Occupational Pensions and Job Mobility in Germany

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Abstract

This paper provides an empirical analysis of the role of occupational pensions for voluntary job mobility in Germany. The analysis is based on individual data from the German Socio-Economic Panel for the period 1995-1998. We estimate the effects of occupational pension coverage and different measures of pension portability loss on voluntary job changes using a sample selection model with bootstrapping. The main findings are that occupational pension coverage reduces worker mobility by imposing a capital loss on those leaving their job early. A higher compensation in pension-covered jobs is important too.

Keywords: Labour mobility, Occupational pensions
JEL-Code: C35, J31, J32, J63, J68

1 Introduction

The role of occupational pensions both for domestic and cross-border mobility is high on the agenda of the European Union. Workers covered by occupational pensions typically suffer losses in pension rights when changing employer. According to the European Council these losses are an obstacle to mobility (Council of the European Union 2003, 88-9). Concerns about hampered mobility have already motivated reforms in occupational pension scheme legislation over the last years in many EU countries, including Germany. Further reforms are currently under way. These generally aim at

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reducing the portability loss associated with leaving a pension scheme before retirement.

However, the empirical evidence of a mobility-deterring effect of portability loss is neither ample nor entirely conclusive. Several decades of research on labour mobility have established that turnover in jobs covered by occupational pensions is indeed lower than in other jobs. The reasons for this are less clear. Three explanations dominate in the empirical literature. Occupational pensions may reduce mobility by imposing a capital loss on those who change jobs. Likewise, pension-covered workers often receive a compensation premium which discourages mobility. Finally, workers who prefer stable employment may sort into jobs covered by pensions and are thus unlikely to change employer. As an alternative explanation Ippolito (2002) proposes that staying is the result of savers who are better, and therefore better paid, workers than non-savers selecting into pension-covered jobs. Research into these hypotheses has so far been supplied for the United States and the United Kingdom. However, there are very few empirical studies of other European countries. This paper is a contribution to filling this gap and providing an empirical basis for reform considerations.

The portability loss suffered by mobile workers depends on the portability options defined by pension regulation. Workers leaving an occupational pension plan before retirement age are usually entitled to a pension only after having completed a vesting period. If they leave before the vesting period is completed, they loose all accrued benefits. Workers whose benefits have become vested are entitled to a deferred retirement pension. Here a real capital loss may occur if the pension is defined benefit. In defined benefit plans, the deferred retirement pension is often based on nominal earnings at the point of job change or on some fixed sum accrued with each year of service. If these benefits are not price or wage indexed, their value erodes over the time until workers are eligible for retirement benefits. Germany has only recently reduced the period until accrued pension benefits become vested from ten to five years to improve portability. Deferred benefits are not indexed, however, and since the vast majority of occupational pension plans are traditionally defined benefit, capital loss is an important issue in Germany.

This paper is one of the first to examine the relationship between occupational pensions and voluntary job mobility in Germany. Using the German Socio-Economic Panel (GSOEP), waves 1995-1998, we estimate the effects of occupational pension coverage and different measures of pension portability loss on voluntary job changes using a sample selection model. In contrast to other studies we distinguish between capital loss of vested and non-vested benefits to account for the long vesting periods in Germany. The indiviual mobility decision is modeled as a function of earnings differentials and mobility costs. We find that pension-covered workers receive a wage premium and are less mobile than other workers. Moreover, portability loss discourages job changes both for workers with vested and non-vested benefits. This finding suggests that reducing the period until benefits become vested may be an effective policy option to enhance mobility. However, it seems equally important to introduce an indexation of preserved benefits.

The paper is structured as follows. Section 2 discusses the findings in the literature about occupational pensions and mobility, briefly describes the pension system in Germany and presents the framework for the calculation of pension portability loss. Section 3 develops the estimation approach taken in this study. Section 4 describes the data used. Section 5 presents the estimation results and section 6 concludes. The appendix provides further information on the data set and the estimation results.

2 Occupational Pensions and Mobility

2.1 Literature

Most studies on job mobility have as a starting point the assumption that differences between wages in current and alternative jobs are the driving force behind changing jobs, and pension portability losses as well as other costs discourage changes of employer. A first generation of papers estimated quit or job change equations which included dummies capturing pension information as well as variables approximating potential wages and/or pension benefits in current and alternative jobs as regressors, e.g. Schiller and Weiss (1979), Mitchell (1982). These early studies for the U.S. found strong and significant evidence of pensions deterring worker mobility, although it was not always possible to relate this effect to specific pension plan characteristics. Two German studies in this vein come to somewhat differing results. They find no effect of pensions plans on voluntary quits (Frick et al. 1999), but a negative effect on voluntary and involuntary firm exits (Schnabel and Wagner 1999).

Some recent studies have tried to capture the effect of occupational pensions on mobility more precisely by explicitly modeling the capital loss incurred by pension-covered mobile workers and thus supplementing the simple dummy-variable approach. This extension of the analysis is often based on the work of Ippolito (1985) who models pension portability loss in the framework of an implicit pension contract. We also adopt this approach which we develop below. In contrast to our paper, most authors concentrate on the capital loss of vested workers. We extend the analysis to non-vested workers to account for the long vesting period in Germany. The studies including measures of capital loss come to differing results. For the U.S., Allen et al. (1993) find a sizable effect of capital loss for layoffs rather than quits, Gustman and Steinmeier (1993) find the effect to be significant but small for job separations, and Andrietti and Hildebrand (2001) find no effect of their capital loss variable on job change. For the U.K., McCormick and Hughes (1984) find a considerable effect of their proxies for capital loss on job separations, while no such effect is apparent in Andrietti (2001 and 2002). A possible explanation is that inflation indexation of deferred benefits - having a huge impact on the magnitude of capital loss - was introduced after the early U.K. study was carried out. However, another post-indexation study does find significant effects (Henley et al. 1994). A comparative European study (Andrietti 2001) shows significant effects of portability loss only for Ireland. It was the only country with no indexation of benefits of early leavers at the time studied.

Many of the more recent papers have looked deeper into the possible causal relationships between pension coverage and mobility and can further account for these somewhat inconclusive results. In addition to portability losses which may cause workers to refrain from changing jobs, two further explanations for the negative relationship between pensions and mobility are being discussed. The first argues that pension covered workers may receive a compensation premium which discourages mobility. Thus it may not be the portability loss (alone) which lowers the mobility rate of pension covered workers. Evidence in favor of this argument is supplied by Gustman and Steinmeier (1993), for example, who show that pension covered workers risk wage losses when they change jobs. Higher compensation in pension-covered jobs may be a result of 'savers' selecting into such jobs. According to Ippolito (2002) savers are better workers and therefore receive higher wages. In this paper we also find clear evidence of the existence of a wage premium for pension-covered workers.

The second argument centers on the possibility that workers who prefer stable employment sort into jobs covered by pensions. These individuals may have a low rate of subjective time discounting and a preference towards provision for old age. Allen et al. (1993) find evidence for sorting on observables: men, whites, union members and those who are married are more likely to sort into pension covered jobs. Based on U.S. data they do not find evidence for sorting on unobservable characteristics. Likewise, the study of U.K. data by Mealli and Pudney (1996) which analyses transitions between pension-covered and non pension-covered jobs find little evidence of sorting due to unobservable characteristics. There is recent evidence of sorting into pension scheme types depending on mobility characteristics in the U.K. (Disney and Emmerson 2002). In our study we control for observables like being married, home ownership, and having dependent children which are taken to proxy preferences for stability.

Methodologically, some studies have made attempts to explicitly incorporate wages in alternative jobs into the analysis of the mobility decision. Since wages in alternative jobs can only be observed for movers, and wages in current jobs can only be observed for stayers, ideally two counterfactual outcomes have to be estimated. McCormick and Hughes (1984) use job satisfaction as a measure to approximate wage and non-wage benefits in alternative employments. In another approach Gustman and Steinmeier (1993) equate the actual wage observed for movers in their new jobs to the (non-observable) alternative wage. This method does not account for the bias arising from selection of workers into stayers and movers, however. A solution to the selection problem commonly used in mobility studies is to employ a sample selection correction (Heckman 1979; Maddala 1983). This approach is chosen, for example, by Mitchell (1983), Andrietti (2001), and Disney and Emmerson (2002), who use a Heckman procedure to produce consistent wage estimates which then enter structural probit equations explaining the mobility choice.

2.2 Occupational Pensions and Portability in Germany

Within the German three-tiered pension system the public first tier is most important. It is mandatory for all employees except civil servants and most self employed. The benefits are earnings-related and provided for 84% of old age incomes in 1999 (cf. Deutscher Bundestag 2001). The second tier consists of public and private occupational pensions which may be offered by employers to supplement the first tier. The third tier includes all forms of private provision like personal pension plans.

In 1999, 21% of German pensioners received a public or private occupational pension which made up approximately 25% of their total monthly income (Euro 319 on average) (Deutscher Bundestag 2001). Occupational pensions are delivered by many different systems, and there are few data sources which provide summary statistics about pension plan characteristics, workers covered by occupational pensions, and benefits accrued.

Pension portability in Germany is rather restrictive compared to the U.S. and to other EU countries. Until recently, the Law to Improve Occupational Pensions (BetrAVG, 1974), arranged that employee pension benefits had to be vested only after a period of 10 years and a minimum age of 35 for early leavers or, alternatively, after 12 years of firm tenure and 3 years of contributions to a scheme with a minimum age of 35. Since the beginning of 2001, contributions to occupational pension schemes are vested after 5 years with a minimum age of 30 years. This vesting period is still longer than in some other European countries (U.K. 2 years, Netherlands 1 year).

In contrast to the U.S., many EU countries have introduced a price or wage indexation of preserved benefits. In Germany, however, there is no legal requirement to index preserved benefits under a defined benefit scheme. Since defined benefit plans prevail in Germany, most workers face real capital losses when leaving pension-covered jobs before retirement age.

Arrangements which allow a transfer of accrued pension rights to new employers' pension schemes are another important portability issue. Again, the German case is comparatively restrictive. Before the 2001 reform portability of accrued benefits was only possible for life insurance retirement bonds (Direktversicherung), and for industry-wide pension plans when changing employer within the same industry. However, the latter existed only in the public sector and the construction industry. The vast majority of German workers covered in a scheme were members of a company pension plan and had no possibility of transfer. Some recent changes in the legal regulation aim at facilitating a transfer of benefits in the future. This endeavor may be supported by the creation of more industry-wide pension plans and more defined contribution plans which we can observe at present.

In summary, capital loss due to lack of indexation of pension rights - the major source of portability loss in the United States - is presently also relevant in Germany. Moreover, in the time period studied in this paper, we can assume considerable portability loss due to the long vesting period of ten years.

2.3 Portability Loss

The framework of an implicit contract between the worker and the firm (Ippolito 1985) is useful to model the capital loss imposed on those who leave jobs early. According to this approach, workers evaluate the package of wage and pension benefits when considering a career with a firm. They forego a portion of their wage throughout their work lives in exchange for a pension at retirement. The workers' implicit pension contributions are equal to the present value of expected pension benefits ("stay pension"). Assume for simplicity that worker i survives to retirement age R with certainty and that his pension benefit is already vested and given in the form of a lump sum (PB_i). The benefit is based on the pension formula $PB_i = bS_iY_i(t)$, where b is a constant reflecting the annual accrual rate of the pension plan, S_i are the years of service (firm tenure), and $Y_i(t)$ are the individual earnings at time t. The stay pension at $t < R$ is

$$SP_i = bS_iY_i(R)e^{-r(R-t)}, \tag{1}$$

where $Y_i(R)$ is expected final earnings and r is the inflation rate. If the worker expects his level of compensation to increase with time at the rate g, we can equivalently formulate the stay pension as a function of current earnings:

$$SP_i = bS_iY_i(t)e^{(g-r)(R-t)}. \tag{1'}$$

A worker leaving the firm prior to his pension age R receives only the present value of his "leave pension" based on his current earnings, $Y_i(t)$:

$$LP_i = bS_iY_i(t)e^{-r(R-t)}. \tag{2}$$

The capital loss is the difference between the stay and the leave pension (1')-(2) for vested workers,

$$CL_i^v = bS_iY_i(t)e^{-r(R-t)}\left(e^{g(R-t)} - 1\right). \tag{3}$$

According to the implicit contract the worker pays for a pension which is indexed with an expected wage path. If he quits, he receives only a nominal pension.

Now consider the case of a worker who leaves the firm at time t^l before the time his accrued benefits are vested ($t^l < t^v < R$). This individual loses the present value of the entire pension capital accumulated up to date. The impact can be assessed by assuming the leave pension to be zero because the mobile worker can expect no pension benefit upon retirement in this case. Thus

$$CL_i^{nv}|t^l < t^v = bS_iY_i(t^l)e^{(g-r)(R-t^l)}. \tag{4}$$

For the empirical analysis we represent the capital loss of vested and non-vested workers by different variables. For one, we directly compute the capital loss functions for vested and non-vested workers, (3) and (4), using the available wage, tenure and age data, and assuming $g = 3.9\%$ and $r = 2.3\%$ which are averages over the period 1985-94. We take the unknown constant b to be $1/60$ which is a value often assumed in other studies.

Figure 1 shows how capital loss of vested and non-vested benefits, CL_i^v and CL_i^{nv}, evolves over time until the retirement age R for an individual who starts to work for an employer at age 20 and has a constant income of $Y_i = 61,200DM/year$. CL_i^{nv} increases with years of service. If wages grow at the same rate as the interest rate ($g = r$), the function increases linearly with years of service, if $g > r$, it is a concave function. CL_i^v is a concave function which is zero at $S_i = 0$ and reaches zero again as the worker approaches retirement age ($R - t = 0$).

- figure 1 about here -

As an alternative representation of capital loss of vested benefits we use an approximation introduced by McCormick and Hughes (1984) who employ the interaction of tenure, S_i, and years to retirement, $R-t$. This approximation captures the functional form of CL_i^v and imposes less assumptions about the pension plan type than direct computation.[1] For non-vested benefits we use years of service, S_i, as an approximation.

3 Model and Estimation Procedure

The individual mobility decision is modeled as a function of earnings differentials and mobility costs such as pension portability losses. A worker will change jobs if the life time earnings gain from moving into a new job exceeds the mobility costs, i.e. if

$$I_i^* \equiv (Y_{mi} - Y_{si}) - C_i > 0, \qquad i = 1, \ldots, n \tag{5}$$

where Y_{mi} is the expected present value of lifetime earnings in a best alternative job, Y_{si} is the expected present value of lifetime earnings in the current job, and C_i is the present value of mobility costs.

In the empirical specification of this model we assume for simplicity that the log of current wages is the best predictor of the log of lifetime earnings. However, we can only observe the wages of movers (w_m) and of stayers (w_s), respectively. The counterfactual, that is the alternative expected wage, is not observable. Likewise, the mobility costs are not directly observable.

[1] In another pension plan type known in Germany, the pension benefit is given by the pension formula $PB_i = fS_i$, where f represents a fixed benefit accrued with every year of service and usually subject to some kind of indexation. f may or may not vary between employees depending on earnings levels or occupational status. An early leaver will suffer capital loss because his deferred benefits are not subject to indexation. Capital loss in this pension plan can be represented by (3) for vested and (4) for non-vested workers by substituting $bY_i(t)$ by f and interpreting g as any indexation rate applied in the plan. Unfortunately, the data set used in this analysis does not allow a distinction of pension plan types by individual, i.e. whether the plans offer a proportion of the final incomes or a fixed sum for every year of service as benefit. Moreover, there is no information available on the relative importance of the two pension formulae. To test the empirical specification, we also calculate the capital loss for the fixed benefit pension plan, omitting the constant f. Estimates using this measure of capital loss yield very similar results to those displayed in section 5 using (3) and (4). They are available from the author on request. The interaction terms which approximate the capital loss functions have the advantage that they capture both pension plan types.

Hence we cannot observe the actual gain from mobility, I_i^*, but only a binary random variable I_i which we can define as

$$I_i = \begin{cases} 1 & \text{if } I_i^* \geq 0 \\ 0 & \text{otherwise} \end{cases} \tag{6}$$

Assuming that mobility costs are determined by a vector of exogenous personal and job specific variables, Z_i, and interactions of these variables, we can describe (5) as a structural probit model such that:

$$I_i^* = \gamma(\ln w_{mi} - \ln w_{si}) + \delta' Z_i + u_i \tag{7}$$

and

$$u_i \sim N(0, \sigma_i^2)$$

To complete the model, it is necessary to estimate the wage differential. The wage differential can be predicted by estimating separate wage equations for movers and stayers

$$\ln w_{mi} = \beta_m' X_i + \varepsilon_{mi} \tag{8}$$

$$\ln w_{si} = \beta_s' X_i + \varepsilon_{si} \tag{9}$$

where X_i is a vector of human capital and personal variables including experience and its square, job tenure as well as dummy variables representing sex, occupational degree, and residence in East Germany to capture the considerable wage differentials between East and West Germany. A further dummy for occupational pensions controls for a wage premium in pension-covered jobs. We derive robust standard errors for clustered data based on the Sandwich estimator (Woolridge 2002) to account for the panel structure of the data.

Separate estimates of the wage equations yield inconsistent parameter estimates if stayers differ in observed and unobserved characteristics from movers. One conventional way of overcoming this selectivity problem is by using a Heckman (1979) two stage procedure. This procedure begins by estimating a reduced form probit equation which contains all variables from X_i and Z_i and examines the effect of individual characteristics on the selection into movers and stayers. Inserting equations (8) and (9) into (7) yields:

$$\begin{aligned} I_i^* &= \gamma(\beta_m' X_i - \beta_s' X_i) + \delta' Z_i + u_i \\ &= \vartheta' W_i + \nu_i \end{aligned} \tag{10}$$

where $\vartheta' = [\gamma(\beta_m' - \beta_s'), \delta']$, $W_i = [X_i, Z_i]$, and $\nu_i = (\gamma(\varepsilon_{mi} - \varepsilon_{si}) + u_i)$. This can be used to calculate the values for the selectivity terms which are the inverse Mills ratios, namely:

$$\lambda_{mi} = \frac{-\phi(\hat{\vartheta}' W_i)}{\Phi(\hat{\vartheta}' W_i)} \qquad \text{if } I_i = 1 \tag{11}$$
$$\lambda_{si} = \frac{\phi(\hat{\vartheta}' W_i)}{1 - \Phi(\hat{\vartheta}' W_i)} \qquad \text{if } I_i = 0$$

$\phi(.)$ being the standard normal density function and $\Phi(.)$ the corresponding cumulative distribution function. This method assumes that the errors in (8), (9), and (10) have a trivariate normal distribution. The selectivity correction terms are included in the wage equations, which are in turn used to predict earnings for stayers and movers. Finally, the difference in predicted income allows estimating the structural probit (7) using maximum likelihood. We use bootstrapping to correct the standard errors for the two-stage procedure. To account for the panel structure of the data we resample individuals instead of observations. Identification is addressed by including more than one variable in the vector Z_i which is not included in X_i. In particular, mobility costs are modeled as a function of a dummy representing the pension status (model 1) as well as of the measures of capital loss described in subsection 2.3 (models 2 and 3). Other variables well known to influence mobility costs include age and house ownership: older workers and home owners are less mobile than younger house tenants. Personal variables like being married and having children may proxy high preferences for a stable job; we can also assume that persons who choose home ownership do so because they tend to be immobile. Further account and justification of the choice of variables is given in the following section.

4 Data Description

The analysis is based on the German Socio-Economic Panel (GSOEP), an annual longitudinal survey of private households in Germany. The GSOEP started in 1984 with interviews in 5,921 households with residence in West Germany. In 1990, another 2,179 households with 4,453 persons from East Germany were added to the panel. Our sample includes the West and East German sub-samples.

We concentrate on the mobility decisions of full-time employees ($\geq$ 35 hours per week), excluding civil servants, self employed and apprentices. Mobility is defined as the first voluntary quit of job and take-up of a new full-time job between annual interviews with or without intervening unemployment. We assume that intervening unemployment after a quit decision is voluntary. In the GSOEP information about employers' provision of occupational pension schemes was asked annually from 1985-1992 and again in 1995. The respondents were asked whether their company offered an occupational pension plan, and - if yes - whether the respondent personally accumulates pension benefits. Unfortunately, there is no information on the type of pension plan or its generosity. This study makes use of the most recent pension information, looking at the mobility decisions of 1995 job holders with and without occupational pensions. Until the time of the next interview in 1996 we observe less than 100 mobile job holders under our mobility definition. Therefore we construct a pooled sample which covers the mobility between 1995-6, 1996-7, and 1997-8. Workers are dropped from the sample after their first job change or when they exit full-time employment.[2] The sample consists of 8,361 observations of which 3,356 are 1996 job holders, 2,756 are 1997 job holders and 2,259 are 1998 job holders. In the whole time span 1995-8 we observe 181 voluntary job changes.

In the longitudinal dataset all variables except pension status, sex, and occupational degree are treated as time varying. Information on mobility costs is based on personal and job characteristics in the year prior to mobility. The earnings equations make use of post-mobility wage, human capital and personal information with the exception that job tenure refers to tenure at the last job for mobile workers. Post-mobility wages are deflated by the German CPI to values of 1995, and implausible cases (13 cases of monthly wages below 1.000 DM, 3 cases of unlikely high wages) as well as cases with missings on any of the variables are deleted. Descriptive statistics of all variables are supplied in the appendix.

The dependent variable of the wage equations is the log of monthly gross wages in the last month before the interview. This variable includes overtime compensation but no other extra payments such as leave pay. Since we look at full-time employees only, there is no need to compute hourly wages where the choice of denominator (hours actually worked vs. hours according to contract) strongly influences the results. Regressors are years of tenure and a dummy for sex. Labor market experience and its square refers to real rather than potential experience. The variables are constructed using spell data, where experience is defined as time spent in full- or part-time employment.

[2] Individuals exiting full-time employment may have different moving probabilities than those not exiting. We have estimated a model including a correction for selectivity in exits, employing regional unemployment rates, years to retirement, gender, and its interaction with the number of dependent children in the household as instruments. We found no significant selection effects, and the results of the model remained the same. Hence we interpret exits from full-time employment as censored responses. Outcomes are treated as a discrete time duration.

We construct 3 dummy variables for the highest formal occupational degree. The reference category codes workers without a completed occupational degree, *vocational degree* stands for workers with a completed apprenticeship or vocational training, and *college/university* codes persons holding a university or technical college degree. A dummy for pension-covered workers (*occupational pension*) is entered into the wage equations to test for a wage premium in their jobs. Furthermore we include a dummy for residents in East Germany to account for the considerable East-West wage differentials. Finally, we use separate dummy variables for each period of observation to control for time-varying factors such as wage rises of persons who change jobs later. We are not able to control for the savings propensity of individuals as proposed by Ippolito (2002) because data on saving activities is are not available in the GSOEP.

Although we could improve the goodness of fit of the wage equations by including further job-specific variables such as occupational status, firm size, and industry, we refrain from doing so. The role of the wage equations in the estimation procedure is to estimate movers' wages for stayers and vice versa. It is not plausible to assume identical job characteristics in the counterfactual situation, as some studies do. Job characteristics of the old job do enter the structural probit as a measure of mobility costs as explained below. By including a pension dummy in the wage equation we assume that each individual would have the same pension status in the counterfactual state. We discuss the implications of this assumption with the empirical results.

Mobility costs are modeled as a function of the pension variables. We employ a pension coverage dummy (model 1) and the capital loss variables described in subsection 2.3 (models 2 and 3). Since the data includes no information on the vesting status of pension rights, we code the pensions of persons aged 35 years and older with a minimum firm tenure of 10 years as vested. Unvested pension rights are coded for persons who are either younger than 35 or have a firm tenure below 10 years.

The choice of further variables modeling mobility costs is guided by the standard results of the mobility literature. Among job-specific variables we include occupational status, firm size and industry in the last job to capture job-specific mobility costs over and above the estimated wage differentials. The mobility literature usually assumes that mobility from large firms is more costly because they offer more and better career opportunities than small firms do. Career opportunities also differ by industry. The occupational status proxies the ability to perform efficient job search, a higher status implying lower transaction costs.

Among the personal variables we include age, home ownership, marital status, and number of children under the age of 16 in the household. Older workers are less mobile than younger workers. A possible explanation is that younger workers benefit from wage increases for a longer time. Thus we include age to capture differences in the pay-off period from mobility between older and younger workers. Home owners are usually less mobile than tenants. This may be because of the costs involved in selling and purchasing property when job changes also involve regional mobility. These costs are particularly high in the German housing market due to considerable taxes on property purchases and fees for property registration. Therefore we can assume that home ownership proxies high preferences for a stable job. We use this variable as well as the marital status and the number of children to control for personal preferences towards mobility vs. stability. [3] We also include dummies for the period of observation to control for possible time-varying effects. Finally, we use a gender dummy to represent the differing mobility behavior of men and women. The ideal solution to estimate separate models for men and women is not feasible because there are too few cases of mobility in the data set.

5 Empirical Results

Before presenting the estimation results we discuss descriptive evidence on occupational pensions, job mobility, and wages. Table 1 shows that mobility among 1995 job holders is low independent of occupational pension status. The average mobility rate over the three year period between 1995 and 1998 is 2.2%. Individuals without pension-covered jobs in 1995 (3/4 of the sample) are two times more likely to change jobs than are individuals on jobs covered by pensions (2.6% vs. 1.3%). This result confirms the findings of studies for other countries that turnover is lower in pension-covered jobs. The table also displays the number of individuals with vested pension benefits among persons covered by pensions. The data show that mobility is much higher among holders of non-vested pensions than among those with pension benefits already vested. The wage data displays averages of CPI-deflated wages observed in 1996, 1997, and 1998 as well as the average wages in the respective prior year. We find considerable wage differences between jobs covered and not covered by pensions. Those who were not a member of an occupational pension scheme in 1995 have far lower earnings than those who were scheme members. In fact, the gross monthly wages in non-pension jobs (de-

[3] Marital status and number of children are somewhat ambivalent proxies for stability preferences, however. We know that males seek better career opportunities once they have become fathers.

flated to values of 1995) were 26% lower than in pension-covered jobs. While individuals leaving jobs not covered by pensions earn about average wages in this group, mobility out of pension-covered jobs concentrates among workers with earnings clearly below the average. The table also shows that movers realize larger wage increases than stayers, irrespective of pension status. Wage increases for movers from pension-covered jobs are almost five times as high as for stayers. However, on average these movers can not compensate the initial wage differentials that separate them from the stayers' earnings. On the other hand, movers from non-pension jobs can realize a wage advantage over the group of stayers in such jobs.

- table 1 about here -

These results may indicate that workers on pension-covered jobs receive a considerable wage premium which makes it more difficult for these individuals to find a better alternative job. This would provide one explanation for the lower mobility among these workers. It would also explain why among pension scheme members those with below-average earnings are more mobile: possibly they face more outside opportunities for better jobs than those already earning above-average wages. The wage relationship of movers and stayers on non-pension jobs seems to be in line with the standard results of mobility studies. These studies usually find better educated, young males with higher occupational status to be more mobile than other workers. According to these findings it comes as no surprise that mobile workers can realize relatively high post-mobility wages. Of course we can only confirm these intuitions when controlling for personal and job characteristics as we do next.

The first column of table 2 presents the results of the reduced-form probit estimate for model 1 (dummy approach). The estimate represents the effect of individual and job characteristics on the selection into stayers and movers, both via the wage differential of moving versus staying and via the mobility costs. In the reduced form of the model the pension status has no significant influence on mobility, although the coefficient has the expected sign. However, the results from the reduced probit cannot be directly interpreted because their effect on mobility depends partly on their influence on the wage differential. The role of this regression in the estimation procedure is to obtain estimates of the selectivity terms which allow estimation of the movers and stayers wage equations corrected for selectivity bias.

- table 2 about here -

Columns two and three of table 2 present the wage equations of movers and stayers using the log of monthly gross wages in values of 1995 as dependent variable. The test statistics werde derived using the Sandwich estimator for clustered data. The coefficients of the variables follow standard expectations. Higher educational degrees are associated with higher earnings; females and individuals with residence in eastern Germany earn significantly less. There is a non-linear relationship between earnings and experience in the labour market, with earnings reaching a maximum after roughly 16 years of experience for movers and after 22 years for stayers. The firm tenure variable is not significant for movers which is in line with the expectation that new employers do not reward tenure at the last employer. The wage equations also include a dummy for pension status. These test for the existence of a wage premium for pension-covered workers after controlling for personal characteristics. The regression results confirm the descriptive evidence of table 1: workers in pension-covered jobs receive significantly higher wages. Finally, the effect of unmeasured characteristics on wages is captured by the selectivity terms λ_m and λ_s. The coefficient of λ_s for stayers is significant, giving evidence of selection bias. The negative coefficient implies a positive selection into the stayers' group: stayers have unobserved characteristics which grant them higher wages than movers would receive in case of immobility. The coefficient of λ_m is also negative - implying positive selection of movers - but insignificant.

The coefficients of the wage equations are applied to each individual's characteristics to calculate a mover's and a stayer's wage. The predicted difference in log wages is then used to estimate the structural probit equation (7) using maximum likelihood. Before presenting the results, we should make some comments on the predicted wage differentials. Somewhat surprisingly we find that average predicted wage differentials are negative for movers and for stayers, the average wage loss being smaller for movers than for stayers. We would have expected a wage gain of moving at least for those changing jobs. A closer look at the data reveals that the effect of unobservables on observed wages is quite substantive. When predicting a mover's and a stayer's wage for each individual, we do not take account of selectivity because the counterfactual selectivity terms are unknown. Thus the predicted wage differentials measure only the differences in returns to human capital and pension status[4] and neglect the selection effects in the observed state as well as in the respective counterfactuals. Bearing this in mind, the predicted wage differentials are consistent with expectations because estimated wage losses are lower for movers than for stayers. Thus we can proceed with

[4]By including pension status in the wage equations we assume this status to be constant over time. This assumption seems plausible for pension holders who are unlikely to move to a non-pensionable job, sacrificing the wage premium. However, individuals without pensions may change jobs to improve their pension status and receive the wage premium connected with it. Hence our wage estimates for movers without pensions are downward biased. This is a second explanation for predicted negative wage differentials.

reporting the results.

In table 3 we report the results of the structural probit mobility equations for models 1-3. Marginal probabilities are shown in the appendix. We do not display the reduced-form and wage equations for models 2 and 3 because they do not differ substantially from model 1. The coefficients of the non-pension variables are quite insensitive to variations in the pension variables. Across all models the results for the non-pension variables confirm most of the standard expectations about mobility behavior. Worker mobility tends to increase with growing wage gains (or diminishing wage losses). However, the coefficient is not significant. A possible explanation is that selection effects - as explained above - are more important for worker mobility than differences in returns to human capital. Older workers are less likely to change jobs than younger workers are. This may be because of older workers' shorter pay-off period from mobility. Women are significantly less mobile than men, possibly because they tend to work in occupations with less outside career opportunities and face more social restrictions in their household backgrounds. At sample means of the other variables being female reduces the probability of changing employer by 0.7 percentage in all models. This is a relative change in mobility probability of roughly 60%. Home ownership also significantly reduces job changes by 0.8 percentage points or approximately 70% at sample means. This may be because stability-oriented persons select into property which is costly to transfer. The data show that only about 1/3 of workers on non-pension jobs are home owners and roughly 1/2 of workers on pension jobs own property. On the other hand, marital status and the number of children have no significant effect on mobility. The estimation results thus suggest that being married and having children are inappropriate proxies for stability preferences. As noted above, responsibility for a family may lead workers to seek more career opportunities and thus may even increase the probability of changing jobs. The estimation shows that the overall effect is undetermined.

- table 3 about here -

Results for firm size and occupational status dummies - not displayed - show that mobility is higher from smaller firms and for managers and white collar workers than from larger firms and for blue-collar workers. Furthermore, mobility costs vary by industry.

Of most interest in this paper are the results for the occupational pension and capital loss variables. Model 1 only uses an occupational pension dummy to measure the effect of pensions on mobility. As expected, the coefficient is negative, showing that pension coverage may deter mobility. However, it is not significant. Thus, having or not having an occupational pension does not, in general, influence mobility behavior over and above the influence of wage premiums on expected wage gains. Introducing measures of capital loss for vested and non-vested workers into the analysis gives a more differentiated picture. Coefficients on both measures of capital loss - the equations derived for the wage-related pension formula (model 2) and their more general approximations (model 3) - are negative and significant both for vested and non-vested workers, although in model 3 capital loss for non-vested workers is significant at the 10 % level only. The larger the capital loss gets for both vested and non-vested workers, the less likely it is that they will change employer. The results show that it is not pension coverage in general which deters mobility. It is the size of capital loss which determines (among the other factors) who changes jobs and who does not. The coefficient on the pension dummy remains insignificant in model 2 and 3 but changes to a positive value. This suggests that the pension status may act, for example, as a proxy for non-pecuniary job attributes not captured in our model. This could explain why the coefficient is insignificant in model 1.

Comparing models 2 and 3, we find that in both estimations the negative coefficients reflect a mobility-deterring effect of pension capital losses. Thus the approximations of capital loss introduced by McCormick and Hughes (1984) work well in our framework. However, the values of the coefficients on the capital loss variables of model 3 can not be directly interpreted because the approximations merely represent functional forms. The capital loss variables in model 2 do capture real capital loss under the assumption that the benefit formula is wage-related and assuming values for expected wage growth, expected inflation rate, and the annual accrual rate (see section 2.3). The results appear plausible because the coefficients on the capital loss variables and thus the magnitude of the effect is almost the same for vested and non-vested benefits. Evaluated at the means of the other variables, in model 2 the marginal probability of job change declines by 0.12 percentage points or a relative change in mobility probability of 11% for a 1000 DM (511 Euro) real loss of pension benefits both for vested and non-vested benefits.

6 Summary and Conclusions

This paper provides evidence that occupational pension coverage in Germany reduces worker mobility through the mechanisms discussed in earlier papers for the U.S. and U.K.: There is a higher level of compensation in pension-covered jobs which makes mobility from such jobs less attractive. Sorting into pension-covered jobs also plays a role in reducing mobility if we accept that home ownership is an adequate proxy for stability preferences. Finally, and most important for policy, we find that pension coverage deters voluntary job transitions by imposing a capital loss on early leavers. Thus our paper contributes to showing that the effects of occupational pensions on mobility do not differ substantively between the anglo-saxon countries thus far studied and Germany whose labor market stands out for its relatively rigid regulation which corresponds with high firm attachment and internal flexibility.

Distinguishing between capital loss of pension benefits which are vested and those which are not yet vested, we find that both sources of capital loss pose an obstacle to job changes. This result holds both for a model where capital loss is computed on the basis of a wage-related benefit formula and for an alternative more general model where capital loss is approximated by interaction terms. According to our estimations, both loss of benefits which are not vested and loss due to erosion of real capital value over time of vested benefits deter worker mobility to the same extent. A relative change in mobility probability of 11% for a 1000 DM (511 Euro) capital loss is considerable, although other aspects like gender and home ownership are more important for the mobility decision.

Regarding pension regulation the results show that decreasing vesting periods to enhance mobility is an important policy option. We would expect that the reduction of the vesting period from 10 to 5 years performed in 2001 was already an important step to foster mobility in Germany. In international comparison a further reduction seems feasible. The mobility-deterring effect of a lacking indexation of preserved benefits (for vested early leavers) is equally important. Hence an indexation of deferred rights could be an effective measure to enhance mobility. Both policy options would reduce retirement income losses of multiple job holders. Against the background of current policy discussions at EU level further research into the relationship between mobility and occupational pensions is required as a basis for well-founded conclusions.

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Table 1: Job Mobility, Occupational Pensions, and Wages

	No occupational pension in 1995 job		Occupational pension in 1995 job	
	Stayer	Mover	Stayer	Mover
Number of observations	6,030	154	2,077	26
Mobility rate (%)		2.6		1.3
Pension benefits vested	-	-	1,215	4
Pension benefits vested (%)	-	-	58	15
Average monthly wage t-1 (DM)	3,788	3,859	5,140	4,396
Average monthly wage t (DM)	3,875	4,181	5,251	4,868
Average wage increase betw. t-1 and t (%)	2.3	8.3	2.2	10.7

Notes: The observations are from the pooled sample and consist of those 3,356 full-time employed individuals observed in t=1996, 2,704 individuals observed in t=1997 and 2,227 individuals observed in t=1998 for whom wage observations are available for the prior year (t-1). Gross monthly wages deflated to values of 1995 by the Consumer Price Index.

Source: German Socio-Economic Panel, 1995-1998; author's calculations.

Table 2: Reduced Form and Selection-corrected Wage Equations, Model 1

	(1) reduced form probit		(2) mover's wage equation		(3) stayer's wage equation	
College/university	0.423	(2.59)**	0.688	(5.60)**	0.564	(25.68)**
Vocational degree	0.225	(1.93)+	0.206	(1.94)+	0.152	(11.72)**
Female	-0.172	(1.98)*	-0.127	(2.50)*	-0.153	(15.07)**
East	-0.194	(2.28)*	-0.269	(4.07)**	-0.271	(22.75)**
Experience	-0.025	(1.29)	0.026	(2.49)*	0.014	(7.04)**
Experience squared/100	0.0094	(0.22)	-0.081	(2.96)**	-0.034	(7.88)**
Tenure	-0.031	(4.03)**	0.0049	(0.69)	0.0040	(5.62)**
Occupational pension	-0.087	(0.86)	0.150	(2.44)*	0.131	(11.11)**
Age	-0.0029	(0.27)				
Home ownership	-0.251	(3.14)**				
Married	0.072	(0.84)				
Number of children<16	0.051	(1.21)				
Period 1995-6	0.097	(1.09)	-0.009	(0.18)	-0.019	(4.00)**
Period 1996-7	0.02	(0.17)	0.093	(1.69)+	-0.013	(3.46)**
λ_m			-0.095	(0.87)		
λ_s					-0.486	(3.24)**
Constant	-1.30	(4.31)**	7.69	(30.41)**	8.09	(288.40)**
Observations	8,361		181		8,180	
log likelihood	-760.53					
R-squared			0.46		0.43	

Notes: + significant at 10%; * significant at 5%; ** significant at 1%. Reduced form probit (1): dependent variable is binary, equalling 1 if mobile and 0 if not. Absolute value of z-statistics in parentheses. Controls for industry and firm size in the year of mobility decision. Wage equations (2) and (3): dependent variable is log of monthly gross wages, deflated to 1995 values by Consumer Price Index. Robust standard errors are derived using the Sandwich estimator for clustered data. T-statistics in parentheses. Results for models 2 and 3 available on request.

Source: German Socio-Economic Panel, 1995-1998; author's calculations.

Table 3: Structural Probit Mobility Equations, Models 1-3

	Model					
	1		2		3	
Log of predicted wage differential	1.645	(1.05)	1.897	(1.26)	1.741	(1.22)
Age	-0.026	(2.69)**	-0.024	(2.90)**	-0.024	(2.90)*
Female	-0.241	(2.02)*	-0.244	(2.15)*	-0.240	(2.02)*
Home ownership	-0.285	(3.58)**	-0.275	(3.36)**	-0.278	(3.51)**
Married	0.026	(0.32)	0.031	(0.35)	0.032	(0.38)
Number of children<16	0.026	(0.63)	0.029	(0.69)	0.029	(0.69)
Occupational pension	-0.150	(1.05)	0.247	(1.26)	0.236	(1.14)
CL_i^v (equation 3/1000)			-0.0402	(2.03)*		
CL_i^{nv} (equation 4/1000)			-0.0427	(2.16)*		
CL_i^v (pens.^{v}*tenure*yrs to retirem.)					-0.0016	(2.09)*
CL_i^{nv} (pension^{nv}*tenure)					-0.0582	(1.87)+
Constant	-0.286	(0.61)	-0.172	(0.43)	-0.291	(0.68)
Observations	8,361		8,361		8,361	
log likelihood	-780.58		-775.34		-776.99	

Notes: + significant at 10%; * significant at 5%; ** significant at 1%. Dependent variable is binary, equalling 1 if mobile and 0 if not. Bootstrapping was used to derive robust standard errors for the two-stage procedure (1000 iterations). T-statistics in parentheses. 8 industry, 4 firm size, 3 occupational status, and 3 time dummies included.

Source: German Socio-Economic Panel, 1995-1998; author's calculations.

Table A1: Descriptive Statistics

Variable	Mean	Std. Dev.	Min	Max
College/university	0.10	0.30	0	1
Vocational degree	0.74	0.44	0	1
Female	0.34	0.47	0	1
Experience	19.40	10.66	0.83	48
Experience squared	489.86	471.78	0.69	2,304
Tenure	11.07	9.21	0	47
East	0.30	0.46	0	1
Age	40.68	10.35	20	67
Home ownership	0.38	0.49	0	1
Married	0.68	0.47	0	1
Number of children<16	0.68	0.93	0	6
Agriculture	0.01	0.12	0	1
Energy	0.03	0.16	0	1
Metal	0.23	0.42	0	1
Construction	0.11	0.32	0	1
Trade	0.16	0.37	0	1
Social	0.20	0.40	0	1
Bank	0.09	0.29	0	1
Other industry	0.16	0.37	0	1
Manager	0.15	0.36	0	1
White collar	0.36	0.48	0	1
Blue collar	0.49	0.50	0	1
Firm size < 20	0.18	0.39	0	1
Firm size 20-199	0.30	0.46	0	1
Firm size 200-1999	0.28	0.45	0	1
Firm size $\geq$2000	0.24	0.43	0	1
Capital loss, vested benefits, model 2	1.87	5.24	0	47.73
Capital loss, non-vested benefits, model 2	1.05	3.72	0	50.35
Capital loss, vested benefits, model 3	48.09	127.63	0	609
Capital loss, non-vested benefits, model 3	0.69	2.29	0	19
Occupational pension	0.25	0.44	0	1
Period 1995-6	0.40	0.49	0	1
Period 1996-7	0.33	0.47	0	1
Period 1997-8	0.27	0.44	0	1
Gross monthly wage (DM) in values 1995	4,232	1,648	1,144	15,916

Table A2: Marginal Probabilities of Structural Probit Mobility Equations, Models 1-3

	Model					
	1		2		3	
Log of predicted wage differential	0.050	(1.05)	0.055	(1.26)	0.052	(1.22)
Age	-0.00078	(2.69)**	-0.00070	(2.90)**	-0.00073	(2.90)*
Female	-0.0068	(2.02)*	-0.0066	(2.15)*	-0.0066	(2.02)*
Home ownership	-0.0082	(3.58)**	-0.0075	(3.36)**	-0.0078	(3.51)**
Married	0.00080	(0.32)	0.00089	(0.35)	0.00093	(0.38)
Number of children<16	0.00080	(0.63)	0.00086	(0.69)	0.00085	(0.69)
Occupational pension	-0.0042	(1.05)	0.0084	(1.26)	0.0081	(1.14)
CL_i^v (equation 3/1000)			-0.0012	(2.03)*		
CL_i^{nv} (equation 4/1000)			-0.0012	(2.16)*		
CL_i^v (pens.^{v}*tenure*yrs to retirem.)					-0.000048	(2.09)*
CL_i^{nv} (pension^{nv}*tenure)					-0.0017	(1.87)+
Observations	8,361		8,361		8,361	
log likelihood	-780.58		-775.34		-776.99	
Observed P	0.022		0.022		0.022	
Predicted P (at sample means of other variables)	0.012		0.011		0.011	

Notes: + significant at 10%; * significant at 5%; ** significant at 1%. Dependent variable is binary, equalling 1 if mobile and 0 if not. Bootstrapping was used to derive robust standard errors for the two-stage procedure (1000 iterations). T-statistics in parentheses. 8 industry, 4 firm size, 3 occupational status, and 3 time dummies included.

Source: German Socio-Economic Panel, 1995-1998; author's calculations.

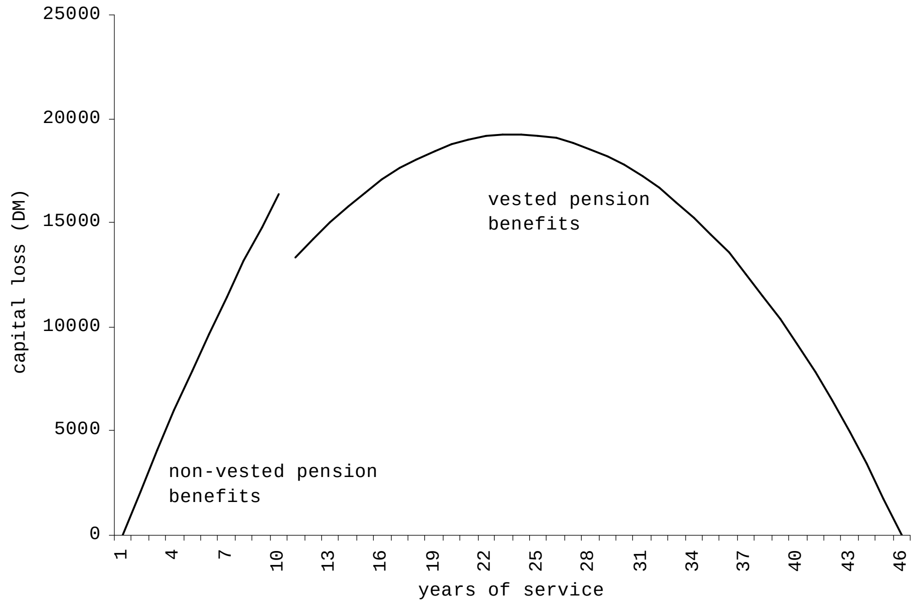

Figure 1: Capital loss of vested and non-vested pension benefits