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ABSTRACT

The Timing of Careers and Human Capital Depreciation*

This paper explores the short and long run effects of career interruptions on wages for young skilled workers in West Germany. The analysis distinguishes four types of career interruptions: unemployment, parental leave for female workers, national service for male workers and other non-work spells. We adopt the human capital model by Mincer and Polachek (1974) with homogenous human capital and test whether net depreciation is equal across types of employment interruptions, and equal in the short and in the long run. The main findings are that timing effects seem important and net depreciation differs across types of interruptions.

JEL Classification: J16, J3, J7

Keywords: wage determination, human capital depreciation, unemployment

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1 Introduction

Previous studies on wage effects due to career interruption have often focused on females (e.g. Corcoran et al. (1983), Kim and Polachek (1994)). In modern economies, however, both males and females experience career interruptions. (e.g. Light and Ureta (1995)) First, interruptions due to unemployment are particularly frequent for young workers in many countries. In addition, interruptions are likely to occur for other reasons such as parental leave, or when taking a sabbatical to go on further education or to take time off. The latter may also be offered by firms to employees who they want to maintain but cannot employ at the very moment due to a gap between project work, for example. From a governmental or firm policy point of view wage effects due to career interruptions are also of interest with regard to incentives for going on a sabbatical, or taking up parental leave for male workers, for example.

This paper investigates the impact of working and non-working periods on wages for young skilled female and male workers in West-Germany. The emphasis is on the estimation of the short and long run effects of different types of non-working periods. The data allow to distinguish three types of career interruptions and a residual group: unemployment, parental leave for female workers, national service for male workers and other non-work spells. Estimates of the short and long run wage effects due to different types of career interruptions are informative in order to investigate the implied income risk. A related issue is the possibility of scarring effects (Heckman and Borjas (1980)), which are especially problematic for young workers who are only at the start of a 30-40 year long career. Generally, unemployment scarring denotes the fact that individuals who have been unemployed once may be more likely to become once more unemployed and, therefore, their wages may be permanently affected.

For the analysis of the wage determination process we adopt the standard human capital model by Mincer and Polachek (1974) that segments the work history into work experience spells and non-working spells. Their model allows to estimate the particular effect of each work history segment, and hence considers timing. Furthermore, it allows to give a structural interpretation to the key parameters, i.e. the coefficients of the work experience variables and non-work variables. The parameters measure the net effect of the return to and depreciation of human capital.(Mincer (1974)

The model of Mincer and Polachek is still the dominant approach in the literature on wages and gender wage gap. Related models have been estimated for several countries. The main finding is that non-working periods result in non-increasing wage growth. In detail, however, little consensus seems to exist on the size of the effect of specific types of employment interruptions on wages. By reviewing the model, we show that these studies implement often the restrictive assumption that the timing does not matter and the loss from non-working is linear. The goal of this paper is to investigate to what extent simplifying assumptions imputed on the Mincer and Polachek (1974) model are justified. That is, whether net depreciation is equal across types of employment interruptions, and whether net depreciation is equal in the short and in the long run. Furthermore, we derive upper and lower bound estimates of gross investment and gross depreciation.

In this paper we use a data sample of young skilled full-time workers drawn from the new release of the IAB employment sample (IABS)[1] for West-Germany, the regional file of the IABS. The IABS is an administrative data set. It is particularly suited for this type of analysis for three reasons: First, it allows to measure the human capital accumulation process in detail and, second, it distinguishes types of time out of work. Third, it is a large and long panel that allows to follow individual work careers over several years, that is up to 15 years. It contains precise information on wages for full-time workers, as well as detailed information on education and human capital acquisition and types and duration of employment interruptions.

The structure of the paper is as follows: The next section reviews the wage model by Mincer and Polachek (1974) and summarises the main studies that have estimated wage effects of employment interruptions. Section three

[1]IABS abbreviates the *Institut für Arbeitsmarkt und Berufsforschung Sample.*

describes the data. In Section four contains institutional details. In section five, the empirical results are presented. The final section concludes.

2 The Model

We specify a simple human capital model that suits to interpret the empirical results in a human capital framework. It is adopted from Mincer and Polachek (1974) where they extend the simple Mincer earnings equation (Mincer (1974)) by allowing to distinguish between gross investment and depreciation of human capital.

Suppose individual's earnings capacity, E, is determined as follows:

$$E_t = E_{t-1} + r_{t-1}C^*_{t-1} - \delta_{t-1}E_{t-1} \tag{1}$$

where r_t is the rate of return to the individual's human capital investment, and C^*_{t-1} is the gross investment in monetary units in period $t-1$, δ_{t-1} is the depreciation rate of the stock of human capital. Using $k^*_t = C^*_t/E_t$, we can rewrite this equation to

$$E_t = E_{t-1}(1 + r_{t-1}k^*_{t-1} - \delta_{t-1}) \tag{2}$$

By recursion and applying the logarithmic approximation of $ln(1 + rk^*) \simeq rk^*$ yields

$$lnE_t = lnE_0 + \sum_{j=0}^{t-1}(r_j k^*_j - \delta_j) \tag{3}$$

Polachek and Mincer (1974) suggested to break up the post schooling period into successive segments of participation and nonparticipation. Assuming constant investment within segments at rate a_j during the segment $\tau - 1, \tau$, - e_j - this leads to the following equation:

$$lnE_t = lnE_0 + r_s s + \sum_{j=s+1}^{t-1} r_j a_j e_j \tag{4}$$

where s is the number of full time years of schooling and $k_s^* = 1$ (in equation (3)) is assumed during full time schooling.[2] Whereas $(a_j e_j) > 0$ denotes positive net investment (ratios), $(a_j e_j) < 0$ represents net depreciation rates, likely in periods of nonparticipation.

To understand the estimates of net depreciation rates it is helpful to express the earnings function in terms of gross investment rates and depreciation rates. As an example we use three segments of working life. Suppose the first segment is a working spell, e_1, the second segment is a non-working segment,e_2, and the third one is another working segment, e_3. To consider the fact that observed earnings correspond more closely to net earnings than earnings capacity we us $Y_t = E_t - C_t^* = E_t(1 - k_t^*)$, where Y are net earnings. Observed earnings in period t can then we written as:

$$lnY_t = \alpha + (r_s s - \delta_s) + (r_1 k_1^* - \delta_1)e_1 + (r_2 k_2^* - \delta_2)e_2 + (r_3 k_3^* - \delta_3)e_3 \tag{5}$$

where $\alpha = lnE_0 + ln(1 - k_t^*)$. Each coefficient in this equation is the compos-

[2] Mincer (1974) has shown that a quadratic function in (potential) work experience is an appropriate approximation.

ite of the return to the investment net of depreciation. The net effects are our parameters of interest. Empirically it is difficult to identify the separate effects. Noteworthy, this approach allows also human capital to depreciate while working due to technological progress, for example. However, the total net wage effect is likely to be positive. In case, for example, of unemployment the depreciation of human capital is likely to dominate. Through special training programmes this effect may be partly offset. Types of employment interruptions could thus be characterised by different investment and depreciation effects if one could identify the separate gross effects.

Several authors have adopted similar approaches to estimate the effect of working and non working spells on wages. One can group them into four models that imply different constraints:

The first approach assumes that net depreciation rates are equal across employment interruption types and timing does not matter; thus, net depreciation in the short and in the long run is the same. As an example, Kim and Polachek (1994) estimate this model for the U.S.. They control for accumulated work experience and home time in period t. From their less intermittent sample of male and female workers they find a loss from 2 to 13 percent per year.[3]

The second approach allows net depreciation rates to vary across types,

[3]These results refer to their first differences estimation results and instrumental variables first difference estimation results, see table 3.

but assume that timing is irrelevant. In Albrecht et al. (1999) this model has been estimated for Swedish samples of males and females. Their logarithmic wage model is quadratic in total accumulated work experience, and includes linear terms in education and time out of work segments. These are parental leave, household time, other time out, unemployment and military service.[4] They restrict parameters, i.e. $(r_j k_j^* - \delta_j)$ in eq (6), to be the same across employment interruption segments of the same type and estimate, therefore, a mean effect across all segments of the same type, assuming that the timing of interruptions do not matter. They find that an additional year of time out of work decreases wages of females by 1.5 percent and wages of males by 4.9 percent. For females, losses seem to be larger due to unemployment, 4.4 percent, than due to parental leave, 1.8 percent. For male workers' losses are smaller due to unemployment, 1.6 percent, than due to parental leave, 7.1 percent. They suggest that this is due to effects of signalling.

The third approach allows net depreciation rates to be different across time, but assumes equality across types. This model, that follows directly from the Mincer and Polachek (1974) application, has been estimated in for the U.S. in Corcoran and Duncan (1979) and Corcoran et al. (1983). Controlling for education, work experience, total out of work time since education and missed hours due to illnesses, Corcoran and Duncan (1979)

[4] They use panel data and take account of unobserved heterogeneity across individuals by applying a fixed effects estimator.

present results for the white male and white female wage gap. Corcoran et al. (1983) look at female wages and the effect of work experience, time out of work, part and full time work, work in female dominated and male dominated jobs. They find a significant short term loss of not working of approximately 3.3 to 4.1 percent per year.

Light and Ureta (1995) test whether the model that allows for timing effects fits the data better than a model that just includes controls for accumulated work experience and time out of work.[5] Hence, they allow the coefficients $(r_j k^*_{jk} - \delta_{jk})$ for $j = t, t-1, ..$ and variable k to differ. k denotes work experience and time out of work. They do not distinguish types of employment interruptions. They find that a more flexible work history model allowing for timing dominates a simple Mincerian earnings equation, which is quadratic in work experience. Focusing on the returns to work experience, they found that about 20-30 percent of the overall experience gap can be explained by male-female differences in the timing of work histories. Furthermore, they find that short run losses from employment interruptions are 13 percent and in the longer run, that is after 2 years, insignificant. The study by Robst and VanGilder (2000) on the U.S. finds similarly that losses due to recent home time and total home time differ. Gregory and Jukes (2001) have shown for men in the U.K. that short run losses from unem-

[5]The authors use the National Longitudinal Survey.

ployment are relatively high, approximately 10 percent per quarter. They decline to approximately 4 percent per quarter after 9 quarters.[6] They conclude that evidence of unemployment scarring is found. *Unemployment scarring* denotes the fact that individuals who have been unemployed once may be more likely to become once more unemployed. (Heckman and Borjas (1980)) In addition unemployment may lead, for example through depreciation of human capital, to lower wages on return in the short run and in the long run.

In summary, these studies suggest that net depreciation varies across types and that timing seems important. Different to our paper none of these studies distinguishes at the same time different types of employment interruptions and short and long effects. In the following we will estimate the unconstraint model and test whether the more parsimonious models are rejected or not.

3 Data

In this paper we use a data sample of young skilled workers taken from the new release of the IAB employment sample (IABS)[7] for West-Germany, the

[6]These results are based on matched data from the New Earnings Survey Panel Dataset (NESPD) and Joint Unemployment and Vacancies Operating System (JUVOS).

[7]IABS abbreviates the *Institut für Arbeitsmarkt und Berufsforschung Sample*.

regional file of the IABS.[8] It is an administrative event history data set. It covers the period 1975 to 1997.

The IABS is a 1 percent random sample drawn from the event history data file of the social security insurance scheme, the employment statistics, collected by the German Federal Bureau of Labour.[9] The regional IABS contains all workers in West-Germany who have had at least one employment spell eligible for the social security insurance scheme. As a result, included are all dependent employees in the private sector, i.e. about 80 percent of total employment in West-Germany.[10] The event history data includes information on every change in working status distinguished into full-time work, part-time work, unemployment and interruption which captures national service and maternity or parental leave. The particular event history data structure implies that a unit of the data is a spell, and not necessarily a yearly spell.

We focus on a sample of full-time skilled workers who are observed from school leaving age, 16, onwards. This ensures that complete labour market careers are observed from entry onwards. This allows the precise measure-

[8]For details on the regional file see Haas (2001), and IABS and the sample young skilled workers see Bender et al. (1996) and Kunze (2001).

[9]The fact that the data was collected for administrative purposes is an obvious advantage and makes the data particularly reliable.

[10]Not included are: civil servants, self-employed, unpaid family workers and people who are not eligible for benefits from the social security system.

ment of human capital accumulation. Our sample contains only records on young full-time[11] workers, who have mostly, 95-98 percent, graduated from school after 10 years of schooling and who are observed afterwards in apprenticeship training. In practice apprenticeships takes 2 to 3 years; and in the data mean duration is 2.1 for females and 2.5 for males. Apprentices have an apprenticeship contract with the firm they are trained with; wages amount to about 20-30 percent of the wage of a skilled blue or white collar worker. We select individuals who have no further vocational training, no technical college or university degree. In the data individuals are followed over early careers. In 1975 the oldest are born in 1959 and in 1997 in 1981. The mean age is 27. We drop workers who are not working throughout age 26 to 30, selecting a sample of more highly attached workers. We use wage histories from 1980 to 1997. Extraction of these workers from the IABS leaves us with a sample containing approximately 17,000 individuals observed in at least two full-time working spells after completion of vocational training, and approximately 220,000 spells.

Main variables in the empirical analysis are logarithmic daily real wages from full time work, actual work experience and the time out of work measures: unemployment, interruption due to parental leave, interruption due to national service, and a residual 'non-work' variable.[12] The wage variable

[11]This rule leads to exclusion of 3 percent of spells for males and 18.6 percent of spells for females.

[12]This variable can be generated from gaps in the data records. Spells are reported

is the logarithm of the daily wage deflated by the CPI index. Wages refer to the main job.[13]

Individuals are in their first spell of full-time work on average 21 year old. Approximately 10 percent of females and 30 percent of males have a gap between completion of apprenticeship training and the first job. This is measured as non-work according to our definition. For males, this is likely to pick up the fact that they go to national service before starting to work. Otherwise we do not know whether this is further education, for example. Mean duration of the gap during the transition is low: 0.11 years for females and 0.31 years for males. To see how work histories of our sample of young female and male workers evolves we present the summary statistics measured at their last wage spell in Table 1. In their last spell individuals are between 30 and 37. They are on average 32 year old. As can be seen, male workers collect more work experience, but are also longer observed in time out of work spells. Interestingly, male workers are approximately 0.2 years longer unemployed than females. Due to national service, 83 percent of males are at least once observed in an employment interruption. Among female workers, the incidence is 62 percent. Hence, this indicates that interruptions in the early labour market career is a relevant issue for

with the calendar date.

[13]Generally, in the IABS wages are top-coded. This, however, does not imply problems for our sample of young skilled workers. None of the wage spells is top-coded.

both males and females.

4 Institutions

Unemployment insurance

Unemployment compensation consists of two parts: unemployment insurance (Arbeitslosengeld) and unemployment assistance (Arbeitslosenhilfe). Unemployment insurance is funded by contributions from workers and employers, while unemployment assistance is funded from general government revenues.

The claim to unemployment insurance is conditional on the claimant being unemployed and registered as such at the Employment Office, and has completed the qualifying period. Duration of claims for unemployment insurance are stated in the *Arbeitsförderungsgesetz* (employment law) in 1969 and subsequent amendments and depend on the duration of work in a job for which social insurance is compulsory. Independent of the year of the amendment of the law, workers younger than 42 years can claim unemployment insurance only up to a maximum of 12 months. After that period an unemployed person may be eligible for unemployment assistance. Due to apprenticeship training that takes 2-3 years all individuals qualify for receipt of unemployment insurance.

In the data we define unemployment as either receipt of unemployment benefits or assistance which differ by approximately 10 percent in the replacement rate. These spells can be identified directly from the data.

National service

For all men national service is compulsory if they live in Germany and are older than 18.[14] Usually, men are drawn to national service at the age of 19. However, men cannot serve before the age of 17[15] and after the age of 25 (in exceptional cases 28) men cannot be drawn into military service. National service in Germany is compulsory in the form of military service (*Grundwehrdienst*) or civil service. Due to bad health men can be released from service completely.

The duration of compulsory service varied over recent decades. In 1972 military service was compulsory for 15 months and civil service took one third longer, 20 months. From October 1989 military service was shortened to 12 months and accordingly civil service to 16 months.[16] Usually, national service implies the right to return to the previous firm and job, unless it was a fixed term contract like in the case of apprenticeship training. In

[14] See: *Wehrdienst - Kriegsdienstverweigerung - Zivildienst. in: Presse- und Informationsamt der Bundesregierung Referat Aussen-, Sicherheits- und Europapolitik*, Feb. 1996.

[15] Before the age of 17 military service can be served only with the consent of the parents.

[16] During natioanl service men receive a very low compensation, which is below the wage of an unskilled worker.

the data, the duration of national service is measured by the interruption variable in the IABS.

Parental leave

We summarise the duration of maternity leave and parental leave in one variable. For the interpretation of the results, it is important to note that the variable is bounded upward. Until 1986 the maximum duration of leave was 6 months, in 1988 it was extended to 12 months, in 1990 to 18 months and since 1991 it has been 3 years. The interesting feature of parental leave is that within the maximum period of duration of parental leave the parent has the right to return to the same firm. On return within a year the person even has the right to return to exactly the same work place.

Maternity and parental leave can be measured in the data only for young females. This is due to the fact that only one variable is available that contains all information on the status when an employment contract is put on hold; that is it continues without wage payments. This category subsumes maternity and parental leave, national service - which is not relevant for females, as well as long sickness leave, for example. We assume that the latter is negligible for young skilled females, and parental leave is the only relevant reason when this variable applies.

5 Estimation Results

The empirical model is specified such that coefficients are allowed to vary across segments - or time - and the type of time out of work variables[17]:

$$lnw_{it} = \beta_0 + \sum_{s=t}^{s=t-6} ex_{is}\beta_{1s} + \sum_{s=t}^{s=t-6} ue_{is}\gamma_{1s} + \sum_{s=t}^{s=t-6} ir_{is}\gamma_{2s} + \sum_{s=t}^{s=t-6} nw_{is}\gamma_{3t} + \nu_i + u_{it} \tag{6}$$

where i indexes individuals and t time. The dependent variable is the logarithmic real daily wage, lnw. The vector ex includes controls for work experience and ue stands for unemployment, ir for interruptions due to national service or maternity leave and nw for a residual group that we refer to as non-work. ν_i is an unobserved individual specific effect and u_{it} idiosyncratic noise. We allow effects to vary for up to six years into the past.[18]

We measure work experience by the percentage of the previous year, one year ago, two years ago, and so on, spent in employment. Furthermore, we include dummies if someone has had a spell of unemployment, of non-work or of interruption in the previous year, one year ago etc.. We let coefficients

[17]We assume that wages are negotiated in the beginning of a working spell, e.g. based on accumulated work experience until then. For technical reasons, therefore, the column of the experience variable, and other work history measures, are shifted by one spell relative to the wage column in our data matrix. As a result we let s start at t.

[18]Hence, we assume that the coefficients are equal for 6, 7 etc. years ago. Tests do confirm that effects further in the past do not vary significantly.

vary across lagged terms. As a result the returns to work experience can be interpreted as yearly effects, while the effect of a spell of unemployment, etc. is measured unweighted by the duration of the spell. The parameters of interest are the coefficients of the *work experience* variable(s) and the coefficients of the three types of *time out of work* variables.

As already mentioned in the discussion of equation (6), we can only identify the net effect of investment and depreciation. Assuming that no investment takes place during non work periods this may give an upper bound estimate of the depreciation effect during work.

We present results from fixed effects estimation taking account of unobserved individual specific factors. Due to the large number of observations and variation in the timing of work histories across individuals we can estimate this model. We acknowledge that in particular in the case of female workers labour supply effects, that is non-random sample selection, are likely to result in a positively selected sample of females. Hence, our estimates might be upward biased estimates of the (population) return to work experience and a too low estimate of losses. The estimates that we show can therefore be interpreted as marginal effects on accepted wages.

In the second column of Table 2 and Table 3 we report estimates using the entire sample of female workers, and male workers, respectively.[19] As ex-

[19]The Hausman statistic lead to the rejection of the random effects model. For the pooled female sample regression the Hausman test statistic is $\chi^2(44) = 897.68$, for the

pected we find positive returns to work experience both for male and female workers. Positive selection of females participating in the labour market is likely to explain the somewhat high return to most recently accumulated work experience. However, human capital accumulated up to six years ago pay 3.9 percent per year for females, and 2.1 percent for males. The return is increasing towards most recent periods to 8.5 percent for females, and 7.2 percent for males. For illustration, to compare these estimates of the return to work experience to the quadratic model (Mincer (1974)) specification we plot predicted wage experience profiles for females, Figure 1, and male workers, Figure 2. Entry wages and the other controls are normalised to zero. For the sample of male workers, estimates from both models are quite similar. Up to four years of experience predicted wages are equal. Then predicted wages from the quadratic model are up to 5 percentage points larger. For female workers, predictions from the model differ more extremely. After two years of experience the flexible model predicts higher wages and by 10 years of accumulated work experience the difference is approximately 20 percent. In contrast to the quadratic model the flexible model estimates suggest that returns are non-decreasing.[20]

More general evidence in favour of the more flexible specification is that

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male sample $\chi^2(44) = 1181.01$.

[20]Due to the fact that in our quadratic model estimation variables measuring time out of work are included linearly, comparison of the two models may render difficult.

the hypotheses of equality of the coefficients across time segments and for the same type can be rejected at the 5 percent significance level for both the female and male sample regresssion. This holds for working and non-working variables. [21]

The estimation results for female workers on employment interruptions show quite strong evidence that not working leads to a wage loss. However, the loss varies in size across types and timing. Losses from unemployment seem small and insignificant in the long run. Unemployment during the most recent year decreases wages on return by 1.9 percent. However, we do not find any 'scarring' effects in the sense that unemployment in the past affects wages. This shows that those skilled workers who return manage to catch up rather quickly. This is contrary to the findings by Gregory and Jukes (2001) for the U.K., who find scarring effects for young workers of all skill groups.

The non work variables, that summarise the residual group of not working, have significant negative coefficients for up to 5 years into the past. Hence, here long run effects seem to matter. The loss decreases from 4.7

[21] The F-statistics (p-values) for the test of equality of coefficients are for the female sample for the unemployment variables in Table 2 column 2 $F(6,66882)=3.53$ (0.0017), for non-work $F(6,66882)=28.96$ (0), and for interruptions due to parental leave $F(6,66882)=13.59$ (0). For males they are for the unemployment variables in Table 3 column 2 $F(6,142204)=3.45$ (0.0021), for non-work $F(6,142204)=27.07$ (0), and for interruptions due to national service $F(6,142204)=39.58$ (0).

percent on return to the job, to 1.7 percent when the spell of non-work has been 5 years ago. Interruptions due to parental leave seem to lead to extremely high losses. On return to the job, females face a 18.3 percent drop in wages. If the interruption has been 2 years ago the drop is 14 percent, and after 5 years 13 percent. Hence, here effects are very large and permanent. Hence, models like Mincer and Ofek (1982) suggesting that wage growth after return is higher than mean wage growth over the life cycle, that is a rebound effect, are not comfirmed.

For males, as shown in Table 3 column 2, results are very different. The effect of a spell of unemployment on wages is insignificant[22], non-work leads to a loss of 1.4 -1.9 percent if the spell has been during the most recent two years and less than 1 percent otherwise. National service leads to a wage gain of 3.2 percent on return to work. If more time has elapsed since than, it may decrease wages slightly at an increasing rate.

Labour markets are generally characterised by strong occupational segregation by gender. Females are more likely to work in services and males in manual jobs. Examples for young female skilled workers are professional clerical workers, sales person, receptionist, hyginist, banking professional, and nurse. For male workers examples are motor vehicle mechanic, electrician, professional clerical workers, machinist, joiner an pipe fitter. In fact

[22]Jointly we can reject the null hypothesis of significance.

more than 60 percent of females are working in the before mentioned occupations. For males, concentration is less strongly. In order to select on "typical" female working careers we reestimate the same model separately for female occupations and Integrated/male occupations. In the data approximately 120 occupations are distinguished. For simplicity, we define a *female occupation* as one where the fraction of female workers in that particular occupation is greater than 60 percent in 1990. All other occupations are summarised as *integrated and male occupations*.

Focusing on the effects due to not working for female workers we find that losses are more distinct in the group of female occupations. Interestingly, the short term loss from unemployment is also significantly negative for male workers in female dominated occupations. For females, the wage loss due to parental leave is also approximately 30 percent smaller in integrated and male occupations than in the comparison group for females. These findings indicate that it is the workers in female occupations who lose in the short run from unemployment. Hence, unobserved job characteristics are captured by the female occupation characteristic. An alternative explanation that this finding is due to the fact that females are on average longer unemployed than males can be excluded, as seen from Table 1. The human capital model by Polachek (1981) predicts that under optimization of life time earnings females will select into careers with relatively flat profiles. It follows that returns and losses are relatively small. The empirical evidence for the early career of young skilled females in Germany contradicts this model. Females in male dominated occupations, that should be those with high returns and on the job training content, are those who are better off.

The strong effect of an interruption spell due to parental leave may be the joint effect of depreciation and firm change. The latter may add to the depreciation by loss of firm specific capital. In Table 4 we show that approximately 22 percent of females do change firm, and approximately 78 percent make use of the option to return to the previous firm. In column 1 and 2 of Table 5 estimation results for females are shown when we control for the fact whether the firm of employment after the interruption is the same as before. While for those on parental leave the option to return to the former employer is guaranteed, this is not the case for the unemployed and those not reported working due to other reasons. Indeed, we do find short run evidence in favour of the firm specific capital hypothesis. In the short run females who change firm face a loss approximately 6 percent higher than in the situation when they would have stayed with the previous employer. In the longer run, however, the effect is negative or insignificant. As has been shown in studies for Sweden (Albrecht et al. (1999)), the impact of national service on wages may even be positive if remunerated similarly to work experience or taken as a signal for commitment, dutyfulness, and responsibility, for example. For the German data we do find similar evidence for the long run in case of firm stayers. Firm movers may face a small gain in the short and a loss in the long run.

6 Conclusions

This paper investigates the wage effects due to employment interruptions and due to the timing of the working career. Our results can be interpreted as effects on accepted wages. We test whether significant net depreciation effects are found for different types of time out of work spells, and whether those differ across time segments of the working career. For the analysis we use a sample of young skilled workers in Germany.

We find that the impact of not working on wages depends on the type. More specifically, the main findings are that the effects in the short and in the long run differ and net depreciation differs across types of interruptions. Therefore, models that allow only for a linear term in the accumulated time out of work spell, or those distinguished into types, are too restrictive. Also models with quadratic terms seem inadequate. We find little evidence for gender differences in losses due to unemployment. However, we find that **short time** losses are experienced in particular in those occupations where more than 60 percent of workers are female. This may proxy job characteristics that are unobserved in the data.

For female skilled workers, in particular, we find large differences in losses across types. While depreciation during unemployment and other non-work spells amounts to approximately 2-5 percent per year, females lose 13 -18 percent per year due to parental leave. Losses may be even larger if on return the individual changes firm. For male workers we also find differences in the impact on wages. However, variation across types is much smaller, and net depreciation seems to be very small, 0.5 to 2 percent.

From this evidence one can derive upper and lower bound estimates of gross depreciation during work and gross investment during non-work. If net depreciation while working is assumed to be between the lower bound, 0.5 to 2 percent, and the upper bound 13-18 percent, then gross investment on the job ranges from 8.5 percent per year to 26 percent per year. Assuming depreciation to be the same across types of employment interruptions and to be near the upper bound, that is 13-18 percent, then gross investment is zero while on parental leave. However, it is positive while unemployed or in another non-working status. While from this analysis it cannot be said to what extent other economic stories such as signalling (Albrecht et al. (1999)) explain these results, these findings may indicate that human capital acquired in employment interruptions may not be complementary with human capital remunerted in the labour market. Hence, it may capture constraints, for example, on effort and flexibility we cannot control for in our analysis.

Appendix: Tables

Table 1: Summary statistics for early career, 1981-1997 IABS data *

	Female Sample mean	(std.)	Male Sample mean	(std.)	Diff. in means
			Total Sample		
age	32.4396	(2.0984)	32.9277	(2.1901)	-.4881*
work experience	10.2313	(3.2503)	10.5538	(3.2828)	-.3224*
interruption					
due to parental leave	.2583	(.5092)			
interruption					
due to national service			.3107	(.4652)	
unemployment	.3176	(.8142)	.5125	(1.0655)	-.1949*
non-work	.8834	(1.9737)	1.0904	(1.8552)	-.2070*
time out of work	1.4593	(2.4335)	1.9137	(2.2815)	-.4543*
# of indiv.	5753		11000		
			Sample excluding spells with zeros in time out of work variable		
age	32.5548	(2.1042)	32.9740	(2.1973)	-.4192*
work experience	9.6449	(3.3703)	10.3131	(3.3374)	-.6681 *
interruption					
due to parental leave	.4166	(.5935)			
interruption					
due to national service			.3740	(.4867)	
unemployment	.5122	(.9846)	.6170	(1.1411)	-.1047*
non-work	1.4248	(2.3478)	1.3126	(1.9625)	.1121*
time out of work	2.3537	(2.7289)	2.3037	(2.3168)	.0500
# of indiv.	3567 (62%)		9138 (83%)		

*Note:** significant at 5 % level. ** Variables are measured at the last wage (working) spell.

Table 2: Fixed Effects Estimates of Wage Equations, 1981-1997 IABS data
Female Young Workers

Variables	Quadratic Model coef. (t-value)	Segmented Work History Model coef. (t-value)	Female Occupations coef. (t-value)	Integrated/Male Occupations coef. (t-value)
Experience	0.0499			
(years)	(16.52)**			
Experience squared	-0.0021			
	(43.97)**			
% of year spent working previous year		0.0847	0.0776	0.0793
		(10.32)**	(8.59)**	(4.43)**
1 year ago		0.0718	0.0704	0.0603
		(21.21)**	(19.94)**	(6.63)**
2 years ago		0.0669	0.0688	0.0435
		(20.00)**	(19.69)**	(4.94)**
3 years ago		0.0618	0.0643	0.0324
		(18.33)**	(18.26)**	(3.70)**
4 years ago		0.0506	0.0546	0.0316
		(14.86)**	(15.36)**	(3.61)**
5 years ago		0.0616	0.0600	0.0567
		(20.84)**	(19.39)**	(7.31)**
6+ years ago		0.0386	0.0418	0.0197
		(45.38)**	(45.11)**	(8.22)**
Unemployment	-0.0297			
(years)	(7.75)**			
1 if in unemployment previous year		-0.0193	-0.0238	0.0016
		(3.62)**	(4.08)**	(0.13)
1 year ago		-0.0092	-0.0089	-0.0049
		(1.89)	(1.66)	(0.46)
2 years ago		0.0049	0.0034	0.0154
		(0.97)	(0.61)	(1.40)
3 years ago		0.0081	0.0105	0.0067
		(1.55)	(1.82)	(0.60)
4 years ago		-0.0041	0.0012	-0.0142
		(0.76)	(0.20)	(1.22)
5 years ago		0.0007	0.0039	-0.0072
		(0.12)	(0.62)	(0.61)
6+ years ago		0.0009	0.0109	-0.0220
		(0.20)	(2.24)*	(2.20)*

Variables		Table 2 - Continued	Female Occupations	Integrated/Male Occupations
Non-Work	-0.0380			
(years)	(12.03)**			
1 if				
in non-work				
previous year		-0.0475	-0.0515	-0.0505
		(10.16)**	(9.90)**	(5.17)**
1 year ago		-0.0502	-0.0615	-0.0390
		(12.43)**	(13.79)**	(4.53)**
2 years ago		-0.0406	-0.0488	-0.0394
		(9.79)**	(10.61)**	(4.53)**
3 years ago		-0.0296	-0.0364	-0.0240
		(6.90)**	(7.64)**	(2.70)**
4 years ago		-0.0288	-0.0343	-0.0167
		(6.45)**	(6.94)**	(1.83)
5 years ago		-0.0172	-0.0212	-0.0047
		(3.70)**	(4.09)**	(0.50)
6+ years ago		0.0059	0.0030	0.0021
		(1.55)	(0.72)	(0.25)
Interruption due to				
parental leave	-0.3181			
(years)	(72.90)**			
1 if				
in interruption				
due to parental leave				
previous year		-0.1837	-0.1899	-0.1228
		(31.39)**	(30.25)**	(8.77)**
1 year ago		-0.1766	-0.1810	-0.1159
		(33.45)**	(32.23)**	(9.12)**
2 years ago		-0.1480	-0.1538	-0.0842
		(26.19)**	(25.50)**	(6.30)**
3 years ago		-0.1406	-0.1446	-0.0949
		(23.02)**	(22.10)**	(6.63)**
4 years ago		-0.1447	-0.1425	-0.1177
		(21.96)**	(20.18)**	(7.76)**
5 years ago		-0.1304	-0.1378	-0.0735
		(18.45)**	(18.16)**	(4.49)**
6+ years ago		-0.1339	-0.1450	-0.0647
		(26.00)**	(25.91)**	(5.50)**
Year Dummies	yes	yes	yes	yes
Constant	4.3840	4.4757	4.4890	4.5352
	(393.51)**	(377.55)**	(351.01)**	(147.96)**
# observations	78009	72679	62997	9682
# individuals	5753	5753	5395	1329
R^2	0.37	0.31	0.33	0.10
F(7, 66882) for ue		3.15		
p-value		0.0025		
F(7, 66882) for ir		57.32		
p-value		0.0		
F(7, 66882) for nw		448.13		
p-value		0.0		

Notes: Absolute value of t-statistics in parentheses. * significant at 5% level; ** significant at 1% level

Table 3: Fixed Effects Estimates of Wage Equations, 1981-1997 IABS data
Male Young Workers

Variables	Quadratic Model	Segmented Work History Model	Female Occupations	Integrated/Male Occupations
	coef. (t-value)	coef. (t-value)	coef. (t-value)	coef. (t-value)
Experience	0.0565			
(years)	(30.53)**			
Experience squared	-0.0018			
	(57.35)**			
% of year spent working previous year		0.0720	0.0844	0.0602
		(18.97)**	(8.07)**	(15.34)**
1 year ago		0.0562	0.0727	0.0468
		(27.67)**	(16.31)**	(21.65)**
2 years ago		0.0444	0.0601	0.0367
		(22.17)**	(13.75)**	(17.25)**
3 years ago		0.0355	0.0552	0.0273
		(17.63)**	(12.64)**	(12.75)**
4 years ago		0.0314	0.0534	0.0214
		(15.52)**	(12.26)**	(9.96)**
5 years ago		0.0402	0.0608	0.0276
		(22.17)**	(15.88)**	(14.22)**
6+ years ago		0.0213	0.0337	0.0160
		(39.21)**	(26.51)**	(26.74)**
Unemployment	-0.0190			
(years)	(9.01)**			
1 if in unemployment previous year		-0.0016	-0.0184	-0.0047
		(0.63)	(2.60)**	(1.76)
1 year ago		-0.0047	-0.0064	-0.0064
		(2.02)*	(1.00)	(2.70)**
2 years ago		-0.0021	-0.0103	-0.0026
		(0.90)	(1.60)	(1.06)
3 years ago		-0.0039	-0.0088	-0.0057
		(1.63)	(1.35)	(2.32)*
4 years ago		0.0030	-0.0147	0.0029
		(1.25)	(2.21)*	(1.16)
5 years ago		0.0008	-0.0059	-0.0011
		(0.34)	(0.86)	(0.45)
6+ years ago		-0.0088	-0.0182	-0.0014
		(4.50)**	(3.36)**	(0.69)

Table 3 - Continued

Variables			Female Occupations Occupations	Integrated/Male Occupations
Non-Work	0.0118			
(years)	(6.15)**			
1 if				
in non-work				
previous year		-0.0193	-0.0093	-0.0186
		(9.41)**	(1.46)	(8.96)**
1 year ago		-0.0146	-0.0142	-0.0132
		(8.19)**	(2.84)**	(7.20)**
2 years ago		-0.0093	-0.0138	-0.0064
		(5.16)**	(2.77)**	(3.48)**
3 years ago		-0.0074	-0.0097	-0.0039
		(4.04)**	(1.92)	(2.09)*
4 years ago		-0.0052	-0.0071	-0.0013
		(2.78)**	(1.39)	(0.69)
5 years ago		0.0017	0.0026	0.0046
		(0.92)	(0.51)	(2.36)*
6+ years ago		0.0061	0.0102	0.0108
		(3.63)**	(2.33)*	(6.15)**
Interruption due to				
national service	0.0098			
(years)	(3.20)**			
1 if				
in interruption				
due to national service				
previous year		0.0328	0.0507	0.0287
		(9.89)**	(6.05)**	(8.33)**
1 year ago		-0.0044	0.0152	-0.0067
		(1.50)	(2.12)*	(2.15)*
2 years ago		-0.0128	0.0032	-0.0122
		(4.21)**	(0.43)	(3.84)**
3 years ago		-0.0186	-0.0022	-0.0178
		(5.96)**	(0.29)	(5.44)**
4 years ago		-0.0157	-0.0022	-0.0150
		(4.91)**	(0.29)	(4.51)**
5 years ago		-0.0139	-0.0093	-0.0117
		(4.29)**	(1.20)	(3.44)**
6+ years ago		-0.0204	-0.0180	-0.0186
		(9.46)**	(3.32)**	(8.20)**
Year Dummies	yes	yes	yes	yes
Constant	4.6295	4.6195	4.5326	4.6441
	(683.91)**	(655.89)**	(261.17)**	(627.11)**
# observations	163503	153248	30632	122616
# individuals	11000	11000	3462	9571
R^2	0.39	0.36	0.54	0.31
F(7, 142204) for ue		3.86		
p-value		0.0003		
F(7, 142204) for ir		32.6		
p-value		0.0		
F(7, 142204) for nw		39.94		
p-value		0.0		

Notes: Absolute value of t-statistics in parentheses. * significant at 5% level; ** significant at 1% level

Table 4: Gender differences in mobility

Percent of firm stayers: In Parentheses fraction of stayers and all individuals

Comparison	Females	Males
before /after...		
unemployment	15.68	32.12
	(487/3105)	(3626/11290)
interruption	78.81	
due to parental leave	(1837/2331)	
interruption		82.74
national service		(3734/4513)
non-work	24.66	26.26
	(1272/5159)	(5056/19252)

Table 5 : Fixed Effects Estimates of time out work variables in the Wage Equations 1981-1997 IABS data

Variables	Female Sample Estimates		Male Sample Estimates	
		FIRMSTAYER interacted with Variables		FIRMSTAYER interacted with Variables
	coef. (t-value)	coef. (t-value)	coef. (t-value)	coef. (t-value)
1 if in unemployment				
previous year	-0.0205	0.0398	-0.0126*	0.0413
	(3.40)**	(3.34)**	(4.23)*	(8.94)**
1 year ago	0.0035	-0.0199	0.0108*	-0.0219
	(0.35)	(1.76)	(2.63)*	(4.59)**
2 years ago	-0.0019	0.0067	-0.0023	-0.0021
	(0.20)	(0.61)	(0.56)	(0.43)
3 years ago	0.0070	0.0008	-0.0009	-0.0037
	(0.66)	(0.07)	(0.20)	(0.73)
4 years ago	-0.0192	0.0189	0.0108	-0.0115
	(1.63)	(1.44)	(2.27)*	(2.12)*
5 years ago	0.0022	-0.0028	0.0020	-0.0009
	(0.17)	(0.20)	(0.40)	(0.15)
6+ years ago	0.0101	-0.0121	-0.0174*	0.0104
	(1.08)	(1.22)	(4.27)*	(2.39)*
1 if in non-work				
previous year	-0.0591	0.0357	-0.0201	0.0281
	(11.10)**	(4.30)**	(8.16)**	(7.17)**
1 year ago	-0.0182	-0.0395	-0.0064	-0.0097
	(2.11)*	(4.16)**	(1.86)	(2.48)*
2 years ago	-0.0289	-0.0156	-0.0108	0.0043
	(3.57)**	(1.70)	(3.26)**	(1.12)
3 years ago	-0.0204	-0.0120	-0.0127	0.0079
	(2.29)*	(1.21)	(3.56)**	(1.96)*
4 years ago	-0.0278	-0.0015	-0.0072	0.0039
	(2.95)**	(0.14)	(1.89)	(0.91)
5 years ago	-0.0163	-0.0011	-0.0041	0.0087
	(1.62)	(0.10)	(1.05)	(1.96)
6+ years ago	-0.0096	0.0171	-0.0083	0.0185
	(1.20)	(2.02)*	(2.40)*	(5.03)**

Table 5 - Continued

Variables	Female Sample Estimates		Male Sample Estimates	
		FIRMSTAYER interacted with Variables		FIRMSTAYER interacted with Variables
	coef. (t-value)	coef. (t-value)	coef. (t-value)	coef. (t-value)
1 if in interruption in parental leave				
previous year	-0.2305	0.0599		
	(22.33)**	(5.47)**		
1 year ago	-0.1450	-0.0350		
	(9.13)**	(2.11)*		
2 years ago	-0.1446	-0.0042		
	(10.02)**	(0.27)		
3 years ago	-0.0955	-0.0522		
	(5.88)**	(3.03)**		
4 years ago	-0.1557	0.0108		
	(8.56)**	(0.56)		
5 years ago	-0.0954	-0.0421		
	(5.02)**	(2.08)*		
6+ years ago	-0.1063	-0.0309		
	(7.47)**	(2.10)*		
constant				
1 if in interruption in national service				
previous year			0.0138	0.0138
			(2.09)*	(1.95)
1 year ago			-0.0113	0.0089
			(1.50)	(1.11)
2 years ago			-0.0251	0.0151
			(3.61)**	(2.01)*
3 years ago			-0.0284	0.0117
			(3.93)**	(1.49)
4 years ago			-0.0231	0.0092
			(3.10)**	(1.14)
5 years ago			-0.0165	0.0032
			(2.11)*	(0.38)
6+ years ago			-0.0218	0.0017
			(4.64)**	(0.34)
constant		4.4636		4.6143
		(374.53)**		(654.19)**
# observations		72679		153248
# individuals		5753		11000
R^2		0.31		0.36

Notes: * significant at 5% level; ** significant at 1% level. All regressions include year dummies.

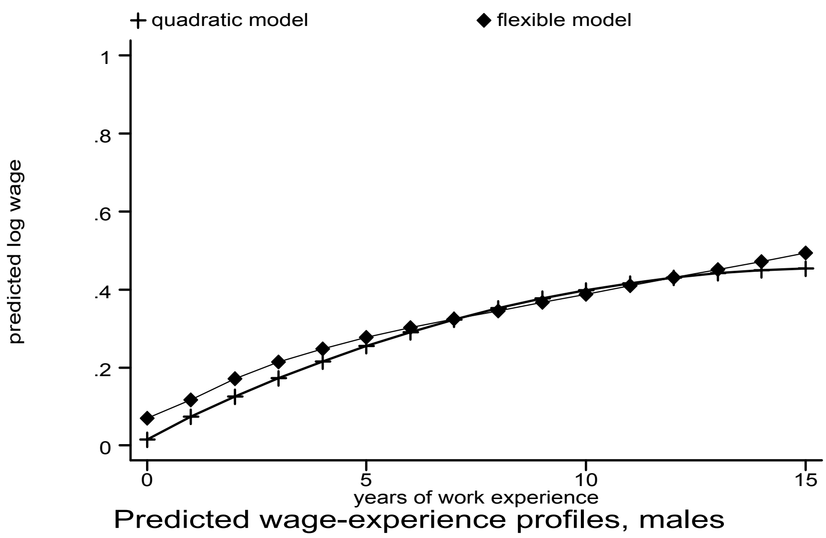

Figure 1:

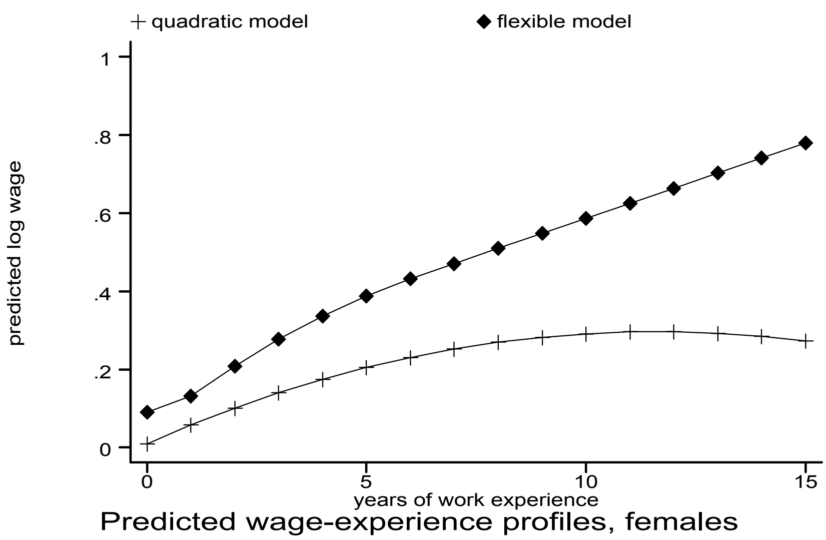

Figure 2:

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