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Returns to Education and to Experience within the EU: Are there Differences between Wage Earners and the Self-employed?

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Abstract: EU countries, distinguishing between wage earners and the self-employed. These returns are estimated by using a comparable data set coming from the European Community Household Panel during the period 1994-2000. To correct for the ability bias and recover the education coefficients, an Efficient Generalized Instrumental Variable technique is applied. Although the results differ across countries, two common features can be observed. First, the earnings-experience profiles indicate certain traits of competitiveness in the labor markets and, secondly, the returns to education show that signaling plays a relevant role in the earnings of workers.

JEL classification: C13, C23, I21, J31.

Keywords: Returns to education, wage earners, self-employed, panel data, European Union

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1. Introduction

Estimating the returns to education and to experience has been the target of labor economists for decades, with a very significant volume of research having been devoted to appraising the causal effect of schooling on earnings. One of the main centers of interest when estimating these returns has been to study whether differences exist across several demographic sectors. In this line, a rapidly growing literature examines the differences in the return to education, distinguishing between the self-employed and wage earners.[1] Fundamentally, these studies have set out not only to investigate earning differentials between the two employment groups *per se*, but also to test competing views about the relationship between earnings and education, on the basis that these groups face different economic incentives. In this context, the self-employed can be used as a control group to discriminate between the human capital or sorting models of wage determination, provided that signaling or screening functions are much less relevant for the self-employed (Wolpin, 1977, Riley, 1979). Returns to education and to experience for the self-employed have also been estimated against those for wage earners in order to test different theories of the labor market, such as those of agency and risk hypotheses, against the learning and matching models, or against the compensating differentials premises, for example. Thus, as long as the self-employed have less incentives to shirk in the job or to quit from it, they should exhibit flatter earnings-experience profiles, since wage earners obtain higher earnings when getting older (see, f.i., Salop and Salop, 1976).[2]

Similarly, a variety of reasons have propelled research into the phenomenon of self-employment during recent years. First, self-employment can be viewed as a potential route out of unemployment and poverty (see Fairlie and Meyer, 1996). It may also be helpful for the disabled and the victims of worker discrimination (Moore, 1983). All this has led states and regions to promote this activity in order to both increase welfare and decrease the scale of public expenditure assigned to benefits. As a consequence, an increase in the self-employment rates –the ratio between the number of self-employed and the total employed population in a country- was observed until the

[1] See, for example, Lazear and Moore (1984), Borjas and Bronars (1989), Evans and Leighton (1989), Rees and Shah (1986), Fairlie and Meyer (1996, 2003) and Hamilton (2000).
[2] However, as discussed below, some other theories predict the opposite behavior (see Hamilton, 2000).

mid-1980s, although this has slightly decreased since that time and nowadays represents an average of 15% in the OECD countries (see Blanchflower, 2000).

Secondly, some interest has also been shown in investigating the relationship between self-employment and the moment of the business cycle. Thus, whereas a number of authors have found that self-employment is an alternative to joblessness in some countries (see Evans and Leighton, 1989, for the US, and Alba-Ramírez, 1994, for the US compared with Spain), others argue that expansionist phases lead to greater enrolling into self-employment activities (see Blanchflower and Oswald, 1990 and 1998, and Taylor, 1996). In this sense, it is a common belief among politicians that self-employment is a potential way to create jobs, although economists usually challenge such statements (see Blanchflower, 2000, for a discussion).

A third reason for the increasing importance of analyzing self-employment activities has relied on assessing whether or not marked differences exist in self-employment rates among countries. Developing countries usually exhibit figures of around 50% (Christofides and Pashardes, 2002). In the OECD countries, by contrast, the rates are much lower, varying from less than 10% in the US and Continental EU countries to about 20% in the Southern EU countries (Blanchflower, 2000). These differences have been investigated in order to identify the institutional factors that may affect the decision to enter self-employment (such as the farming occupation in the country in question, accessibility to financial resources, immigration policies or tax structures).[3]

A final reason why self-employment is increasingly analyzed is precisely because many salaried workers would like to become self-employed. The studies by Freeman and Blanchflower (1997), Blanchflower and Oswald (1998) and Blanchflower (2000) show that the self-employed of the OECD countries are more satisfied with their job than are paid workers. Not surprisingly, this finding has led to research considering the determinants of why workers should wish to be either self-employed or paid.[4]

Against this general background, in the paper we set out to estimate the returns to education and to experience for the self-employed and wage earners, with our aim

[3] Fuchs (1982), Borjas (1986), Blau (1987), Borjas and Bronars (1989), Evans and Jovanovic (1989), Blanchflower and Oswald (1998) and Schuetze (2000).
[4] Some the studies focused on this topic are Aronson (1991), Alba-Ramírez (1994), Blanchflower and Oswald (1998) and Blanchflower (2000).

being to cast further light on the issues raised above.[5] In doing so, we provide evidence on such returns for a set of EU countries, namely the 15 pre-enlargement EU countries, using panel data information. Studying different countries is helpful in identifying common features that are not considered in a single-country analysis. This paper is an extension of that by García-Mainar and Montuenga-Gómez (2004), in which the analysis was only carried out for Portugal and Spain. Using a homogeneous database (the European Community Household Panel, from 1994 through 2000) and applying an Efficient Generalized Instrumental Variables, it contains an exhaustive analysis of the returns of the experience and education for the wage earners and self-employed in an ample set of EU countries. The main contribution of our paper is to present efficient estimates of the returns to education and to experience for a set of EU countries, rather than analyzing one single country. To this end, information for more than a year provided in panel data form, which allows for controlling unobserved heterogeneity, is used. The results obtained allows us to draw conclusions about the magnitude of such returns, to derive the implications of the estimation method and to cast some light about the labor market behavior in the EU countries.

The rest of the paper is structured as follows. In the next section, we consider the theoretical aspects of the returns to education and to experience, as well as the phenomenon of self-employment. Section 3 is devoted to the empirical model and to a discussion of the estimation procedure. In Section 4 we describe the data and present some information about self-employment rates and education attainments in the countries under study. Section 5 offers the estimates of the rates of return for the countries under study. Here, we also examine the results across countries in the hope of casting some light on the sector-employment differences. Finally, Section 6 closes the paper with a summary of the main results.

2. Theoretical aspects of returns to schooling and to experience in relation to self-employment

A new-born child enjoys an initial endowment of human capital (a conglomerate of intelligence, ability, motivation, characteristics of the social and economic

[5] However, we do not pursue an explanation for the sources of the differences in the returns to education across countries. Recent attempts in this area can be seen in Ashenfelter et al. (1999), Acemoglu (2003) and Denny et al. (2002).

environment, etc.) that can be improved upon by its accumulation both during the schooling period and through on-the-job experience. According to the human capital theory (Becker, 1964, 1967), there exists a positive relationship between the investment in human capital and earnings, in such a way that a greater accumulation of human capital is rewarded in the labor market with higher earnings. A simple economic model relating initial human capital, the optimal choice of schooling and earnings may then be useful, and is now sketched.

An individual is assumed to maximize his/her expected lifetime utility. The individual derives utility both from schooling, through non-monetary benefits, and from working, through lifetime earnings. More investment in education implies a lower number of years obtaining earnings, but very likely these earnings will be higher than those obtained if the individual had left school in previous periods. Assuming an infinite horizon, a constant discount rate (ρ), and that the individual cannot stay in school forever (ρ is sufficiently large), the individual's problem consists of choosing the years of schooling, *Edu*, such that the following expression is maximized:

$$E_0\left\{ \int_0^{Edu} \psi(A,t)\, e^{-\rho t} dt + \int_{Edu}^{\infty} w_t e^{-\rho t} dt \right\}$$

subject to the relationship between earnings and human capital:

$$ln(w_t)=g(A,Edu,Exp_t,X_t) + v_t$$

whenever $t \geq Edu$. $\psi(A,t)$ represents the non-monetary benefits from attending school, which are assumed to depend on initial human capital, A, say ability hereafter, and time, t. w_t is the earnings obtained at period t, which is a function of ability and schooling (both time-invariant), experience, Exp_t, (which is time-varying) and a set of personal and labor characteristics (such as gender, age, occupation, type of contract, etc.) which can be time-constant or time-varying, X_t. It is also assumed that there are no earnings during the schooling period ($t<Edu$), and non-monetary benefits during the working life ($t\geq Edu$). It is further assumed that once an individual leaves school, he/she does not study any more, such that the education level cannot be increased later on. Similarly, once an individual starts to work, he/she accumulates experience at a rate of one per year.

The individual chooses to stay in school until the expected marginal benefit equals the expected marginal costs of one additional year of schooling. Differences in

ability across individuals causes the schooling choices to differ across them. In line with most of the previous work that has studied the returns to education and to experience, and with the aim of keeping things tractable and making comparisons possible, we finally assume a linear relationship between the log of the earnings and the set of regressors.[6] This implies that ability influences only the intercept of log-earnings. That is to say, we apply the widely-used Mincer (1974) wage equation, that can be expressed as:

$$ln\ w_t = \alpha f(A_t) + \beta\, g(Edu_t) + \gamma h(Exp_t) + \delta' X_t + \varepsilon_t \qquad (1)$$

Since ability is usually unobservable to the researcher, this must be included in the error term. However, this ability may be correlated with schooling, such that standard least squares yield biased estimates (Griliches, 1977). This issue will be further discussed in Section 4.

Although specification (1) has been derived on the grounds of human capital theory, competing perspectives may generate similar conclusions. In particular, the sorting model also predicts that higher earnings are positively related with higher educational attainments. However, in this case, greater human capital does not lead to higher productivity (and thus, higher earnings), but that greater human capital is acquired in order to signal for higher productivity (Spence, 1973, Stiglitz, 1975). In other words, firms do not reward productivity in a direct way because this is not observed *a priori*; rather, they infer productivity from education, and students choose an education level to signal their productivity to potential employers. Similarly, firms offer higher wages for the highly educated because it acts as a screening device, as long as education is positively correlated with the unobserved productivity.

As a consequence, estimating equation (1) does not help to discriminate between the human capital and the sorting models. Whilst it may be viewed as a good approach to assessing the effect of schooling on earnings, it is not completely satisfactory in elucidating which view prevails in the process of wage determination (see Weiss, 1995, for a thorough discussion on this matter). However, considering the self-employed as a control group may serve as a device to investigate such a question, since signaling and

[6] This assumption may not be innocuous. Card (1999, 2001) derives a non-linear specification between the log of wages and schooling to consider the possibility of heterogeneous returns to education across individuals, such as differences in the discount rate and in the schooling parameter. He concludes, nevertheless, that the linear approximation still seems to be a good approach for computing the average

screening purposes seem to be unimportant for this group of workers. In this line, the earlier works by Wolpin (1977) and Riley (1979) settled the basis for making comparisons between both groups of workers. The null hypothesis adopted in these papers is that returns to education will be higher in those occupations that exhibit signaling. On the basis that is difficult to reconcile the idea that education for the self-employed could act as a sorting mechanism, returns to schooling for those in paid employment should be higher since they reap the dual effect of education: the productive and the informative functions. By contrast, the self-employed are only remunerated for the productive nature of education and, thus, returns are lower.[7]

However, although the theoretical implications seem quite clear-cut, the empirical evidence reported for a variety of countries shows very different results. Thus, a number of papers, such as those of Chiswick (1977), Soon (1987), Rees and Shah (1986), Taylor (1996), Cohn et al. (1987), Gill (1988), Grubb (1993) and Hamilton (2000) report a similar finding, namely that self-employed earnings are less responsive to human capital variables than wage employed earnings, thereby favoring the sorting hypothesis. Others, for example, Fields and Schultz (1982), for Colombia, and Henderson (1983), for Malaysia, find that the coefficients of the human capital variables between the two samples are similar. Finally, the studies by Borjas and Bronars (1989), Evans and Jovanovic (1989), Evans and Leighton (1989), Fairlie and Meyer (1996), Clain (2000) and Kawaguchi (2003) for the US, Alba-Ramírez (1994) and Alba-Ramírez and Sansegundo (1995), for Spain, and Simpson and Sproule (1998), for Canada, find that self-employed earnings equations have larger schooling coefficients than those corresponding to the wage employed, rejecting the sorting hypothesis.

For its part, distinguishing between self-employed and wage earners' returns may be helpful in providing insights into the features of the theoretical labor market models. Thus, studying the experience profile in earnings may serve to ascertain whether agency issues, learning and matching models or compensating differentials theories, for example, better fit the labor market. A number of studies predict that earnings-experience profiles are flatter for the self-employed. Within the agency or risk theories (see Lazear, 1979, 1981, Lazear and Moore, 1984), employers should pay less

return to education in the US. Belzil and Hansen (2002), by contrast, find that a non-linear relationship based on splines to different levels of education yields a better fit, rejecting the linearity specification.

[7] The same argument is presented in more recent contributions, such as those of Altonji (1995) and Brown and Sessions (1998).

than the marginal productivity to workers when they are young, and more when they grow older, to avoid shirking in the job, contrary to the case of the self-employed, given that these do not have any incentives to shirk. Similarly, asymmetric information models (Salop and Salop, 1976, Guasch and Weiss, 1982) argue that because employers are interested in minimizing quits of more productive workers, they offer tilted-up wage profiles as a screening device, in such a way that only workers with low probabilities of quitting apply for jobs. By contrast, since the self-employed are not willing to quit, they have flatter earnings profiles than those of wage earners.

In the same vein, learning models claim that, due to sector-specific abilities that are unknown for the individual, workers may not match themselves to the appropriate sector. Those who realize they have a poor match quit their jobs, and only those with relatively good matches stay. This situation causes experience profiles to increase over time (Jovanovic, 1979, 1982). Furthermore, since the self-employed habitually invest strongly at the start-up of their businesses, they are not able to move out of their poor match, and therefore their experience profiles are flatter (see, for instance, Dunn and Holt-Eakin, 2000).[8]

The absence of differences in the returns to experience between the two employment sectors would support the competitive approach of the labor market. However, this is not as clear as it may seem. Thus, compensating differentials theories may explain the earnings differences and the distinct performance of returns to education and to experience between both employment sectors.[9] Hamilton (2000) estimates returns to education for the two groups in the US, finding evidence of steeper experience profiles for wage earners, as well as the existence of non-pecuniary rewards that compensate the self-employed for receiving lower earnings. This led him to claim that compensating differentials are at work.

Finally, Kawaguchi (2003) shows that even human capital theory may predict a flatter earnings-experience profile for the self-employed, always provided that their earnings are subject to more variation and that their returns to education are higher than for the case of wage earners (which holds for the US). Opposite results are argued to be found by the investment model, for example, which justifies that the self-employed obtain steeper earnings profiles because physical and human capital investments are not

[8] However, when such costs do not exist, bad-matched self-employed leave their occupation and enroll in paid employment, reverting the performance of the profiles.

shared with an employer (Hashimoto, 1981). Similarly, Rosen (1981) claims that average returns may be distorted by the existence of a few, but very successful, entrepreneurs, "superstars", with the bulk of them staying with low returns or leaving for paid employment, in such a way that earning profiles are steeper for the self-employed.

In summary, undisputed conclusions about the magnitude of the returns to education and to experience for the self employment and for salaried employment have not been achieved. Most of the analyses has focused on investigating only one country, without offering any kind of comparative study. Furthermore, only a limited number of papers have used information for more than one year. Even when they have done so, they have estimated returns by pooling the data, an approach which does not allow them to control for the unobserved characteristics of the individuals, or the movements into or out of self-employment. The aim of this article is precisely to address some of these gaps in the literature by computing returns to education and to experience for a set of EU countries using information provided in panel data form.

3. The empirical model and estimation procedure

This section focuses on the empirical specification of the earnings equation and the methodology used for its estimation. The first sub-section is devoted to arguing which is the most appropriate empirical model for our study, whilst the second describes the reasons that have led us to use the Hausman-Taylor procedure in the estimation.

3.1 Empirical specification

As discussed in Section 2, estimates of the returns to education and to experience for the self-employed and wage earners are habitually obtained from Mincer-type wage regressions. Dating from the mid 20th century, a body of empirical work has investigated these returns across countries on the basis of such a specification.[10] During recent years, one line of specialization that has rapidly become

[9] See, for instance, Kanburn (1982), Evans and Leighton (1989) and Blanchflower and Oswald (1998).
[10] See the set of continuous updating works by Psacharopoulos (1973, 1981, 1985, 1994) and Psacharopoulos and Patrinos (2002). See also the recent cross-country comparative study by Trostel et al. (2002) and the surveys by Harmon et al. (2003) and Heckman et al. (2003).

more popular is to study the performance of returns across different demographic groups. From amongst these, the analysis of returns for the self-employed as against wage earners has gained importance essentially, and as mentioned earlier, in order to discriminate between alternative theories of the labor market. Cross-sectional information has habitually been used in the estimation of the returns to education. Furthermore, an IV approach to take care of endogeneity and ability biases, as well as measurement errors, has progressively substituted for the traditional OLS estimation. This has resulted in estimates of the rates of return well above those obtained from OLS (see Card, 1999, 2001, for recent overviews).

In this paper, the estimated model is an extended version of the Mincerian-baseline equation (1), in which earnings rewarding more education can be seen as the combined effect of human capital accumulation and the effect of being identified as a graduate rather than as a dropout. It takes the following form:

$$ln\ w_{it} = \beta\ Edu_i + \mu_1\ Exp_{it} + \mu_2\ Exp^2_{it}/100 + X'_{it}\ \delta + Z'_i\ \gamma + u_{it}, \qquad (2)$$

where *i* and *t* stand for the *N* individuals and the *T* time periods, respectively. As indicated before, *w* denotes earnings; *Edu* is the education variable (that is considered time-invariant); *Exp* is the experience; *X* is a set of time-varying regressors and *Z* a set of time-invariant regressors. The β coefficient expresses the rate of return to education; μ_1 and μ_2 represent the earnings-experience profile, whereas δ and γ are the set of parameters accompanying the rest the regressors.

Because of the structure of the analysis, a random effects-type model is selected. This model assumes that the error term u_{it} consist of the sum of two components, i.e. $u_{it} = \alpha_i + v_{it}$, where α_i represents the random individual-specific effect that characterizes each worker and is constant through time, and v_{it} is a random disturbance varying through time and individuals. This latter stochastic term is assumed to be uncorrelated with all included variables, whereas some variables (education, experience) may be correlated with unobserved variables contained in α_i such as, for instance, ability. Similarly, it is also assumed that the random disturbance is a sequence of i.i.d. random variables with mean zero and variance σ^2_v; v_{it} and α_i are mutually independent, and that α_i is i.i.d. over the panels with mean zero and variance σ^2_α. Thus, the variance-covariance matrix of the system has the random effects structure that can be represented

as $E(UU') = \sigma^2_\alpha (i_T i_T' \otimes I_N) + \sigma^2_v (I_T \otimes I_N)$, where i_T is a Tx1 vector containing ones and I_N (I_T) is the identity matrix of rank N (T), and U is an NTx1 vector of disturbances.

The dependent variable is the natural log of net earnings, where these are defined as gross earnings less tax, expressed in per hour real terms. The habitual education measure employed in the literature, years of schooling, is discarded because the existence of some errors in the collection of data. Instead, we use the educational attainment by each worker, which presents two clear advantages: first, it does not impose the annual marginal effect of schooling to be the same in each year of education; secondly, the level of education is a more appropriate measure, since multiple education streams characterize European countries, and salary profiles use to be largely linked to the education category attained (see the discussion in Hungeford and Solon, 1987, about the "sheepskin effect"). Thus, the educational attainment, which is considered time-invariant in our sample, represents the last completed type of schooling and is classified into the three levels described earlier: primary, secondary and high.[11].

The earnings-experience profiles are analyzed by considering the number of years that an individual has been working, and its squared value divided by 100 to take care of the decreasing returns. Specifically, experience is measured as the difference between the current age and the age of initiation at work, thereby expressing the actual experience. The rest of independent variables considered in the estimation, and that are represented in equation (2) by X and Z, are the following. First, gender is a dummy variable, with the value 1 for male and 0 for female; second, several dummy variables that indicate the marital status: married, single, divorced, widow or separated; third, a dummy variable that specifies whether the worker has realized some course of occupational training or not; forth, nine dummies referring to the occupation of the worker; fifth, a dummy variable that indicates whether the individual works in the private or in the public sector; sixth, three dummies for seniority: less than two years, between 2 and 10 years and more than 10 years; and, finally, year fixed effects.

3.2 Estimating the earnings equation: the hausman-taylor procedure

[11] In this case the fragment "$\beta\, Edu_i$" in equation (2) would be represented by "$\beta_1\, EduS_i + \beta_2\, EduH_i$". The category of reference is $EduP_i$ which is omitted in the estimation.

As stated earlier, the estimation of the Mincerian earnings function to determine the rate of return to education is not problem-free. The presence of measurement errors and unobserved variables, such as ability, motivation, etc., that may be correlated with schooling, bias OLS estimates. Specifically, it has been shown that measurement error bias downwards the OLS estimates (Griliches, 1977, Angrist and Krueger, 1999) although recent evidence (Card, 2001) only attributes a ten per cent gap, at most, to this source of bias. By contrast, since schooling and the unobserved ability may be positively correlated, omitting measures of ability results in the schooling coefficient being biased upwards (Griliches, 1977). Consequently, some effort must be made to alleviate such an ability bias as much as possible. When a direct indication of ability, such as IQ score tests, or information from twins or siblings, is not available (see Ashenfelter and Krueger, 1994, and Miller et al., 1995), the most appropriate exercise is to select an instrumental variables estimator by way of which schooling is instrumented with variables that are correlated with it, but not with errors. A broad range of instruments have been proposed in the literature. Typical examples are those known as natural experiments (see Rosenzweig and Wolpin, 2000, for a summary) which include: i) school reforms and features of the school system (Harmon and Walker, 1995); ii) the proximity to college in the place of residence (Card, 1995b); iii) other supply-side instruments capturing features of the education system (see Card, 2001 for a survey of the literature); and iv) the season of birth of the individual (Angrist and Krueger, 1991). Other possibilities include family background (Blackburn and Neumark, 1995) and the absolute degree of risk aversion (Brunello, 2002). Excellent surveys about this matter can be found in Card (1995a, 1999, 2001).

When using IV, an habitual finding is that estimates are 20% higher, or even more, than OLS estimates. This is a rather unexpected result, since OLS is already believed to provide upward biased estimates because of the ability bias. Some reasons have been argued to explain such a result. Thus, Ashenfelter et al., (1999) show that researchers tend to report the highest IV estimates (and the most precisely estimated), leading to an upwards publication bias. An alternative explanation is based on the idea that the downward bias in OLS due to measurement errors dominates the upward ability bias (Angrist and Krueger, 1991). However, as indicated by Card (1999), whilst measurement errors can account for, at most, 10 per cent of the gap between OLS and IV estimates, they do not fully explain the total existing gap.[12] In the same line,

[12] Unless these measurement errors would be higher in the groups affected by the treatment when applying IV (see Kane et al., 1999).

differences in family background, or in the average ability in the different groups may generate an upward grouping-data bias (see Angrist and Imbens, 1995).

Bound and Jaeger (1996), consider that IV estimates are biased upwards further than OLS because of the existence of unobserved differences between the characteristics of the treatment and comparison groups implicit in the IV scheme. Here, Card (1999, 2001) claims that returns to education are heterogeneous across individuals and the IV estimates tend to recover the returns to education of the population group most affected by the intervention.[13] Similarly, IV estimates will tend to be biased towards the returns to schooling attainments that are most common in the sample data (see Belzil and Hansen, 2002).

Both the available data structure and the existence of problems associated with the choice of instruments have influenced the procedure applied in this study. On the one hand, the dataset used here, obtained from the European Community Household Panel, ECHP hereafter, is in panel data form, but does not provide information about IQ tests and the presence of twins is not especially accounted for. On the other, although the number of alternative instruments habitually considered in the literature is sufficiently wide, their application to our data is quite complex. Specifically, as regards the choice of family background as an instrument for education, there is information in the ECHP only about the educational level attained by the spouse, but this information is only available for married individuals. Nor is data provided on the proximity to a college. Furthermore, thirteen countries are analyzed, which would make the consideration of institutional shocks as instruments really daunting. With respect to other natural experiments, information about the season of birth is included in the data, but after some exploratory exercises (not shown, but available from the authors upon request), this variable emerged as a weak instrument which would produce biased estimates (see Bound et al., 1995). All this has led us to consider an alternative procedure for estimation, in which the availability of panel data is taken into account, namely the IV-type model proposed by Hausman and Taylor (1981), a choice we will now discuss in more detail.

Our opting for this procedure is motivated by several reasons. As is well known, the availability of panel data allows us to control for individual unobserved heterogeneity, possibly correlated with other included variables, since this may be eliminated by mean or time-differencing, i.e. by applying a fixed effects-type

estimator.[14] Although this within estimator is probably not fully-efficient, it produces consistent estimates. However, when operating in this way, coefficients of the time-constant variables (f.i. the level of education) cannot be estimated, because they disappear when mean or time-differences are constructed. For its part, a pure random effects estimator, the Generalised Leasts Squares (GLS) estimator, produces biased and inconsistent estimates because it assumes that there is no correlation between any of the regressors and the individual effects. In our case, the GLS estimator is not valid because at least one of the regressors, education, is endogenously determined.[15]

One possibility to obtain consistent estimates of the returns to education and to experience would be to find instruments for these variables which are potentially correlated with the individual effects. The choice of the appropriate instruments is, however, not an easy task, since the use of instruments that are weakly correlated with endogenous variables may produce downward biased estimates, even with large samples (see Bound et al., 1995, Staiger and Stock, 1997, Chamberlain and Imbens, 2004), which generates uncertainty in the selection of instruments. Consequently, what we require is a procedure that controls for the endogeneity of education (and possibly other variables), but which is still able to recover the coefficient of time-invariant regressors. Hausman and Taylor (1981) propose a model where some of the regressors may be correlated with the individual effects, as opposed to the random effects model, where no regressor can be correlated with the individual effect, and to the fixed effects model, where all the regressors may be correlated with the individual effects. If, in addition, this procedure does not require instruments excluded in the regression but, by contrast, the instruments used are precisely those included in the wage regression, the Hausman-Taylor estimator is, potentially, the best choice.

This Hausman-Taylor estimator is an instrumental variables estimator that uses both the between and within variation of the strictly exogenous variables as instruments. More specifically, the individual means of the strictly exogenous regressors are used as instruments for the time invariant regressors that are correlated with the individual effects. This procedure is implemented in the following steps. First, equation (2) is estimated by pooled Two Stages Least Squares (2SLS), where the set of variables

[13] IV estimates are hence a better approximation for the returns to education of the affected group rather than for the whole population.
[14] Some examples of studies on earnings and panel data are those of Baltagi and Khanti-Akom (1990), Polachek and Kim (1994), Rosholm and Smith (1996), Kalwij (2000).
[15] As mentioned earlier, the Mincerian human capital earnings function may be interpreted as stemming from the theoretical conceptions by Becker (1964, 1967) where human capital investment is an endogenous choice by the individual (see Willis, 1986, Card, 1999).

mentioned above act as instruments. Secondly, the pooled 2SLS residuals are used to obtain estimates of σ^2_α and σ^2_v, which can then be used to construct the weights for a Feasible Generalized Least Squares estimator. Thirdly, these weights are used to transform (by quasi-time demeaning) all the dependent variables, explanatory variables and instrumental variables. Finally, the transformed regression is again estimated by pooled 2SLS, where the individual means over time of the time-varying regressors and the exogenous time-invariant regressors are the instruments. Under the full set of assumptions mentioned in the previous sub-section, this Hausman and Taylor estimator is the Efficient Generalized Instrumental Variables (EGIV) and coincides with the efficient GMM estimator (see Appendix for more details).

Specifically, in the case under consideration, education is a potentially endogenous, time-invariant regressor, whereas the experience variables may be also endogenous, but time-varying. Since we are interested in the coefficients of these variables, all the exogenous variables (either time-invariant or time-varying), plus the individual means over time of the all time-varying regressors, can be used as instruments to obtain consistent estimates of the returns to education and to experience. Additionally, the variance-covariance structure can be taken into account to obtain more efficient estimators.[16] Consequently, the Hausman-Taylor procedure is a good alternative to pure IV estimation when panel data is available. With the advent of new panel data sets, this method has become increasingly common in studies devoted to the job market in general (f.i. Greenwood et al., 1999, 2003, and Contoyannis and Rice, 2001), as well as to the analysis of the return to education (see Hansen and Wahlberg, 1997, Kalwij, 2000 and García-Mainar and Montuenga-Gómez, 2004). For a recent assessment of the adequacy of the Hausman-Taylor method, see Baltagi et al. (2003).

One final problem in our estimation derives from the fact that the sample used in each country is selected by employing a non-random scheme, since only individuals with observed earnings are considered. This gives rise to sample selection, leading to biased parameter estimates. Furthermore, the Mincer equation we use is estimated for sub-populations with given characteristics, i.e. wage earners as against the self-employed, in such a way that the final sample may not be representative of the whole population. We have chosen not to correct explicitly for these two possible sources of

[16] Amemiya and MaCurdy (1986) and Breusch et al. (1989) proposed alternative variations to this model in order to obtain more efficient estimates. However, evidence in favor of one particular approach has been not consistently adduced (see Baltagi and Khanti-Akom, 1990, and Boumahdi and Thomas, 1992). More recently, Gardner (1998) and Im et al., (1999) have proposed some modifications to the benchmark Hausman-Taylor approach to take care of specific circumstances.

sample selection bias and, in making this choice, we rely on a number of arguments in support.

First, when using cross-section information, there are inconclusive results about the influence of the sample bias in estimating returns to education and to experience. Frequently, the irrelevance of such a bias is observed (see Evans and Leighton, 1989, Alba-Ramírez, 1994, Hamilton, 2000).[17] Secondly, only limited efforts have been made to investigate this problem in a panel data framework, and only recently have estimation procedures been suggested for panel data sample selection models. These correspond to two lines of research: two-step estimators, following the idea of Heckman (1979);[18] and maximum likelihood estimators (Husted et al., 2001). These estimators differ in the prior specification of both the equation of interest and the selection process. Each of these may produce consistent estimates provided that certain assumptions are satisfied. However, since some, or many, of the assumptions are likely to be violated in applied work, uncertainty in the choice of the appropriate estimator clearly arises. Additional drawbacks are the need for exclusion restrictions for the sample selection, the difficulty in finding variables included in the selection process that do not enter the equation of interest, and the fact that these estimators are derived under the assumption of exogenous regressors.[19]

Although some recent attempts have been made to compare different panel data sample selection models, no outstanding procedure has yet been found. Jensen et al., (2002) undertake a Monte Carlo study to show that, whilst maximum likelihood estimators tend to offer the best performance, they are computationally very demanding, so that the two-step estimators (in particular, the one proposed by Kyriazidou, 1997), are preferable. They conclude, however, that more possibilities should be considered. Dustmann and Rochina-Barrachina (2000) compare three two-step estimators in an environment similar to ours, namely the estimation of wage equations. They find that the Kyriazidou estimator is indeed difficult to apply in such a specification, since it imposes a conditional exchangeability assumption that is difficult to meet. Finally, the

[17] This is an habitual finding in countries where self-employment is not prevalent, that is to say, countries where self-employment rates are lower than 20% (see Christofides and Pashardes, 2002). However, some other studies have found a negative influence of the selection bias.

[18] See Wooldridge (1995), Kyriazidou (1997), Vella and Verbeek (1999), Rochina-Barrachina (1999) and Lee (2001).

[19] Dustmann and Rochina-Barrachina (2000) extend three estimators (Wooldridge, 1995, Kyriazidou, 1997, and Rochina-Barrachina, 1999) to cover the possibility of endogenous regressors. However, they point out the difficulty, in some cases, in finding proper instruments excluded both from the equation of interest and from the selection process. Moreover, in our case, we have both types of regressors,

procedures habitually suggested in the literature to correct for the bias arising from attrition (see Wooldridge, 2002, for a summary), apply first differences to remove unobserved heterogeneity. This hinders the estimation of the coefficients of time-invariant regressors, which constitute the main target of our study and, thus we are not in a position to apply such procedures.

Given all these shortcomings, the researcher is faced with two choices: either to apply some, or all, of these estimators and to present estimates for all the possible cases and compare; or simply to ignore this source of bias, in the hope that either the selection process is time constant or that the biases affect all the countries in the study a similar way, such that the final estimates are still comparable between them. Opting for the first possibility supposes a disproportionate burden of work, generating a huge amount of information that, nevertheless, does not assure success. Furthermore, we should bear in mind that this is not the main aim of our research, which is to carry out a comparative study across a set of EU countries. Consequently, we have decided to choose the second possibility and not to apply any of the proposed estimators. As a way to alleviate the sample selection bias as much as possible, we have decided instead to introduce the largest achievable set of regressors which, apart from improving the robustness of all the coefficients and the overall fit, helps to control for several factors that might influence the workers' decisions.

As regards the source of sample bias arising from the split of the sample between wage earners and self-employed, equation (2) has been estimated, first, by separating both samples and, secondly, by pooling the self-employed and wage earners into one single sample and including dummy variables to differentiate both types of workers. Nevertheless, the estimates obtained when separating both samples are more reliable than the pooled ones, since pooling estimation imposes common coefficients for all the regressors, except education, which is not necessarily true in reality.

4. The data and descriptive statistics

The data used in this study come from the ECHP for the period 1994-2000. This is the only database that provides individual information that is comparable for all the EU countries, since the design and organization of the survey is coordinated by EUROSTAT. Individual or micro data is preferred to more aggregate data both because

exogenous and endogenous, and thus finding accurate instruments may become a really intractable

they provide more flexibility in creating sample restrictions, and because they allow us to directly control for individual-level characteristics in our regression.

At the time of the interview, individuals are requested to indicate whether they are working in a job for at least 15 hours a week. If so, workers identify themselves as either self-employed or employee when asked about their main labor market activity. As a consequence, the job status of a particular worker may vary from year to year. In the sample, we have selected those workers, either self-employed or wage earners, that have provided information for all variables under consideration. These variables include personal and labor characteristics such as gender, marital status, schooling, experience, earnings, seniority, occupation, whether the individual works in the private or in the public sector, the number of hours worked per week, and if the worker has taken some training course during the last year.

The number corresponding to wage earners in the sample ranges from about 3,500 in Belgium to more than 8,000 in Germany and Spain. For the self-employed, the figures are considerably lower, varying between less than 1,000 in Austria, Belgium, Denmark, France, Germany and The Netherlands to almost 3,000 in Greece. The following tables illustrate some characteristics of self-employment for the pre-enlargement EU-15 in the last two decades. Table 1 shows the official self-employment rates at three moments in time during the last 20 years. Up to the 1980s, self-employment displayed a marked rate of growth. During the sample period, however, they have exhibited a global trend of maintenance or slight decrease.[20] Within this general behavior, significant differences across countries can be appreciated. Those countries where rates have reduced more sharply (Portugal, Spain, Ireland and Greece), show the highest values (about 20%, with Greece being the highest).[21] In the set of countries with high self-employment rates, Italy is the exception, since its rate has kept more or less unaltered at 24%. By contrast, countries that in the mid-1980s exhibited the lowest values of the self-employment rate (Germany, Denmark, Luxembourg, The Netherlands, Austria and Sweden) have tended to maintain these values, with the

problem. Also see Vella and Verbeek (1999).

[20] Most of this downward trending behavior has been due to the decreasing importance of farming activities. When analyzing non-agricultural self-employment rates, a clear upward trend over time can be observed in most of the countries (see Blanchflower, 2000).

[21] These countries are those which have exhibited the highest unemployment rates within the EU. Some authors (Meager, 1992, Alba-Ramírez, 1994) have suggested that in these types of countries, self-employment is a plausible alternative to joblessness. However, Blanchflower (2000, p. 488) finds no clear-cut relationship between self-employment and unemployment, especially in these four countries. Besides, these countries are also those which started from the levels of a less developed economy and where farming activities were more relevant.

countries that were situated in the medium distribution showing a slowly decreasing path in their rates (Belgium, Finland, UK and France). Overall, we can note a general process of timid convergence towards quite low values of self-employment rates, according to which those of the "peripheral" countries (characterized by the highest rates) tend to decrease over time, whereas the "core" countries maintain their levels around 10%.

(Table 1 about here)

Table 2 presents information, for thirteen out of the fifteen sample countries[22] and distinguishing between the self-employed and wage earners, about average earnings per hour, the years of experience and the educational attainments of workers. Bearing in mind that self-employed earnings are habitually believed to be underreported,[23] no significant differences between their earnings and those of the wage earners seem to be observed. Globally, wage earners appear to earn a little more than the self-employed in the Southern EU countries (Portugal, Spain, Italy and Greece) and in Finland, whereas the opposite occurs in the core countries and in the UK. Note that dispersion in earnings is higher for the self-employed, reflecting the great heterogeneity in these types of activities, from low-ability jobs (retailers and basic services) to those of professionals, such as doctors or lawyers. Obviously, people living in the above-average per capita income countries obtain higher earnings than those resident in below EU average states.

(Table 2 about here)

The years of experience are higher, in average terms, in the self-employed sector than in that of wage earners, especially in the Southern EU countries and Ireland. Table 2 also shows the percentage of workers that have achieved a certain level of education for both samples. The levels under consideration are labeled as primary, secondary and high, where primary includes elementary and below elementary school, secondary includes vocational and middle school, and high includes university studies (either in short or in long cycles). Some general results emerge. In the Southern EU countries (Greece, Italy, Portugal and Spain) and Ireland, more than one half of the self-employed have only attained primary level. At the other extreme, in the UK and in Belgium most of the self-employed have attained the high level, whilst in the remaining countries the highest proportion of the self-employed have achieved a secondary level. That is to say,

[22] Luxembourg and Sweden are excluded from the analysis because of the lack of adequate data for some variables.
[23] Arguments such as the tax structure and the lack of some types of compensation for the self-employed are usually advocated (see Hamilton, 2000).

in those countries which enjoy the higher self-employment rates, the self-employed themselves are less educated, given that they have basically achieved only primary-level studies. By contrast, countries with low self-employment rates exhibit a higher proportion of workers, either wage earners or the self-employed, who have obtained at least a secondary level of education.[24] In summary, most of the self-employed in the peripheral countries have only primary studies, with a quite low proportion of the highly educated. On the other hand, the structure of educational attainment among wage earners is more equally distributed. In the case of the core countries, secondary levels of education prevail for both categories of employment.

5. Estimation results

This section presents the empirical evidence adduced from our study. The estimated results are first shown using alternative specifications and are then assessed in the light of the aspects mentioned in Section 2, with our aim being to provide some insights into the functioning of the European labor markets. We present two sets of results. The first corresponds to the pure GLS estimation, where no correlation between the regressors and individual effects is permitted; that is to say, the education and experience variables are taken as exogenous. Secondly, we do consider the possibility of endogeneity in such variables.

The results of the GLS estimations for all the sample countries are shown in Table 3. Self-employed and salaried earnings are estimated separately. Only the estimated coefficients of the variables of interest are displayed. Most of the coefficients are significant at the 1% level and have the right signs. The use of qualifications as independent variable allows to test the linearity hypothesis. Growing returns are found as we move up the qualification ladder, especially from secondary to higher education, which supports a convex configuration of earnings on the returns to education.[25] Although the results differ across countries, two common ideas emerge. First, when returns to secondary education are found for the case of the self-employed, these are usually higher than for the wage earners. Secondly, it appears that in most of the sample countries, returns to higher education are usually greater for the self-employed.

[24] The case of the UK is especially appealing since workers enjoying a secondary level are clearly less than those of primary or high education, indicating some kind of a bi-modal distribution. For its part, Portugal presents very low levels of above-primary education (only 10% among the self-employed and 20% among the wage earners).

[25] The percentage change in wage for group i relative to the base group, say d_i, can be calculated by $d_i=e^{\beta i}-1$, where β_i is the coefficient for the dummy variable for group i.

However, we should note that in some countries the samples are quite short for the case of the self-employed. In order to circumvent this possible source of inefficiency, equation (2) has been re-estimated, pooling the self-employed and wage earners into one single sample and including dummy variables expressing jointly the level of education and the employment status. This specification considers that the effects on wages of all the variables, except education, are the same across both types of workers. Table 4 presents the results. The category of reference is that of a self-employed worker with primary level of education. In eight countries, returns associated to primary education are higher for the wage earners, as the coefficients for the self-employed that are attained this educational level are significant and negative. In two countries, Ireland and UK, returns are similar for both groups, whereas in three countries, France, Germany and The Netherlands, the return is higher for the self-employed. For the secondary and higher education we compare the coefficients of the same level education for both type of workers. In the majority of cases, the returns are higher for wage earners than for self-employed of the same education level, except in Germany for secondary level and France and Ireland for the high level. In general, we can observe that, when computed in this way, returns to education are usually slightly higher for wage earners than for the self-employed.

(Tables 3 and 4 about here)

As regards the earnings-experience profile, on-the-job training increases human capital accumulation along the life cycle, as expected, attaining the maximum return when the worker has around 30 years experience, albeit with differences across countries.[26] To facilitate the comparison of results for the different countries being studied, we have computed the rate of return as $\mu_1 + \mu_2 \, Exp/50$, evaluated at the sample average in each country. Looking at the first column in Table 5, it can be observed that, except in the case of France, this rate is higher for wage earners than for the self employed.

(Table 5 about here)

Nevertheless, some problems may be biasing the estimated coefficients of interest. As argued earlier, the EGIV Hausman-Taylor procedure can be applied in order to control for such biases, and the results obtained when using this technique are shown in Table 6. We have previously carried out a pair of Hausman tests to investigate which

[26] The point where experience stops adding positively to earnings is defined by $\partial lnw/\partial Exp$, from earnings function (2). This is equal to $\mu_1+\mu_2 \, Exp/50=0$; $\mu_2<0$.

is the most adequate estimator and to identify the appropriate instruments. Following Baltagi et al. (2003), a first Hausman test is the standard one to distinguish between the random and fixed effects estimators. In almost all of the cases, the random effects hypotheses is rejected in favor of the fixed effect estimator (see column H1 in Table 6). A second Hausman test contrasts the Hausman-Taylor against the fixed effects model. Although the fixed effects estimator is not an option in our study, since it does not allow the estimation of the coefficients of the time invariant regressors, it is useful in order to test the strict exogeneity of the regressors that are used as instruments in the Hausman-Taylor estimation. Thus, when strict exogeneity for a set of regressors is rejected, others must be considered in the estimation to act as instruments. Once the second Hausman test has identified which are the regressors that are strictly exogenous, they are used as instruments in the Hausman-Taylor estimation, see column H2. Again, due to the short data samples for the self-employed in some countries, and with the aim of controlling for selection biases, a pooled estimation has been made, and the corresponding results are presented in Table 7. We will consider these jointly.

(Tables 6 and 7 about here)

Comparing the coefficients of Tables 6 and 7 with those set out in Tables 3 and 4, we can note that the Hausman-Taylor estimation provides coefficients of education and experience that, in general, are consistently much higher than those obtained by GLS. This is in accordance with the typical finding reported in the literature when using instrumental variables. However, a more interesting exercise is to compare the results obtained across countries, which allows us to draw a number of conclusions.

Focusing first on returns to experience, these are very similar for both type of workers, even though greater returns seem to appear for the wage earners**.** However, this may simply be an indication of the fact that these workers usually exhibit fewer years of experience (than their self-employed counterparts), thereby reflecting a higher valuation of the scarce resource. As a consequence, there is a certain amount of evidence that competitive functioning of the labor market may be at work in these countries. Whilst the different theories cannot be compared one with another in the absence of a more detailed analysis, it nevertheless appears that imperfections in the labor market play a less relevant role than expected. A clearer impression can be derived from Figure 1, which displays the earnings-experience profiles for both types of workers in the sample countries. These profiles have been constructed from the coefficients estimated in Table 6, and have also been used to obtain the returns to experience evaluated at the sample

average in each country, which are presented in column 2 of Table 5. In the cases of Belgium, Denmark and France, the results must be taken with care because of the short samples available for the self-employed producing non-significant coefficients. In the rest of the countries, distinct patterns of behavior can be observed. Thus, the profiles in Germany and Spain for both employment statuses are very similar. In Ireland and Italy, the profiles are always steeper for the self-employed, with a similar result being found in Austria, Finland, The Netherlands and the UK, albeit only when experience is greater than 18 years. The evidence for all these countries seems to point to a certain degree of competitiveness in the labor markets. Finally, in Greece and Portugal, the profiles for wage earners are clearly steeper than for the self-employed, indicating some kind of non-competitive environment in the labor market. However, in both these countries, this conclusion must again be treated with care, since their self-employment rates are the highest in the EU, and further bearing in mind that the proportion of the highly educated among the self-employed is quite reduced. Overall, the body of evidence we can present seems to indicate that competitive aspects should not be discarded when analyzing wage determination in the EU labor markets.

(Figure 1 about here)

Turning now to the returns to education, we find that there are few differences between wage earners and the self-employed, with returns usually being higher in the case of wage earners. Only in France are returns clearly higher for the self-employed. Similarly, in most of the sample countries the percentage changes across educational categories are marked, giving support to a certain relevance on the part of the sheepskin effect. This points to some degree of sorting or signaling role being played by education. However, as regards the returns for the different education levels, the regularity noted when estimating by GLS is no longer observed. Indeed, the variability is now so high that no common features are detected.

In summary, as regards the functioning of the labor market in the set of EU countries considered in this paper, two basic ideas emerge. First, according to the evidence shown by the earning-experience profiles, we can note certain traits of competitiveness, given that profiles tend to be steeper for the case of the self-employed. Secondly, returns to education are, in general, found to be higher for wage earners, which can be interpreted as an indication of the relevance of the signaling role of education in determining earnings. This latter result was expected bearing in mind the

prevalent payment schedules in the EU countries, where wages are usually linked to the education level attained by the worker.

6. Conclusions

In this paper we have estimated the rates of return to education and to experience in a set of 15 pre-enlargement EU countries, distinguishing between the self-employed and wage earners. We should first recall that self-employment activities have attained a certain degree of importance in the EU, with the self-employment rate averaging some 15%, albeit with marked differences among the Member States. Against this background, the aim of this paper has been to extend the existing research on the returns to human capital accumulation that differentiates between the self-employed and wage earners in various directions. First, by providing evidence in a cross-country framework using a homogenous database, which mitigates the problems associated with the existence of different data sources across countries. Secondly, by using a panel data approach that is useful in dealing with endogeneity and selectivity biases, as well as the unobserved heterogeneity. Thirdly, by applying an efficient estimation method that caters for the correlation between individual effects and the time-invariant regressors, and that avoids the insecurity associated with the choice of the appropriate instruments.

Additionally, the self-employed have been used as a control group to help in assessing the true impact of credentials achieved in the process of wage determination, as well as in determining which type of theoretical structure underlies labor market behavior. We have operated under the premise that, on the basis that signaling is of much less relevance for the self-employed, comparing across both types of employment statuses should show that, for the sorting hypothesis to be accepted, returns to education for wage earners are significantly higher than those for the self-employed, as well as possibly increasing in a non-linear way. However, if the returns to education of wage earners are of a similar magnitude to those of the self-employed, then we may be willing to conclude that the human capital hypothesis is a good approximation to reality. Similarly, most of the labor market models based on imperfect information predict steeper experience-earnings profiles for wage earners, whereas competitive traits in the labor market would imply similar or flatter profiles for this category of worker.

Information from the ECHP for the period 1994-2000 has been used and this panel data availability has allowed us to apply a random effects-type model that

provides consistent estimates of the rates of return to education and to experience. Education has been represented by dummies of qualification levels (primary, secondary and higher), whereas experience has been measured as the difference between the current age and the age of initiation at work, thereby expressing actual experience. The results have been presented in a reduced form, with the aim being to provide both comparisons across countries about the earnings differentials between the two employment statuses analyzed, as well as evidence as to whether such differences are consistent with the predictions issued by a variety of theoretical models. Two sets of results are presented. First, we have estimated the Mincerian earnings equation, both for countries and for the self-employed and wage earners, without considering the possible endogeneity of the education and experience variables, that is to say, a GLS estimation (random effects model). We have then considered the possibility of endogeneity in some of the regressors, and applied the Hausman and Taylor method, checking such endogeneity with a pair of Hausman-type tests.

With respect to the first set of results, differences arise when estimating the returns to education and experience for the two samples of workers and when estimating for all workers jointly. Although the results differ across countries, when returns to secondary and high education are found for the case of the self-employed, these are usually higher than for the wage earners in the estimation of two samples and the opposite occurs when using the single sample containing all the workers. We can also observe that, the experience rate return is higher for wage earners than for the self employed, except in the case of France.

The second set of results are different. The Hausman-Taylor estimation provides coefficients of education and experience that are, in general, consistently much higher than those obtained by GLS, and this is in accordance with the typical finding reported in the literature when using instrumental variables.

The returns to experience are very similar for both type of workers, even though greater returns seem to appear for the wage earners. This result, joined to the fact that these workers usually exhibit fewer years of experience, reflect a higher valuation of the scarce resource, which points to a certain evidence that competitive functioning of the labor market may be at work. In Greece and Portugal the results are different. Overall, the body of evidence we can present seems to indicate that competitive aspects should not be discarded when analyzing wage determination in the EU labor markets.

With respect to the returns to education, we find that there are few differences between wage earners and the self-employed, with returns usually being higher in the case of wage earners, with the exception of France, which can be interpreted as an indication of the relevance of the signaling role of education in determining earnings. This latter result was expected bearing in mind the prevalent payment schedules in the EU countries, where wages are usually linked to the education level attained by the worker.

Appendix

The Hausman and Taylor (1981) model can be represented in its most general form as follows:

$$ln\ w_{it} = \alpha_i + X'_{it}\,\delta + Z'_i\,\gamma + v_{it}, \qquad \text{(A1)}$$

where $i = 1, \ldots, N$ and $t = 1,\ldots, T$. The Z_i are individual time-invariant regressors, whereas the X_{it} are time-varying. α_i is assumed to be i.i.d.$(0, \sigma^2_\alpha)$ and v_{it} i.i.d.$(0, \sigma^2_v)$, both independent of each other and among themselves. The matrices X and Z can be split into two sets of variables $X=[X_1, X_2]$ and $Z=[Z_1, Z_2]$, such that X_1 is NT x k_1, X_2 is NT x k_2, Z_1 is NT x g_1, and Z_2 is NT x g_2. The X_1 and Z_1 are assumed exogenous and not correlated with α_i and v_{it}, while X_2 and Z_2 are endogenous due to their correlation with α_i but not with v_{it}. Hausman and Taylor (1981) suggest an instrumental variables estimator which pre-multiplies expression (A1) by $\Omega^{1/2}$, where Ω is the variance covariance term of the error component $\alpha_i + v_{it}$, and then perform 2SLS using $[Q, X_1, Z_1]$ as instruments. Q is the within transformation matrix with $X^* = QX$ having a typical element $X^*_{it} = X_{it} - \overline{X}_i$ and $\overline{X}_i$ is the individual mean. As Baltagi et al. (2003) argue, this is equivalent to running 2SLS with $[X^*, X_1, Z_1]$ as the set of instruments. If the model is identified, in the sense that there are at least as many time-varying exogenous regressors X_1 as there are individual time-invariant endogenous regressors Z_2, i.e. $k_1 \geq g_2$, this Hausman-Taylor estimator is more efficient than fixed effects. If the model is under-identified, i.e. $k_1 < g_2$, then one cannot estimate γ and the Hausman-Taylor estimator of δ is identical to fixed effects.

In the estimation carried out in this paper, the only time-invariant (potentially) endogenous variable is education, whereas there are several time-varying exogenous regressors. Some Hausman-based specification tests (Hausman, 1978) have been applied to choose the more appropriate set of regressors, as discussed in the text. For more details, see Hausman and Taylor (1981), Wooldridge (2002) and Baltagi et al. (2003). The 8.0 version of Stata includes the Hausman-Taylor procedure and is used to obtain the estimates presented in this paper.

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Table 1
Self-employment rates in the 15 pre-enlargement EU States 1987-2000

	1987	1995	2000
Austria	-	10.8	10.5
Belgium	15.3	15.4	13.6
Denmark	9.2	8.4	8.0
Finland	-	14.3	12.6
France	12.7	11.6	10.0
Germany	9.1	9.4	9.7
Greece	35.4	33.8	31.3
Ireland	21.8	20.8	16.5
Italy	24.4	24.5	23.6
Luxembourg	9.2	10.0	8.7
Netherlands	10.1	11.5	10.0
Portugal	27.2	25.8	20.2
Spain	23.5	21.8	18.0
Sweden	-	11.3	9.8
United Kingdom	12.5	13.0	10.9
EU	15.9	15.0	13.6

NOTE: Percentage of self-employed persons over total employed.
SOURCE: Eurostat Labour Force Survey.

Table 2
Mean Values of Earnings, Experience and Years of Schooling. Percentage of Workers in the Three Education Levels

		Earnings/hour	Experience	Primary education	Secondary education	Higher education
Austria	Self-employed	5.70	26.36	21.50	72.02	6.48
		(8.45)	(11.50)			
	Wage earner	8.59	20.54	18.78	72.41	8.81
		(5.05)	(11.16)			
Belgium	Self-employed	9.13	20.49	28.59	31.97	39.44
		(21.73)	(11.82)			
	Wage earner	8.89	18.31	29.52	30.43	40.05
		(4.25)	(10.37)			
Denmark	Self-employed	11.48	28.52	24.98	39.84	35.18
		(16.11)	(12.76)			
	Wage earner	10.10	22.15	19.72	40.57	39.71
		(4.46)	(11.84)			
Finland	Self-employed	8.22	27.05	34.28	44.06	21.66
		(9.18)	(11.41)			
	Wage earner	11.90	21.36	19.26	40.59	40.15
		(8.49)	(11.15)			
France	Self-employed	10.19	25.41	30.69	44.60	24.71
		(17.15)	(11.45)			
	Wage earner	9.28	20.42	32.85	40.61	26.54
		(7.65)	(11.74)			
Germany	Self-employed	9.06	22.94	12.31	46.19	41.50
		(8.26)	(10.98)			
	Wage earner	8.48	20.48	19.48	57.88	22.64
		(4.30)	(11.16)			
Greece	Self-employed	3.40	25.49	61.46	22.96	15.58
		(5.13)	(14.08)			
	Wage earner	4.49	16.89	34.39	33.29	32.32
		(3.09)	(11.10)			
Ireland	Self-employed	9.01	31.45	56.27	31.40	12.33
		(32.56)	(14.34)			
	Wage earner	7.66	18.46	34.30	42.49	23.21
		(4.84)	(12.37)			
Italy	Self-employed	6.21	23.69	56.82	31.95	11.23
		(5.31)	(13.34)			
	Wage earner	6.73	18.03	44.13	44.63	11.24
		(3.55)	(11.07)			
Netherlands	Self-employed	8.86	23.51	20.22	53.78	26.00
		(8.82)	(12.35)			
	Wage earner	9.09	19.55	25.49	51.01	23.50
		(6.32)	(11.26)			
Portugal	Self-employed	2.27	31.96	90.24	6.11	3.65
		(2.68)	(15.89)			
	Wage earner	3.32	19.97	79.11	13.27	7.62
		(2.77)	(13.56)			
Spain	Self-employed	4.51	26.44	65.62	16.87	17.51
		(4.70)	(13.71)			
	Wage earner	5.59	19.53	50.30	20.22	29.48
		(3.63)	(12.48)			
United Kingdom	Self-employed	8.71	25.33	46.48	13.74	39.78
		(9.73)	(13.03)			
	Wage earner	7.88	20.02	45.99	13.70	40.31
		(5.59)	(12.86)			

NOTE.- Standard errors between parentheses. Earnings are expressed in terms of the PPP. Luxembourg and Sweden are excluded from the analyisis because of the lack of adequate data for some variables.

Table 3. Estimated Coefficients of Mincerian Earnings Function by GLS by Employment Status

		Experience	Experience^2/100	Secondary education	Higher education	Number of observations	Number of groups
Austria	Self-employed	0.029*	-0.051*	0.405**	0.987**	2076	689
		(2.36)	(-2.38)	(4.02)	(4.94)		
	Wage earners	0.043**	-0.062**	0.197**	0.483**	15563	4134
		(17.70)	(-11.98)	(9.90)	(14.54)		
Belgium	Self-employed	0.055**	-0.101**	-0.031	0.309**	1648	609
		(4.23)	(-4.09)	(-0.27)	(2.59)		
	Wage earners	0.064**	-0.116**	0.059**	0.266**	12534	3649
		(27.34)	(-20.67)	(3.00)	(12.98)		
Denmark	Self-employed	0.064**	-0.118**	0.194	0.315*	1225	408
		(4.78)	(-5.67)	(1.55)	(2.29)		
	Wage earners	0.062**	-0.105**	0.133**	0.278**	17225	4306
		(32.92)	(-28.00)	(6.98)	(13.30)		
Finland	Self-employed	0.027**	-0.039**	0.108	0.342**	3877	1291
		(3.37)	(-2.83)	(1.80)	(4.50)		
	Wage earners	0.075**	-0.116**	0.133**	0.490**	15704	5032
		(29.60)	(-20.13)	(5.05)	(16.71)		
France	Self-employed	0.033*	-0.044	-0.030	0.640**	1064	485
		(2.26)	(-1.66)	(-0.31)	(4.27)		
	Wage earners	0.026**	-0.045**	0.157**	0.355**	24362	6444
		(18.34)	(-13.80)	(9.45)	(16.86)		
Germany	Self-employed	0.017*	-0.033*	-0.048	-0.035	2697	840
		(2.15)	(-2.27)	(-0.66)	(-0.46)		
	Wage earners	0.034**	-0.066**	0.040**	0.220**	34740	8066
		(24.98)	(-21.90)	(3.04)	(13.10)		
Greece	Self-employed	0.034**	-0.067**	0.225**	0.459**	10936	2942
		(10.72)	(-11.92)	(6.45)	(9.42)		
	Wage earners	0.052**	-0.091**	0.205**	0.391**	16324	4488
		(27.69)	(-21.15)	(10.97)	(17.98)		
Ireland	Self-employed	0.028**	-0.043**	0.212**	0.635**	4191	1239
		(5.96)	(-6.10)	(4.69)	(9.52)		
	Wage earners	0.045**	-0.078**	0.170**	0.419**	16146	4949
		(28.63)	(-23.45)	(12.35)	(23.60)		
Italy	Self-employed	0.026**	-0.040**	0.094**	0.291**	8555	2494
		(9.33)	(-7.79)	(3.67)	(6.61)		
	Wage earners	0.041**	-0.071**	0.203**	0.480**	32064	7865
		(35.66)	(-27.24)	(18.72)	(27.32)		
Netherlands	Self-employed	0.033**	-0.056**	0.097	0.395**	1400	544
		(3.94)	(-3.54)	(1.17)	(3.89)		
	Wage earners	0.033**	-0.059**	0.162**	0.374**	23964	6475
		(25.41)	(-19.87)	(12.88)	(23.86)		
Portugal	Self-employed	0.037**	-0.070**	0.313**	0.580**	7129	2209
		(7.20)	(-9.01)	(3.38)	(4.13)		
	Wage earners	0.050**	-0.085**	0.307**	0.796**	28654	6906
		(35.58)	(-31.71)	(14.99)	(28.20)		
Spain	Self-employed	0.056**	-0.090**	0.127*	0.390**	6607	2218
		(9.99)	(-8.94)	(2.15)	(5.85)		
	Wage earners	0.060**	-0.092**	0.224**	0.417**	28313	8274
		(37.34)	(-28.39)	(12.74)	(22.56)		
United Kingdom	Self-employed	0.024**	-0.045**	0.042	0.117**	3515	1053
		(5.34)	(-5.30)	(0.75)	(2.77)		
	Wage earners	0.041**	-0.076**	0.135**	0.237**	27287	6433
		(33.75)	(-30.44)	(7.73)	(19.17)		

NOTE.- t-ratios between parentheses. Both panels are unbalanced, since the employment status may vary across individuals over time. (Controls used. Gender: 1 for male and 0 for female. Marital status: married, single, divorced, widow or separated. Occupational training: if the worker has realized some course of occupational training. Dummies that indicate occupation. Dummies that indicate whether the individual works in the private or public sector. Dummies that indicate seniority: less than two years, between 2 and 10 years and more than 10 years. Dummies that indicate the year.) * Significant at the 5% level. ** Significant at the 1% level.

Table 4
Estimated Coefficients of Mincerian Earnings Function by GLS

	Experience	Experience^2/100	Secondary education x wage earner	Higher education x wage earner	Primary education x self employed	Secondary education x self employed	Higher education x self employed	Number of observations	Number of groups
Austria	0.042**	-0.062**	0.190**	0.476**	-0.697**	-0.190**	0.436**	17639	4687
	(16.11)	(-11.65)	(8.11)	(12.22)	(-12.28)	(-5.26)	(5.63)		
Belgium	0.064**	-0.117**	0.055*	0.258**	-0.274**	-0.330**	0.056	14182	4120
	(24.09)	(-19.41)	(2.26)	(10.23)	(-6.12)	(-7.55)	(1.32)		
Denmark	0.063**	-0.107**	0.132**	0.269**	-0.268**	-0.102**	0.182**	18450	4542
	(31.86)	(-28.23)	(6.34)	(11.82)	(-5.63)	(-2.77)	(4.64)		
Finland	0.069**	-0.106**	0.126**	0.492**	-0.254**	-0.125**	0.158**	19581	6037
	(28.83)	(-20.67)	(4.64)	(16.45)	(-6.75)	(-3.52)	(3.73)		
France	0.026**	-0.045**	0.160**	0.367**	0.110*	0.079	0.604**	25426	6863
	(18.25)	(-13.79)	(9.33)	(17.00)	(2.20)	(1.86)	(10.89)		
Germany	0.033**	-0.060**	0.044**	0.214**	0.198**	0.080**	0.122**	37437	8551
	(23.84)	(-21.20)	(3.27)	(12.56)	(6.03)	(3.79)	(4.94)		
Greece	0.046**	-0.084**	0.184**	0.384**	-0.240**	0.027	0.235**	27260	6812
	(27.91)	(-25.74)	(8.67)	(15.57)	(-12.32)	(1.06)	(7.60)		
Ireland	0.040**	-0.064**	0.174**	0.426**	-0.031	0.156**	0.548**	20337	5960
	(27.44)	(-23.83)	(11.37)	(21.64)	(-1.33)	(6.44)	(15.83)		
Italy	0.037**	-0.062**	0.209**	0.477**	0.034**	0.108**	0.274**	40619	9735
	(35.03)	(-27.16)	(18.93)	(26.74)	(2.74)	(6.83)	(10.83)		
Netherlands	0.033**	-0.059**	0.158**	0.363**	-0.260**	-0.156*	0.307**	25364	6823
	(25.03)	(-19.91)	(11.83)	(21.89)	(-7.60)	(-2.13)	(9.03)		
Portugal	0.049**	-0.087**	0.274**	0.712**	-0.587**	-0.211**	0.259**	35783	8378
	(31.82)	(-31.93)	(11.16)	(20.60)	(-36.07)	(-4.24)	(4.19)		
Spain	0.059**	-0.091**	0.214**	0.424**	-0.383**	-0.213**	-0.023	34920	9809
	(33.90)	(-26.90)	(10.72)	(20.22)	(-19.24)	(-6.34)	(-0.68)		
United	0.039**	-0.073**	0.137**	0.233**	0.013	0.029	0.159**	30802	7000
Kingdom	(32.86)	(-30.04)	(7.75)	(18.64)	(0.76)	(0.94)	(8.17)		

NOTE.- Same as Table 3.

Table 5
Returns to experience evaluated at the sample average

		GLS	Hausman-Taylor
Austria	Self-employed	0,21	2,33
	Wage earner	1,75	2,33
Belgium	Self-employed	1,36	-1,83
	Wage earner	2,15	2,55
Denmark	Self-employed	-0,33	-2,66
	Wage earner	1,55	-0,79
Finland	Self-employed	0,59	1,12
	Wage earner	2,54	-0,46
France	Self-employed	1,06	1,30
	Wage earner	0,76	-0,34
Germany	Self-employed	0,19	2,18
	Wage earner	0,70	2,36
Greece	Self-employed	-0,02	0,23
	Wage earner	2,13	3,17
Ireland	Self-employed	0,10	-0,10
	Wage earner	1,62	1,06
Italy	Self-employed	0,70	1,61
	Wage earner	1,54	1,49
Netherlands	Self-employed	0,67	-0,50
	Wage earner	0,99	-0,17
Portugal	Self-employed	-0,77	-1,25
	Wage earner	1,61	3,67
Spain	Self-employed	0,84	1,59
	Wage earner	2,41	3,84
United Kingdom	Self-employed	0,12	2,51
	Wage earner	1,06	2,06

NOTE.- Own calculations from the estimated coefficients obtained in Tables 3 and 6.

Table 6
Estimated Coefficients of Mincerian Earnings Function by Hausman-Taylor

		Experience	Experience^2/100	Secondary education	Higher education	H1 Hausman test (p-value)	H2 Hausman test (p-value)
Austria	Self-employed	0.057	-0.064	1.588*	2.674**	19.04	2.01
		(1.89)	(-1.22)	(2.22)	(3.49)	(0.7500)	(1.0000)
	Wage earners	0.068**	-0.099**	1.136**	1.534**	461.71	11.50
		(15.36)	(-10.74)	(12.52)	(16.47)	(0.0000)	(0.9778
Belgium	Self-employed	0.071	-0.218**	-0.685	-0.133	26.94	1.25
		(0.92)	(-3.04)	(-0.20)	(-0.05)	(0.3587)	(1.0000)
	Wage earners	0.091**	-0.179**	-0.179	0.511	370.02	4.23
		(2.77)	(-17.19)	(-0.09)	(0.45)	(0.0000)	(1.0000)
Denmark	Self-employed	0.031	-0.101	-0.848	-0.304	29.05	2.95
		(0.82)	(-1.91)	(-1.49)	(-0.49)	(0.2618)	(1.0000)
	Wage earners	0.051**	-0.133**	0.019	0.550**	774.42	0.64
		(14.21)	(-22.16)	(0.12)	(5.11)	(0.0000)	(1.0000)
Finland	Self-employed	0.062**	-0.092**	-0.290	0.590	62.18	21.49
		(3.46)	(-3.13)	(-0.58)	(1.92)	(0.0000)	(0.4905)
	Wage earners	0.080**	-0.198**	-1.713**	-0.126	1380.15	6.59
		(17.95)	(-18.06)	(-6.18)	(-0.68)	(0.0000)	(0.8564)
France	Self-employed	0.133*	-0.245**	0.195	0.252	25.77	6.70
		(2.36)	(-2.06)	(0.13)	(0.11)	(0.0070)	(0.8231)
	Wage earners	0.015**	-0.045**	2.109**	1.745**	1439.18	3.85
		(8.14)	(-9.79)	(18.28)	(19.66)	(0.0000)	(1.0000)
Germany	Self-employed	0.053**	-0.068**	0.654	1.243*	234.96	8.44
		(3.72)	(-2.83)	(0.89)	(2.45)	(0.0000)	(0.9986)
	Wage earners	0.065**	-0.084**	0.848**	1.291**	1117.89	10.51
		(22.24)	(-22.21)	(8.83)	(17.31)	(0.0000)	(0.9921)
Greece	Self-employed	0.057**	-0.112**	0.566**	0.951**	108.31	28.91
		(7.63)	(-8.72)	(2.76)	(6.74)	(0.0000)	(0.2679)
	Wage earners	0.051**	-0.120**	0.655**	0.986**	423.74	19.33
		(11.32)	(-16.41)	(2.11)	(9.49)	(0.0000)	(0.7809)
Ireland	Self-employed	0.062**	-0.104**	0.217	1.234**	183.95	7.03
		(6.45)	(-7.26)	(0.84)	(6.24)	(0.0000)	(0.9998)
	Wage earners	0.056**	-0.123**	0.759**	1.081**	1362.05	22.35
		(8.30)	(-23.85)	(6.69)	(15.29)	(0.0000)	(0.06157)
Italy	Self-employed	0.037**	-0.053**	0.339**	0.631**	138.47	34.66
		(6.30)	(-4.66)	(2.78)	(6.35)	(0.0000)	(0.0946)
	Wage earners	0.047**	-0.089**	0.551**	0.847**	2103.26	49.63
		(27.52)	(-23.60)	(15.09)	(19.07)	(0.0000)	(0.0752)
Netherlands	Self-employed	0.019	-0.051	0.856**	0.753*	47.68	25.55
		(1.04)	(-1.64)	(2.79)	(2.29)	(0.0120)	(0.2716)
	Wage earners	0.028**	-0.076**	0.955**	1.077**	1390.22	8.21
		(14.27)	(-16.65)	(21.83)	(24.93)	(0.0000)	(1.0000)
Portugal	Self-employed	0.064**	-0.118**	3.517**	0.762*	187.69	22.93
		(4.31)	(-5.43)	(5.23)	(2.00)	(0.0000)	(0.5818)
	Wage earners	0.066**	-0.103**	0.420	2.201**	1585.12	21.94
		(17.83)	(-24.81)	(0.70)	(4.46)	(0.0000)	(0.6394)
Spain	Self-employed	0.075**	-0.127**	1.023*	0.696**	60.61	26.01
		(4.93)	(-4.98)	(2.45)	(3.63)	(0.0001)	(0.4069)
	Wage earners	0.063**	-0.139**	-1.694	1.019**	931.17	30.43
		(16.96)	(-24.85)	(-1.89)	(4.29)	(0.0000)	(0.2088)
United Kingdom	Self-employed	0.053**	-0.055**	0.629	0.714**	275.90	5.92
		(5.16)	(-3.41)	(0.74)	(3.54)	(0.0000)	(0.9999)
	Wage earners	0.063**	-0.106**	1.524**	0.706**	2064.11	4.58
		(27.27)	(-27.28)	(8.41)	(15.14)	(0.0000)	(1.0000)

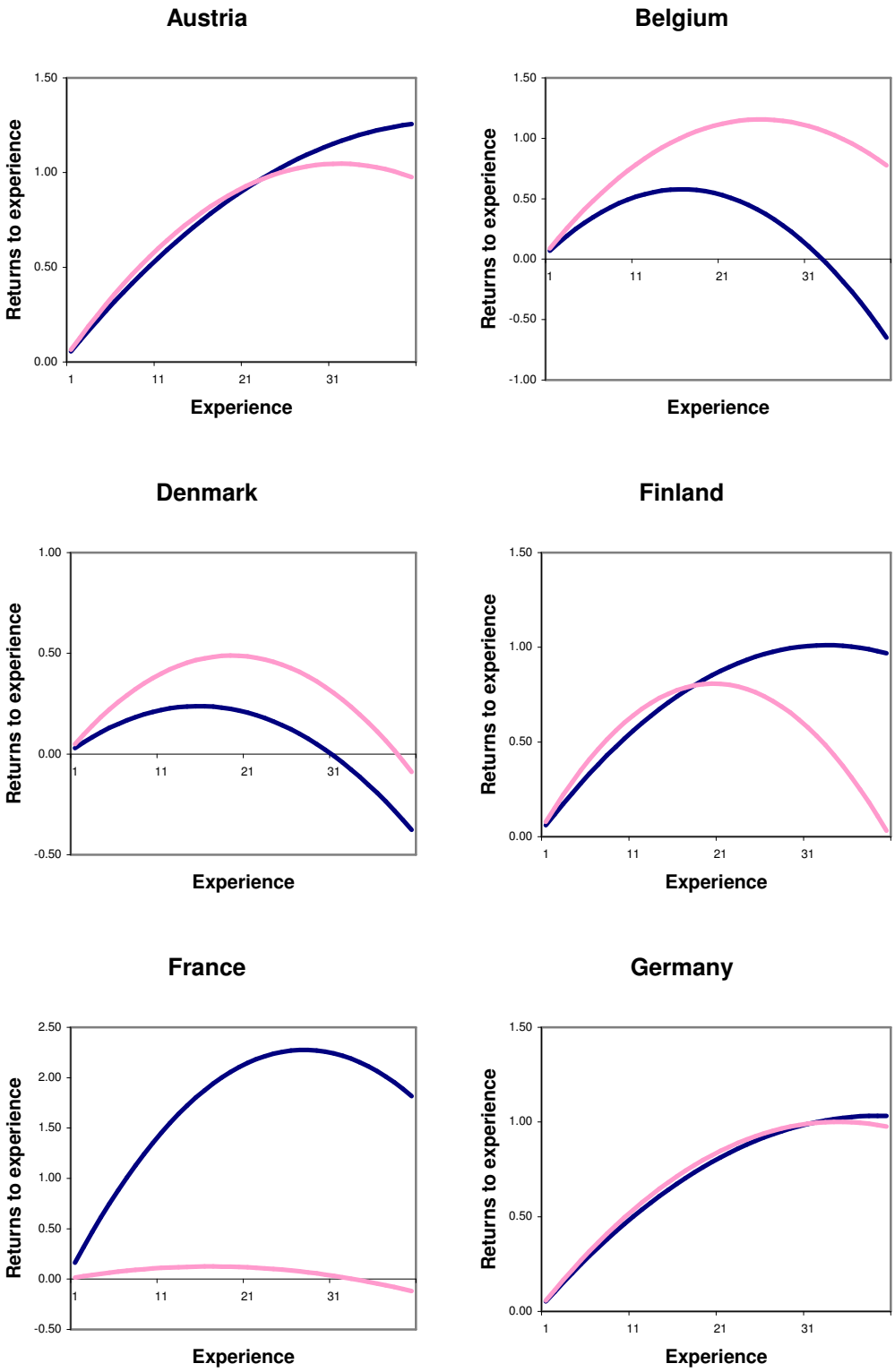
NOTES.- Same as Table 3. H1: This tests the random effects estimator against the fixed effects. H2: This tests the Hausman-Taylor estimator against the fixed effects.

Table 7
Estimated Coefficients of Mincerian Earnings Function by Hausman-Taylor

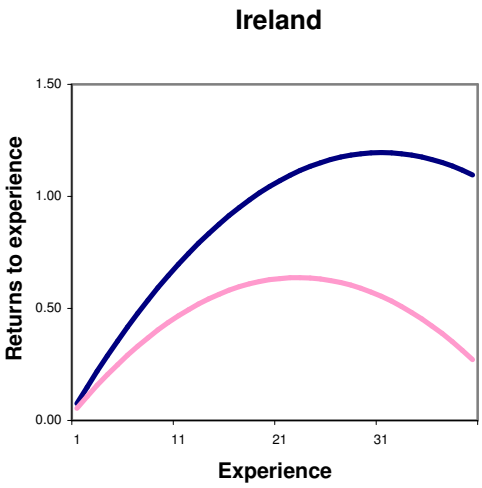
	Experience	Experience^2/100	Secondary education x wage earner	Higher education x wage earner	Primary education x self employed	Secondary education x self employed	Higher education x self employed	H1 Hausman test (p-value)	H2 Hausman test (p-value)
Austria	0.061**	-0.093**	1.206**	1.602**	-0.551**	1.118**	1.670**	296.29	38.26
	(9.93)	(-9.99)	(3.45)	(5.07)	(-4.91)	(3.17)	(5.10)	(0.0000)	(0.0673)
Belgium	0.089**	-0.190**	-0.229	0.422	0.023	-0.351	0.396	342.19	17.34
	(3.55)	(-14.93)	(-0.15)	(0.46)	(0.35)	(-0.24)	(0.44)	(0.0000)	(0.0201)
Denmark	0.060**	-0.128**	0.042	0.503**	-0.101	-0.115	0.506**	574.62	3.98
	(16.51)	(-19.49)	(0.29)	(5.14)	(-1.57)	(-0.77)	(4.83)	(0.0000)	(1.0000)
Finland	0.030**	-0.182**	-4.400**	-1.791**	-0.106*	-4.518**	-1.905**	778.73	24.67
	(6.20)	(-17.31)	(-9.13)	(-5.58)	(-1.98)	(-9.33)	(-5.89)	(0.0000)	(0.4809)
France	0.014**	-0.048**	1.151**	1.225**	0.549**	1.293**	1.264**	1336.07	4.50
	(8.46)	(-11.88)	(4.05)	(6.50)	(5.58)	(4.44)	(6.12)	(0.0000)	(1.0000)
Germany	0.057**	-0.078**	0.911**	1.335**	-0.244	0.477**	0.983**	986.26	7.21
	(23.92)	(-20.18)	(9.57)	(18.42)	(-0.80)	(6.61)	(11.07)	(0.0000)	(1.0000)
Greece	0.056**	-0.111**	0.337**	0.755**	-0.234**	0.238**	1.615**	390.64	14.82
	(16.06)	(-17.25)	(2.52)	(10.26)	(-9.47)	(1.74)	(7.82)	(0.0000)	(0.9803)
Ireland	0.044**	-0.121**	-0.003	0.744**	-0.030	0.030	0.788**	1318.97	8.55
	(13.99)	(-25.84)	(-0.01)	(3.65)	(-0.98)	(0.10)	(3.81)	(0.0000)	(0.9998)
Italy	0.028**	-0.070**	0.464**	0.748**	0.036*	0.337*	0.618**	1247.16	21.67
	(14.60)	(-19.18)	(2.80)	(4.13)	(2.09)	(2.02)	(3.39)	(0.0000)	(0.7964)
Netherlands	0.026**	-0.072**	0.981**	1.037**	-0.154**	0.931**	1.081**	2398.04	18.18
	(13.14)	(-15.69)	(15.10)	(17.96)	(-3.22)	(13.11)	(15.79)	(0.0000)	(0.8692)
Portugal	0.062**	-0.112**	1.521**	1.422**	-0.454**	0.001**	1.032**	1558.36	9.42
	(16.27)	(-20.16)	(8.41)	(11.21)	(-22.26)	(5.25)	(7.75)	(0.0000)	(0.9996)
Spain	0.075**	-0.130**	0.640**	0.989**	-0.277**	0.419	0.624**	652.14	19.74
	(20.02)	(-20.05)	(3.90)	(13.95)	(-10.52)	(2.48)	(7.72)	(0.0000)	(0.8738)
United	0.036**	-0.100**	0.039**	0.437*	0.015	0.393**	0.370	1790.54	18.70
Kingdom	(9.14)	(-19.15)	(3.74)	(2.15)	(0.59)	(3.64)	(1.81)	(0.0000)	(0.9070)

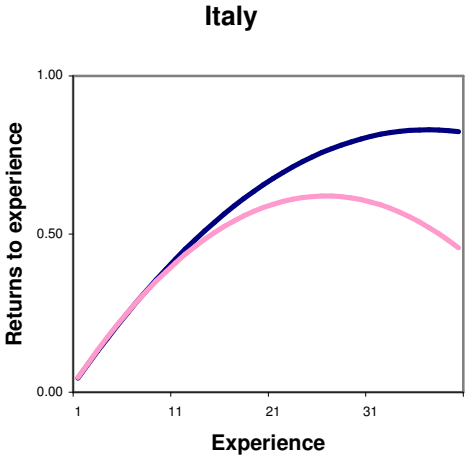
NOTE.- Same as Table 6.

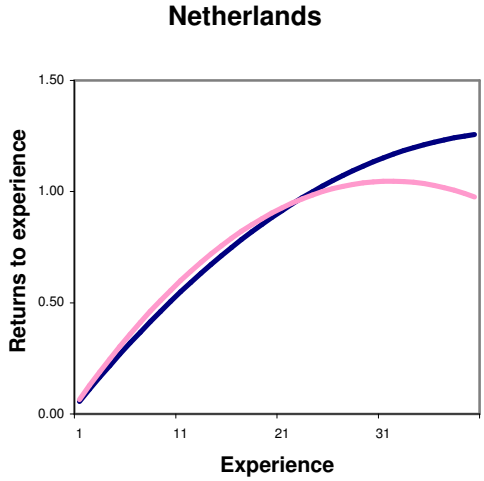
Figure 1. Earnings experience profiles for the 13 sample countries

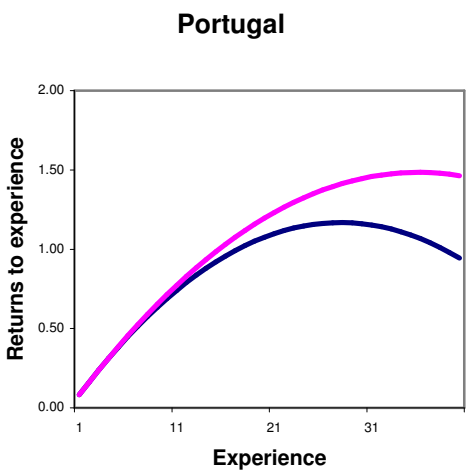


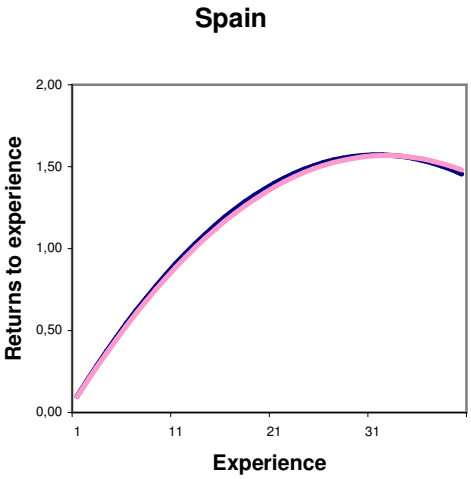

Greece
Returns to experience
1.50
1.00
0.50
0.00
1
11
21
31
Experience

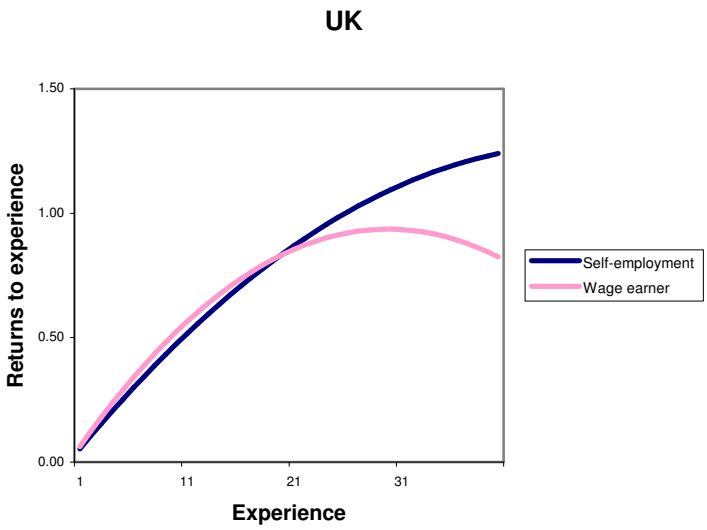

Ireland
Returns to experience
1.50
1.00
0.50
0.00
1
11
21
31
Experience


Italy
Returns to experience
1.00
0.50
0.00
1
11
21
31
Experience


Netherlands
Returns to experience
1.50
1.00
0.50
0.00
1
11
21
31
Experience


Portugal
Returns to experience
2.00
1.50
1.00
0.50
0.00
1
11
21
31
Experience


Spain
Returns to experience
2,00
1,50
1,00
0,50
0,00
1
11
21
31
Experience


UK
Returns to experience
1.50
1.00
0.50
0.00
1
11
21
31
Experience
Self-employment
Wage earner

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