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Abstract

This study investigates the re-employment hazard of displaced German workers. It uses data from the first fourteen sweeps of the German Socio-Economic Panel (GSOEP) survey for the purpose. The paper employs both parametric and non-parametric discrete-time models to study the re-employment hazard. Alternative mixing distributions have also been used to account for unobserved heterogeneity. Results based on single risk models show that the average hazard rate of exit via re-employment declines with the duration of time in unemployment. Accounting for unobserved heterogeneity does make a difference, but the crux of the results in terms of duration dependence remains largely unchanged. In terms of covariate effects, those at the lower end of the skills ladder, those who had been working in the manufacturing industry and those with previous experience of inactivity are found to have lower hazard of exit via re-employment. That those at the lower end of the skills ladder and those with previous experience of inactivity have difficulty getting re-employed calls for appropriate intervention to ameliorate the lot of the 'disadvantaged'.

Theme: Microeconomics of unemployment
Key words: Unemployment duration, job displacement, Germany
JEL classification: J6, J63, J64, C41

1. Introduction

The problem of unemployment has been featuring top of the European labour market literature for sometime now. Two of the most prominent discourses in this regard relate to the persistence of high unemployment in recent decades (Nickell, 2003; Heckman, 2002; Blanchard and Wolfers, 2000) and the contrast between the levels of unemployment in the US and the European labour markets (Nickell, 1997; 2003; Blanchard and Portugal, 2001; Heckman, 2002; Siebert, 1997; Blank, 1994). Both these aspects of the unemployment situation in Europe have been attributed to adverse shocks, adverse institutions and the interactions of the two. What has become common explanation more recently, particularly in relation to European labour markets, has to do with the role of institutions and how they respond to adverse shocks. The consensus in this regard is that labour market rigidities[1] have, at least in some of these labour markets, led to the rise in the level of unemployment by affecting the equilibrium level of unemployment as well as deviations of actual unemployment from its equilibrium level. The system of unemployment benefit and the type of employment protection scheme in place are, in particular, regarded to be important factors behind the high level of unemployment in these countries.

Labour market rigidities are likely to influence the equilibrium level of unemployment in several ways. They can affect the way in which unemployed individuals can be matched

[1] Labour market rigidities refer to a number of labour market characteristics including the presence of strong unions, minimum wages, generous unemployment benefit system, high payroll taxes, high employment protection and the likes (Blanchard and Wolfers, 2000; Nickell, 2003).

to available job vacancies. They also tend to raise the wage rate even when there is excess supply of labour. By lowering the search intensity of the unemployed, for example, the system of unemployment benefit in place - its level and the duration it lasts - reduces the readiness of the unemployed to fill available vacancies. Employment protection laws, on the other hand, are likely to make firms more cautious regarding filling available vacancies and, therefore, may lower the speed with which the unemployed may take up jobs. Because such laws are primarily meant to protect jobs, however, they also have the tendency to curtail involuntary separations and, therefore, lower inflows into unemployment. As such, therefore, the effect of employment protection schemes on the equilibrium level of unemployment may not be clear-cut. Nevertheless, at least in the countries that have been experiencing high levels of unemployment, one typical observation has to do with the prevalence of long-term unemployment. This means that the system of benefits in place and the type of employment schemes in place have made the labour markets of these countries rather stagnant.

The German labour market has been regarded as a typical case of the European labour markets exhibiting most signs of rigidity. The evidence in most recent studies that look into the nature of unemployment in Europe supports this claim. Nickell (1997; 2003), Bender *et al* (2002), Heckman (2002) and Blanchard and Wolfers (2000) are some of the most recent studies that look into labour market rigidities in terms of high employment protection, generous benefit schemes, strong presence of labour unions, wage-setting arrangements and the likes. The rise in the unemployment rate in Germany is also linked to a substantial increase in the share of long-term unemployed people who have been unemployed for over 12 months (Steiner, 2001; Nickell, 2003).

Long-term unemployment has particularly adverse effects to individuals experiencing it and society at large in many respects. To begin with, the long-term unemployed are likely to be discouraged to carry on searching for jobs. Because of the likely scaring effect of long-term unemployment, firms may not also be willing to take on such workers. It is also possible that such workers lose, at least in part, whatever skill they have if they stay unemployed for long. The combined effect of these will be to make unemployment even more persistent and possibly bring about poverty and social exclusion to the particular segment of the labour force that experiences such long-term unemployment (Arulampalam *et al*, 2001).

The types of workers that are most likely to suffer from the adverse effects of long-term unemployment are displaced workers who separate from their jobs involuntarily. The presence of high long-term unemployment due to labour market rigidities means that there are barriers to re-entry into the job market, and such difficulty of re-entry may particularly be relevant to displaced workers. It may also be the case that some segment of the displaced may fare particularly worse. Displaced workers that are at the lower level of the skill/qualification ladder may bear the brunt of the unemployment problem as a result of labour market rigidities. In the context of the German labour market, such workers form the bulk of displaced workers (Haile, 2002). The combination of employment protection schemes and minimum wage legislation may prove especially deadly for these workers, as firms may not be willing to take the risk of hiring them (Blanchard, 1998).

In this study I investigate how displaced workers fare in terms of the duration of time that they spend unemployed. There are very few studies[2] that investigate the duration of unemployment in Germany (Hunt, 1995; 1997; Steiner, 1994; 2001) and even fewer that focus on the unemployment duration of displaced workers in particular (Couch, 2001; Bender *et al*, 2002). It is therefore evident that not much is known regarding the duration of unemployment that unemployed workers in general experience. Most importantly, there is a huge gap in our knowledge pertaining to the duration of unemployment that displaced workers experience in Germany.

What is equally important is that there is lack of consensus regarding whether the outflow rate from unemployment declines as the duration of unemployment lengthens. Broadly speaking, the evidence on negative duration dependence in Europe is mixed (Machin and Manning, 1999). Moreover, results on negative duration dependence seem to be sensitive to the way unobserved heterogeneity is accounted for. Taking these into account, this study attempts to fill the gap in the literature by studying the duration of displacement unemployment in Germany.

This study has six parts and has the following structure. In section 2 a review of the literature on the duration of unemployment will be given focusing mainly on the

[2] This is excluding studies that are written in German. Steiner (2001) and Hunt (1995) cite some of the studies in German and claim that they find broadly similar results.

unemployment duration literature in the context of the German labour market. Section 3 is devoted to the description of the data and sample used in the empirical exercise carried out in this study. Section 4 gives an account of the econometric specifications and methods of estimation used for the purpose of studying the duration of displacement unemployment. Section 5 discusses the estimation results obtained and the final section concludes the paper.

2. Literature Review

As stated earlier, there are few empirical studies that look into the duration of unemployment in Germany in general and that of the duration of displacement unemployment in particular.[3] Bender *et al* (2002) is one of the most recent studies that look into the duration of non-employment in Germany. Using the administrative social security data file (IAB) and focusing on separations due to plant closure and other types of reasons, they analyse the duration of time that displaced workers, those that left job due to plant closure, and other separators, such as those that were dismissed for cause, spend in non-employment. They find that displaced workers leave non-employment at a faster rate than workers who separated for other reasons. This finding is in line with previous findings (such as Gibbons and Katz, 1991) and is partly the result of the way in which they define displaced workers. As is generally the case, identifying displaced workers using administrative data, which relies on whether or not plants have been

[3] This might have to do with the fact that displaced workers in Germany do not experience a spell of unemployment as much as their counterparts in other less regulated labour markets such as the US do. Bender *et al* (2002) argue that provisions such as advance-notice in Germany are likely to reduce or eliminate the likelihood of experiencing a spell of unemployment following job loss.

closed, is likely to lead to selectivity problem. This definition ignores those workers that get displaced from declining but still operating plants. Although Bender *et al* (2002) go some distance by way of explaining the duration of non-employment that displaced workers experience, their study is different from this study in a number of ways. First, it looks at the duration of non-employment as opposed to the duration of unemployment, which this study is primarily about. Secondly, the way displaced workers have been identified is likely to suffer from the problem of selectivity bias. Finally, their study does not address the issue of unobserved heterogeneity, which has been found to play an important role in explaining the duration of unemployment that displaced and other unemployed workers experience. Heckman and Singer (1984) and Keifer (1988), among others, have shown that not accounting for unobserved heterogeneity gives rise to (downward) biased estimates of duration dependence.

Using the first twelve waves of the German Socio-Economic Panel (GSOEP) data, Steiner (1994; 2001) investigates whether or not there is unemployment persistence in the West German labour market. Steiner states that the persistence of high unemployment has been a serious problem in Germany for many years and argues that the rise in the unemployment rate has mainly been due to substantial increase in the share of the long-term unemployed - those who have already been unemployed for at least one year. Using discrete-time approach with flexible baseline hazard specified as random-effects logit and accounting for unobserved heterogeneity, Steiner tests whether or not unemployment persistence in the West German labour market can be explained by negative duration dependence or sorting. The estimation results indicate that declining hazard rates from unemployment are due to unobserved heterogeneity. He finds that once unobserved heterogeneity is accounted for, negative duration dependence in the employment hazard rate disappears. In fact, Steiner finds the unobserved heterogeneity controlled hazard rate for men to be positive.

Steiner (2001)'s study differs from this study for a number of reasons. First, the focus of his study is on all unemployment spells as opposed to displacement unemployment spells, which are the prime focus of this study. Secondly, the discrete-time logit specification used for the hazard of exit from unemployment represents a possible drawback of his study. Allison (1982) and Vermunt (1996) argue that this specification is sensitive to the choice of the length of the time intervals, and also necessitates that these intervals be of equal length. This is because the length of the time interval influences the probability that an event will occur in a particular interval and it, therefore, influences the hazard that the event of interest takes place in the interval. The complementary log-log specification is more appropriate in this context particularly when there are no time-covariate interactions and with proportional hazard specification. As will be discussed in Section 4 below, the complementary log-log specification has been used in this study and it is likely to give rise to better results. Third, unlike Steiner (2001), alternative specifications will be used in this study by way of accounting for unobserved heterogeneity. This is likely to give better results in terms of duration dependence and the effect of covariates on the hazard of exit from unemployment.

Couch (2001) investigates the duration of unemployment that displaced workers experience using the GSOEP data over the period 1988 – 1996. He estimates the annual months of unemployment that displaced workers experience using Tobit regression. Accordingly, the estimated number of months that displaced workers experience in the year of displacement range from .31 to .48 moths while no significant effect is found in the years preceding the year of displacement. The approach used in this study is vitally different from the one used by Couch. In this study, the hazard specification is used to assess the cost of job displacement in terms of the duration of unemployment displaced workers experience. As such, therefore, Couch's findings cannot be compared directly to findings of this study.

Hunt (1995) uses the GSOEP data over the period 1983 – 1988 to analyse the effect of unemployment compensation on unemployment duration in Germany. Using the Cox partial likelihood proportional hazards model, Hunt estimates competing risks of transitions to employment and inactivity for both men and women. The results from this study indicate that changes in the law which took the form of increasing the potential duration of unemployment insurance was found to be an important factor explaining differences in the patterns of exits to employment and inactivity for men and women in Germany and also why German unemployment spells are so much longer than American spells. Hunt (1995) focuses on overall unemployment spells and the effect of the policy change on the unemployment spells. As such, Hunt's study differs from this study, which focuses on displacement unemployment. Methodologically, Hunt (1995) uses the Cox partial likelihood proportional hazards model, which is more appropriate for continuous-time event history data (Vermunt, 1996). As will be detailed in Section 4 of this study, more appropriate methodology is used in this study to model the duration of unemployment for displacement and overall unemployment spells in Germany.

3. The Data and Sample

The data used in this study come from the German Socio-Economic Panel (GSOEP). The first fourteen waves of the GSOEP data covering the period 1984 to 1997 have been sued. In addition to the contemporaneous information collected at the interview date of each wave of the GSOEP, recall information is collected on labour market activities of respondents in each month in the preceding calendar year. Combining the contemporaneous/yearly information on labour market status of respondents with the recall/monthly information on labour market activity, a sequence of monthly labour market status has been constructed for each subject included in the duration analysis made in this study.

This data construction process has two main parts. The first part involves carefully selecting individuals on the basis of some sample selection criteria that were applied to the yearly information of the GSOEP. Accordingly, individuals of working age (16 – 64) from samples A and B of the GSOEP[4] that have been interviewed successfully at each wave over the period considered in this study have been chosen first. Then selection was

[4] Samples A and B are the initial samples of the GSOEP representing West Germans and foreigners residing in West Germany, respectively. The GSOEP has been extended to include more samples, such as Sample C representing East Germans that started in 1990. Details on the various samples of the GSOEP and other relevant information can be obtained in English from the GSOEP web site at http://www.diw-berlin.de/english/sop/index.html

made on the basis of some criteria. The first such criterion involves excluding individuals that reported to have worked in activities/industries such as agriculture that are regarded to represent heavy subsidisation over the years in a way that does not conform to the normal operation of the labour market. The second criterion involves excluding those individuals that were in self-employment. Individuals in this category too are not generally considered to make up an ideal sample for the purpose of studying the costs of job displacement stemming from the workings of the labour market. Individuals with gaps in observation are the next group of people that have to be excluded from the sample used for the duration analysis. After imposing these restrictions a final sample has been obtained by matching the different person and household level data files of the GSOEP.[5] The yearly panel constructed in this way consists of a total of 8,055 individuals that have been covered by the GSOEP in any year over the study period. Of this total, 2,824 have appeared in every wave over the period considered in this study.

The second part of the data construction process to get the sample of displacement unemployment spells used in this study involved merging the yearly panel briefly described earlier with the recall information from the calendar data files. This gave rise to an initial panel consisting of 6,081 individuals and 660,288 person-month observations. Two restrictions have been imposed to this initial sample. First, those individuals with gaps in person-month information have been removed from the sample. Second, those individuals with severe inconsistencies between their contemporaneous and recall labour

[5]The Desktop Companion to the German Socio-Economic Panel (GSOEP) study (available at http://www.diw-berlin.de/english/sop/service/dtc/dtc.pdf) gives a detailed account of how to match the different data files of the GSOEP. See Chapter 4 of the Desktop Companion, in particular.

market status were also eliminated from the sample. These two restrictions and the elimination of left-censored non-employment labour market status gave rise to the second panel consisting of 4,913 individuals and 508,565 person-month observations.

The next stage of construction of the sample of GSOEP unemployment spells involved copying the contemporaneous personal, household, regional and labour market information to each and every person-month observation using a series of rules.[6] Keeping only unemployment spells that were preceded by a spell of employment and with valid contemporaneous information on relevant covariates gave rise to the final sample of 1022 individuals and 16,620 person-month observations. Of these, 949 individuals with 13,974 person-month observations of unemployment spells make up the sample of displacement unemployment spells. The remaining 191 individuals with 2,646 person-month observations of unemployment spells represent non-displacement unemployment spells. The total number of people in the two samples of unemployment indicates that there are some people who experienced unemployment as a result of both job displacement and 'other' reasons during the sample period.

Table A1 in the appendix gives a summary of the variables used in the duration analysis made in this study. As can be seen from the table, most of the unemployment spells are the result of job displacement. Most of those that experience displacement unemployment are men, married and over 45 years of age. Most of the displaced unemployed are also

[6] The series of rules applied for the purpose of copying contemporaneous information to each person-month observation very much resembles that used by Upward (1999).

Germans residing in rented houses and with no/incomplete apprenticeship and/or higher-level qualification. Most of the displaced unemployed also have some health condition that hinders their day-to-day activity/work. Most also come from Western Germany which includes the regions North Rhine Westphalia; Hesse and Rhineland-Palatinate; as well as Saarland. Most of the displaced unemployed had skilled manual job and were working in large firms in the manufacturing industry in their previous employment. Moreover, well over 50 per cent of the displaced unemployed had some prior experience of unemployment.

4. Model specifications and methods of estimation

That the duration variable of interest to this study is measured to the nearest month means that the appropriate approach to modelling the duration of unemployment is the discrete-time hazard model. The estimation of discrete-time duration models requires expanded or person-period data set organised in such a way that there will be as many data rows for each individual in the sample as there are time intervals over which the individual in question is at risk of experiencing the event of interest - re-employment here. The creation of such expanded person-period (person-month) data is a crucial part of the discrete time hazard modelling exercise as it ensures that the likelihood contribution of each individual is properly accounted for (Jenkins, 1997; 2003).

Following Cox (1972), Prentice and Gloecker (1978) and Meyer (1990), the discrete time hazard of exiting the state of unemployment can be modelled using the discrete-time proportional hazards model. Accordingly, the hazard of re-employment in the jth month, $h(t_j)$, for individual i with a vector of covariates, $\mathbf{x}$, having spent t months in unemployment and given that re-employment has not occurred before t_{j-1} can be given by

$$h_{ij}(t_j \mid \mathbf{x}_i) = 1 - \exp(-\exp[\gamma_j(t) + (\mathbf{x}_i'\boldsymbol{\beta})]) \tag{1}$$

Rearranging the discrete-time hazard given in equation (1) gives what is known as the complementary log-log transformation of the conditional probability of exiting the state of unemployment at time t_j as

$$\ln(-\ln(1 - h_{ij}(t_j \mid \mathbf{x}_i))) = \mathbf{x}_i'\boldsymbol{\beta} + \gamma_{\mathrm{j}}(t) \tag{2}$$

Given this complementary log-log transformation, the parameter $\boldsymbol{\beta}$ is interpreted as the effect of covariates in $\mathbf{x}$ on the hazard rate of re-employment in interval j, assuming the hazard rate to be constant over the jth interval. The log-likelihood function for the sample of individuals used in this study can be given by

$$\begin{aligned}\log L &= \sum_{i=1}^{n}\sum_{j=1}^{t} y_{it} \log\left(\frac{h_{ij}}{1-h_{ij}}\right) + \sum_{i=1}^{n}\sum_{j=1}^{t} \log(1-h_{ij}) \\ &= \sum_{i=1}^{n}\sum_{j=1}^{t} \left[y_{it} \log h_{ij}(t \mid \mathbf{x}) + (1-y_{it})\log(1-h_{ij}(t \mid \mathbf{x}))\right]\end{aligned} \tag{3}$$

As stated earlier, it is well established in the duration literature that not accounting for unobserved heterogeneity might lead to biased estimates of the baseline hazard as well as the covariate effects on the hazard of exit from the sate of unemployment (Heckman and Singer, 1984; Lancaster, 1990). Taking this into account, an attempt has been made in this study to control for unobserved heterogeneity. The standard practice in the literature is to introduce a positive-valued random variable (mixture), *v*, into the hazard specification. In the context of the discrete-time approach, the augmented hazard function, which incorporates a multiplicative mixture term, is given by

$$h_{ij}(t_j, \mathbf{x}_i \mid v_i) = h_0(t_j)\exp(\mathbf{x}_i'\boldsymbol{\beta})\, v_i \tag{4}$$

The complementary log-log version of equation (4) is then given by

$$h_{ij}(t_j, \mathbf{x}_i \mid v_i) = 1 - \exp(-\exp(\mathbf{x}_i'\boldsymbol{\beta} + \gamma_j(t) + u_i)) \tag{5}$$

where, as before, $u_i = \log v_i$ and $\gamma_{\mathrm{j}}(t) = \int_{t_{j-1}}^{t_j} h_0(u)du$.

The discrete-time likelihood function that incorporates the unobserved heterogeneity term is obtained by summing the discrete-time likelihood functions of each individual *i* that can be given by

$$L_i(\beta,\gamma,\sigma) = \int_{-\infty}^{+\infty}\left[\prod_{j=1}^{t} h_j(t,\mathbf{x}_i \mid u_i)^{y_{ij}}[1 - h_j(t,\mathbf{x}_i \mid u_i)]^{1-y_{ij}}\right] g_u(u_i)du_i \tag{6}$$

where $h_j(t, \mathbf{x}_i \mid v) = 1 - \exp[-\exp(\mathbf{x}_i^{'}\boldsymbol{\beta} + \gamma_j(t) + u_i)]$ and σ is the vector of unknown parameters in $g_u(u)$. The unobserved heterogeneity term is assumed to be independent of observed covariates, $\mathbf{x}_i$, and the random duration variable, *T*, and have density $g_v(v)$. It is possible to solve the integral in expression (6) to obtain the appropriate density for the monthly duration information that is organised in a sequential binary response format (Stewart, 1996; Andrews, *et al*. 2002; Dolton and van der Klaauw, 1995; Wooldridge, 2002).

Solving the mixing distribution specified in equation (6) necessitates making specific distributional assumption for the density of the mixing distribution. The distributional assumption may either be parametric or non-parametric. The parametric approach specifies a particular functional form for the mixing distribution.[7] The non-parametric approach, on the other hand, uses the mass point approach pioneered by Heckman and Singer (1984), where the mixing distribution is approximated by a finite discrete distribution of unrestricted form. In the absence of theoretical justification for using one or the other approach for the purpose of approximating the mixing distribution, it may be reasonable to try and employ both approaches. The parametric distribution assumed in this study in order to approximate the mixing distribution is the Gaussian distribution[8]

[7] There are several candidates for the parametric mixing density distribution. However, the choice of a particular parametric distribution is generally harder to justify than the choice of functional form for the baseline hazard. This is due to the fact that economic theory may suggest a particular functional form for the baseline hazard but not for the mixing distribution (Van den Berg, 2001).

[8] An attempt has been made to estimate Gamma mixture model using PGHAMZ. However, the model fails to converge when used to estimate the mixing distribution for the sample of all unemployment spells. As a result estimates based on the Gamma mixing model are not reported. This is rather unfortunate given that the Gamma distribution is the most commonly used distribution in the literature due to the fact that it is

while the non-parametric approach follows the mass pint technique of Heckman-Singer. The Gaussian distribution does not yield a closed form solution. However, its use is justified if one views the heterogeneity term as being a combination of a 'vast number of minor characteristics of the unemployed individual that are not observed by the investigator' (Stewart, 1996). In the case of the non-parametric technique, the unknown distribution of the unobserved heterogeneity is approximated using discrete distribution. The mass points and the associated probabilities of the discrete distribution are estimated jointly with other parameters of the model.[9]

The second issue of importance in relation to estimating alternative models has to do with the way the baseline hazard is specified. The baseline hazard can be specified either parametrically or semi-parametrically. In the case of parametric specification for the baseline hazard, a particular functional form is assumed for the same. Although there is no strong theoretical justification for it, the Weibull is the commonly used parametric specification for the baseline hazard in the unemployment duration literature. Taking this into account, the first model estimated assumes Weibull for the baseline hazard and this variant is estimated with and without consideration for unobserved heterogeneity. Semi-parametrically, the baseline hazard is estimated together with other parameters of the model. Imposing a particular functional form for the baseline hazard may lead to the problem of misspecification. A way round this possible problem is to estimate the

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analytically tractable and gives closed form solution for the relevant likelihood function (Lancaster 1990; Meyer 1990; Stewart 1996; Van den Berg 2000).

[9] See Stewart (1996) and Andrews *et al* (2001) for the likelihood functions of the Gaussian. For the likelihood function of the non-parametric mixing distribution, see Stewart, 1996; Van den Berg, 2001; Andrews, *et al*., 2002.

baseline hazard semi-parametrically in line with Han and Hausman (1990), Meyer (1990, 1995) and others. Because there are no events in some months, the monthly time intervals in this study had to be regrouped into just seven time periods for the sake of identification. The piece-wise constant baseline hazard specification is therefore the preferred non-parametric specification for the baseline hazard estimated in this study.

5. Estimation Results and Discussion

In this section discussion of results from the estimation exercise will be made. The first set of results in this study is that which is based on the Weibull specification for the baseline hazard while the second set of results is from the piece-wise constant baseline hazard specification. Results from the Weibull model are given in Table 1. Results from the piece-wise constant model are given in Table 2 and Table 3. In what follows discussion of these results will be made.

5.1. Results from Weibull model

The first set of estimation results is from the most commonly used Weibull specification for the baseline hazard. The Weibull model imposes a particular monotonic shape for the baseline hazard. Both homogeneous and mixing discrete-time Weibull models have been estimated. The mixing model estimated assumes that the heterogeneity term is distributed normally. As can be seen from the estimation results in Tables 1, the estimated coefficients in the mixing models are slightly large in absolute value terms. Likelihood ratio test of zero unobserved heterogeneity is also rejected decisively indicating the importance of accounting for unobserved heterogeneity. The coefficient on $ln(t)$ in the context of discrete-time duration models is an estimate of the parameter describing the baseline hazard. The estimated coefficient shows that the baseline hazard declines with time, indicating negative duration dependence. In other words, the longer that an unemployed individual stays unemployed, the more difficult it will be (for the individual) to leave the state of unemployment. Accounting for unobserved heterogeneity does make a difference in the sense that the parameter describing the baseline hazard is less negative in the mixing Weibull model than its homogenous counterpart. This indicates that although accounting for unobserved heterogeneity does not eliminate negative duration dependence completely as in Steiner (2001), it is quite important in explaining whether or not there is negative duration dependence in the re-employment pattern of displaced workers.

Referring to the estimated coefficients of the covariates included in the models reveal how they affect the hazard of re-employment. Accordingly, older displaced workers are found to have a lower hazard of re-employment compared with their younger counterparts. In particular, those displaced unemployed that are over 45 years of age have a 62 percent less hazard of exit to re-employment compared with those that are between 30 and 45 years of age. On the other hand, those displaced unemployed that are less than 30 years of age have a higher hazard of exit to re-employment compared with their older counterparts who are between 30 and 45 years of age. These findings are in line with expectation. Older people are more likely to receive fewer job offers given that they are a less attractive investment for firms that seek to invest in younger workers that are capable of working for longer periods. Young workers are also more capable to learn new skills that best suit changing demand situations compared with their older counterparts who may be relatively less suited when it comes to learning new tricks. Also older workers may decline to accept more jobs than their younger counterparts for various reasons. Mobility problem as a result of family and other responsibilities, for example, may force older people to reject some job offers.

Men who are unemployed due to displacement have a 56 percent higher hazard of re-employment compared with women. There can be different explanation for this. To start with, men account for more than 60 per cent of the displaced unemployed as can be seen from Table A1 in the appendix. That men are dominantly represented in the sample of displacement unemployment spells should mean that they face a greater risk of exit from unemployment. Another explanation relates to the type of previous jobs that the displaced unemployed had. More than 70 per cent of previous jobs left are manual type, mostly involving men. Assuming re-employment to more or less similar type of employment as before, it would not be surprising to find that men have a higher hazard of exit to re-employment. Another explanation that best fits the labour economics literature is, of course, that which relates to the gender difference in the labour market behaviour of workers. Men are generally expected to receive more job offers than women do mainly due to the labour market behaviour of women that is characterised by (or perceived to be) frequent interruptions.

Those that are married displaced unemployed have a 24 percent less hazard of getting re-employed compared with their single counterparts. Germans who lost their job as a result of displacement have a 35 percent higher hazard of getting re-employed compared with EU nationals residing in Germany. On the other hand, foreigners are found to have a lower hazard of exiting via re-employment although this effect is found to be statistically insignificant. This means that non-Germans have a longer unemployment duration compared with Germans. This is in line with expectation given that Germans are likely to receive more job offers vis-à-vis their non-German counterparts. It should also be noted that Germans are dominantly represented in the sample of unemployment spells. Those displaced unemployed who have some health condition/problem have a 27 percent less hazard of getting re-employed. Those displaced unemployed that do not have own dwelling have a 16 percent less hazard of securing re-employment. Such workers are more likely to have a higher offer acceptance rate, and the lower hazard of getting re-employed should stem largely from lower arrival rate of job offers.

Type of qualification of the unemployed is found to be an important factor explaining the hazard of exit from the state of unemployment. Accordingly, those displaced workers without apprenticeship and/or college level training who experienced a spell of unemployment following job displacement have a 23 percent less hazard of being re-employed compared with their counterparts with completed apprenticeship and/or college level training. This finding is in line with what one would expect. Those with the least qualification are less likely to receive many job offers and hence are less likely to exit unemployment via re-employment. It can of course be the case that individuals with the least qualification are more likely to have lower reservation wage and hence are more likely to accept offered jobs. The net effect therefore depends on which effect is stronger. In this case, the result seems to imply that the former effect is stronger.

The estimated results also suggest strong regional variation in the patterns of exit from unemployment via re-employment for the unemployed. Those displaced unemployed in Northern and Western Germany have longer duration of unemployment compared with their counterparts in the south of the country. In particular, those workers in the north and west of the country have a 29 percent lower hazard rate of re-employment compared with their counterparts in Southern Germany. The regional variables used here serve as proxy for local labour market conditions that are usually captured using local unemployment and vacancy rates. Although the region variables hide lots of variations that may exit among the 10 regions that the GSOEP samples come from, this result can be interpreted in terms of the respective unemployment and vacancy rates in the regional groups considered here.

Previous job and labour market history related covariates are the other most important factors explaining the hazard of exit from unemployment. Accordingly, those displaced unemployed individuals who had unskilled manual job have longer durations of unemployment compared with their counterparts who had managerial, technical or professional job. In terms of the hazard of exit to re-employment, these workers have a 35 percent lower hazard of re-employment compared with their counterparts. Another interesting result is that those who had been working in small size firms have a higher hazard of finding re-employment vis-à-vis those that had been working in large firms. Focusing on the displaced unemployed, those who had been working in small and medium size firms have 32 percent and 24 percent higher hazards of re-employment, respectively, compared with those who had been working in large size firms. This could be attributed to the strong possibility that those with some experience working for smaller firms are more mobile given that there are relatively more firms of the smaller/medium type. Institutional factors such as the presence of labour unions and labour protection schemes in place may tend to be stronger in large firms, making re-employment in the same (similar size) firm difficult.

Those who got displaced from the manufacturing as well as the trade, transport and communication industries have a lower hazard of exit via re-employment. This is likely to be the result of fewer job offers coming from the manufacturing sector, in particular, which is generally regarded as a sector in decline. Those who were trade union members in their previous job are found to have a lower hazard of exit via re-employment. Specifically, displaced workers who were trade union members have a 17 percent lower hazard of securing re-employment compared with displaced unemployed workers who were not trade union members. This might be attributed to a relatively lower offer acceptance rate that former trade union members may have. Such workers might have been getting some wage premium in their previous job and could have a higher reservation wage. Another interesting result is that those who got displaced and had been previously unemployed are likely to experience longer unemployment duration compared with other displaced unemployed individuals with no prior unemployment experience. Previous experience of inactivity also reduces the hazard of exit from unemployment.

5.2. Results from non-parametric models

As discussed earlier, parametric models such as the Weibull assume a particular shape for the baseline hazard. Assigning a particular shape for the baseline hazard may prove to be a major shortcoming in duration analysis. In the context of proportional hazards models, a number of studies including Meyer (1990); Han and Hausman (1990) and Trussell and Richard (1985) have shown that assigning a specific parametric density for the baseline hazard can lead to a more serious problem of misspecification than that caused due to disregarding unobserved heterogeneity. In the face of such evidence, it is reasonable to adopt semi-parametric specification for the baseline hazard. Such specification has an additional advantage in that parameter estimates will be less sensitive to the distributional assumptions made for unobserved heterogeneity. A likelihood ratio test comparing differences in likelihood scores between the Weibull and the piece-wise constant models also suggest significant improvements in fit of the piece-wise constant models. Given these, the estimation results presented in Tables 2 and 3, which are based on piece-wise constant specification for the baseline hazard, are the preferred results explaining the duration of displacement unemployment in Germany.

The first piece-wise constant baseline hazard model estimated is the homogeneous proportional hazards model following Andrews *et al* (2001), which rules out unobserved heterogeneity. Then two mixing proportional hazards models have been estimated. The mixing proportional hazards models estimated are the Gaussian mixing[10] and the non-parametric Heckman-Singer mixing models.[11] Estimating alternative mixing distributions enable us to assess the sensitivity of estimated parameters across the models and whether or not unobserved heterogeneity is worth considering.

5.2.1. Baseline hazards

Estimates of piece-wise constant baseline hazards from the homogeneous and the two mixing distributions are given in Table 2. These results are obtained using the complementary log-log transformation given in equations (1) above by setting all covariate values equal to zero.[12] The first important result worthy of a note has to do with the rejection of the null hypothesis of zero unobserved heterogeneity. The likelihood ratio test of zero unobserved heterogeneity for the mixing distributions is strongly rejected with a P-value of almost zero. As can be seen from the estimated results in Table 2, the estimated baseline hazards are strikingly similar for the Gaussian and Non-parametric distributions, lending support to Meyer (1990)'s suggestion that using a flexible specification for the baseline hazard removes the sensitivity of estimated parameters to

[10] The default number of quadrature points in STATA is 12. The default quadrature points have been used in this study but checks have been made using quadrature check and the results remain more or less the same.

[11] The empirical estimation of the discrete mass point approach made in this study has been carried out using GLLAMM. GLLAMM is a computationally efficient program that fits a large class of multilevel latent variable models including multilevel generalised linear mixed models (Rabe-Hesketh *et al*., 2002)

[12] As stated earlier, the monthly time period has been regrouped to get only seven time periods The re-arranged time periods are: Month 1; Months 2-3; Months4-6; Months7-9; Months10-12; Months13-18 and Months >18.

the type of distribution assumed for unobserved heterogeneity. For the homogeneous distribution, the baseline hazard estimates are higher than those obtained using the mixing distributions for the first two time periods. After the second period, however, the baseline hazard estimates from the homogeneous model are found to be less than those of the mixing distributions. These patterns are shown in figure 1 below where the plot of the homogeneous baseline hazard drops faster than those of the mixing distributions. Although there are some differences in the magnitude of the estimated hazards from the three models, the general patterns observed are more or less similar. Accounting for unobserved heterogeneity does reduce the observed (negative) duration dependence. As such, therefore, these results are in line with the common claim in the literature that accounting for unobserved heterogeneity reduces negative duration dependence.

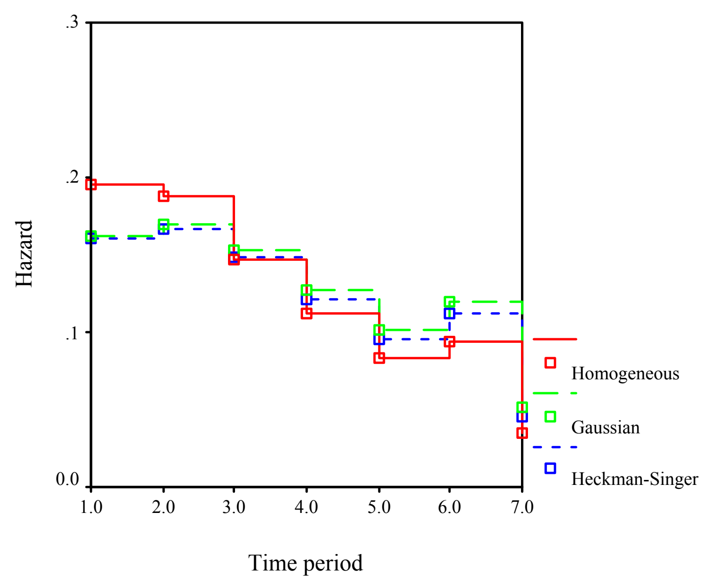

Figure 1: Re-employment baseline hazard, displacement unemployment

Baseline hazard estimates for the homogeneous model exhibit more or less continuous decline in the estimated hazards of exit via re-employment. Results from the mixing distributions, on the other hand, tell a slightly different but more appealing story. For the displacement unemployment sample, the hazard estimates form the mixing distributions increase initially and then decline more or less continuously afterwards. This might indicate that displaced unemployed workers are likely to find re-employment more difficult if they fail to secure re-employment in the first three months following displacement.

5.2.2. Effects of covariates

Results showing the estimated effects of covariates on the hazard of re-employment are given in Table 3. The effects of covariates on the hazard of exit via re-employment are more or less similar across the three models estimated, with only marginal differences. The results from the non-parametric models also show some similarity to the earlier results from the Weibull model. Comparing the maximum of the log-likelihoods from the piece-wise constant models shows that the Gaussian model has an edge over the other two models. As a result, discussion of the covariate effects on the hazard of re-employment made below relies on the Gaussian model.

Starting with the effect of personal characteristics on the hazard of exit via re-employment, older displaced workers have a 59 percent lower hazard of re-employment compared with those that are between 30 and 45 years of age. On the other hand, those displaced unemployed individuals that are less than 30 years of age have a 22 percent higher hazard of re-employment compared with the reference group of displaced unemployed individuals between 30 and 45 years of age. Same reasoning as given earlier in relation to results from the Weibull model can be given here in relation to these results.

Men who are unemployed due to displacement have a 53 percent higher hazard of re-employment compared with their women counterparts. Those that are married displaced unemployed have a 21 percent lower hazard of re-employment. Germans have a higher hazard of re-employment while there is hardly a difference between the re-employment hazards of foreigners and EU nationals residing in Germany. Displaced unemployed individuals with some health problem have a longer duration of unemployment compared with their counterparts with no such problem. Those displaced unemployed that do not own their own dwelling have a 16 percent lower hazard of re-employment. In terms of the type of qualification of the unemployed, those displaced unemployed workers without apprenticeship and/or college level training have longer duration of unemployment compared with displaced unemployed workers with completed apprenticeship and/or college level training. In terms of region, those residing in Northern and Western Germany tend to have longer unemployment duration compared with their counterparts in the South of the country.

As before, previous job and labour market history related covariates are important determinants of the re-employment hazard. Accordingly, those displaced unemployed individuals who had unskilled manual job have longer duration of unemployment with a 33 percent lower hazard of re-employment compared with their counterparts who had managerial, technical or professional job. In comparison, those displaced workers who were working skilled non-manual job have a 25 percent lower hazard of re-employment. As in the earlier result, firm size too has been found to have an important role in explaining the hazard of exit via re-employment. Accordingly, those who had been working in small and medium size firms have shorter unemployment durations with a 28 percent and 18 percent higher hazard of re-employment compared with their counterparts who had been working in large firms.

Workers who got displaced from the manufacturing and the trade, transport and communication industries have longer duration of unemployment with a 25 percent and 23 percent lower hazards of re-employment, respectively, compared with those previously working in the finance, insurance and services industry. On the other hand, those who were working in the mining, energy and construction industry have a 13 percent higher hazard of exiting unemployment for job. Those who were trade union members in their previous job are found to have a lower hazard of exit from unemployment. Specifically, displaced workers who were trade union members have 14 percent lower hazard of re-employment compared with displaced unemployed workers who were not trade union members. Another interesting result is that those who got displaced and had been previously unemployed have 15 percent lower hazard of re-employment compared with other displaced unemployed individuals with no prior unemployment experience. On the other hand, those with previous experience of inactivity have an even longer duration of unemployment. Accordingly, those displaced unemployed individuals who had been out of the labour market at some point in the past have a 19 percent lower hazard of re-employment.

6. Conclusion

This paper attempted to study the duration of unemployment in Germany as part of the drive to establish on the costs, in terms of unemployment, that displaced workers experience. The focus of the study has been on unemployment spells that were initiated as a result of job displacement as a result. Parametric and non-parametric discrete-time models have been used to study the duration of displacement unemployment spells. In addition, alternative mixing distributions have also been employed to account for unobserved heterogeneity. The results obtained indicate that there is evidence of negative duration dependence in the hazard of exit via re-employment. Accounting for unobserved heterogeneity does matter, but the main finding of this study with regards to duration dependence remains unchanged.

With regards to the effect of covariates on the hazard of re-employment, those displaced workers who are at the lower end of the skills ladder, those who had been working in the manufacturing industry and those with previous experience of inactivity are found to have particularly lower hazard of exit via re-employment. The fact that those at the lower end of the skills ladder and those with previous experience of inactivity have difficulty exiting unemployment calls for appropriate intervention to ameliorate the condition of the 'disadvantaged'.

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Table 1: Re-employment hazard of displacement unemployment, proportional hazards models

	Weibull (Homogeneous)			Weibull (Mixing)		
	Coeff.	Risk ratio	P-value	Coeff.	Risk ratio	P-value
Personal and region variables						
Age <=30	0.162	1.176	0.028	0.213	1.237	0.031
Age >45	-0.810	0.445	0.000	-0.975	0.377	0.000
Male	0.322	1.380	0.000	0.446	1.562	0.000
Married	-0.153	0.858	0.119	-0.278	0.758	0.027
Married & Children < 16 in the house	0.142	1.153	0.314	0.208	1.231	0.250
German	0.249	1.282	0.015	0.303	1.354	0.030
Foreigner	-0.043	0.958	0.675	-0.021	0.979	0.883
Has some health problem	-0.277	0.758	0.000	-0.310	0.733	0.000
Tenant	-0.168	0.845	0.017	-0.173	0.842	0.072
Children under 16 in the household	-0.076	0.927	0.493	-0.094	0.910	0.528
No apprenticeship or college	-0.114	0.893	0.112	-0.256	0.774	0.010
Northern Germany	-0.303	0.739	0.000	-0.346	0.708	0.003
Western Germany	-0.357	0.699	0.000	-0.348	0.706	0.000
Previous job & labour market history						
Unskilled manual	-0.419	0.657	0.003	-0.435	0.647	0.016
Skilled manual	-0.080	0.923	0.511	-0.033	0.967	0.836
Skilled non-manual	-0.306	0.737	0.011	-0.293	0.746	0.068
Small firm	0.219	1.244	0.007	0.279	1.322	0.008
Medium firm	0.201	1.223	0.013	0.218	1.244	0.036
Mining, energy and construction	0.167	1.182	0.126	0.114	1.120	0.436
Manufacturing	-0.282	0.754	0.006	-0.310	0.733	0.018
Trade, Transport & communication	-0.290	0.749	0.011	-0.291	0.747	0.050
Trade union member	-0.101	0.904	0.134	-0.187	0.829	0.054
Previously unemployed	-0.161	0.851	0.009	-0.127	0.880	0.082
Previous experience of inactivity	-0.167	0.847	0.098	-0.198	0.821	0.122
Ln(t)	-0.609	0.544	0.000	-0.232	0.793	0.002
Constant	-1.024		0.000	-1.322		0.000
Variance				0.542		0.000
No of observations/groups	1187			949		
No of person-period observations	13974			13974		
Log-likelihood	-3545.9			-3512.5		

Table 2: Re-employment baseline hazard estimates, displacement unemployment

	Homogeneous			Gaussian			Heckman-Singer		
	Coef. Est.	Hazard	P value	Coef. Est.	Hazard	P value	Coef. Est.	Hazard	P value
Month1	-1.223	0.196	0.000	-1.399	0.162	0.000	-1.355	0.161	0.000
Months2-3	-1.274	0.188	0.000	-1.349	0.170	0.000	-1.316	0.167	0.000
Months4-6	-1.539	0.147	0.000	-1.465	0.153	0.000	-1.446	0.148	0.000
Months7-9	-1.836	0.112	0.000	-1.666	0.127	0.000	-1.659	0.121	0.000
Months10-12	-2.131	0.084	0.000	-1.901	0.102	0.000	-1.904	0.096	0.000
Months13-18	-2.022	0.094	0.000	-1.731	0.119	0.000	-1.744	0.112	0.000
Months>18	-3.041	0.035	0.000	-2.617	0.051	0.000	-2.668	0.046	0.000

The baseline hazard is computed as $h(t) = 1 - \exp(-\exp(\gamma(t)))$

Table 3: Re-employment hazard of displaced unemployed individuals, proportional hazards models

	Homogeneous			Gaussian Mixing			Heckman-Singer Mixing		
	Coefficient	Risk ratio	P-value	Coefficient	Risk ratio	P-value	Coefficient	Risk ratio	P-value
Personal and region variables									
Age <=30	0.153	1.166	0.038	0.196	1.216	0.037	0.211	1.235	0.027
Age >45	-0.753	0.471	0.000	-0.888	0.412	0.000	-0.794	0.452	0.000
Male	0.332	1.393	0.000	0.425	1.530	0.000	0.394	1.483	0.000
Married	-0.144	0.866	0.143	-0.242	0.785	0.044	-0.244	0.784	0.040
Married & Children < 16 in the house	0.142	1.153	0.313	0.195	1.215	0.256	0.170	1.185	0.317
German	0.233	1.262	0.022	0.277	1.319	0.034	0.238	1.268	0.069
Foreigner	-0.032	0.968	0.751	-0.017	0.983	0.898	-0.059	0.943	0.656
Has some health problem	-0.254	0.775	0.000	-0.283	0.753	0.000	-0.290	0.748	0.000
Tenant	-0.169	0.845	0.016	-0.169	0.844	0.064	-0.174	0.840	0.043
Children under 16 in the household	-0.076	0.927	0.496	-0.099	0.906	0.482	-0.055	0.946	0.694
No apprenticeship or college	-0.104	0.902	0.145	-0.217	0.805	0.020	-0.243	0.785	0.012
Northern Germany	-0.311	0.733	0.000	-0.333	0.717	0.002	-0.287	0.751	0.007
Western Germany	-0.339	0.712	0.000	-0.335	0.715	0.000	-0.318	0.728	0.000
Previous job & labour market history									
Unskilled manual	-0.399	0.671	0.004	-0.409	0.665	0.017	-0.471	0.625	0.007
Skilled manual	-0.076	0.927	0.532	-0.039	0.962	0.795	-0.063	0.938	0.676
Skilled non-manual	-0.299	0.742	0.013	-0.282	0.754	0.061	-0.301	0.740	0.045
Small firm	0.203	1.225	0.012	0.250	1.284	0.012	0.250	1.284	0.009
Medium firm	0.198	1.218	0.014	0.209	1.232	0.034	0.167	1.182	0.083
Mining, energy and construction	0.169	1.185	0.119	0.121	1.129	0.377	0.113	1.120	0.406
Manufacturing	-0.272	0.762	0.008	-0.294	0.745	0.018	-0.258	0.773	0.040
Trade, Transport & communication	-0.254	0.776	0.025	-0.267	0.766	0.057	-0.218	0.804	0.113
Trade union member	-0.086	0.917	0.199	-0.151	0.860	0.096	-0.119	0.888	0.191
Previously unemployed	-0.188	0.829	0.002	-0.158	0.854	0.027	-0.151	0.860	0.032
Previous experience of inactivity	-0.195	0.822	0.053	-0.210	0.811	0.083	-0.214	0.807	0.069
Variance				0.376		0.000	0.287		
Mass point 1 location							-0.425		

Mass point 1 probability			0.614
Mass point 2 location			0.676
Mass point 2 probability			0.386
No of observations	1187	1187	1187
No of person-period observations	13974	13974	13974
Log-likelihood	-3511.1	-3491.5	-3493.8

Appendix

Table A1: Descriptive statistics by type of unemployment spell

Variables	Displacement Unemployment		All Unemployment	
	Mean	Std. Dev.	Mean	Std. Dev.
Personal & household				
Age <=30	0.202	0.402	0.210	0.407
Age 31-45	0.320	0.466	0.329	0.470
Age >45	0.478	0.500	0.461	0.499
Male	0.613	0.487	0.589	0.492
Female	0.387	0.487	0.411	0.492
Married	0.747	0.435	0.741	0.438
German	0.474	0.499	0.502	0.500
EU National	0.179	0.383	0.181	0.385
Non-EU foreigner	0.347	0.476	0.317	0.465
No apprenticeship/College	0.500	0.500	0.469	0.499
Hindered by health	0.420	0.494	0.438	0.496
Tenant	0.751	0.432	0.739	0.439
Children under 16 in the household	0.446	0.497	0.437	0.496
Region of residence				
Northern Germany	0.210	0.407	0.232	0.422
Western Germany	0.428	0.495	0.412	0.492
Southern Germany	0.363	0.481	0.357	0.479
Previous job and labour market history				
Previous job unskilled Manual	0.206	0.404	0.193	0.394
Previous job skilled Manual	0.512	0.500	0.491	0.500
Previous job skilled Non-manual	0.222	0.416	0.243	0.429
Previous job Managerial, tech or profess.	0.060	0.237	0.073	0.261
Previous experience of unemployment	0.576	0.494	0.570	0.495
Previous experience of inactivity	0.151	0.358	0.168	0.374
Previous firm/industry of employment				
Previously in small firm	0.319	0.466	0.316	0.465
Previously in medium firm	0.306	0.461	0.294	0.456
Previously in large firm	0.375	0.484	0.390	0.488
Previously in Mining, Energy & Const.	0.169	0.375	0.154	0.361
Previously in Manufacturing	0.502	0.500	0.494	0.500