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Skills, Employment, and Labor Market Institutions: Evidence from PIAAC

Jon Marius Vaag Iversen[1] — Bjarne Strøm[2]

Abstract. Hanushek et al (2015, 'Returns to Skills Around the World: Evidence from PIAAC', European Economic Review 73: 103) find a weak wage–skill relationship in countries with limited skill reward possibilities due to high union density, strict employment protection, and large public sector. If these factors also restrict employment possibilities and the incentives to join the labor market, a possible mirror image of the weak wage–skill relationship is a steeper employment–skill gradient. We use PIAAC data to estimate the employment–skill association, and the results for the whole sample of individuals give some indication that the employment–skill gradient is steeper in countries with strict employment rules and centralized bargaining. Our results for subgroups show imprecisely estimated employment–skill gradients for immigrants. For individuals with poor health conditions and low formal education, the estimated gradient is somewhat higher than in the whole sample in countries with high bargaining coverage, a large public sector, and centralized collective bargaining systems.

1. Introduction

Workers' skills are key determinants of economic growth in modern economies. Empirical micro studies have traditionally considered the relationship between wages and skills as measured by formal education, usually years of education (Mincer equations), whereas research at the macro level has analyzed the relationship between GDP growth and aggregate country level of formal education. However, formal education levels may hide important elements of educational quality especially in a cross-country context. Skills vary within schooling categories used to define formal education, and the quality of schooling institutions will to an unknown extent vary across countries. Thus, a more recent line of research has studied the relationship between wages, economic growth, and direct measures of skills as measured by test scores (Hanushek and Woessmann (2008), Hanushek *et al.* (2015), and Hanushek *et al.* (2017)). Controlling for fixed country effects, Hanushek *et al.* (2015) find that the relationship between wages and skill measures based on newly available data from the 'Programme for the International Assessment of Adult Competencies (PIAAC)' varies across countries with different labor market institutions. In particular, the link between wages and skills is weaker in the public than in the private sector, and weaker

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in countries with high union density, large public sector shares, and strict employment protection laws. Their evidence is consistent with the view that such institutional factors generate a compressed wage structure, high wage floors, and restrict the skill reward possibilities.[1] Whereas the cross-country differences in wage returns to skills are interesting and important, less attention is paid to the relationship between the employment–skill gradient and labor market institutions.

A mirror image of the weak relationship found between skills and wages in countries with strong trade unions, large public sectors, and strict employment protection laws may appear as a stronger relationship between skills and employment propensity in such countries in general and in particular for vulnerable groups in the labor market. This may partly be a consequence of employers that requires high productivity employees to compensate for high and inflexible wages generated by labor market institutions.

Most of the previous empirical literature has studied how *levels* of unemployment and employment shares vary across and within countries with different labor market institutions and findings and conclusions vary across studies.[2] In contrast, our study is one of the few that investigates in a systematic way how the employment–skill *gradient* varies across countries with different labor market institutions.[3] By estimating an employment equation at the individual level, controlling for fixed country effects, we account for all observable and unobservable country-specific variables that may affect the association between individual employment probability and country-specific institutions. In other words, the employment *level* effects of labor market institutions are not identified in our model, whereas the model enables us to isolate the association between the *employment–skill gradient* and institutional characteristics.

Further, compared with earlier studies, our paper contributes to the literature by exploiting the rich comparable micro cross-country data provided by PIAAC linked with information on labor market institutions to investigate how the employment–skill gradient varies across a relatively large number of countries with different institutions. In particular, the use of direct measures of skills based on comparable tests of key competencies across individuals in the countries represents an improvement relative to traditional studies with skill measures solely based on formal education and years of schooling that does not account for differences in school quality across countries. We also estimate the employment–skill gradient for specific subgroups traditionally thought to be vulnerable in the labor market, that is, individuals with weak health, immigrant background, and individuals with low formal education. Whereas the estimation results vary somewhat in precision and across specifications and subsamples, the results are broadly consistent with the hypothesis that the employment–skill gradient is larger in countries with strict employment protection legislation, large public sector, and centralized collective bargaining.

The rest of the paper is organized as follows: Section 2 presents theoretical considerations and Section 3 describes the data. Section 4 presents the empirical approach. Sections 5 and 6 contain empirical results, whereas Section 7 concludes.

2. Institutions and the relationship between wages, employment, and skills

This section gives a broad overview of the possible mechanisms through which labor market institutions can affect the employment–skill gradient directly or indirectly through the wage-setting system and government interventions in the labor market.

2.1. Trade unions and wage bargaining

Trade union power and wage bargaining systems may affect the employment–skill relationship. A simple view is that a compressed wage structure due to high wage floors set in negotiated contracts between unions and firms may reduce the probability to be employed for some individuals as employers require sufficiently high productivity or skill level in the first place to match the high wage. However, several authors argue that unionized labor markets may imply efficiency gains. Hirschman (1970) and Freeman and Medoff (1984) argue that unions are associated with both efficiency increasing behavior by exerting collective voice, and traditional efficiency reducing behavior through rent-seeking. Barth *et al.* (2014) argue that strong trade unions and coordinated wage bargaining combined with high welfare spending and social security safety nets lead to better macroeconomic performance, high employment rates and sustained long-run economic growth as evidenced by economic development in the Scandinavian countries. Thus, the sign of the association between the employment–skill gradient and trade union strength and collective bargaining institutions is an empirical question.

2.2. Statutory minimum wages

According to conventional theory, government-imposed statutory minimum wages may function similar to union-imposed wage floors above the competitive wage and limit demand for low-skilled workers, decrease the wage–skill relationship, and increase the employment–skill gradient. However, Boeri (2012) argues theoretically and finds empirical evidence consistent with the hypothesis that statutory minimum wages are typically set lower than union-imposed wage floors. Further, the conventional view of a negative effect of minimum wages on employment has been questioned in empirical studies and the issue is still highly controversial, see the recent discussion between Allegretto *et al.* (2017) and Neumark *et al.* (2014). Hanushek *et al.* (2015) find that the wage–skill gradient is actually positively associated with statutory minimum wages when a minimum wage indicator is included separately in their wage model. Thus, to what extent the employment–skill gradient is positively or negatively associated with the statutory minimum wage is an open question that deserves empirical analysis.

2.3. Public sector employment share

Public sector employment may stabilize total employment and increase wages and employment levels of disadvantaged groups. However, recent research has shown that large public sector employment may have more subtle side effects, see the overview in Caponi (2017). The overall wage distribution is found to be more compressed in the public than in the private sector and centralized pay setting with little regional flexibility seems to be the norm for important groups of public sector employees like nurses, teachers, and military personnel in many countries.[4] Countries with large public sectors may thus have a more compressed wage structure that restricts the skill reward possibilities. The mirror image of this may be a positive relationship between the employment–skill gradient and the share of workers in the public sector.

2.4. Employment protection regulations

Employment protection regulations can affect employment, wages, and the employment–skill gradient. For given wages, stricter employment protection implies higher firing costs and increases the likelihood of incumbent workers to retain their jobs. At the same time, as expected future firing costs increase, firms become more reluctant to hire workers in the first place and to require higher skills and productivity to compensate for higher future firing costs. Thus, the net effect on employment and the employment–skill gradient for given wages is theoretically uncertain. However, stricter employment protection may also affect wage setting. On one hand, it can decrease wages, as workers are willing to accept lower wages in return for job security. On the other hand, it can lead to higher wages, because stricter employment protection gives incumbent workers a stronger bargaining position. Recent empirical research has found that the wage response is highly heterogeneous (Leonardi and Pica 2013) and international empirical evidence Koeniger *et al* (2007) suggests that stricter employment protection legislation decreases wage differentials. Consistent with this, Hanushek *et al.* (2015) find that wage return to skills is lower in countries with stricter employment legislation. It is an empirical question to what extent the low wage–skill gradient in such countries leads to a high employment–skill gradient.

3. Data

PIAAC is a large survey initiated by the OECD (2013) consisting of individual indicators of skills based on tests results covering key competencies in several dimensions (numeracy, literacy and problem-solving in technology-rich environments) as well as information of key employment, earnings, education and background characteristics for around 5000 individuals per country in a number of countries around the world. The analysis in this paper explores data for 21 of these countries.[5] Below, we describe the definitions of the key variables used in the analysis.

3.1. Employment

The key-dependent variable in our empirical analysis is employment which in our baseline models and descriptive statistics is defined by an indicator for whether the individual has been employed last year or not. In PIAAC, there is also an alternative indicator defined as whether the individual has been employed the last week. Analysis using this alternative measure is used in robustness checks in section 4.3.

3.2. Skills

The PIAAC survey assesses cognitive skills in three domains Numeracy, Literacy, and Problem Solving in Technology-Rich Environments (PS-TRE). Our primary choice is to use Numeracy as our skills measure similar to Hanushek *et al* (2015) among others. Literacy is used in robustness checks, whereas we have chosen not to use PS-TRE, partly because Spain, Italy, and France did not participate in the test of PS-TRE and partly because individual respondents were allowed to opt out of this test, see also p. 113 in Hanushek *et al.* (2015).[6] Detailed descriptions of the skill domains can be found in OECD (2019).[7] To help interpretation, the skill measures are normalized into variables with mean zero and standard deviation 1.

3.3. Individual controls

We follow Hanushek *et al.* (2015) and include some standard individual control variables; experience defined as the number of years in paid work, experience squared, a gender dummy equal 1 if individual is male. In addition, we control for immigrant status for the individuals as well as their parents by including dummies for first- and second-generation immigrants. In robustness analysis, we also include controls for educational attainment measured by years of completed education and replace the experience variable with the respondents age. In some of the specifications below, we also include individual health condition indicators as additional control variables. The background questionnaire in the PIAAC survey included a question about respondent's health conditions: The respondents could answer: 'Excellent', 'very good', 'good', 'fair' and 'poor'. We choose to use a more aggregated representation of this variable. We define persons answering 'fair' and 'poor' health into one category named 'Weak health', and those answering 'Excellent' or 'very good' are defined as one category named 'very good health'. 'Good health' represents respondents reporting 'good' in the original questionnaire. Subjective health indicators are far from perfect, and Jürges (2007) gives an extensive discussion of the problems with the use of such indicators in cross-country studies.

In most models estimated, we also include age group dummies (5-year intervals) as reported by the variable AGEG5LFS in the PIAAC database.

3.4. Labor market institutional variables

As our primary research question is to what extent labor market institutions affect the employment–skill gradient, we take as point of departure the labor market characteristics included in Hanushek *et al* (2015). The institutional labor market variables included in their study were *union density*, index of *employment protection strictness*, an indicator for statutory *minimum wage,* and the *share of employment in the public sector*.[8] As the statutory minimum wage dummy will not account for the extent to which the minimum wage bite, we have instead used the level of the minimum wage relative to the median wage in the country as reported by OECD similar to Koeniger *et al* (2007). Further, we include the *bargaining coverage* variable included in the database in the online Supplemental material in Hanushek *et al* (2015), but not used as an explanatory variable in the published version of their paper. In addition to the variables in their dataset, we also include two variables that characterize the collective bargaining institutions in different countries, motivated by the large literature on the effect of such institutions on macroeconomic outcomes. These are the indexes of *Collective bargaining centralization* and *Collective bargaining coordination.* The centralization index describes the level at which bargaining typically takes place, that is, at the plant, industry or national level and we use the variable in Visser (2015). The coordination index captures in addition to the level where bargaining takes place, the amount of implicit coordination between independent unions and employers, see Boeri and van Ours (2008 p, 56). This type of variables has been widely used to explain wages and unemployment in cross-country studies, see Calmfors and Driffill (1988), Nickell *et al.* (2005), Nunziata (2005), Nymoen and Sparrmann (2015), and chapter 3 in Boeri and van Ours (2008). While Calmfors and Driffill (1988) argue that there is a hump-shaped relationship between unemployment and the degree of centralization, Di Tella and McCulloch (2005) suggest a monotonic positive relationship between unemployment and bargaining centralization. Nunziata (2005) finds a negative relationship between aggregate labor costs and the degree of bargaining coordination. Appendix A gives detailed descriptions and sources for the institutional indicators, whereas Table A1 reports the actual number for the indicators for the countries in our analysis.

3.5. *Descriptive statistics*

Figure 1 presents employment shares in 21 countries according to the definition used above. The Nordic countries Norway, Denmark, Sweden, and Finland stand out with high employment shares, whereas countries in eastern and southern Europe have the lowest employment shares.

Figure 2 shows simple plots of the relationship between employment shares and institutional characteristics of the labor market in the countries. The broad picture is that the employment share is positively correlated with public sector share, union density, and the index of bargaining coordination and negatively correlated with the index of employment protection strictness and the minimum wage variable. However, as noted in the Introduction, our focus in the empirical analysis below is on the employment–skill **gradient** and the relationship between the gradient and institutional characteristics of the labor market.

4. Empirical approach

The ideal approach in empirical research is to estimate causal relationships between variables. However, in our case and with the type of data and variables available in the cross-country setting, we make no claim that the relationships we estimate can be interpreted causally. Rather the relationships should be interpreted as conditional correlations. However, we investigate to what extent the associations between individual employment performance and measured skills are robust to including different sets of control variables as well as other robustness checks.

Point of departure for the empirical study is equation (1) where i, c, and j denote individual, country, and cohort (age group), respectively:

$$y_{icj} = \alpha_0 + \alpha_1 skills_{icj} + \mathrm{X}_{icj}\alpha_2 + \beta_c + \gamma_j + \varepsilon_{icj} \qquad [1]$$

The dependent variable y represents employment outcomes. *Skills* represents the competence measures in the PIAAC survey, numeracy, and literacy. The parameter of interest is α_1 and measures the employment–skill gradient. In addition, we include a vector X of individual and family characteristics with a corresponding coefficient vectorα_2. β_c represents country fixed effects. γ_j is an age group dummy variable and accounts for the fact that employment outcomes vary with age group j. In one specification reported below, we also

Figure 1. Employment shares across countries (PIAAC) [Colour figure can be viewed at wileyonlinelibrary.com]

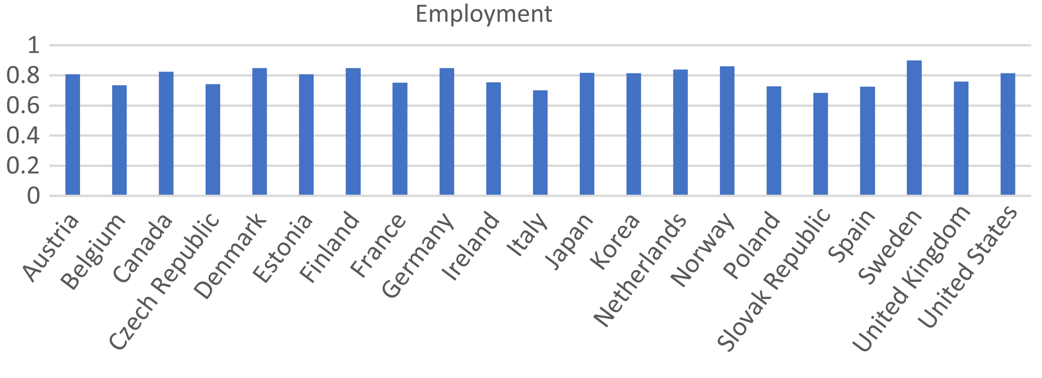
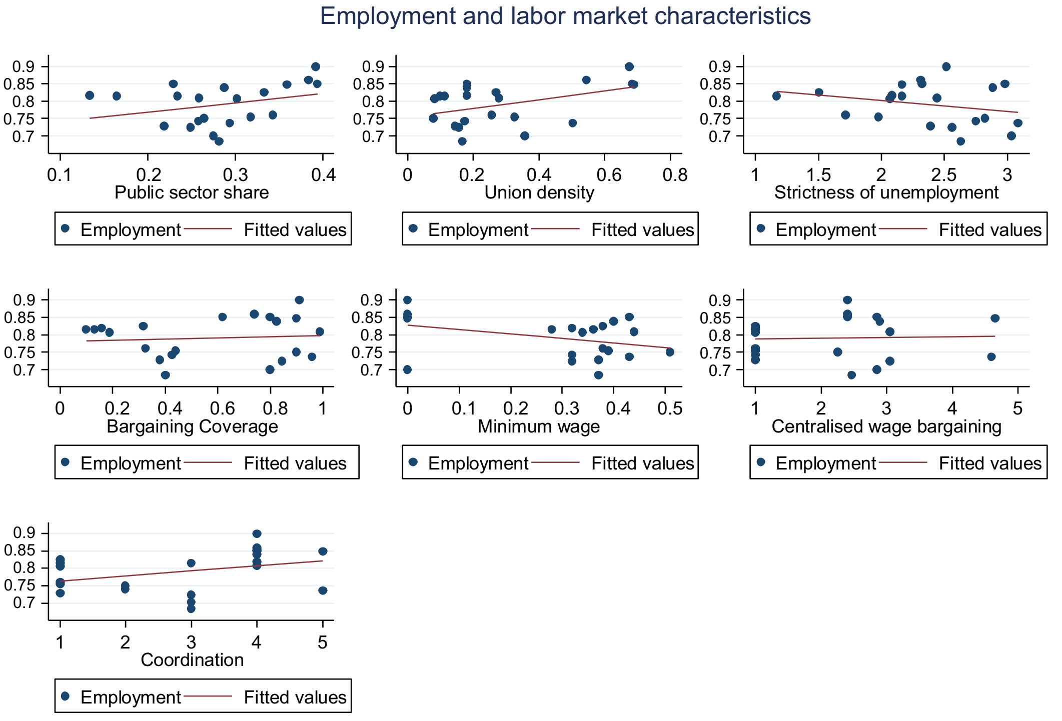


Figure 2. Cross-plot of employment shares and labor market characteristics [Colour figure can be viewed at wileyonlinelibrary.com]



allow these age group differences to vary across countries by including a term $\beta_c \times \gamma_j$. This can be justified by several arguments. Individuals in different cohorts may have experienced country-specific shocks and external conditions affecting labor market outcomes and educational investments. Examples are reforms in the education system (extension of years of compulsory education and expansion of higher education), changes in the pension system, and retirement age. Another example is that the labor market situation in the country back in time can have long-lasting effects on labor market outcomes as it may have affected individuals' educational decisions made in their teenage years as well as the wage rates and type of jobs at the start of their labor market career.[9]

We run OLS regression models corresponding to different versions of (1) using whole sample of individuals and countries in the PIAAC survey as well as subsamples of individuals with weak health, low formal education, and immigrants.

5. Baseline results

5.1. Employment–skill gradient

We first present the results from the baseline pooled regression models in Table 1.[10] This provides the broad picture of the average relationship between employment and skills in the included 21 countries.

Column (1) reports the estimated relationship between employment and skills when no controls or fixed effects are included, that is, the raw correlation between the two variables.

The estimated coefficient implies that one standard deviation increase in numeracy competence is associated with 10 percentages point higher employment probability. Columns (2) add standard observed individual characteristics and age group fixed effects that reduce the employment–skill gradient by nearly one half. According to the results in column (2), one standard deviation increase in numeracy is associated with around 5.8 percentages points increase in the probability to be employed. Column (3) adds country fixed effects, whereas column (4) adds country by age group fixed effects. In column (5), we control for subjective health conditions. The reported estimates in column (5) reveal that employment status is strongly associated with individual health conditions. Relative to the reference category (*Weak health*), having *Good* and *Very good* health is associated with around 13 and 17 percentages point higher employment probability, respectively. However, the employment–skill gradient does not change dramatically by adding health conditions as the estimated coefficient in front of the skill variable drops from 0.057 in column (3) to 0.048 in column (5). Moreover, the skill estimate is fairly unaffected by the inclusion of country fixed effects and age group by country effects. In columns (6), we also add completed years of schooling to the model. Adding this variable reduces the skill effect by around one half, although it is still a significant predictor of employment. In the following, we mostly estimate variants of the model specification in column (3). In addition, we also include years of education in the model in further robustness analysis.[11]

5.2. *Employment–skill gradient and labor market institutions*

Hanushek *et al.* (2015) show that the wage–skill gradient is low in countries with high union density, large public sector shares, and strict employment protection legislation.

Table 1. Baseline model results

	(1)	(2)	(3)	(4)	(5)	(6)
Numeracy	0.100***	0.0579***	0.0575***	0.0572***	0.0477***	0.0259***
	(0.00506)	(0.00430)	(0.00432)	(0.00395)	(0.00352)	(0.00342)
Good health					0.134***	
					(0.0140)	
Very good health					0.173***	
					(0.0135)	
Years of schooling						0.0219***
						(0.00154)
Observations	119,192	109,116	109,116	109,116	109,116	107,466
R-squared	0.066	0.211	0.228	0.264	0.251	0.249
Individual controls	No	Yes	Yes	Yes	Yes	Yes
Age group fixed effects	No	Yes	Yes	Yes	Yes	Yes
Country fixed effects	No	No	Yes	Yes	Yes	Yes
Age group by country fixed effects	No	No	No	Yes	No	No
Number of countries	21	21	21	21	21	21

Notes: Dependent variable is employment last 12 months. Estimated standard errors corrected for clustering at country level. *,**, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included Belgium, Canada, Check Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, work experience, work experience squared and immigration status, and parents immigrant status.

These institutional variables are usually associated with countries with strong unions, highly compressed wage structure, and a high general wage level. It does not come as a surprise that the wage–skill gradient is less steep in such countries than in countries with weak trade unions as USA and UK.[12] It is a possibility that the relationship between skill gradient and the labor market institutions is quite different when we look at the employment margin. Employers may rationally require a highly productive workforce to compensate for the relatively high wage level, limited skill reward possibilities due to strong unions and centralized wage setting, and strict employment protection legislation. This may be particularly important for vulnerable groups in the labor market. However, as argued in section 2 above, the theoretical relationship between the employment–skill gradient and labor market institutions is ambiguous, and hence, the question can only be answered by empirical analysis.

We now estimate an extended version of the baseline model where we add interaction terms between the skill variable and indicators for the labor market institutions discussed above. Our interaction models include fixed country effects that account for all observable and unobservable country-specific variables that may affect employment probabilities.

Formally, we estimate the following extended version of equation (1), where Z_c represents labor market institutions:

$$y_{icj} = \alpha_0 + \alpha_1 skills_{icj} + \alpha_3 skills_{icj} \times Z_c + X_{icj}\alpha_2 + \beta_c + \gamma_j + \varepsilon_{icj} \qquad [2]$$

Thus, the employment *level* effects of labor market institutions are not identified in our model, whereas the effects of institutions on the *employment–skill* gradient are measured by the coefficients in front of the interaction term. In these respects, our model is similar to the model used to investigate the relationship between the *wage–skill* gradient and labor market variables in Hanushek *et al.* (2015).

As most of the institutional variables are heavily correlated, Table 2 reports results when the interaction variables are included one by one. In columns (1)–(6), we include institutional variables available in the dataset in Hanushek *et al* (2015): the indicators for union density, strictness of employment protection, bargaining coverage, public sector share, and minimum wage. Columns (7) and (8) include the variables characterizing the collective bargaining institutions in the countries. Whereas only two of the interaction variables included in columns (1)–(6) are statistically significant (strictness of employment protection and bargaining coverage), the signs of the interactions between employment protection strictness, union density, and public sector share are opposite to those reported in Table 9 in Hanushek *et al.* (2015). To illustrate the numerical relationship between the employment–skill gradient and bargaining coverage in column (2), an increase in bargaining coverage by 60 percentage points, which is approximately the difference between USA and Norway, increases the probability of employment by 1.8 percentage points per standard deviation increase in numeracy skills. Among the two additional characteristics of the bargaining institutions, the interaction with the centralization index has a positive sign and is significant at the 10% level. The numerical result implies that an increase in the index of degree of bargaining centralization from the countries with the lowest (UK and USA) to the highest (Finland) is associated with an increase in probability of employment by 2.5 percentage points per standard deviation increase in numeracy skills. The effect of the bargaining coordination index is however negative, numerically small and far from statistically significant at conventional levels. It is a question why the effect of these two indicators of collective bargaining institutions is so different. It is not clear from the literature what is the best measure of these institutions as pointed out on p.56 and p.71 in Boeri and van Ours (2008).[13]

5.3. Robustness

5.3.1. The influence of specific countries. As our institutional variables only vary at the country level, a reasonable question is to what extent the results are dependent on specific countries included in the sample. To investigate this question, we run the model in Table 2 excluding the countries one by one. Figure 3a-g shows the estimated interaction effects along with the upper and lower bounds of the 95% confidence intervals using this procedure. The estimates for the interaction between the skill effect and the centralization of bargaining and the index of employment protection strictness are fairly independent of which of the countries that are left out. The size and precision of the estimate for the bargaining coverage and public sector share interaction are somewhat dependent on whether Japan and Korea are included in the sample although the point estimates are positive in all cases.

5.3.2. Using literacy competence instead of numeracy as skill measure. Until now, we have used numerical competence as our skill measure. Table 3 shows the results for the effect of labor market institutions on the employment–skill gradient when we instead measure skills by literacy competence. The results turn out to be very similar to those reported in Table 2. The most substantial differences are that the interaction effects between skills and bargaining centralization becomes somewhat larger and more precisely estimated (significant at 5 percent level), whereas the estimated effect of the minimum wage variable is negative. Whereas the latter result is at odd with the effect of the other labor market characteristics, it should be noted that Hanushek *et al.* (2015) find that the wage–skill effect is *positively* associated with the minimum wage variable.[14]

5.3.3. Controlling for years of schooling. Table 4 shows results for the effect of labor market institutions on the employment–skill gradient when we add years of schooling to the model and use numeracy as the skill measure. In general, the estimated numerical effects of the interaction variables are quite similar to those in the baseline model in Table 2. The most substantial difference is that the interaction effects with union density and public sector share become numerically higher, whereas the effects of the interactions with bargaining coverage and bargaining centralization are a bit reduced and less precisely estimated.

5.3.4. Controlling for age instead of actual work experience. It could be argued that actual work experience is endogenous and thus should not be included in the employment equation. On the other hand, work experience is an important channel for learning and thus likely to be correlated with measured skills. Thus, excluding this variable could lead OLS estimates to suffer from omitted variable bias. So far, we have followed Hanushek *et al.* (2015) and controlled for a quadratic polynomial in work experience in the estimated equations. As a robustness check, we nevertheless report in Table 5 the results when the quadratic polynomial in actual experience is replaced by a quadratic polynomial in age. As the public use files for USA, Canada, Germany, and Austria only contain data on age intervals for the respondents, these countries are omitted in the regressions, leading to a substantial reduction in observations relative to Table 2. Whereas the precision of the estimates is somewhat reduced due to lower sample size, the results are numerically quite similar to the results in Table 2.

Table 2. Employment–skill gradient interacted with labor market institutions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Numeracy	0.0569***	0.0214	0.0395***	0.0241	0.0550***	0.0440***	0.0637***
	(0.00955)	(0.0161)	(0.0127)	(0.0286)	(0.00421)	(0.0149)	(0.00539)
Numeracy*Union density	0.00205						
	(0.0199)						
Numeracy*Strictness of employment		0.0154***					
		(0.00594)					
Numeracy*Bargaining coverage			0.0313*				
			(0.0181)				
Numeracy*Public sector share				0.118			
				(0.0935)			
Numeracy*Minimum wage					0.00867		
					(0.0151)		
Numeracy*centralized wage bargaining						0.00688*	
						(0.00396)	
Numeracy*coordination wage bargaining							−0.00215
							(0.00233)
Observations	109,116	109,116	109,116	109,116	109,116	109,116	109,116
R-squared	0.228	0.228	0.229	0.228	0.228	0.228	0.228
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age group fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Dependent variable is employment last 12 months. Estimated standard errors corrected for clustering at country level. *, **, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included Belgium, Canada, Check Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, work experience, work experience squared and immigration status, and parents immigrant status.

Figure 3. (a) Centralized wage bargaining. Estimated interaction effects and corresponding upper and lower bounds of 95% confidence interval when excluding separate countries one by one. (b) Strictness of employment. Estimated interaction effects and corresponding upper and lower bounds of 95% confidence interval when excluding separate countries one by one. (c) Bargaining coverage. Estimated interaction effects and corresponding upper and lower bounds of 95% confidence interval when excluding separate countries one by one. (d) Union density. Estimated interaction effects and corresponding upper and lower bounds of 95% confidence interval when excluding separate countries one by one. (e) Collective bargaining coordination. Estimated interaction effects and corresponding upper and lower bounds of 95% confidence interval when excluding separate countries one by one. (f) Minimum wage. Estimated interaction effects and corresponding upper and lower bounds of 95% confidence interval when excluding separate countries one by one. (g) Public sector share. Estimated interaction effects and corresponding upper and lower bounds of 95% confidence interval when excluding separate countries one by one [Colour figure can be viewed at wileyonlinelibrary.com]

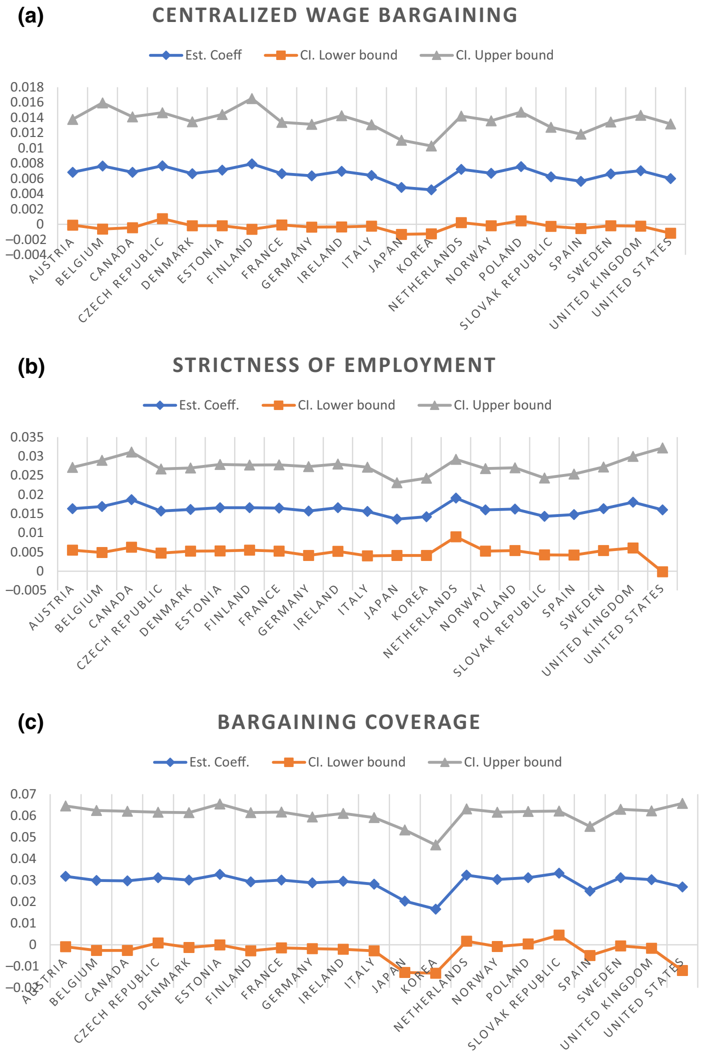


Figure 3. Continued

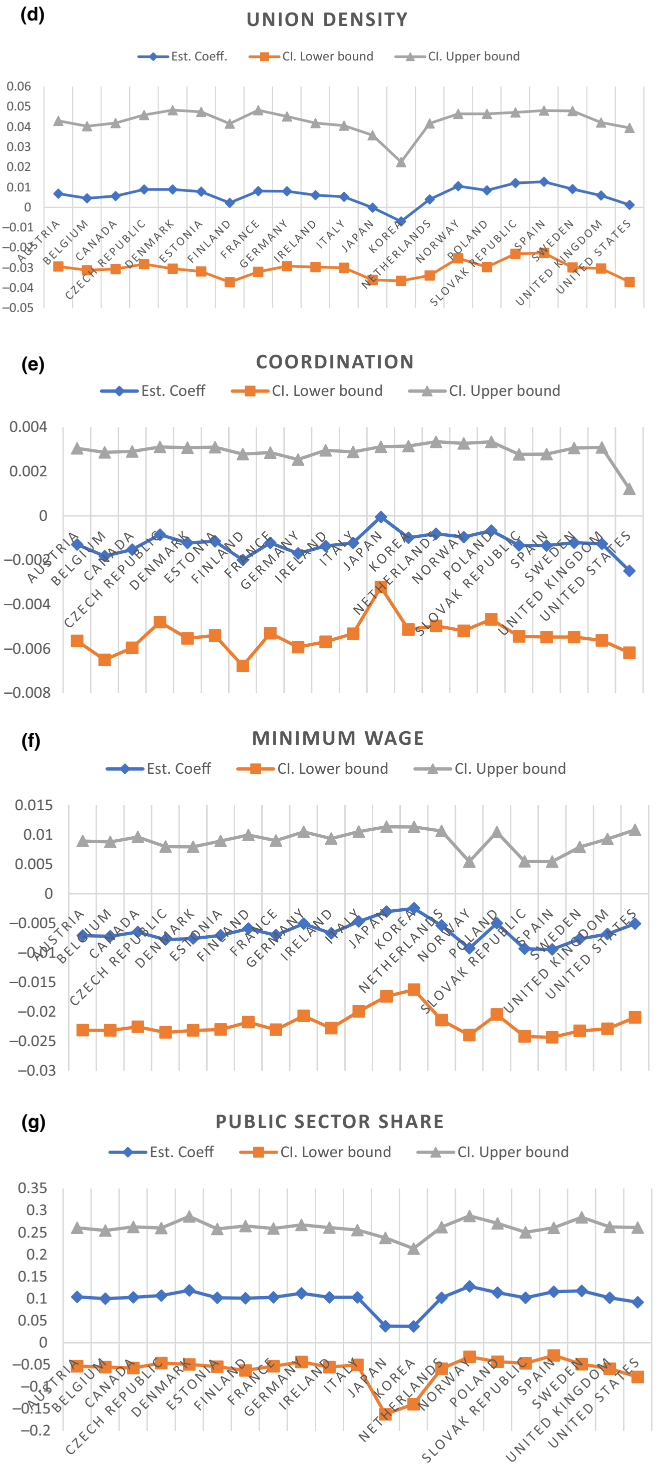


Table 3. Employment–skill gradient interacted with labor market institutions. Literacy as skill measure

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Literacy	0.0441***	0.0167	0.0252*	0.00408	0.0559***	0.0423**	0.0536***
	(0.00946)	(0.0183)	(0.0127)	(0.0309)	(0.00676)	(0.0154)	(0.00562)
Literacy*Union density	0.0279						
	(0.0213)						
Literacy*Strictness of employment		0.0151**					
		(0.00650)					
Literacy*Bargaining coverage			0.0471**				
			(0.0186)				
Literacy*Public sector share				0.170			
				(0.103)			
Literacy*Minimum wage					-0.0128		
					(0.0204)		
Literacy*centralized wage bargaining						0.00992**	
						(0.00421)	
Literacy*coordination wage bargaining							−0.000492
							(0.00291)
Observations	109,116	109,116	109,116	109,116	109,116	109,116	109,116
R-squared	0.225	0.225	0.226	0.225	0.224	0.224	0.224
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age group fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Dependent variable is employment last 12 months. Estimated standard errors corrected for clustering at country level. *, **, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included Belgium, Canada, Check Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, work experience, work experience squared and immigration status, and parents immigrant status.

5.3.5. Using employment status last week as dependent variable. So far, we have presented estimation results with employment status last year as dependent variable. An alternative is to use the respondent's employment status in the week proceeding the interview. Table 6 presents the results for the estimated interaction between employment–skill gradient and labor market institutions using this alternative dependent variable. The estimated coefficients for the interaction terms increase slightly in numerical terms, but the qualitative results are very similar to those obtained in Table 2 using employment status last year as dependent variable.

6. The employment–skill gradient for vulnerable groups

Some individuals are believed to be particularly vulnerable in the labor market. This includes people with disabilities or poor health, immigrants and ethnic minorities, and people with little formal education. Several studies find that poor health reduces labor force participation, see Cai (2010) and the references therein. Evidence also suggests that immigrants perform poorer in the labor market in terms of both employment and wages, see Bisin *et al.* (2011), Bratsberg *et al.* (2014), and Sarvimäki (2011). Labor force participation is generally lower for people with low formal education and Kahn (2000) finds that relative employment rates of males with low education vary inversely with bargaining coverage in the country.

In general, the literature has focused on variation in employment and earnings levels for these vulnerable groups across countries with different labor market institutions. Much less attention has been given to variation in the association between employment and skills across countries for these groups. One exception is Bratsberg *et al.* (2013). Using data from the ALL survey, they find evidence that the association between employment and skills for immigrants and individuals with poor health is stronger in Norway compared to that in USA and Canada. They argue that this is consistent with the hypothesis that compressed wage structure due to centralized collective bargaining institutions, strict employment protection rules, and generous social insurance in Norway compared with USA and Canada may create adverse employment effects for low-productivity groups. Our data set contains a much larger number of countries and enables us to investigate to what extent their findings based on a comparison of these three countries represents a general pattern.

In this part of the paper, we therefore split the sample between three vulnerable subgroups defined by health conditions, immigration status, and formal education and investigate to what extent the employment–skill gradient for these particular subgroups varies with the institutional characteristics of the labor market. In Table 7, columns (2), (3), and (4) present the estimated coefficients for the interaction terms for the different institutional variables. To facilitate comparison with results for the whole sample, Column (1) contains the estimated interaction coefficients from Table 2.

6.1. Individuals with weak health

In Table 7, column 2 reports the relationship between the employment–skill gradient and our measures of labor market institutions for the group of individuals with weak health. The interaction terms for the union power proxies reveal a mixed picture. Whereas the estimate for the union coverage interaction term is significantly positive for the whole sample,

Table 4. Employment–skill gradient interacted with labor market institutions. Numeracy as skill measure. Controlling for years of education

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Numeracy	0.0216**	−0.0054	0.00948	-0.0162	0.0260***	0.0122	0.0294***
	(0.00799)	(0.0159)	(0.0124)	(0.0236)	(0.00436)	(0.00975)	(0.00571)
Years of education	0.0219***	0.0218***	0.0218***	0.0220***	0.0219***	0.0218***	0.0219***
	(0.00148)	(0.00155)	(0.00154)	(0.00145)	(0.00153)	(0.00153)	(0.00154)
Numeracy*Union density	0.0143						
	(0.0197)						
Numeracy*Strictness of employment		0.0132**					
		(0.00630)					
Numeracy*Bargaining coverage			0.0285				
			(0.0196)				
Numeracy*Public sector share				0.148*			
				(0.0812)			
Numeracy*Minimum wage					−0.00046		
					(0.0146)		
Numeracy*centralized wage bargaining						0.00642	
						(0.00405)	
Numeracy*coordination							−0.000122
							(0.00242)
Observations	107,466	107,466	107,466	107,466	107,466	107,466	107,466
R-squared	0.249	0.249	0.249	0.250	0.249	0.249	0.249
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age group fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Dependent variable is employment last 12 months. Estimated standard errors corrected for clustering at country level. *, **, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included Belgium, Canada, Check Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, work experience, work experience squared and immigration status, and parents immigrant status.

Table 5. Employment–skill gradient interacted with labor market institutions. Quadratic polynomial in experience replaced with q-uadratic polynomial in age

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Numeracy skills	0.0715***	0.0164	0.0512**	0.0261	0.0746***	0.0590***	0.0837***
	(0.0141)	(0.0361)	(0.0187)	(0.0301)	(0.00545)	(0.0151)	(0.00720)
Numeracy*Union density	0.00752						
	(0.0244)						
Numeracy*Strictness of employment		0.0236					
		(0.0137)					
Numeracy*Bargaining coverage			0.0384				
			(0.0235)				
Numeracy*Public sector share				0.166*			
				(0.0919)			
Numeracy*Minimum wage					−0.00272		
					(0.0216)		
Numeracy*centralized wage bargaining						0.00686	
						(0.00551)	
Numeracy*coordination							−0.00336
							(0.00223)
Observations	80,245	80,245	80,245	80,245	80,245	80,245	80,245
R-squared	0.157	0.157	0.157	0.158	0.157	0.157	0.157
Individual controls	yes	yes	yes	yes	yes	yes	yes
Country fixed effects	Yes	yes	yes	yes	yes	yes	yes

Notes: Dependent variable is employment last 12 months. Estimated standard errors corrected for clustering at country level. *, **, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included Belgium, Check Republic, Denmark, Estonia, Finland, France, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, age, age squared and immigration status, and parents immigrant status.

Table 6. Employment–skill gradient interacted with labor market institutions. Employment status last week as dependent variable. Numeracy as skill variable

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Numeracy score	0.0512***	0.00159	0.0353**	0.0112	0.0589***	0.0376***	0.0554***
	(0.0106)	(0.0139)	(0.0142)	(0.0280)	(0.00411)	(0.0112)	(0.00564)
Numeracy*Union density	0.0174						
	(0.0207)						
Numeracy*Strictness of employment		0.0176**					
		(0.00657)					
Numeracy*Bargaining coverage			0.0364*				
			(0.0198)				
Numeracy*Public sector share				0.159*			
				(0.0888)			
Numeracy*Minimum wage					−0.0922		
					(0.0147)		
Interaction - Numeracy*centralized wage bargaining						0.00875**	
						(0.00418)	
Interaction - Numeracy*coordination							0.000305
							(0.00198)
Observations	109,116	109,116	109,116	109,116	109,116	109,116	109,116
R-squared	0.196	0.196	0.196	0.196	0.196	0.196	0.196
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age group fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Dependent variable is employment last week. Estimated standard errors corrected for clustering at country level. *, **, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included: Belgium, Canada, Check Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, work experience, work experience squared and immigration status, and parents immigrant status.

the estimate is low and statistically insignificant for the group with weak health. For the union density variable, the opposite picture emerges. The strictness of employment protection interaction term is imprecisely estimated and lower in numerical terms compared with the whole sample. For the public sector share, the estimated interaction term is larger and statistically significant for individuals with weak health compared with the estimate for the whole sample. The numerical estimate implies that an increase in the public sector share by 15 percentage points (The difference between USA and Norway) is associated with 5.4 percentage points increase in employment probability per one standard deviation increase in numeracy skills for people with weak health. The association between the employment–skill gradient and centralization of collective bargaining is also substantially higher for individuals with weak health than for the whole sample, whereas the association with the coordination index is positive, but small and not statistically significant. According to the numerical result, an increase in bargaining centralization from the lowest (1 in UK and USA) to the highest in the sample (5 in Finland) is associated with 5.7 percentage points increase in employment probability per standard deviation increase in numeracy skills for individuals with weak health, whereas the comparable result for the whole sample is 2.8 percentage points.

The estimated effect of the interaction between cognitive skills and the minimum wage variable is negative and significant using this subsample. Taken literally, this result implies that the association between employment and skills for individuals with poor health is

Table 7. Estimated Interaction terms between numeracy and institutional variables for different subgroups

Interaction with numeracy	**(1)**	**(2)**	**(3)**	**(4)**
	Whole sample	Weak health	Immigrants	Low formal education
Union density	0.00205	0.0820**	0.0447	0.0341
	(0.0199)	(0.0313)	(0.0458)	(0.0215)
Strictness of employment	0.0154***	0.00972	0.0195	0.0169*
	(0.00594)	(0.0144)	(0.0130)	(0.00860)
Bargaining coverage	0.0313*	0.0473	0.0350	0.0385*
	(0.0181)	(0.0282)	(0.0286)	(0.0207)
Public sector share	0.118	0.360***	0.210	0.209***
	(0.0935)	(0.0694)	(0.159)	(0.0689)
Minimum wage	0.00867	−0.0734**	−0.0438	−0.0204
	(0.0151)	(0.03375)	(0.0511)	(0.0182)
Centralization of wage bargaining	0.00688*	0.0144**	0.00660	0.00986**
	(0.00396)	(0.00641)	(0.00746)	(0.00448)
Coordination of wage bargaining	-0.00215	0.00231	0.00706	0.00244
	(0.00233)	(0.00530)	(0.00529)	(0.00294)
Age group fixed effects	Yes	Yes	Yes	Yes
Individual controls	Yes	Yes	Yes	Yes
Country fixed effects	Yes	Yes	Yes	Yes
Observations	109,116	19,081	12,700	51,461

Notes: Dependent variable is employment last 12 months. Estimated standard errors corrected for clustering at country level. *, **, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included Belgium, Check Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, work experience, work experience squared and immigration status, and parents immigrant status.

weaker in countries with statutory minimum wages. This somewhat surprising result is however consistent with the finding of a *positive* association between minimum wage and wage returns to skills in Hanushek *et al.* (2015).

Broadly speaking, our findings in Table 7 column (2) are consistent with the hypothesis that countries with high share of employment in the public sector and centralized collective bargaining system, are assosicated with reduced employment possibilities for persons with a combination of weak health and low cognitive skills. To some extent, this finding resembles the results for individuals with poor health in Bratsberg *et al.* (2013).

6.2. Immigrants

Column (3) in Table 7 reports the estimated interaction terms for the employment–skill gradient for the subsample of individuals defined as immigrants (not born in the country). Whereas the estimated coefficients have the same sign and most of them are numerically in the same ballpark as for the whole sample, none of the estimated interaction terms are statistically significantly different from zero at conventional levels. Thus, our results for immigrants do not give any clear answer whether the findings for immigrants in Bratsberg *et al.* (2013) based on data from Norway, Canada, and USA can be generalized to a larger sample of countries. A problem with the immigration variable in our data set is that the composition of this group is likely to differ a lot across host countries in terms of source country background. It would be interesting to explore the relationship across individuals from different groups of source countries. The fact that information on source country background is not available for many countries prevents further study along these lines. Further, the number of individuals not born in country is already very small in a number of countries included in the regressions.

6.3. Individuals with low formal education

Column (4) in Table 7 reports the estimated interaction terms for the employment–skill gradient for the subsample of individuals with low formal education defined as persons with no more than 12 years of schooling. For all institutional variables except the minimum wage variable, the sign of the interactions is positive. Whereas the coefficient for the strictness of employment protection interaction term was imprecisely estimated for the two former subgroups, the estimate for the individuals with low education is significant at a 10% level and numerically in the same range as for the whole sample.[15]

Turning to the interaction between skills and public sector employment share, the coefficient is precisely estimated, almost the double of that found for the whole sample, but smaller than for the group with weak health. The coefficient estimate implies that an increase in the public sector share by 15 percentage points (the difference between USA and Norway) is associated with 3.1 percentage points increase in employment probability per one standard deviation increase in numeracy skills for people with weak health. This is larger than for the whole sample, while smaller than for the subgroup with poor health. Finally, the estimated coefficient for the interaction term for bargaining centralization is also between estimates from individuals with poor health and the whole sample. According to the estimate, an increase in bargaining centralization from the lowest (1 in UK and USA) to the highest in the sample (5 in Finland) is associated with 3.9 percentage points increase in employment probability per standard deviation increase in numeracy skills for the individuals with low education.

7. Conclusions

This paper exploits the rich comparable micro cross-country skill data provided by PIAAC to investigate the association between employment propensity and skills across countries with different labor market institutions. Whereas results vary somewhat between specifications, for the whole sample of individuals, most of our regressions indicate that the employment–skill gradient is steeper in countries with strict employment protection rules, high bargaining coverage, and centralized collective bargaining systems. When analyzing subsamples of individuals belonging to vulnerable groups in the labor market, the results are somewhat mixed, but we find that centralized bargaining is associated with a higher employment–skill gradient for individuals with weak health and individuals with low formal education. For individuals with little formal education, we find a steeper employment–skill gradient than the whole sample, in countries with high public sector employment share and centralized collective bargaining system. Broadly speaking, our findings are consistent with the hypothesis that in countries with a compressed wage structure due to a high public sector share and centralized bargaining, the employment–skill gradient is particularly high for some groups with weak attachment to the labor market in the first place. This may represent a challenging situation in such countries as the demand for high-skilled (low-skilled) workers are likely to increase (decrease) in the future due to rapid technological change and globalization.

Appendix A

Definitions, sources, and data for labor market institutional variables

Definitions and sources of indicators for labor market institutions.

Union density measures the share of wage and salary earners who are trade union members Source and description: Hanushek *et al.* (2015) footnote 47, p. 121.

Union coverage measures the number of employees covered by the collective agreements, divided by the total number of wage and salary earners. Source and description: OECD (2019) and Hanushek *et al.* (2015).

Employment protection index is a composite index indicating the strictness of employment protection against individual and collective dismissal of employees on regular contracts. It is a weighted sum of subindicators concerning the strictness of employment protection against individual (weight 5/7) and collective dismissal (weight 2/7). Source and description: Hanushek *et al.* (2015) footnote 47, p. 121.

Minimum wage is a dummy taking the value 1 if the country has a statutory minimum wage. Source and description: Hanushek *et al.* (2015) footnote 47, p. 121. We extend this variable by the minimum wage relative to the median wage in the country in 2011, source: OECD (2019).

Public sector share is the share of wage and salary workers employed in the public sector (calculated from PIAAC data). Source and description: Hanushek *et al.* (2015) footnote 47, p. 121.

Index of wage bargaining centralization measures the level at which bargaining takes place taking into account additional enterprise bargaining, articulation, legal status of derogation and existence, and use of general opening clauses and ranks the countries on a scale from a minimum of 1 (low) to a maximum of 5.75 (high). Year: 2011. Source and description: Visser (2015)

Index of wage bargaining coordination measures the degree of ‘intentional harmony’ in the wage-setting process and the degree to which minor actors (sectors or unions) follow major actors in the wage-setting process, see Kenworthy (2001) p. 75. The index we use ranks the countries on a scale from 1 (low) to 5 (high) according to the degree of wage bargaining coordination. Year: 2011. Source and description: Visser (2015)

Table A1. Data for the labor market institutional variables

Country	Public sector share	Minimum wage	Product market regulation	Strictness of employmnt protection index	Bargaining Coverage	Union density	Coordination	Centralized wage bargaining
Austria	0.2593209	0.44	1.381	2.442177	.99	0.27769651	4	3.05
Belgium	0.2942857	0.43	1.369	3.083333	.96	0.50374782	5	4.6
Canada	0.3332307	0.38	.957	1.505811	.316	0.26801603	1	1
Czech Re	0.2581967	0.32	1.557	2.751134	.425	0.17336349	2	1
Denmark	0.3944172	0	.992	2.320295	.8	0.68507832	4	2.4
Estonia	0.3025894	0.34	1.24	2.066327	.19	0.08066065	1	1
Finland	0.3587054	0	1.116	2.166667	.9	0.69048735	5	4.65
France	0.2640212	0.51	1.391	2.822562	.9	0.07840701	2	2.25
Germany	0.2292994	0.43	1.267	2.977891	.62	0.18045888	4	2.85
Ireland	0.3178622	0.39	.86	1.978458	.44	0.3259815	1	1
Italy	0.2752894	0	1.318	3.032313	.8	0.35611085	3	2.85
Japan	0.1332933	0.32	1.14	2.085034	.16	0.18084057	4	1
Korea	0.1648213	0.36	1.478	2.168367	.1	0.09886762	3	1
Netherlands	0.2873065	0.40	.905	2.884354	.823	0.18163291	4	2.9
Norway	0.3839581	0	1.154	2.309524	.74	0.54556427	4	2.4
Poland	0.2187422	0.37	2.202	2.39059	.38	0.14583882	1	1
Slovak Republic	0.28158	0.37	1.54	2.634921	.4	0.16687737	3	2.475
Spain	0.2494024	0.32	.961	2.557823	.845	0.15558882	3	3.05
Sweden	0.3922764	0	1.235	2.517007	.91	0.67496803	4	2.4
United Kingdom	0.3432592	0.38	.789	1.713152	.327	0.25615764	1	1
United States	0.2345938	0.28	.836	1.171429	.131	0.11329488	1	1

Appendix B

Employment–skill gradients by country

Figure B1. Estimated employment–skill gradients by country [Colour figure can be viewed at wileyonlinelibrary.com]

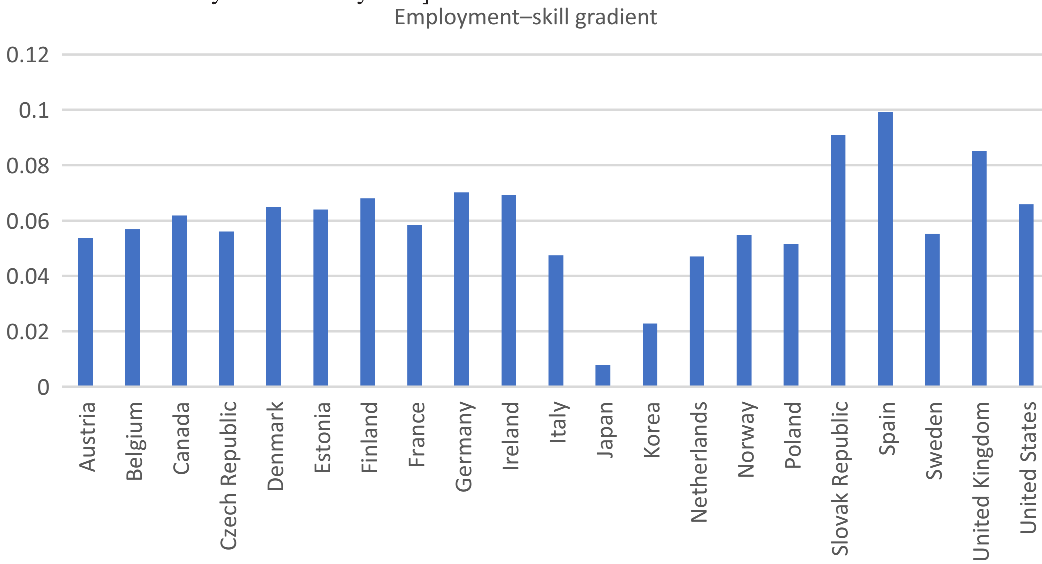
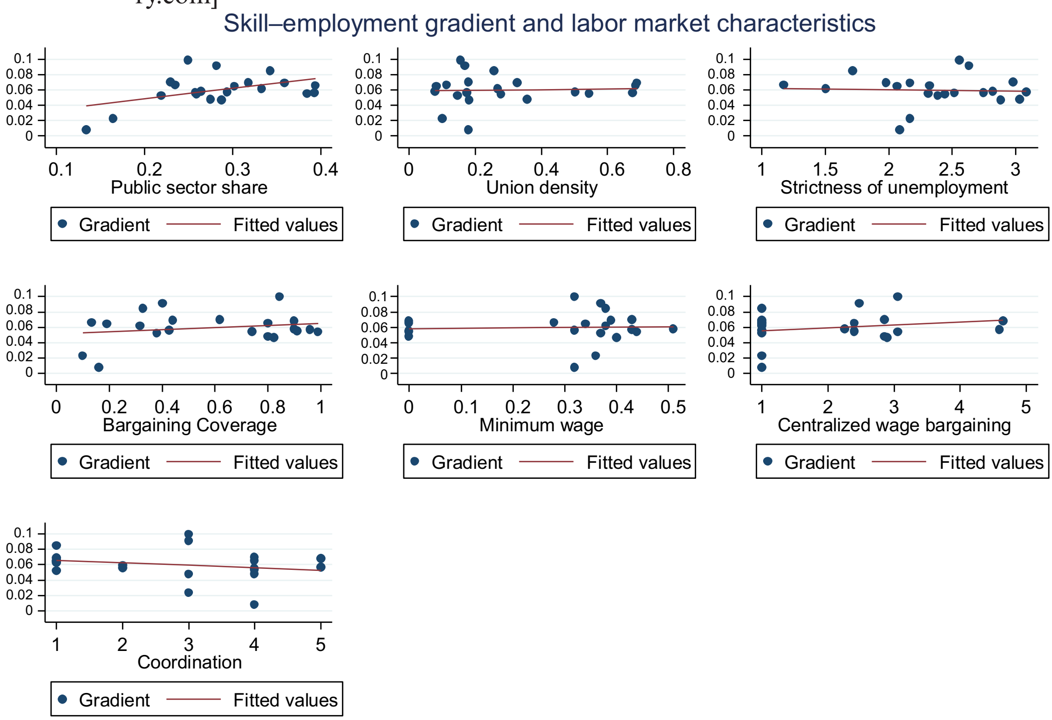


Figure B2. Cross-plot of estimated employment–skill gradients by country and labor market institutional variables [Colour figure can be viewed at wileyonlinelibrary.com]



Appendix C

Probit estimation of the model in Table 1 in main text

Table C1. Baseline model results. Probit estimation of the model in Table 1. Marginal effects reported

	(1)	(2)	(3)	(4)	(5)	(6)
Numeracy	0.0967***	0.0561***	0.0554***	0.0550***	0.0455***	0.0277***
	(0.0056)	(0.0051)	(0.0039)	(0.0036)	(0.0034)	(0.0035)
Good Health					0.104***	
					(0.0080)	
Very good health					0.1499***	
					(0.0098)	
Years of schooling						0.0189***
						(0.0012)
Observations	119,192	109,116	109,116	109,116	109,116	107,466
Individual controls	No	Yes	Yes	Yes	Yes	Yes
Age group fixed effects	No	Yes	Yes	Yes	Yes	Yes
Country fixed effects	No	No	Yes	Yes	Yes	Yes
Age group by country fixed effects	No	No	No	Yes	No	No

Notes: Dependent variable is employment last 12 months. Estimated standard errors corrected for clustering at country level. *, **, and *** denote statistical significant at 10%, 5%, and 1% levels, respectively. Sample weights used and normalized such that each country has equal weights as in Hanushek *et al.* (2015). Countries included Belgium, Canada, Check Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovak Republic, Spain, Sweden, UK, and USA. Age: between 25 and 65. Individual controls are gender, work experience, work experience squared and immigration status, and parents immigrant status.

Notes

[1] Koeniger *et al.* (2007) provide evidence that wage inequality measured by the 90–10 percentile ratios is negatively associated with union density, unemployment benefit generosity, size of the public sector, strictness of employment protection law, and the *size* of the minimum wage relative to the median wage.

[2] Kahn (2000) finds that greater union coverage and membership led to higher relative pay and lower relative employment for less skilled men as measured by educational attainment. Nickell *et al.* (2005) claims that shifts in institutional factors including unemployment benefit systems, systems of collective bargaining, and measures of union strength can explain broad movements in equilibrium unemployment rates from the 1960s to the 1990s. Bertola *et al.* (2007) find that unionization decreases employment–population ratios for young and old individuals relative to prime-aged and raise unemployment rates for prime-aged women and young men compared to prime-age men. On the other hand, Freeman and Schettkat (2001) and Barth and Moene (2012) find that countries with strong trade unions and compressed wage structures do not experience adverse employment effects for vulnerable groups defined by gender and age.

[3] Bratsberg *et al.* (2013) use data from the 2003 Adult Literacy and Life Skills Survey (ALL) to study the employment–skill gradient for groups with low employment possibilities (immigrants and individuals with poor health conditions) in Norway, Canada, and the USA. They find that the employment–literacy skill gradient is higher for these groups in Norway than in the USA and Canada and conclude that this finding is consistent with the hypothesis that low-skilled immigrants

and people with poor health conditions may find it particularly difficult to gain employment in a country like Norway with compressed wage structure due to centralized bargaining and strict employment protection regulation.

[4] Several authors note that public sector wages appear to be geographically rigid and analyze the potential negative effects on the quality of public workers and public services. See Propper and van Reenen (2010) for evidence for nurses in England, Britton and Propper (2016) for English teachers, Bonesrønning *et al.* (2005) for teachers in Norway, and Carrell (2007) for military personnel in the USA.

[5] The countries included in the main analysis are Belgium (Flanders), Canada, Czech Republic, Denmark, Estonia, Finland, France, Germany, Ireland, Italy, Japan, South Korea, Netherlands, Norway, Austria, Poland, Slovakia, Spain, Sweden, UK (England and Northern Ireland), and the USA.

[6] OECD (2019) uses the following definitions of the literacy and numeracy domains competencies: Literacy: Ability to understand, evaluate, use, and engage with written texts to participate in society, to achieve one's goals, and to develop one's knowledge and potential. Numeracy: Ability to access, use, interpret, and communicate mathematical information and ideas in order to engage in and manage the mathematical demands of a range of situations in adult life.

[7] We follow Hanushek *et al.* (2015) who use 'plausible value 1' to represent the skills. We have also conducted analysis with 'plausible value 2' with very similar results (not reported).

[8] Hanushek *et al.* (2015) also include an index for the extent of product market regulation in their study of the wage–skill gradient, but they do not find any effect of this variable.

[9] Clark (2011), Reiling and Strøm (2015) and von Simson (2015) show that enrollment in and completion of upper secondary school is countercyclical. Oreopoulos *et al* (2012) show that wages are positively correlated with the unemployment rate experienced at the start of the labor market career.

[10] A possible concern is that our use of linear probability model is not fully adequate in this setting. We have therefore also estimated the basic model in column (1)–(6) in Table 1 by probit, and the results are presented in Appendix C Table C1. As can be seen from the table, the marginal effects estimated by probit are very similar to those obtained by the linear probability model.

[11] We have also estimated the model version in column (3) separately for each country. The estimated country-specific employment–skill gradients are shown in Figure B1 in Appendix B, while Figure B2 shows the cross plots between these estimates and the labour market institutional variables. However, we base our discussion of the relationship between the employment–skill gradient and country-specific labor market institutions on the more rigorous model analysis using the pooled micro data controlling for country-specific fixed effects.

[12] However, a potential weakness of the Hanushek *et al.* (2015) and other studies of the institutional determinants of the wage–skill gradient is that no indicator for the financing system of higher education in the countries is included. It can be argued that the returns to skills and/or years of education is lower in countries with generous public subsidies in terms of highly subsidized loans and absence of tuition fees for the students in higher education. In countries without such arrangements, it is natural that the required returns to education and skills are higher just to compensate for the high individual costs related to human capital investments.

[13] One possible reason for the frequent use of bargaining coordination measures in analysis of the effect of collective bargaining institutions on wage setting and unemployment outcomes using cross-country panel data, [Nunziata (2005) and Nymoen and Sparrmann (2015)] may be that this measure varies more over time within countries than measures based on the formal centralization of collective bargaining.

[14] See the results in column (3) in Table 9, page 122 in Hanushek *et al* (2015).

[15] We have also estimated the model for individuals with no more than 10 years of schooling (not reported), and the results are qualitatively similar.

References

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