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Heterogene Arbeit: Positive und Normative
Aspekte der Qualifikationsstruktur der Arbeit

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Returns to Type or Tenure?

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Abstract:

This paper takes a closer look at the way unobserved characteristics of individuals confound wages and firm tenure. In turn of our structural analysis, which is built on estimating a reduced form for tenure and a structural wage equation, we disentangle returns to a latent type variable from estimates of general returns to tenure. The obtained results for Germany indicate that the type plays a crucial role in the remuneration of employees. As compared to types who stay, quitters earn less on average, but obtain steeper wage profiles in tenure. This nonseparability has previously remained unnoticed in the literature.

JEL Classification : J31
Keywords : Returns to tenure, unobserved heterogeneity, control function approach, nonseparable model
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Returns to Type or Tenure?*

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Abstract

This paper takes a closer look at the way unobserved characteristics of individuals confound wages and firm tenure. In turn of our structural analysis, which is built on estimating a reduced form for tenure and a structural wage equation, we disentangle returns to a latent type variable from estimates of general returns to tenure. The obtained results for Germany indicate that the type plays a crucial role in the remuneration of employees. As compared to types who stay, quitters earn less on average, but obtain steeper wage profiles in tenure. This nonseparability has previously remained unnoticed in the literature.

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1. Introduction

The length of an individual's working time span in a firm is influenced by a variety of factors including individual, firm, and match specific factors as well as macroeconomic crises or booms. Once there are different types of employees and employer-employee matches are formed endogenously, match quality is likely to vary across matches. Theoretical work by Jovanovic (1979) and search models such as the one by Burdett (1978) show that a high-productive job match, once found, is unlikely to end and hence, firm tenure may be positively correlated with an employee's productivity and—following human capital theory—his wage. Moreover, tournament models such as the one by Lazear and Rosen (1981) and screening models such as the one by Amann (2004) predict that highly productive employees obtain steeper wage profiles than their less productive colleagues. For empirical work, the former implies that the job match component of the error term in a wage equation is most likely to be correlated with tenure and actual experience, whereas the latter implies that wage-tenure profiles are steeper for individuals who stay longer than average in a given firm.

One empirical approach that has been taken in order to deal with the endogeneity of tenure is instrumental variables (IV) estimation. Altonji and Shakotko (1987) calculate the deviation of observed tenure from its mean and use it as the principal instrument for tenure. A similar instrument is used by Abraham and Farber (1987).[1] As an alternative to this IV approach Topel (1991) implements a two-stage procedure. In the first stage, he estimates within-job wage growth from first differences relying on the subsample of individuals who do not change firms. Using these first stage estimates, he calculates initial wages and estimates the return to initial experience on the job in a second stage.[2]

[1] In Section 2, we comment in detail on this approach and compare it to the techniques used in this paper. For applications of the IV approach see, for instance, Mascle-Allemand and Tritah (2005) comparing wage profiles between states with or without employment protection legislation, Bratsberg and Terell (1997) analyzing the difference in wage growth between young black and white men, and Dustmann and Pareira (2005) investigating the wage growth in Germany and the UK.

[2] Intuitively speaking, he sets tenure to zero by subtracting within-job wage growth from observed

A third approach is to exploit information on firm closures. Assuming that a firm closure is exogenous conditional on observables it could be used in order to identify a subgroup of individuals whose tenure is exogenously set to zero. For those individuals, the returns to tenure can in principle be estimated by OLS once all individuals are equally likely to find a new job. Therefore, one could argue that the approach has advantages over the one proposed by Topel (1991), which hinges on the assumption that individual firm changes, rather than firm closures, are exogenous conditional on observables.

In more recent work, Dustmann and Meghir (2005) use a control function approach in order to control for the additional endogeneity of the time it takes to find a new job after a firm closure. One concern with respect to their approach is that, since they use a subsample of displaced workers, it uncovers the average effect of firm tenure on wages for the subpopulation of individuals who were *forced* to start a new job instead of the population average effect. For instance, it could be that the highly productive employees left the firm before it finally closed and only the less productive are therefore laid off. Another concern is that conditional on observables, a firm closure is in fact not endogenous. In both cases, the subpopulation of laid-off employees would be a selected sample.

Therefore, rather than relying on a subsample of individuals whose plant was closed, we use the full data set and pursue an Imbens and Newey (2003) type control function approach. Although this approach is similar to the approach taken by Dustmann and Meghir (2005) we use it in addition to carefully characterize the impact of the latent type variable on wages. Both analyses are for German data. However, Dustmann and Meghir (2005) use data from German Social Security records (IAB data), while we use data from the German Socio Economic Panel (GSOEP). Therefore, both sets of results can be seen as complementary.

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wages. In an application of this method Connolly and Gottschalk (2001) uncover different tenure effects for different levels of education.

Our approach consists of estimating a reduced form for tenure in order to recover the reduced form error term—the type variable. The type variable is then included into the wage equation in order to control for factors confounding wages and tenure. This allows for type specific matching patterns. In our specification, we explicitly allow for interaction effects between unobservable factors determining tenure and observable covariates in the structural wage equation. Our main findings are the following. First, those types who stay shorter have higher returns to tenure but earn less on average. This nonseparability between the type variable and the effect of tenure on wages has previously remained unnoticed in the literature. It stems from the positive correlation between the error term and firm tenure in a wage equation. Our result challenges theoretical results by Jovanovic (1979) and Burdett (1978). In particular, we argue that the error term in the tenure equation mirrors not only workers' productional performance. Sector, firm, or turnover specific factors may also affect the commonly obtained results. Our argument is bolstered by the result that the error term affects the return to tenure negatively. Considering theoretical approaches, this effect can not be explained by the assumption that the unobserved characteristics corresponds entirely to productivity.

The remainder of the paper is organized as follows. In Section 2, we present our econometric approach in detail. The used data set is described in Section 3. In Section 4, we carefully chose appropriate functional form assumptions and present the empirical results. Section 5 concludes.

2. Econometric Approach

2.1. Econometric Model

The approach taken in this paper is to estimate identifiable features of a structural wage equation in which tenure enters as an endogenous variable. 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}, exogenous covariates, x_{ijt}, and a structural vector valued error term, ε_{ijt}, which will be allowed to be correlated with tenure. We include the individual's age into the wage equation in order to contrast the returns to tenure to the returns to general experience. We decided to include age rather than actual experience since the latter is likely to be endogenous.[3]

For simplicity, we present the model with only a linear term in tenure and one exogenous covariate, say age, denoted by x_{ijt}. The generalization to the specification we take to the data in Section 4 is trivial. Under this simplifying assumption our wage equation is given by

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

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 structural error term. We think of this error term 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, ν_{ijt}, and an economy wide wage disturbance, ε_t.

The specification in (1) nests a generalized Mincer (1974) type wage equation as the special case in which $\beta_1(\varepsilon_{ijt}) = \varepsilon_i + \varepsilon_{ij} + \varepsilon_{it} + \varepsilon_t + \nu_{ijt}$ and $\beta_k(\varepsilon_{ijt}) = \beta_k$ for $k = 2, 3$. Altonji and Shakotko (1987) estimated the returns to tenure for such an equation using instrumental variable (IV) techniques. In particular, they use the variation of tenure and its square over a given job match, $\tilde{t}_{ijt}$ and $\tilde{t}^2_{ijt}$, as IVs for a linear and a quadratic term in tenure.[4] Formally, denote the average of i's tenure in firm j in the sample by $\overline{t_{ij}}$. Then, $\tilde{t}_{ijt} \equiv t_{ijt} - \overline{t_{ij}}$ is the deviation from this average, and $\tilde{t}^2_{ijt} \equiv t^2_{ijt} - \overline{t^2_{ij}}$ is the deviation of the

[3] Additionally, an endogenous control for general experience would possibly influence the coefficient of the tenure variable. See also Altonji and Shakotko (1987) for details.

[4] Once a quadratic function in tenure is estimated, two instruments are of need since both are potentially endogenous.

squared tenure from its average. These instruments are, by construction, uncorrelated with individual match quality, ε_{ij} and individual specific components, ε_i, 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. Moreover, Altonji and Shakotko (1987) argue that $\tilde{t}_{ijt}$ is likely to be independent of time specific idiosyncratic and macroeconomic shocks, ν_{ijt} and ε_t, since they do not affect the wage in the current job relative to other jobs. This is also validated by previous studies such as Topel (1991) which show that ν_{ijt} follows a random walk. For these reasons, Altonji and Shakotko (1987) are confident that $\tilde{t}_{ijt}$ is uncorrelated with ε_{ijt} which qualifies it as an instrument for tenure.

Our random coefficients specification is more general in two ways. First, we do not restrict the slope coefficients $\beta_2(\varepsilon_{ijt})$ and $\beta_3(\varepsilon_{ijt})$ to be degenerate. Second, we will allow the intercept, $\beta_1(\varepsilon_{ijt})$ and the effect of tenure on wages, $\beta_2(\varepsilon_{ijt})$, to be correlated with tenure. To see that this allows for a considerable amount of unobserved heterogeneity let

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

be a stylized version of the reduced form for tenure. η_{ijt} is the reduced form error term—our type variable—for which we assume that $\mathbb{E}[\eta_{ijt}] = 0$.[5] If it is positive, then we face a type which stays longer than expected in a given firm. The model is flexible because it allows ε_{ijt} and η_{ijt} to be correlated and hence allows tenure and the returns to tenure to depend on each other. Under the conditions stated below this error term can be consistently estimated which allows us to investigate the dependence of wages on the type.

As for *stochastic restrictions*, we assume that the observations are independently distributed across ijt.[6] Moreover, we assume that the vector of unobservables, $(\varepsilon'_{ijt}, \eta_{ijt})'$,

[5]This is a normalization given that we include an intercept term.

[6]This assumption is stronger than it is actually needed. For example, our estimator would still be consistent, though not efficient, if η_{ijt} was in fact serially correlated across t for a given i and j. In our analysis, since the panel is unbalanced and very short for a substantial part of the observations,

is mean independent of the vector of observables, $(\tilde{t}_{ijt}, x_{ijt})$. Importantly, this allows for tenure and the return to tenure to be correlated with the structural error term ε_{ijt} via the reduced form error term η_{ijt}. In particular, we assume that

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

for some constants ϕ_k.[7] In the empirical analysis, we show evidence for such a dependence of the expected returns to tenure on the type.

To the contrary, a more classical IV type correlated random coefficient specification such as the one by Heckman and Vytlacil (1998) would additionally assume that η_{ijt} is independent of ε_{ijt} conditional on $(\tilde{t}_{ijt}, x_{ijt})$ and hence prohibit this kind of endogeneity. The characterization of the endogeneity of tenure based on the proposed structure will be at the center of the empirical analysis.

2.2. *Parameters of Interest and Identification*

In our analysis, we are interested in the expected value of wages given covariates and the reduced form error which we interpret as a latent type variable, $\mathbb{E}[y_{ijt}|x_{ijt}, \eta_{ijt}]$. 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 coefficients on the reduced form error term, i.e. the constants ϕ_1 and ϕ_2. Imbens and Newey (2003) show that those features of the structural wage equation are identified from observations once we control for the endogeneity of tenure by including the reduced form error into the wage equation.[8] Essentially, invert-

techniques other than random effects panel estimators were not feasible. However, standard errors will be bootstrapped so that the loss we risk is mainly a loss of efficiency.

[7]This assumption could 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. However, we experienced for our data that the simple specification in (3) worked best.

[8]They show nonparametric identification of several features of the outcome equation and propose a nonparametric two step series estimator. This is in contrast to Newey, Powell, and Vella (1999) who

ibility of the reduced form equation in its scalar disturbance ensures identification. In our case, this condition is satisfied since (2) is strongly increasing in its error term.

In particular, it follows directly from (1), the stochastic restrictions, and (3) that[9]

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

so that estimates of $\bar{\beta}_1$, $\bar{\beta}_2$, $\bar{\beta}_3$, ϕ_1 and ϕ_2 can be obtained from an OLS regression of y_{ijt} on a constant term, tenure, age, the fitted reduced form error term, as well as the interaction thereof with tenure. This proceeding requires that the rank of the matrix of right hand side variables to be 5, a condition which holds in our data.

Finally, an estimate for the expected value of wages given covariates can be obtained by integrating over the sample distribution of the fitted reduced form error term. Trivially, since the mean of the reduced form error term is zero by assumption, we have that by the stochastic restrictions $\bar{\beta}_1$ and $\bar{\beta}_2$ are the population mean intercept and the population mean return to tenure, respectively.

3. Data and Descriptive Evidence

The data we use for our analysis stem from the German Socio-Economic Panel (GSOEP), a longitudinal database that started in 1984. We use all samples up to sample F. We only

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consider additive structures. Blundell and Powell (2003) survey the recent literature for such models and form the terminology "average structural function" for a prominent identifiable feature which is linked to the average treatment effect parameter in program evaluation. Earlier work by Garen (1984) is more similar to our specification but is built on normality as an identifying assumption.

[9] For every right hand side variable, say e.g. x_{ijt}, we have that

$$\mathbb{E}[\beta_3(\varepsilon_{ijt}) \cdot x_{ijt}|t_{ijt}, x_{ijt}, \eta_{ijt}] = \bar{\beta}_3 \cdot x_{ijt}$$

by the uncorrelatedness between x_{ijt} and ε_{ijt} conditional on η_{ijt}.

Variable	*Mean*	*Std.*	*Min.*	*Max.*
annual income in Euros of year 2000	34,351.38	18,785.25	613.55	697,524
annual hours worked	2,221.16	438.78	352	5,196
log wage	2.66	0.41	-0.75	6.32
age	42.07	8.94	28	60
tenure	13.96	9.05	1	52
year	1994.05	6.05	1984	2003
mean number of obs. per individual	10.05	5.72	1	20
number of firms per individual	1.15	0.44	1	6

6831 individuals and 39,200 observations across individuals and time.

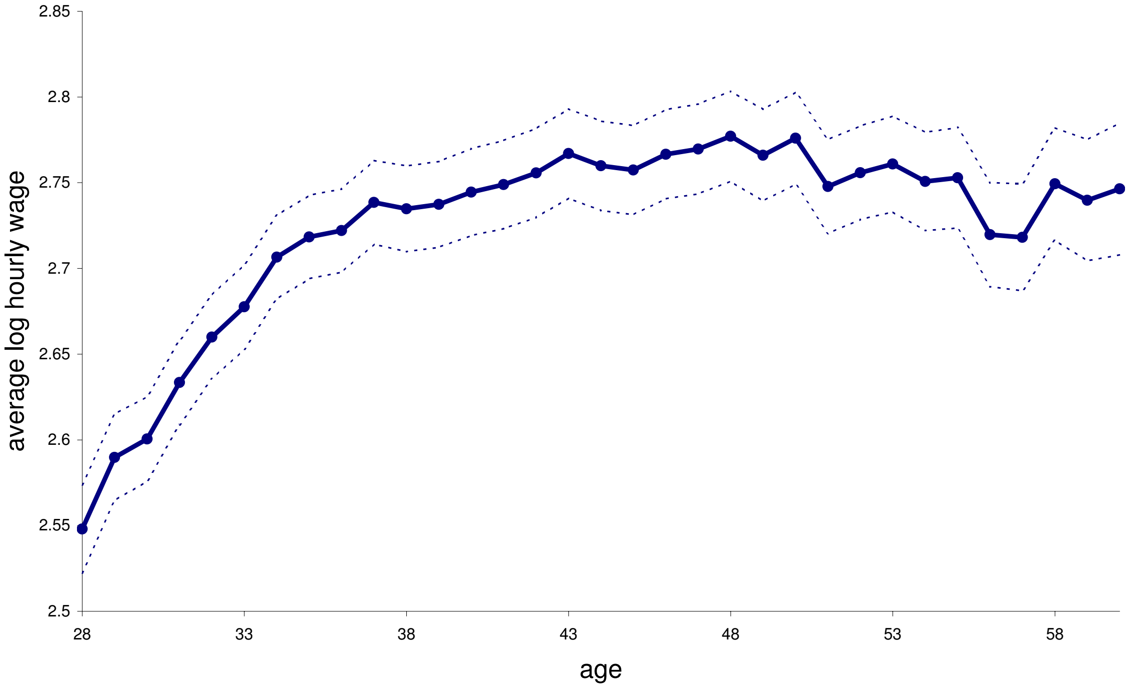
Table 1: Summary statistics.

select privately employed men from West Germany as women and individuals working in East Germany decide upon their career based on systematically different circumstances. The outcome of interest is the real log annual wage, deflated by the consumer price index. Furthermore, we focus only on individuals who are between 28 and 60 years old, and whose weekly hours in their work contract is at least 35. Moreover, we require them to work for at least 350 hours in a given year. We select observations with at least one year of firm tenure because we fear that wages are measured with error in years firm tenure is equal to zero.[10] Some summary statistics are reported in Table 1.

As a first piece of descriptive evidence, we regressed real log hourly wages on a full set of age and year indicators. Figure 1 contains the wage profile against age. Apparently, wages rise sharply until the age of about 43. Here, we remain silent about the sources of wage growth as we do not condition on tenure or firm changes, for example. Figure 2, in turn, shows that real wages for a given age have been increasing considerably over the last 20 years.

In Figure 3 we juxtapose the total effect of age on wages, the dash-dotted line, and the partial effect of age on wages once we condition on firm tenure. Interestingly, the partial

[10]This is innocuous since identification in our case does not rely on observations with zero firm tenure, as opposed to the approach taken by Topel (1991), e.g. Moreover, it comes in handy as we will decide in favor of a specification in which we include log tenure on the right hand side of the regression equation.

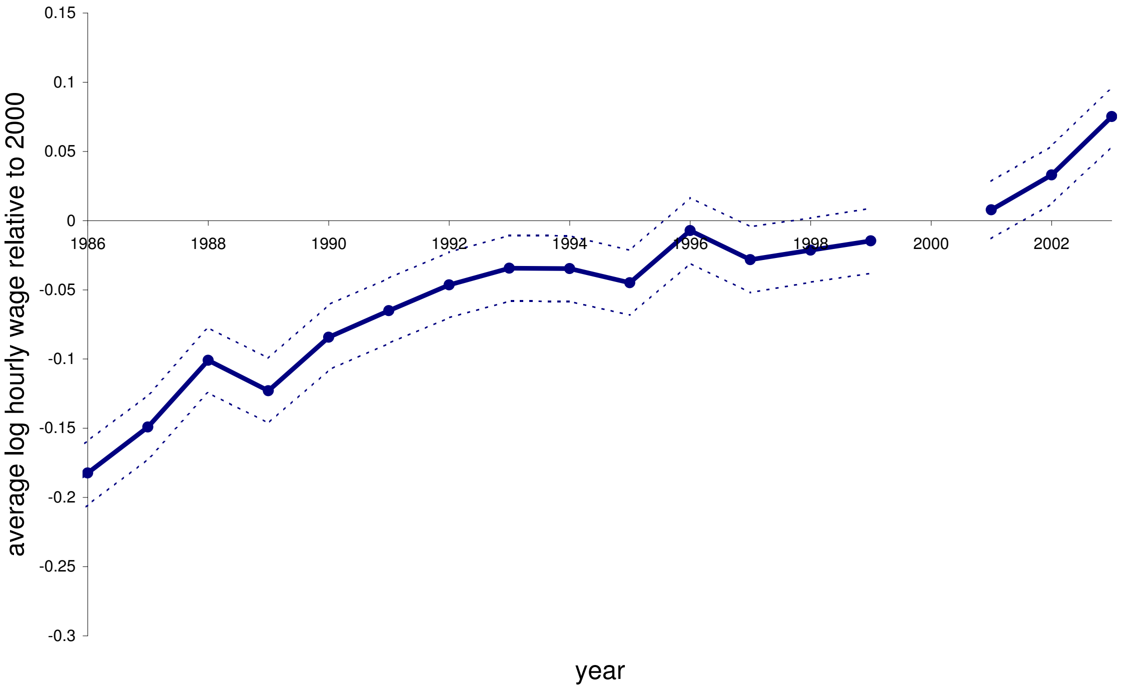


Notes: Estimates were obtained by regressing real log hourly wages on a full set of indicator variables for age and year. The dotted lines are pointwise 95% confidence intervals obtained from least squares standard errors.

Figure 1: Average real log hourly wage by age.

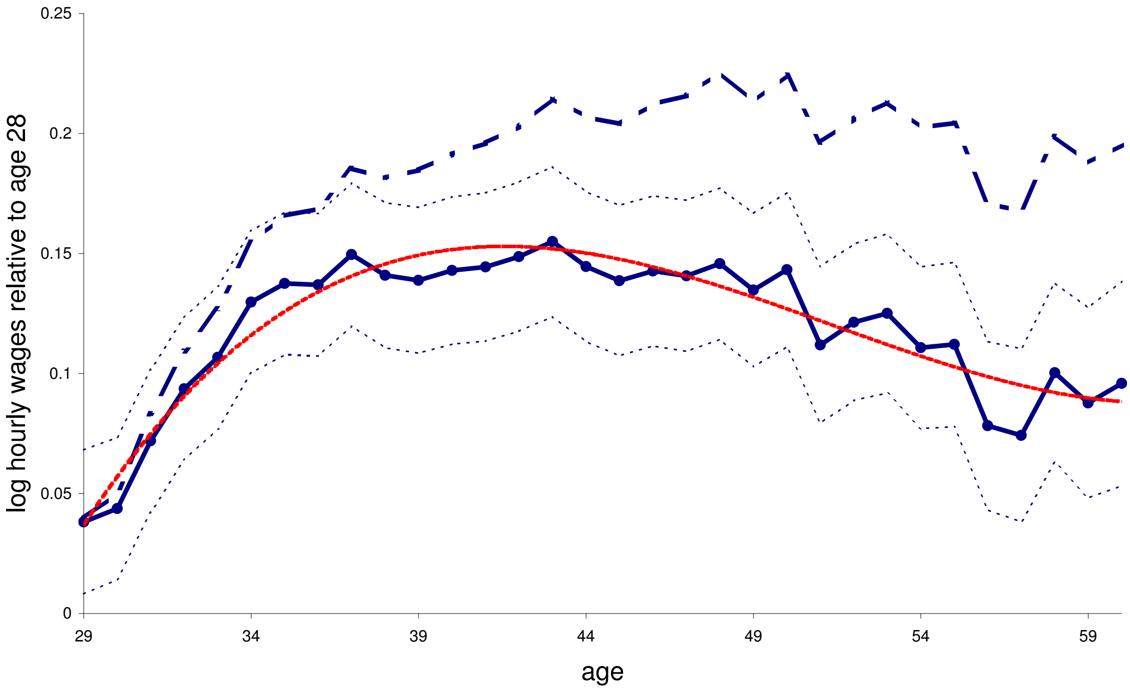
effect of age on wages is negative from the age of 43 on. This is because we now control for the wage growth which is associated with longer firm tenure. Figure 4, in turn, shows the partial effect of tenure on wages once we control for age differences.[11]

[11]We will refer to the solid lines in turn of Section 4 below.



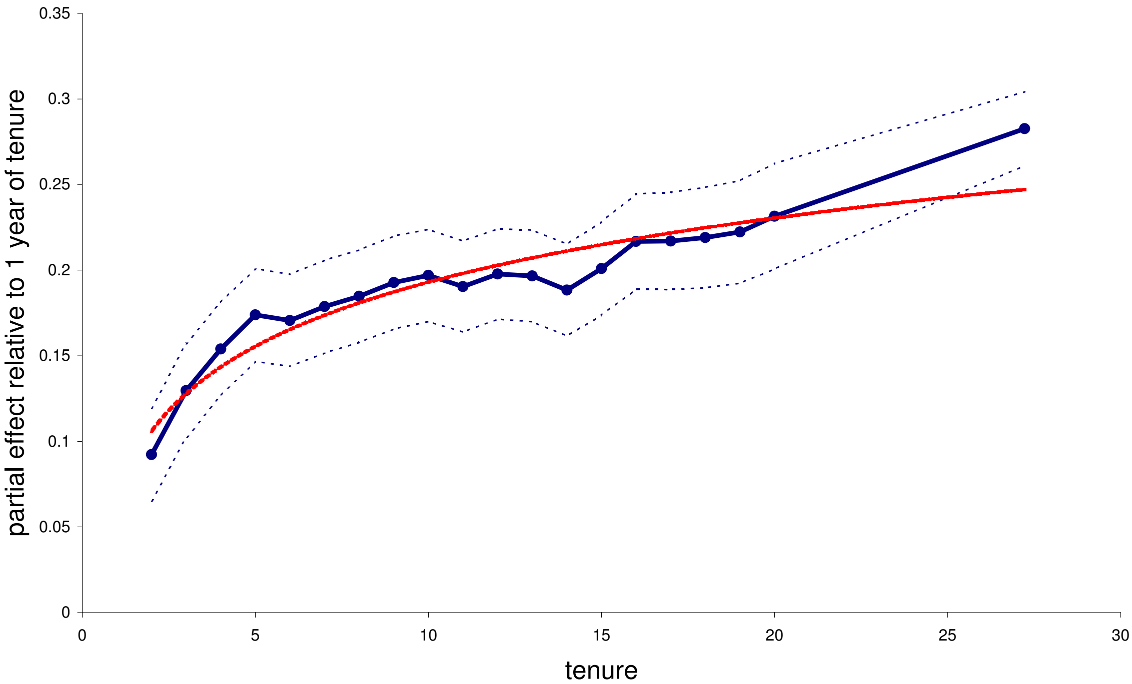
Notes: Estimates were obtained by regressing real log hourly wages on a full set of indicator variables for age and year. The dotted lines are pointwise 95% confidence intervals obtained from least squares standard errors.

Figure 2: Average real log hourly wage relative to year 2000.



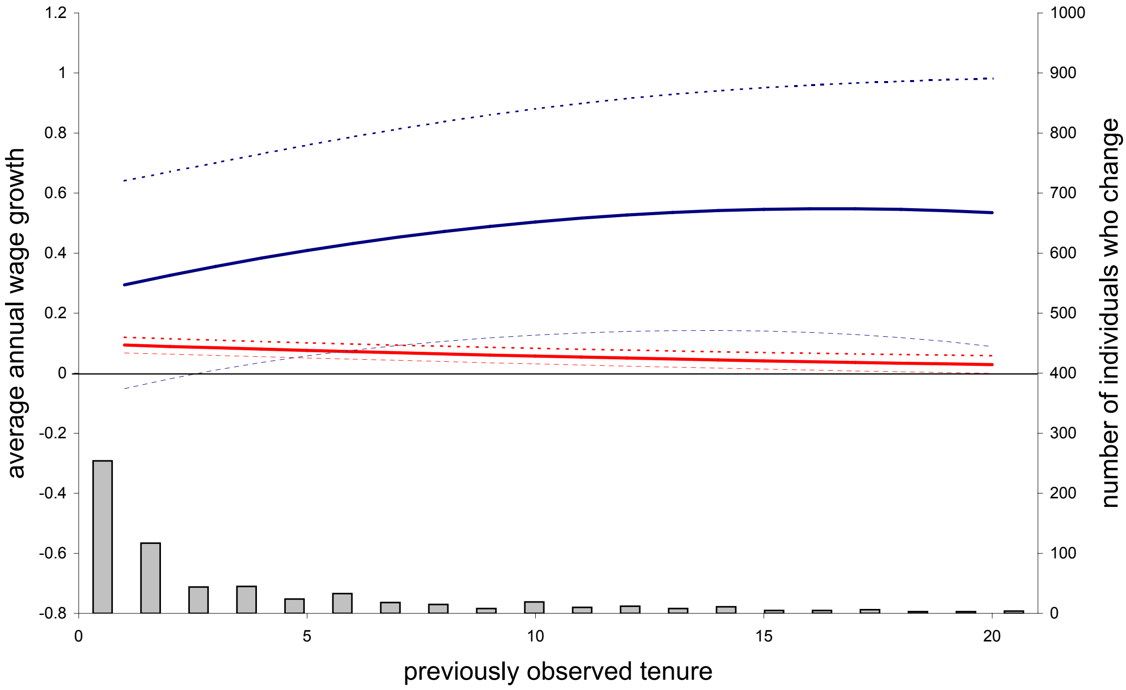
Notes: The dash-dotted line is the difference between the average real log hourly wage at a given age and the real log hourly wage at the age of 28. The connected dots are the partial effect of age on wages once we control for differences in tenure. Estimates were obtained by regressing real log hourly wages on a full set of indicator variables for age (and tenure for the connected dots) as well as region indicators and a linear time trend. The dotted lines are pointwise 95% confidence intervals. The solid line is a third order polynomial fit through the dots.

Figure 3: Partial and total effect of age on wages.



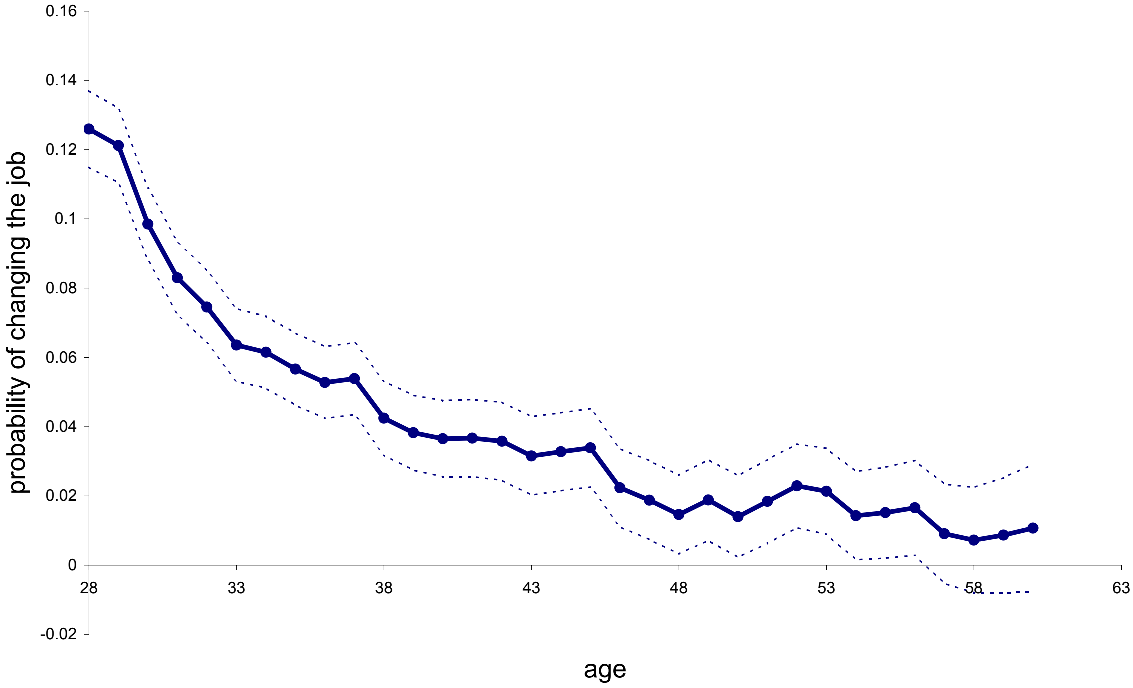
Notes: The connected dots are the partial effect of tenure on wages once we control for differences in age. Estimates were obtained by regressing real log hourly wages on a full set of indicator variables for age and tenure as well as region indicators and a linear time trend. The dotted lines are pointwise 95% confidence intervals. The solid line is a exponential fit through the dots.

Figure 4: Partial effect of tenure on wages.



Notes: The upper solid line is the estimated average annual wage growth for those who changed the firm (656 observations) plotted against previously observed tenure. The lower solid line is the estimated average wage growth for those who stayed in a firm (31,713 observations) against previously observed tenure. The dotted lines are pointwise 95% confidence intervals. The histogram shows frequencies of previously observed tenure for job changers. Estimates were obtained by regressing average annual wage growth on a second order polynomial for the respective subsample with previously observed tenure, controlling for age. See column (1) and (4) in Table 4 in the Appendix.

Figure 5: Annual wage growth by previously observed tenure.



Notes: Estimates were obtained by regressing an indicator for a firm change on age indicators. The dotted lines are pointwise 95% confidence intervals.

Figure 6: Probability of a job change by age.

Next, we investigate whether job changes constitute a source of wage growth in our data. Figure 5 contains estimates of the average annual wage growth for those individuals who changed the firm (inter-firm wage growth) and those who did not (intra-firm wage growth). Average wage growth is plotted against previously observed tenure, controlling for differences in age. It shows that the wage gain from changing jobs relative to staying in a given job is the higher the longer an individual has worked in a given firm. Clearly, job changes are endogenous as individuals only change jobs once such a change is advantageous. Therefore, we shall abstain from giving this association a causal interpretation.

The underlying estimates are contained in column (1) and (4) of Table 4 in the Appendix. Whereas column (2) and (5) are for comparison column (3) and (6) indicate that the difference between inter-firm wage growth and intra-firm wage growth is increasing in age as the coefficient on age is positive for those who change jobs and negative for those who don't. By a similar argument it is increasing in previous tenure, as we have already seen in Figure 5. Last, notice that there is a level effect as the previous log hourly wage enters negatively. This level effect if bigger for those who change jobs—however, wages of changers grow considerably more at the same time.

We conclude our descriptive analysis with Figure 6 which shows the probability to observe a job change for a given age which is constantly declining over the life span.

In turn of next section's empirical analysis, we implicitly control for endogenous firm changes by controlling for the endogeneity of tenure. Job changers should thereby, by the arguments made in turn of the discussion of the econometric model in Section 2, be characterized by a negative value of η_{ijt} once such a job change cannot be predicted by the observables in the reduced form equation, (2).

4. Specification and Results

Our data include a host of information about employers and their employees. For example, there is information on the size of the firm and on the type of position, high or low. Moreover, there are variables containing the number of years the employee was part time employed, full time employed, and unemployed. However, most of these variables are potentially endogenous. Therefore, and since we are interested in juxtaposing the returns to tenure, the returns to type, and the returns to general experience, we do not include the above mentioned variables in our regressions. Additional to the variables in Table 1 our set of variables includes indicator variables for the current state of residence (the "Bundesland").

Throughout, we estimate the first stage regressions by ordinary least squares and obtain fitted values of the residuals which we include in the second stage as a control function for unobservable factors confounding wages and tenure. We then compare these estimates to OLS and IV estimates. Standard errors thereof are obtained from $1,000$ bootstrap replications. As for the specification, we draw on the descriptive analysis in Section 3. The solid line in Figure 3 is a third order polynomial fit through the estimated coefficients of the indicator variables of the effect of age on wages conditional on tenure. Similarly, the solid line in Figure 4 is an exponential fit. As for formal testing, we can reject the null that the coefficients of the indicator variables on age indicators are jointly zero, with a P-value of 0.0027, once we include a second order polynomial in age into the regression of wages on indicator variables for age, tenure, a set of region indicators, and a linear time trend. However, once we include a third order polynomial in age, with a P-value of 0.8764 we cannot reject the null any more. Therefore, we feel that we should include a third order polynomial in age into the regression.

A similar result does not hold for the partial effect of tenure on wages. It turns out that the test statistic for the joint significance of the tenure indicators is 9.13, with a P-value of zero, once we include a second order polynomial in tenure following Altonji and Shakotko (1987) in addition to the above mentioned set of variables. However, once we include log tenure instead of a second order polynomial in such a regression, the test statistic is 4.84, also with a P-value of zero. Based on the fit in Figure 4 we nevertheless decided to go on with a specification in which we include a third order polynomial in age and the log of tenure into our regressions because this proceeding has one key advantage. It allows us to directly compare the returns to tenure, as estimated by OLS, to the estimates obtained from an Altonji and Shakotko (1987) type IV procedure, and the control function estimates obtained using the procedure outlined in Section 2.[12]

Reduced form estimates for the first stage are reported in Table 2. The first column is the reduced form for log tenure which was used for obtaining the Altonji and Shakotko (1987) type IV estimates for comparison.[13] Here, we replaced $\tilde{t}_{ijt}$ by $\log(t_{ijt}) - \overline{\log t_{ij}}$ because this newly constructed variable is by construction orthogonal to the fixed match specific component in the wage equation once we include log tenure instead of a linear and a quadratic term as a right hand side variable. Column (2) contains the reduced form for tenure which we use in the sequel in order to obtain the fitted residuals $\hat{\eta}_{ijt}$ for the control function estimates. Obviously, $\tilde{t}_{ijt}$ is strongly correlated with tenure. For comparison, since tenure is restricted to be at least equal to one, we report Tobit estimates. Respective coefficient estimates are very similar to the OLS estimates in column (2) so that we decided to base the subsequent control function estimates on the specification in column (2).

Notice that according to the reduced form, equation (2), those individuals with a high value of η_{ijt}, the fitted residual from the specification in column (2) of Table 2, are

[12]Instrumental variables estimation is only feasible if the number of instruments exceeds the number of endogenous variables. In our data, we did not manage to find enough instruments for a full set of tenure indicators. Moreover, estimating more flexible structures solely based on indicator variables turned out to impose not enough structure and resulted in imprecise estimates throughout which, apart from that, yield conclusions similar to the ones we reach below.

[13]Estimating the original specification with a second order polynomial in tenure resulted in qualitatively similar results.

	(1) Reduced Form log tenure	(2) Reduced Form tenure	(3) Tobit tenure
$\log(t_{ijt}) - \overline{\log t_{ij}}$	0.918**		
	(0.012)		
$t_{ijt} - \overline{t_{ij}}$		0.690**	0.718**
		(0.015)	(0.015)
age	-0.118**	-1.200**	-0.904**
	(0.032)	(0.326)	(0.340)
age sq.	0.004**	0.038**	0.032**
	(0.001)	(0.008)	(0.008)
age cu.	-0.000**	-0.000**	-0.000**
	(0.000)	(0.000)	(0.000)
R^2	0.32	0.35	

39,200 observations across individuals and time. The first two columns are OLS estimates. Standard errors in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. We also included a set of indicator variables for the state of residence as well as a time trend.

Table 2: Reduced form first stage regressions.

more likely to be of a type that stays longer in a given firm, as compared to the average. Importantly, and by construction, η_{ijt} includes individual and match specific factors since $\tilde{t}_{ijt}$ is uncorrelated with them by construction. Therefore, once we include fitted values for η_{ijt} as a control function in the structural wage equation, we can thereby not only control for the endogeneity of tenure but can, at the same time, assess the impact of high values of this control function on expected wages.

For the second stage estimates we implemented a standard generalized least squares (GLS) random-effects panel estimator, a similar two stage IV estimator in which we include the fitted tenure from the specification in column (1) of Table 2 as a regressor instead of observed tenure, and the control function estimator based on the residuals obtained from the specification in column (2) of Table 2.[14]

[14] For some individuals there is just one observation in our data. Therefore, a fixed effects model could not be implemented. In fact, it follows from our restrictions that the model should be estimated using pooled OLS since the observations are assumed to be independent across ijt. However, the stochastic restrictions stated in Section 2 are actually stronger than needed and stated in that way for the ease of the exposition. They can readily be relaxed.

	(1) GLS	(2) IV	(3) Structural I	(4) Structural II	(5) Structural III
log tenure	0.069**	0.045**	0.032*	0.069**	0.033**
	(0.005)	(0.007)	(0.014)	(0.008)	(0.011)
year	0.011**	0.012**	0.012**	0.011**	0.012**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
age	0.069**	0.084**	0.075**	0.068**	0.075**
	(0.019)	(0.018)	(0.021)	(0.019)	(0.019)
age sq.	-0.001**	-0.002**	-0.001**	-0.001**	-0.001**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
age cu.	0.000*	0.000**	0.000+	0.000*	0.000*
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\mathbb{1}\{\hat{\eta}_{ijt} \in (-\infty, -6)\}$				-0.008	
				(0.023)	
$\mathbb{1}\{\hat{\eta}_{ijt} \in [-6, -3)\}$				0.005	
				(0.020)	
$\mathbb{1}\{\hat{\eta}_{ijt} \in [-3, -1)\}$				-0.006	
				(0.018)	
$\mathbb{1}\{\hat{\eta}_{ijt} \in [1, 3)\}$				-0.013	
				(0.018)	
$\mathbb{1}\{\hat{\eta}_{ijt} \in [3, 6)\}$				-0.020	
				(0.020)	
$\mathbb{1}\{\hat{\eta}_{ijt} \in [6, \infty)\}$				-0.023	
				(0.028)	
$\hat{\eta}_{ijt}$			0.010**		0.009**
			(0.003)		(0.002)
log tenure $\times$ $\hat{\eta}_{ijt}$			-0.002*	0.001	
			(0.001)	(0.001)	
log tenure $\times$ $\mathbb{1}\{\hat{\eta}_{ijt} \in (-\infty, -6)\}$					0.020*
					(0.010)
log tenure $\times$ $\mathbb{1}\{\hat{\eta}_{ijt} \in [-6, -3)\}$					0.020*
					(0.009)
log tenure $\times$ $\mathbb{1}\{\hat{\eta}_{ijt} \in [-3, -1)\}$					0.007
					(0.008)
log tenure $\times$ $\mathbb{1}\{\hat{\eta}_{ijt} \in [1, 3)\}$					-0.011
					(0.008)
log tenure $\times$ $\mathbb{1}\{\hat{\eta}_{ijt} \in [3, 6)\}$					-0.013+
					(0.008)
log tenure $\times$ $\mathbb{1}\{\hat{\eta}_{ijt} \in [6, \infty)\}$					-0.020*
					(0.009)
Observations	39200	39200		39200	39200
Number of persnr	6831	6831		6831	6831

Bootstrapped standard errors from 1,000 (100 for column (4)) replications in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. The dependent variable is the real log annual wage in prices of 2000. We also included a set of indicator variables for the state of residence as well as a time trend.

Table 3: Second stage estimates.

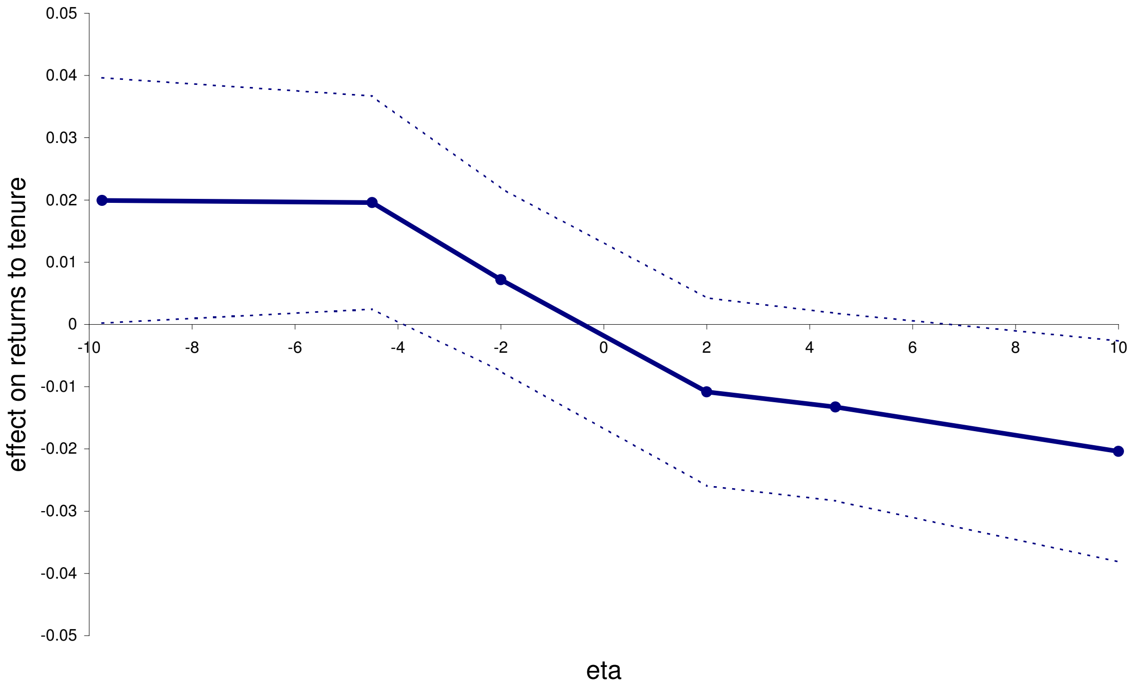


Figure 7: The impact of η_{ijt} on the returns to tenure.

The GLS estimates in column (1) of Table 3 have the interpretation of observed differences and derivatives in the direction of the covariates. The observed wage increase from a duplication of tenure is about 7 per cent. IV estimates in column (2) are lower and, compared to GLS estimates, attribute wage increases to a larger extent to increases in age. A similar observation is made in the analysis of Altonji and Shakotko (1987) and discussed by Topel (1991) who argues that such estimates are misleading. In light of the modern literature on program evaluation, an important aspect in such a discussion is the heterogeneity of the effect of tenure on wages. It can be argued that the instrument uncovers the local average treatment effect of Imbens and Angrist (1994), the difference in wages of those who are induced by the instrument, $\log t_{ijt} - \overline{\log t_{ij}}$ to stay in a given job. In general, it does not uncover the population average effect of tenure on wages which is sought to be estimated in this paper.

Column (3) through (5) contain control function estimates built on the econometric model laid out in Section 2 which do, under the assumptions made, in fact uncover the population average effect, as shown in Section 2. Column (3) contains our baseline specification. Column (4) contains a specification in which we include indicator variables for $\hat{\eta}_{ijt}$ in order to get an idea about the impact of the type on wage levels without imposing too much structure. It turns out that the set of indicator variables is jointly insignificant with a P-value of 0.9218. We therefore like to think of the specification in column (4) as a robustness check confirming the specifications in column (3) and (5).[15]

Notably, our most preferred specification in column (5) indicates that wages are estimated to be increasing in the value of $\hat{\eta}_{ijt}$ which means that those types who stay longer in a firm have higher wages on average. At the same time, the returns to tenure are estimated to be decreasing in $\hat{\eta}_{ijt}$.[16] The effect, which is depicted in Figure 7, is that those individuals with high values of $\hat{\eta}_{ijt}$ have lower returns to tenure on average.[17]

Taking the robustness check in column (4) aside, we interpret our results as evidence for population average returns to tenure to be lower than observed returns to tenure. We estimate the effect from doubling firm tenure to be a real wage increase of about 3 per cent. Average wages increase by about 1 per cent per year for those types who stay longer than expected in a given firm. At the same time, as shown in Figure 7, they exhibit lower returns to tenure.

[15] We experimented with several specifications. It turned out that more flexible specifications did not reveal a specific pattern. For example, we included a third order polynomial in $\hat{\eta}_{ijt}$ interacted with log tenure instead of the set of indicator variables interacted with log tenure in column (4).

[16] The interaction terms between $\hat{\eta}_{ijt}$ indicators and log tenure are jointly significant with a P-value of 0.0187.

[17] See Topel (1991) for an explanation why highly productive employees may stay shorter at firms.

5. Concluding Remarks

In this paper we have analyzed wage tenure profiles for different types of employees. Using a flexible control function approach, we have disentangled returns to observable tenure from returns to an unobservable scalar type variable and have shown that there exist interaction effects between the two. In particular, we find that stayers obtain flatter wage profiles in tenure than quitters. However, on average, job seniority yields higher wages. Considering the theoretical literature, the obtained results suggest that the employees' type variable incorporates not only the workers' productivity. Hence, the interpretation of the error term in the wage equation seems to be not clear-cut.

In future work, we plan to incorporate information on the individuals' undergraduate majors into the analysis. Furthermore, we plan to exploit the panel structure of our data in a more efficient way.

Appendix

	(1) change	(2) change	(3) change	(4) no change	(5) no change	(6) no change
previous tenure	0.035*	0.012*	0.017**	-0.005**	-0.003**	-0.001*
	(0.015)	(0.006)	(0.006)	(0.001)	(0.000)	(0.000)
previous tenure sq.	-0.001+			0.000**		
	(0.001)			(0.000)		
age	0.008+	0.007	0.008+	0.001*	0.001*	0.001**
	(0.005)	(0.005)	(0.005)	(0.000)	(0.000)	(0.000)
previous log hourly wage			-0.375**			-0.277**
			(0.067)			(0.006)
Constant	0.261	0.339*	1.218**	0.098**	0.083**	0.794**
	(0.178)	(0.172)	(0.230)	(0.013)	(0.012)	(0.019)
Observations	656	656	656	31713	31713	31713
R-squared	0.02	0.01	0.06	0.00	0.00	0.07

Standard errors in parentheses. + significant at 10%; * significant at 5%; ** significant at 1%. Estimates were obtained by regressing annualized wage growth, for the respective subsample, on previous tenure, its square in column (1) and (4), age, and the previous log hourly wage in columns (2), (3), (5), and (6). The annualized wage growth is defined as the difference between the log wage in t and t' over $t - t'$ for all $t > t'$.

Table 4: Wage growth by previous tenure.

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