27 inequality in schooling and inequality in current competences (south-west quadrant), because the
28 latter are clearly cumulated following the permanence in school. However current competences are
29 also correlated with competences when in school (north-west quadrant). But we have to remind that
30 inequality in competences when young is an underestimate of actual inequality (since in the current
31 exercise it comprises only the between-group component). This may explain its low correlation
32 with schooling (south-west quadrant). The raw data do not present (unconditional) correlation
33 between inequality in achievements and inequality in earnings (south-east quadrant).
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39 Moving to the replication of the inequality model, the estimates referred to the largest
40 dataset are reported in table 8. Column 1 of table 8 regresses the inequality on earnings (computed
41 over relevant deciles for 20 countries) onto inequality of schooling and inequality in current
42 competences for the relevant age cohort (25-29 year old); column 4 replicates the same structure
43 extending the sample to a larger age cohort. In both cases inequality of current competences has a
44 higher correlation with earnings inequality. However, if we replace current with (imputed) past
45 competences (columns 2 and 5), the correlation becomes insignificant for the younger cohort, while
46 it retains significance and magnitude for the largest sample.²¹ Finally we include both measures of
47 quality of education (past and current competences) in addition to quantity of education in columns
48 3 and 6, but the effect of current competences dominates over the effect of past ones, which does
49 not come out significant. So, while the earlier analysis pointed to stronger effects of skills
50 inequality than of attainment inequality, the present analysis shows that contemporary skills
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2 inequality is more strongly associated to earnings inequality than skills inequality during the high
3 school years. Yet, the fact that (adult) competences are more predictive of earnings inequality than
4 attainment inequality confirms our findings on multiple cohorts.
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8 [Table 8 here]
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10 11 **6. Conclusions and discussion** 12 13

14 We studied the relationship between educational inequality and earnings inequality. Adopting a
15 framework in which educational distributions have both a qualitative and a quantitative dimension,
16 with quality referring to student performance on standardised tests, and quantity referring to the
17 attained level of education, we perform two exercises where we examine the correlation existing in
18 the available data using two alternative strategies of matching different datasets. In both exercises
19 we find that inequality in quality of education affects inequality in earnings, even controlling for
20 inequality in quantity (proxied by schooling). And the effect of quality is stronger than the effect of
21 attained level. We employed instrumental variable models using information on educational
22 reforms. Exploiting cross-country and cross-year variations in the policies, we are able to show that
23 educational reforms have an effect on the distribution of the quality and the quantity of education.
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31 Interpreting these results from different theoretical perspectives concerning how educational
32 distributions are related to changing earnings inequalities, our findings are more consistent with a
33 human capital/functionalist explanation than with an explanation derived from closure theory.
34 Based on the alliance between human capital and functionalist theories, that hold that education is
35 increasingly rewarded because of the skills that are needed in today's economies, a stronger effect
36 of skill inequality is expected (as we found). Closure theory would hold that earnings inequality can
37 only be explained by educational factors to the extent that educational performance is translated in
38 educational attainment. Skills inequality should have a minor role once educational attainment
39 inequality is held constant (opposite to what we found). Nevertheless, also educational attainment
40 inequality had an independent effect on earnings inequality, which does support the claim that the
41 contribution of education to explaining earnings inequality is not restricted to a human
42 capital/functionalist model. In line with closure theory, educational attainment inequality has an
43 independent effect on earnings inequality in most of our models.
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52 One issue that we have not explicitly dealt with, but is potentially important for our findings,
53 is within-group inequality. It is well-known that the rising earnings inequality cannot fully be
54 attributed to rising inequalities between educational categories (supported in our analyses), because
55 also within-group inequality is rising. If the within-education-group inequality is strongly
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1 overlapping with measured skills, then this could explain why we find stronger skill effects (which
2 would still be consistent with our interpretation of the neo-classical/functionalist model).²²
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5 Note that the human capital/functionalist perspective fares better to explain associations
6 between inequalities, but this does not mean that human capital theory also fares better to explain
7 average earnings. Our analysis showed that educational attainment is strongly correlated with
8 average earnings even after holding constant on (adult) competences. Following the approach by
9 Bowles and Gintis (2002) this would point to relatively weak evidence for a human capital
10 interpretation of how mean earnings are determined.
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15 In line with earlier studies (Bedard and Ferrall 2003; Solga 2014), our results demonstrated
16 that dispersions in educational outcomes are associated to dispersions in incomes. This central
17 finding has important policy implications. Even if income policies have started to include an
18 orientation towards education, our results more explicitly call attention for educational distributions
19 as a cause of income inequality. While governments are increasingly concerned with students'
20 performance in international tests, these concerns are mostly motivated by an efficiency concern of
21 promoting economic growth. Our results would suggest that policy makers who are concerned
22 about the growing income inequality in most Western societies should more explicitly incorporate
23 educational distributions into their agenda setting. Education policy and income policy go hand in
24 hand. Our approach has however limitations, since we cannot consider life-cycle implications of
25 changes in current distributions of educational inequality. Due to the lack of longitudinal data we are
26 unable to assess whether measured inequality in current earnings will expand or contract with years.
27 Given existing evidence on the differential decay of competences by educational attainment, we
28 suspect that inequality in earnings may increase overtime. But this is topic for future research.
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41 ¹ It is common procedure when there is more than one suspected endogenous variable to regress
42 them on all exogenous variables and all instruments, since there is no statistical test indicating
43 which instrument is more appropriate for which variable. In principle we could have selected some
44 instruments as more theoretically appropriate for either inequality in skills or in education, but we
45 would have not been able to reject alternative models.

46 ² The micro-data on the second and third surveys (the first one is only publicly accessible on paper)
47 are accessible through the International Association for the Evaluation of Educational Achievement,
48 IEA, at www.iea.nl.

49 ³ <http://epp.eurostat.ec.europa.eu/portal/page/portal/microdata/echp>.

50 ⁴ http://epp.eurostat.ec.europa.eu/portal/page/portal/microdata/eu_silc

51 ⁵ Belgian data are available for the Francophone and Flemish communities separately, but the same
52 does not apply in the case of earnings data, therefore we were forced to consider Belgium as a
53 unique entity.

54
55 ⁶ In order to account for potential country differences in retention rates, we faced the problem of
56 attributing a standard duration (in years) of the maximal educational attainment. We have obtained
57 an estimate of median duration of studies by using the individual age of leaving school (variable
58 AGEDU in ECHP or variable PE030 "Year when highest level of education was attained" in SILC),
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3 and then we have assigned 7 years for ISCED=1, 10 years for ISCED=2, 13 years for ISCED=3, 15
4 years for ISCED=4 and 18 years for ISCED=5 in SILC. Corresponding values are 8 for EDU=1, 12
5 for EDU=2 and 17 for EDU=3 in ECHP.

6⁷ Although since 1995, item response theory has been used to measure five plausible values of the
7 true mathematics performance, for the earlier cohorts this was not possible. In TIMSS 1995 the
8 correlation between our test score and the five plausible values is, in all five cases, 0.89. For reasons
9 of comparability we chose to operationalize math in the same way for all three cohorts.

10⁸ One may notice that the inequality in years of education has lower values than inequality in test
11 scores. This may partly be attributed to the fact that education is computed on a group-aggregate
12 variable (standard duration for educational attainment) while test score is computed on a (bounded)
13 continuous variable. Thus our inequality measure on quantity has lower mean and variance
14 (corresponding to the between-group component) than the inequality measure on quality; other
15 things constant the former should obtain a higher regression coefficient than the latter when
16 correlated to inequality in earnings.

17⁹ As further robustness check we estimated our model separately for men and women, retaining
18 some statistical significance for inequality in years of education, but losing it for inequality in tests
19 score. But the reliability of this check is limited, since country/cohort/survey fixed effects use 27
20 out of 41 degrees of freedom.

21¹⁰ This is only partially related with the different scales of measurement of inequalities in quantity
22 and quality of human capital. If we were to replace the coefficients of column 5 in table3 with beta
23 coefficients (i.e. standardizing each variable), we would obtain a coefficient of 0.55 for inequality in
24 quality and 0.27 for inequality in quantity.

25¹¹ We can look at the sign and the statistical significance of the estimated correlation between
26 inequality in education and educational reforms, whereas when considering the magnitude we
27 should distinguish according to whether we are dealing with step-dummies or continuous variables

28¹² This latter effect contradicts standard findings in the literature on determinants of test scores (see
29 for example Hanushek and Woessmann 2005) while it is consistent with the findings for inequality
30 in years of education (Brunello and Checchi 2007).

31¹³ Notice that all regressions control for country, birth year, age and survey fixed effects, so that
32 reform impacts are identified by time variation within the countries and/or gender differential
33 impacts within each country/year.

34¹⁴ While we included all policies as instruments for both educational variables, one could also
35 partition the policies as instruments for either test score inequality or attainment inequality. As a
36 robustness check we did this, with public pre-primary, standardized tests, school accountability and
37 teacher autonomy as instruments for math test inequality, and the other reforms as instruments for
38 attainment inequality. The results (available upon request) showed significant positive effects of test
39 score inequality, and non-significant positive effects of attainment inequality, just like in the 2SLS
40 and IV GMM models of table 4. The effect sizes are larger, but that follows from the use of fewer
41 instruments. As the choice of instruments for either one of the two educational variables is arbitrary,
42 we chose to report the full instrumental variable models.

43¹⁵ We have multiplied the first stage coefficients - from bottom part of table 4 - by the impact of the
44 supposedly endogenous variable - from upper part of table 4. Take for example a unitary increase of
45 the variable "reform on public pre-primary schooling": it produces a decrease of skill inequality by
46 -0.102 and education inequality by -0.096. The overall decline in dependent employment earnings is
47 the given by $[-0.102 \times 1.519 - 0.096 \times 0.730] = -0.225$. The effect on total labour earnings is computed
48 in a similar way.

49¹⁶ See www.oecd.org/pisa/.

50¹⁷ The PIAAC, see <http://www.oecd.org/site/piaac/>

51¹⁸ We cannot use the exact age because 5 countries in PIAAC report age in 5-year brackets. Another
52 information we would have liked to use for a coarser match would have been the type of secondary
53 school attended, which is available in PIAAC but in PISA is only available starting from 2003.

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3 Thus we are left with gender (2) × native/foreign born (2) × highest parental education (recoded in 3
4 items for comparability across surveys – this now corresponds to the ECHP classification, indicated
5 in footnote 6) × books at home (recoded in 4 items for comparability across surveys), yielding an
6 overall number of 48 possible combinations, which are obviously unevenly distributed in the
7 population.

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9 ¹⁹ In analogy to the previous exercise, the lack of a proper longitudinal structure of the data forces
10 us to consider only between-type inequality in test score of the young population (PISA), because
11 we do not have a convincing strategy to reconstruct the within-type inequality of young test score in
12 the adult population. It is therefore not surprising the correlation between young and adult test
13 scores takes such a low value, since it is just measuring the between-type component.

14 ²⁰ This variable is absent for Germany, and has been reconstructed from (median) leaving school
15 age.

16 ²¹ Since inequality in test score captures inequality between-types, if there have been relevant
17 compositional changes in the distribution of types across age-cohorts within each country (the
18 average being cleaned away by the country fixed effect), this may be reflected in the statistical
19 significance of inequality of competences at the younger age.

20 ²² However, dealing with within- and between-group inequality would require a different research
21 design in which educational attainment, skills and luck are included in a structural model.
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Table 1: Construction of the sample – matching rule

Birth year	Aged 14	Aged 28	Aged 43-44	Aged 59	matched cohorts/countries
1950	1964 (from FIMS: BE,FI,FR,DE,NL,UK)	1978 (data not available)	1994 (from ECHP1994: BE,FR,DE,NL,UK)	2009 (from SILC2009: BE,FI,FR,DE,NL,UK)	11
1966	1980 (from SIMS: (BE,FI,FR,HU,NL,SE,UK)	1994 (from ECHP1994: BE,FR,NL,UK)	2009 (from SILC2009: BE, FI,FR,HU,NL,SE,UK)		11
1981	1995 (from TIMS: AT,BE,CZ,DK,FR,DE, GR,HU,IE,IT,LV,NL,NO, PT,SK,SI,ES,SE,UK)	2009 (from SILC2009: AT,BE,CZ,DK,FR,DE, GR,HU,IE,IT,LV,NL,NO, PT,SK,SI,ES,SE,UK)			19

Table 2: Descriptive statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
Gini index on dependent employment gross earnings (including non labour force with zero earnings)	82	0.466	0.110	0.254	0.721
Gini index on gross earnings (including self-employed and non labour force - negative earnings set to 0)	82	0.433	0.109	0.229	0.698
Gini index on years of education (computed from ISCED attainments)	82	0.120	0.032	0.072	0.243
Gini index on dependent employment gross earnings (only positive values - excluding unemployed with zero earnings)	82	0.310	0.061	0.193	0.452
Gini index on gross earnings (including self-employed but excluding non labour force - negative earnings set to 0)	82	0.316	0.061	0.209	0.472
Gini index on years of education (computed from ISCED attainments - only population with positive earnings)	82	0.115	0.029	0.074	0.216
Gini index on math test scores	82	0.175	0.031	0.124	0.244
age of individuals (when interviewed about occupational status)	82	37.04	11.40	28	59
reform on public pre-primary schooling	82	0.508	0.445	0	1
compulsory education (start age)	82	6.024	0.608	5	7
compulsory education (end age)	82	15.415	1.440	12	18
tracking age	82	13.476	2.263	10	16
introduction of standardised test	82	0.341	0.451	0	1
reform on school accountability	82	0.293	0.458	0	1
reform on school teacher autonomy	82	0.549	0.494	0	1
reform of university access	82	0.606	0.430	0	1

Table 3: Gross earnings and educational inequality – Gini indices – OLS with country fixed effects

	1	2	3	4	5	6
	dep.empl. earnings robust se	total earnings robust se	dep.empl. earnings clustered se	total earnings clustered se	dep.empl. earnings >0 clustered se	total earnings >0 clustered se
inequality in math test scores	1.631 [0.555]***	1.716 [0.546]***	1.631 [0.815]*	1.716 [0.817]**	1.084 [0.508]**	1.079 [0.560]*
inequality in years of education (from isced attainments)	0.849 [0.371]**	0.825 [0.354]**	0.849 [0.370]**	0.825 [0.377]**	0.570 [0.153]***	0.519 [0.194]**
male component	-0.076 [0.012]***	-0.103 [0.012]***	-0.076 [0.013]***	-0.103 [0.012]***	-0.037 [0.009]***	-0.033 [0.010]***
Observations	82	82	82	82	82	82
R-squared	0.84	0.85	0.84	0.85	0.77	0.74

col.1-2: robust standard errors in brackets - col.3-6: standard errors in brackets clustered by country - constant, country, age, birth year and survey controls included - * significant at 10%; ** significant at 5%; *** significant at 1%

Table 4: Gross earnings and educational inequality – Gini indices – OLS and IV estimates with educational reforms as instruments

	1 OLS		3 IV 2SLS		4 IV GMM	
	dep.empl. earnings	total earnings	dep.empl. earnings	total earnings	dep.empl. earnings	total earnings
inequality in math test scores	1.631 [0.815]*	1.716 [0.817]**	1.073 [1.029]	1.269 [0.998]	1.519 [0.579]***	1.426 [0.690]**
inequality in years of education (from isced attainments)	0.849 [0.370]**	0.825 [0.377]**	1.277 [1.261]	1.544 [1.272]	0.73 [0.843]	1.314 [0.916]
Observations	82	82	82	82	82	82
R-squared	0.84	0.85	0.84	0.83	0.83	0.84
			1st stage:		1st stage:	
			Gini math test	Gini yrs education	Gini math test	Gini yrs education
reform on public pre-primary schooling			-0.102 [0.020]***	-0.096 [0.020]***	-0.102 [0.021]***	-0.096 [0.072]
compulsory education (start age)			-0.066 [0.012]***	-0.075 [0.015]***	-0.066 [0.017]***	-0.075 [0.041]*
compulsory education (end age)			0.01 [0.002]***	0.009 [0.002]***	0.01 [0.002]***	0.009 [0.005]*
tracking age			0.009 [0.002]***	0.007 [0.002]***	0.009 [0.003]**	0.007 [0.055]
introduction of standardised test			-0.093 [0.015]***	-0.074 [0.014]***	-0.093 [0.018]***	-0.074 [0.055]
reform on school accountability			0.015 [0.027]	0.051 [0.024]*	0.015 [0.032]	0.051 [0.075]*
reform on school teacher autonomy			0.03 [0.008]***	0.029 [0.009]***	0.030 [0.008]***	0.029 [0.018]
reform of university access			0.082 [0.012]***	0.038 [0.011]***	0.082 [0.015]***	0.038 [0.037]
R-squared			0.94	0.77	0.94	0.77
F test 1st stage [p-value]			169.5 [0.00]	1932.2 [0.0]	29.29 [0.0]	1.08 [0.38]
Test for overidentifying restrictions (Sargan for 2SLS, Hansen for GMM)			10.17 [0.11]	9.51 [0.14]	10.17 [0.11]	9.51 [0.14]

Standard errors in brackets clustered by country [2sls] or robust against heteroscedasticity [gmm and 1st stage] – * significant at 10%; ** significant at 5%; *** significant at 1% - constant, gender, age, country, survey and year controls included

Table 5: Reduced form multipliers computed from table 4: effects of policies on earnings inequality

	estimated from reduced form		computed from columns (5) and (6) of table 4	
	Gini index dependent employment earnings	Gini index on total labour earnings	Gini index dependent employment earnings	Gini index on total labour earnings
reform on public pre-primary schooling	-0.346 [0.078]***	-0.407 [0.083]***	-0.225	-0.272
compulsory education (start age)	-0.200 [0.053]***	-0.226 [0.056]***	-0.155	-0.193
compulsory education (end age)	0.001 [0.007]	0.008 [0.008]	0.022	0.026
tracking age	-0.007 [0.008]	-0.005 [0.008]	0.019	0.022
introduction of standardised test	-0.178 [0.076]**	-0.232 [0.088]**	-0.195	-0.230
reform on school accountability	0.176 [0.099]*	0.232 [0.102]**	0.060	0.088
reform on school teacher autonomy	0.100 [0.031]***	0.125 [0.032]***	0.067	0.081
reform of university access	0.077 [0.052]	0.104 [0.055]*	0.152	0.167
Observations	82	82		
R-squared	0.83	0.85		

* significant at 10%; ** significant at 5%; *** significant at 1%

Standard errors in brackets clustered by country - constant, gender, age, country, survey and year controls included

Table 6. Distribution of population by background information – PISA 2000 and PIAAC 2012

individual characteristics (per cent points)	PISA 2000 (sample weights)	PIAAC 2012 Aged 25-29 (sample weights)	PIAAC 2012 Aged 20-34 (sample weights)
male	49.40	50.91	50.21
female	50.60	49.09	49.79
Either native-born or native-language at home	98.34	91.99	91.89
Foreign-born and foreign-language at home	1.66	8.01	8.11
both parents less than secondary completed	29.38	15.90	16.53
at least one parent with secondary degree	30.66	43.22	43.74
at least one parent with college degree	39.96	40.88	39.72
books at home (when 14 yrs old): 0-10	9.93	12.39	12.21
books at home (when 14 yrs old): 11-100	40.69	47.96	48.23
books at home (when 14 yrs old): 101-500	37.63	32.06	32.39
books at home (when 14 yrs old): >500	11.74	7.60	7.18

Table 7. Determinants of (log) earnings – population aged 25-29 - PIAAC 2012

	1	2	3	4	5	6
	OLS	OLS	OLS	oprobit	oprobit	oprobit
adult competences	0.001**		0.001***	0.005***		0.004***
	[0.000]		[0.000]	[0.001]		[0.000]
young competences (from PISA cells' means)		0.001	0.000		0.003***	0.002***
		[0.001]	[0.001]		[0.000]	[0.000]
Years of education	0.036***	0.041***	0.036***	0.099***	0.117***	0.093***
	[0.008]	[0.008]	[0.010]	[0.022]	[0.030]	[0.023]
female	-0.124***	-0.123***	-0.116**	-0.287***	-0.299***	-0.259***
	[0.035]	[0.038]	[0.039]	[0.068]	[0.049]	[0.062]
age	-0.203	-0.255	-0.248	0.769	0.586	0.651
	[0.580]	[0.599]	[0.596]	[0.691]	[0.692]	[0.718]
age ²	0.004	0.005	0.005	-0.013	-0.01	-0.011
	[0.011]	[0.011]	[0.011]	[0.013]	[0.013]	[0.013]
Observations	5605	5485	5485	8345	8206	8206
Countries	15	15	15	20	20	20
R ² -Pseudo R ²	0.92	0.919	0.92	0.04	0.04	0.04

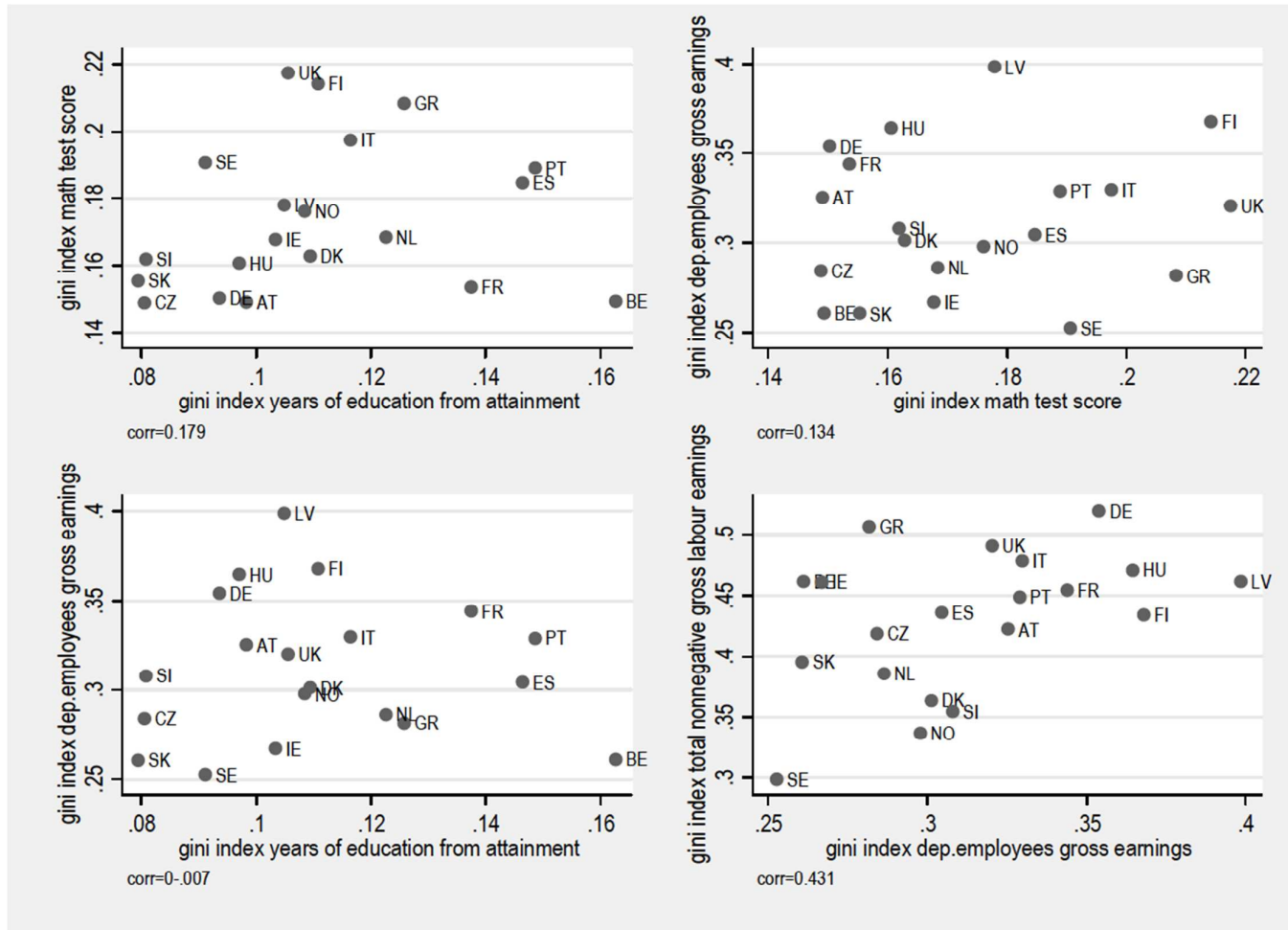
Robust standard errors in brackets *** p<0.01, ** p<0.05, * p<0.1 errors clustered by country - weighted – country fixed effect included

Table 8. Inequality in education and earnings – population aged 25-29 and 20-65 -
PIAAC 2012

	1	2	3	4	5	6
VARIABLES	Gini hourly wage decile	Gini hourly wage decile	Gini hourly wage decile	Gini hourly wage decile	Gini hourly wage decile	Gini hourly wage decile
	aged 25-29			aged 20-65		
Gini index years of education	0.487** [0.227]	0.657*** [0.212]	0.484** [0.230]	-0.209 [0.177]	-0.007 [0.175]	-0.209 [0.177]
Gini index adult competences – PIAAC 2012	0.676** [0.250]		0.716** [0.295]	1.074*** [0.277]		0.981*** [0.289]
Gini index youth competences (imputed) - PISA 2000		0.395 [0.370]	-0.139 [0.443]		1.025** [0.448]	0.595 [0.458]
female population share	0.033*** [0.007]	0.032*** [0.008]	0.033*** [0.008]	0.050*** [0.004]	0.046*** [0.004]	0.049*** [0.004]
Observations	40	40	40	360	360	360
Countries	20	20	20	20	20	20
R ²	0.467	0.384	0.468	0.65	0.639	0.652

Robust standard errors in brackets *** p<0.01, ** p<0.05, * p<0.1 - constant, country controls included – columns 4 to 6 also include birth cohort controls

Figure 1: Inequality in competences, years of schooling, gross labour earnings (dependent employment and total employment – From SILC2009)



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Figure 2: Earnings inequality impact of educational reforms

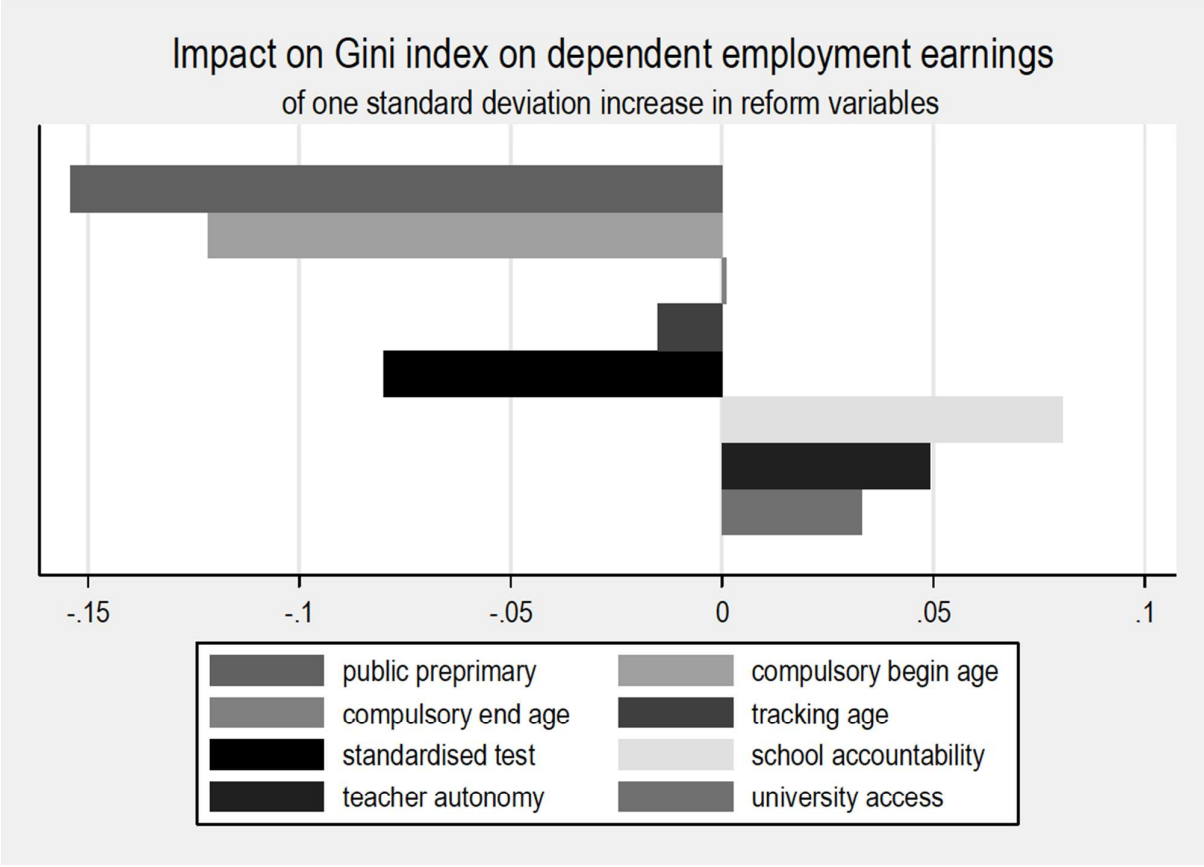


Figure 3: Inequality in current and past competences, years of schooling and gross labour earnings

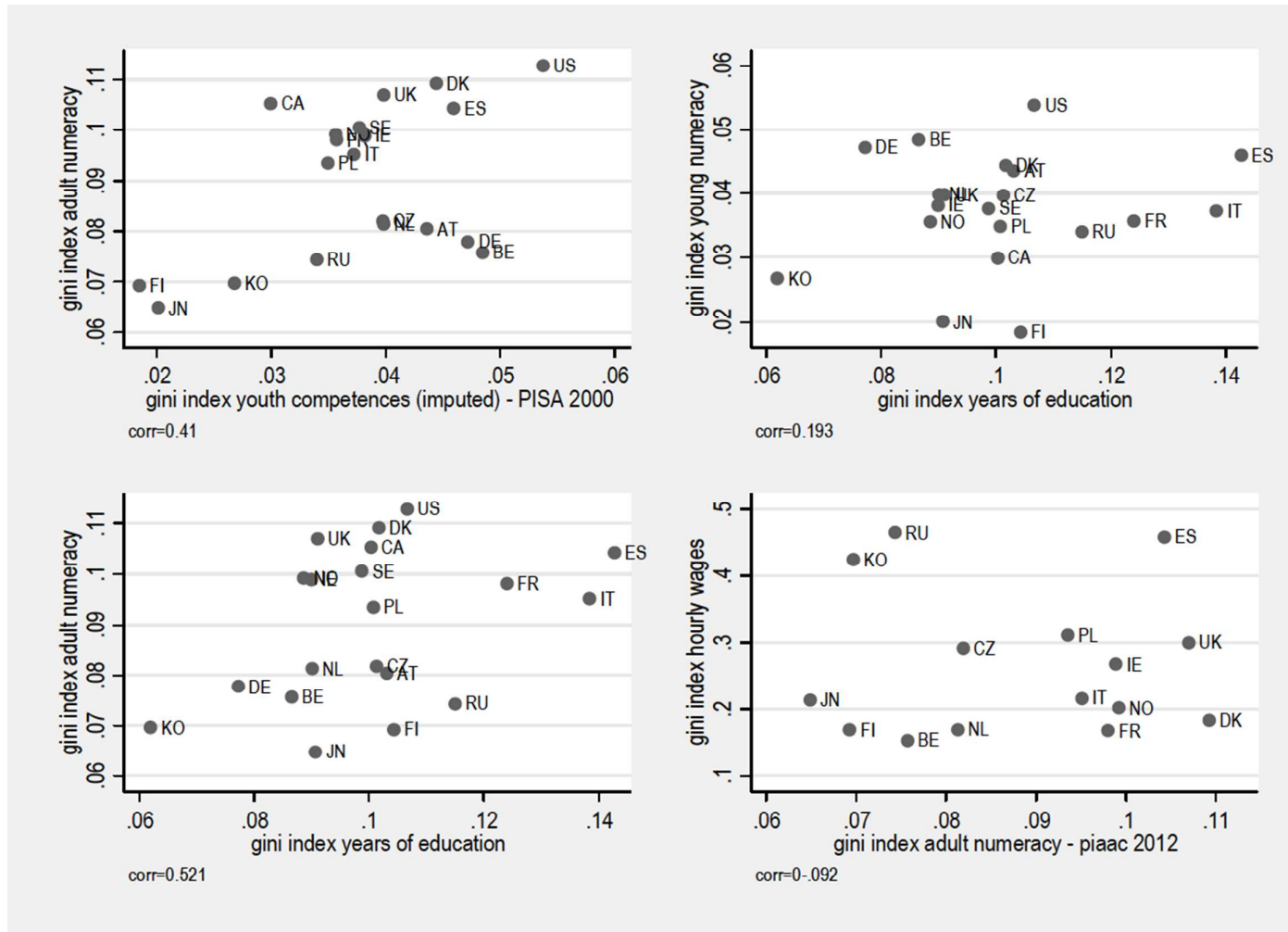


Table A1 – Inequality in earnings and educational attainment – averages across gender and age cohorts

1st number: Gini index on gross total labour earnings of employed –

2nd number: Gini index on years of education of the same population (from maximal educational attainment) –

3rd number: Gini index on math test score –

4th number: individuals with positive earnings

5th number: observations available in the sample

	birth year			Total
	1950	1966	1981	
Austria			0.32	0.32
			0.10	0.10
			0.15	0.15
			134	134
			2	2
Belgium	0.32	0.25	0.21	0.27
	0.15	0.12	0.10	0.13
	0.17	0.15	0.12	0.15
	202	282	136	620
	4	4	2	10
Czech Republic			0.30	0.30
			0.08	0.08
			0.15	0.15
			188	188
			2	2
Denmark			0.30	0.30
			0.11	0.11
			0.16	0.16
			97	97
			2	2
Finland	0.37	0.35		0.36
	0.12	0.10		0.11
	0.21	0.22		0.21
	363	361		724
	2	2		4
France	0.41	0.33	0.26	0.35
	0.17	0.12	0.10	0.14
	0.18	0.15	0.14	0.16
	456	545	246	1247
	4	4	2	10
Germany	0.35		0.36	0.35
	0.12		0.09	0.11
	0.14		0.16	0.15
	638		185	823
	4		2	6
Greece			0.31	0.31
			0.11	0.11
			0.21	0.21
			168	168
			2	2
Hungary		0.35	0.37	0.36
		0.09	0.09	0.09
		0.16	0.16	0.16
		266	234	500
		2	2	4
Ireland			0.26	0.26
			0.10	0.10
			0.17	0.17

			85	85
			2	2
Italy			0.33	0.33
			0.11	0.11
			0.20	0.20
			421	421
			2	2
Latvia			0.39	0.39
			0.10	0.10
			0.18	0.18
			134	134
			2	2
Netherlands	0.32	0.28	0.22	0.28
	0.12	0.11	0.10	0.11
	0.18	0.18	0.14	0.17
	406	482	196	1084
	4	4	2	10
Norway			0.30	0.30
			0.11	0.11
			0.18	0.18
			107	107
			2	2
Portugal			0.33	0.33
			0.15	0.15
			0.19	0.19
			106	106
			2	2
Slovak Republic			0.26	0.26
			0.08	0.08
			0.16	0.16
			198	198
			2	2
Slovenia			0.30	0.30
			0.08	0.08
			0.16	0.16
			420	420
			2	2
Spain			0.31	0.31
			0.14	0.14
			0.18	0.18
			361	361
			2	2
Sweden		0.23	0.28	0.25
		0.09	0.09	0.09
		0.22	0.16	0.19
		260	155	415
		2	2	4
United Kingdom	0.37	0.36	0.29	0.35
	0.14	0.13	0.09	0.13
	0.23	0.24	0.18	0.22
	387	554	130	1071
	4	4	2	10
Total	0.35	0.3	0.3	0.32
	0.14	0.11	0.1	0.11
	0.18	0.19	0.17	0.18
	2452	2750	3701	8903
	22	22	38	82

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For Peer Review

Table A2 – Areas of educational reform (from Braga et al. 2013)

area of reform	expected impact on schooling inequality	available information on reforming activity
pre-primary education	reduction (through increased educational attainment of students from disadvantaged background)	<ul style="list-style-type: none"> ✓ extending access to (public preprimary education) ✓ making preprimary education compulsory ✓ raising training requirements for preprimary teachers
expansion of compulsory education	reduction (through increased educational attainment of students from disadvantaged background)	<ul style="list-style-type: none"> ✓ compulsory education (start age) ✓ compulsory education (end age)
school tracking	ambiguous (vocational tracks have shorter duration, prevent academic enrolment but have lower drop-out rates)	<ul style="list-style-type: none"> ✓ raising tracking age
school autonomy	ambiguous (adaptability to social environment, increased competition in presence of centralised control)	<ul style="list-style-type: none"> ✓ national testing for other purposes than career advancement ✓ teacher autonomy in designing curricula
school accountability	increase (school differentiation, screening and sorting of students)	<ul style="list-style-type: none"> ✓ national testing for career advancement (ISCED 1 and 2)
teacher qualification	ambiguous (better quality benefits students from poorer backgrounds but allows for greater differentiation)	<ul style="list-style-type: none"> ✓ strengthening teacher training requirements for primary school teachers ✓ strengthening teacher training requirements for secondary school teachers
university autonomy and selectivity	increase (increased signalling value of tertiary education requires a more intensive selectivity in university admissions)	<ul style="list-style-type: none"> ✓ selectivity in admissions ✓ financial support to student ✓ size of grant/loan component/interest rates ✓ university autonomy (budget, recruitment, organization, logistics, course design, self-evaluation, planning)

Table A3 – Sample size of PISA 2000 and PIAAC 2012 for countries participating to both surveys

country	PISA 2000	PIAAC 2012 Aged 16- 19	PIAAC 2012 Aged 20- 24	PIAAC 2012 Aged 25- 29	PIAAC 2012 Aged 30- 34	PIAAC 2012 Aged 35- 39
AUSTRIA	2401	424	462	465	464	497
BELGIUM	3453	432	478	398	442	442
CANADA	16010	2146	2320	2084	2370	2530
CZECH REPUBLIC	3002	716	747	702	617	593
DENMARK	2247	514	542	452	566	614
FINLAND	2529	409	462	474	534	458
FRANCE	2371	493	517	536	555	599
GERMANY	2498	493	542	482	471	448
IRELAND	2071	332	409	563	795	803
ITALY	2680	222	301	339	439	588
JAPAN	2816	356	393	405	492	626
KOREA, REPUBLIC OF	2759	569	496	600	653	749
NETHERLANDS	1308	423	456	377	406	460
NORWAY	2123	480	468	423	490	517
POLAND	1749	1052	3377	1643	451	447
RUSSIAN FEDERATION	3573	423	762	463	324	250
SPAIN	3184	457	512	521	627	714
SWEDEN	2362	372	443	389	402	413
UNITED KINGDOM	4748	456	652	758	942	917
UNITED STATES	1842	343	467	516	502	464
Total	65726	11112	14806	12590	12542	13129

Figure A1 – Country averages of educational reform measures with inclusive orientation (from Braga et al. 2013)

