BUDAPEST WORKING PAPERS ON THE LABOUR MARKET

BWP – 2006/7

Workplace Literacy Requirements and Unskilled Employment in East-Central and Western Europe

Evidence from the International Adult Literacy Survey (IALS)

JÁNOS KÖLLŐ

INSTITUTE OF ECONOMICS, HUNGARIAN ACADEMY OF SCIENCES
DEPARTMENT OF HUMAN RESOURCES, CORVINUS UNIVERSITY OF BUDAPEST
BUDAPEST, 2006

Budapest Working Papers On The Labour Market
BWP – 2006/7

Institute of Economics, Hungarian Academy of Sciences
Department of Human Resources, Corvinus University of Budapest

Author:
János Köllő
senior research fellow
Institute of Economics Hungarian Academy of Sciences
E-mail: kollo@econ.core.hu

ISBN 963 9588 90 3
ISSN 1785-3788

December 2006

Publisher:
Institute of Economics, Hungarian Academy of Sciences

Workplace Literacy Requirements and Unskilled Employment in East-Central and Western Europe

Evidence from the International Adult Literacy Survey (IALS)

János Köllő

Abstract

Primary degree holders have extraordinarily low employment rates in Central and East European (CEE) countries, a bias that largely contributes to their low levels of aggregate employment. The paper looks at the possible role for skills mismatch in explaining this failure. The analysis is based on data from the IALS, an international skills survey conducted in 1994-98. Multiple choice models are used to study how educational groups and jobs requiring literacy and numeracy were matched in the CEEs (Czech Republic, Hungary, Poland and Slovenia) and two groups of West-European countries. The results suggest that selection to skill-intensive jobs was more severely biased against the less-educated in the CEEs than in the rest of Europe including countries hit by high unskilled unemployment at the time of the survey (UK, Ireland, Finland). The paper concludes that the skill deficiencies of workers with primary and apprentice-based vocational qualification largely contribute to the unskilled unemployment problem in the former Communist countries, more than they do in mature European market economies.

JEL: J01, J24

Keywords: Labor Economics, Human Capital, Skills

Acknowledgement

This research was supported by the Europe-Japan Cooperation Fund and the EBRD.
I thank John Micklewright and Sylke Schnepf for instructions to use the IALS data set, Tomasz Mickiewicz and Milan Vodopivec for help with publicly unavailable Polish and Slovenian data, and workshop participants in Cambridge, Massachussets and Szeged, Hungary for helpful comments.

1. INTRODUCTION

This paper looks at the possible role for skills mismatch in explaining an exceptionally strong bias against low-educated workers in CEE labor markets, relying on data from 4 CEEs and 9 West-European countries. In particular, it analyses how males with different levels of educational attainment and age, on the one hand, and jobs requiring different levels of literacy and numeracy, on the other, were matched East and West in the late 1990s. Section 2 summarizes the most important stylized facts and findings on unskilled unemployment in the CEEs. Section 3 proposes a simple framework for analyzing the unique but in many ways second-best, cross-section data at hand. These are introduced in detail in Section 4. Section 5 presents the results and Section 6 concludes.

The findings suggest that education-specific skill differentials were substantially larger in the CEEs than in Western Europe including countries fighting high unskilled unemployment. While the low-educated East-Europeans compared unfavorably to their West-European counterparts in terms of skill endowments, workplace skill requirements in their typical occupations were also less skill-intensive at the time of the survey. Looking at how jobs and workers were matched we find that in the domain of *high-skill jobs* the links between workplace literacy requirements and educational composition of the workforce were similar East and West. In the domain of *low-skill jobs*, the adverse impact of higher skill requirements on the demand for low-educated labor was significantly stronger in the CEEs that in the West. The results hold after controlling for compositional differences by sectors, occupation and firm size. The paper concludes that poor literacy skills and insufficient workplace literacy experience may have largely contributed to the exclusion from employment of less-educated East-Europeans in a period of de-industrialization and gaining ground of Western technologies. The findings do not support a similar conclusion for the UK, Ireland and Finland, western countries where unskilled unemployment was high at the time of the IALS survey: the distribution of skill endowments and skill requirements, and the patterns of matching were similar to those found in other West-European countries.

2. UNSKILLED EMPLOYMENT IN THE CEES

By the end of the post-communist transition unskilled employment rates in the CEEs fell to levels unprecedented in the OECD. In 2002, the employment ratios of workers with primary school background (*P* henceforth) ranged between 30 per cent in Slovakia and 47 per cent in the Czech Republic (CzR). This compared to the 51-63 per cent standard deviation band

around the 57 per cent mean in the western OECD. While in the last decade *P* employment was slightly rising in the western OECD it fell by two-digit percentages in the CEEs except in Hungary where the shock came a few years earlier (Table 1).

Table 1: Employment by education in five CEEs

	Below upper secondary ISCED 0-2	Upper secondary ISCED 3	Tertiary ISCED 5-7
Levels in 2002			
OECD non-CEE, unweighted mean	57	73	82
(standard deviation)	(8)	(6)	(8)
Czech R	47	76	88
Hungary[1]	37[1]	72	83
Poland	41	65	84
Slovakia	30	70	87
Slovenia[2]	40	74	87
Log changes in 1995-2002			
OECD non-CEE, unweighted mean	.013	.022	.007
(standard deviation)	(.051)	(.029)	(.025)
Czech R	-.175	-.076	-.044
Hungary	-.027	.007	.023
Hungary 1992-2002	*-.217*	*-.067*	*-.011*
Poland	-.198	-.074	-.012
Slovakia	-.262	-.069	-.011
Slovenia	n.a.	n.a.	n.a.

OECD Employment Outlook, 2003 except: 1) Hungary 1992, 1995: author's calculation from the Labor Force Survey. 2) Slovenia: Employment and labour market in Central European countries, European Commission, 2003, p. 56, population aged 15-64

The low employment ratio of *P* workers accounted for a large fraction of the OECD-CEE gap in aggregate employment. In *Table 2* the gaps of five countries are decomposed using data on their population's educational composition and education-specific employment ratios. In Hungary, Slovenia and Slovakia the gaps were almost entirely accounted for by the exceptionally low employment rates of *P* workers. In the Czech Republic this was the only component having a negative contribution while in Poland the poor job prospects of workers with secondary school attainment also added to the country's low aggregate employment level.

Table 2. Deviation of the employment-population ratio from the OECD average – Decomposition, 2002

	Aggregate employment, deviation from the OECD mean per cent	Weigthed with the educational composition of the OECD		Weighted with the educational composition of the country	
		Parameter effect of low-skilled employment	All other effects	Parameter effect of low-skilled employment	All other effects
Hungary	-7.8	-6.4	-1.2	-5.8	-2.0
Poland	-10.7	-6.1	-4.6	-3.4	-7.3
Slovakia	-5.0	-9.3	4.3	-4.1	-0.9
Slovenia	-6.0	-5.4	-0.6	-5.8	-0.2
Czech R	2.6	-3.8	6.4	-1.4	4.0

Data source: Education at a Glance, OECD, 2004. 'Low-skilled' stands for lower than upper secondary degree

Transition from a centrally planned to a market economy conveyed the two major risk factors usually held responsible for high/rising unskilled unemployment: rapid changes in the industrial composition and changes of technology and work content. The risk that these changes hurt unskilled employment can be particularly high if the educational background and previous work experience of less-educated workers hinder them in adaptation and/or their wages do not adjust sufficiently to accommodate the adverse shifts in demand. How relevant are these consideration in the transition setting?

Changes in the industrial composition. The low-skilled labor market was severely affected by a major shift of demand from agriculture and manufacturing to services and trade immediately after the collapse of communism. Demand for blue-collar workers fell substantially when industrial and agricultural production was cut by two-digit percentages while many low-educated workers were restricted in entering the tertiary sector. Workers trained for and historically employed in positions requiring no written communication and/or contact with suppliers and customers usually face difficulties in tertiary sector jobs, which entail literacy and numeracy, the use of ICT and fluent oral and written communication. The presence and implications of skills mismatch arising from the lack of literacy-based and communication skills of the former factory workers were extensively analyzed in many crisis-hit industrial areas of the West. For the probably closest counterpart see the case of the Detroit area studied in depth in Holzer (1999) and Danzinger et al. (2000).

The effects of industrial restructuring were strong in the early years of the transition but faded away in recent years. By the end of the 1990s the share of the non-tradable sector was stabilized at western levels in most countries. The available evidence on the demand effects of trade specialization furthermore suggested that in later stages of the transition demand shifts in the tradable sector were not unambiguously detrimental for unskilled labor. Woerz (2003) and Landesmann and Stehrer (2002) indicated shifts toward higher-tech and higher-skill sectors within the Visegrad-5 group, and Aturupane *et al*. (1999) and Landesmann and

Stehrer (2002) also suggested that the more developed CEEs were increasing the unit value ratios of their products. In the same time, new results by Dulleck *et al.* (2005) pointed to a move towards the low-quality product segment within low-tech industries. A paper by Egger and Stehrer (2003) on 14 manufacturing industries in the CzR, Hungary and Poland suggested that since 1993 intermediate goods trade with the EU has accounted for a considerable *reduction* of the skilled-to-unskilled wage bill ratio. Firm survey data analyzed in Köllő (2006) suggested that FDI as well as large increases in exports had significant *positive* impact on demand for *P* labor in Hungary and insignificantly positive in Romania. Assembly plants employing low-educated labor were the fastest increasing segment of the economy after 1995 in Hungary, and this probably applied to other CEEs.

Changes of technology. The available production and cost function estimates provide evidence that technological changes did not favor low-skilled labor. Based on her macro-level translog cost function estimates for Hungary 1980-2002 Tarjáni (2005) concluded that white collars and capital were used as absolute complements in production while blue collars and capital (as well as blue collars and white collars) were substitutes, implying skill biased technological change (SBTC). Using enterprise-level data for translog estimation Halpern *et al.* (2004) found relative capital-skill complementarity in large Hungarian firms, 1996-99, with the elasticity of substitution between capital and blue collars being significantly higher than that between capital and white collars.

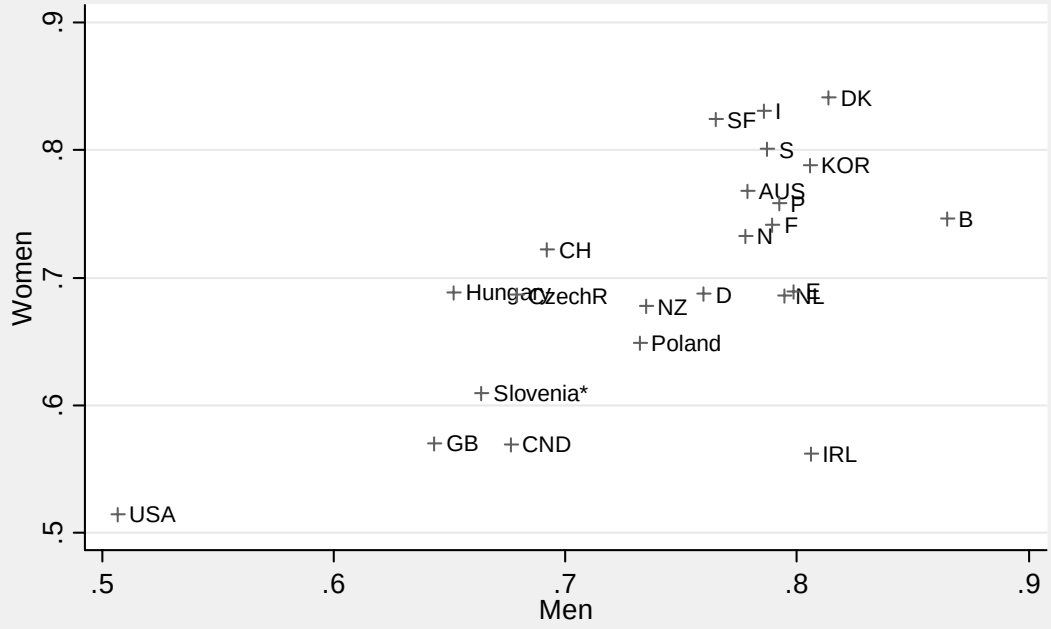
These results capture the effect of technological change on the demand for white collars versus blue collars. Less is known about how the content of work was changing within blue collar occupations. We have ample anecdotal evidence and some research results suggesting that the skill content of work did change substantially even within continuing blue collar jobs. Fifteen years ago an East-European truck driver was expected to drive his vehicle from one place to another. Today, even if employed by a firm, he is expected to deal with invoices and order lists, organize his route, communicate with shopkeepers and, quite often, have his truck maintained. Longitudinal firm survey data suggest that the returns to higher generic skills (general secondary education) grew in the blue collar occupations of continuing firms in the late 1990s in Hungary and Romania (Commander and Köllő 2002).

It is worth noting that the nature of technological change affecting unskilled labor in the CEEs was different of what is analyzed in the current western literature of SBTC, for at least two reasons. First, as shown in Autor *et al.* (2003) recent technological changes in the developed countries are dominated by the automation of routine cognitive tasks, putting repetitive white collar jobs at the highest risk. Accordingly, most papers studying the impact of SBTC are concerned with the effects of computers and R&D. As much as 41 out of 78 empirical SBTC papers reviewed by Sanders and ter Weel (2000), for instance, looked at the effect of computers and IT, and 23 addressed the impact of R&D. Most studies investigate

the impact of technological change on high school versus college graduates and even those studying the production versus non-production division deal with relatively skilled labor. In their account of what is a production worker in US manufacturing Berman, Bound and Machin (1997) showed, for instance, that in the mid-1990s 58 per cent of the production workers had high school attainment, 30 per cent had some college, and 8 per cent had college or university background. Second, technological changes in the CEEs had some specific components connected with the elimination of the socialist 'shortage economy' and the rise of competition. Quality upgrading, smaller batches and increased need for adaptation challenged the practice of employing *P* workers in core production jobs[1] while the elimination of bottlenecks and shortages obstructing the production process under communism reduced the demand for unskilled auxiliary workers.

Figure 1

The relative earnings of workers with ISCED 0-2 attainment



Data source: Education at a Glance, OECD, 2002

Wage adjustment was apparently unable to offset the decline in demand for unskilled labor in the CEEs despite the fact that the relative wages of *P* workers are significantly lower

[1] In Hungary 1986 *P* and *V* workers had roughly equal shares among cooks, waiters, bakers, tailors, upholsterers, jewelers, electricians, printers, smelters and welders, and *P* workers had high shares in dozens of further skilled blue collar occupations including carpenters (33%), bricklayers (35%), engine crews (29%), metal turners (29%), locksmiths (26%) or shoemakers (19%). Altogether, only about half of the *P* workers were employed in elementary occupations including assemblers, machine operators and farm workers while 1/3 were employed in skilled blue collar positions. (Data drawn from the Hungarian Wage Survey, May 1986)

here than in continental Europe, and fall close to levels in the English-speaking countries (except for the US where the wages of those with ISCED 0-2 attainment are extremely low). See Figure 1. Looking at the whole wage distribution we observe huge differences across educational categories in the CEEs (Sabirianova 2003). Estimates by Carbonaro (2002) using comparable measures of education and earnings from the IALS found that the association between education and wages is by far the strongest in the transition economies (CzR, Hungary, Poland and Slovenia): returns to education are roughly 2.2 times greater in the former communist nations than in long established capitalist economies. If low-skilled wages are 'too high' they are high relative to the expected productivity of less-educated workers in the millions of jobs created in the emerging market economy.

3. FRAMEWORK AND DATA REQUIREMENTS

The cross-section data at our disposal are clearly insufficient for analyzing the *process* of matching, and testing the relevance of different scenarios predicting growing unskilled unemployment in response to SBTC, import competition, or changes in the supply of skills. (Mortensen and Pissarides 1994, McKenna 1996, Acemoglu 1998, Albrecht and Vroman 2002, among others). Our efforts are simply directed at assessing how differently educated workers were matched with jobs requiring different skills in what we think of as equilibria in Western Europe and a snapshot of a changing world in the CEEs.

Let y_{ij} denote the expected productivity yield of j-educated workers (j=1,2,...,J) when employed in job type i (i=1,2,...,I), and the w_j-s their reservation wages, assumed to vary with educational attainment but not with the type of job.[2] Assuming that wages are set as a weighted average of reservation wages and the productivity yield of a given match – with $0 \leq \beta \leq 1$ standing for the relative bargaining power of employers in a country or region – the firm solves:

$$(1) \quad \max_j \pi_{ij} = \max_j (y_{ij} - w_{ij}) = \max_j [y_{ij} - (\beta w_j^* + (1-\beta) y_{ij})]$$

Suppose that job types can be characterized with a continuous or ordinal measure of complexity (R) so that $R_1 < R_2 < ... < R_I$, and that the productivity yields from employing a j-educated worker in a job of R-level complexity can be approximated with the linear projection $y_{ij} = \propto_j R_i$. Equation (1) can be re-written as:

$$(2) \quad \max_j \pi_{ij} = \max_j \left(\beta \alpha_j R_i - \beta w_j^* \right)$$

When employers decide on hiring an individual their choices are based on wages and expected productivity that they predict on the basis of the applicant's education and further proxies of his/her skills. These may be observed by the employer but not by the researcher

and are therefore summarized in a residual term ξ satisfying E(ξ)=0, cov(ξ,*w*)=0 and cov(ξ, *R*)=0. For an applicant of *j*-level education expected profit is:

$$(3a) \quad \pi_{ij} = \beta\alpha_j R_i - \beta w_j^* + \xi_{ij}$$

For an applicant for the same job with education J:

$$(3b) \quad \pi_{iJ} = \beta\alpha_J R_i - \beta w_J^* + \xi_{iJ}$$

Subtracting 3b from 3a we have:

$$(4) \quad \pi_{ij} - \pi_{iJ} = \beta(\alpha_j - \alpha_J)R_i - \beta(w_j^* - w_J^*) + (\xi_{ij} - \xi_{iJ}) = v_{ijJ} \ ,$$

and the probability that *J* is chosen for job type *i* is:

$$(5) \quad \Pr(J \ is \ chosen \ for \ i) = \Pr(v_{i1J} \le 0, v_{i2J} \le 0, ..., v_{i,J-1,J} \le 0) = F(R_i, w_j^*)$$

While the educational categories may be ordered in the sense that $y_{i1}<y_{i2}<...<y_{iJ}$, the alternatives in equation (5) are unordered since *J*>*j* does not imply $\pi_{iJ}>\pi_{ij}$. Our main concern is how the α_j-s relate to each other within a country or a region. The coefficients on *R* contain the effect of bargaining power (unless wages are set by the employers at their reservation levels implying β=1) but their relations to each other are informative of how the choice of one educational category relative to another is affected by an increase in R holding wages constant. [3] For the estimation (discussed in Section 4) we need internationally comparable and preferably continuous measures of workplace literacy requirements, and comparable indicators of educational attainment and relative wages.

4. DATA AND ESTIMATION

The IALS, conducted by the OECD and Statistics Canada in 20 countries in 1994-98 was primarily aimed at measuring the adult population's endowments with practical skills important for work as well as everyday problem-solving in a modern society. The interviews also provided information on the respondents' social background, education, labor market status, wage levels and job content. The IALS results were presented in a summary report (OECD 2000) and supplemented with instructions for micro-data users (Statistics Canada 2000). Micklewright and Brown (2004) provided a profound discussion of methodological problems arising in the IALS and some other, school-based skill surveys. Relatively few academic papers have used the IALS data as yet. Devroye and Freeman (2000) and Blau and Kahn (2000) compared the skills and wage distributions of Americans and Europeans. Micklewright and Schnepf (2004) studied the consistency of the results of different skill surveys in English-speaking countries versus the rest of the world. Several papers including

[2] In Section 5 we look at the choice of age-education interactions in additions to choice of educational levels.

Denny et al. (2004) and Carbonaro (2002) estimated augmented Mincer-type wage equations using the respondents' wage quintile position on the left hand and various literacy indicators on the right hand. McIntosh and Vignoles (2000) estimated both wage regressions and employment probits with literacy measures included on the right hand.

In contrast to the latter group of papers I do not use the literacy test scores as explanatory variables in modeling employment probabilities, wages or matching. The IALS provides information on *current* skills, ones that carry the effect of work experience, and the survey does not contain information that could be used to eliminate the resulting endogeneity bias. The variables, which may appear to be proper instruments at first sight such as parents' education, reading habits or the frequency of attending cultural events are likely to affect employment probabilities through channels other than their impact on literacy skills[4] The problem, as mentioned in McIntosh and Vignoles (2000), could best be solved by using childhood ability measures, which are regrettably not available in the survey. This paper looks at literacy performance as the output of school-based and on-the-job training rather than an input of the matching decision – an indicator that has importance in assessing future developments but not in explaining how the status quo came into being. In analyzing the latter we rely on the following key variables.

Workplace skill requirements. The interview had 13 questions on the frequency of reading, writing and quantitative tasks occurring in the respondent's job (Table 3). Workers were asked to tell if they perform these tasks (i) every day, (ii) a few times a week, (iii) once a week, (iv) less than once a week, (v) rarely or never.

Table 3: Workplace literacy requirements observed in the IALS

Reading in main job	**Writing in main job**	**Use of maths in main job**
Letters, memos	Letters, memos	Measuringt the size or weight of objects
Forms, bills, invoices, budgets	Forms, bills, invoices, budgets	Calculating prices, costs or budgets
Reports, reference books, catalogues	Reports, articles	
Diagrams, schematics	Estimates, technical specifications	
Bills, invoices, budget tables		
Material in foreign languages		
Instruction of use, recipes		
Options: 1) every day 2) a few times a week 3) once a week 4) less than once a week 5) rarely or never		

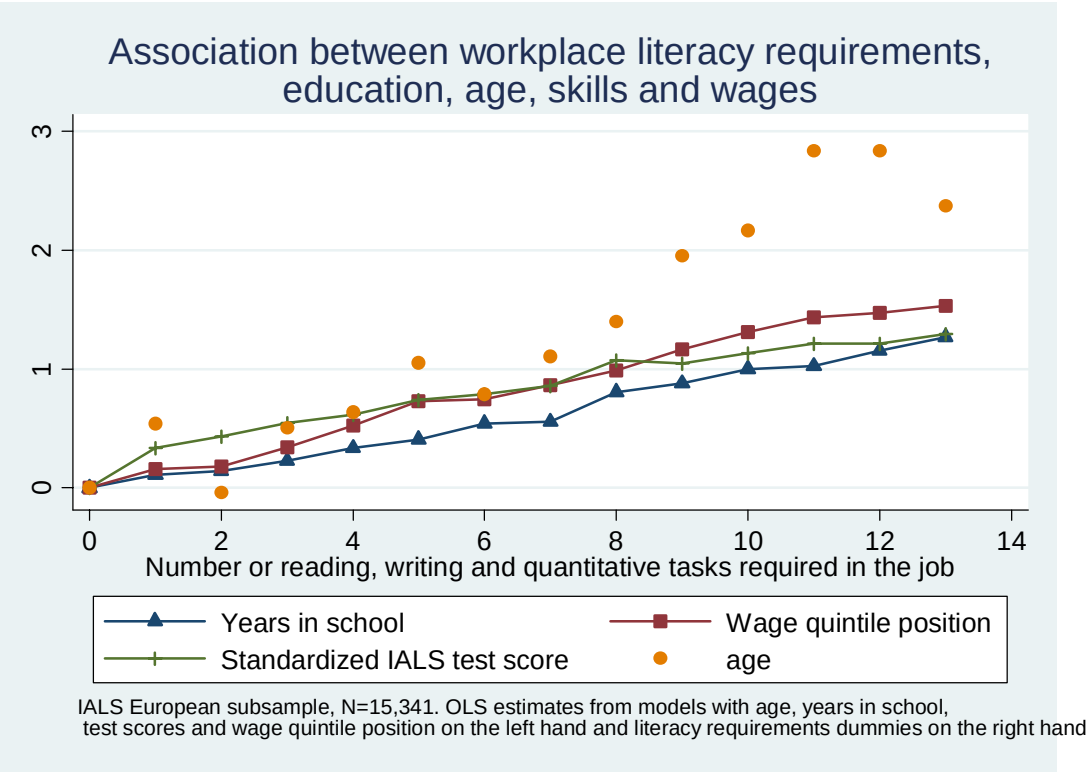
In this paper complexity will be approximated with the *number* of reading and writing tasks. "Rarely or never" was coded zero, all other valid responses were coded 1, and the 13 dummies were summed up to create a continuous measure of skill intensity (R) running from 0 to 13. This is one of a large number of possible scales that need to be tested before

[3] Assuming a skill-neutral payroll tax rate τ-1>0 the coefficient of R in equation (4) becomes $(1+\tau-\tau\beta)(\alpha_j-\alpha_J)$ and that of the wage becomes $\tau\beta$.

[4] This relates to wages, too. The IALS data suggest that workers employed in jobs with higher literacy requirements are paid higher wages, all else equal, while they are also likely to achieve higher levels of literacy over time. Their current test scores and current wages are simultaneously affected by their skills and reservation wages at the time they were selected for their jobs.

being used. In Figure 2 several proxies of individual skills such as years in school, literacy test scores and wages are regressed on dummies of the skill requirements indicator, with the resulting coefficients capturing differences in the skills of those employed at levels R=0, 1, ...,13. The estimates relate to the whole European male sample.

Figure 2



IALS European subsample, N=15,341. OLS estimates from models with age, years in school, test scores and wage quintile position on the left hand and literacy requirements dummies on the right hand

The results suggest that employees' skills increase monotonically and nearly linearly as we move from low to high levels of R. This also relates to workers' age up to R=8 but age increases at different rates below and above this level. It seems that the continuous measure constructed in the above-mentioned way can be a useful proxy of complexity. "Literacy requirements" obviously represent only a part of the competencies required and rewarded by employers - firms pay a premium for a series of further skills and even non-skills such as loyalty, friendliness and outlook (Bowles *et al.* 2001). However, most of the marketable competencies grow out of some *basic skills* closely tied to literacy as discussed in Murnane and Levy (1996, 30-32) and OECD (2000). Endogeneity issues (the question of whether R is indeed an attribute of the job rather than a choice of the employee) are discussed after presenting the benchmark estimates.

Education. Instead of using the ISCED codes available in the survey I measure education with years in school not counting repeated years. The ISCED classification, which takes into consideration the type of education a person received is, in principle, superior to the years in school measure. Unfortunately, differences in the national interpretation of the classification rules seem to destroy comparability. Countries with identical educational systems like Hungary and Poland, for instance, report sharply different shares of ISCED categories depending on how they classify short-term post-primary training. The cross-country differences in classification are largest in the range of 10-13 years in school, as shown in Table A3 of the Data Appendix. Instead of adopting this noisy categorization I use the years in school measure and distinguish between groups having 0-10 years, 11 years, 12-14 years and 15 or more years in school. Treating 11 years as a separate category is justified by the outstanding role it historically played in the CEEs. Under socialism, between one third and one half of the primary school graduates continued their studies in apprentice-based vocational training typically lasting for three years after an eight-grade primary school and providing about one year of further general education as opposed to four years in general and vocational secondary schools. Apprentice-based vocational training did not conclude in a so-called 'maturity exam', a certificate required for further studies. As much as 28 per cent of the CEE sub-sample of the IALS had 11 years in school, the largest fraction within the population. The shares were 35 per cent in Hungary, 29 per cent in Poland and Slovenia albeit only 17 per cent in the CzR. [5]

Wages. The public file of the IALS records the wage quintile position of the respondent. This measure has the advantage of being comparable across countries but also has the major disadvantage of being non-continuous. Ideally, we would need estimates of reservation wages rather than observations on actual wages since the latter are affected by the distribution of educational categories across R and therefore not exogenous to the matching decision. In estimating equation 5 I first use the observed country-and-education-specific means of the wage quintile variable. Second, I use the coefficients of education dummies from country-specific wage quintile regressions controlled for R. Obviously, both measures are crude making the respective coefficients difficult to interpret. All we expect is that the wage term has a negatively signed parameter.

Literacy test scores. Respondents of the IALS were tested in three areas of functional literacy: the understanding of simple prose like news articles and ads, the understanding and interpreting of simple documents like time-tables, prescriptions or instructions of use, and the solving of simple quantitative tasks like understanding balance sheets or invoices. In this

[5] In the CzR the primary school was 9 years long in 1960-78 and after 1990. Adjusting for these shifts suggests that the V share must have been about 30 per cent in this country, too, with 17 per cent completing 8+3 years and about 13 per cent completing 9+3 years. See Table A7 and the added note in the Data Appendix

paper literacy proficiency is approximated with the mean of the 15 'plausible values' measuring performance in the three main fields of testing. This measure has a correlation of 0.98 with the first principal component of the 15 values (Denny at al, 2004), and changes little in response to changes in the weighting scheme. Experiments with alternative measures (such as a dummy for 'low literacy' being 1 in case the respondent achieved Level 1 or 2 out of 4 levels at all three tests) suggested that the qualitative results are robust to the choice of indicator. (For descriptive statistics on the test scores see Table A4 of the Data Appendix).

Sample. Throughout the paper the analysis is restricted to men aged 15-59 excluding students, and a subset of the participating countries. Women were excluded because their labor market states were recorded in sharply different ways in the participating countries, with the share of a vaguely defined 'home-maker' category varying between 1 per cent and 40 per cent (Table A2). Older people were excluded to mitigate the impact of variations in the mandatory retirement age, a problem that falls outside the scope of this paper. The overseas countries were dropped because a comparison of Europe with Chile, Canada, New Zeeland and the US would have involved the discussion of institutional differences leading away from the paper's key problem. The three regional (German, French and Italian) sub-samples responding the questionnaire in different years in Switzerland were also disregarded. Sweden had to be dropped because of missing data on literacy requirements. The price paid for comparability is a substantial reduction of the sample size from an original 64,049 to 14,364 implying an average national sample size of 1,104. (See Table A1 of the Data Appendix.) [6]

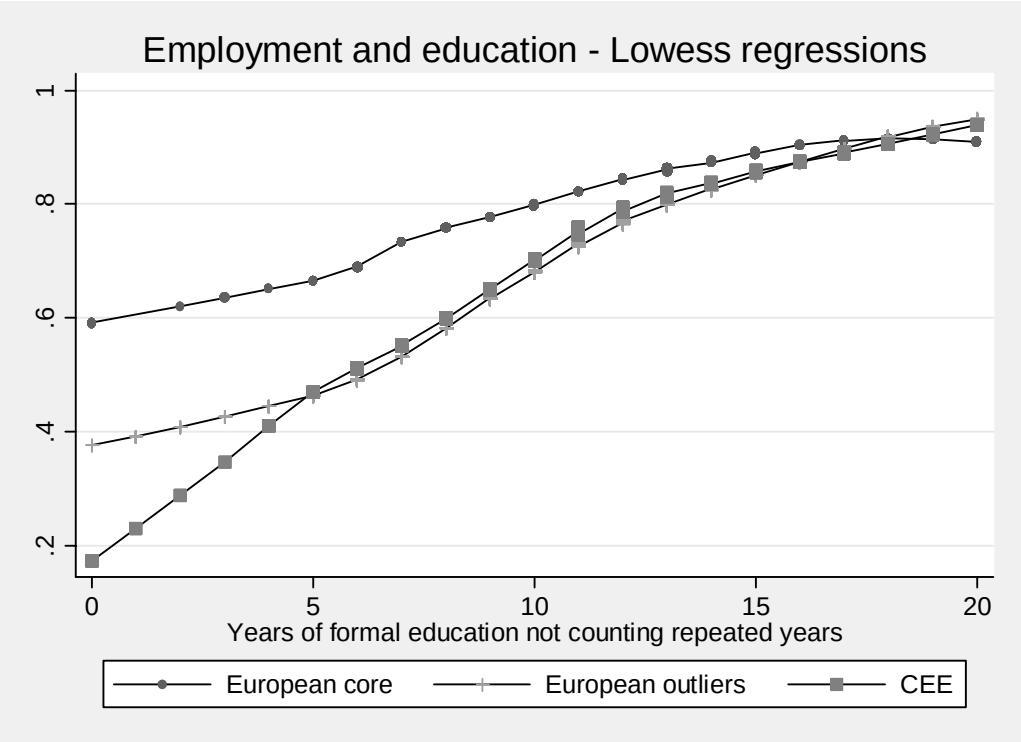
Grouping countries. The CEEs are compared to two groups of European countries selected on the basis of within-country link between education and employment. The first group (European core or EUC) is composed of Norway, Denmark, Belgium, The Netherlands, Germany, and Italy while the second (European outliers, EUO) consists of the UK, Ireland and Finland.[7] The latter group is distinguished from the core by a much closer link between education and employment. In the IALS sample, an additional year of schooling increased the probability of being employed by 0.7-1.9 per cent in the EUC countries, 3.1-4.2 per cent in EUO and 2.5-5.3 per cent in CEE. (For estimates of the education-employment relationship see Table A5 and an added note in the Data Appendix.) Non-linearities in the education-employment relationship are checked by lowess regressions for the three groups

[6] Results for Portugal were published in OECD (2000) but the country was excluded from the public file. By contrast, Italy was excluded from the survey report but included in the public file.

[7] Comparison with the 'European core' were made in two variants, one including Denmark and Germany and another excluding them. The reason was that the IALS-based employment figures for these two countries were inconsistent with their OECD Employment Outlook statistics so their data were treated with suspicion. The region-level results with and without Germany and Denmark were practically identical, however, and the two countries are included in this paper. See Table A5 in the Data Appendix.

of countries (Figure 3). The curves for the EUO and CEE are almost identical in the range of 5 or more years in school and hint at significantly lower employment compared to EUC countries in the range of 5-10 years in school.

Figure 3



The patterns observed in the IALS sample are consistent with suggestions of the OECD statistics for the CEEs, the UK and Finland (all of them had permanently low unskilled employment rates in the past decade) but not for Ireland. According to the ISCED-based statistics the level of unskilled employment was relatively high in this country, 74 per cent compared to the 68 per cent OECD average in 2002 (OECD 2004). However, the ISCED 0-2 category comprised over 42 per cent of the Irish male population at the time of the IALS survey, with this 'below upper secondary' group being rather heterogeneous itself. According to the survey data, members of the ISCED 0-2 group were almost evenly distributed between 6 and 10 years in school with 9 years being the modal level of schooling. While an additional year of schooling increased the risk of being employed by 4.2 per cent in the whole Irish sample its effect was 5 per cent within the ISCED 0-2 categories, suggesting that the ISCED-based statistics hide substantial variations.

Treating the 'Czech miracle'. Czech respondents performed much better at the IALS tests than did Hungarians, Poles and Slovenes (ranked 4th as opposed to 18th, 19th and 20th out of

20 countries), had more years in school (a median of 12 years as opposed to 11 in other countries), were employed in jobs requiring more literacy tasks (R=7.6 in contrast to 5.6, 5.0 and 6.4), and generally had higher probability of being employed (0.87 compared to 0.72, 0.67 and 0.8). Despite these differences the CzR is included in the CEE group. The skill distribution and the matching of educational categories and jobs followed similar patterns here and other CEEs therefore including or excluding the CzR did not affect the qualitative conclusions drawn from the cross-region comparisons.

Weighting. The survey covered representative samples of the population with weights attached to the respondents. The sum of weights adds up the size of the relevant population for each country. When several countries are pooled the patterns characteristic of large countries may strongly affect the region-level results. In order to assess the typical patterns in each of the regions under examination the weights were normalized so as their sum be 1 for each country, and were used as analytical weights in the estimations.

Steps of the analysis. As a first step I estimate OLS and probit equations to study how the level of education affected skill endowments on the one hand and employment probabilities, on the other in the three groups of countries. This is followed by a comparison of differentials in skill requirements. Finally, differences in matching jobs and educational levels are analyzed. Further results are presented on matching jobs and education-age interactions.

Equation (5) calls for estimation with alternative-specific multiple choice models such as conditional logit or conditional probit able to cope with the problem that R_i is a case-specific variable while w_j is alternative-specific. For a discussion of these models originating in McFadden (1974) see Wooldridge (2002, 497-503). For at least four reasons reducing the capability and reliability of the alternative-specific estimates I also estimated (5) with multinomial logit after dropping the wage term. First, the cost of doing so may not be very high as the wage variable at our disposal may be too crude to effectively control for labor cost differentials. Second, the alternative-specific models do not allow the inclusion of covariates which do not vary across possible matches within a case, such as industry or occupation dummies. (The effect of the continuous skill requirements proxy can be estimated by interacting R with dummies for the educational levels). Third, the weighting scheme available for the conditional logit and probit models (*clogit* and *asmprobit* in Stata 9.0) distort the standard errors of the coefficients. Finally, it is not possible to compute marginal effects conditional on one positive outcome per group after clogit[8]. In view of the virtues and shortcomings of the available methods I estimated several specifications starting with unweighted clogit and evaluating the significance of the estimates. Second, the model is estimated with mlogit without the wage term, controlling for 6 sectors, 7 occupations and 6

firm size categories, and weighting the observations using the normalized weights. Third, I estimate mlogit, clogit and asmprobit models in order to check how variations in the functional form, weighting and the wage proxies affect the coefficients.[9]

Note that in the alternative-specific models the units of observation are potential matches between jobs and j-educated workers (so the dataset contains 4N rather than N cases) with a choice dummy indicating which of the four possible matches was actually consummated. The equation have the choice dummy on the left hand, education level dummies (d_j) and interactions of R and the education level dummies (d_jR) on the right hand. In the mlogits the equations have the educational category variable (j=1,2,3,4) on the left hand and *R* without or with controls on the right hand.

5. RESULTS

Unskilled employment. The IALS data confirm that less-educated males had particularly low probability of being employed in the CEEs compared to more educated workers in their countries as well as their West-European counterparts. This is shown in Table 4 presenting employment probits. The reference category for the age-education interactions considered in the model were people with more than 14 years in school and younger than 36. The young college and university graduates comprised in this group had practically equal employment probabilities in the three groups of countries (91, 90 and 92 per cent, respectively). Compared to them, primary degree holders had lower employment rates by 6-10 per cent in EUC, 25-28 per cent in EUO and 34-36 per cent in CEE, depending on age. The disadvantage of those who had 11 years in school amounted to 12-14 per cent in EUC and EUO but 19-23 per cent in CEE. Young secondary school graduates (12-14 years in school) were in similar position East and West but their older counterparts had significantly lower employment probabilities in the CEEs. College and university graduates had similar employment rates East and West, irrespective of their age.

[8] For an explanation see stata.com/support/faqs/stat/mfx_unsuit.html

[9] The choice between conditional logit and the hard-to-fit simulated maximum likelihood multinomial probit depends on whether the independence from irrelevant alternatives (IIA) assumption implicit in the conditional logit holds. Since the sets of available choices are identical in all countries we have no reason to believe that the IIA assumption is hurt in a way affecting cross-region comparison.

Table 4: Employment probits - marginal effects

	European core		European outliers		CEEs	
	dy/dx	SE	dy/dx	SE	dy/dx	SE
0-10 years in school, old	-.0999	.0271***	-.2492	.0460***	-.3628	.0619***
0-10 years in school, young	-.0580	.0297*	-.2843	.0563***	-.3424	.0675***
11 years in school, old	-.1413	.0411***	-.1352	.0525***	-.2252	.0599***
11 years in school, young	-.1219	.0455***	-.1345	.0515***	-.1939	.0605***
12-14 years in school, old	-.0217	.0236	-.0547	.0457	-.1771	.0586***
12-14 years in school, young	-.0946	.0282***	-.1334	.0482***	-.1292	.0582**
>14 years in school, old	.0265	.0199	-.0106	.0437	-.0401	.0520
Rural (population<20,000)	.0295	.0112***	.0372	.0168**	.0113	.0155
Born in another country	-.0678	.0368*	-.0125	.0446	-.1179	.0662*
Mother tounge different	-.0806	.0373**	-.0866	.0576	.0616	.0461
Never attends movie, play, concert	-.0861	.0137***	-.1423	.0214***	-.0427	.0163***
Never reads books, periodicals	.0229	.0117**	-.0208	.0213	-.0055	.0184
Never reads news, journals, mags	-.1228	.0621***	-.3711	.0725***	-.1910	.0530***
Low-educated father (ISCED 0-2)	-.0069	.0150	-.0027	.0193	-.0254	.0173
Number of observations		5,844		4,382		4,068
Pseudo-R2		.0614		.0913		.0899

Dependent variable: employed at the time of the interview (0/1).
Sample: males aged 15-59 excluding students. Significant at the ***) 0.01 **) 0.05 *) 0.1 level. Young stands for people aged 15-35. European core: Norway, Denmark, Netherlands, Belgium, Germany, Italy. European outliers: UK, Ireland, Finland. CEEs: Czech Republic, Hungary, Poland, Slovenia

The data suggest a sharp dividing line between two groups of the CEE population. Young secondary school graduates and college/university graduates (of any age) had much better labor market prospects than did people without secondary education and/or older people with secondary school background. A dividing line at 12 years in school does exist in Western Europe but the employment differentials are smaller in the EUO and much smaller in the EUC than in former communist countries.

Skills of the unskilled. Second, the literacy score regressions in Table 5 suggest that the education-specific within-country differentials in functional literacy were significantly larger in the CEEs than elsewhere, with the less-educated East-Europeans performing very poorly at the tests. Similarly to the previous case, young higher education graduates were chosen as the reference category, whose test performance was similar in absolute terms in the three regions (316, 318 and 309 points on a scale of 0-500, respectively).

Table 5: Literacy test score regressions

	European core		European outliers		CEEs	
	dy/dx	SE	dy/dx	SE	dy/dx	SE
0-10 years in school, old	-1.0775	.0527***	-.9701	.0625***	-1.4780	.0791***
0-10 years in school, young	-.9498	.0613***	-1.0110	.0758***	-1.4297	.0873***
11 years in school, old	-.7287	.0627***	-.5234	.0698***	-1.0243	.0751***
11 years in school, young	-.6397	.0740***	-.6525	.0661***	-.9304	.0726***
12-14 years in school, old	-.4596	.0467***	-.3147	.0619***	-.7239	.0724***
12-14 years in school, young	-.4160	.0469***	-.3063	.0595***	-.4626	.0729***
>14 years in school, old	-.0705	.0438	-.0002	.0616	-.3213	.0732***
Rural (population<20,000)	-.0482	.0291*	-.0786	.0322**	-.0723	.0319**
Born in another country	-.4123	.0995***	-.2345	.0905***	.0452	.0884
Mother tounge different	-.5135	.0982***	-.4194	.1203***	-.2009	.0941**
Never attends movie, play, concert	-.3033	.0327***	-.4080	.0422***	-.2078	.0331***
Never reads books, periodicals	-.2716	.0310***	-.5105	.0451***	-.3275	.0400***
Never reads news, journals, mags	-.3879	.1634**	-.8928	.1642***	-.2617	.1023**
Low-educated father (ISCED 0-2)	-.1550	.0379***	-.1183	.0389***	-.1621	.0366***
Number of observations		5,844		4,382		4,068
R2		.2942		.3627		.3176

Dependent variable: standardized literacy test score, country mean=0, country standard deviation=1. Sample: males aged 15-59 excluding students. Significant at the ***) 0.01 **) 0.05 *) 0.1 level. Young stands for people aged 15-35. European core: Norway, Denmark, Netherlands, Belgium, Germany, Italy. European outliers: UK, Ireland, Finland. CEEs: Czech Republic, Hungary, Poland, Slovenia

Let us compare the CEEs to EUC first. Compared to the reference group, young secondary school graduates performed similarly in both regions. The disadvantage of CEE workers with 11 years of education was roughly as large as that of the primary degree holders in EUC, while the least educated East-Europeans lagged behind the reference category by almost 1.5 times the standard deviation of the test score. Another difference was that while in the EUC there was no significant difference between the older and younger college/university graduates, in the CEEs the elderly had a remarkable disadvantage of 0.32 standard deviation. Note that the education-specific skill differentials were larger in the young than in the old generation in the East.

In the CEEs, the dividing line in functional literacy separated young-skilled people (with at least secondary education and younger than 35) from low-skilled people in general and the skilled-old, who acquired their secondary or higher education degree under communism. This division was different from the one found in the employment probit at one point: older university graduates had relatively low skills but had high employment probabilities. Otherwise the positions of the age-education groups in terms of relative skills and relative employment rates were similar.

In the EUO group – western countries with high unskilled unemployment – skill differentials *did not* follow the CEE patterns: the disadvantages of the less-educated groups were either similar or smaller than in the EUC group. One remarkable difference is that there

seems to be a closer link between low literacy and lack of some basic cultural activities (reading printed material, attending movie, play or concert) here than in continental Europe.

Other variables included in the employment probits and the literacy score regressions do not reveal surprising patterns. The sons of low-educated fathers have lower literacy skills in all regions, by similar rates, and their employment probabilities do not differ from those of other respondents. Immigrants have lower probability of being employed in EUC and CEE and lower skills in EUC and EUO. Respondents whose mother tongue was different from the language of the interview performed significantly worse at the tests in all regions but only in the EUC had this group lower employment probability. The rural-urban differences appeared to be relatively small and proved insignificant in most cases.

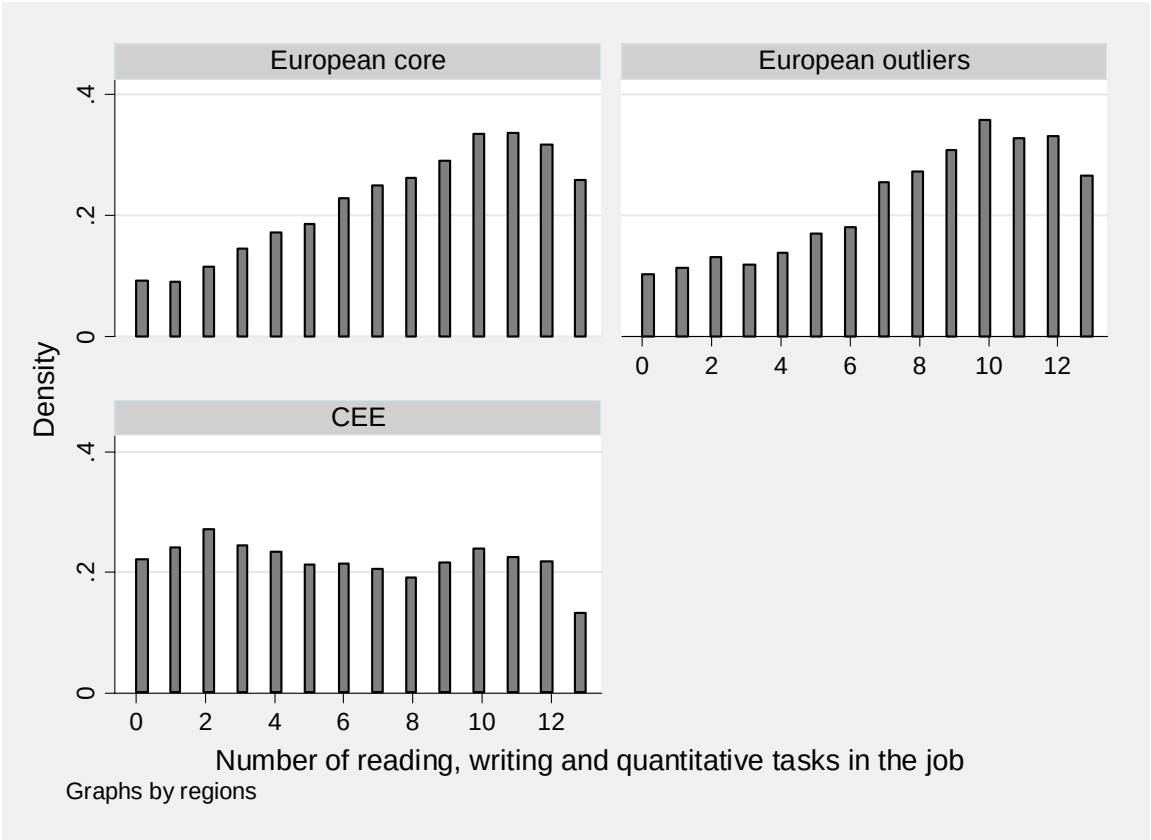
In evaluating the education-specific skill differentials detected in the test score regressions it should be borne in mind that these may contain the effect of innate abilities, so they do not necessarily hint at deficiencies in school-based education. In a country where only a few people stop at the primary level of education they are expected to have lower measurable skills, on average, than their counterparts in a country, where *P* education is prevalent. Assuming ability-based selection in the school system, logistic ability distribution, and linear relationship between ability and the test score, for instance, the expected relative score of the median primary degree holder would be proportional to $-ln[(2-P)/P]$ if *P* is the fraction of *P* workers.[10] The relative skills of *P* workers in the CEEs may reflect such a selection effect as the share of *P* attainment is relatively low in European comparison (33 per cent in the typical EUC country, 36 per cent in EUO and 23 in CEE). This, however does not apply to P+V qualifications together (40, 55 and 51 per cent) while we find this larger population to have significantly lower skills, too. The education-specific skill differentials also convey the effect of unequal access to work-based literacy experience as discussed in the forthcoming paragraphs.

Literacy requirements. Turning to the patterns of demand for literacy-based skills East and West, Figure 4 reveals characteristic differences between the CEEs and the rest of Europe. In both the EUC and EUO the size distribution of jobs by R is skewed on the right with the mode being at 10-12 tasks. In CEE the distribution is bi-modal with the upper mode located at 10-12 tasks and the lower one located at R=2.

[10] Denoting ability with $A \in (0,1)$, that of the median *P* worker with A_P and the test score with $S=\beta A+u$ we have $\Pr(A \leq A_P)=1/(1+exp(-A_P))=P/2$. From $A_P = -\ln[(2-P)/P]$ we have $E(S)=-\beta\ln[(2-P)/P]$ for the median *P* worker.

Figure 4

The size distribution of jobs by literacy requirements (unweighted)



On average, CEE jobs require less literacy tasks by only 1.5 or about 20 per cent of the mean. (In the whole sample R has a mean of 7.4 with a standard deviation of 3.8). The difference varies largely across sectors and occupations, however. (Table 6). It is negligible and statistically insignificant in high-skill occupations such as managers, professionals and assistant professionals, modest in services and trade, and large in the industrial blue-collar occupations. Similarly, in a breakdown by sectors the difference appears to be small in the tertiary sector, larger in manufacturing and very substantial in agriculture, mining and quarrying. Skill requirements in small firms (less than 20 employees) are closer to the European mean than those in medium-sized and large firms.

In countries of the EUO skill requirements *do not* differ markedly from those in the EUC. In some sectors and occupations as well as in a region-level comparison (controlled for industrial composition) they appear to be marginally *higher*.

Table 6: Literacy requirements – Deviations from the European core

	European outliers	CEE
Sectors		
Retail trade, hotels and restaurants	0.1	-0.9***
Services	0.3*	-1.1***
Manufacturing, gas, water, construction	0.4***	-1.5***
Agriculture, mining and quarrying	0.2	-2.8***
Sector unknown	-0.4	-1.6***
Occupations		
Legislators, senior officials, managers	0.3*	0.2
Professionals	0.6***	0.2
Technicians, associate professinals	0.0	-0.4
Service, shop and market sales workers	-0.1	-0.8***
Craft and related trades workers	0.3*	-1.6***
Plant and machine operators, elementary occupations	-0.2	-1.1***
Occupation unknown	0.0	-1.3***
Firm size		
<20	0.3	-1.1***
20-100	0.2	-1.7***
101-200	0.1	-1.7***
201-500	0.3	-1.8***
Over 500	0.3	-1.5***
Firm size unknown	-0.6*	-3.4***
Regional differences		
Uncontroled	0.2**	-1.5***
Controlled for sector, occupation and firm size	0.2**	-0.9***
The figures are coefficients from regressions estimated separately by sectors, occupations and firm size categories, respectively, with the number of literacy tasks on the left hand (scale 0-13) and two country-group dummies on the right hand. Those in the last two rows are coefficients estimated for all workers without/with controlling for sector, firm size and occupation. N=10,695. Significant at the ***) 0.01 **) 0.05 *) 0.1 level		

Matching jobs and educational categories. Table 7 summarizes the results of the multiple choice models. In each model secondary education was treated as the reference. The coefficients of the education dummies and other controls are not displayed.

Starting with our preferred specifications, both the unweighted clogit and the controlled mlogit suggest that low-educated East-Europeans were strongly disperefered in skill-intensive jobs. In the CEEs, a unit change in *R* reduced the log odds of primary degree holders relative to secondary school graduates by about -0.3 in the clogit and -0.24 in the controlled mlogit. In the EUC and EUO groups the respective coefficients were about -0.15 and -0.11, just about half of the CEE values. Similarly, East-Europeans having 11 years in school were disprefered in skill-intensive jobs, although in this case the coefficients for the EUO group are halfway between the EUC and CEE estimates. Higher levels of R increased the probability that a college or university graduate (as opposed to a secondary school graduate) was chosen for the job. This applied to all regions but the effect was stronger in the CEEs than elsewhere. The clogit coefficients are 0.25 in the CEEs and 0.15 in the EUC and EUO groups, while in the controlled mlogit the respective coefficients are about 0.13 versus 0.09 in the EUC and 0.07 in the EUO.

Table 7: Matching jobs and levels of education – Estimates of equation (5)

	Weights	Variables	EUC	EUO	CEE
Conditional logit	No	Wage [a]	.5197	-1.7362	-1.3401
			(.1102)	(.1633)	(.1904)
		0-10 years	-.1534	-.1380	-.3186
			(.0105)	(.0131)	(.0173)
		11 years	-.0632	-.0796	-.1810
			(.0156)	(.0134)	(.0129)
		>14 years	.1438	.1849	.2496
			(.0109)	(.0159)	(.0173)
		Observations	20,880	14,424	13,316
		Pseudo-R2	.1359	.0765	.1533
Multinomial logit controlled for sector, occupation and firm size	Yes	0-10 years	-.1104	-.1235	-.2429
			(.0133)	(.0179)	(.0211)
		11 years	-.0168	-.0832	-.1107
			(.0180)[n]	(.0184)	(.0158)
		>14 years	.0956	.0689	.1336
			(.0146)	(.0236)	(.0214)
		Observations	5,220	3,606	3,329
		Pseudo-R2	.1409	.1457	.2246
Conditional logit	Yes [b]	Wage	.3222	-1.4174	-1.6992
		0-10 years	-.1489	-.1579	-.3035
		11 years	-.0477	-.1077	-.1633
		>14 years	.1414	.1448	.2599
Conditional probit	Yes [b]	Wage	.3072	-.9302	-1.4215
		0-10 years	-.1327	-.0714	-.2283
		11 years	-.0393	-.0191	-.1305
		>14 years	.1482	.1333	.2382
Conditional logit	No	Adjusted wage [c]	.8107	-1.4404	-1.9045
		0-10 years	-.1547	-.1392	-.3180
		11 years	-.0613	-.0735	-.1727
		>14 years	.1456	.1973	.2414
Conditional logit	Yes [b]	Adjusted wage [c]	.4236	-.9212	-1.6853
		0-10 years	-.1487	-.1612	-.3035
		11 years	-.0461	-.1032	-.1618
		>14 years	.1432	.1551	.2481
Multinomial logit [d]	Yes	0-10 years	-.1464	-.1664	-.3026
		11 years	-.0445	-.1097	-.1708
		>14 years	.1428	.1549	.2401

For the specifications see the text. All coefficients are significant at the 0.01 level. In the conditional logits standard errors were adjusted for clustering by individuals.

a) Mean wage quintile position of j-educated workers in country k (13x4=52 unique values)

b) Standard errors questionable

c) Regression-adjusted wage quintile position of j-educated workers in country k measured with the coefficients of education dummies in country-specific wage quintile regressions controlled for experience, literacy requirements, rural/urban residence, sectors and occupations, treating 12-14 years in school as the reference. (13x3=39 unique values with the wage of secondary school graduates set to zero in all countries)

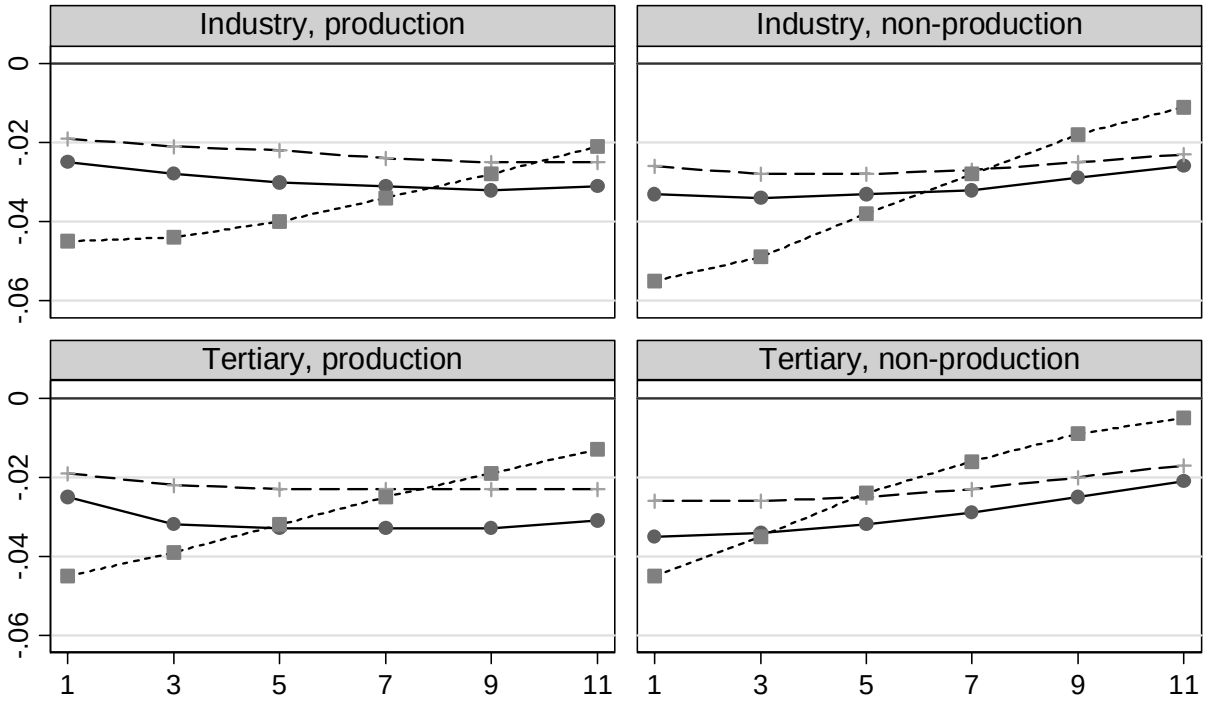
d) Uncontroled, with only the literacy requirements proxy (R) on the right hand

The coefficients on the wage terms included in the clogits and the asmprobit are negative, as expected, in the CEEs and the EUO but consistently positive in the EUC. This is a counter-intuitive result, which may arise if wages do not reflect scarcities, are mismeasured, or the estimates are biased by omitted variables capturing the constraints and costs of adjustment. It seems, however, that the inclusion/exclusion of the imperfect wage measures have no remarkable impact on the coefficients on R. The weighted and unweighted clogit models yield similar coefficient estimates. The conditional logit and conditional probit estimates are similar in the EUC and the CEEs but differ at some points in the EUO (the coefficients on R×11 years in school and the wage are smaller in absolute terms). However, the similarity of EUC to EUO and the dissimilarity of CEE to Western Europe as a whole are clearly reflected in both models.

The charts in Figure 5 look at the marginal effects of R. These may differ at low and high levels of R so I estimated them at R= 1, 3, 5, 7, 9, 11 and different values of other controls using the coefficients of a controlled mlogit containing R, a sector dummy, an occupation dummy, firm size dummies and a rural/urban dummy on the right hand. Retail trade, transport and private and public services were treated as 'tertiary' and other sectors were labeled as 'industry'. Legislators, managers, officials, professionals, assistant professionals and technicians were treated as non-production workers. The Eastern and Western results are compared by type of job in Figure 5.

Figure 5: The marginal effect of workplace literacy requirements on the probability of employing differently educated workers (multinomial logit)

0-10 years in school



11 years in school

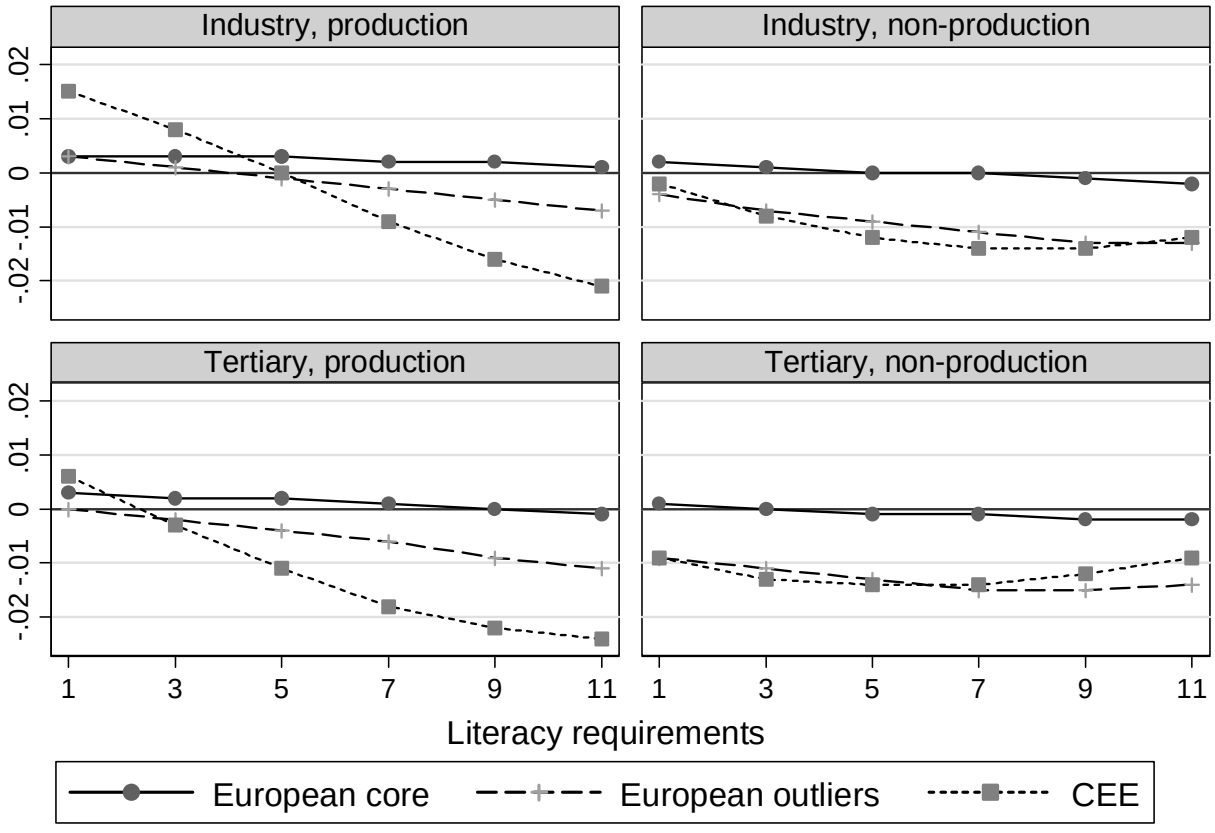
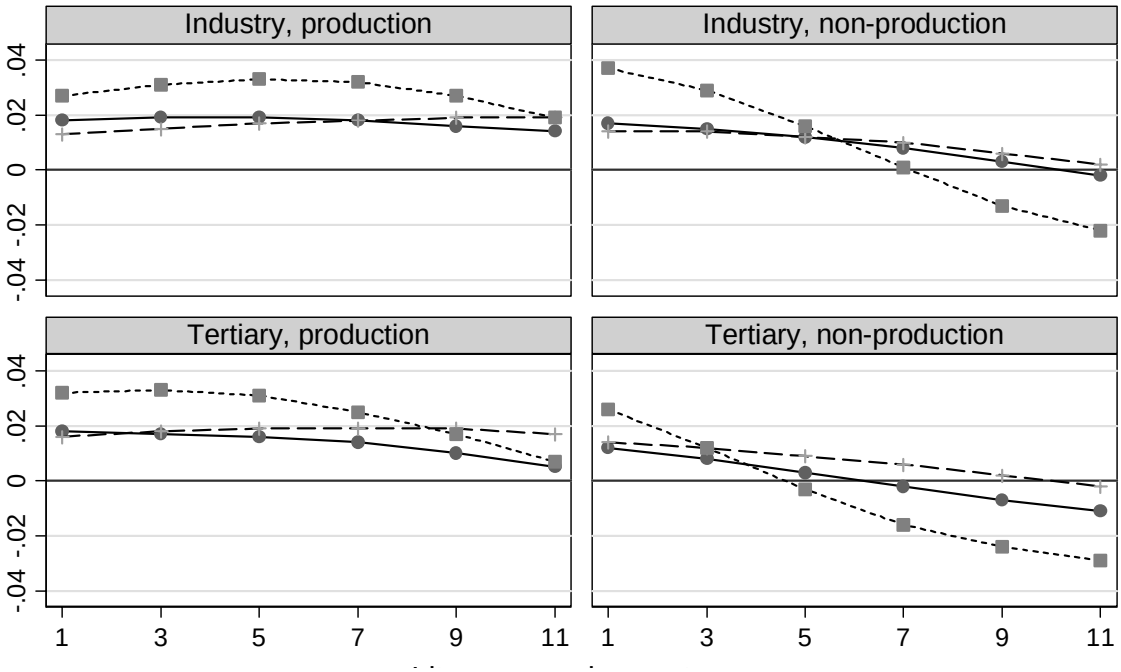
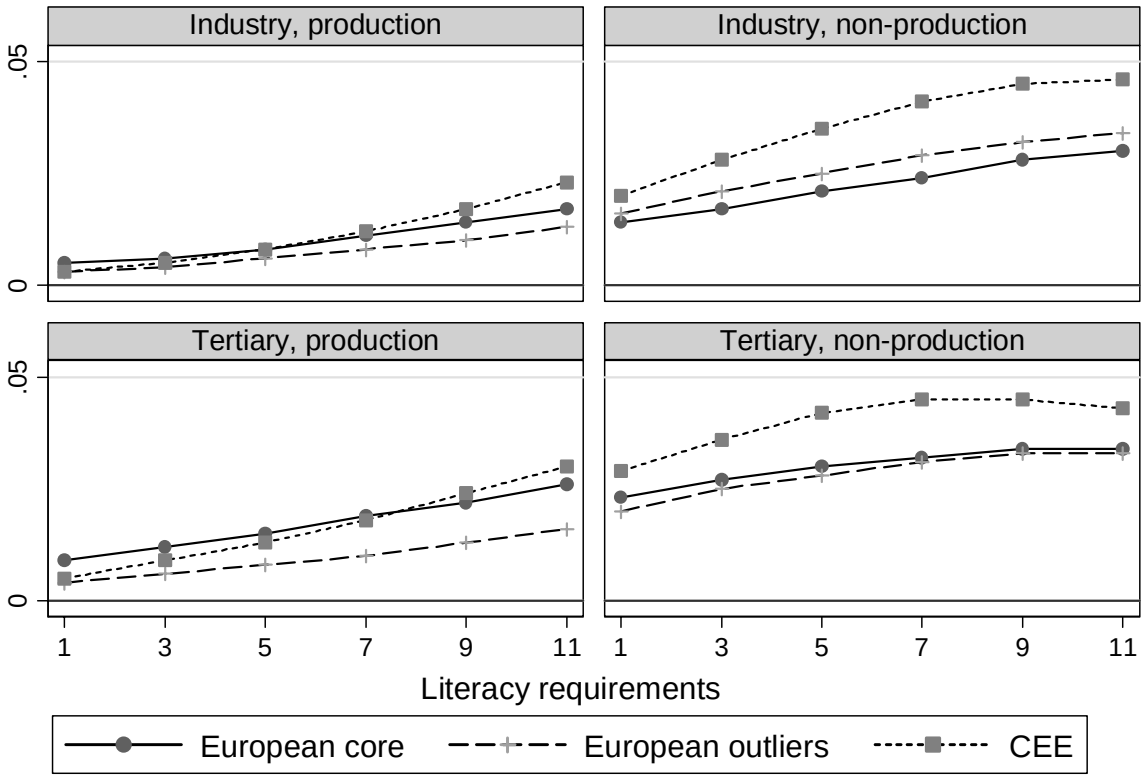


Figure 5 continued

12-14 years in school



15 or more years in school



Note: The models used to estimate the marginal effects were similar to those in block 2 of Table 7 except using a tertiary sector dummy and a production/non-production dummy in place of the industry and occupation dummies. The marginal effects were evaluated at firm size below 20 workers and urban firms.

Starting with the charts for *primary degree holders*, we observe particularly strong adverse effects of changes in R at low levels of R in the CEEs. The marginal effects are similar in the EUC and EUO and substantially weaker than in the CEEs (least so in tertiary non-production jobs). As R changes from 1 to 2 the implied fall in the share of primary degree holders amounts to 4-6 per cent in the CEEs as opposed to 2-3 per cent in Western Europe. At higher levels of R the differences become far smaller.

The East-West differentials are also remarkable with respect to demand for workers with *11 years in school*, at least in production jobs. In the CEEs, at low levels of R an increase in R has a *positive* impact while further increases in R tend to reduce the share of vocationally trained workers. The marginal effect curves are similarly sloped but less steep in the EUO. In non-production jobs the CEE and EUO curves are almost identical and different from that of the EUC. While in the EUC an increase in R has no effect at all, in the other two regions a unit change in R tends to reduce the share of this educational category by about 1.5 per cent throughout the range of R.

In the case of production jobs skill intensity has positive impact on the share of *secondary school graduates* (12-14 years in school) in all regions, especially in the CEEs. In non-production jobs the marginal effect curves are steeper in the CEEs than elsewhere, suggesting that higher skill requirements have positive impact on the demand for secondary school graduates in the range of relatively simple jobs while the effect is negative in the range of more complex jobs. The marginal effect curves are closest to each other in the case of *university/college graduates* hired for production jobs. In non-production jobs, increases in skill intensity affect the demand for college graduates positively in all regions and at all ranges of R but the effects are stronger in the East than in the West.

Before drawing conclusions two issues relevant for the interpretation are touched upon.

In the CEEs, where skill intensity is most probably increasing over time, the correlation between education and R may be driven by the age of the observed matches. Pairs of low-skilled workers and low-R jobs may be heavily over-reperesented among older matches, concluded under communism, while we may not observe strong bias against the less-educated in more recently consummated matches. In order to assess if the links between literacy requirements and educational attainment vary with age the controlled mlogit estimations were repeated for interactions of age and education, distinguishing between unskilled and skilled workers (0-11 versus 12 or more years in school) and 'young' and 'old' workers (aged 15-35 versus 36-59). As shown in Table 8 old and young unskilled workers were roughly equally disprefered in the high-skill jobs of the CEEs, so we have no reason to

believe that the East-West contrast discussed in the previous paragraphs is a statistical artifact arising from correlation between the level of education and age.[11]

Table 8: The effect of literacy skill requirements on the probability of choosing education-age categories

Multinomial logits controlled for sector, occupation and firm size. Coefficients and standard errors

	EUC	EUO	CEE
Unskilled-old	-.1076	-.0838	-.1475
	(.0147)	(.0190)	(.0198)
Unskilled-young	-.1122	-.1158	-.1804
	(.0178)	(.0202)	(.0197)
Skilled-old	.0181	.0626	.0231
	(.0131)[x]	(.0206)	(.0168)[x]

Reference category: skilled-young (at least 12 years in school, aged 15-35. The coefficients of 5 sector, 6 occupation and 6 firm size dummies and the constants not displayed.
x) not significant at 0.1 level. Unmarked coefficients are significant at 0.01 level

Finally, the validity of the results depend on whether *R* (as observed) is indeed an exogeneous attribute of the job or determined by the worker. In a *laissez-faire* regime we expect strong correlation between education and *R* as those lacking the necessary skills may simply choose not to read and write as much as do their more educated colleagues. If their employers face soft budget constraints they may accept their employees' choices without harming the firm's survival. Further, in self-employment and some small firms it is clearly the worker who determines *R* by starting a business matching his/her school-based skills. We have three reasons to believe that endogeneity does not bias the cross-region differentials observed in this paper. First, *R* was based on the *incidence* of certain reading and writing tasks. While workers have some degree of freedom in choosing the intensity of dealing with these tasks they are hardly in the position to neglect them completely. Second, ten years after the start of transition CEE firms did not face soft budget constraints. Third, as shown in Table 9 the patterns of cross-region differentials do not change markedly if we exclude self-employed workers and SME employees from the sample except that the EUO and CEE estimates get closer to each other in terms of the effect of *R* on choosing workers with 11 years in school.

[11] Unfortunately, the IALS has no data on tenure that would enable the study of how the composition of the workforce varies with the age of the match East and West.

Table 9: Multinomial logit coefficients for the whole sample and a reduced sample excluding SME employees (<100 workers) and self-employed persons

	European core Estimation sample		European outliers Estimation sample		CEE Estimation sample	
	Whole	Reduced	Whole	Reduced	Whole	Reduced
0-10 years	-.1104	-.0971	-.1235	-.1404	-.2429	-.2187
	(.0133)	(.0188)	(.0179)	(.0232)	(.0211)	(.0265)
11 years	-.0168	.0033	-.0832	-.1043	-.1107	-.1148
	(.0180)[n]	(.0277)[n]	(.0184)	(.0227)	(.0158)	(.0214)
>14 years	.0956	.0659	.0689	.0735	.1336	.1406
	(.0146)	(0209)	(.0236)	(.0269)	(.0214)	(.0287)
Observations	5,220	1,811	3,606	1,678	3,329	1,638
Pseudo-R2	.1409	.1290	.1457	.1473	.2246	.2493

Note: The specifications are identical with that in block 2 of Table 7. All coefficients are significant at the 0.01 level except n: insignificant at 0.1 level

6. CONCLUSIONS

The newly created market economies of CEE apparently failed at absorbing less-educated labor after the 'transformational recession'. This is a common feature of countries conducting different employment policies, running different benefit regimes, privatizing fast or slow, having different shares of the informal economy[12] , different degrees of ethnic diversity and different rates (even sign) of net migration. While the CEEs followed divergent development paths they all started with a common legacy of distortions in technology and education. Isolated from the rest of the world, millions were trained for and employed in jobs requiring little or none of a series of basic literacy-based skills, which became increasingly important in the post-industrial age.

We looked at traces of this legacy in the supply and demand for skills and the way workers and jobs were matched about ten years after the start of the transition. We found that the low-skill labor market was still rather large in the CEEs in the second part of the 1990s. This was only partly explained by the industrial composition of their economies. Agriculture, mining and manufacturing continued to offer blue-collar jobs with rather low skill requirements compared to similar sectors in the EU while jobs in the 'upper tiers' of the labor market required literacy levels comparable to the West.

In the domain of high-skill jobs the pattern of matching educational levels and workplace skill requirements proved to be rather similar East and West. Consistent with this, we found the literacy proficiency of high-educated East-Europeans (especially the young) to be relatively close to the European mean. In the 'lower tiers' of the labor market we found the

[12] Johnson at al (1997) estimate that in 1989-95, on average, the share of the unofficial economy was 0.12 in Slovakia, 0.13 in the CzR, 0.18 in Poland and 0.29 in Hungary. Other estimates also indicate huge variations in the share of the informal economy.

patterns of matching to be different. In the CEEs, small differences in complexity were associated with huge differences in the composition of the labor force. The data suggested that in this segment of the market higher levels of skill intensity led to substitution away from primary degree holders in favor of vocational and secondary school graduates who possess higher levels of generic skills.

Whereas we found marked East-West differences in skill endowments, skill requirements and matching we did not find the EUC-EUO difference to be analogous with the East-West contrast. In the EUO we found no evidence of huge education-specific variations in skill endowments, remarkable deviations from the EUC in the distribution of literacy requirements, or major differences in the patterns of matching jobs and educational categories, except for one category of workers (11 years in school) in one segment of the economy (non-production jobs). The reasons underlying high unskilled unemployment in the UK, Ireland or Finland may be different from those in the CEEs, with skill deficiencies probably playing less crucial role.

The findings of this paper are based on almost ten year-old data but they have a clear message for the ongoing and future developments of the CEE unskilled labor market. As the CEEs continue to adopt western technologies and organizational schemes the distribution of their jobs in terms of skill intensity is expected to move closer to the European pattern (unless they specialize in low-skill, low-tech products that seems unlikely in view of the current trends). The lower mode of the skill requirements distribution observed in Figure 4 is expected to vanish in two ways. On the one hand, many low-skill jobs will be closed affecting P workers directly. On the other hand, we can expect further increases in the skill requirements of 'simple jobs'. The finding that at low levels of complexity a small increase in complexity induces large shifts away from primary degree holders predicts continuing squeeze out of employment of the least educated East-Europeans.

REFERENCES

Aturupane, C., S. Djankov and B. Hoekman (1999): Horizontal and vertical intra-industry trade between Eastern Europe and the European Union, Weltwirtschaftliches Archiv, 135(1), 62-81

Autor, David H., Frank Levy and Richard Murnare (2003): The skill content of recent technological change: An empirical exploration, Quarterly Journal of Economics, 118(4)

Berman, Eli, John Bound and Stephen Machin (1997): Implications of skill-biased technological change: International evidence, NBER WP 6166

Blau, Francine D. and Lawrence M. Kahn (2000): Do cognitive test scores explain higher US wage inequality?, mimeo

Bowles, Samuel, Herbert Gintis and Melissa Osborne (2001): The determinants of earnings: A behavioral approach, Journal of Economic Literature 39 (4) 1137-1176

Carbonaro, William (2002): Cross-national differences in the skills-earnings relationship: The role of skill demands and labor market institutions, Department of Sociology, University of Notre Dame, November

Denny, Kevin, Colm Harmon and Vincent O'Sullivan (2004): Education, earnings and skills: A multi-country comparison, The Institute for Fiscal Studies, Dublin, WP 04/08

Devroy, Dan and Richard B. Freeman (2001): Does inequality in skills explain inequality in earnings across advanced countries?, NBER WP 8140

Dulleck, Uwe, Neil Foster, Robert Stehrer and Julia Woerz (2005): Dimensions of quality upgrading, Economics of Transition 13 (1), 51-76

Egger, P. and Robert Stehrer (2003): International outsourcing and the skill specific wage bill in Eastern Europe, The World Economy 26(1), 61-72

Halpern, László, Miklós Koren, Gábor Kőrösi and János Vincze (2004): The budgetary effect of the minimum wage (in Hungarian with English summary), Közgazdasági Szemle 51, 325-345

Danzinger, Sheldon H., Reynolds Farley and Harry J. Holzer (2000): Detroit divided, Russell Sage Foundation

Holzer, Harry J. (1999): What employers want: Job prospects for less-educated workers, Russell Sage Foundation

Johnson, Simon, D. Kaufmann and Andrei Schleifer (1997): Politics and enterpreneourship in transition economies, Working Paper Series 57, The William Davidson Institute at the University of Michigan Business School, Ann Arbor

Landesman, Michael and Robert Stehrer (2002): Evolving competetiveness of CEEs in an enlarged Europe, Rivista di Politica Economia, 92 (1), 23-87

Machin, Stephen, John Van Reenen (1998): Technology and Changes in Skill Structure: Evidence from Seven OECD Countries, Quarterly Journal of Economics, 113(4), 1215-44

McFadden, Daniel I. (1974): Conditional logit analysis of qualitative choice analysis, in Frontiers of Econometrics, ed. P. Zarembka, New York, Academic Press, 105-142

McIntosh, Steve and Anna Vignoles (2000): Measuring and assessing the impact of basic skills on labour market outcomes, Centre for the Economics of Education, London School of Economics and Political Science, July

Micklewright, John and Georgina Brown (2004): Using international surveys of achievement and literacy: A view from the outside, UNESCO Institute for Statistics, Montreal

Micklewright, John and Sylke Schnepf (2004): Educational achievement in English-speaking countries: Do different surveys tell the same story?, IZA DP 1186

Murnane, R.J. and F. Levy (1996): Teaching the new basic skills. Principles for educating to thrive in a changing economy, Free Press, New York

OECD and Statistics Canada (2000): Literacy in the information age, OECD, Paris

OECD (2004): Education at a Glance, Paris

Sabirianova, Klara (2004): xx

Sanders, Mark and Bas ter Ween (2000): Skill-biased technical change: Theoretical concepts, empirical problems and a survey of the evidence, Paper presented at the DRUID Conference, Copenhagen, Denmark, 6-8 January 2000

Statistics Canada (2001): International Adult Literacy Survey Guide for Micro Data Users, Montreal

Tarjáni, Hajnalka (2004): Estimating some labour market implications of skill biased technological change and imports in Hungary, Hungarian National Bank, MF 2004/3

Woerz, Julia (2003): Skill upgrading in Central and Eastern European manufacturing trade, The Empirical Economics Letters, 2(6), 247-56

Wooldridge, Jeffrey M. (2002): Econometric analysis of cross section and panel data, Camridge Mass and London, MIT Press

APPENDIX

Table A1: Number of men aged 15-59 excluding students in the IALS

	Years in school				
	0-10	11	12-14	15-	Total
Sweden	323	148	312	194	977
Norway (bokmal)	369	114	439	357	1,279
Netherlands	268	95	309	457	1,129
Belgium (flanders)	120	17	247	263	647
Italy	352	86	338	259	1,035
EU-5 sub-total	**1,432**	**460**	**1,645**	**1,530**	**5,067**
Denmark	231	78	426	363	1,098
Germany	344	54	176	115	689
EU-7 sub-total	**2,007**	**592**	**2,247**	**2,008**	**6,854**
UK *	478	597	828	566	2,469
Ireland	450	132	154	81	817
Finland	345	156	313	297	1,111
CzR	86	133	519	271	1,009
Poland	400	307	256	124	1,087
Hungary	182	332	214	179	907
Slovenia	311	326	319	131	1,087
Total	**3,684**	**2,443**	**4,248**	**3,179**	**13,554**

*) Great Britain and Northern Ireland pooled

Table A2. The labor market status of women in the IALS (Europe)

	Employed	Retired	Unempl	Student	Home-maker	Other	Total
UK	60.86	5.35	8.77	2.97	13.04	9.01	100.00
Ireland	38.36	1.15	6.41	11.18	40.03	2.87	100.00
Sweden	51.62	27.23	4.84	9.90	3.75	2.66	100.00
Norway (bokmal)	69.80	1.53	2.81	9.73	6.33	9.80	100.00
Denmark	62.18	11.94	6.44	13.68	1.19	4.58	100.00
Finland	59.65	11.39	9.30	13.67	4.61	1.38	100.00
Netherlands	43.77	9.73	2.53	6.61	33.57	3.78	100.00
Belgium (flanders)	43.68	6.63	9.74	11.55	25.79	2.62	100.00
Germany	44.45	8.91	6.55	6.06	25.79	8.24	100.00
Italy	39.89	8.91	6.67	10.48	34.05	0.00	100.00
Czech R	58.94	16.25	4.24	9.67	0.14	10.76	100.00
Poland	46.52	19.04	9.41	11.15	10.70	3.17	100.00
Hungary	52.13	21.22	6.84	6.33	5.39	8.10	100.00
Slovenia	54.62	17.22	7.93	13.29	6.63	0.30	100.00
Total	48.95	10.63	6.95	7.70	20.36	5.41	100.00

Table A3. Per cent classified as ISCED 3 (upper secondary) by years in school

	Years in school not counting repeated years			
	10	11	12	13
UK*	9	16	22	34
Belgium	0	0	75	91
Poland	5	13	74	71
Germany	9	43	59	66
Netherlands	12	27	50	61
Slovenia	5	89	96	94
Italy	6	69	65	99
Hungary	10	92	95	86
Czech R	40	94	95	96
Ireland	24	74	64	37
Denmark	24	63	92	87
Sweden	47	76	78	66
Finland	46	63	92	76
Norway	99	99	99	4

Table A4: Selected indicators of literacy proficiency in the IALS sub-sample used

	Mean score	St. deviation	At least one test Level 1 or 2	All tests Level 1 or 2
UK	278.208	61.904	0.527	0.351
Ireland	263.982	59.974	0.616	0.424
Sweden	310.463	48.098	0.318	0.145
Norway (Bokmal)	297.299	42.804	0.388	0.192
Denmark	295.286	39.309	0.471	0.180
Finland	288.952	47.135	0.454	0.262
Netherlands	291.061	43.103	0.445	0.228
Belgium (Flanders)	284.011	50.557	0.488	0.276
Germany	290.105	42.138	0.514	0.219
Italy	252.067	55.690	0.727	0.516
Czech R	287.789	45.732	0.564	0.247
Poland	233.002	61.646	0.839	0.628
Hungary	255.969	47.570	0.831	0.444
Slovenia	233.994	60.312	0.831	0.622
Total	272.653	56.765	0.598	0.361

Table A5: Employment to population ratios for persons aged 15-64 in the IALS and the OECD Employment Outlook (same year or nearest), and link between education and employment in the IALS

	Year	OECD [a]	IALS	dE/dY [d]
Sweden [b]	1994	71.5	68.8	0.7
Norway [b]	1998	78.2	75.1	1.9**
Denmark	1998	75.3	68.1	1.5*
Netherlands	1994	63.0	62.1	0.8
Belgium (Flanders)	1996	57.0 [e]	58.2	1.5
	1998	64.1	57.4	1.5
Germany				
Italy	1998	50.8	54.1	1.4
UK	1996	67.0	68.0	3.1***
Ireland	1994	52.3	50.6	4.2***
Finland	1998	64.8	62.4	3.2***
Czech R	1998	67.5	66.3	2.7***
Hungary	1998	55.3	56.9	5.3***
Poland	1994	58.3	53.7	3.7***
Slovenia	1998	63.4[c]	61.4	2.5***

a) OECD Employment Outlook 1999, 225. except for Belgium
b) Persons aged 16-64
c) Refers to 2001. Source: Employment and labour market in Central European countries, European Commission, 2003, 56
d) The marginal effect of years in school on the probability of being employed, in per cent, estimated with univariate probit. The stars denote the significance of F-test for the equality of the within-country probit parameter with one estimated for a pooled sample of continental EU countries (first group).
e) Refers to 1997. Source: www.oecd.org, Database on Labour Force Statistics

Note. In the IALS, employment was measured in three different ways: employed at the time of the interview, had labor income in the preceding year, weeks of employment in the preceding year. This paper adopts the first measure. Table A5 gives an overview of employment data in the IALS versus OECD publications. Because of differences in the definition of employment we do not expect the IALS and the OECD figures to be equal. However, in Denmark and Germany the difference exceeds 6 percentage points, a deviation I cannot explain. The table also presents the marginal effect of an additional year in school on the probability of being employed as observed in the IALS, estimated with univariate probit.

Table A6: Selected indicators by educational attainment in the CzR (IALS, males)

Education	Population shares	Recent workplace literacy experience[1]	Literacy test score (standardized for the CzR)[1]
0-10 years	9.8	-4.12	-1.64
11 years	17.4	-3.29	-1.19
12 years, probably V	13.5	-3.29	-0.95
12-14 years, probably S	38.3	-2.37	-0.65
>14 years	21.0	Ref.	Ref.
1) Results from regressions with potential experience and education level dummies on the RHS. The number of reading/writing tasks in the current or last job and the standardized literacy test score were on the LHS, respectively.			

Note. In the CzR the primary school was 9 years long in 1960-78 and after 1990. Therefore people born in 1954-64 or after 1975 and completing 12 years in school may have vocational rather than general secondary education. (This may also apply to people born in September-December 1953 or 1974 but there is no information on the month of birth.) Table A7 suggests that those belonging to these cohorts indeed had lower skills than the members of other cohorts completing 12 years. After adjustment the V share is estimated to be about 30 per cent.

Budapest Working Papers on the Labour Market published since 2001

2005

Kertesi G. – Varga J.: Foglalkoztatottság és iskolázottság Magyarországon. BWP 2005/1

Köllő János: A nem foglalkoztatottak összetétele az ezredfordulón. BWP 2005/2

Kertesi G.- Köllő J.: Felsőoktatási expanzió „diplomás munkanélküliség" és a diplomák piaci értéke. BWP 2005/3

Kertesi Gábor: Roma foglalkoztatás az ezredfordulón –a rendszerváltás maradandó sokkja. BWP 2005/4

Kertesi G. – Kézdi G.: A foglalkoztatási válság gyermekei – roma fiatalok középiskolai továbbtanulása az elhúzódó foglalkoztatási válság idején. BWP 2005/5

Zsombor Cs. Gergely: County to county migration and labour market conditions in Hungary between 1994 and 2002. BWP 2005/6

Szilvia Hámori: Comparative Analysis of the Returns to Education in Germany and Hungary (2000). BWP 2005/7

Gábor Kertesi – Gábor Kézdi: Roma Children in the Transformational Recession - Widening Ethnic Schooling Gap and Roma Poverty in Post-Communist Hungary. BWP 2005/8

John Micklewright - Gyula Nagy: Job Search Monitoring and Unemployment Duration in Hungary: Evidence from a Randomised Control Trial BWP 2005/9

J. David Brown – John S. Earle – Álmos Telegdy: Does Privatization Hurt Workers? Lessons in Comprehensive Manufacturing Firm Panel Data In Hungary Romania, Russia and Ukraine. BWP 2005/10

2006

Köllő János: A napi ingázás feltételei és a helyi munkanélküliség Magyarországon. Újabb számítások és számpéldák. BWP 2006/1

J. David Brown - John S. Earle - Vladimir Gimpelson - Rostislav Kapeliushnikov - Hartmut Lehmann - Álmos Telegdy - Irina Vantu - Ruxandra Visan - Alexandru Voicu: Nonstandard Forms and Measures of Employment and Unemployment in Transition: A Comparative Study of Estonia, Romania, and Russia. BWP 2006/2

Balla Katalin – Köllő János – Simonovits András: Transzformációs sokk heterogén munkaerő-piacon. BWP 2006/3

Júlia Varga: Why to Get a 2nd Diploma? Is it Life-Long Learning or the Outcome of State Intervention in Educational Choices?. BWP 2006/4

Gábor Kertesi – Gábor Kézdi: Expected Long-Term Budgetary Benefits to Roma Education in Hungary. BWP 2006/5

Kertesi Gábor – Kézdi Gábor: A hátrányos helyzetű és roma fiatalok eljuttatása az érettségihez. Egy különösen nagy hosszú távú költségvetési nyereséget biztosító befektetés. BWP 2006/6

Budapest Working Papers on Labour Market is jointly published by the Labour Market Research Department, Institute of Economics of the Hungarian Academy of Sciences and the Department of Human Resources, Budapest Corvinus University

Papers can be downloaded from the homepage of Institute of Economics: http://www.econ.core.hu

Publisher:
Institute of Economics, Hungarian Academy of Sciences