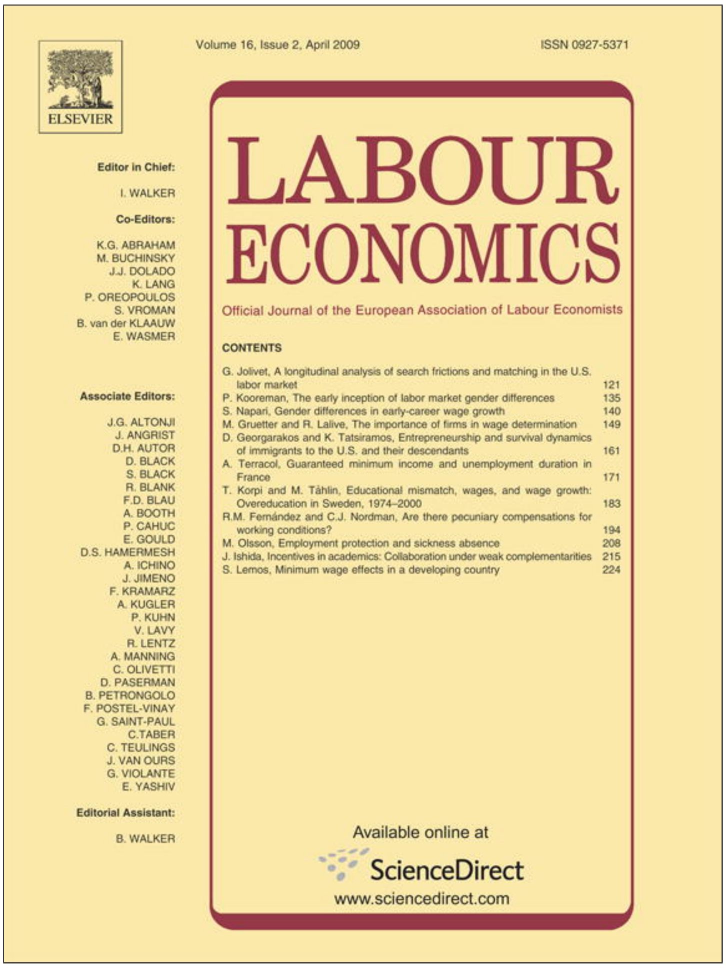

Volume 16, Issue 2, April 2009

ISSN 0927-5371

ELSEVIER

LABOUR ECONOMICS

Official Journal of the European Association of Labour Economists

CONTENTS

Available online at
ScienceDirect
www.sciencedirect.com

Labour Economics 16 (2009) 183–193

Contents lists available at ScienceDirect

Labour Economics

journal homepage: www.elsevier.com/locate/econbase

Educational mismatch, wages, and wage growth: Overeducation in Sweden, 1974–2000

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ARTICLE INFO

Article history:
Received 15 March 2007
Received in revised form 15 February 2008
Accepted 18 August 2008
Available online 29 August 2008

JEL codes:
J24
J31
J62

Keywords:
Educational mismatch
Overeducation
Wages

ABSTRACT

We examine the impact of educational mismatch on wages and wage growth in Sweden. The empirical analyses, based on cross-sectional and panel data from the Level of living surveys 1974–2000, are guided by two main hypotheses: (a) that educational mismatch reflects human capital compensation rather than real mismatch, and (b) that educational mismatch is real but dissolves with time spent in the labour market, so that its impact on wages tends toward zero over a typical worker's career. Our findings do not support these hypotheses. First, significant differences in contemporaneous economic returns to education across match categories remain even after variations in ability are taken into account. Second, we find no evidence that the rate of wage growth is higher among overeducated workers than others. Our conclusion is that the overeducated are penalized early on by an inferior rate of return to schooling from which they do not recover.

1. Introduction

It is well-known that the average level of education has risen significantly in advanced industrial nations in recent decades (Barro and Lee, 2001; Bassanini and Scarpetta, 2001). During the same period, the average skill level of jobs has risen as well (Acemoglu, 2002; Green, 2006). An important issue, *inter alia* for the evolution of economic skill premia, is how these two trends are related to each other. There are two main strands in the literature. The first is the upgrading view, i.e., that skill demand is increasing at a higher rate than skill supply (education). The supposed excess demand for skills is widely used as an explanation for the increase in wage dispersion across skill or education categories that has been observed in several (but not all) countries (Acemoglu, 2003). The main rationale behind such a growth in demand is skill-biased technological change (SBTC), i.e., changes in production processes and work organization favoring employment of high-skill workers. In addition to SBTC, globalization – in particular increased international trade – is viewed as a cause of skill bias in the evolution of labour demand in OECD countries.[1]

The second strand in the literature is the overeducation perspective. The sense that educational expansion is outstripping the demand for skills in the labour market dates back at least as far as the 1940s (Harris, 1949). In the wake of the rapid growth of student enrollment at colleges and universities in the 1960s, this impression became a wide-spread view (Berg, 1970; Freeman, 1976). An empirical literature on overeducation and earnings started with Duncan and Hoffman (1981) and has since become substantial. There is by now a large body of international evidence on the incidence and wage effects of what has been called educational mismatch, overeducation, or overschooling. A fair amount of educational mismatch appears to exist in Western labour markets. Between 20 and 50% of all workers seem to have more schooling than their job requires, with American rates tending to be higher than European. Trends in overeducation are not well established. Available evidence indicates that overeducation has increased in Europe in recent decades, but not in the United States.[2] This pattern might partly explain why research on overeducation is active and growing in Europe (see, e.g., Borghans and de Grip, 2000; Büchel et al., 2003) but is currently not a major topic in American education and labour studies.

How to reconcile the seemingly contradictory upgrading and over-education perspectives? We consider two hypotheses that solve the contradiction by explaining away an observed oversupply of skills. According to the first, all overeducation is apparent rather than real. As pointed out by Duncan and Hoffman (1981), the estimation of a gap

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[1] Feenstra and Hanson (2003) and Acemoglu (2002) review the trade and technology literatures, respectively. Autor et al. (2003) document a link between the expansion of information technology and a rise in skill demand in the U.S. Dissenting accounts regarding the technology-inequality link include Bernstein and Mishel (2001) and Card and DiNardo (2002).

[2] On the upward trend in Britain, see, e.g., Green (2006:40f); recent analyses of the U.S. case are Gottschalk and Hansen (2003) and Handel (2003).

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doi:10.1016/j.labeco.2008.08.004

between the skill requirements of jobs and the skills of workers is based on the "idea of relatively inflexible, technologically determined educational job requirements" (1981:76), as in assignment models (Sattinger, 1993). This runs counter to the traditional analysis of firms' production decisions, in which labour processes are assumed to be continuously adapted to available input factors, including human capital. With such flexibility on the part of firms, mismatch between job tasks and employee skills seems unlikely. Implicit in this traditional standpoint, consistent with a standard human capital model, is thus the view that 'overeducation' is largely due to measurement error. Naively observed educational mismatch, then, would simply reflect unobserved heterogeneity across individuals: apparently overeducated workers are in reality less able than others on other dimensions (while apparently undereducated workers are more able than others). When all relevant ability differences are taken into account, wage returns to education are expected to be independent of the skill requirements of the job. We call this the human capital compensation hypothesis.

An alternative line of argument is that the recorded skill gap may be real, but only in the short run. If the upward shift in overeducation rates is confined to workers in early career phases, it need not be incompatible with increasingly steep skill-wage gradients in the labour market as a whole. Human capital investment strategies as well as labour market search may generate initial mismatch that will, however, typically dissolve through career processes and job mobility. Higher rates of wage growth are therefore expected to compensate for initial wage losses. We call this the career mobility hypothesis.

In what follows, these two explanations of the apparent contradiction between the upgrading and overeducation perspectives are assessed empirically by examining the impact of educational mismatch – in particular overeducation – on wage levels and wage growth. We start from the general finding in the literature that contemporaneous economic returns to education are significantly larger for years of schooling up to the required level of the job than for education above that level (Rubb, 2003). The two hypotheses as to why this difference appears are then considered. To examine the human capital compensation hypothesis we apply explicit measures of some otherwise unobserved skill dimensions as well as fixed effects and instrumental variable models to account for unobserved heterogeneity. The career mobility hypothesis is tested by including measures of the training and career dimensions of jobs as well as by analyzing the connection between educational mismatch and wage growth.

We use cross-sectional and panel data from four waves of the Swedish Level of living surveys, 1974 to 2000. Sweden is an interesting test case displaying many of the characteristics pertinent to the upgrading vs. overeducation debate. Educational expansion since the 1970s has resulted in a significant rise in the average level of education, and the skill requirements of jobs also increased in the same period. The trend in skill supply (education) has been stronger than the growth in skill demand (job requirements), with rising overeducation as a consequence (le Grand et al., 2001, 2004; Åberg, 2003). In 2000, about one third (35%) of all workers had an amount of schooling at least two years in excess of their job requirements. This rate has more than doubled since the 1970s. With regard to wage inequality, a trend toward decreasing wage differences by education was reversed in the 1980s and Sweden has since experienced a significant increase in inequality (Katz and Autor, 1999; le Grand et al., 2001; Kenworthy and Pontusson, 2005).

Our empirical findings indicate that overeducation has substantial real wage effects, both in the short and long run. First, significant differences in contemporaneous economic returns to education across match categories remain even after variations in ability are taken into account. Second, we find no evidence that the rate of wage growth is higher for overeducated workers than for others. We conclude that surplus education carries an early wage penalty in the form of an inferior rate of return to schooling from which the overeducated do not recover. The simultaneous occurrence of a rise in overeducation – with real and persistent wage effects – and an increase in wage dispersion across education categories indicates that market forces alone cannot explain the evolution of economic inequality. Institutional factors must have mattered as well.

The remainder of the paper is organized as follows. Next section contains an overview of previous research on the impact of educational mismatch on wages and discusses some implications and limitations of past findings. Following a description of the data and variables we use in the analysis, we test the two hypotheses outlined above, devoting one empirical section to each of them. We conclude with some reflections on our findings.

2. Educational mismatch and wages: the ORU model

Duncan and Hoffman (1981) decompose attained education into three parts as expressed by the equation

$$\mathrm{AE} = \mathrm{RE} + \mathrm{OE} - \mathrm{UE} \tag{1}$$

where AE denotes attained education, RE is the required amount of education in the job that the worker holds, OE is the amount of education attained by the worker that is in excess of what the current job requires, and UE is the amount of education required by the job that is in excess of what the worker has attained. Hence, OE is zero for correctly matched and undereducated workers, while UE is zero for correctly matched and overeducated workers. The equation thus reduces to AE = RE for the correctly matched, to AE = RE + OE for the overeducated, and to AE = RE − UE for the undereducated.

This decomposition into over, required, and undereducation, known as the ORU model, has two attractive traits. First, conceptually, it combines the information on attained and required education while fully retaining the continuous character of both dimensions. This allows an assessment of separate payoffs to years of attained education dependent on the nature of the job match as revealed by earnings (or other rewards) regressions. Second, empirically, the main pattern of results from this model has turned out to be remarkably robust across both time and countries (Rubb, 2003).

The three schooling components defined above are then inserted into a standard Mincer wage equation, as in

$$W_{ti} = \beta_1 \mathrm{RE}_{ti} + \beta_2 \mathrm{OE}_{ti} + \beta_3 \mathrm{UE}_{ti} + \gamma X_{ti} + \varepsilon_{1i} \tag{2}$$

where X is a vector of independent variables including a constant, γ is a corresponding vector of coefficients, and ε_1 an error term. For correctly matched workers, β_1 indicates the total schooling return. For mismatched workers, the effects β_2 and β_3 interpreted in conjunction with β_1 yield estimates of the total impact of their education. The total return to schooling among overeducated workers is thus β_1 for the years of schooling corresponding to the job requirements together with β_2 for the additional years. Among undereducated workers the total return to schooling is given by β_1 again indicating the return to the years of schooling corresponding to the job requirements but less β_3 for the missing years of schooling.

The following results from cross-sectional wage regressions have been found in virtually all published studies, regardless of time and place (Rubb, 2003a): β_1 and $\beta_2 > 0$ while $\beta_3 < 0$ and $|\beta_3| < \beta_1 > \beta_2$. Put differently, overeducated workers earn more than correctly matched workers in the same kind of jobs ($\beta_2 > 0$), but less than correctly matched workers with the same amount of education ($\beta_1 > \beta_2$). The converse pattern holds for undereducated workers: they earn less than correctly matched workers in the same kind of jobs ($\beta_3 < 0$), but more than correctly matched workers with the same amount of education ($\beta_1 + \beta_3 > 0$).

These consistent empirical results would seem to clearly contradict structural accounts of reward attainment in the labour market, such as the job competition model in Thurow (1975), as well as standard human capital models. The most straightforward interpretation from a standard human capital viewpoint is to emphasize that formal schooling is an obviously incomplete measure of individual productive capacity. Apparently overeducated workers might in reality use their "surplus" schooling to compensate for deficient human capital in other respects. Therefore, the "over-educated" are less productive than others with the same amount of education and so receive a smaller payoff to each year of schooling than the correctly matched workers do. The converse argument would apply to "undereducated" workers.

Allen and van der Velden (2001) find empirical support on Dutch data for this hypothesis by showing that wages are less strongly tied to survey respondents' subjective assessment of educational mismatch than to external classifications based on educational attainment and the skill requirements of jobs. Their subjective indicator of over-education is the degree of disagreement with the statement "My current job offers me sufficient scope to use my knowledge and skills" (2001:438). While useful for many purposes, it is less than obvious that indicators of this kind are relevant to our concerns. One problem is that what counts as 'sufficient' for the respondent may look very different from the employer's point of view, and it is reasonable to suppose that the latter is more important than the former in the determination of wages. Bauer (2002) tests the human capital compensation hypothesis by estimating fixed effects models on German data. His results indicate that unobserved worker hetero-geneity explains a large part of the difference across match categories in wage returns to education, especially for women. However, the measures of required education are based on the educational level of individual job incumbents rather than on the educational require-ments of jobs, and therefore conflate supply and demand side indicators of skill.

Along the same lines as in the human capital compensation hypothesis, but from the viewpoint of jobs rather than workers, educational requirements are obviously not a perfect measure of the skill level of jobs. Hence, due to unobserved job heterogeneity the true skill requirements level may be underestimated for "overeducated" workers and overestimated for "undereducated" workers, so that the mismatch is again overestimated which attenuates the OE and UE coefficients.

The original formulation by Duncan and Hoffman (1981) of the mismatch model contained no reference to the idea of human capital compensation. Rather, the model was proposed in an *ad hoc* fashion without explicit theoretical underpinning (see Hartog, 2000). A rationale was, however, later provided by Sicherman and Galor (1990; see also Sicherman, 1991) with a model in which over-education is seen as part of a human capital investment strategy. In a classical formulation of human capital theory, Mincer (1974) models investment in education and on-the-job training, and also discusses differences in the return to education based on differential invest-ment in on-the-job training. In the career mobility model of Sicherman and Galor, holding a job for which one is overeducated is seen as part of a long-term strategy in which the currently reduced rate of return is compensated by future wage growth. The apparently overqualified may, for example, hold jobs offering extensive on-the-job training or superior promotion prospects. This alternative to unobserved worker characteristics could thus be said to focus on unobserved job characteristics. Based on the career mobility perspective, overeducated workers might be seen as being in an early phase of an upwardly oriented job career. Therefore, the payoff to their attained education is underestimated by only considering current rewards. That is, overeducation should be associated with greater than average wage growth, and the gap in returns to schooling relative to correctly matched workers will over time decrease and eventually go to zero.

The evidence regarding the career mobility hypothesis is mixed. Sicherman and Galor (1990) and Sicherman (1991) find empirical support for catch-up over time among the initially overschooled in relation to better matched workers, in that the overeducated are more likely to display upward occupational mobility.[3] Some subsequent longitudinal studies have reached the same conclusion (e.g., Robst, 1995; Hersch, 1991).

As shown by Büchel and Mertens (2004), however, the findings from these longitudinal studies (all on U.S. data) are not convincing, due to specification problems of the empirical models. In essence, the difficulty lies in distinguishing between the impact of mismatch and the effects of other attributes of the starting position on future reward attainment. In models of occupational mobility, which is the focus of Sicherman (1991) and several followers, one should control for the reward level of the current job when examining reward changes, which these studies did not. This omission leads to mismatch effects being confounded with the impact of vertical scale limits (floor and ceiling) and of regression to the mean. Specifically, Sicherman (1991, Table 3) controls for schooling when estimating the effect of overeducation (as a dummy variable). This means that the over-education indicator will reflect a low job level rather than mismatch (although mismatch may also be involved), and it is well-established from a large amount of previous research (see, e.g., the overview in Rosenfeld, 1992) that the starting job level is inversely related to the direction of subsequent job shifts for purely technical reasons.

Büchel and Mertens (2004) first replicate the U.S. findings on German data, and then show how the positive impact of over-education on upward occupational mobility disappears when starting occupation is controlled. They then proceed to examining wage growth as a further test of the career mobility argument. We agree that wage growth is a better measure of reward change than occupational mobility. However, we are not convinced that the empirical wage growth model used by Büchel and Mertens is properly specified. They conclude that overeducated workers have a signifi-cantly lower wage growth than correctly matched employees, who in turn have a lower growth rate than undereducated workers. Surprisingly enough, given the authors' arguments, this result seems driven by insufficient attention to starting position attributes. The problem is, once again, that schooling is included in the empirical model, so that the overeducation indicator (just like in Sicherman, 1991) reflects low occupational rank rather than mismatch. And although workers in low occupational positions have relatively high rates of upward job mobility (which makes Sicherman's result hard to interpret and potentially spurious), their average wage growth rates are typically weaker than others (see, e.g., Tåhlin, 2007) which makes Büchel and Mertens' finding questionable.

In sum, previous attempts at testing the hypotheses of human capital compensation and career mobility have produced interesting but inconclusive results. We now turn to empirically examining these issues on the basis of Swedish data (described in the next section). We first consider the static case of contemporaneous wages, and in this connection evaluate the hypothesis of human capital compensation. In a second step, we proceed to the dynamic case of wage growth, in the context of which we assess the career mobility hypothesis. In both cases, we start out by providing some descriptive information, and then go on to more explicit tests of the different explanatory perspectives that we seek to evaluate.

3. Data and variables

The data come from the Swedish Level of living surveys (*Levnadsnivåundersökningarna*, LNU) from 1974, 1981, 1991, and 2000. At each occasion, a national probability sample of about 6000 adults (15–75 years 1974 and 1981, 18–75 years 1991 and 2000)

[3] More specifically, Sicherman (1991) examined promotion probabilities and the implied positive relationship between the degree of overeducation and promotion. This would seem to be at best a partial test, since promotions are only one (albeit important) aspect of wage growth. Wages may grow even without a job shift, and promotions may come about for other reasons.

Table 1
Descriptive statistics, Swedish Level of living surveys (LNU), 1974–2000

	Mean	SD	Mean	SD
Year	1974 (N=2828)		1981 (N=3135)	
Wage (SEK/h)	89.85	34.08	89.70	31.60
Ln wage	4.45	0.30	4.45	0.30
Education (years)	1.67	2.40	2.38	2.68
Required education (years)	1.80	2.47	2.03	2.44
Undereducation (years)	0.67	1.42	0.51	1.20
Overeducation (years)	0.53	1.17	0.86	1.49
Female	0.42	0.49	0.47	0.50
Experience (years)	18.90	13.16	18.37	12.63
Verbal ability	1.56	1.30	1.93	1.38
Health problems	5.53	5.14	5.45	4.99
Year	1991 (N=3207)		2000 (N=2877)	
Wage (SEK/h)	92.45	31.10	116.55	54.88
Ln wage	4.48	0.28	4.70	0.31
Education (years)	2.76	2.78	3.81	2.84
Required education (years)	2.39	2.50	3.08	2.63
Undereducation (years)	0.62	1.34	0.59	1.23
Overeducation (years)	0.98	1.54	1.32	1.80
Female	0.50	0.50	0.49	0.50
Experience (years)	18.59	12.02	19.65	12.32
Verbal ability	2.14	1.36	2.14	1.33
Health problems	5.77	4.92	6.02	5.18
Tenure (years)	10.06	9.66		
Formal training	5.51	21.34		
Informal training	13.14	13.68		
Learning new things	3.44	1.14		
Job satisfaction	4.25	0.80		
Advancement prospects			0.20	0.40

Note: Descriptive statistics using listwise deletion among all variables within each year. Note that the sample used in the analyses will differ depending on the combination of variables used in any specific analysis.

residing in Sweden were interviewed (by personal visits) about their living conditions along several dimensions, such as education, working conditions, health, housing, and family life. The non-response rate was 14.8% in 1974, increasing to 23.4% in 2000. The samples have a panel structure, such that all individuals in the sample at t_1 (1974–1991) are included in the sample at t_2 (1981–2000) if still within the targeted age range and residing in Sweden. New members of the sample are drawn at each time-point, entering either through age or immigration. In the analyses in this paper, we use data on 6426 individuals who were employed respondents aged 19–65 on at least one of the four interview occasions: 3112 in 1974, 3285 in 1981, 3326 in 1991, and 3060 in 2000. 2622 of these respondents have participated (and been employed) once, 1944 twice, 1167 three times, and 693 individuals have responded (and been employed) at all four occasions. The analyses pertain to all employees, where employees are defined as those working for somebody other than themselves at least 10 h per week.

Table 1 shows descriptive statistics for all variables used in the empirical analyses below. ED_t (t=1974–2000) is the respondent's attained number of years of full-time education beyond compulsory school. RE_t is the required ED in the worker's current (t) job, according to the respondent's own assessment. RE is thus a crucial variable, on which all of the empirical results depend. The variable is based on two interview questions, phrased: (a) "Is any schooling or vocational training above elementary schooling necessary for your job?" (Yes–No.) (b) "About how many years of education above elementary school are necessary?" (Number of years, ungrouped.) Respondents answering 'No' to (a), or 'Yes' to (a) but less than '1' to (b), are assigned RE=0, while respondents answering 'Yes' to (a) and at least '1' to (b) are assigned RE=x, where x is the response to (b). This information is of high quality, as indicated by both reliability and validity tests. First, in LNU 1991 re-interviews were made with a random subsample of respondents. The outcome of the double interviews showed that indicator (a) had a Cohen's kappa of 0.82 (N=133), while indicator (b) had a Pearson's r of 0.88 (N=76), not much less than ED (r=0.95); see Bygren (1995). Second, RE correlates highly with external judgments of the respondents' occupation (r=0.83 with the SEI code of Statistics Sweden, indicating typical educational requirements by occupation as listed by the Swedish public employment exchange; N=183 occupations). Third, RE is a very strong predictor of wage rates (r=0.51 with ln(wage/h) in LNU 1991), significantly stronger than the corresponding value for ED (r=0.33), and even as strong by itself as a full Mincer model (ED plus experience and its square; R=0.51).

With regard to the other variables used in the analyses, they include a set of standard indicators, viz. hourly wage, total employment experience, tenure with current employer, and sex. Wage is measured as the sum of earnings (as reported in the interview) in the current job during a specific time period (usually one month) divided by the number of hours worked in the same period, adjusted for inflation (all values in 2000 prices). Two other measures are then intended to capture crucial dimensions of what in previous studies has fallen under the rubric unobserved individual heterogeneity; namely health and verbal ability. The former is an index based on 44 indicators of subjective physical and mental health problems, encompassing ailments from headaches to anxiety to cancer, in each case distinguishing between no, minor and severe problems. With no problems coded as zero and the two degrees of severity coded as 1 and 2 respectively, we have an additive index with a range from zero to 88.[4] This index correlates (in LNU 1991) −.15 with log of hourly wages and −.10 with years of education. Verbal ability is also measured through an index, this time constructed from five items indicating various aspects of verbal competence. They include (a) self-rated capacity to write a complaint to a government agency; (b) self-rated frequency of book reading; whether one has (c) spoken to a meeting or (d) written an article; and (e) number of books at home.[5] These have here been coded as zero-one indicators, with the additive index ranging from zero to five. This measure correlates (again in LNU 1991) .36 with the log of hourly wages, .44 with years of education, and has a highly significant positive effect if added to a standard Mincer regression. Thus, although these two measures clearly are less than complete measures of productive capacity, they nevertheless seem to capture important wage relevant characteristics that have not previously been incorporated in analyses using the ORU model.

As for the career based explanations of overeducation, the LNU surveys contain several measures of on-the-job training, advancement prospects and job satisfaction that are useful to illuminate this issue empirically. First, there is a standard indicator of formal employer provided training, measured as the (self-reported) number of days (full-time equivalent) that the worker spent in formal training (education provided by or paid for by the employer) during the last (prior to the interview) twelve months. Second, there is an indicator of informal on-the-job training, measured as the (self-reported) time required from entry into the current job until the worker has learnt to carry out the job tasks "reasonably well". Third, there is a standard job quality indicator, measuring the (self-reported) extent to which the worker's current job requires that s/he keeps learning new things. This variable has five response alternatives: 1=not at all, 2=to a small extent, 3=to some extent, 4=to a large extent, and 5=to a very large extent. Fourth, there is an explicit prospective career mobility indicator, measuring the (self-reported) extent to which the worker's

[4] The questions were "Have you during the preceding twelve months experienced …?"

[5] The actual wording of the questions were; "Could you take it upon yourself to write a letter appealing against a decision made by a public authority?" (yes/no), "Do you read books in your spare time?" (yes, often=1/yes, sometimes and no=0), "Have you ever spoken to a meeting of an association or an organisation?" (yes/no), "Have you ever written a letter to the editor or an article in a journal or a newspaper?" (yes/no), and "Do you have at least 5 running metres of books [at home], not counting works of reference?" (yes/no).

Table 2
Experience, tenure, verbal ability and health by educational match category

	1	2	3	4	5
	Experience	Tenure	Tenure	Verbal ability	Health problems
UE	3.030***	1.684***	0.190	−0.150***	0.084
	(0.161)	(0.137)	(0.118)	(0.018)	(0.077)
RE	−0.611***	−0.054	0.244***	0.282***	−0.217***
	(0.083)	(0.070)	(0.058)	(0.009)	(0.038)
OE	−1.699***	−0.892***	−0.054	0.208***	−0.117*
	(0.128)	(0.109)	(0.091)	(0.014)	(0.060)
R-sq	0.19	0.09	0.40	0.28	0.04
N	3291	3279	3276	3291	3291

OLS regressions. B-coefficients, standard errors in parentheses.
Notes: RE = years of schooling matched by job requirements; UE = years of deficit schooling; OE = years of surplus schooling. All models include a sex dummy. Models 3 to 5 also include experience and its square. Significance levels: ***<=.001, *<=.05. Swedish Level of living survey 1991.

current job contains prospects of advancement, either within the current firm, outside that firm, or both. This measure is coded 1 if both internal and external options are good, and 0 otherwise. Finally, there is an indicator of job satisfaction, namely the extent to which the employee is happy, generally speaking, with his/her current job. The five response alternatives to this question are: 1=very unsatisfied, 2=fairly unsatisfied, 3=neither satisfied nor unsatisfied, 4=fairly satisfied, and 5=very satisfied.[6]

4. Human capital compensation and wages

We now turn to empirically examining the two hypotheses, beginning with the human capital compensation idea, and then – in the next section – looking at career mobility. As a general strategy, we use the ORU model to examine the empirical association between educational mismatch and other human capital related dimensions. We start by relating mismatch to four indicators of human capital. The first two are standard items: labour market experience, indicating general skills, and tenure (time spent with current employer), indicating firm-specific skills. The third and fourth indicators are the measures of health problems and of verbal ability discussed above.

Table 2 shows the model results, all based on cross-sectional data from LNU 1991. We see, firstly, that undereducation is positively related to experience, while overeducation is negatively related to this measure of general skills (model 1). For each year of "deficit" schooling, experience increases by about three and a half years (the UE effect is evaluated relative to the estimate for matched education, RE, not relative to zero). Conversely, each year of "surplus" schooling is associated with on average 1.7 years shorter experience than among otherwise similar workers (the OE effect is evaluated relative to zero). These associations accord well with previous empirical findings from other countries and time-points (e.g., Sicherman, 1991 on US data from the PSID), although they have not previously been estimated on the basis of the ORU specification.

The results on firm-specific skills, indicated by years of tenure, are slightly different. In this case as well, there is some evidence that apparent educational mismatch works as human capital compensation: in model 2, undereducation is associated with relatively long tenure and overeducation with comparatively short tenure. There is no relation, however, between tenure length and matched years of education. But obviously tenure is a subset of total experience, and in order to isolate the specific-skill part it is useful to compare tenure levels at similar levels of experience. When experience is held constant (model 3), matched years of education are significantly related to tenure. The association is positive, indicating that the amount of firm-specific skills is larger among the well-educated, but only to the extent that their job requires the achieved level of education. These results thus indicate that educational mismatch has a compensatory function with respect to general but not to firm-specific skills.

Next, we examine the link between educational mismatch and verbal ability; see model 4. The pattern of results resembles the standard wage results (as described above). There is thus a strong positive association between education and ability up to the schooling level required by the job, and somewhat less strong but still significantly positive above that level, while workers with deficit schooling are significantly less able than correctly matched workers at the same job level but by a smaller amount than expected from their education alone. Finally, in model 5 we show the result with regard to ill health. When one takes into account that we here study disabilities rather than abilities, and therefore have a reversal of signs, these results display clear substantive similarities to the previous ones. There is in other words a strong negative relationship between the level of matched education on the job and health problems, and a similar yet weaker negative association for education above that level. The difference is not however very well established; the *p*-value is .15. In contrast, there is no significant difference between matched and deficit education (excess requirements) in their association with ill health, implying that the health of the undereducated is better than that of workers with similar educational qualifications and equal to that of their matched co-workers.

Overall, according to these results the overeducated workers seem to be more able than correctly matched workers at similar job levels (and with similar amounts of experience), but by a slightly smaller amount than their achieved level of schooling indicates. The undereducated on the other hand appear to be less capable than their matched co-workers, but more able than other workers with similar levels of education. The results would thus seem to support the human capital compensation thesis.

What consequence does this have for the ORU results with respect to wages? Estimates of the standard ORU model, based on the combined samples from the four surveys 1974–2000, are presented in Table 3, model 1. (The OLS model has here been estimated with robust standard errors to take account of the fact that we have multiple observations per respondent.) These Swedish results replicate the

Table 3
Educational match and wages

	1	2	3	4	5
	OLS	OLS w/ add. human capital contr.	Fixed effects	2SLS-IV	Fuller LIML-IV; $a=1$
UE	−0.026***	−0.023***	−0.018***	−0.371	−0.431
	(0.003)	(0.002)	(0.003)	(0.206)	(0.265)
RE	0.068***	0.061***	0.034***	$0.207^{\dagger\dagger\dagger}$	$0.222^{\dagger\dagger\dagger}$
	(0.001)	(0.002)	(0.003)	(0.053)	(0.068)
OE	0.027***	0.023***	0.009***	−0.176	−0.203
	(0.002)	(0.002)	(0.003)	(0.099)	(0.124)
R-sq	0.40	0.41	0.35		
Anderson can. corr., *p*-value				0.07	0.07
Hansen's *J*, *p*-value				0.37	
Cragg-Douglas, *F*-stat				1.40	1.40

Pooled cross-sections, 1974–2000. Standard errors in parentheses.
Notes: Dependent variable ln Wage_t. RE = years of schooling matched by job requirements; UE = years of deficit schooling; OE = years of surplus schooling. No. of respondents 6233, no. of observations 12,124. In addition to the variables shown, all models include sex, experience and experience squared. Model 2 also includes health and verbal ability. Swedish Level of living survey 1974–2000. Sign. levels: ***<=.001. Significance levels with biased standard errors: †††<=.001.

[6] While all of these indicators are based on self-reports, and therefore subject to the various kinds of biases tied to such measures, this should be of less concern here: the career mobility explanation would appear to be based precisely on job characteristics as subjectively assessed by the workers themselves.

results found in other countries. In the OLS version the effect of RE is positive, the effect of OE is also significantly positive although much smaller than the effect of RE, while the effect of UE is significantly negative and thus smaller than the effect of RE, but still significantly above zero (0.067−0.025=0.042).

In model 2 we then show the results when the health and ability variables are included in the analyses. Despite the evidence of human capital compensation presented above, the inclusion of these additional measures of human capital leaves all the educational estimates unaffected. Although these two variables clearly capture ability variations (as shown above) among employees with different educational backgrounds in different jobs, this seems to be of basically no importance when it comes to wage differences across match categories. While the two variables do not cover a complete range of potentially important skills, this result would still seem to question the unobserved heterogeneity story.

4.1. An FE model of ORU

Does this mean that the human capital compensation perspective is misconceived? Before jumping to this conclusion, it seems useful to examine results from other standard approaches to controlling for unobserved differences. One such approach is of course to apply a fixed effects model. When estimating Eq. (2) above, unobserved productivity differences become part of the error term ε_1. Decomposing the error term ε_1, we can write

$$W_{ti} = \beta_1 \mathrm{RE}_{ti} + \beta_2 \mathrm{OE}_{ti} + \beta_3 \mathrm{UE}_{ti} + \gamma X_{ti} + (\rho_{ti} + \varepsilon_{2ti}) \tag{3}$$

with ρ being an indicator of productivity. With a negative correlation between ρ and OE and a positive between ρ and UE the estimates of the educational effects produced by the OLS analyses of the ORU specification (2) would be biased, with the absolute magnitude of both β_2 and β_3 being underestimated.

If ρ is a time invariant person specific factor (i.e. $\rho_{ti}=\rho_{t+1i}$) unbiased estimates could be obtained through the estimation of a standard fixed effects model. One way of specifying the fixed effects model is the first difference model, where

$$\begin{aligned} W_{t+1} - W_t = {} & \beta_1(\mathrm{RE}_{t+1} - \mathrm{RE}_t) + \beta_2(\mathrm{OE}_{t+1} - \mathrm{OE}_t) + \beta_3(\mathrm{UE}_{t+1} - \mathrm{UE}_t) \\ & + \gamma(X_{t+1} - X_t) + (\rho - \rho) + (\varepsilon_{2t+1} - \varepsilon_{2t}) \end{aligned} \tag{4}$$

and the individual index i has been dropped to simplify the notation. The effect of the time invariant factor ρ here cancels out, and the resulting estimates are unbiased.

The results obtained from the fixed effects specification, estimated in its deviations-from-mean form rather than the first difference model of Eq. (4), are presented in model 3.[7] These resemble those obtained from the previous two, although the absolute values of all the estimates decrease noticeably. That the point estimates from the between-person comparison in the OLS model are greater than the estimates from the within-person comparison in the fixed effects model may suggest that unobserved time invariant factors, personal and/or others, are part of the explanation for the differences in the rate of return to the three types of skill match. An alternative interpretation when comparing the estimates such as these is that the changes over time examined in the fixed effects model may involve such a slow process that the time span between interviews may not be enough to capture the full effect of changes in mismatch (Petersen, 2004). Another reason for lower within effects may be measurement error, which generally tends to induce attenuation bias in the estimates (Wooldridge, 2002: 312).[8]

However, there are some more subtle differences between the two models that also may explain the differences in the estimates. While the FE model has the desirable property of controlling for unobserved fixed effects, the focus on changes over time in the independent variables is not unproblematic. Recall that the attractiveness of the ORU model was based in the decomposition of attained education, AE=RE+OE−UE, and the interpretation of the schooling estimates in relation to this decomposition. The move to the FE model involves a shift from between-person comparison to within-persons comparison among individuals who have changed educational level or job. Take the case of a change in RE. This would entail a change in the required qualifications together with a change in educational level of equal magnitude. The typical case would probably be further education matched by a promotion. A change in OE would instead imply an increased educational level without a corresponding increase in required qualifications, i.e. further education without a subsequent promotion. Variation in UE would instead be a promotion without any corresponding addition to educational qualifications.

The differences in the RE estimates could thus also be interpreted such that the promotional payoff to further education in the FE model is less than the return to schooling in terms of starting level evident in the OLS model. The negligible OE estimate in the FE-specification in turn indicates that the return to further education is almost nil if one fails to land a promotion. The reduced UE estimate finally suggests that the UE-deduction is less among those who have proven their mettle on the job.

The relatively small reduction of the UE estimate is in this context nonetheless somewhat surprising. The undereducated workers are often thought of as employees who lack the formal qualifications but who still are highly productive along some unobserved dimension. This unobserved productivity is believed to have allowed them to make a career despite their insufficient qualifications. If this is the case, it would seem reasonable to expect these employees who have demonstrated their ability to receive the same payoff to a promotion as the correctly matched employees (in which case we would have $\beta_3=0$). Yet they still get a lower bonus, so formal qualifications thus still seem to matter somehow.

These considerations also point to some other drawbacks related to the FE model. First, if education level remains unchanged changes in the three match variables involve vertical job mobility. This would produce a change in RE and therefore also in OE or in UE. While vertical mobility such as promotions is one aspect of the attainment process, the specification fails to model wage growth occurring without an occupational shift. Second, in the remaining cases changes in the match variables involve changes in educational level, i.e. further education among adults. Here it can be argued that this is a qualitatively different variable than the cross-sectional measure of educational level, since the latter will tend to focus on youth education. There are two issues involved: the type and the timing of education. Whereas youth education tends to be more general in nature, adult education tends to be vocationally oriented.

[7] The data do not enable us to identify cases where people, such as students, work part-time jobs “on the side”. While the employment definition in the data does exclude short part-time jobs involving few hours per week, some casual employment may still be included. This could induce some “spurious” changes in educational requirements, and to examine this we have conducted two analyses in which we have excluded individuals with less than two or less than five years of employment experience. The results from these analyses were basically identical to the ones shown.

[8] An example of such measurement error is respondents who (implicitly) report that their level of education has decreased between interviews. It is important to note that the attenuation in the estimated effects induced by such errors would tend to reduce significance levels. That the results obtained here are highly significant indicates that measurement error is a relatively benign problem in relation to our conclusions. This is also supported by the fact that analyses of various influence diagnostics (residuals, leverages and combinations thereof) as well as experiments with different recoding and deletion schemes produced substantively similar results.

The years of schooling included in the analyses would thus measure two different types of education, with the impact of one type not necessarily relevant for conclusions regarding the impact of the other. Similarly, the timing of education may be important if age affects the rate of return, for instance through an effect on motivation.

4.2. An IV model of ORU

In the human capital compensation perspective the choice of educational attainment is related to unobserved personal characteristics: some individuals compensate for traits and skills they lack through greater investment in education. In addition to fixed effects, an alternative and common way of handling unobserved heterogeneity involves the use of instrumental variables (IV).

There are numerous examples of IV analyses of the return to education (see, e.g., Harmon et al., 2003 for a recent overview). These share the characteristic of having one endogenous variable measuring education that is instrumented using one or more variables. Here the situation is somewhat different. Recall that the definitions of the various mismatch categories are OE=AE−RE if OE>0, UE=RE−AE if UE>0 and RE=AE if OE=0 and UE=0. Since attained education appears in all three mismatch categories we have to treat all three as endogenous. As we for identification need at least one instrument per endogenous regressor we thus need at least three exogenous instruments.

In order for IV analyses to yield acceptable estimates, these instruments have to fulfil two conditions: relevance and exogeneity. Relevance implies a correlation between the instrument and the endogenous regressor, whereas exogeneity implies no correlation between the instrument and the error term in the main regression. Important is also the predictive power of the instruments, as weak instruments yield less reliable estimates (see Murray, 2006 for a didactic overview of these and other issues in IV estimation).

When considering potential instruments we are limited to the information available in the surveys. We have focused on variables measuring aspects of childhood circumstances, as these seem likely to influence educational careers (i.e., relevant) while at the same time are less likely to be correlated with adult wages given education (i.e., exogenous). Specifically, four instruments are used in the analyses; sibship size, place of residence during childhood, economic problems in family of origin and disruption in family of origin. These instruments are all examples of instruments that have been applied in previously published analyses of the return to education and/or analyses of educational choice.[9]

Results from two different IV analyses are reported in Table 3, models 4 and 5, together with the results from various specification tests. In model 4 we present the 2SLS version of IV. We begin by focusing on the specifications tests reported at the bottom of the table. The Anderson canonical correlations likelihood-ratio test indicates the relevance of the instruments, with the null hypothesis being that they are not. Although not entirely conclusive, the *p*-value indicates that the hypothesis can be rejected. Hansen's *J*-test in turn indicates exogeneity, with the null hypothesis that the instruments are exogenous. Fortunately, this hypothesis is supported. Less satisfactory is the result from the last specification test, the Cragg–Donald *F*-statistic. This is a test of the strength of the instruments, and the null hypothesis is that the instruments are weak. Although there are no critical values available for this type of application, the statistic nevertheless suggests that our instruments are weak.[10]

With weak instruments, both the point estimates and the standard errors obtained using 2SLS may be biased. However, it has been suggested that the bias in the point estimates may be dealt with by estimating the model using Fuller LIML (Andrews and Stock, 2005; Hahn et al., 2004; Hansen et al., 2005). These estimators differ among each other by a positive parameter a, and with $a=1$ the estimator is approximately unbiased. Although such estimates still leave us with the problem of (downward) biased standard errors, the unbiased point estimates will provide grounds for a rough assessment of our hypotheses. Results from this analysis are shown in model 5.

While the results thus come with some notable caveats, they provide little support the human capital compensation hypothesis. The point estimate for OE is negative, so although the biased standard errors prohibit unambiguous conclusions it seems questionable if it is equal to the almost equally strong but positive RE effect. The standard error for the OE estimate is quite large, so the conclusion suggested here is in other words that years of overeducation have no effect on pay. The UE coefficient is numerically large and negative, yet not significantly different from the impact of matched education (RE).[11] On the basis of model 5, it can thus not be ruled out that only the type of job held is of any real importance.

Summarizing the results so far, the differences in results between the FE and the two OLS models may be related to the elimination of unobserved fixed factors in the FE model, but they may also be due to other differences between the models. Whether or not unobserved factors are part of the story, the fact that the larger pattern of results ($|\beta_3|<\beta_1>\beta_2$) remains the same indicates that unobserved heterogeneity in any case is not the essence of the story. This is also the tentative conclusion drawn from the IV analyses. Even if the instruments are less powerful than desired, the analyses provide no indication that the OLS pattern is the result of a compensation mechanism among the overeducated. What then is the main story?

5. Careers and wage growth

In addition to unobserved ability, a second explanation for the pattern of findings from the standard ORU models such as the one examined in Table 3 is specifically related to the difference in size between the returns to required education and to overeducation. As discussed above, Sicherman and Galor (1990) and Sicherman (1991) argued that overeducation may be seen in a career perspective, as part of a human capital investment strategy. To move forward on this issue, we apply the standard ORU model used for cross-sectional estimation

[9] Sibship size simply indicates the number of siblings, disruptions in family is a dummy variable coded one for those who did not grow up with both biological parents present, economic problems is also a dummy coded one for those who state that there were such problems in the family during their childhood, while place of residence during childhood is represented through two dummies with the first coded one if the respondent grew up abroad and the second coded one if the respondent grew up in a town with a population of at least 10,000.

[10] In a recent paper, Stock and Yogo (2002) discuss and present two types of critical values for the Cragg–Donald statistic. The first refers to the bias in the significance levels, whereas the second denotes the bias in the point estimates. Both the critical values related to both types of test are dependent on the number of endogenous variables and on the number of instruments. Unfortunately, the values reported in the paper do not encompass our case (3 endogenous and 5 instrumental variables), yet extrapolating the values in the tables clearly indicate that our instruments are weak.

[11] As discussed by Harmon et al. (2003), in comparison to OLS models IV models frequently yield larger educational effects. This is often explained with reference to local average treatment effects, i.e. the fact that the instruments will identify the effect for a specific group of individuals. If educational effects differ across the population and the instruments identify individuals with a relatively high return the IV model will generate higher educational estimates than the population average produced by the OLS model. Our instruments would seem likely to identify less educated individuals, and estimates of the returns to education using quantile regression have also shown that the rate of return is higher among this group. A point of reference may here be provided by standard Mincer regressions. An OLS model generates a rate of return to education of 5.2, while an IV model using our instruments produces a return of 8.8. Both the first estimate and the increase associated with the IV model are fairly close to the meta-analytic results reported by Harmon et al. (2003).

Table 4
On-the-job training, advancement prospects, and job satisfaction by educational match category

	1 Formal training	2 Informal training	3 Learning opportunity	4 Advancement prospects	5 Job satisfaction
UE	−0.149	−0.328	0.010	0.010	−0.038**
	(0.345)	(0.183)	(0.017)	(0.044)	(0.012)
RE	0.727***	2.059***	0.139***	0.146***	0.048***
	(0.170)	(0.091)	(0.008)	(0.019)	(0.006)
OE	0.007	0.252	−0.009	0.020	−0.041***
	(0.266)	(0.141)	(0.013)	(0.029)	(0.010)
R-sq	0.010	0.286	0.106	0.091	0.031
N	3288	3272	3283	2983	3285

OLS (models 1–3, 5) and logistic (model 4) regressions; standard errors in parentheses. Notes: Formal on-the-job training is number of days during last 12 months spent in employer provided education. Informal OJT is number of months needed after entry to current job before carrying out tasks 'reasonably well'. Opportunity to learn new things on the job is a scale 1–5. Prospects for advancement in current job is an indicator variable with 'good prospects both within and outside current firm' = 1; otherwise = 0. Job satisfaction is a scale 1–5. (See further section Data and variables.) RE = years of schooling matched by job requirements; UE = years of deficit schooling; OE = years of surplus schooling. All models include a sex dummy plus experience and its square. Swedish Level of living survey (LNU) 1991; except model 4: LNU 2000. Significance levels: ***<=.001, **<=.01, *R*-sq, model 4 = Nagelkerke.

of wage effects of educational mismatch to the dynamic case of wage growth. More precisely, we examine wage change in the form of

$$W_{t+1}-W_t = \beta_1 \mathrm{RE}_t + \beta_2 \mathrm{OE}_t + \beta_3 \mathrm{UE}_t + \gamma X_t + \varepsilon_{3t} \quad (5)$$

Note that this specification differs from the first difference specification discussed above. Although the left hand side is identical, the right hand side is different in that we would not be examining the effect of changes in OE, RE, and UE. In other words, this model would focus on the impact of being (mis-) matched on all forms of wage growth, both in connection to promotions and otherwise. If overeducation is regarded as an investment strategy this should be a more relevant model as it takes all forms of investment return into consideration. If overeducation involves investment, initial overeducation should in Eq. (5) be associated with greater than average wage growth, $\beta_2>\beta_1$, and the gap in returns to schooling relative to correctly matched workers should over time decrease and eventually go to zero.

A third explanation discussed by Sicherman (1991) is that the mismatch is temporary, not in the planned career sense discussed above but rather as a result of job search with imperfect information. If the employee is aware that the current job match is less than perfect this would imply further search and subsequent job mobility conditional on better offers being located. In this scenario we would also expect that estimates based on Eq. (5) would yield $\beta_2>\beta_1$.

These explanations of mismatch predict that overeducated workers should experience greater than average wage growth. In contrast, Büchel and Mertens (2004) argue that overeducation should be seen as an indicator of underachievement: the overeducated are not able to land a job at their supposed skill level. There is thus no career investment nor is there any mismatch. Rather than expecting the "proven" underachievers to suddenly become over-achievers, they should be expected to remain underachievers. This is another version of the unobserved ability argument above, but now applied to wage growth. The expectations regarding the parameters in the model would in this case be the reverse of the above, i.e., using Eq. (5) we should obtain $\beta_2<\beta_1$.

Before turning to the wage growth models, we provide some cross-sectional evidence on the career mobility perspective. The human capital compensation explanation of the ORU earnings regression results is based on (usually unobserved) *individual* heterogeneity: that the differences in ability (true productive capacity) between skill match categories of workers are smaller than the education based indicators imply. We saw in a previous section that there is something to this explanation, especially as concerns the (apparently) undereducated. The career mobility explanation, by contrast, is based on (usually unobserved) *job* heterogeneity: the career prospects are hypothesized to be better in jobs held by (apparently) overeducated workers than measures based on schooling requirements imply. The training provided on the job may thus add significantly to the human capital of the job incumbent, and the sojourn in a seemingly low-skill job is therefore an investment that the worker undertakes in order to enhance her future career.

Models 1 to 3 of Table 4 show the results of ORU regression models with the on-the-job training indicators as outcomes, while model 4 focuses on advancement prospects. All regressions are based on cross-sectional data from LNU 1991, except the prospective career mobility model which uses data from LNU 2000 (the advancement indicator was not available until that survey). According to all models in the table, the career mobility explanation is unsupported. The entire association between schooling and on-the-job training and advancement operates via the skill level of the job as indicated by its educational requirements.[12] Net of these requirements, there is no significant linkage between the worker's attained level of schooling and the amount of training she receives on the job. In all the four cases considered, the number of matched years of education is positively and strongly (significant by a good margin) related to training opportunities, while the association between training and surplus education is never significantly different from zero. But note that it is the job rather than the match that is crucial: years of deficit education (excess skill requirements) have the same impact on training content as matched years of education (the UE coefficient, which is evaluated against the matched component RE, is not significant). The conclusion of these simple cross-sectional analyses is thus that jobs do not seem to be heterogeneous with respect to training content in the way that the career mobility hypothesis supposes. All the skill variation in on-the-job training and advancement opportunities appears to occur between jobs, with no net variation at all between schooling categories.

The final model in Table 4 estimates the relation between educational mismatch and workers' satisfaction with their job. The logic behind this model is the following: the ability explanation of earnings variation across match categories as well as the career mobility explanation states that, due to (usually unobserved) individual and job heterogeneity, respectively, the degree of mismatch is overestimated in simple ORU models. Given that overeducated (undereducated) workers are in reality less (more) able than others, or that the jobs held by overeducated workers are in reality better (have a larger training content) than their schooling requirements indicate, apparently mismatched workers should not be less satisfied with their jobs than other workers. The mismatch explanation, by contrast, states that the mismatch is real (even if temporary). Therefore, job satisfaction should differ significantly between match categories. The overeducated might in this view be expected to be less satisfied than correctly matched workers. In the case of undereducated workers, however, the prediction is less clear. Having attained a job level above one's educational credentials is in one sense a positive achievement, but having less schooling than the position normally requires may also

[12] In distinction to formal and informal on-the-job training, which are both measured with interval scale time measures, the measure of learning opportunities is an ordinal scale item with five values and the indicator of advancement prospects is a dichotomy (see the Data and variables section above). The advancement model is therefore a logistic regression. The learning model should ideally be estimated by ordinal regression rather than – as in Table 4 – by OLS. However, there are several variants of specifying the ordinal model, all of which are mere approximations to the "true" latent continuous distribution. In practice, none of these variants produces results that deviate significantly from the OLS specification.

Table 5
Educational match and wage changes

	1	2	3	4
	OLS 1974–2000	OLS 1991–2000 w/ job controls	OLS 1991–2000, w/ job controls & exp<=15	OLS, 1974–2000 w/ lagged wage
UE	−0.009**	−0.006	−0.018	−0.020***
	(0.003)	(0.004)	(0.010)	(0.003)
RE	0.009***	0.007**	0.013***	0.042***
	(0.002)	(0.003)	(0.004)	(0.002)
OE	0.012***	0.010**	0.010	0.022***
	(0.003)	(0.004)	(0.006)	(0.003)
R-sq	0.06	0.14	0.12	0.27
N resp	3474	1895	957	3474
N obs	5715	1895	957	5715

Pooled cross-sections, 1974–2000. Standard errors in parentheses.
Notes: Dependent variable ln Wage_{t+1} − ln Wage_t. RE = years of schooling matched by job requirements; UE = years of deficit schooling; OE = years of surplus schooling. In addition to the variables shown, all models include sex, experience and experience squared. Model 2 furthermore includes formal on-the-job training, informal OJT, learning opportunity, and job satisfaction. Swedish Level of living survey 1974–2000. Significance levels: ***<=.001, **<=.01.

be problematic in several ways. The net outcome of these conflicting mechanisms is an empirical matter.

We find that the mismatch explanation is supported. Education is strongly and positively related to job satisfaction up to the schooling level required by the job, but strongly and *negatively* related to satisfaction beyond that point. This pattern means that both undereducated and overeducated workers are significantly less satisfied than matched workers.[13] So in distinction to the case of on-the-job training (models 1 through 4 in Table 4), job satisfaction is not accounted for simply by variation across jobs. Nor is satisfaction explained simply by variation across individuals. Consistent with the mismatch explanation, it is the *interaction* between individual and job characteristics that is crucial.[14] Hence, mismatch to a significant extent seems to be real rather than merely apparent. The degree to which the mismatch is long-term rather than only temporary is, however, another matter.

In Table 5, we move from these descriptive results to estimates based on the dynamic ORU models described in Eq. (5). This is presented in model 1, and according to these estimates both matched years of education (RE) and years of surplus education (OE) pay off significantly in wage growth. While the point estimates indicate higher returns to OE than to RE, the difference between them is not significant (p-level = 0.29). Hence, in this specification, the rate of return to overeducation is identical to that of required (and attained) education. This result does not directly contradict the career and search theories, but provides no clear support either since we still do not observe the *greater* wage growth impact of OE than RE implied by these hypotheses. As in the static models, years of deficit schooling (UE) have a significantly less positive impact on wages than years of matched education; in fact the UE coefficient is not significantly different from zero. But in distinction to the static case, the magnitude of $|\beta_3|$ is not smaller than that of matched education. In sum, the OLS estimation of this model of wage growth implies that $|\beta_3| = \beta_1$, consistent with a conventional human capital perspective. But note that this result *combined with* the finding $|\beta_3| < \beta_1$ from the contemporaneous wage regression in Table 3 means that the static result holds also in the longer run: the overeducated are (on average) penalized early on by an inferior rate of return to schooling from which they (on average) do not recover. In other words, compared to other workers with the same amount of education, their wage growth curve starts below and then runs parallel to the curve of matched workers.

Although only available in the last two surveys, that is for the period 1991 to 2000, the variables examined in Table 4 allow us to account for some of the differences between jobs that are supposed to generate wage growth differentials. These have here been included as control variables in the analysis and the results are shown in model 2. (The advancement prospects indicator is only available in the last survey, and therefore it cannot be included in these analyses of wage growth.) The only substantive difference in relation to those in model 1 is that the effect of undereducation relative to matched education here is insignificant, in part probably due to the reduced sample size. More importantly, the point estimate for overeducation is still not significantly different from that of matched education. There is thus still no clear evidence of any catch-up among the overeducated.

Since the career mobility theory focuses on career investment decisions, it would seem plausible that the return to investment would be greater among the relatively young. They would simply have the greatest opportunity to recuperate any short term losses in the return to education. This idea is examined in model 3, where we analyze wage growth among those with less than 15 years of experience. However, the estimate for overeducation is not greater than that for matched education, and here it is not even significantly greater than zero. Support for the theory of career mobility is again lacking.

For theoretical reasons, however, the specification in Eq. (5) should be extended. Wage careers clearly involve state dependence: the wage rate at t_1 causally affects the wage rate at t_2, primarily due to downward stickiness but also to differential growth rates across starting wage levels. One possibility is for example that the equality of the RE and OE estimates in the OLS model is due to a combination of two different offsetting growth processes. The career and search theories thus postulate greater wage growth among the overeducated. However, the overeducated tend to earn more than correctly matched employees with the same job requirements (see Table 3), and growth rates tend to decrease with starting wage (i.e., a ceiling effect). This would imply a downward pressure on the growth rates of the overeducated, suggesting that the equality just documented may be the outcome of two counteracting processes. In order to reach firmer conclusions, we therefore need to examine the importance of the starting wage level more closely.

Proceeding from Eq. (5) and taking the starting level of wages into account yields

$$W_{t+1} - W_t = \beta_1 \text{RE}_t + \beta_2 \text{OE}_t + \beta_3 \text{UE}_t + \gamma X_t + \beta_4 W_t + \varepsilon_{4t} \quad (6)$$

Estimates of Eq. (6) using OLS are presented in model 4, Table 5. These show a by now very familiar pattern, namely the standard ORU result $|\beta_3| < \beta_1 > \beta_2$. The estimates may be interpreted in relation to the OLS estimates presented in model 1. For example, among the correctly matched workers in model 1 those in jobs requiring higher qualifications could look forward to greater wage growth. We obtain the same result when we compare employees with the same wages, as we do in model 4, although here the difference is even greater. The latter can be thought of as a comparison of young university graduates with older industrial workers; while wages at t_1 may be identical the former can expect greater wage growth. This type of comparison across models is particularly interesting in relation to the OE estimates. Above we discussed the possibility that the equality of

[13] The UE coefficient implies that each year of excess educational requirements has a zero impact on job satisfaction (the difference between the UE and RE coefficients is 0.10, which is not significantly different from zero). Hence, the overall implication is that (a) requirements matched by education are positive for satisfaction, (b) excess requirements (i.e., unmatched by education) do not matter for satisfaction, and (c) excess education (i.e., unmatched by requirements) are negative for satisfaction.

[14] As in the case of learning opportunities, ordinal regression models of satisfaction (which is measured with an ordinal rather than interval scale item) in several different specifications produce results that are identical to (i.e., not substantively different from) the estimates from the reported OLS model.

the RE and OE estimates in model 1 was due to offsetting processes, potentially concealing a greater wage growth among the over-educated. However, as is evident in model 4 this was not the case. Although the OE estimate increases in relation to the results in model 1, wage growth among the overeducated is now significantly lower than among matched workers. There is thus no indication of greater wage growth associated with overeducation.

6. Conclusion

We have examined the impact of educational mismatch on wages in Sweden in the context of static and dynamic versions of the ORU model. The empirical analysis based on cross-sectional and panel data from the Level of living surveys (LNU) 1974–2000 have been guided by two main hypotheses attempting to explain away the incidence or importance of overeducation: (a) that educational mismatch reflects human capital compensation rather than real mismatch, and (b) that educational mismatch is real but dissolves with time spent in the labour market so that its impact on wages tends toward zero over a typical worker's career.

Our findings do not support these two hypotheses. First, while there are some indications that overeducated (undereducated) workers are less (more) able than correctly matched workers, significant differences in contemporaneous economic returns to education across match categories remain even after variations in ability are taken into account. This has here been done through the inclusion of explicit indicators of ability, or by using fixed effects or instrumental variable estimation. Although each of these analyses have their weaknesses, they nevertheless represent improvements over much of what has been done previously. The fact that none of them dramatically alters the classical ORU results strongly suggests that these results are not entirely due to unobserved heterogeneity. Second, there is some evidence that rates of wage growth are not lower for mismatched workers than for others, but we find no evidence that their growth rate is higher. Our main conclusion is thus that the overeducated are (on average) penalized early on by an inferior rate of return to schooling from which they (on average) do not recover.

A possible extension of these analyses would be to distinguish between different sub-categories among the mismatched, each linked to theoretically informed hypotheses as to why mismatch appears and how it affects labour market rewards. The overeducated are most probably a quite heterogeneous group; some workers compensating for weak human capital in other dimensions than schooling, others in the beginning of an upwardly oriented career path, yet others with temporary low-skill jobs between education spells, and some stuck in undesirable positions for which they are genuinely overqualified. Several of our results indicate that the last of these groups (and the list could of course be extended) is fairly large, but we so far have no precise estimates, nor any clear theoretical explanation for why such negative states endure.[15] And apart from examining the relative size of various sub-categories in a cross-section, it is important to assess the changes in their size over time. Such an assessment is crucial to the interpretation and policy implications of the upward trend in overeducation, in Sweden and elsewhere.

How do our findings relate to the apparent contradiction beween the upskilling and overeducation perspectives that we discussed at the outset of the paper? The analyses here do of course indicate that overeducation, and mismatch in general, is a real phenomenon with important economic effects. At least for Sweden, then, our results imply that the rise in skill demand has been sufficiently met by an increase in skill supply; indeed, perhaps more than sufficiently. Although a significant upgrading of the job structure has occurred in most (if not all) OECD countries, shifts in the *balance* between skill demand and supply need not have changed in a way that enhances wage inequality. The fact that the dispersion of wages has increased in several countries (Sweden among them) that simultaneously show rising rates of overeducation indicates that market factors alone cannot explain the evolution of economic inequality. Institutional factors – such as the character of wage bargaining – must be taken into account as well.

[15] Assignment models may be useful in this regard (see, e.g., Sattinger, 1975, 1993; Teulings, 1995), but have so far proved difficult to combine with the ORU-type model specification (see Hartog, 2000: 140f).

Acknowledgments

We have benefited from comments from several individuals and one anonymous reviewer, as well as from seminar participants in Mannheim, New York, Stockholm, and Uppsala. Korpi furthermore gratefully acknowledges financial support from the Sweden-America Foundation, from the Wenner-Gren Foundations, and from the Swedish Council for Working Life and Social Research. Any errors are our own.

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