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RETURNS TO TYPE OR TENURE?

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Returns to Type or Tenure?

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Abstract

A regression of wages on firm tenure is likely to yield biased estimates of the returns to tenure because tenure and wages are confounded by unobserved attributes of the job and the unobserved quality of the match between the firm and the employee. Previously, the within-job variation in tenure has been used as an instrument to estimate the average returns to tenure. In this paper, we propose to use instead an easy-to-implement control function estimator for the returns to tenure and their dependence on unobserved heterogeneity. The obtained results for Germany indicate that there is a substantial amount of unobserved heterogeneity in the returns to tenure and that good job matches are characterized by higher returns to tenure in the first five years and lower returns thereafter.

JEL Classification: J31.

Keywords: Wage growth, returns to tenure, selection on unobservables, control function approach, nonseparable model.

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1. Introduction

Economists often distinguish between individual wage growth that is due to the acquisition of general skills over time and wage growth that is due to the acquisition of firm specific skills. The former make an individual more productive in any job and are called returns to experience while the latter are called returns to tenure. The important difference between the two is that the returns to tenure are lost when individuals change firms so that from a policy perspective, it is of considerable interest whether they are sizable. For example, if the returns to tenure are high and reflect the productivity of the employee in his job, then a policy maker might want to design an employment program in such a way that matches between employers and employees are not only formed quickly, but that the ensuing employment relationships will also last long.

When trying to estimate the returns to tenure, it is important to keep in mind that they are likely to depend on job characteristics because of the differences in the amount of skills that are accumulated over time (Miller, 1984; Gathmann and Schönberg, 2010, e.g.). Although we may be able to control for observed characteristics of the employee, as well as for firm and job characteristics, there may still be unobserved factors that give rise to a relationship between the returns to tenure and tenure that has to be taken into account when estimating the former.

Besides, we can think of jobs as being ‘search’ or ‘experience’ goods, or both. We call them search goods when match quality is observed immediately. Then, wage levels and tenure are positively related from the beginning of the employment relationship onwards (Jovanovic, 1979a; Farber, 1994). We call them experience goods when workers and employers learn about the quality of a job match over time. Then, wages increase the more with tenure, the higher the (unobserved) initial uncertainty about the quality of the job match. Also, if the promise of wage growth is used by the employer to motivate individuals or to attract the most suitable ones for the job (Carmichael, 1989; Hutchens, 1989), then it is likely that wage levels and the returns to tenure are related because the extent to which this is the case will again depend on unobserved characteristics of the job. In either case, when estimating the returns to tenure, it is important to

take the possibility into account that the returns to tenure, wage levels, and tenure are related.

In this paper, we demonstrate that an easy-to-implement control function estimator can not only be used to estimate the population average returns to tenure but also to characterize the joint determination of wages and tenure. We argue that the latter characterization yields additional valuable insights into the economic mechanisms that are at play. Our main findings for Germany are the following. First, population average returns to tenure are positive, but low. We estimate them to be 2.4 percent per year in the first five years, and 0.3 percent thereafter. Second, returns are heterogeneous and correlated with tenure. In particular, individuals staying in a firm for longer than average (conditional on observed characteristics, most importantly age) experience lower returns to tenure from the sixth year onwards. Before that, their returns to tenure are higher, and they generally earn more on average.

This means that individuals who change jobs often and whose tenure is therefore shorter than average, experience steeper wage-tenure profiles from the sixth year onwards while earning less on average because their wages grow less in the first five years. This is consistent with the idea that in those employment relationships, effort has to be induced by promising higher wages in the future (Lazear, 1979, 1981). Our findings imply that those job-matches are less productive in the sense that wages are lower on average even if we control for a host of job characteristics such as the required job qualifications, the industry, and the occupational prestige.

The remainder of the paper is organized as follows. In Section 2 we discuss related results, focussing on approaches that have been taken to estimate the average returns to tenure and on recent studies for Germany. We present the econometric approach in Section 3 and describe our data in Section 4. Section 5 presents and discusses the results and Section 6 concludes.

2. Related Literature

2.1. Methods

Altonji and Shakotko (1987) and Abraham and Farber (1987) were among the first to point out that the error term in a wage equation is likely to be positively correlated with tenure and that, consequently, a regression of wages on tenure yields upward biased estimates of the effect of firm tenure on wages. This is supported by the theoretical model of Jovanovic (1979a) and search models such as the one by Burdett (1978), in which individuals receive job offers when they search. These models imply that a highly productive job match, once found, is unlikely to end. This is because the better the job match, the less likely it is individuals receive better job offers when they search. Topel and Ward (1992) and Farber (1994) provide empirical support for this.

To estimate the average returns to tenure, Altonji and Shakotko (1987) calculate the deviation of observed within-job tenure from its mean (within the sample) and use it as the principal instrument for tenure. This is a Hausman and Taylor (1981)-type instrument which is, by construction, uncorrelated with the unobserved and fixed individual specific and the job-match specific component of wages, respectively.

As an alternative to this instrumental variables (IV) estimator, Topel (1991) uses the subsample of individuals who do not change jobs to estimate within-job wage growth from first differences. He then subtracts within-job wage growth from observed wages and estimates the returns to general experience. The return to tenure is then given by the within-job wage growth per year minus the return to general experience. See also Altonji and Williams (2005) for a discussion of this approach and a comparison to Altonji and Shakotko (1987).

Topel (1991, p. 153) points out himself that his point estimates are only unbiased if job changes are exogenous. Otherwise, they are upper bounds on the returns to experience and lower bounds on the returns to tenure. In this situation, it is useful to exploit information on firm closures (Kletzer, 1989; Gibbons and Katz, 1992; Neal, 1995, e.g.), provided that it can

be assumed that conditional on observables, the firm closure itself is exogenous and that job finding rates are not related to tenure, which are both strong assumptions. However, if they hold, then a firm's closure exogenously sets the employees' tenure to zero so that the returns to experience and tenure can in principle be estimated by ordinary least squares (OLS) using the subsample of individuals who experienced a firm closure.

Dustmann and Meghir (2005) also use firm closures as a source of exogenous variation in tenure, but additionally control for unobserved individual differences in job finding rates, tenure and actual labor market experience using a control function approach (Heckman, 1978; Garen, 1984; Heckman and Robb, 1985; Florens et al., 2008). The control functions are error terms from reduced forms for actual labor market experience, firm tenure and current labor market participation, and they are included as additional regressors. For identification, Dustmann and Meghir (2005) need exogenous variation in actual labor market experience and firm tenure. They use a sample of individuals that are at most 35 years old and argue that age can be excluded from the wage equation. This then allows them to use variation in age interacted with education. Their paper is related to Topel (1991) in that they first estimate the returns to actual labor market experience and sector tenure using only observations with zero tenure, but controlling for labor force participation, and only then estimate the returns to tenure, after subtracting the previously estimated wage growth that is due to actual labor market experience from wages.

In this paper, we also use a control function estimator. However, instead of age we use the within-job variation in tenure as the source of exogenous variation. This is the Hausman and Taylor (1981)-type instrument that has been used by Altonji and Shakotko (1987). It is well known that IV estimates can either be obtained by replacing the endogenous variable by its fitted value, where the fitted value is obtained in a first-stage regression of the endogenous variable on the exogenous variables and the instrument, or by including the first-stage residual into the set of regressors while keeping the endogenous variable in it. Our estimator can be seen as a direct extension of the Altonji and Shakotko (1987) estimator as we propose to additionally include the interaction term between the endogenous variable and the first-stage residual. Dustmann and

Meghir (2005) also include such a term, but do not provide an interpretation of the coefficient on the interaction between tenure and the error term. One of the main contributions of this paper is to provide such an interpretation and to link our findings to economic theory. This provides additional valuable insights into the mechanisms that are at play.

2.2. Returns to Tenure in Germany

There are a number of studies for Germany, including most recently Dustmann and Meghir (2005), Dustmann and Pereira (2008), and Orłowski and Riphahn (2008). The study by Dustmann and Meghir (2005) stands out in that they use German administrative data (IAB data, see Bender et al., 1996, for a description). Most other studies, including ours, use data from the German Socio-Economic Panel (GSOEP, see Wagner et al., 1993, for a description).

There are advantages and disadvantages to using either of those two data sets. The IAB data are a 1 percent sample of the German dependent employees who pay social security contributions. They contain information on plant closures, and the large number of observations allows Dustmann and Meghir (2005) to use those as a source of exogenous variation that sets tenure to zero. Other advantages to using the IAB data are that there is no attrition and the measure of tenure is very accurate. However, although average daily wages are measured very accurately, the data do not contain information on hours worked so that they can only use daily wages, and not hourly wages, as the dependent variable. Moreover, wages are top-coded. Dustmann and Meghir (2005) therefore use a sample of employees who at most completed apprenticeship training and are at most 35 years old. These individuals can be expected to work “normal” hours, and for them top-coding is not an issue. A drawback, however, is that they are not representative of the German population.

By contrast, the GSOEP data are self-reported by individuals. Therefore, it is more likely that there is some measurement error in the data. However, we use a corrected tenure measure (see below for details). Tenure is measured up to a tenth of a year. Wages are not top-coded,

and working hours are available, so that hourly wages for the month preceding the interview can be calculated. Two additional advantages of the GSOEP data are that they contain detailed information on job characteristics and that they are (roughly) representative of the German population.

Dustmann and Meghir (2005) use IAB data from 1975 to 1995 and perform the analysis separately for skilled and unskilled employees, defining skilled employees as the ones who completed apprenticeship training and unskilled employees as the ones who obtained only a lower qualification. For the unskilled individuals, they estimate the returns to tenure to be 4 percent in each of the first 5 years, and zero thereafter. For the skilled (with completed apprenticeship training), they estimate the returns to be 2.4 percent in each of the first five years, and zero thereafter. As pointed out before, individuals in their data set are at most 35 years old. Using GSOEP data from 1984 to 1999 for non-self-employed white men aged between 18 and 60 and the Altonji and Shakotko (1987) IV estimator (denoted ‘IVten2’ in their paper), Dustmann and Pereira (2008) estimate 10-year returns to be around 6 percent for both of those skill groups, and zero for university graduates. Finally, Orłowski and Riphahn (2008) use GSOEP data from 2002 to 2006 on male individuals aged 25 to 60 and find, using the same estimator, that returns are not significantly different from zero for individuals working in the private sector.

3. Econometric Approach

We model the real log hourly wage, y_{ijt} , of individual i in firm j at time t as a function of tenure, t_{ijt} , covariates, x_{ijt} , and a structural vector-valued error term, ε_{ijt} , which will be allowed to be correlated with tenure. It is instructive to think of this vector as being composed of several components including a fixed individual-specific error component, ε_i , a fixed job match specific error component, ε_{ij} , an individual specific transitory component, ε_{it} , a transitory match specific component, v_{ijt} , and an economy-wide wage disturbance, ε_t .

For the ease of the exposition, we present the model with only a linear term in tenure. The

wage equation is then given by

$$(1) \quad y_{ijt} = \beta_1(\varepsilon_{ijt}) + \beta_2(\varepsilon_{ijt}) \cdot t_{ijt} + \beta_3(\varepsilon_{ijt}) \cdot x_{ijt}.$$

This is a correlated random coefficient model with random coefficients $\beta_1(\varepsilon_{ijt})$, $\beta_2(\varepsilon_{ijt})$ and $\beta_3(\varepsilon_{ijt})$, which are functions of the vector-valued error term ε_{ijt} .

The model we estimate is richer in that we use a piecewise linear function in tenure. x_{ijt} includes a third order polynomial in age and a host of control variables for observed differences in jobs and occupations across individuals and time that could confound wages and tenure. The coefficients on the control variables have a causal interpretation whenever the control is exogenous in the sense of being independent of ε_{ijt} . Otherwise, the controls are valid for estimating the returns to tenure if an external manipulation of tenure does not alter the value of the control. See Heckman and Vytlacil (2005, p. 677) for a discussion of this condition in the context of a binary treatment indicator instead of tenure. This condition is likely to hold for individual characteristics as they are pre-determined, and job characteristics as an external manipulation of tenure will not induce individuals to change jobs.

The specification in (1) nests a generalized Mincer (1974) type wage equation as the special case in which $\beta_1(\varepsilon_{ijt}) = \beta_1 + \varepsilon_i + \varepsilon_{ij} + \varepsilon_{it} + \varepsilon_t + v_{ijt}$ and non-random coefficients $\beta_k(\varepsilon_{ijt}) = \beta_k$ for $k = 2, 3$. Altonji and Shakotko (1987) estimate the returns to tenure for such an equation using the variation of tenure and its square over a given job match, \tilde{t}_{ijt} and \tilde{t}_{ijt}^2 , as instruments for a linear and a quadratic tenure term. Formally, we denote the average of i 's tenure in firm j in the sample by \bar{t}_{ij} . Then, $\tilde{t}_{ijt} \equiv t_{ijt} - \bar{t}_{ij}$ is the deviation from this average, and $\tilde{t}_{ijt}^2 \equiv t_{ijt}^2 - \bar{t}_{ij}^2$ is the deviation of the squared tenure from its average. These instruments are, by construction, uncorrelated with the job-match specific component ε_{ij} and the individual specific components ε_i of the error term since \tilde{t}_{ijt} sums to 0 over the sample years in which individual i is in job j and ε_i as well as ε_{ij} are constant for i in job j . Crucially, Altonji and Shakotko (1987) assume that \tilde{t}_{ijt} is uncorrelated with time-varying components, ε_{it} , ε_t and v_{ijt} , of the error term in the wage

equation, and we make the same assumption here (in addition they make the same assumption for \tilde{t}_{ijt}^2). Using within-variation to construct instruments in this way has been suggested by Hausman and Taylor (1981), who considered IV estimation of a panel data model with fixed effects.

In this paper, instead of thinking of the slope coefficients $\beta_2(\varepsilon_{ijt})$ and $\beta_3(\varepsilon_{ijt})$ as being constants, we explicitly model them as random coefficients and allow the intercept, $\beta_1(\varepsilon_{ijt})$, and the effect of tenure on wages, $\beta_2(\varepsilon_{ijt})$, to be correlated with tenure. To describe our approach in detail let

$$(2) \quad t_{ijt} = \gamma_1 + \gamma_2 \cdot \tilde{t}_{ijt} + \gamma_3 \cdot x_{ijt} + \eta_{ijt}$$

be the reduced form for tenure. Here, we include \tilde{t}_{ijt} , the instrument of Altonji and Shakotko (1987), on the right hand side so that η_{ijt} is the idiosyncratic—possibly time varying—match-specific component of individual tenure that is orthogonal to the within-job variation of tenure. We impose the normalization that $\mathbb{E}[\eta_{ijt}] = 0$, so that if η_{ijt} is positive we face a type which is, at the time of the observation, longer than expected in a given firm.

The model is flexible since it allows ε_{ijt} and η_{ijt} to be correlated, and thereby allows tenure and the unobserved components of the wage equation to depend on each other. We are able to investigate the dependence of wages on the type if this error term can be consistently estimated. For this, we assume that the vector of unobservables, $(\varepsilon'_{ijt}, \eta_{ijt})'$, is mean-independent of the vector of observables, $(\tilde{t}_{ijt}, x_{ijt})$. Importantly, ε_{ijt} and η_{ijt} are still allowed to be correlated. This is important as the returns to tenure and tenure are thereby allowed to be related.

Finally, we assume that

$$(3) \quad \begin{cases} \mathbb{E}[\beta_k(\varepsilon_{ijt}) | \eta_{ijt}] = \bar{\beta}_k + \phi_k \cdot \eta_{ijt} & \text{for } k = 1, 2 \\ \mathbb{E}[\beta_3(\varepsilon_{ijt}) | \eta_{ijt}] = \bar{\beta}_3, \end{cases}$$

where $\bar{\beta}_k$, $k = 1, 2, 3$, and ϕ_k , $k = 1, 2$ are parameters. This functional form assumption is testable and can easily be relaxed. For example, the linear functional form could be replaced by higher order polynomials and more coefficients could be allowed to depend on the reduced form error term.

In our analysis, we are interested in the expected value of wages given tenure, covariates and the reduced form error term, $\mathbb{E}[y_{ijt}|t_{ijt}, x_{ijt}, \eta_{ijt}]$. Moreover, we are interested in average partial effects, which are given by $\bar{\beta}_1$, $\bar{\beta}_2$ and $\bar{\beta}_3$ in our model, as well as the dependence of the returns to tenure on the reduced form error term, i.e. the constants ϕ_1 and ϕ_2 . Chesher (2003) and Imbens and Newey (2009) show that those features of the structural wage equation are identified once we control for the endogeneity of tenure by including the reduced form error in the wage equation. Invertibility of the reduced form equation in its error term ensures identification, but it need not be additive in it. In our case, this condition is satisfied since (2) is strictly increasing in its error term.

To show identification, notice that for every right hand side variable, e.g. t_{ijt} , we have that

$$\mathbb{E}[\beta_2(\varepsilon_{ijt}) \cdot t_{ijt}|t_{ijt}, x_{ijt}, \eta_{ijt}] = (\bar{\beta}_2 + \phi_2 \cdot \eta_{ijt}) \cdot t_{ijt}$$

by the functional form assumptions and the uncorrelatedness between t_{ijt} and ε_{ijt} conditional on η_{ijt} . Therefore, it follows directly from (1), the stochastic restrictions, and (3) that

$$\begin{aligned} (4) \quad & \mathbb{E}[y_{ijt}|t_{ijt}, x_{ijt}, \eta_{ijt}] \\ &= \mathbb{E}\left[\beta_1(\varepsilon_{ijt}) + \beta_2(\varepsilon_{ijt}) \cdot t_{ijt} + \beta_3(\varepsilon_{ijt}) \cdot x_{ijt} \middle| t_{ijt}, x_{ijt}, \eta_{ijt}\right] \\ &= (\bar{\beta}_1 + \phi_1 \cdot \eta_{ijt}) + (\bar{\beta}_2 + \phi_2 \cdot \eta_{ijt}) \cdot t_{ijt} + \bar{\beta}_3 x_{ijt} \end{aligned}$$

so that estimates of $\bar{\beta}_1$, $\bar{\beta}_2$, $\bar{\beta}_3$, ϕ_1 and ϕ_2 can be obtained from a regression of y_{ijt} on a constant term, tenure, the covariates, the fitted reduced form error term, as well as the interaction thereof with tenure. This procedure requires the rank of the matrix of right hand side variables to be

equal to the number of right hand side variables in this regression, a condition that holds in our data.

4. Data

We use data from the German Socio-Economic Panel (GSOEP), a longitudinal database that started in 1984, in particular sample A through H up to 2008. Our study population are full-time dependent employees that live in West Germany and work in the private sector. Moreover, we select only men that are between 28 and 60 years old. We drop the bottom and the top 5 percent of the observations in terms of their working hours so that individuals in our sample work at least 36 and less than 60 hours per week. Previous month's income is measured in Euros and deflated by the consumer price index with base year 1984.

The outcome of interest is the real gross log hourly wage and is calculated from the information on income earned in the previous month, without extra payments, and the corresponding working hours. Instead of the reported tenure we use the corrected measure that is available in the GSOEP (see SOEP Documentation, 2007, for details). Tenure is reported up to a tenth of a year. We drop observations with a value of less than a month, because then we lack information on the log hourly wage as it is only available for the previous month.

The data include a host of information about employers and their employees. Throughout, we control for the type of blue and white collar job, the qualification required for the current job, the industry as measured by the NACE classification, the occupational prestige measure of Wegener (1988), and the autonomy in the job. We do so by means of indicator variables, except for the measure of occupational prestige, which is a number between 0 and 1000. Categories for blue-collar workers are untrained, semi-trained, trained, foreman, and master craftsman. For white collar workers they are industry foreman, untrained, trained, qualified professional, highly qualified professional, and managerial. For autonomy in the job, the categories are being an apprentice and 5 categories from low to high. The idea is that we thereby control for

Table 1: Summary statistics.

variable	mean	std.	min.	max.
log hourly wage in Euros of 1984	2.348	0.345	1.427	3.938
age	42.505	8.862	28.000	60.000
tenure	13.090	9.755	0.100	46.000

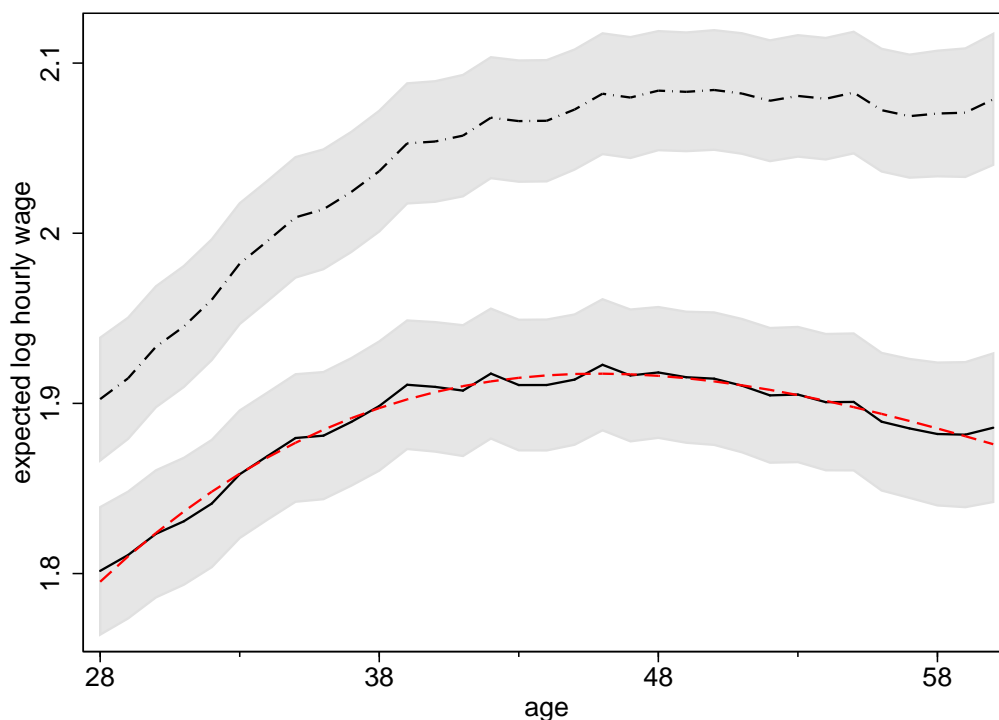
8,305 individuals and 45,621 observations across individuals and time.

differences in wage levels that are related to an employee's qualification and type of job. In addition, we control for time effects by including year indicators, differences in the sampling schemes across subsamples by means of subsample indicators, and geographical differences by including indicators for the state of residence.

Summary statistics for the log hourly wage, age and tenure are reported in Table 1. To graphically analyze the dependence of wages on age we regressed real log hourly wages on a full set of age and year indicators, as well as the aforementioned controls. The dash-dotted line in Figure 1 shows the wage profile against age. Average wages increase most in early ages. This is a stylized fact that is well established in the literature and has been linked to job-shopping in the early phase of a career (Johnson, 1978; Topel and Ward, 1992; Farber, 1994). In the same figure, we compare the total effect of age on wages to the partial effect of age on wages once we subtract wage growth that is associated with tenure, the solid line. Figure 2 shows the partial effect of tenure on wages once we artificially set the age to 42 years. Based on those graphs we specify average log hourly wages to be a third order polynomial in age (see dashed line in Figure 1) and piecewise linear in tenure (see dashed line in Figure 2).

5. Implementation and Results

We estimate the unknown parameters both in the first and in the second stage by ordinary least squares. As fitted values of the first-stage residuals enter the second stage, we need to correct the second stage standard errors for the first stage estimation error. We do so by calculating them

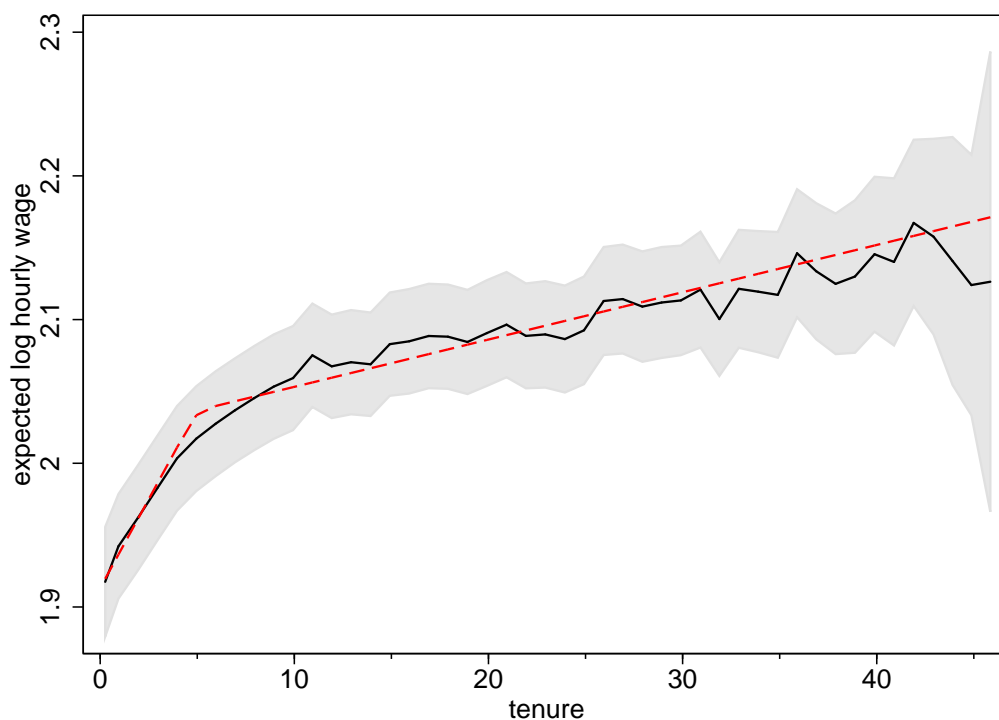


Notes: The dash-dotted line is the average real log hourly wage in Euros of 1984. The solid line is the average real log hourly wage when we subtract wage growth that is due to tenure. Estimates were obtained by regressing the real log hourly wage on a full set of indicator variables for age, and tenure, as well as job characteristics and year, sample and region indicators. The shaded areas depict pointwise 95 percent confidence intervals. The dashed line is the estimated average log hourly wage net of wage growth that is associated with tenure, if we impose that it is a third order polynomial in age.

Figure 1: Partial and total effect of age on wages.

from 1,000 bootstrap replications, estimating the residuals within each bootstrap replication anew. We bootstrap clusters so that the standard errors are also cluster robust and thereby account for correlation of the error terms across periods. Throughout, as described above, we control for job characteristics, year, sample, and region.

First stage estimates are reported in Table 2. The first column is the reduced form for tenure that we use to obtain the fitted values of the residuals, $\hat{\eta}_{ijt}$, that enter the second stage. Notice that \tilde{t}_{ijt} is strongly correlated with tenure. The second and third column are first stages for an IV estimator with piecewise linear function in tenure and a knot at 5 years of tenure. We implement this estimator below for comparison. The functional form specification is based



Notes: The solid line is the wage once we subtract the wage growth that is associated with aging, by setting the age to the median of 42 years. Estimates were obtained by regressing the real log hourly wage on a full set of indicator variables for age and tenure, as well as job characteristics and year, sample, and region indicators. The shaded area depicts the pointwise 95 percent confidence interval. The dashed line is the wage-tenure profile once we impose that it is a continuous piecewise linear function in tenure with a knot at 5 years of tenure.

Figure 2: Partial effect of tenure on wages.

on Figure 2. Formally, the two endogenous variables in this case are $(t_{ijt} - 5) \cdot 1\{t_{ijt} < 5\}$ and $(t_{ijt} - 5) \cdot 1\{t_{ijt} \geq 5\}$, where $1\{\cdot\}$ is the indicator function that takes on the value one if the statement in braces is true, and zero otherwise. Following Altonji and Shakotko (1987), we use the within-job variation in those variables, denoted with a tilde, as instruments. They use a specification that is quadratic in tenure and use the within-individual-firm variation of tenure and its square as excluded instruments. In the table, we use shorthand notation for $1\{t_{ijt} < 5\}$ and $1\{t_{ijt} \geq 5\}$, respectively, which is defined in the notes to the table. The F test for the hypothesis that the coefficients on those two variables are jointly zero indicates that these excluded instruments are strongly related to the piecewise linear function in tenure. Finally, we report Tobit estimates in

Table 2: First stage estimates.

	(1)	(2)	(3)	(4)
	OLS	OLS	OLS	Tobit
	t_{ijt}	$(t_{ijt} - 5) \cdot 1_{5-}$	$(t_{ijt} - 5) \cdot 1_{5+}$	t_{ijt}
\tilde{t}_{ijt}	0.546** (0.020)			0.548** (0.020)
$(t_{ijt} - 5) \cdot 1_{5-}$		0.967** (0.004)	-0.379** (0.021)	
$(t_{ijt} - 5) \cdot 1_{5+}$		-0.014** (0.003)	0.557** (0.019)	
age/10	-11.752* (4.649)	1.981** (0.710)	-13.911** (4.326)	-11.393* (4.661)
age sq./100	3.768** (1.128)	-0.299+ (0.164)	4.105** (1.053)	3.692** (1.130)
age cu./1,000	-0.266** (0.089)	0.016 (0.012)	-0.286** (0.083)	-0.261** (0.089)
F	767.13	50,975.32	1,410.74	769.31
R^2	0.440	0.343	0.445	

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$. 45, 621 observations across individuals and time. The first three columns contain ordinary least squares (OLS) estimates, the last column contains Tobit estimates. There are 8, 305 individuals and standard errors are clustered at the individual level. We also control for job characteristics, year, sample, and region. 1_{5-} and 1_{5+} is shorthand notation for $1\{t_{ijt} < 5\}$ and $1\{t_{ijt} \geq 5\}$, respectively. F is the F -test for the hypothesis that the coefficients on the excluded variables in the first three rows are jointly zero. R^2 is the conventional goodness-of-fit measure.

the fourth column. These are obtained by maximum likelihood estimation, assuming that the error term is normally distributed, and account for the fact that tenure is always greater than one month in our sample. Respective coefficient estimates are very similar to the OLS estimates in the first column.

Table 3 contains estimates of the returns to tenure. We report results for a specification of (1) that is piecewise linear in tenure, as in Dustmann and Meghir (2005), and cubic in age. This choice is based on Figure 1 and Figure 2. We formally tested whether the coefficients on the quadratic terms of a piecewise quadratic function in tenure are significantly different from zero, and they are not (at the 5 percent level). Moreover, we tested whether the square of η_{ijt} as well

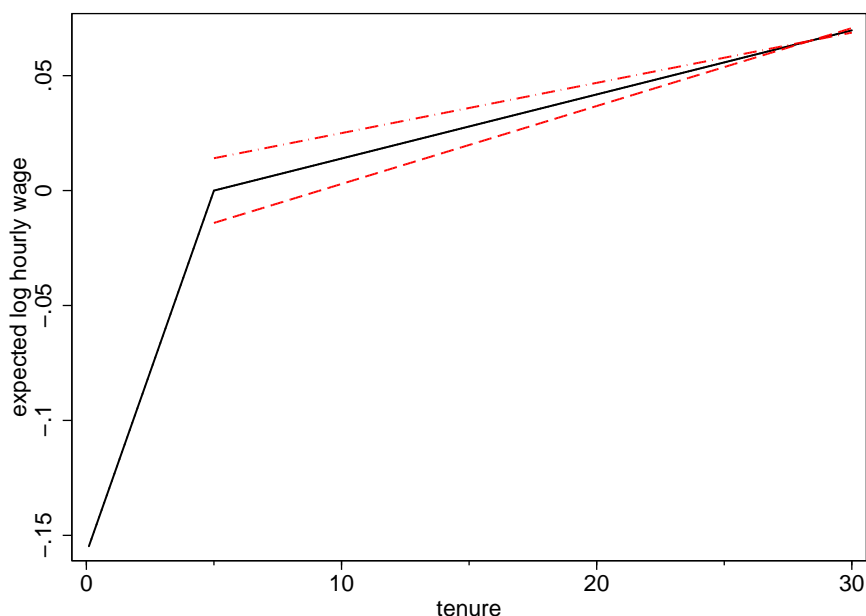
Table 3: Second stage estimates.

	(1)	(2)	(3)
	OLS	IV	CF
$(t_{ijt} - 5) \cdot 1_{5-}/10$	0.247** (0.014)	0.115** (0.017)	0.316** (0.034)
$(t_{ijt} - 5) \cdot 1_{5+}/10$	0.033** (0.003)	0.033** (0.011)	0.028** (0.006)
$\eta_{ijt}/10$			0.028** (0.006)
$(t_{ijt} - 5) \cdot 1_{5-} \cdot \eta_{ijt}/100$			0.129** (0.034)
$(t_{ijt} - 5) \cdot 1_{5+} \cdot \eta_{ijt}/100$			-0.012** (0.003)
age/10	0.648** (0.131)	0.732** (0.132)	0.679** (0.131)
age sq./100	-0.112** (0.031)	-0.129** (0.031)	-0.122** (0.031)
age cu./1,000	0.006* (0.002)	0.007** (0.002)	0.007** (0.002)
observations	45,621	45,621	45,621
clusters	8,305	8,305	8,305

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$. 45,621 observations across individuals and time. The first column contains ordinary least squares (OLS) estimates, the second column instrumental variables (IV) estimates, and the last column control function (CF) estimates. OLS and IV standard errors are analytic and CF standard errors are bootstrapped with 1,000 replications and clustered at the individual level (8,305 individuals). We also control for job characteristics, year, sample, and region.

as its interactions with tenure should be included, and again we found that the coefficients on those additional terms were not statistically different from zero (again at the 5 percent level).

The first column of Table 3 contains OLS estimates. Notice that coefficients have been rescaled. The observed wage increase that is associated with an increase of tenure by 1 year is 2.47 percent in the first 5 years. Thereafter, it is 0.33 percent. The second column contains two stage least squares IV estimates, where the first stage regressions are given in the second and third column of Table 2. As compared to OLS estimates, these IV estimates of the returns to tenure are lower in the first five years, but equal to them thereafter. Yitzhaki (1989), Angrist



Notes: Estimates were obtained using the results in column (4) of Table 3. For the dash-dotted line we use $\eta_{ijt} = 5$, for the middle solid line we use $\eta_{ijt} = 0$, and for the dashed line we use $\eta_{ijt} = -5$. The figure shows the expected log hourly wage relative to a firm tenure of five years.

Figure 3: The impact of η_{ijt} on the returns to tenure.

et al. (2000) and Heckman et al. (2006) show that the IV estimand is a weighted average of the individual returns to tenure (in our case), $\beta_2(\varepsilon_{ijt})$. However and unfortunately, their results are not directly applicable to the situation with two endogenous variables in the sense that they are not useful in predicting which average is actually estimated. One key advantage of the control function estimator used in this paper, relative to IV estimation, is that it allows us to avoid this problem by controlling for the dependence between the returns to tenure and tenure. This is done by including the interaction term between tenure and the first-stage residual as an additional regressor.

The third column contains our main results. We estimate the average annual returns to tenure to be 3.16 percent in the first five years and 0.28 percent thereafter. These estimates are higher than the ones reported by Dustmann and Meghir (2005) for both skill groups. They are also higher than the zero returns estimated by Orłowski and Riphahn (2008), and the implied

returns for the first 10 years of tenure are higher than the 6 percent reported by Dustmann and Pereira (2008).

The coefficient on η_{ijt} indicates that wages are 0.28 percent higher for individuals who have been one year longer with a firm than other individuals with the same characteristics, i.e. those who are, in particular, of the same age and work in similar jobs. We argued earlier that these individuals are good types in the sense that good matches tend to survive. The returns for those individuals are 0.129 percent bigger in the first five years and 0.012 percent smaller thereafter. Figure 3 plots the wage profile for an individual who has been with a firm for five years more than could be expected (the top dash-dotted line; here we use $\eta_{ijt} = 5$), for as long as could be expected (the middle solid line; here we use $\eta_{ijt} = 0$), and for five years less than could be expected (the lower dashed line; here we use $\eta_{ijt} = -5$). We only make this distinction after five years of tenure since an individual who spent five years more than expected with a given firm cannot have a firm tenure equal to a value less than five. In this figure, differences in the returns to tenure are reflected in differences in the slopes. It shows that good types with a positive value of η_{ijt} earn higher wages on average and experience lower returns to tenure from the sixth year onwards. This is consistent with the view that these are indeed good matches between the employer and the employee (Hall, 1982; Topel and Ward, 1992; Farber, 1994), that jobs are experience goods in the first five years (Jovanovic, 1979b), and that in worse job matches effort is induced by the employer using a steep profile of wages in tenure (Lazear, 1979, 1981). Conversely, bad types earn lower wages and experience higher returns to tenure.

Next, in Table 4, we report the same second stage results for different subgroups. In particular, following Dustmann and Meghir (2005), we report the results for unskilled and skilled individuals, where being skilled means that individuals have at least completed an apprenticeship. For both groups, Dustmann and Meghir (2005) find returns that are not statistically different from zero from the sixth year onwards. Although our results are statistically different from zero, they are estimated to be very small, 0.21 percent per year. Our estimates for the returns in the first five years are similar for skilled and unskilled individuals, slightly above 3

Table 4: Second stage estimates for subgroups.

	(1)	(2)	(3)	(4)
	unskilled	skilled	small firm	big firm
$(t_{ijt} - 5) \cdot 1_{5-}/10$	0.327** (0.049)	0.315** (0.039)	0.265** (0.038)	0.417** (0.048)
$(t_{ijt} - 5) \cdot 1_{5+}/10$	0.024** (0.009)	0.024** (0.007)	0.014 (0.009)	0.036** (0.007)
$(t_{ijt} - 5) \cdot 1_{5-} \cdot \eta_{ijt}/100$	0.101+ (0.053)	0.154** (0.040)	0.124** (0.044)	0.166** (0.046)
$(t_{ijt} - 5) \cdot 1_{5+} \cdot \eta_{ijt}/100$	-0.008+ (0.005)	-0.015** (0.003)	-0.006 (0.004)	-0.010** (0.003)
$\eta_{ijt}/10$	0.026* (0.010)	0.034** (0.007)	0.014 (0.010)	0.014+ (0.008)
age/10	0.739** (0.197)	0.662** (0.164)	0.690** (0.190)	0.867** (0.181)
age sq./100	-0.154** (0.046)	-0.109** (0.039)	-0.124** (0.045)	-0.169** (0.043)
age cu./1,000	0.010** (0.004)	0.006+ (0.003)	0.007* (0.003)	0.011** (0.003)
average log hourly wage	2.161	2.437	2.219	2.440
average tenure	13.090	13.420	10.836	13.740
observations	14,713	30,908	18,411	25,447
clusters	3,416	6,327	4,413	5,275

+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$. Standard errors are bootstrapped with 1,000 replications and clustered at the individual level. We also control for job characteristics, year, sample, and region.

percent per year, lower than the 4 percent Dustmann and Meghir (2005) report for unskilled individuals, and higher than the 2.4 percent they report for skilled individuals. We also report the results separately for small firms (with up to 100 employees) and big firms. Here we lose a few observations because firm size is sometimes missing. Individuals in small firms earn lower wages, and average tenure is lower, as compared to individuals in big firms.

Interestingly, controlling for job characteristics, year, sample, and region, we still find a relationship between tenure and the returns to tenure when we perform the analysis separately for unskilled and skilled employees, and separately for employees in small and big firms, respectively. In particular, as before, we find that good types experience higher returns to tenure in the

first five years and lower returns thereafter.

6. Concluding Remarks

In this paper, we propose to use an easy-to-implement control function estimator to estimate the returns to tenure and their dependence on an unobservable type variable. We demonstrate that this does not only allow us to estimate the average returns to tenure but also the dependence between the returns to tenure and an unobservable type variable, which yields additional valuable insights into the economic mechanisms that are at play.

Among equally old individuals with similar jobs, we find that those who work longer for a firm than average experience higher returns to tenure in the first five years and lower returns thereafter.

Results like ours are not only useful for supporting economic theory, but also provide useful information for policy makers. Chetty (2008) shows that individuals search too little because many of them are liquidity constrained. In such a situation, it may be worthwhile to subsidize further search by extending unemployment benefits so that better, longer lasting job matches are established. Our results imply that such longer lasting job matches are characterized by high wage growth in the beginning and higher wages on average, and are thus desirable. This information enables the policy maker to compare the costs of establishing such a longer lasting job match to the benefits.

The estimator we use can of course also be employed in other contexts where the endogenous variable is continuously distributed and where IV estimation is feasible. For example, if we treat years of schooling as (approximately) continuous, then it can be used to characterize the dependence between the returns to schooling and schooling choice, and to investigate, for example, whether these selection patterns are related to the socio-economic background of the individuals by performing the analysis separately for different groups.

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