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Essays on Unemployment Duration in Germany

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Abstract

Unemployment duration is critical when evaluating economic policies. In my dissertation, I examine: i) the gender differences in unemployment duration, ii) the effects of housing tenure on unemployment duration, and iii) the differences in determinants of unemployment between East and West Germans. Despite the evidence of gender differences in duration of unemployment, its determinants remain not well understood. This paper aims to address the gap in the literature on determinants of unemployment by studying the differences in unemployment duration between genders and within-gender groups of East and West Germans. I construct unemployment spells from the German Socio-Economic Panel for 1990–2011 to shed more light on the differentials and their association with a partner and children. A simple comparison of median durations and survival functions shows that women's unemployment spells are twice as long as men's, and that these differences nearly triple during the child–bearing and –rearing ages, regardless of location. I find that the differences in unemployment duration between genders are associated with the presence of young children and a partner. Women's unemployment hazard rates are between 40 and 80 percentage points lower in the presence of children aged 0–4 than comparable men's hazard rates. In addition, West German men and women have hazard rates 20 and 25 percentage points lower than comparable East Germans, even after controlling for non-labor income. I find that the differential in unemployment duration between Eastern and Western females is associated with the differential effect of young children and a partner. These findings are robust after accounting for unobserved heterogeneity.

Homeownership is associated with positive outcomes such as urban sprawl, enacting zoning laws, and lowering crime rates, but there is also evidence of homeownership reducing labor mobility. This paper examines the effects of homeownership on unemployment lengths in Germany. I use unemployment spells from the German-Socio Economic Panel for the period 1990–2011 to estimate the effect of homeownership on unemployment length while controlling for possible self-selection into homeownership. The data suggests that homeowners experience two months shorter duration of unemployment which translates to a 13 percent higher rate of exits from unemployment after controlling for observables. However, after controlling for self-selection into homeownership, I find that homeownership significantly decreases the exit rate out of unemployment by 56 percent compared to renters.

Since the unification of East and West Germany, East Germany underwent economic and institutional transitions which were followed by soaring and persistent unemployment. Previous literature shows that there are tenacious differences between East and West Germans in unemployment duration and its determinants during first nine years after unification. In this paper, I use the data from the German Socio-Economic Panel to shed light on the differences in duration of unemployment and its determinants for 1990–2012 period. My findings show that there are differences between East and West Germans in unemployment duration and its determinants, in particular, education and age, during the early periods of transition. Easterners have shorter spells of unemployment, and the difference is especially large during longer spells. However, I find that these differences disappear over time. In addition, I find evidence of an emergence of the educational differential between East and West Germans with a higher levels of education shortening East German unemployment duration more than West German unemployment durations.

Dedication

I dedicate my dissertation work to Mirsad and Mirsada (my parents), Samir (my brother), and Aida (my wife-to-be) who have been a constant source of support and encouragement during the challenges of graduate school and life.

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Chapter 1

Explaining Gender Differences in Unemployment Duration: Evidence from Reunified Germany

1.1 Introduction

Understanding unemployment duration is critical when evaluating economic policies such as unemployment, especially in countries with extensive safety nets. Previous literature focuses on the effects of unemployment insurance on unemployment duration.[1] Furthermore, the evidence shows that longer unemployment lengths have long-term negative effects on future outcomes (Ruhm, 1991; Raphael and Ebmer, 2001, Arulampalam; 2001; Mroz and Savage, 2006), and that females experience longer unemployment spells than males (Podgursky and Swaim, 1987; Farber et al., 1997; Abbring et al., 2002; Keltzer and Fairlie, 2003; Hu and Taber, 2011). Recent empirical evidence sheds some light on the gender differential in duration being associated with the presence of young children in the U.S. (Kunze and Troske, 2012), but this differential is still not well understood.

In this paper, I study differences in unemployment duration between genders, as well as

[1] These are some examples of literature studying the effects of the unemployment insurance on duration for the U.S.: Meyer (1990), Katz and Meyer (1990), Card and Levine (2000); for Europe: Hunt (1995), Røed and Zhang (2003), Lalive (2008), van Ours and Vodopivec (2006), Card et al. (2007), Caliendo et al. (2013), Card et al. (2015).

within-gender differences between East and West Germans.[2] I use periods after reunification when safety nets and other formal institutions have been homogenized between East and West Germany. This guarantees that differences in the welfare state and other formal institutions are not responsible for any observed differential in unemployment duration between East and West Germans. Existing evidence indicates differences in post-unified labor market outcomes between East and West Germans such as wages, likelihood of employment, or receiving a college degree (Krueger and Pischke, 1992; Hunt, 2002; Hunt, 2004; Riphahn and Trübswetter, 2013; Mesella and Fuchs-Schündeln, 2015). Thus, in this paper I explore whether duration of unemployment is different between East and West Germans.

In Section 1.2, I describe the construction of the sample and perform a preliminary analysis of gender and within-gender differences in unemployment duration. I construct unemployment spell data using the German Socio-Economic Panel (GSOEP) for 1991–2011. The first advantage of the GSOEP over official German administrative data is the ability to identify a location of origin prior to reunification. The second advantage of my data set is the availability of information on non-labor income (including public transfers and partners' wages), as well as data on the presence of spouses and young children, which might be associated with the differential in unemployment length between genders. My sample consists of individuals in their prime working years, ages 18–55, from the moment they transition to unemployment until the time at which they either exit unemployment to work or are censored. The prime–age workers are more similar with respect to decisions affecting their job search than, for example, workers reentering the labor market after inactivity, workers coming out of education, or workers preparing to end their working life.

I find that female unemployment spells are longer than male spells. The median duration of unemployment for females is 14 months, compared to 7 months for males. Examining duration across age cohorts indicates that the largest difference in duration of unemployment between genders is during the childbearing and childrearing ages (25–35). Furthermore, female median durations are up to three times longer in the presence of young children when compared to either male durations with children or childless female durations. In addition, I find that both East German males and females have median durations shorter by one month than comparable West Germans of the same gender. For East and West German males, the median durations within and across age cohorts are similar, while I find that East German females have shorter median durations than West German females

[2] I identify whether an individual is East or West German by domicile in 1989, just prior to reunification.

across all age cohorts. However, I find that the difference in unemployment durations between East and West German females are only significant during childbearing and childrearing ages. These findings indicate that the gender differences in unemployment duration might be related to the role of children, and that the role of children might not be the same across East and West Germans.

In Section 1.3, I present my empirical strategy that examines the differentials in the duration of unemployment while taking into account the role of other determinants, such as non-labor income or local labor market conditions. My empirical model is based on Mortensen's (1986) seminal work on job search, where the probability of exiting unemployment is a product of the probability of receiving a job offer and the probability of then accepting that job offer. Since my primary interest is the effects of determinants, I use the Extended Cox model to estimate effects of a partner and children on the probability of leaving unemployment while controlling for observed and unobserved heterogeneity, as well as censoring.[3] This semi-parametric hazard model is a reasonable compromise between the Kaplan-Meier non-parametric model, and the possibly excessively structured parametric models. Additionally, I use time-varying explanatory variables to relax restrictive assumptions about the proportionality between hazards in the Cox Proportional Hazard model.

In Section 1.4, I present estimates for the differences in unemployment between genders and between East and West Germans. I find individuals older than 46 have significantly lower hazards regardless of gender, while the differential in unemployment duration between genders is only present in age groups 18–25, 26–30 and 31–35. Females in these cohorts are approximately 35, 37, and 32 percentage points less likely to leave unemployment than comparable males, *ceteris paribus*. However, this differential between genders during ages 18–35 ceases to exist after I control for the presence of children and a partner. I find that females with children ages 0–1 and ages 2–4 have hazards which are 80 and 40 percentage points lower, respectively, than those of comparable males. In addition, I find that females have lower hazards in the presence of a partner regardless of the partner's employment status.

The results also indicate that East German males and females have probabilities of leaving unemployment spells that are 20 and 25 percentage points higher, respectively, than West Germans of the same gender. Furthermore, the results show that these gaps are not due to the differences in non-labor income. My estimates indicate that 22 percent of the gap in unemployment hazards between East and West German females is due to the presence of children ages 2–4 in the household.

[3]The Extended and the Proportional Cox models differ in that former allows time–varying regressors.

After controlling for the presence of a partner and young children in the household, I find that the gap between unemployment hazards of East and West German males diminishes; however, I find that the effects of these determinants are not significantly different between East and West German males.

Duration models are sensitive to the presence of unobserved heterogeneity, which can contribute to negative duration dependence (Heckman and Singer, 1984b). In my main results, the baseline is allowed to vastly vary with individual unemployment duration which accounts for some heterogeneity between individuals. In addition, in Section 1.5, I check the robustness of the main results to the presence of individual unobserved heterogeneity. I estimate the Extended Cox model with a flexible baseline, and I assume that the unobserved heterogeneity follows the Gaussian distribution. Estimates indicate that my main results are robust to the presence of individual unobserved heterogeneity.

A possible explanation for the difference in unemployment duration between the genders is in female comparative advantage in childbearing and childrearing (Becker, 1993). The presence of young children could either increase female reservation wages or reduce search intensities. Consequently, the duration of female unemployment is expected to vary with age, i.e. female duration of unemployment rises during childbearing and childrearing years, and falls afterward. Alternative explanations for the difference in unemployment duration between the genders are discrimination and productivity differences (Black, 1995; Bowlus, 1997; Bowlus and Eckstein, 2002). These explanations imply that the gender differential in unemployment duration persists across age groups, and that it does not disappear after controlling for children. My estimates, on the other hand, provide evidence suggesting that the differential in unemployment duration between genders is driven by the division of labor in the household.

Finding a differential effect of young children in the duration of unemployment between East and West German females could be due to differential investment in early childhood human capital. The Becker et al. (2015) model of intergenerational transmission of human capital predicts that wealthier households invest more in their children's human capital. If investment in children's human capital is associated with the female length of unemployment during early childhood of their children, then females from wealthier households have either higher reservation wages or lower search intensities, and consequently have longer unemployment spells.

East Germans are considerably worse off economically. For example, the East German average annual wage in 1993 was €10,740, compared to €15,892 in West Germany; in 2011, the average East German wage was €14,536 and the average West German wage was €17,113.[4] The economic differences between East and West Germans are even wider if one compares the net worth of average East Germans (€76,000) and West Germans (€153,000) (Connolly, 2015).

I contribute to the literature in three ways. First, I show that the gender differences in unemployment durations vary with age in both East and West Germany and that these differences are associated with the presence of children and a partner. This distinguishes my paper from Kunze and Troske (2012), who only show that the gender differences in unemployment duration vary across age cohorts in West Germany. This also contributes to the literature studying differences in labor market outcomes between genders. Second, in the context of unemployment duration literature, I show strong evidence of fertility decisions being another source of variation in unemployment durations, in addition to unemployment insurance. Third, my findings that a partner and young children have differential effects on East and West German durations of unemployment contribute to the literature on differences in labor outcomes between Germans. Finally, I provide concluding remarks in Section 1.6.

[4] These wages are calculated using the wages for full-time employed workers from entire GSOEP sample, and are in real 2005 €.

1.2 Unemployment Duration in Germany

1.2.1 Data

I use data on unemployment spells from the German Socio–Economic Panel (GSOEP) which collects data on households and individuals in Germany annually. In 1984, the survey began by collecting data on 4,500 households in West Germany, and in 1990, the survey expanded its scope to East Germany by including an additional 2,179 households. As of the 2011 update, the survey included 12,290 households and 21,069 individuals.[5] The GSOEP contains a large array of socio-economic variables including information on children's age, partners' wages, and non-labor income, which are not available in administrative data. Additionally, it follows individuals through their employment biographies, and can distinguish between East and West Germans by their residence during 1989, just before reunification. In addition, I collect data on job vacancies and registered unemployed individuals at the state-yearly level from the Federal Statistics Office (Statistisches Bundesamt).

I use data for the period 1991–2011 to construct unemployment spells. The GSOEP survey asks two types of labor questions: contemporaneous and retrospective. In a monthly calendar section individuals retrospectively declare their labor status for each month in the previous year. I construct unemployment spells using this part of the survey. In the contemporaneous part of the survey, individuals provide information on the exact date when they lost their last job during the previous year or during the survey year. I use this information to cross-reference beginnings of the constructed spells. After cross-referencing the dates, I obtain a sample of 9,321 spells for 5,817 individuals, of which 57 percent are female spells. There are 4,509 East German spells, of which 52 percent are female spells, and 4,812 West German spells, of which 61 percent are female spells.[6]

According to the International Labor Organization's 1982 definition of unemployment, a person is considered to be unemployed if he or she is without employment while currently available and seeking work. Unfortunately, the GSOEP does not collect information on a person's availability for work or whether the person is actively looking for a job on a monthly basis. To circumvent this problem, I examine the spells of individuals in their prime working age (18–55) who entered unemployment either from employment (95 percent of spells) or after completing an apprenticeship

[5]For further information about the GSOEP, see Rahmann and Schupp (2013).
[6]My sample contains only individuals who are born in Germany.

education (5 percent of spells), since these individuals are most likely to be available and looking for work.

In Table 1.1, I provide an overview of the spells by location and gender (West German males, West German females, East German males, East German females). I use the Kaplan and Meier (1958) estimator to estimate median durations, i.e. the time until which at least half of the spells end in an exit from unemployment. Male median unemployment lengths are approximately two times shorter than female median durations, regardless of location. There are differences in the types of exits from unemployment between genders and between East and Wester Germans. The East German male share of unemployment spells ending in full-time employment is 95 percent, exceeding the West German male, East German female, and West German female shares by 6, 28, and 57 percentage points (respectively). Similarly, the East German male share of unemployment spells ending in part-time employment is approximately 5 percent, which is lower than the West German male, East German female, and West German female shares by 5, 27, and 56 percentage points (respectively).

Additionally, an individual in my sample can have more than one spell of unemployment. I find that 36 percent and 33 percent of West German male and female individuals (respectively) have more than one spell. In contrast, 49 percent and 45 percent of East German male and female individuals (respectively) have more than one spell. In summary, Table 1.1 indicates the drastic differences in length of unemployment and types of exits from unemployment between genders. Men enjoy shorter spells of unemployment and are more likely to return to full-time employment. East Germans are more likely to have multiple spells of unemployment with shorter durations, and East German females are considerably more likely to return to full-time employment than West German females.

1.2.2 Demographic Characteristics

In Table 1.2, columns 3–4 and 7–8, I present the composition of unemployment spells by demographic characteristics for West and East Germans. I find that variation in age at which individuals enter unemployment is similar between West German males and females, with approximately 60 percent of the spells originating before age 36 for both genders. However, there are differences between West German males and females in the distribution of spells across the three youngest age groups. There are 31.4 percent, 15.5 percent, and 14.8 percent of West German male spells in age groups 18–25, 26–30, and 31–35 (respectively), while 20.4 percent, 19.6 percent, and 20.1 percent of West German female spells are in the same age groups (respectively). In contrast, 49 percent and 53 percent of East German male and female spells (respectively) begin before age 36. I also note that the distribution of spells within each age cohort is much smaller between East German males and females, especially for those younger than 36. The comparison of the sample within gender across age groups for East and West Germans indicates that there are more spells starting at later ages in East Germany than in West Germany for both males and females.

During their unemployment spells, West German males are predominately childless or have a child older than seven, accounting for 74 percent of their spells. In contrast, West German females more often have a child during their unemployment spells. There is a child younger than eight in 55 percent of West German female spells, while there is a child younger than five in 44 percent of spells. East German males are also predominately childless or have a child older than seven during unemployment spells, accounting for 72 percent of their spells. East German females have a child younger than eight in 44 percent of their spells, while in 33 percent of their spells is there a child younger than five. A comparison of children indicators within genders and across locations shows that in 72 percent and 74 percent of East and West (respectively) German male spells, the males are childless or have a child older than seven. A similar comparison between females indicates that 33 percent and 44 percent of East and West (respectively) German female spells include a child younger than five, while there is a child younger than eight in 44 percent and 55 percent (respectively) of spells.

During unemployment spells, East and West German males are more likely to be single than East and West German females. Furthermore, an employed spouse is present in 34 percent and 67 percent of West German male and female (respectively) spells; in 21 percent and 7 percent of West German male and female (respectively) spells, the partner is also unemployed. In contrast, an employed spouse is present in 43 percent and 61 percent of East German male and female (respectively) spells; in 23 percent and 13 percent of East German male and female (respectively) spells, the partner is also unemployed. In terms of education, West German males have equal or lower educational levels than an apprenticeship in 56 percent of spells, while West German females have 35 percent of spells in the same categories. In contrast to the West, East German males have equal or lower educational levels than an apprenticeship in 29 percent of spells, while East German females have 19 percent of spells in the same categories.

In summary, a typical West German male spell in my sample would originate during ages 18–25 while the individual is single and childless. In contrast, a typical West German female spell would originate during ages 18–35 while the individual is in a relationship with an employed partner and has a child younger than eight. Similarly, a typical East German male spell would originate during ages 18–25 while the individual is childless but in a relationship with an employed partner. A typical East German female spell would originate during ages 18–25 while the individual has a child younger than eight and is in a relationship with an employed spouse.

1.2.3 Differences in Duration of Unemployment

In Table 1.2, I further examine differences in unemployment durations between genders and between East and West Germans by demographics. Columns 1–2 and 5–6 summarize median durations of unemployment for genders by location of origin. I present median durations as summary statistics, while I use the Peto and Peto (1972) and Prentice (1978) log-rank test (henceforward the Peto-Peto-Prentice test) to examine statistical differences in survivor functions between genders and between East and West Germans.[7] Additionally, I estimate survivor functions, i.e. the probability that an individual stays in unemployment for another month, using the Kaplan and Meier (1958) estimator.

First, I compare the median unemployment durations by seven age groups. In columns 1 and 2, the median durations show that West German males have shorter spells than West German females at any age. I also note two patterns for West Germans: first, median durations increase after age 50 for both genders; and second, only female median durations increase during childbearing and childrearing ages (26–35). In columns 5 and 6, I compare East German males and females and observe similar patterns as between West Germans, with the exception of East German males. I find that East German males only experience a one-month increase in duration after age 50 (compared to a 6-month increase for West German males). In figures 1.1 and 1.2, I plot the survival functions for both genders within each age cohort. I compare these survivor functions using the Peto-Peto-Prentice test and I find that females have significantly higher survivor functions than males, regardless of age or location.

I also compare the median durations and survivor functions within gender across locations.

[7] Unlike other log-rank tests of equality between survivor functions, the Peto-Peto-Prentice's log-rank test is appropriate even when hazards do not vary proportionally and the test is not affected by differences in censoring patterns across groups. All the comparisons of the survival functions are performed at a 5% significance level.

East and West German males have similar median durations of unemployment for all age cohorts except for the oldest group. In the oldest cohort, West German male durations exceed East German male durations by six months. However, I find that the differences between East and West males are only significant for age groups 31–35 and 51–55. In contrast to male durations, East German female median durations are lower than the West German female durations for all age groups except ages 18–25, where the median durations are same. Nevertheless, the Peto-Peto-Prentice test shows that differences between East and West German females are only statistically significant for age groups 26–30 and 31–35.

Second, I compare the median durations and survivor functions between genders in the presence of children. In column 1, West German females have 22 month-long and 18 month-long median durations in the presence of children ages 2–4 and 5–7 (respectively) which exceed West German male median durations by 15 months and 12 months (respectively); see column 2. Similarly, in column 5, East German females with children 2–4 and 5–7 have median durations of 15 months and 14 months (respectively) which exceed East German male durations by 10 and 9 months (respectively); see column 6. A comparison of survivor functions between genders in both East and West Germany shows that these differences are significant. Although the differences in median durations between genders in the presence of children older than seven or without children are small in both East and West Germany, the comparison of survivor functions shows that these differences are significant.

I compare East German median durations to West German durations within gender and across children indicators. I find no significant difference in survivor functions between East and West German male survivor functions for any of the variables indicating the presence of the children. In contrast to males, I find that the West German female median durations exceed the East German female median durations by 7 months and 4 months for children's age groups 2–4 and 5–7 (respectively), and the differences between corresponding survivor functions are significant. Additionally, I find no significant difference in survivor functions between East and West German males with children older than seven or without children. A similar comparison for females shows that East German females with children older than seven or without children have a longer median duration than comparable West German females, but these differences are only significant for females without children.

Third, I compare median durations and corresponding survival functions between genders in East and West Germany by relationship status. I find no difference in median durations or survivor functions between single West German males and females, while I find significant differences between married West German males and females, regardless of their partner's employment status. For example, West German female median durations are 17 months and 21 months in the presence of an employed and an unemployed partner (respectively), which exceeds West German male median durations by 12 months and 15 months (respectively). In contrast to West Germans, I find significant differences in survivor functions between genders in East Germany, regardless of relationship status. For example, East German married females have median durations of 13 months and 16 months in the presence of an employed and an unemployed partner (respectively), and these exceed comparable East German male durations by 9 months in the presence of both an employed and an unemployed partner.

The East-West comparison of median durations and survivor functions within genders shows that married East German females have a shorter median duration than married West German females, regardless of their partner's employment status. In contrast, single West German females have a shorter median duration than single East German females. Nevertheless, the Peto-Peto-Prentice test shows that the differences in survivor functions are significant between single East and West German females, as well as between married East and West German females with an employed partner. A similar comparison between East and West German males shows significant differences in duration only between married individuals with an employed partner. The median unemployment duration is longer for all groups if both partners are unemployed. This might be related to public transfers that are determined on the household level.

Fourth, I compare unemployment durations between genders across education levels. In columns 1 and 2, I find that West German males have shorter median durations than West German females for any education level, but the difference in median duration between genders falls with higher levels of education. For example, West German males with an apprenticeship degree have a median duration of eight months, which is seven months shorter than the comparable median duration for West German females. The difference in median durations between genders in West Germany is smallest for holders of tertiary degrees. The comparison of survivor functions between genders in West Germany shows that these differences in duration are significant for each educational level. The differences in median durations and survivor functions between genders in East Germany are similar to the West German differences. An exception is for East Germans with an incomplete degree, since there seems to be no significant difference in survivor functions between males and females.

Finally, I compare within-gender East-West differences across education levels. I find that East German males have shorter median durations than West German males, with the exception of those with incomplete degrees. The comparison of survivor functions between East and West German males indicates that these differences in duration are only significant for those with an incomplete degree. In contrast to males, I find that East German females have shorter median durations than West German females only for levels of education higher than an apprenticeship. The comparison of survivor functions indicates that the differences in duration are only significant between East and West Germans with secondary education.

1.2.4 Differences in Non-Labor Income

In Table 1.3, I present summary statistics for non-labor income by location of origin and gender, as well as their differences and corresponding standard errors. I include the following measures of non-labor income: income assets, public transfers, and partner's wages. All three variables are recorded yearly, and the first two variables are reported at a household level. Income assets include yearly income from interest, dividends, and rent, while public transfers include yearly income received from the government by all individuals in the household as well as benefits received at the household level.[8]

I find significant differences in non-labor income between genders in both West and East Germany. In both regions, females have on average higher non-labor income from income assets during their unemployment spells, whereas males on average receive more in public transfers. Additionally, during a female's unemployment spell, her partner on average earns more in wages than a male's partner. For example, West German females' partners will earn €19,171 more than West German males' partners. Similarly, East German females' partners earn €7,172 more than East German males' partners. These comparisons indicate that the female durations might be extended because females have significantly higher non-labor income, especially in terms of a spouse's wages. Thus, females might have higher reservation wages or lower search intensities than males.

Additionally, I make comparisons within genders between East and West Germans. I find

[8]Individual benefits include student grants, maternity benefits, unemployment benefits, unemployment assistance, subsistence allowances, and transition pay over all individuals in the household. Household benefits include housing allowances, child benefits, nursing care insurance, direct housing subsidies, subsistence assistance, support for special circumstances, social assistance for elderly, and benefits from the "unemployment benefit II" program.

that East Germans have significantly lower non-labor income than West Germans. For example, average asset income of an East German male is approximately €336–three times lower than the amount from asset income of a West German male, but the difference in the variation of income from assets is eight times larger in West. East German males also receive on average €9,453 in public transfers, and this exceeds West males' public transfers by €500. West German males' partners earn on average €6,367, which exceeds East German males' partners' wages by €1,206. A similar comparison of average non-labor income for East and West German females shows the same trend as for males, but with larger differences. On average, West German females receive €1,739 in non-labor income from assets, which exceeds East German females' averages by €1,268. East German females receive on average €8,401 in public transfers during their spells of unemployment, which exceeds West German females' public transfers by €2,452. Also, the average West German female partner's wage is €25,538 which is approximately two times larger than East female's partner's wage. Hence, the East Germans duration of unemployment might be shorter because their non-labor income is significantly lower than West Germans which means that either their reservation wages are lower or their search intensity is higher.

1.2.5 Differences in Labor Market Conditions

Besides the role of non-labor income in unemployment duration, local labor market conditions are important factors for unemployment length. I use a measure of labor market tightness and average real state wages at the yearly level to control for labor market conditions. Using the data from the Federal Statistics Office in Germany, I construct a measure of labor market tightness, a ratio of the number of job vacancies to the number of unemployed individuals. In Figure 1.3, I plot labor market tightness for East and West Germany. The figure indicates that labor market conditions are better in West Germany than in East Germany. Additionally, the figure indicates that local labor conditions in East and West Germany move together, but the gap between them does not shrink over time. Finally, the figure indicates that local labor market conditions began improving around 2005 in both East and West Germany.

In addition, I construct average real state wages as another control of local labor market conditions in Germany. I use real wages of full-time employed individuals from the entire GSOEP to construct average state wages. In Figure 1.4, I plot densities of wages for four periods: 1990–1995, 1996–2000, 2001–2005. and 2006–2010. The figure shows that average wages are significantly lower in East Germany during the first period, and that over time the East German wage converges towards the West German wage. I also note that the variance in average wages is increasing over time.

In summary, I find that females have a 14 month-long median duration of unemployment, which is two times longer than the median duration for males. East Germans have shorter durations than West Germans by a month. A decomposition of median durations across demographic characteristics shows that the differences between genders are largest during the childbearing and childrearing ages. In addition, gender differences are enlarged up to three times in the presence of young children. I also find that gender differences in unemployment durations between genders are smaller in East Germany than in West Germany.

One reason for the smaller differences in unemployment duration between genders in East Germany may be due to significantly lower non-labor income, which translates into lower reservation wages. However, the measures of labor market conditions show that East Germans are faced with a significantly tighter labor market than West Germans. Consequently, I would expect that East Germans are less likely to receive a job offer, which extends their unemployment durations. Thus, the difference in median durations and survivor functions between genders and between East and West Germans might be due to the lack of controls for factors such as non-labor income or local labor market conditions. In the next section, I use the Extended Cox model to see whether these differences in duration persist after controlling for other variables.

1.3 Empirical Strategy

1.3.1 The Job Search Model

The foundation of my empirical model is the Mortensen (1986) job search model. In the model, risk-neutral individuals look for a job while facing exogenous labor market conditions. During each period, an unemployed person receives a job offer with some probability. The job offer comes with a wage offer, which is identically and independently drawn from a known distribution function, $F(w)$. Given these assumptions, Mortensen (1986) shows that the optimal search strategy is characterized by a reservation wage, w^R, which depends on unemployment benefits, search costs, the discount rate, the probability of receiving a job offer (i.e. the contact rate), and the wage distribution function. Although Mortensen's (1986) model refers to the wage offer as a monetary reward, it could be generalized to capture some measure of the desirability of the job, such as the benefits, location, prestige, etc. Furthermore, the model's definition of unemployment benefits can be expanded to include the value of leisure or home production.[9] Comparative statics show that w^R increases with the value of benefits and decreases with an increasing discount rate or increasing search costs (Mortensen, 1986).

Mortensen's (1986) model also implies that in a two-state (employment and unemployment) labor market, the probability of leaving unemployment (or the unemployment hazard) is a function of the probability of receiving a job offer and the probability of accepting an offer that has been made.[10] Let the duration of unemployment be represented by a positive continuous random variable T. Then the unemployment hazard for an individual i at time t, $\lambda_i(t)$, is defined by

$$\lambda_i(t) := \lim_{h \to 0^+} \frac{Pr(t + h > T_i \geqslant t | T_i \geqslant t)}{h} \tag{1.1}$$

I parameterize person i's hazard $\lambda_i(t)$ as a function of a baseline hazard $\lambda_0(t)$ and a function of regressors x and parameters β, $\phi(x_t, \beta)$, where x can be either time-constant or time-varying.[11]

[9] See Rogerson et al. (2005) for a survey of search-theoretical models of labor markets.

[10] This reduced-form model implies that the total effects of the variables on the unemployment hazard are estimated rather than the distinct effects on the reservation wage and the probability of receiving a job offer. Thus, while this is not a direct test of search theory, it does have the advantage of not imposing further restrictive distributional assumptions required for structural analysis.

[11] The most common choice of $\phi(x_t, \beta)$ is the exponential form, which permits coefficients to be easily interpreted and ensures $\phi(x_t, \beta) > 0$.

Thus, the hazard function is

$$\lambda_i(t|x_t, \beta) = \lambda_0(t)\ \phi(x_{it}, \beta) \tag{1.2}$$

Duration models are sensitive to unobserved heterogeneity, which can bias the estimated hazard toward negative duration dependence if ignored, as shown by Heckman and Singer (1984b); this means that the probability of leaving unemployment will appear to be lower than it is. In this paper, I control for unobserved heterogeneity in two distinct ways. First, the Extended Cox model allows the baseline to be vastly different across unemployment durations thereby controlling for unobserved heterogeneity.[12] Second, I introduce an individual's unobserved heterogeneity, θ_i, as a random variable independent of regressors $x_i(t)$ and which enters the unemployment hazard multiplicatively, so that

$$\lambda_i(t|x_t, \beta, \theta_i) = \lambda_0(t_i)\ \phi(x_{it}, \beta)\ \theta_i \tag{1.3}$$

The model in (1.3) is identified if variation in observed survival times can be uniquely decomposed by contributions of the covariates, unobserved heterogeneity, and the duration dependence.[13] One must either make an assumption about the shape of the distribution $\mu(\theta)$ or use Heckman and Singer's (1984b) methodology to avoid the distributional assumption for $\mu(\theta)$. I use Meyer's (1990) methodology to estimate a discrete-time version of the Cox model with the baseline specified as a piecewise constant function, i.e. a dummy for each month of unemployment.[14] I also assume that unobserved heterogeneity, θ, follows the Gaussian distribution. Nicolettia and Rondinelli (2010) show that if the normality assumption for the unobserved heterogeneity distribution is incorrect, then neither the duration dependence nor the covariate coefficient estimation is biased.

1.3.2 Estimation

My hypothesis is that unemployment hazard varies differently between genders during a woman's childbearing and childrearing years. I use variations of the hazard function (1.2) to test

[12]This estimator is somewhat similar to Chamberlain's estimator for the logit model with panel data in that a conditioning operation is used to remove the heterogeneity (Green, 2012).

[13]See Melino and Sueyoshi (1990) for the conditions and proof of identifiability.

[14]Meyer (1990) shows that complementary log-log model (cloglog) is equivalent to the discrete-time version of the Cox model.

whether unemployment hazards vary with age differently between genders, and to test whether the differences in unemployment hazards between genders are associated with the presence of young children and a partner. Thus, I specify the hazard function as

$$\begin{aligned}\lambda_i(t|X\beta) = \lambda_g(t)\ \exp\Big[\ \vec{age}_i\ \vec{\alpha} + \vec{kids}_i(t)\ \vec{\gamma} + \vec{partner}_i(t)\ \vec{\delta} + female_i * \vec{age}_i\ \vec{\zeta} \\ + female_i * \vec{kids}_i(t)\ \vec{\eta} + female_i * \vec{partner}_i(t)\ \vec{\theta} + \vec{Z}_i(t)\ \vec{\mu}\ \Big]\end{aligned} \tag{1.4}$$

where $g = \{male, female\}$, and $\vec{age}_i$ is a vector of time-invariant indicators determining the age group at which an individual enters unemployment; $female$ is an indicator of an individual's gender; $kids_i(t)$ is a vector of time-varying dummy variables indicating the presence of children in a specific age cohort; $partner$ is a vector of indicators determining a partner's employment status. The set of regressors in $Z_i(t)$ includes personal controls (partner's employment status, education, public transfers, income assets, partner's real wage), and variables for local labor conditions (average real wage, labor market tightnesses), state and year fixed effects. In (1.4), I allow the baseline $\lambda_g(t)$ to differ across genders. In this specification, the coefficients of interest are $\vec{\zeta}$, $\vec{\eta}$, and $\vec{\theta}$, which indicate the differences in unemployment hazards within age groups between genders, the differences in unemployment hazards between genders in the presence of children, and the differences in unemployment hazards between genders in the presence of a partner (respectively).

I also test whether the differences in unemployment hazards between East and West Germans are associated with the presence of young children. In order to test the hypothesis, I specify the hazard function as

$$\begin{aligned}\lambda_i^g(t|X,\beta) = \lambda^g(t)\ \exp\Big[west_i\ \alpha^g + \vec{kids}_i(t)\vec{\gamma_g}\ + \vec{partner}_i(t)\ \vec{\delta^g} \\ + west * \vec{kids}_i(t)\ \vec{\zeta^g}\ + west * \vec{partner}_i(t)\ \vec{\eta^g} + \vec{Z}_i(t)\ \vec{\mu^g}\ \Big]\end{aligned} \tag{1.5}$$

where $g = \{male, female\}$, and $west$ is a dummy variable equal to one if an individual lived in West Germany during 1989 (just prior to reunification), and zero otherwise. In this specification, I allow the baseline hazard $\lambda^g(t)$ and the effect of covariates to differ across genders. The main coefficients of interest are α^g, $\vec{\zeta^g}$, and $\vec{\eta^g}$, which capture the conditional difference in unemployment hazards within genders between East and West Germans, the difference in unemployment hazards between East and West Germans due to children, and the difference in unemployment hazards between East and West Germans due to a partner (respectively).

I use the Cox (1972, 1975) partial likelihood method to estimate coefficients in (1.4) – (1.5). The hazard functions are semi-parametric, since the baselines remain unspecified. The Cox model is a series of comparisons of those subjects who leave unemployment to those subjects who are in unemployment; I informally refer to the latter set as the risk set. Thus, the partial likelihood largely depends on the order of events. When there are multiple exits from unemployment during the same instance, referred to as a tie, one must decide how to calculate the risk sets for these tied observations. Because the data on unemployment are collected on a monthly level, there are multiple exits from unemployment during the same month of unemployment, or multiple ties. I use Efron's (1977) method to correct for the presence of ties.[15] In both specifications, the effects of covariates are modeled parametrically as a multiplicative effect on the hazard, and the effects of covariates are independent of the baseline in both specifications.

In the estimation, I assume that an unemployment spell is a random draw. I obtain standard errors via bootstrap with 1000 replications, and I cluster standard errors to account for the possibility of an individual having more than one spell. Because the Extended Cox Model estimates relative risk between groups, I report hazard ratios (exponentiated coefficients), which can be interpreted as a change in the hazard at time t when the variables are increased by one unit. For interpretation, a hazard ratio (HR) greater than one indicates a shortening of the spells of unemployment as a covariate increases, and a HR less than one indicates an extension of the spells as a covariate increases.[16]

[15]Efron (1977) outperforms the Breslow and the Kalbfleisch-Prentice methods (see: Hertz-Picciotto and Rockhill, 1997); at same time, it is significantly less computationally demanding than exact marginal-likelihood methods.
[16]$HR = exp[b]$, HR>1 if b>0, and HR<1 if b<0.

1.4 Main Results

1.4.1 Gender Differences in Unemployment Duration

In Table 1.4, I report estimation results from the Extend Cox model specified in (1.4). First, I present regression results without controls for the presence of children or a partner in column 1. I then sequentially include dummies for children and a partner as well as their interactions with the indicator $female$ in columns 2 and 3 (respectively). I control for education, non-labor income, local labor market conditions, and year-state fixed effects in all three specifications. I find that higher levels of education are associated with significantly higher hazard rates across specifications. In addition, I find that partner's wage and public transfers significantly decrease the hazard rates, and that the local labor market tightness has a negative significant effect on the hazards.

Analysis of median durations and survivor functions shows marginal differences in durations of unemployment between men across age cohorts. In contrast, the estimates in column 1 of Table 1.4 indicate that men ages 26–40 have higher unemployment hazards than the omitted group (males 18–25). The results also show that the unemployment hazards start increasing after age 50, but the increase is significant only for the 51–55 age group. The interactions between the indicator $female$ and dummies for age categories capture differences in duration of unemployment between women and men within an age group. The results show that gender differences are largest in the 25–30 age group (37 percentage points) and smallest in the 41–45 age group (10 percentage points). The difference between genders falls by cohorts. However, the differences in unemployment hazards between genders are only significant in the 18–25, 26–30 and 31–35 age groups, where the females are 35, 37, and 32 (respectively) percentage points less likely to leave unemployment than comparable males. Additionally, I find the differences in unemployment hazards between genders for ages 18–35 are jointly significant while the differences between genders for age groups 45-50 and 51-55 are jointly insignificant. Thus the estimates of these interactions, given in column 1, provide evidence that the differences in unemployment durations between genders varies with age.

Becker (1993) suggests that women have a comparative advantage in childbearing and childrearing, which is why men specialize in market work and women in non-market work. Family formation typically happens during the late 20s and early 30s for men and women, while the prime childbearing and childrearing years are between ages 18 and 35.[17] Consequently, unemployed mar-

[17] In the entire GSOEP sample, there are 21, 64, 79, and 83 percent of individuals ages 18–25, 25–30, 31–35, and

ried men in this age range either have lower reservation wages or higher search intensities than single men, while unemployed married women either have higher reservation wages or lower search intensities than single women. Thus, the results in column 1 of Table 1.4 might reflect division of labor in the household since (i) men aged 26-40 have significantly higher hazards than men younger than 25, and (ii) women during their childbearing and childrearing years (ages 18–35) have significantly lower hazard rates than men, while there is no significant difference between genders at old age.

I am able to empirically test whether these differences in hazard rates between genders are due to the presence of young children and a partner. In column 2 of Table 1.4, I present estimates of the specification with the dummies for the presence of children and a partner. In this specification, I constrain the effects of children and a partner to be the same across genders. The results indicate that the presence of children ages 0–1 and 2–4 decreases the unemployment hazard 70 and 21 percentage points (respectively) while the presence of children ages 5–7 increases the unemployment hazard by 22 percentage points. Additionally, I find the presence of a partner increases the unemployment hazard by 53 and 71 percentage points if the partner is employed and unemployed, respectively. Finding that individuals are more likely to leave unemployment in presence of an employed partner is counterintuitive, which is why I test the equality of the effects of a partner's employment status on the unemployment hazard. I find no significant difference in the effects of an employed and an unemployed partner on the unemployment hazard.

In contrast to column 1, column 2 shows that differences in unemployment hazards between men ages 18–25 and men ages 31–40 disappear after controlling for children and a partner. The coefficients on interaction between $female$ and indicators for age cohorts indicate that women in any age group have significantly lower hazards than comparable men, and that these differences between genders are largest in age groups 25–30 and 31–35. Because I constrain the effects of children and a partner to be the same across genders, these findings do not come as a surprise, since the coefficients on age groups and their interactions with $female$ might be capturing differential effects of children and a partner on the unemployment hazard between genders.

In column 3 of Table 1.4, I present the results of the specification which allows the effects of children and a partner to differ across genders. The estimates on interactions between the indicator $female$ and age groups become insignificant for age groups 25–30 and 31–35. Thus, the difference in

36–40, respectively, living with a partner. At the same time, 17, 45, 52, and 34 percent of females ages 18–25, 25–30, 31–35, and 36–40, respectively, have children younger than eight.

unemployment duration between genders during childbearing and childrearing age disappears after including controls for children and a partner. Furthermore, the coefficients of interactions between $female$ and dummies for the presence of children show that the effects of children and a partner are different across genders. I find that young children significantly increase female unemployment durations, and that this effect falls with children's age. For example, women with children ages 0–1 have hazards 80 percentage points lower than comparable men, while women with children ages 2–4 and 5–7 have hazards 43 and 11 percentage points lower (respectively). In contrast to women, I find no significant difference in unemployment hazards between men with and without children ages 0–4. On the other hand, men in the presence of children 5–7 have a hazard rate 26 percentage points higher than men without children or with children older than seven.

The estimates in column 3 indicate that male unemployment durations are significantly shorter in the presence of a partner, regardless of the partner's employment status. In contrast, the interactions between the indicator $female$ and dummies for a partner's employment status show that female unemployment spells are significantly longer in the presence of a partner. For example, women with an employed partner and women with an unemployed partner have a chance of leaving unemployment 45 and 46 (respectively) percentage points lower than men with a partner, *ceteris paribus*.

In contrast to the results in columns 1 and 2, these results show that the hazard rates fall with age for both genders, but the effect is significant only for individuals older than 40. This finding is consistent with the theory of human capital investment. Older individuals are less likely to switch industry or occupation because they experience a greater loss in specific human capital than younger workers. Older workers are also less likely to invest in new human capital because the return is expected to be lower at an older age, since older workers are closer to the end of their working life. The interaction between $female$ and age groups 41–45 and 46–50 shows that females experience an additional decrease in unemployment hazards by 16 and 25 (respectively) percentage points, relative to comparable males. This finding indicates that females are less likely to leave unemployment at a later age, which might indicate that women are even less likely to switch occupations or industries at an older age than men.

In summary, the results in column 1 of Table 1.4 are similar to Kunze and Troske's (2012) findings for West Germany, and show that gender differences in durations of unemployment vary with age. In columns 2 and 3 of Table 1.4, I present new evidence indicating that differences in duration of unemployment between genders are associated with the presence of young children and a partner.[18] These findings are consistent with the theory of division of labor in the household and are similar to empirical findings by Kunze and Troske (2015) in the U.S.

1.4.2 Differences in Unemployment Durations between East and West Germans

In Table 1.5, I report estimation results from the Extended Cox regressions which test the differences in unemployment duration between East and West Germans. In columns 1–4, I present estimates for women, starting with the simplest model that controls only for personal characteristics (age, education level), local labor conditions, and state-year fixed effects. Then subsequently I add controls for non-labor income and indicators for children and a partner, as well as interactions between the dummy *west* and indicators for children and a partner. In columns 5–8, I present results for men in a similar order.

In columns 1 and 5, the coefficients on *west* indicate that West German females and males have significantly lower hazards, approximately by 25 and 19 percentage points (respectively), than their East German counterparts, *ceteris paribus.* These findings reinforce results from the analysis of median durations and survivor functions between East and West Germans. After I introduce controls for non-labor income (public transfers, partner's wage, income asset transfers) in columns 2 and 6, the differences between East and West Germans increase to 26 and 20 percentage points for females and males (respectively). These results show that differences in non-labor income are not primary drivers of differences in unemployment durations between East and West Germans.

In columns 3 and 7, I introduce controls for the presence of young children and a partner, but I constrain the effects of these variables to be the same across East and West Germans. These estimates show that West German women have a lower hazard by 21 percentage points than East German women. Similarly, I find that West German men have a lower unemployment hazard by 13 percentage points than East German men, but this difference is not significant. I also find that the presence of children ages 0–1 and 2–4 significantly reduces female unemployment hazards, while the presence of children ages 2–4 and 5–7 significantly increases male unemployment hazards. These estimates, in contrast to estimates in columns 1–2 and 5–7, indicate that the differences in duration

[18] I also estimate the same models separately by location of origin, and I find results similar to those in Table 1.4. These results are available on request.

of unemployment between Germans are related to the presence of children and a partner.

In columns 4 and 8, I show the final specifications that allow the effects of children and a partner to differ across East and West Germans, corresponding to equation (1.5). In columns 4 and 8, I find that West German females have hazards nine percentage points higher than East females, and that West German males have hazards eight percentage points lower than East German males, but these differences are insignificantly different. Furthermore, the coefficients on the indicators for children's age show that both East and West German females have lower unemployment hazards in the presence of children ages 0–4. Similarly, I find that the presence of children ages 2–7 increases both East and West German male hazard rates. Finally, the effect of a partner on unemployment hazard is positive regardless of partner's employment status, but the effects are more substantial for males. For example, male hazard in the presence of a partner is 97 percentage points greater than single male hazard, while a similar comparison between females shows a difference of 37 percentage points.

The coefficients on interactions of *west* with children indicators in column 4 indicate that there is a differential effect of young children on the hazards between East and West Germans. I find that West German females with children ages 0–1 have hazards 65 percentage points higher than comparable East German females. This finding might reflect higher opportunity costs of staying in unemployment for West Germans, but it should be taken with caution due to the small number of exits from unemployment in the presence of children ages 0–1 in East Germany.[19] I also find that West German females in the presence of children ages 2–4 have hazards 24 percentage points lower than respective East German females, while in the presence of children ages 5–7 there is neither an economic nor a statistical difference. In contrast to the differential response to the presence of children and a partner between East and West German females, I find no significant difference between East and West German males. The only exception is for the interaction of *west* with the indicator of children ages 0–1, where the West German hazard rate is found to be smaller than the East German hazard rate by 29 percentage points. However, this result should be also taken with caution, due to a small number of exits from unemployment in the presence of children ages 0–1.[20]

The coefficients on interactions of *west* with dummies for a partner indicate that West German women have hazards 31 and 40 percentage points lower in the presence of an employed

[19] There are 311 female spells that end in the presence of children age 0–1, of which 111 are East German.

[20] There are 155 male spells that ended in the presence of children ages 0–1, of which 69 are East German.

partner and an unemployed partner (respectively) than their East German counterparts. This result is counterintuitive, but reasonable in the context of institutional settings, since some welfare programs (such as unemployment insurance) are determined on the household income level, which starts after expiration of unemployment insurance . Additionally, I test the equality between these coefficients and find no significant difference. In contrast to females, I find no significant difference in the effect of a partner on unemployment hazards between East and West German males. Therefore, these findings reinforce our results from previous section that gender differences in unemployment durations are driven by children and a partner, and that they are present in both East and West Germany.

Finding the differential effects in unemployment duration in the presence of young children and a partner between East and West females can be associated with the differential in intergenerational transmission of human capital. In Becker et al. (2015), the level of intergenerational transfer of human capital between parents and children is determined by the credit constraint which depend on earnings, i.e. wealthier parents tend to invest more in the human capital of their children. If the female unemployment duration in the presence of young children is put in the context of intergenerational investment in human capital, then the differential in unemployment durations between East and West Germans might reflect higher human capital investment by West German parents. This mechanism is in line with the empirical findings of lower intergenerational educational mobility in East Germany (Riphahn and Trübswetter, 2013).

I find that the effects of other variables do not change across specifications in Table 1.5, and that their effects on the duration of unemployment are similar to findings the in literature. My results also show evidence of longer spells being associated with higher public transfers and a higher partner's wage. I also find that labor market tightness (the ratio of the number of job vacancies to the number of the unemployed) is significantly more important for women than men. The results indicate that the duration of unemployment varies with age and with education. Older individuals have longer unemployment spells, while better educated individuals have shorter unemployment spells.

In summary, conditional analysis confirms the conjectures from section 1.2: (i) unemployment durations are shorter in East Germany for both genders; (ii) the gap in unemployment hazards between East and West German women is associated with the presence of young children and a partner; (iii) there is no differential effect of children and a partner on the unemployment hazards of

East and West German males. These results are consistent with Becker et al. (2015) that there is more intergenerational transfer of human capital in West Germany.

1.5 Robustness

1.5.1 Unobserved heterogeneity

In Tables 1.6 and 1.7, I present estimates of the differences in unemployment hazards between genders and between East and West Germans from the Extended Cox model that includes the presence of individual unobserved heterogeneity. I assume that males and females share a common baseline in the specifications presented in Table 1.6, and it is assumed to be a piecewise constant function for each month, i.e. an indicator variable for each month of unemployment. There are no benefits in allowing the baseline to differ across genders, since I have modeled individual unobserved heterogeneity. In addition to the coefficients on covariates, I report the estimates of the standard deviation of the unobserved heterogeneity variance, σ_u, and the estimate of the portion of the total variance, ρ, that is due to the panel-level variance.[21] I test for the presence of the unobserved heterogeneity using a likelihood-ratio test where ρ is equal to zero under the null hypothesis. I perform the likelihood-ratio test for all three specifications in Table 1.6, and I find evidence of unobserved heterogeneity.[22] I perform similar tests for the specifications in Table 1.7, and find evidence of unobserved heterogeneity in all specifications except for the first specification in column 5.[23]

In table 1.6, the estimates in column 1 show a difference in unemployment durations between genders that varies with age. I find that the differences in unemployment hazard rates between genders are 66 and 62 percentage points for age groups 25–30 and 31–35 (respectively). However, the results also indicate that the gender differential in unemployment hazards is significant for older age groups except the oldest age group, 51–55. This finding might be due to omission of variables accounting for the presence of children and a partner. In contrast to column 1, the estimates in column 2 show that the differentials in unemployment hazards are smaller but still present across

[21] When ρ is zero, the panel-level variance component is not important, and the panel estimator is no different from a pooled estimator.

[22] The likelihood-ratio tests statistics for $H_0 : \rho = 0$ are $LR_{col1} = 81.91$, $LR_{col2} = 130.55$, $LR_{col3} = 159.29$, and the critical value with one degree of freedom is 3.842.

[23] The likelihood-ratio tests statistics for $H_0 \rho = 0$ are $LR_{col1} = 74.08$, $LR_{col2} = 124.10$, $LR_{col3} = 53.84$, $LR_{col4} = 46.85$, $LR_{col5} = 2.43$, $LR_{col6} = 100.60$, $LR_{col7} = 84.55$, $LR_{col8} = 81.91$, and the critical value with one degree of freedom is 3.842.

all age cohorts after including variables for the presence of children and a partner. In the final and most flexible specification, I allow the effects of children to vary across genders and find that the differential in unemployment duration does not vary with age. In addition, I find that the differences in unemployment hazards between genders are driven by the differential effect of young children and a partner. For example, the female hazard rates are 86 and 65 percentage point lowers than comparable male hazards in the presence of children age 0–1 and 2–4 (respectively). In the presence of a partner, regardless of employment status, the female hazard rate is 57 percentage points lower than the male hazard rate. Therefore, I find that my main results are robust to individual unobserved heterogeneity.

In Table 1.7, I present the estimates of the differences between East and West Germans in the presence of unobserved heterogeneity. These estimates indicate differences in unemployment duration between East and West Germans. In columns 1 and 5, I find that West German women and men have hazards 24 and 27 percentage points lower than respective East Germans, although the difference in hazards between East and West women is not significant. The difference in hazards between East and West Germans increases after I account for the presence of non-labor income (columns 2 and 6). In the last two specifications for both genders, I first control for presence of the children and a partner, and then I allow the effects of children and a partner to vary across East and West Germans. I find that the differences in unemployment hazards between East and West females are due to differential responses to young children and a partner, even after controlling for individual unobserved heterogeneity, and the estimates of these effects are very close to the main results. In contrast to the differences between East and West German females, I find West German males have a hazard rate that is 20 percentage points lower than the East German male hazard rate, but this difference is insignificant. I find no differential effect of children or a partner on the hazard rates between East and West German males, except for the interaction of *west* with a dummy variable for the presence of children ages 2–4. Hence, I my main findings regarding the differences in unemployment hazards between East and West German females is not affected by the presence of individual unobserved heterogeneity, while the results are somewhat weaker for males.

1.6 Conclusions

Using the German Socio-Economic Panel, I investigate the differences in unemployment durations between genders, and between East and West Germans for the period of 1991–2011. I use unemployment spells of prime age workers (18–55) to control for possible unobserved factors that might induce differential search between genders and between East and West Germans. A strength of my data are the availability of information on children, non-labor income, and ability to identify East and West Germans.

My estimates show that (i) there is a significant difference in unemployment duration between genders; (ii) upon return to work, females are more likely to take part-time jobs, especially West German females; (iii) the difference in unemployment durations between the genders varies with age, i.e. women have significantly longer durations of unemployment during the childbearing and childrearing ages; (iv) these differences in durations of unemployment between genders are associated with the presence of young children and a partner's employment status; (v) the difference in unemployment hazards between genders falls from 80 percentage points in the presence of children ages 0–1 to 40 percentage points in the presence of children ages to 2–4, and to an insignificant difference in the presence of children ages 5–7; (vi) both East German men and women have lower durations of unemployment; (vii) these differences between East and West Germans cease to exist after I account for the presence of children and a partner's employment status. My main results are robust to the presence of individual unobserved heterogeneity.

Findings in this paper provide new evidence on the differences in unemployment duration between genders associated with young children and a partner in Germany. This evidence is similar to Kunze and Troske's (2015) findings for the U.S., and suggest that this trend is not unique to the U.S. This paper also presents empirical evidence at odds with existing theoretical literature of employer based discrimination (e.g. Black, (1995); Bowlus and Eckstein, 2002), which predicts that women of all ages experience longer unemployment spells. My results indicate that the differences between East and West German female unemployment duration are due to the differential effect of children and a partner. A possible mechanism behind the differential effect of children on the duration of unemployment between East and West Germans might be the intergenerational transmission of human capital. Since average West Germans are significantly wealthier than average East Germans, they invest more in human capital for their children. However, further research is needed to examine if the intergenerational transmission of human capital is the true mechanism.

Table 1.1: Summary of unemployment spells

	West Germans		East Germans	
	Male	Female	Male	Female
Median duration (months)	7	14	6	13
Exits from unemployment (percent)				
Full-time	89.86	38.89	95.20	67.06
Part-time	10.14	61.11	4.80	32.94
Total spells exiting unemployment	69.85	60.89	74.60	64.60
Distribution of number of spells				
1	64.59	67.54	51.03	55.57
2	21.83	23.17	24.09	25.94
3	7.83	6.34	12.23	10.51
4	3.17	2.14	5.91	4.83
>4	2.58	0.81	6.74	3.15
Number of spells [Exits]	1864[1302]	2948[1795]	2150[1604]	2359[1524]
Number of individuals	1232	2063	1145	1377
Spell per individual	1.51	1.43	1.88	1.71
Total Observations (person-month)	17191	45297	20216	32408

Table 1.2: Unemployment durations by characteristics and sample composition

	West Germans				East Germans			
	Median Duration (months)		Sample Composition (percent)		Median Duration (months)		Sample Composition (percent)	
	Male (1)	Female (2)	Male (3)	Female (4)	Male (5)	Female (6)	Male (7)	Female (8)
Age								
18–25	7	11	31.4	20.4	6	11	22.0	21.4
26–30	5	18	15.5	19.6	5	14	13.6	17.4
31–35	8	19	14.8	20.1	5	13	13.6	15.0
36–40	6	12	12.7	15.1	5	10	12.0	14.6
41–45	6	13	10.3	10.6	6	11	13.5	11.6
46–50	8	13	7.9	7.4	7	15	13.2	10.8
51–55	14	19	7.4	6.8	8	15	12.1	9.3
Children								
Age 0–1	7	–	6.4	22.3	8	–	5.7	15.5
Age 2–4	7	22	10.3	22.1	5	15	10.6	17.1
Age 5–7	6	18	9.0	10.7	5	14	10.6	11.8
Age>7	7	9	24.2	23.6	6	10	17.7	21.6
Childless	6	8	50.1	21.2	6	9	55.4	34.0
Partner's employment status								
Single	8	9	45.8	25.3	8	11	33.4	24.8
Employed Partner	5	17	33.6	67.4	4	13	42.8	61.3
Unemployed Partner	6	21	20.7	7.3	7	16	23.8	13.8
Education								
Incomplete	8	20	18.5	11.1	10	23	5.2	3.9
Apprenticeship	8	15	38.0	24.2	7	19	24.2	15.0
Secondary	6	15	31.7	50.4	5	13	60.1	62.6
Tertiary	10	12	11.9	14.3	7	11	10.4	18.6

Note: Female median durations in the presence of children age 0–1 can not be estimated. The 25^{th} percentile durations are 18 and 14 months for East and West Germans, respectively.

Table 1.3: Mean and standard deviations for non-labor income by region and gender

	West Germans			East Germans		
	Males	Females	Differnce	Males	Females	Differnce
	Mean		(Males–Females)	Mean		(Males–Females)
	(std. deviation)		(std. error)	(std. deviation)		(std. error)
Income Assets	1040	1739	-699	336	471	-135
	(7273)	(11756)	(81)	(836)	(1817)	(12)
Public Transfers	8955	5949	3006	9453	8401	1052
	(6566)	(4857)	(58)	(6032)	(5430)	(54)
Partner's Wage	6367	25538	-19171	5161	12334	-7172
	(18465)	(25478)	(193)	(8773)	(13783)	(102)

Note: All variables of income are in real 2005 €. Averages are take across spells. I report standard deviations and standard errors in parentheses.

Table 1.4: Unemployment hazards between genders

Variables	(1)		(2)		(3)	
	Hazard ratio	Standard Error	Hazard ratio	Standard Error	Hazard ratio	Standard Error
Age cohorts						
26–30	1.249***	(0.0778)	1.162**	(0.0724)	0.978	(0.0624)
31–35	1.218***	(0.0884)	1.037	(0.0752)	0.909	(0.0671)
36–40	1.345***	(0.108)	1.126	(0.0915)	0.988	(0.0824)
41–45	1.100	(0.0864)	0.879*	(0.0688)	0.805***	(0.0660)
46–50	1.026	(0.0823)	0.796***	(0.0651)	0.734***	(0.0633)
51–55	0.758***	(0.0664)	0.563***	(0.0511)	0.521***	(0.0502)
Female*Age cohort						
18–25	0.653***	(0.0406)	0.723***	(0.0456)	1.086	(0.0729)
26–30	0.627***	(0.0527)	0.676***	(0.0557)	0.900	(0.0763)
31–35	0.674***	(0.0638)	0.716***	(0.0660)	0.893	(0.0855)
36–40	0.813**	(0.0821)	0.774**	(0.0797)	0.902	(0.0952)
41–45	0.904	(0.0954)	0.811**	(0.0845)	0.858	(0.0925)
46–50	0.822*	(0.0884)	0.727***	(0.0789)	0.774**	(0.0881)
51–55	0.891	(0.109)	0.794*	(0.100)	0.864	(0.116)
Children age						
0–1			0.297***	(0.0191)	0.930	(0.0942)
2–4			0.784***	(0.0339)	1.090	(0.0792)
5–7			1.221***	(0.0515)	1.273***	(0.0869)
Partner						
Employed			1.531***	(0.0669)	1.982***	(0.114)
Unemployed			1.710***	(0.0886)	1.987***	(0.142)
Female*Child age						
0–1					0.198***	(0.0257)
2–4					0.598***	(0.0528)
5–7					0.883	(0.0756)
Female*Partner						
Employed					0.552***	(0.0397)
Unemployed					0.532***	(0.0553)
Education						
Apprenticeship	1.187***	(0.0722)	1.197***	(0.0706)	1.211***	(0.0716)
Secondary	1.359***	(0.0784)	1.397***	(0.0782)	1.409***	(0.0782)
Tertiary	1.739***	(0.115)	1.895***	(0.123)	1.896***	(0.124)
Non-labor income	YES		YES		YES	
Local labor market condtions	YES		YES		YES	
State and year fixed effects	YES		YES		YES	
ln(L)	-51874		-47227		-47046	
Individiuals[Spells]	5817[9321]		5817[9,321]		5817[9,321]	
Exits from unemployment	6225		6225		6225	
Observations(month-person)	105791		105791		105791	

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Note: Results are obtain using the Extend Cox model. I report hazard ratios, exponentiated coefficients, with corresponding standard errors. Standard errors are obtained via bootstrap with 1000 replications, and I clustered standard errors at individual level to account for presence of multiple spells per individual. Non-labor income controls include public transfers, income asset transfers, and partner's wage, while local labor market conditions include average state wage and a measure of labor market tightnesses.

Table 1.5: Differences in unemployment hazard between East and West Germany

Variable	Females (1)	(2)	(3)	(4)	Males (5)	(6)	(7)	(8)
West	0.753***	0.732***	0.786***	1.099	0.801**	0.792**	0.866	0.928
	(0.0506)	(0.0526)	(0.0546)	(0.106)	(0.0786)	(0.0842)	(0.0898)	(0.108)
Children age								
0–1			0.177***	0.128***			0.970	0.922
			(0.0150)	(0.0194)			(0.103)	(0.136)
2–4			0.588***	0.694***			1.144*	1.337***
			(0.0299)	(0.0482)			(0.0849)	(0.130)
5–7			1.087	1.099			1.293***	1.277***
			(0.0605)	(0.0833)			(0.0931)	(0.110)
Partner								
Employed			1.136**	1.369***			1.977***	2.063***
			(0.0668)	(0.102)			(0.136)	(0.168)
Unemployed			1.070	1.349***			1.974***	1.929***
			(0.0861)	(0.138)			(0.140)	(0.184)
West* Children age								
0–1				1.653***				0.714**
				(0.287)				(0.100)
2–4				0.777***				1.095
				(0.0732)				(0.227)
5–7				0.995				1.041
				(0.107)				(0.152)
West* Partner								
Employed				0.683***				0.877
				(0.0676)				(0.0914)
Unemployed				0.599***				1.076
				(0.0942)				(0.147)
Personal controls	YES	YES	YES	YES	YES	YES	YES	YES
Non-labor income	NO	YES	YES	YES	NO	YES	YES	YES
State and year fixed effects	YES	YES	YES	YES	YES	YES	YES	YES
ln(L)	-26295	-25989	-25627	-25608	-21766	-21491	-21370	-21364
Individuals[Spells]		3440[5307]				2377[4014]		
Exits from unemployment		3319				2906		
Observations(month-person)		72398				33393		

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Note: Results are obtain using the Extend Cox model. I report hazard ratios, exponentiated coefficients, with corresponding standard errors. Standard errors are obtained via bootstrap with 1000 replications, and I clustered standard errors at individual level to account for presence of multiple spells per individual. Personal controls include education and age variables. Non-labor income controls include public transfers, income asset transfers, and partner's wage, while local labor market conditions include average state wage and a measure of labor market tightnesses.

Table 1.6: Unemployment hazards between genders with unobserved heterogeneity

Variables	(1)		(2)		(3)	
	Hazard ratio	Standard Error	Hazard ratio	Standard Error	Hazard ratio	Standard Error
Age cohorts						
26-30	2.114***	(0.345)	1.674***	(0.191)	1.288**	(0.147)
31–35	2.536***	(0.543)	1.782***	(0.269)	1.437**	(0.219)
36–40	3.521***	(1.008)	2.376***	(0.480)	1.952***	(0.399)
41–45	2.688***	(0.946)	1.846**	(0.472)	1.647*	(0.430)
46–50	2.210*	(0.983)	1.655	(0.544)	1.494	(0.505)
51–55	1.305	(0.729)	1.100	(0.465)	0.978	(0.425)
Female*Age cohort						
18-25	0.496***	(0.0631)	0.625***	(0.0544)	1.082	(0.101)
26-30	0.339***	(0.0692)	0.526***	(0.0657)	0.834	(0.103)
31–35	0.379***	(0.0769)	0.576***	(0.0745)	0.845	(0.112)
36–40	0.570***	(0.105)	0.663***	(0.0881)	0.866	(0.119)
41–45	0.711*	(0.137)	0.744**	(0.106)	0.815	(0.123)
46–50	0.644**	(0.132)	0.666***	(0.1000)	0.714**	(0.114)
51–55	0.699	(0.154)	0.724*	(0.120)	0.802	(0.142)
Children age						
0–1			0.266***	(0.0204)	1.022	(0.132)
2–4			0.612***	(0.0398)	1.179*	(0.116)
5–7			1.278***	(0.0725)	1.307***	(0.124)
Partner						
Employed			1.771***	(0.108)	2.682***	(0.234)
Unemployed			2.095***	(0.152)	2.594***	(0.265)
Female*Child age						
0–1					0.142***	(0.0235)
2–4					0.352***	(0.0472)
5–7					0.897	(0.106)
Female*Partner						
Employed					0.427***	(0.0456)
Unemployed					0.434***	(0.0614)
Education						
Apprenticeship	1.438***	(0.164)	1.325***	(0.106)	1.343***	(0.110)
Secondary	1.880***	(0.233)	1.651***	(0.129)	1.673***	(0.133)
Tertiary	2.994***	(0.491)	2.556***	(0.244)	2.592***	(0.249)
Non-labor income	YES		YES		YES	
Local labor market condtions	YES		YES		YES	
State and year fixed effects	YES		YES		YES	
σ_u	1.794	(0.294)	1.057	(0.099)	1.099	(0.089)
ρ	0.662	(0.073)	0.404	(0.045)	0.424	(0.039)
ln(L)	-21880		-21633		-21441	
Individuals[Spells]	5817[9321]		5817[9,321]		5817[9,321]	
Exits from unemployment	6225		6225		6225	
Observations(month-person)	105791		105791		105791	

*** p<0.01, ** p<0.05, * p<0.1

Note: Exponentiated coefficients are reported with corresponding standard errors. Standard errors are clustered at individual level to account for presence of multiple spells per individual. I estimate a random coefficient clog-log model, a discrete counter version of the Extended Cox model, in which unobserved heterogeneity is assumed to follow Gaussian distribution. I provide estimates of the standard deviation of unobserved heterogeneity variance, σ_u and the portion of the total variance that is due to the panel-level variance, ρ. Also, note that in these specifications I do not allow the baseline to differ across genders.

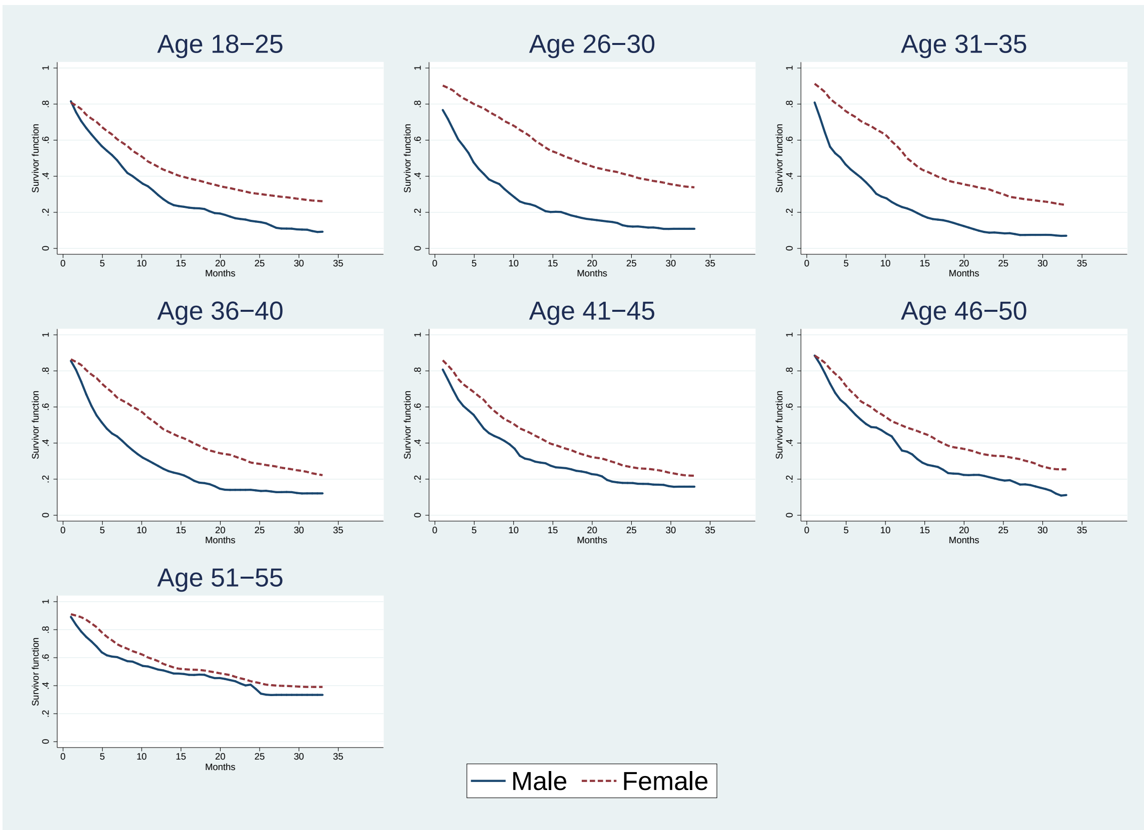
Table 1.7: Unemployment hazards between East and West Germans with unobserved heterogeneity

Variable	Females				Males			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
West	0.760	0.575***	0.757***	1.050	0.727**	0.636**	0.765*	0.812
	(0.326)	(0.0880)	(0.0661)	(0.124)	(0.108)	(0.118)	(0.116)	(0.136)
Children age								
0–1			0.143***	0.103***			1.090	1.086
			(0.0140)	(0.0167)			(0.151)	(0.205)
2–4			0.404***	0.491***			1.264**	1.568***
			(0.0383)	(0.0526)			(0.137)	(0.224)
5–7			1.122*	1.089			1.353***	1.333**
			(0.0771)	(0.101)			(0.138)	(0.170)
Partner								
Employed			1.161**	1.409***			2.885***	2.940***
			(0.0854)	(0.130)			(0.344)	(0.399)
Unemployed			1.090	1.383***			2.754***	2.599***
			(0.107)	(0.166)			(0.329)	(0.383)
West* Children age								
0–1				1.697***				0.990
				(0.326)				(0.268)
2–4				0.768**				0.619**
				(0.0864)				(0.125)
5–7				1.072				1.046
				(0.141)				(0.214)
West* Partner								
Employed				0.680***				0.910
				(0.0816)				(0.139)
Unemployed				0.592***				1.152
				(0.113)				(0.228)
Personal controls	YES	YES	YES	YES	YES	YES	YES	YES
Non-labor income	NO	YES	YES	YES	NO	YES	YES	YES
State and year fixed effects	YES	YES	YES	YES	YES	YES	YES	YES
σ_u	0.244	1.836	0.953	0.907	0.722	1.656	1.249	1.225
	(5.648)	.529	(0.115)	(0.115)	(0.466)	(0.378)	(0.169)	(0.163)
ρ	0.0349	0.672	0.356	0.333	0.241	0.625	0.487	0.477
	(1.560)	(0.127)	(.055)	(0.056)	(0.236)	(0.107)	(0.068)	(0.067)
ln(L)	-13077	-12756	-12425	-12408	-9399	-9077	-8962	-8958
Individuals[Spells]	3440[5307]				2377[4014]			
Exits from unemployment	3319				2906			
Observations(month-person)	72398				33393			

*** p<0.01, ** p<0.05, * p<0.1

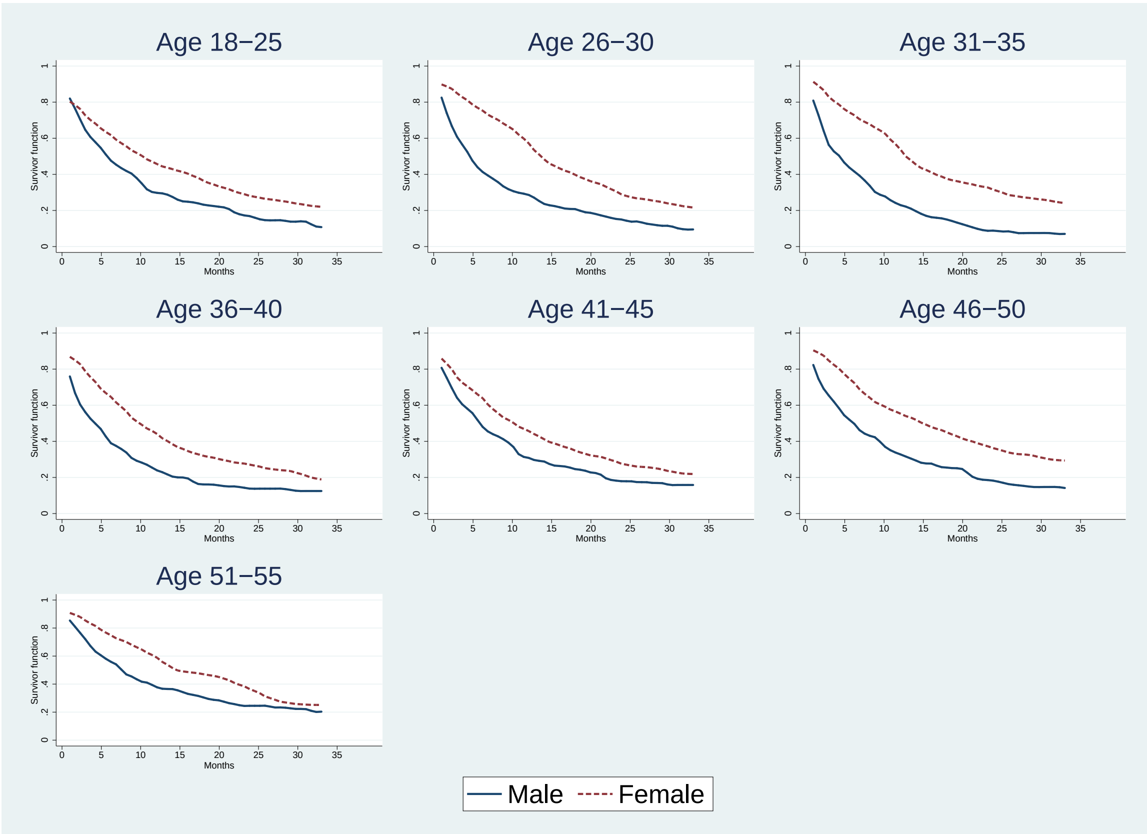
Note: Exponentiated coefficients are reported with corresponding standard errors. Standard errors are clustered at individual level to account for presence of multiple spells per individual. I estimate a random coefficient clog-log model, a discrete counter version of the Extended Cox model, in which unobserved heterogeneity is assumed to follow Gaussian distribution. σ_u is the standard deviation of the heterogeneity variance, and ρ is the portion of the total variance that is due to the panel-level variance.

Figure 1.1: Unemployment survivor functions by genders and age cohort, West Germans



Source: Author's calculations.
Note: The survivor functions are calculated using the Kaplan–Meier estimator. For all age groups, the Peto–Peto–Prentice rank test of equality between survivor functions yields p=0.001

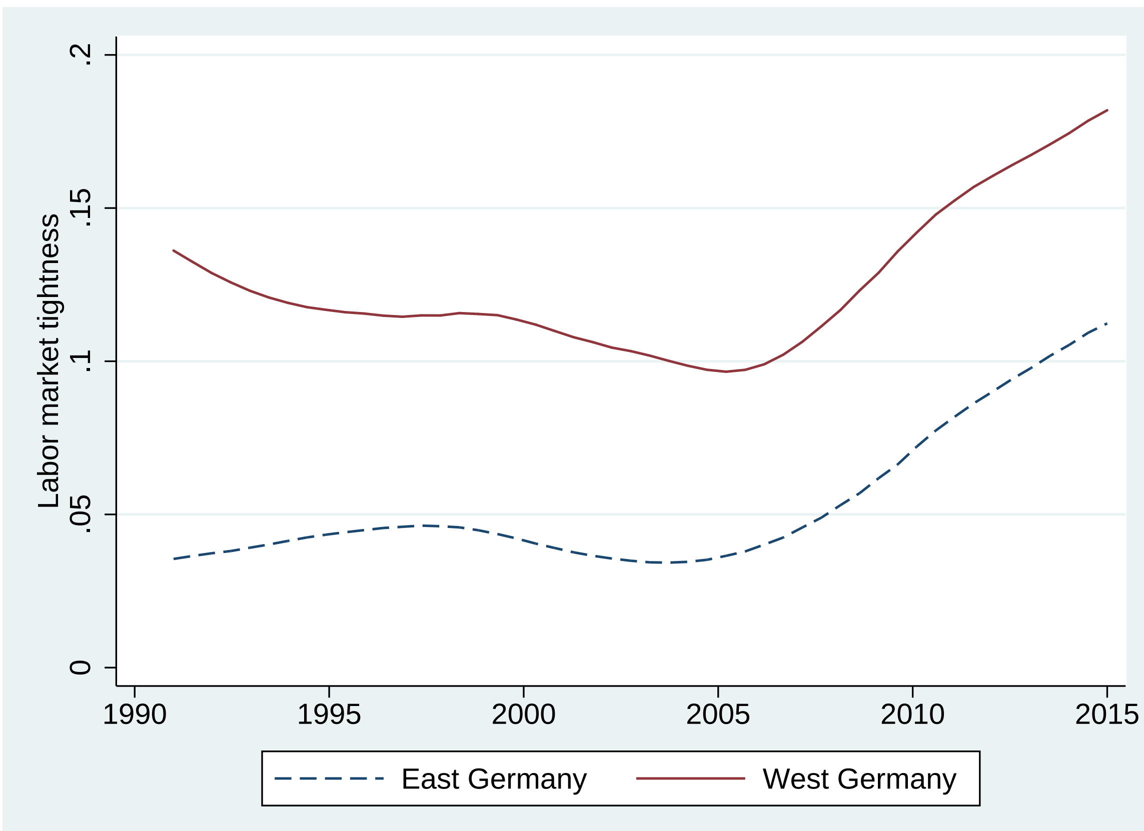
Figure 1.2: Unemployment survivor functions by genders and age cohort, East Germans



Source: Author's calculations.

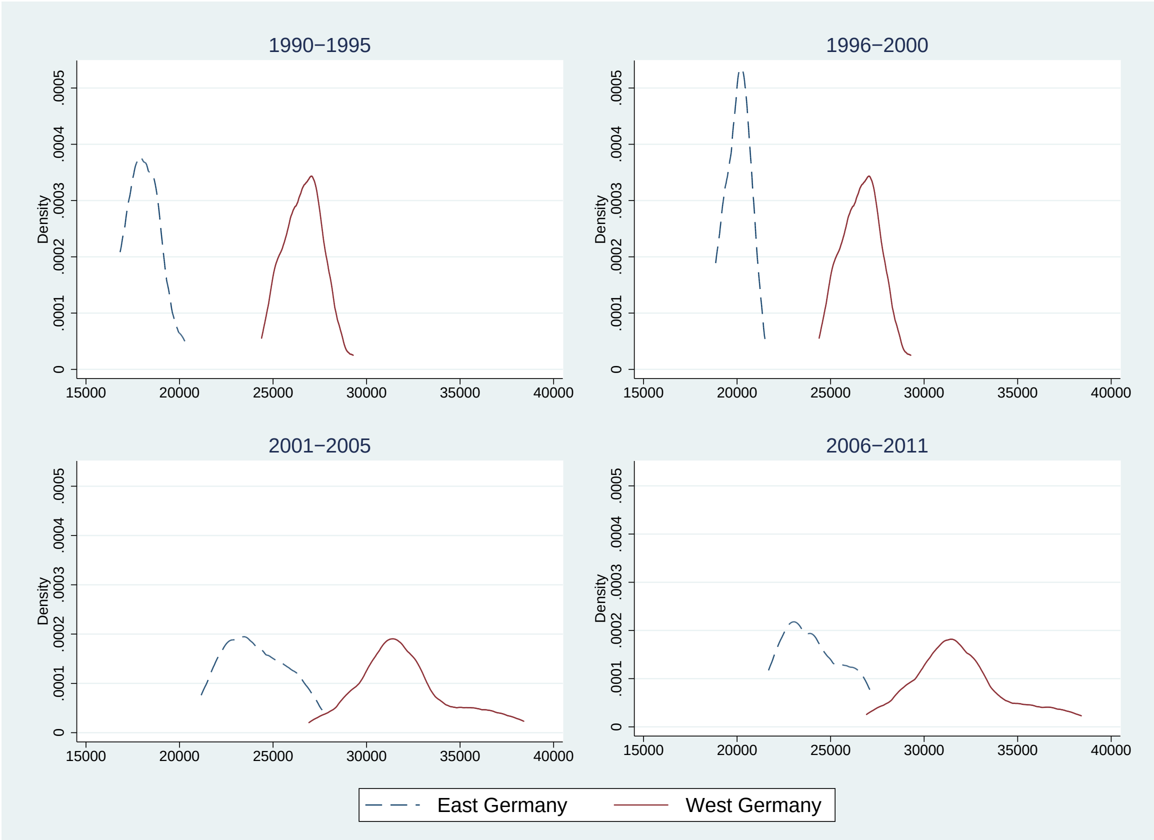
Note: The survivor functions are calculated using the Kaplan–Meier estimator. For all age groups, the Peto–Peto–Prentice rank test of equality between survivor functions yields p=0.001

Figure 1.3: Labor market tightness in East and West Germany, 1990-2015



Source: Author's calculations using data from the Federal Statistics Office in Germany (Statistisches Bundesamt).
Note: Labor market tightness is ratio of vacancies to number of registered unemployed individuals.

Figure 1.4: Densities of real wages in East and West Germany



Source: Author's calculations using the entire German Socio-Economic Panel sample of full-time employed individuals.

Chapter 2

The Hidden Cost of Homeownership in Germany: Unemployment Length

2.1 Introduction

Homeownership is associated with a number of positive outcomes such as urban sprawl or enacting zoning laws. There is evidence that homeowners have better physical and mental health, are less likely to be victims of crime, and are more politically and socially involved in local affairs than renters (Dietz and Haurin, 2003). However, homeownership negatively affects labor market mobility, for example, literature finds that homeownership is associated with higher rates of unemployment and lower labor mobility in the OECD countries (Oswald, 1996; Nickell and Layard, 1999; Green and Hendershott, 2001a; Blanchflower and Oswald, 2013)[1]. Additionally, understanding the interaction between homeownership and the labor market is important for policy evaluation such as unemployment benefits and homeownership subsidies.

In this paper, I study the effect of homeownership on the unemployment duration in Germany between 1990 and 2011. I use micro-data to study the effects of homeownership on unem-

[1] A 10 percentage point increase in homeownership rate is associated with a rise in unemployment rate between 1.3 and 2 percentage points (Oswald, 1996; Nickell and Layard, 1999).

ployment duration, and I address self-selection into homeownership by using the full information maximum likelihood and exclusion restriction[2]. Homeownership can increase unemployment duration, namely by decreasing geographical mobility. In the context of the job search model, this mechanism would increase reservation wages for jobs requiring moves, which would reduce the rate of exits from unemployment (unemployment hazard rate) to the non-local labor market. Even if homeowners increase search efforts in local labor markets to compensate for higher reservation wages to non-local jobs, Munch et al. (2006) shows that homeowners have lower chances of leaving unemployment than renters under fairly general assumptions. In addition, Munch et al. (2006) shows that the difference in the rate of exits from unemployment between homeowners and renters is an empirical question that depends on factors such as the size of a country, distribution of industries across a country, as well as cultural and linguistic differences, etc. I find that homeownership increases unemployment duration after I control for self-selection into homeownership.

Previous studies find mixed evidence on the effects of homeownership on unemployment as well as on labor mobility, and this evidence varies with the type of data used (micro vs macro), with a country, and the econometric approach[3]. In Germany, previous literature using regional-level data finds that homeownership is associated with higher rates of unemployment in Germany (Lerbs, 2011; Wolf and Caruana-Galizia, 2015) while only Wolf and Caruana-Galizia (2015) control for self-selection into homeownership using the instrumental variable approach.

In Section 2.2, I provide some information about my sample and examine the differences in unemployment duration between homeowners and renters using median duration and non-parametric survivor functions. My data comes from the German-Socio Economic Panel (GSOEP) for 1990–2011. The advantage of the GSOEP over official German administrative data is the availability of data on homeownership status, partner status, and presence of young children, which are important determinants of both labor mobility and homeownership. My sample consists of unemployed men in their prime working years, ages 25–55, from the moment they transition to unemployment until the time at which they either exit unemployment to work or drop out of the survey. I use prime-age workers because they are more similar with respect to decisions affecting their job search than workers reentering the labor market after inactivity, workers coming out of education, or workers

[2]This approach is similar to Heckman's sample selection model but it is more appropriate for duration analysis and more flexible.

[3]Evidence of the positive effects or no effects of homeownership includes: van Leuvensteijn and Koning (2004), Munch et al. (2006), Taşkın and Yaman (2016), while these studies find negative effects: Oswald (1996), Munch et al. (2008), Blanchflower and Oswald (2013).

that are close to retirement. Finally, I choose to use only men to avoid potential issues with women staying in unemployment due to childbearing and childrearing.

When I examine the median duration of unemployment and probability of remaining in unemployment, I find that homeowners have shorter spells of unemployment than renters, and this difference is especially large during longer spells. The median duration of unemployment for homeowners is five months compared to seven months for renters. Additionally, the difference in the third quartile of unemployment duration between homeowners and renters is six months, i.e. homeowners and renters spend 13 and 19 months in unemployment, respectively. I also compare differences in probability of remaining in unemployment between homeowners and renters and find that homeowners have significantly lower probabilities of remaining in unemployment than renters. Finally, a comparison of homeowners and renters across observables shows that the two groups are similar with an average homeowner being slightly older and better educated, and being more likely to have a partner who on average earns more than an average renter's partner.

In Section 2.3, I present the empirical strategy which controls for the differences in the observables and corrects for self-selection into homeownership. Because of high transaction costs, unfavorable financing terms and lack of tax breaks or subsidies for homeowners in Germany, most households in Germany purchase homes only if they plan to settle down in a certain region for an extended period of time[4]. Thus, men choosing homeownership, for example, probably have stronger preferences for location compared to renters (e.g. proximity to family or stability). Then in case of unforeseen unemployment, homeowners might be less willing to leave unemployment for a job that requires a move because they are inherently less mobile and not for their choice of homeownership. Thus, this would produce an upward bias in the estimate. It is also possible for the bias to be downward. For example, it is possible that some men are more productive and that these men are more likely to be homeowners while at the same time they face better opportunities in labor market in case of unemployment.

I address self-selection into homeownership by simultaneously estimating the duration model and the homeownership model while allowing the models to be correlated by specifying a bivariate distribution for two unobserved variables; one affects the duration model and the other affects the

[4] In Germany, transaction costs: including taxes, realtor and notary fees are approximately 10 percent of the purchase price (Catte et al., 2004). Additionally, the German housing market is inflexible, mainly due to the unfavorable property loan market compared to neighboring countries. The market is characterized with the long term mortgage loans and fixed interest rates for a period of 5 to 10 years, as well as with high early repayment penalties in the case of terminating the mortgage earlier (Voigtländer, 2009).

homeownership model[5]. The bivariate distribution is estimated non-parametrically to minimize the biasing impact of the distributional assumption (Heckman and Singer, 1984a). The identification is achieved using the exclusion restriction, i.e. there is a variable or a set of variables that affect the probability of homeownership but not the exit rate out of unemployment. Similar to van Leuvensteijn and Koning (2004), I use state-level homeownership rates as the excluded variable since it is reasonable to expect that homeownership rates are correlated with the individual probability of owning a home, but not with the unemployment duration.

In Section 2.4, I summarize estimation results for the effect of homeownership on unemployment duration. I estimate the model with and without correction for self-selection into homeownership. I use the model without correction as a baseline and find that homeowners have a 13 percent higher rate of exists from unemployment than comparable renters, and this difference is statistically significant. This finding is similar to the evidence found using median duration and non-parametric estimate of the probability of remaining in unemployment. However, these findings need to be taken with caution because i) I do not control for unobserved heterogeneity, which in duration models can create negative duration dependence, and ii) I do not control for self-selection into homeownership.

After I correct for unobserved heterogeneity and self-selection into homeownership, I find that homeowners have a 56 percent lower exit rate out of unemployment than renters. This is statistically significant at 5 percent level. The bias-corrected 95 percent confidence interval shows that homeowners have between an eight and 62 percent lower exit rate out of unemployment than comparable renters[6]. Finding that homeownership reduces the rate of exits from unemployment suggests that homeownership comes with the costs of lower labor mobility as argued by Oswald (1996), and more recently, Blanchflower and Oswald (2013). Furthermore, my results are also in line with evidence from German regional level data (Lerbs, 2011; Wolf and Caruana-Galizia, 2015), and with micro-evidence from France presented by Burnet and Lesueur (2003). Evidence in this paper is at odds with evidence from Denmark and the U.S., where homeowners are more likely to leave unemployment sooner than renters (Munch et al., 2006; Taşkın and Yaman, 2016).

I contribute to the literature by providing a micro evidence of homeownership increasing unemployment duration in Germany. The use of micro-data instead of macro or meso-data allows me to correct for the spurious relationship between homeownership and unemployment present in

[5]This methodology was successfully in van Leuvensteijn and Koning (2004) and Munch et al. (2006).

[6]Standard asymptotic inference is not appropriate in this case due to left skewness in the empirical distribution of my estimate.

regional and macro data, and to identify the causal effect of homeownership on unemployment duration. Furthermore, the evidence on the effects of homeownership on unemployment duration in this paper is more robust than the evidence presented in Green and Hendershott (2001b), Coulson and Fisher (2002), and Burnet and Lesueur (2003) because I control for possible self-selection into homeownership, which proves to be important due to the large bias in the estimate. Finally, I provide concluding remarks in Section 2.5.

2.2 Unemployment Duration and Homeownership in Germany

2.2.1 Data and Sample

I use data on unemployment spells from the German Socio–Economic Panel (GSOEP) which annually collects data on households and individuals in Germany. In 1984, the survey began by collecting data on 4,500 households in West Germany, and in 1990, the survey expanded its scope to East Germany by including an additional 2,179 households. As of the 2011 update, the survey includes 12,290 households and 21,069 individuals.[7] The GSOEP contains a large array of socio-economic variables including information on homeownership, childrens' ages, partners' wages, and non-labor income, which are not available in administrative data.

I use data for the period 1991–2011 to construct unemployment spells. The GSOEP survey asks two types of labor questions: contemporaneous and retrospective. In a monthly calendar section individuals retrospectively declare their labor status for each month in the previous year. I construct unemployment spells using this part of the survey. In the contemporaneous part of the survey, individuals provide information on the exact date when they lost their last job during the previous year or during the survey year. I use this information to cross-reference beginnings of the constructed spells. After cross-referencing the dates, I obtain a sample of 9,321 spells for 5,817 individuals, of which 57 percent are female spells.

According to the International Labor Organization's 1982 definition of unemployment, a person is considered to be unemployed if he or she is without employment while currently available and seeking work. Unfortunately, the GSOEP does not collect information on a person's availability for work or whether the person is actively looking for a job on a monthly basis. To circumvent this problem, I study the spells of individuals in their prime working age (25–55) who entered unemployment either from employment (95 percent of spells) or after completing an apprenticeship education (5 percent of spells), since these individuals are most likely to be available and looking for work. In addition, I restrict the sample to men in order to avoid potential issues with women staying in unemployment due to childbearing and childrearing. This leaves 3,077 spells for 1,820 individuals.

[7] For further information about the GSOEP, see Rahmann and Schupp (2013).

2.2.2 Unemployment Duration

In Table 2.1, I summarize unemployment duration by the status of homeownership. Examining the survival time indicates that homeowners and renters have the same duration in the 25th percentile, and that homeowners have a shorter duration of unemployment in the upper percentiles of the survivor function, probability of staying in unemployment beyond month t. For example, at the median, homeowners are unemployed for two months less than renters, while in the 75th percentile of unemployment duration, homeowners spend six months less in unemployment than renters. In Figure 2.1, I plot separately the survivor functions for homeowners and renters estimated using the Kaplan-Meier estimator, and find that the renters' survivor function is higher than the homeowners' survivor function. I find that this difference between survivor functions is significantly different using the Peto-Peto-Prentice rank test of equality between survivor functions. Despite differences in duration of unemployment, I find a lack of difference in the types of exits from unemployment between homeowners and renters. Both groups have similar rates of exits from unemployment to both full- and part-time employment. Finally, I do not observe exits out of unemployment for 20 percent of homeowners' spells and 26 percent of renters' spells.

In Table 2.2, I present spell composition by characteristics including age, education, partner status, children, and non-labor income. I find that homeowners are approximately equally likely to enter unemployment spells across age groups, while renters are more likely to be young when entering spells of unemployment. In particular, I find that approximately 30 percent of homeowners' spells originate at age age 35 or younger, while 43 percent of renters' spells originate at age 35 or younger. I also find that renters are slightly less educated than homeowners. Homeowners are five percentage points more likely to have high school as a terminal degree than renters, and about four percentage points less likely to have only elementary education as their terminal degree.

Furthermore, I find that homeowners are five percentage points more likely to have a partner. Interestingly, homeowners are 11 percentage points more likely to have an employed partner than renters and seven percentage points less likely to have unemployed partners than renters. Homeowners have one or more children in 57 percent of unemployment spells, while renters have one or more children in 52 percent of unemployment spells. Additionally, homeowners are nine percentage points more likely to have children age 13–18 during their unemployment spells. These differences between homeowners and renters can be indicative of homeowners' preferences for stability or location. Finally, homeowners on average have access to € 2,133 and € 425 more than renters coming from partners' wages and public transfers, respectively.

In summary, homeowners have shorter lengths of unemployment than renters and the differences are especially large above median duration of unemployment. Additionally, homeowners have significantly lower survivor function than renters. Thus, this preliminary evidence suggests that homeowners are more likely to end unemployment spells sooner than renters, which contradicts the Oswald hypothesis. However, I also find that homeowners are older, more likely to have a partner and children, and have higher non-labor income.

2.2.3 Homeownership in Germany

When it comes to an international comparison in homeownership rates, Germany's homeownership rate of 43 percent is lower than in Spain (80 percent), Italy (80 percent), United Kingdom (70 percent), Austria (60 percent), France (55 percent), etc. (Voigtländer, 2009). Voigtländer points at five key factors that are responsible for the low homeownership rate: i.) the relative size and quality of the rental market, ii.) the less favorable tax treatment compared to neighboring countries, iii.) the unfavorable terms of financing (substantial down payments, long-term fixed-rate loans, and high transaction costs), iv.) the relative costs of owned and rented units, and v.) slow appreciation of the houses. Additionally, Boehm and Schlottmann (2014) argue that demographic composition in Germany is an important factor of low homeownership rate. In particular, they argue that marital status, a primary demographic determinant of housing choice, is a reason for the low homeownership rate in Germany since 57 percent of adults are single.

I summarize average homeownership rates by the state over period 1990 to 2011 in Table 2.3. In addition, I provide time series of homeownership by the state in Figure 2.2. I generate these statistics using all the data from the GSOEP for 1990–2011, and I use individual weights in the aggregation process to ensure that the rate is representative of Germany. Table 2.3 shows large variation in homeownership rates across states with the average lowest rate of 11 percent in Berlin and the average highest rate of 62 percent in Saarland. Although one might think that the large variation in homeownership rates is the product of GSOEP data being non-representative of Germany, in Table 2.4 I show that these average rates are similar to averages collected by the German Microcensus[8].

[8]Although German Microcensus is a more representative data set, I do not use German Microcensus to obtain homeownership rates because the data is only available for a limited number of years.

Finally, Figure 2.2 shows developments in homeownership rates across German states over time. The figure points at the variation in the time trends across states. Thus, the aggregated statistics suggest the presence of large variations in homeownership rates across German states, and in time-trends of homeownership across the states. The large regional variation in the homeownership rate is associated with regional differences in relative housing prices and capital requirements needed for the purchase of a housing unit, regional household composition, level of urbanization, employment levels, and neighborhood characteristics (Lerbs and Oberst, 2014).

2.3 Empirical Strategy

2.3.1 The Job Search Model

I use the Munch et al. (2006) job search model as a foundation for my empirical model. Munch et al. extend Mortensen's (1986) job search model by introducing two types of exits from unemployment: local and non-local exits. In Munch et al.'s model, risk-neutral individuals look for a job while facing exogenous labor market conditions. It also assumed that individuals must live and work in the same region. Additionally, it is assumed that renters do not have costs associated with moving, while homeowners face the cost of selling their homes. During each period, an unemployed person receives job offers in local and non-local labor markets with some probabilities. The job offers come with the wage offers, which are identically and independently drawn from a same known distribution function, $F(w)$.

Given these assumptions, Munch et al. (2006) show that an unemployed renter is indifferent between accepting a job in a local and non-local market, since reservation wage, w^R, is the same for accepting a job offer in the two markets. Moreover, Munch et al. show that an unemployed homeowner has a larger reservation wage for job offers on the non-local market, w_n^R, than the reservation wage for jobs in the local market, w_l^R, since the homeowner needs to be compensated for costs associated with moving. Additionally, a homeowner is willing to reduce w_l^R below w^R to avoid incurring moving costs. Thus, a homeowner has a lower reservation wage in the local labor market, while a homeowner has higher reservation wage than a renter in the non-local labor market, $w_l^R < w^R < w_n^R$.

The model also implies that the hazard rate out of unemployment is a product of the arrival rate of the job offers and the probability of accepting the job offer. Hence, a renter's hazard rate to a job in the local labor market, λ_l^R, and a renter's hazard rate to a job in the non-local labor market, λ_n^R are

$$\lambda_l^R = \alpha_l \left[1 - F(w^R)\right] \text{ and } \lambda_n^R = \alpha_n \left[1 - F(w^R)\right] \tag{2.1}$$

where α_l and α_n are job arrival rates in the local and non-local labor markets. Similarly, a homeowner's hazard rates are

$$\lambda_l^H = \alpha_l \left[1 - F(w_l^R)\right] \text{ and } \lambda_n^H = \alpha_n \left[1 - F(w_n^R)\right]. \tag{2.2}$$

Given $w_l^R < w^R < w_n^R$, it follows that homeowners have higher hazard rate of unemployment in the local labor market than renters, and that renters have higher hazard rates out of unemployment in the non-local labor market than homeowners. Finally, the overall hazard rates out of unemployment, $\lambda_l^H + \lambda_n^H$ and $\lambda_l^R + \lambda_n^R$, depend on the relative sizes of arrival rates, α_l and α_n, and the relative sizes of $F(w_l^R) - F(w^R)$ and $F(w_n^R) - F(w^R)$ [9].

2.3.2 Empirical Model

I employ a hazard model to estimate the effect of homeownership on the rate at which individuals leave unemployment. One limitation of my study is the inability to identify if individuals leave unemployment for jobs in or out of the local labor market due to the lack of information. Let the duration of unemployment be represented by a positive continuous random variable T. Then the unemployment hazard for an individual i at time t, $\lambda_i(t)$, is defined by

$$\lambda_i(t) := \lim_{h \searrow 0^+} \frac{Pr(t + h > T_i \geqslant t | T_i \geqslant t) + \nu}{h}. \tag{2.3}$$

I parameterize person i's hazard $\lambda_i(t)$ as a function of duration dependence $\lambda_0(t)$, the binary variable for homeownership status *homeowner*, a matrix of observable variables X, and unobserved time-invariant characteristic ν_u. In the estimation, I specify $\lambda_0(t)$ as a cubic polynomial rather than piecewise constant function to reduce the number of parameters from 56 to 33. Thus, the individual hazard rate out of unemployment is

$$\lambda_i(t|X, \mathit{homeowner}) = \exp(\lambda_0(t) + \delta\, \mathit{homeowner} + X\,\beta + \nu_u). \tag{2.4}$$

I use a discrete hazard model to accommodate the discrete feature of my data, i.e. the spell length is reported in months. Camron and Trivedi (2005) show that corresponding discrete hazard function λ^d analogous to (2.4) and discrete survivor function S^d for individual i during $[t_{k-1}, t_k)$ are

[9] In estimation, I only concern myself with the overall hazard rates due to lack of data on location of employment after leaving the unemployment spell.

$$\lambda_k^d(t_k|X, homeowner) = 1 - \exp\left[-\exp\left(\lambda_0(t) + \delta\; homeowner + X\beta + \nu_u\right)\right], \text{and}$$
$$S_k^d(t_k|X_i) = 1 - \lambda_k^d(t_k, X). \tag{2.5}$$

Let d be a binary variable equal to one if an individual's observed leaving unemployment for employment during the final month of observation, and zero otherwise, then the contribution of individual i's spell of unemployment lasting T months to the likelihood function is

$$L_i = (1 - S_T)^d\, S_T^{1-d} \prod_{j=1}^{T-1} S_j\ . \tag{2.6}$$

I refer to (2.6) as the unemployment model.

The concern with estimating the effect of the homeownership on the hazard rate this way is a possible selection bias into homeownership. For example, some households might be less mobile, because of their preferences for location due to proximity to family, size of the family, etc. Then, such households are more likely to choose to homeownership. In the case of unemployment, these households might be less willing to move for job opportunities outside of the local labor market. However, this is not due to their choice of homeownership but rather due to their preferences for location. To control for the possible presence of self-selection, I estimate simultaneously the hazard model and the homeownership model while allowing the two models to be correlated via bivariate distribution for two unobserved variables from each model (van Leuvensteijn and Koning, 2004; Munch et al., 2006; Munch et al., 2008).

In the homeownership model, I assume that the probability of an individual choosing homeownership is a logit function of observables X and an unobserved component ν_h,

$$Pr_i(home = 1|X_i, \nu_h) = \frac{\exp\left[X_i\,\gamma + \nu_h\right]}{1 + \exp\left[X_i\,\gamma + \nu_h\right]}. \tag{2.7}$$

Then the contribution to the homeownership likelihood function from an individual i is

$$L_i(t|X_i, \nu_h) = Pr(X_i, \nu_h)^{z_i}\left[1 - Pr(X_t, \nu_h)\right]^{(1-z_i)} \tag{2.8}$$

where z_i is equal to one if individual is homeowner and 0 otherwise.

If I assume that that correlation between the hazard model and homeownership model, beyond correlation captured by explanatory variables, can be represented by the individual-specific heterogeneity terms that are time-invariant and constant across repeated spells, then the complete likelihood function for individual i is

$$L = \prod_{m=1}^{M} \int_{\nu_u} \int_{\nu_h} L_u\left(t|X_m, \nu_u\right)\ L_h\left(t|X_m, \nu_h\right)\ dF\left(\nu_u, \nu_h\right) \tag{2.9}$$

where M is the number of spell individual experiences in the sample period, and $F\left(\nu_u, \nu_h\right)$ is the joint distribution function for the unobserved heterogeneity.

2.3.3 Identification

There are two possible strategies to identify parameters in equation (2.9): i) using multiple spells of unemployment per individual with variation in homeownership status during the spells and ii) using exclusion restriction. The first identification method uses Honoré's (1993) results for duration models with multiple spells which show that parameters are identified if at least a subset of individuals has multiple spells of unemployment across which homeownership status changes, i.e individuals are homeowners in some and renters in other spells. Conditional on $F\left(\nu_u, \nu_h\right)$, the equations are independent and identification depends on intra-person-spell variation in homeownership status. In my sample, only four percent of individuals are observed as homeowners and renters over different spells, which is why I opt for the second methodology. This type of identification was used in the literature on duration models, as well as in a similar context by Munch et al. (2006, 2008).

The second identification methodology uses an exclusion restriction, i.e. a variable or set of variables that affect the choice of homeownership but do not have a direct impact on the labor market. I use state-level homeownership rate as an excluded variable in the duration model, since it is reasonable to expect that the state homeownership rates affects the probability of being a homeowner but does not affect the labor market outcomes. This strategy was successfully used in a similar context by van Leuvensteijn and Koning (2004) and Taşkın and Yaman (2016).

2.3.4 Estimation

I use a finite mixture model proposed by Heckman and Singer (1984a) to estimate equation (2.9), which assumes that the unknown distribution $F\left(\nu_u, \nu_h\right)$ is discrete and that it can be represented with a finite number of points of support. It is assumed that each ν_u and ν_h can take two values with one of the support points being normalized to zero for both ν_u and ν_h because the baseline acts as a constant term in both equations. There are four possible combinations of this bivariate unobserved heterogeneity distribution and their respective probability weights. One can think of each of these values corresponding to a subsample of the individuals in the data set that are not observable to the econometrician. The individual likelihood in equation (2.9) can be rewritten as

$$L_i = \prod_{m=1}^{M} \sum_{j=1}^{4} \left[\pi_j \; L_u \left(t|X, \nu_u\right) \; L_h \left(t|X, \nu_h\right)\right] \tag{2.10}$$

where π_j is parameterized using the logit function. Thus, I estimate the vectors of parameters β and γ, the support points ν_u and ν_h, and the associated probability weights π_1 to π_3[10]. Because the likelihood in (2.10) is not globally concave, I estimate this likelihood via simulated annealing, a global optimization routine[11].

[10] This is a nonparametric maximum likelihood estimator since both points of support and its respective probability weights are estimated.

[11] I use GenSA package available in R to perform simulated annealing (Xiang et al., 2013).

2.4 The Effect of Homeownership on Unemployment Duration

In this section, I present estimation results: first, for the model which does not correct for selection into homeownership in Table 2.5; and second, for the model which corrects for selection into homeownership in Table 2.6. Estimating the model without correction for selection into homeownership is equivalent to estimating the unemployment hazard and homeownership choice models separately. In Table 2.5, the first two columns show coefficient estimates and their respective standard errors for the unemployment hazard model, while the third and fourth columns show estimates and standard errors for the homeownership choice model.

In the context of my econometric model presented in Section 2.3, estimating the model without correction for selection into homeownership is the same as assuming that there are no unobserved differences in exits from unemployment between homeowners and renters after controlling for observable. Results in Columns 1 and 2 of Table 2.5 show that homeowners have significantly higher exit rates out of unemployment than renters. In fact, homeowners are 13 percent ($= \exp(0.122) - 1$) more likely to leave unemployment spells than renters, *ceteris paribus*, which is a similar finding to the univariate analysis of unemployment duration using the Kaplan-Meier estimator. This finding is contrary to the regional evidence presented in Wolf and Caruana-Galizia (2015), who find a positive effect of homeownership on the unemployment rate.

I also find that the exit rate out of unemployment increases with age (Table 2.5). Moreover, I find that better-educated individuals have higher hazard rates compared to those individuals with only an elementary education, but these coefficients have large standard errors and are not statistically significant. Furthermore, I find that individuals with a partner have higher exit rates from unemployment than single individuals, but this effect is only significant for men with an unemployed partner. In particular, men with an unemployed partner are 23 percent more likely to leave unemployment than single men ($= \exp(0.212) - 1$). In regards to non-labor income, I find that a public transfer significantly lower exits out of unemployment while a partner's income increases exits out of unemployment. Finally, I note that a decrease in labor market tightness, a ratio of vacancies to unemployed individuals, increases the exits out of unemployment.

Estimates for the homeownership model, Columns 3 and 4 in Table 2.5, indicate that state-level homeownership rate is an important predictor of being a homeowner. Also, I find that the probability of becoming a homeowner significantly falls with age, which is expected given that purchasing a home often involves taking a mortgage and banks are less likely to approve mortgages for older individuals. On the other hand, the probability of being a homeowner increases with the level of education. A man with an employed partner is more likely to be a homeowner than single men, while a man with an unemployed partner is less likely to be a homeowner when compared to single man. I find that non-labor income coming from public transfers is positively associated with the likelihood of being a homeowner, although it is not statistically significant. Finding positive qualitative effects of public transfers on the probability of being a homeowner might be surprising, but it is expected since the largest portion of non-labor income are unemployment benefits, which are a function of previous wages. Thus, it is expected that men with higher earnings are more likely to become homeowners. I also find that the partner's income is negatively associated with the probability of being a homeowner, but this effect is statistically insignificant. Finally, I find that an increase in labor market tightness has a negative effect on the probability of becoming a homeowner.

The effect of homeownership on unemployment hazard without control for self-selection needs to be taken with caution because survival models are prone to a severe bias in the presence of self-selection (Heckman and Singer, 1984a). It is conceivable to have both upward and downward bias in the homeownership estimate. For example, more productive individuals are more likely to leave unemployment sooner (see column 1 in Table 2.5), and at the same time, they will be more likely to be homeowners due to higher earnings (see column 2 in Table 2.5). A downward bias in the estimate of homeownership might come because some individuals might have strong preferences for location (proximity to family or friends). Then these individuals might be more likely to become homeowners, and in cases of unemployment, these individuals might be less willing to consider employment opportunities outside of local labor markets. Thus, their unemployment spells would be longer. This is why I employ joint estimation of two models and allow two models to be correlated through unobserved effects, while I use exclusion restriction to achieve identification of the homeownership parameter.

In Table 2.6, I present the estimates of homeownership on exit rates out of unemployment. The standard errors are obtained via bootstrap with 1000 replications. I report 95% bias-corrected confidence intervals (Efron and Tibshirani, 1986) because standard inference that relays on asymptotic normality is not appropriate given skewness of the empirical distribution of my estimate for homeownership (see Figure 2.3)[12]. After introducing unobserved heterogeneity in both equations and accounting for selection into homeownership by allowing the unobserved effects to be correlated, I find that homeowners have 56 percent lower unemployment hazards ($= \exp(-0.832) - 1$) and this is statistically significant at a 5 percent level.

The correlation between unobserved effects ν_u in hazard model, and ν_u in unemployment model, is positive and statistically significant which implies the need to account for selection into homeownership. Contrasting the effects of homeownership on exit rates from unemployment without and with correction for self-selection, 0.122 vs. -0.832, indicates that unobserved characteristics that make individuals more likely to be homeowners also increases the rate of exits from unemployment. This is expected because individuals with better labor market opportunities are also more likely to be homeowners. Thus, finding that homeowners have 56 percent lower unemployment hazards than comparable renters after accounting for self-selection supports the Oswald hypothesis that homeownership reduces labor mobility. My finding is also consistent with evidence from the German regional data, which shows that higher rates of homeownership are associated with higher levels of unemployment (Lerbs, 2011; Wolf and Caruana-Galizia, 2015).

In the international context on the effects of homeownership on labor mobility, my results are similar to Burnet and Lesueur (2003), who find that homeownership increases unemployment duration in France. In contrast to Burnet and Lesueur (2003), Munch et al. (2006) find that homeowners have shorter unemployment duration even after correcting for self-selection in Denmark. After distinguishing between types of exits from unemployment, local labor market vs. non-local labor market, Munch et al. and van Leuvensteijn and Koning (2007) find that homeowners have lower exit rates from unemployment to the non-local labor market, and higher exit rates from unemployment into local labor markets in Denmark and the Netherlands. Thus, their findings provide empirical support for Oswald's hypothesis, mainly that homeownership hinders geographical labor mobility, but this mechanism is not strong enough in these two countries. A possible explanation for differences between Germany and France, on one side, and Denmark and the Netherlands, on the other side, might be in the countries' geographical sizes, which also affects the size of the local and national labor market.

[12] I use bias-corrected confidence intervals over accelerated bias-corrected confidence intervals due to computation difficulty with estimating acceleration factor.

2.5 Conclusion

I study the effect of homeownership on unemployment duration using unemployment spells from the German-Socio Economic Panel for the period 1990–2011. Although the economic theory predicts that homeownership will reduce exits from unemployment to non-local jobs, i.e. jobs that require a move, the theory does not provide a clear prediction about the total rate of exits from unemployment, since this depends on both local and non-local labor markets. In addition, there is self-selection into homeownership that can produce both upward and downward bias. For example, some households might be inherently less mobile, e.g preferences for proximity to family, in which case these households choose homeownership rather than renting. Then in the case of unemployment, these men might stay longer in unemployment spells but not due to homeownership. It is also possible that some men are more productive and are more likely to be a homeowner, and in case of unemployment, these men have better labor market opportunities. I correct for possible self-selection into homeownership by simultaneously estimating the hazard model, with the homeownership model and at the same time I allow the two models to be correlated by specifying a bivariate distribution for two unobserved variables, one of which affects the hazard model while the other affects the hazard rate. Additionally, I minimize the biasing impact of distributional assumption by estimating it nonparametrically. I use state-level homeownership rates as an exclusion variable in the hazard model to identify the parameter.

My initial findings suggest that homeowners have a median unemployment length that is two months shorter than renters. After introducing other observable controls, I find that homeowners have a 13 percent higher rate of exits from unemployment than renters. However, I find that homeowners are 56 percent less likely to leave unemployment compared to renters when I correct for self-selection into homeownership, and this effect is statistically significant at a 5 percent level. These findings are in line with previous research that used regional level data, and indicates a potential negative consequence of homeownership on the unemployment duration. Furthermore, a potential extension of this paper is to examine how homeownership affects unemployment exits into a local labor market vs. how it affects the unemployment exits that require the change of residency. Finally, it would be interesting to evaluate the joint effects of homeownership and unemployment benefits on the unemployment hazard due to policy implications.

Table 2.1: Summary of Unemployment Spells

	Homeowners	Renters
Survival time (months)		
25 percentile	2	2
50 percentile	5	7
75 percentile	13	19
Exits from unemployment (percent)		
Full-time	93.2	92.9
Part-time	6.8	7.1
Total spells ending in employment	80.3	73.5
Distribution of number of spells (percent)		
1	57.5	60.1
2	21.8	22
3	9.9	9.5
4	4.7	4.5
>4	6.1	3.9
Spell [Exits]	1134[911]	1943[1428]
Individuals	709	1189
Observations (person-spell)	1134	1943

Note: Author's calculation.

Table 2.2: Descriptive Statistics

	Homeowners	Renters
Age (%)		
25–30	16.58	26.45
31–35	13.14	16.11
36–40	18.87	19.92
41–45	17.72	14.51
46–50	16.67	12.45
51–55	17.02	10.55
Education (%)		
Elementary	0.06	0.10
Apprenticeship	0.30	0.32
High-school	0.49	0.44
College	0.15	0.14
Partner (%)		
Single	0.25	0.30
Employed partner	0.54	0.43
Unemployed partner	0.21	0.28
Children (%)		
0-12	0.31	0.35
13-18	0.26	0.17
Non-labor income (Real Euros)		
Partner's wage	15,974	13,841
Public Transfers	7,486	7,061

* Means are taken over the spells.

Note: Author's calculation.

Table 2.3: Homeownership Rates by State, 1990–2011

State	Mean	Standard Deviation
Baden-Wüerttemberg	54.51	2.17
Bavaria	53.35	1.70
Berlin**	11.24	3.05
Brandenburg*	40.05	4.12
Bremen	36.71	9.18
Hamburg	20.02	4.11
Hessen	45.57	3.33
Lower Saxony	40.69	7.14
Mecklenburg-Western Pomerania*	54.94	2.84
North Rhine-Westphalia	42.47	3.25
Rhineland-Palatinate	55.91	2.61
Saarland	61.95	7.49
Saxony*	33.67	5.25
Saxony-Anhalt*	41.56	5.58
Schleswig-Holstein	57.11	4.27
Thuringia*	45.09	3.64

Source: German Socio-Economic Panel

* State was part of the Former German Democratic Republic (East Germany)

** Berlin was dived between East and West Germany.

Note: Author's calculation.

Table 2.4: Homeownership Rates by State From Microcensus

State	1998	2002	2006	2010
Baden-Wüerttemberg	48.3	49.3	49.1	52.8
Bavaria	47.6	48.9	46.4	51.0
Berlin	11.0	12.7	14.1	14.9
Brandenburg	35.5	39.8	39.6	46.2
Bremen	37.5	35.1	35.4	37.2
Hamburg	20.3	21.9	20.2	22.6
Hessen	43.3	44.7	44.3	47.3
Mecklenburg-Western Pomerania	32.2	35.9	33.2	37.0
Lower Saxony	48.9	51.0	49.0	54.5
North Rhine-Westphalia	37.4	39.0	38.7	43.0
Rhineland-Palatinate	55.0	55.7	54.3	58.0
Saarland	58.1	56.9	54.9	63.7
Saxony	28.7	31.0	29.5	33.7
Saxony-Anhalt	36.5	39.6	37.9	42.7
Schleswig-Holstein	46.8	49.4	47.1	49.7
Thüringia	39.2	41.8	40.6	45.5
Germany	**40.9**	**42.6**	**41.6**	**45.7**

Source: German Microcensus

Table 2.5: Estimates of the Unemployment Hazard and Homeownership Choice without Unobserved Heterogeneity

Variable	Unemployment Hazard		Homeownership Choice	
	Coefficient	Standard Error	Coefficient	Standard Error
Homeowner	.122	.050	–	–
State Homeownership Rate	-	-	2.738	.439
Age	.064	.008	−.134	.016
Age^2	−.001	.0003	.002	.0002
Education				
Apprenticeship	−.046	.097	.173	.163
High school	−.127	.095	.455	.159
College or Higher	−.058	.109	.374	.181
Partner				
Employed Partner	.153	.091	.405	.149
Unemployed Partner	.212	.073	−.171	.120
Non-labor Income				
Public Transfer	−.031	.012	.020	.019
Partner's Income	.014	.008	−.025	.013
Labor Market Tightness	−2.131	.432	−2.600	.834
ln(L)	−1330.3		−1926.0	
Individuals	1898			
Observations	3077			

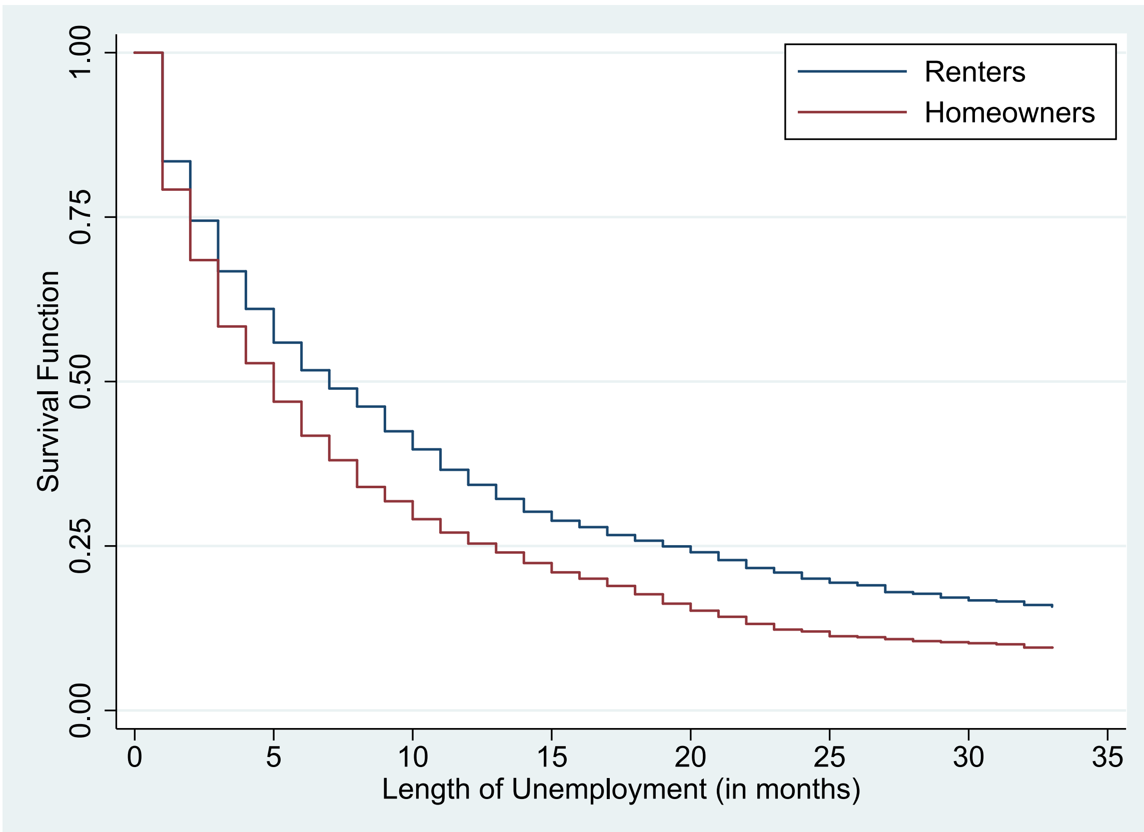
Note: Estimating the model without correction for selection into homeownership is equivalent to estimating the unemployment and homeownership models separately. The unemployment model is specified as Cox duration model, while the homeownership model is logit. The standard errors are clustered at individual level to account for the presence of multiple spells per individual. Estimate of the baseline are not reported.

Table 2.6: Main Estimation Results with Exclusion Restriction

Variable	Unemployment Hazard Coeff.	95% CI		Homeownership Choice Coeff.	95% CI	
Homeowner	$-.832$	-2.460	$-.464$	–	–	–
Homeownership Rate	–	–	–	1.923	$-.096$	2.700
Unobserved Effects						
ν_u	2.775	.331	4.350			
ν_h	1.960	.457	4.500			
$Corr(\nu_u, \nu_h)$	.069	.007	.130			
Support Points						
π_1	.300	.133	.718			
π_2	.115	.002	.302			
π_3	.056	.001	.295			
π_4	.529	.329	.789			
Log			3258.15			
Individuals			1898			
Observations (person-spell)			3077			

Note: Estimation is performed using simulated annealing, global optimization technique, in R with package GenSA. I reported 95% CI are biased-corrected intervals as described in Efron and Tibshirani (1986) because standard inferences that rely on asymptotic properties is not appropriate due to skewness in the empirical distribution of the estimators. The standard error for the correlation coefficient is calculated based on 1000 draws from the multivariate normal distribution with mean and covariance matrix set to the estimated parameter vector and covariance matrix.

Figure 2.1: Unemployment Survivor Functions by Homeownership Status

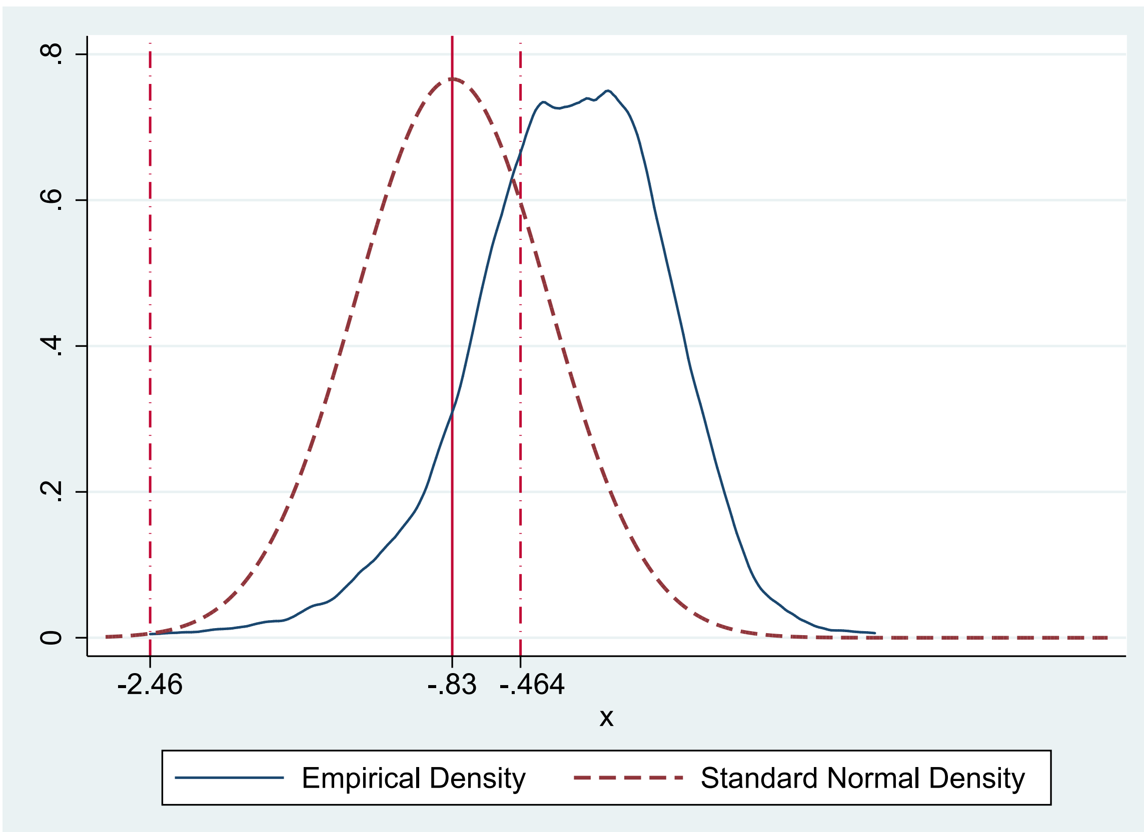


Source: Author's calculations.
Note: The survivor functions are calculated using the Kaplan-Meier estimator. The Peto-Peto-Prentice rank test of equality between survivor functions yields p=0.001.

Figure 2.2: Homeownership Trends by State, 1990–2012

Source: Author's calculations using the GSOEP.
Note: Individual data for Saarland is only available from 2000 in the GSOEP.

Figure 2.3: Empirical Distribution of Homeownership Estimate



Note: The empirical distribution of the homeownership estimate is obtained via kernel estimator. The solid vertical line represents point estimate, while the dash-point vertical lines represent 95% bias-corrected confidence interval (Efron and Tibshirani, 1986).

Chapter 3

Convergence in Unemployment Duration and Its Determinants Between East and West Germans: Is the Berlin Wall Gone?

3.1 Introduction

All Eastern European countries with Communist pasts underwent economic and institutional transitions during which unemployment soared, and the German Democratic Republic (GDR) was no exception to this rule despite institutional and financial help from the Federal Republic of Germany, later referred to as West Germany. In the former GDR, later referred to as East Germany, unemployment rate for all civilians in labor force was 15 percent in 1994, rising to an all-time high of 19 percent in 2005, and then falling to about 10 percent in 2014[1]. High and persistent unemployment rates are detrimental for long-run labor market outcomes such as wages, future spells of unemployment, crime rates, and health outcomes (Ruhm, 1991; Hamilton, Merrigan and Dufresne,

[1]The initial effect of German unification on the East German labor market was a dramatic fall in employment, which declined by 35 percent between 1989 and 1992, according to official statistics of the Bundesagentur für Arbeit (the Federal Employment Agency).

1997; Raphael and Ebmer, 2001; Arulampalam, 2001; Mroz and Savage, 2006). Hence, it is important to understand the differences in duration of unemployment and its determinants between Eastern and Western Germans after unification and how these differences evolve over time for labor market policies. In this paper, I study how duration of unemployment and its determinants related to age, gender, education, and family are different between Eastern and Western Germans after unification, and how these differences evolve over time.

The strengths of this study are in examining the longer time period after unification and distinguishing between unemployed individuals "in" and "out" of the labor force. Hunt (2004), for example, uses the 1990–1999 period and finds persistent differences in duration of non-employment between East and West Germany, and a weak convergence in the role of determinants of non-employment durations. Because economic transition involves changes of both formal and informal institutions, a longer time horizon is essential for assessing the transition. Thus, I use the period 1990–2012 to study East German transition. Furthermore, in this study, I differentiate between unemployed individuals "in" or "out" of the labor force. It is important to differentiate between the two types of unemployment, since these two unemployment states are behaviorally distinct. During the early days of unification, the majority of unemployed Easterners involuntary transitioned from employment to unemployment, which means that they are behaviorally more comparable to unemployed West Germans "in" the labor force than those "out" of the labor force.[2] Thus, to make more accurate comparisons of differences in duration and its determinants between East and West, I investigate unemployment spells of individuals 18–55 whom I observe transitioning from employment to unemployment.

Despite the financial assistance and "know-how" help, East Germany received in the process of setting up institutions, the unemployment rate in East Germany was consistently high. Previous literature attributes the lethargic performance of the unemployment rate in East Germany to market distortions, policy interference, and high adjustment costs (von Hagen and Strauch, 1999; Snower and Merkl, 2006; Burda, 2006). Others argue that the Western labor unions' assistance to the Eastern labor unions in wage negotiation led to the fast growth of wages during early unification, which ultimately acted as a wage floor (Sinn and Sinn, 1992; Siebert et al., 1991; Link, 1993; Lange

[2]The East German labor force participation rate was 10 percentage points higher than West German rate for workers 16–65 (Bonin and Zimmermann, 2001, pp.18). The difference in the labor force participation rates was even larger between females, 75.7 percent in East Germany compared to 55 percent in West Germany (Bonin and Zimmermann, 2001, pp.19).

and Paugh, 1998). Zimmermann (1993) and Vogler-Ludwig (1997) argue that supply-side factors, such as high female labor force participation, are a major reasosn for high unemployment in East Germany.

In addition, literature points at the long-lasting effects of Communist institutions on economic outcomes. For example, Alesina and Fuchs-Schündeln (2007) show the evidence of communism having effects on not only outcomes, but also economic preferences in Germany. In addition, Fuchs-Schündeln and Masella (2016) find that an additional year of socialist education substantially decreases the probability of obtaining a college degree and adversely affects labor market outcomes such as employment, working hours, and wages, even 15 years after the collapse of the East Germany. In section 3.2, I provide an in-depth review on the effects of East-West unification on the German labor market and the key labor market developments over the last 28 years.

In Section 3.3, I summarize information about the dataset, and examine changes in unemployment duration and its determinants over time. I use the German Socio-Economic Panel from 1990 until 2012 to study unemployment duration and its determinants. I narrow my sample down to individuals who are 18–55 years old and whom I observe transitioning from employment to unemployment. These individuals are more likely to reflect the behavior of the unemployed "in" the labor force. I also separate unemployment into three groups by spell's origination time: 1990–1995, 1996–2003, and 2003–2012. I separate these spells using two important milestones: i.) ending of privatization of publicly owned enterprises in East (1995), and ii.) beginning of labor market reforms (2002).

Comparing composition of these samples across determinants and time periods shows that there are differences between East and West Germans regardless of gender. For example, unemployed East females are more educated, less likely to be married, and less likely to have children that are infants than West females during the period 1990–1995. Similarly, I find that East males are older, better educated and more likely to be married than their West counterparts. I find that over time differences in the composition of the samples between East and West Germans diminish.

Furthermore, a comparison of unemployment duration shows that East Germans remain in unemployment for a shorter time than West Germans, especially during long spells of unemployment. For example, there are no differences in median duration of unemployment between East and West males regardless of the time period, while I find that East males have six months shorter unemployment duration at the 75th percentile of unemployment duration. Similarly, I find that East females' median duration is about one month shorter regardless of time period, and that they have nine months shorter unemployment duration at 75th percentile. Interestingly, both East and West females spend less time in unemployment during later time periods while the East-West difference persists, i.e.the West female unemployment spells are shortening but the East females still have shorter spells.

In Section 3.4, I develop an econometrics framework to analyze the roles of age, gender, education, and family in unemployment duration, and asses how the effects of these detriments on unemployment duration change over time. I use the panel logit estimator to estimate the effects of determinants on unemployment duration as described by Kiefer (1988) and Sueyoshi (1991). This model is more appropriate than the Proportion Cox Model, given the grouped nature of my data, i.e. when exits from unemployment are observed or recorded at aggregated time intervals such as weeks or months.

In Section 3.4.3, I summarize the effects of the role of determinants on unemployment duration, and how these effects change over time. I find that the role of the determinants changes over time for East Germans, both males and females, and West females, while the effects do not change significantly for West males. Furthermore, I find that the effects of the determinants are significantly different across East and West Germans. My results show that initially education and age have the largest disproportional effect on unemployment lengths of East and West Germans. I also find that the role of education in the determining unemployment duration experiences the largest change over time in East Germany, followed by the role of age. During the early period after unification, education plays no role in the East unemployment duration of both genders, which is not surprising since the East economy was in total disarray after unification. Education becomes an important determinant of the hazard function in later periods for Eastern Germans. Additionally, I find the emergence of the educational differentials between East and West Germans with education being significantly more important in reducing unemployment duration for East Germans.Finally, I find that the role of children and relationship with a partner plays a similar role in unemployment duration since unification.

My findings contribute to the literature on the long-lasting effects of Communist institutions on economic outcomes and to the literature which explains persistently high unemployment rates in East Germany during the 1990s and early 2000s. Unlike Hunt (2004), I show that the differences in unemployment duration and its determinants exist during the initial period of transition but these differences shrink over time. My findings suggest that the effects of Communist institutions and transition at least do not persist in the duration of unemployment and its determinants between East and West. Furthermore, I show that the role of age, education, children, and relationship status are not drivers of the differences in unemployment in the late 1990s and 2000s. Finally, in Section 3.5, I present some final remarks with potential questions for further research.

3.2 German Unification and Labor Market

East and West Germany entered monetary, economic, and social union on 1st July 1990. At the time of unification, the East German population was 20 percent of unified Germany; in terms of area, the East makes up 30 percent of all Germany; and the East's pre-unification GNP was 15 percent of the combined total (Sinn and Sinn, 1992, pp.22)[3]. East Germany had poor infrastructure compared to the West, outdated Soviet technology, non-existing free-market institutions, and lack of experience with prices and markets. For example, 70 percent of the housing stock had been constructed before World War II, more than twice the proportion in the West. The East's industrial production was dirtier than in West Germany or in some Eastern European countrie; sulfur dioxide emission per capita was 7 times higher than West Germany. In comparison to the Eastern European Communist countries, sulfur dioxide emission per capita were 50 percent worst than in Czechoslovakia and 100 percent worse than in Hungary and Poland. About a thousand of districts did not have access to drinkable water, and a third of the industrial wastewater was untreated when it left plants. Chemical contamination was reaching extremely high levels; up to 50,000 industrial sites were polluted by chemicals. In terms of capital stock, 50 percent of it was obsolete, and no less than 67 percent would have been written off under Western accounting rules[4].

The Western institutions, including the systems of justice, regulations, industrial relations, banking, education, social security, and welfare were set in place after unification, and only a small number of institutions were subject to the gradual introduction. For example, property rights over public firms were resolved by the end of 1995 through privatization. Unification brought institutional shock that caused a sharp drop in Eastern GDP and employment; East GDP fell by 30 percent and employment fell by 35 percent between 1989 and 1992. In the years after 1992, the East's economy began with rapid growth; the ratio of Eastern to Western GDP per capita grew from 49 to 66 percent between 1992 and 1995.[5] However,the ratio of Eastern to Western GDP per capita stagnated around 66 percent during 1996–2004, indicating the East's economy had reached a plateau. The driving component of growth was financial aid coming from West Germany which was mainly spent on infrastructure and welfare. From 1990 to 2015, this assistance accumulated to about USD 2 trillion, and today this assistance is around USD 80 billion per year (Gehler, 2011).

[3] Assuming 1:1 exchange rate.

[4] The statistics on conditions of housing stock, pollution, and capital stock in East Germany are from Sinn and Sinn (1992).

[5] The GDP statistics are obtained from Statistische Ämter Bundes und der Länder.

The East labor market went through turbulent phases especially during the early years of transition. The employment to population ratio fell by 17 percent between 1990 and 1992, and it was 65 percent compared to 73 percent in West Germany in 1999. Examining the unemployment rate in East Germany (excluding East Berlin) indicates a similar trend as GDP statistics. The rate increased from 10 percent to 19 percent between 1991 and 1995, where it remained until 2006. The lack of search requirements for the registered unemployed is partly responsible for the high unemployment rates. However, these rates would probably be higher if there were no early retirement program during the early days of unification or active labor policies such as retraining or public work. The active labor policies have decreased unemployment rates since the participants of these programs would have been unemployed had they not taken early retirement or entered retraining programs and public works.

For example, the temporary early retirement program allowed individuals 55 or older to retire with Western-level benefits, and the number of retirees peaked at 854,000 in 1993. The public works jobs peaked at 388,000 in 1992 and remained around 300,000 until 1998. The retraining programs peaked at 372,000 in 1992 and declined to 147,000 by 1998. The portion of the labor force involved in the programs peaked at 18 percent in 1992 and fell to 7.5 percent in 1998[6].

In the rest of this section, I discuss the events that contributed to the weak performance of the East labor market during unification's early period and the effects of labor reforms known as Hartz I–IV that were implemented in the early 2000s. These were the events that contributed to the weak performance of the labor market during 1990-2004: aggregate demand shocks, the one-for-one exchange rate between Ostmarks and the Deutschmarks, wage hikes induced by unions, privatization of public firms, and social welfare. The reasons behind the sharp drop in East GDP and employment in the early years of transition are attributed to a large fall in aggregate demand shock caused by the removal of price controls, substitution of the Eastern goods with the Western, and the collapse of trading arrangements with the former Communist Eastern European countries. From 2002 until 2004, the German labor market went through the Hartz I–IV reforms with the aim of decreasing structural (long-term) unemployment.

In the East Germany, wages rose fast after unification especially from 1990 to 1993 because of the assistance Eastern unions received from Western unions in negotiation with managers of public

[6]The unemployment and employment rates are from the Bundesagentur für Arbeit. Statistics on the programs for unemployed and early retirement are from Sachverständigenrat various years.

firms, who did not have any incentive to resist wage hikes. These wages acted as a wage floor and are considered to be the single most important factor that caused the fall in East German employment (Sinn and Sinn, 1992; Akerlof et al., 1991). After unification, West unions feared that there would be a large migration of workers from East to West, and as a result, helped East unions negotiate high wages. The data from the German Socio-Economic Panel indicates that the median real monthly wage for workers 18–55 rose 83 percent between 1990 and 1993 in East Germany. The growth in wages considerably slowed down after 1993. In 1993, the East-West ratio of nominal monthly wage was at 71 percent, and the ratio in 1999 was only 76 percent.

As already mentioned, a part of the reason behind the fast rise in wages were the lack of property rights over firms during the early years of transition. By the end of 1992, the majority of small firms were privatized, while the large industrial firms were privatized by the end of 1995. There were 8,500 large firms with 44,000 plants and 45 percent of the workforce (Carlin, 1994). Unprofitable firms and plants were closed while the remaining firms were sold mainly to the Western firms that were in the same industry (Dyck, 1997). The Western firms obtained 74 percent, the Eastern individuals 20 percent, and the foreign firms obtained 6 percent of Eastern firms when weighted by employment (Carlin, 1994). Despite the increase in unemployment in the early days of transition, in general, privatization is considered to be successful because it promoted employment in the medium and in the long run.

von Hagen and Strauch (1999) pointed at the welfare system as a reason for high unemployment in East Germany in the medium run and the long run. Indeed, prior to 2003 and Hartz reforms unemployed individuals would receive between 60 percent and 67 percent of the previous income for up to 32 months depending on age. Following the expiration of unemployment benefits, the individual would receive unemployment assistance between 53 percent and 57 percent of the previous income indefinitely. An individual was also qualified for welfare benefits, means-tested benefits, to ensure decent living conditions. These benefits were 70 percent of the lowest wage in the industrial sector, and 100 percent of the lowest wage in low-paid sectors. Besides the generous unemployment benefits, individuals were not incentivized to search for jobs because there were no search requirements. Thus, the generous welfare system increased the reservation wages, which in turn raised the duration of unemployment.[7]

Hartz reforms were enacted in stages: Hartz I and II in January 2003, Hartz III in January

[7] The facts about the German unemployment benefits originate from Gaskarth (2014).

2004, and Hartz IV in January 2005. The primary goal of these labor supply-side reforms was reducing structural (long-term) unemployment in Germany through reorganization of the Federal Labor Offices, promotion of part-time jobs and self-employment, the introduction of penalties if individuals rejected a reasonable offer to work, and reduction of long-term unemployment benefits. Among these reforms, the reduction of long-term unemployment benefits was probably the most important since the old system of benefits was replaced with a flat-rate benefit of 359 euros a month (not including rent). Nagl and Weber (2014) find that hazard rates increased by 24 percent after reform and that the effect is more prominent in East Germany than West Germany. Their results also indicate that Hartz reforms were more beneficial to short-term unemployed rather than the long-term unemployed individuals.

Although the Eastern labor market and income levels have not yet converged to the Western, it is reasonable to expect that the East labor market will resemble the West labor market quite closely in the long run. There are no geographical barriers or significant differences between East and West Germany; East Germany is not behind in terms of education; the two regions share the same language and similar culture; and there are no differences in the formal institutional framework; there are no restrictions on the mobility of labor or capital. Given the facts about the institutions, there is no reason to believe that income levels will not converge as well as the behavior of unemployed individuals in the labor market.

3.3 Differences in Unemployment Duration and Its Determinants Between East and West Germans

3.3.1 Data Description

The German Socio-Economic Panel (GSOEP) is a representative annual panel survey of households and individuals. It began collecting survey data in 1984 in West Germany, and East Germany was included in 1990. The main advantages of these data are the ability to identify whether the individuals are East or West German, a large array of socio-economic variables, and the panel structure which allows following individuals through their employment biographies. The survey identifies an individual's residence during 1989 just before unification, which is how I defer between East and West Germans.

The calendar section of the panel follows an individual's labor status during the 12 months prior to the year of survey. In the calendar section of the GSOEP, individual chose among the following categories of labor status: full-time employed, short-work hours, part-time employed, vocational training, unemployed, retired, maternity leave, school or college, military or community service, housewife or husband, second job, other, first job–training (apprenticeship), continuing education (retraining), and mini–job (up to EUR 400).

The survey does not limit the number of labor statuses during each month, so 70 percent of the observations have multiple statuses. For example, an individual could declare to be in full-time employment and at the same time to be a housewife (husband), or in vocational training and an at the same time to be full-time employed during the same month. It is also possible to have individuals declare being unemployed and full time employed during the same month. I resolve this issue of multiple statuses during each month by sorting employment statuses and dropping irrelevant statuses, so that each individual has a unique spell for each month. The employment statuses dominate other statuses, while the observations with the conflicting statuses such as full–time employed and unemployed during the same month are dropped from the analysis.

After establishing a unique labor status during each month for every individual, I keep only unemployment spells for which I observe the starting date. This approach has two advantages: first, the sample is more representative of unemployed persons since the individuals are more likely to be searching for a job immediately after becoming unemployed; and second, I avoid making assumptions about durational dependence. This is very important since the force statuses of the unemployed “in” and “out” of the labor force are behaviorally distinct (see: Heckamn and Flinn, 1983; Tano, 1991), which means that studying their behavior jointly can produce misleading conclusions. Combining the two unemployment states is even more problematic in the context of East German transition, because the majority of Easterners went to unemployment involuntary after unification. Thus, East German sample of unemployed individuals is more likely to have a greater portion of individuals actively looking for jobs than West German sample, which means that such comparison is deemed to find more persistent differences in the duration and determinants of unemployment.

My sample has 7,813 individuals: 3,116 are Easterners and 4,697 are Westerners. The sample has 16,319 spells of which 6,755 are Eastern spells and 8,911 are Western spells (see Table 3.2). Thus, on average Easterners have 2.2 spells and Westerners 1.9 spells during the period from 1990–2012. Because I am interested in the effect of transition on the duration of unemployment and its determinants, I divided spells into three groups by time of spell origination: 1990–1995, 1996–2002, and 2003–2012. These periods were chosen according to two important milestones in the German labor market. First, privatization of public enterprises in East Germany ended in 1995; thus, the period 1990–1995 is an era of institutional adjustment. Second, in 2002, Germany began overhauling its welfare system related to unemployment and both parts of Germany were equally affected by these changes.

In the sample, it is possible for an individual to have multiple spells of unemployment over this period of 22 years. Table 3.1 summarizes the number of individuals by type of spells. Table differentiates between individuals who had one spell of unemployment and those with multiple spells of unemployment in different periods. Approximately 30 percent of East Germans and 23 percent of West Germans have multiple spells of unemployment. Furthermore, approximately 76 percent of East Germans and 80 percent of West Germans with multiple spells experience unemployment in adjacent periods.

3.3.2 Differences in Unemployment Determinants and Their Evolution over Time

In Tables 3.3 and 3.4, I present averages of the following unemployment determinants: age, level of education, the presence of children, the status of a relationship with a partner, and the distance from the closest metropolitan area by gender for East and West. All determinants are time-invariant binary variables except for the time-varying indicators for the presence of a child in the family.

The indicators of age at which women enter unemployment, in Table 3.3, show that the largest share of female unemployment spells occurs during ages 26–35 regardless of origin. I also observe that Western females are 13.6 percentage points more likely to be in unemployment during age 26–35 than Eastern females during 1990–1995. This differential falls over time and in the period 2003–2012 is only two percent. In the last period, 2003–2012, Eastern females are nine percentage points more likely to be in unemployment during age 18–25 while they are 11 percentage points less likely to be unemployed in the age group 36–45 than their Western equivalents. Hence, there is convergence in the composition of unemployment spells by age between females.

Furthermore, both Eastern and Western females have either apprenticeship or vocational education in approximately 82 percent of unemployment spells during 1990–1995. Eastern females with higher education are 7.8 percentage points more likely to have a spell of unemployment than comparable Western females during 1990–1995. In 2003–2012, the share of spells of females, both Eastern and Western, with an apprenticeship or vocational education falls by approximately 13 percentage points, while there is a rise in the share of female spells with higher education, six percentage points in the East and ten percentage points in the West. Also, the East-West female differential in the share of spells with higher education falls to four percent in 2003–2012. Thus, the differences in education level between East and West females during unemployment spells fall over time.

Examining the differences in relationship status during unemployment spells in the 1990–1995 period indicates that: i.) Eastern females are five and seven percentage points more likely to be unemployed while single or while in cohabitation than comparable Western females; and ii.) Eastern females are 11 percentage points less likely to be married while in unemployment. In the 2003–2012 period, these differences increase. For example, Eastern females are 11 and 15 percentage points more likely to be single or in cohabitation than comparable Western females. Also, Eastern females are 23 percentage points less likely to be married while in unemployment. Although the East-West female differential in relationship status is increasing over time, the general trend is that both East and West females are more likely to be single or in cohabitation while unemployed.

There is at least one child ages 0–6 in 42 and 65 percent of East and West female unemployment spells during 1990–1995. This differential is almost completely eliminated for unemployment spells that originated during 2003–2012. Examining this trend over time indicates an increase in the share of spells with a child of ages 0–1 in the East by about 18 percentage points, while in the West it remains approximately the same. In contrast, the share of spells with a child of ages 2–6 does not change in East while it falls by seven percentage points in the West over time.

Finally, there is no significant differences between East and West females in regards to the location of living, measured as a distance in kilometers from a city center. Furthermore, it is important to note that there is an increase in the number of East females living in the West German states over time by around 11 percentage points. In conclusion, while there are some differences in the determinants of unemployment duration between East and West females during the initial period, these differences fall over time. I also note that the role of children and relationship status are changing over time in both East and West.

In Table 3.4, East males ages 26–35 make up 30 percent of unemployment spells while other age groups make up the rest of unemployment spells in approximately the same amount during the 1990-1995 period. In contrast, during the same period, Western males less than 25 years old make up 40 percent of unemployment spells, while Western males ages 26–35 are in 31 percent of spells. These differences between East and West males disappear in 2003–2012 period. Moreover, Eastern males are eight percentage points less likely to be in the age group 36–45 and 8 percentage points more likely to be in the age group 46–55 than their Western equivalents during the 2003–2012 period. Thus, both East and West males are more likely to be in unemployment if older than 36.

In 76 percent of unemployment spells, both East and West males had either apprenticeship or vocational education during the period 1990–1995. East males in 16 percent of unemployment spells have a higher education, while West males have only 5 percent during the same period. In the period 2003–2012, East males with an apprenticeship or vocational education make up 86 percent of unemployment spells, i.e. an increase of 10 percentage points from the 1990–1995 period.In contrast, the share of Western males with apprenticeship or vocational education did not change in later periods. Interestingly, East males with higher education account for approximately 6 percent of unemployment spells, a fall of 10 percentage points compared to the 1990–1995 period. At the same time, West males with higher education make up 11 percent of unemployment spells, an increase of six percentage points compared to the 1990–1995 period. Thus, there is a slight divergence in education composition of unemployment spells between East and West males.

During the 1990–1995 period, married East males account for 57 percent of spells followed by single East males with 32 percent of spells. In contrast, married Western males make up only 35 percent of total spells while single West males make up about 50 percent of total unemployment spells during the same period. In the period 2003–2012, the share of married East males fell by 22 percentage points, while the share of single East males and the share of those in cohabitation increased by approximately 11 percentage points each. In contrast, the share of married West males and the share of West males cohabiting increased by 4.6 and 2.7 percentage points from 1990–1995, while the share of single West males decreased by 7.3 percentage points compared to 1990–1995. Thus, there is convergence in the composition of the samples in terms of relationship status with a partner between East and West males.

In 42 and 31 percent of unemployment spells, East and West males had a child present in the household during 1990–1995. There was a decrease in the presence of children in the household for both East and West unemployed males by 20 and 4 percentage points, respectively during 2003–2012. For both East and West males, this fall came because of a decrease in the shares of spells where a child is 2–6 years old and 7–11 years old. Thus, there is convergence in the make up of unemployment spells with respect to the presence of children in the household. Finally, I want to note that no clear pattern exists in differences in terms of location of living measured as a distance to a city center between East and West males. Additionally, between 91 and 97 percent of East males live in the former GDR East depending on the period while approximately 98 percent of West males live in the former FRG .

To summarize, the determinants of unemployment duration for the most part suggest convergence, i.e. composition of East and West unemployment spells looks more similar over time. The composition of unemployment spells for East and West Germans in 2002–2013 is much more similar than in 1990–1995 by age, education, relationship status, and the presence of children in the household.

3.3.3 Convergence in Unemployment Duration between Eastern and Western Germans

In this section, I examine the differences in unemployment duration using percentiles and unconditional probability of leaving unemployment. Table 3.2, columns 1–3, shows the estimates of unemployment duration for 25th, 50th, and 75th percentiles over the three periods for the Easterners. In the same table, columns 4–6 present the same information for the Westerners. In addition, columns 7–8 show estimates of unemployment duration for the 25th, 50th, and 75th percentile during the 1990–2012 period. During 1990–1995, there are no differences in the 25th percentile of unemployment duration between East and West Germans, while the median and 75th percentile durations are lower for Eastern Germans regardless of gender. This suggests that the Easterners return to work faster than the Westerners. Comparing the median and 75th percentile for East and West Germans within gender over time shows that unemployment length is falling in both the East and West, as well as convergence in the duration of unemployment. These preliminary statistics of unemployment duration suggest that the unemployment durations between Easterners and Westerners are more alike than different.

Furthermore, using the Kaplan-Meier estimator, the Wilcoxon (Breslow-Gehan), the Tarone-Ware, and the Log-rank test statistics, I examine differences in unconditional probability of leaving unemployment in East and West Germany over time. Figure 3.1 presents survivor functions by gender for the Easterners and Westerners for 1990–2012. Irrespective of whether I look at East or West, female survivor functions are higher than male survivor functions, i.e. females are less likely to return to employment at any duration. All three tests at the 10 percent level show no significant differences in the survivor functions between East and West males. In the case of female survivor functions, the Log-rank and the Wilcoxon test imply that female survivor functions are significantly different at a 6 percent level.

Pooling spells together ignores the effects of changes in institutions over time and assumes that spells beginning immediately after the fall of the Berlin Wall are the same as the spells that began in the late 2000s. To better understand the changes in unemployment duration, I estimate the unconditional probability of leaving unemployment for East and West Germans for the spells that originated in 1990–1995, 1996–2002, and 2003–2012, separately. Figures 3.2 and 3.3 plot survival functions with respect to the time spells originated for males and females separately.

Figure 3.2 indicates that female survivor functions are falling over time in both the East and West. All three tests reject the joint hypothesis of female survivor functions being the same at 1 percent level within both the East and West. The estimates of the median and 75th percentile durations drop by five months and 13 months respectively for Eastern females, while the same estimates drop by 5 months and 18 months for the Western females (see Table 3.2). Comparing

East to West female survivor functions for each period shows that the largest difference was during the 1990–1995 period when the Log-rank and the Tarone-Ware tests reject the null hypothesis at 1 percent confidence level. The differences between Eastern and Western female survivor function falls over time while these differences rename significantly different.

Figure 3.3 shows that the East male survival functions do not change for shorter durations of unemployment, but the functions increase for longer durations over time. The figure also shows that the West male survivor functions are falling over time for shorter spells. The Tarone-Ware test fails to reject the hypothesis that these survivor functions are changing over time for both the East and West. The Log-rank and Tarone-Ware tests suggest that there are significant differences between East and West male survivor functions at 7 percent confidence level for spells originating during 1990–1995, and fail to reject hypothesis for other periods. Thus, differences between male survivor functions disappear over time.

In summary, East Germans stay shorter in unemployment than West Germans, and the difference is especially drastic at longer lengths of unemployment. Furthermore, I find that for both East and West Germans the unemployment duration is falling over time, which indicates some evidence for convergence in unemployment duration. I find interesting developments with the differences in unemployment duration between the East and West females. The differences between East and West females survivor functions are not falling in the expected directions with East survival functions becoming more like the West; but rather, I find that the Western female survivor functions are falling towards Eastern female survivor functions.

3.4 Convergence in the Role of Unemployment Determinants

3.4.1 Econometrics Model

In this section, I present the econometric models used to estimate the effects of the determinants on unemployment duration and to assess the evolution of these effects on unemployment duration. I use the Discrete-Time Binary Choice model, a binary response model for discrete (grouped) duration (Kiefer, 1988; Sueyoshi,1991), rather than the Cox Proportional Model (Cox, 1972). The Discrete-Time Binary Choice model is more appropriate for the discrete (grouped) data and it is less computationally demanding when modeling unobserved heterogeneity.

When it comes to modeling the effects of determinants on the duration of unemployment, a common practice is to use the Cox Proportional Model (CPM). Although the CPM is a reasonable compromise between the nonparametric Kaplan-Meier estimator and structural parametric models, using the CPM with grouped data will produce asymptotic bias in point estimates and standard errors (Cox and Oakes, 1984, pp.99; Kalbfleisch and Prentice, 1980, pp.75). Modeling heterogeneity in the CPM model is computationally intensive, and the model restricts the effects of the explanatory variables to be proportional through the progression of unemployment spells. The proportional hazard assumption fails if either the effect of explanatory variables is non-proportional, or if there is unobserved heterogeneity causing the effects of explanatory variables to depend on duration.

Grouped data in survival analysis occur when failure times are observed or recorded at aggregated time intervals such as weeks or months. In my case, a labor status of an individual is aggregated at a monthly level, so the duration of the event of interest is in discrete units. This implies that there are ties in data especially for shorter spells of unemployment. The literature deals with this problem by approximating exact marginal likelihood (see Breslow, 1974; and Efron 1977Efron). The Breslow method is inaccurate in cases where there are heavy ties, so the method leads to an increasing asymptotic bias of parameter estimates (Prentice and Gloeckler, 1978; Hsieh (1995), 1995). The Efron method is more accurate than Breslow, but in the presence of the large amount of ties remains inaccurate.

In each discrete time interval, two outcomes are possible: the spell either ends or it does not. Let duration of unemployment follow a positive continuous random variable T. Each individual spell of unemployment is grouped at monthly level t_j, $j = 1...J$, i.e. I observe the number of months an individual is unemployed. Conditional on an individual remaining unemployed up to the beginning of interval t_{j-1} and the set of observable determinants of unemployment, $X(t_{j-1})$, the discrete probability that the individual's unemployment spell ends in a given time interval $[t_{j-1}, t_j)$ where $j = 1...J$, is given by

$$\lambda^d(t_j|X) = Pr\left[t_{j-1} \leq T < t_j | T \geq t_{j-1}, X(t_{j-1})\right]. \tag{3.1}$$

Hence, the probability that an individual stays unemployed beyond t_j is

$$S^d(t_j|X) = Pr[T \geq t_{j-1}|X] = \prod_{k=1}^{j-1}(1 - \lambda^d(t_k|X)). \tag{3.2}$$

The discrete-time hazard in (3.1) is the probability of leaving unemployment in $[t_{j-1}, t_j)$, and (3.1) can be restated using survivor functions, $S(.)$, as

$$\lambda^d(t_j|X) = \frac{S(t_{j-1}|X) - S(t_j|X))}{S(t_{j-1}|X)}. \tag{3.3}$$

Within each interval $S(t|X) = \exp(-\int_{t-1}^{t} \lambda(s)ds)$, and after simplifying (3.3) I obtain

$$\lambda^d(t_j|X) = 1 - \exp\left(-\int_{t-1}^{t} \lambda(s)ds\right). \tag{3.4}$$

Since with the discrete data $S(t)$ is evaluated at $t = t_j$, (3.4) can be written as

$$\lambda^d(t_j|X) = G\left(\gamma\left(t_j\right) + X\left(t_{j-1}\right)\beta\right). \tag{3.5}$$

G(.) is a distribution function that ensures $0 \leqslant \lambda_{ij} \leqslant 1$ for all i, j, while $\gamma(t_j)$ is a function of interval time that allows hazard rate to vary across periods. $\gamma(t_j)$ can be approximated with a set of dummy variables, or with a flexible functional form. The determinants of unemployment duration are either time variant or time invariant.

If each individual leaves unemployment spell during jth interval $[t_{J-1}, t_J)$, then the likelihood function will have two parts: the first part is the contribution of the individual i being unemployed t_{J-1} months, and the second part is the contribution of the individual i leaving the unemployment spell before t_J ends, so the likelihood function is

$$L = \prod_{i=1}^{N} \left[\prod_{j=1}^{J_i - 1} \left(1 - \lambda_{ij}^d(X(t_{j-1}))\right) \right] \lambda_{iJ}^d(X(t_{J-1})). \tag{3.6}$$

The likelihood equation (3.6) does not take into account the presence of either censoring or multiple spells. Estimating (3.6) consistently requires that censoring is non-informative of T_i, that it occurs at boundaries, and that spells are independent. Since the goal of this study is to examine the behavior of the unemployed, i.e. those who are available for work and actively searching for jobs, I only concentrate on the individuals from whom I can observe the transition from employment to unemployment. For left-censored spells, I cannot distinguish whether an individual is unemployed or out of the labor force, so I do not consider those spells. From an empirical point of view, the left censored spells should be disregarded in order to avoid any restrictive *a priori* assumptions about the duration dependence of the hazard rate. On the other hand, right censored spells are not an issue in estimation because I do not need to make assumptions about the duration dependence of the hazard rate. Keeping the right censored spells also implies that I am implicitly assuming that individuals with right censored spells remain in the labor force, i.e. they are continuing to search for a job.

Taking into account presence of the right censored spells requires modification of the likelihood function in (3.6). Let δ_{ij} be an indicator function such that $\delta_{ij} = 0$ if the individual left an unemployment spell during $[t_{j-1}, t_j)$, and 0 otherwise. Then (3.6) can be written as follows:

$$L = \prod_{i=1}^{N} \left[\prod_{j=1}^{J_i} \left(1 - \lambda_{ij}^d(X(t_{j-1}))\right)^{1-\delta_{ij}} \right] \lambda_{iJ}^d(X(t_{J-1}))^{\delta_{ij}}. \tag{3.7}$$

Clearly, the only difference between equations (3.6) and (3.7) is in cases of right censored spells where contribution to likelihood function is only made by the first term of likelihood. The likelihood functions in (3.6) and (3.7) are equivalent to likelihood function for a binary panel regression model with dependent variable δ_{ij}.

3.4.2 Estimation

Estimating the likelihood function (3.7) requires assumptions for the functional form of the hazard rate, λ_{ij}^{d}, and for the functional form for the baseline, $\gamma_j(t)$. The most common functional specifications of hazard function are the normal, logistic and extreme-value minimum distribution, which yield a probit, logit, and cloglog model. In my estimation, I assume logistic distribution and estimate panel-logit model[8]. Baseline, $\gamma_j(t)$, is assumed to be a cubic function unemployment duration. In addition, I include binary variables for federal states and binary variables for years to account for any state-time specific effects that might be correlated with determinants. Because of multiple spells per individual, I use the Huber Sandwich Estimator to obtain conditional independence between spells.

3.4.3 Results

Firstly, I examine the effects of determinants on the duration of unemployment between East and West Germany by estimating a model where the effects of the covariates including the grouped baseline vary across four groups: East, West, male, and female. For all pairs, I test if the joint effect of the determinants of unemployment is the same across these four groups. The results show that the effects of covariates are significantly different across groups at one percent confidence level for all pairs, which is why I perform separate regressions for East, West, females, and males[9].

Secondly, I test for each group separately if the effects of covariates on the duration of unemployment are the same across three different periods: 1990–1995 , 1996–2002, and 2003–2012. For each group, under the null hypothesis, i.e. the effects of covariates are the same across three periods, I estimate a model in which there is an interaction between binary variables for the period in question and the covariates including the baseline. I find that the effects of covariates across time periods are different or both the East and the West females at 10 percent and 1 percent, respectively. The results for males are mixed because I find evidence of determinants of unemployment being significantly different across the time periods for East males and lack of the same evidence for West males[9]. Hence, I present the results across the time periods for females and East males separately,

[8]Assuming the normal, logistic and extreme-value minimum distribution produces similar results, but these three distributions differ in the type of proportional effects covariates have on the hazard. Sueyoshi (1991) shows that the probit model generates non-proportional effects of the variables on discrete hazard while the logit and cloglog model tend to show only slight non-proportionality because derivatives of the log discrete hazard differ substantially between these distributions.

[9]These results are available upon request.

while I pool three periods for the West males.

Tables 3.5 and 3.6 summarize the effects of unemployment determinants on hazard rates in terms of odds ratios by gender for each period. In Table 3.5, columns 1–3 and columns 4–6 report the effects of the determinants on the Eastern and the Western females' hazard rates for the three periods: 1990–1995, 1996–2002, and 2003–2012. In Table 3.6, columns 1–3 summarize the effects of unemployment determinants on the Eastern males for the same periods, while column 4 reports the effects on the Western males for the period 1990–2012.

Comparing the coefficients on age groups for East females between 1990–1995 and 2003–2012 periods (columns 1 and 3 in Table 3.5) shows that East females older than 36 years entering unemployment in the period 2003–2012 have lower odds of leaving unemployment than comparable females in the period 1990–1995. In contrast, West females in all age groups have higher odds of leaving unemployment if they entered unemployment in the period 2003–2012 as opposed to 1990–1995, columns 4 and 6 in Table 3.5. A comparison between East and West females during the period 2003–2012 suggests that West females are more likely to leave unemployment if older than 36. Hence, there is a divergence in the role of age in unemployment hazards between East and West females.

In Table 3.5, for the East, I find that education was not important for the hazard rate during the 1990–1995 and 1996–2002 periods. This is somewhat expected since all the industries and sectors were affected by the initial shock from unification, while the recovery was slow during the medium run. However, in the latest period, the individuals with at least a vocational education have 2.4 times higher odds of leaving unemployment than those who did not complete their degrees, the omitted group. In contrast, education is an important determinant in the West regardless of period and the effect of education on West female hazards is roughly constant over time. Comparing the coefficients between the East and the West for the period 2003–2012 indicates that education has a higher impact on the hazard rate in the East. These results indicate that there is convergence in the role of education in unemployment duration between East and West females. In addition, the results indicate the emergence of an educational differential in the hazard rate between East and West females with education being more important for East females.

Examining the role of children's age and relationship status in Table 3.5 shows that i.) infants lower odds of leaving unemployment for both East and West females regardless of the period; ii.) East females with an infant entering unemployment during 2003–2012 have lower odds of leaving unemployment than comparable East females in 1990–1995; iii.) West females with a child age 2–6 have lower odds of leaving unemployment in the period 2003–2012 than they did in 1990–1995; iv.) children older than seven years seem to be unimportant for the duration of unemployment for both East and West females regardless of time period. On the other hand, I find that the role of relationship status in unemployment duration of East and West females is similar and does not change over time. Thus, there is evidence of divergence in the role of children, in particular the role of infants, in unemployment duration between East and West females with West females having higher odds of leaving unemployment in the presence of infants.

In Table 3.6, I find that age becomes an important determinant of unemployment duration over time. In the 1990–1995 period, only East male ages 46–55 have 0.76 times lower odds of leaving unemployment than the youngest group, 18–25. In contrast, East males that enter unemployment in the period 2003–2012 and are older than 36 have lower odds of leaving unemployment than younger males. In contrast to East males, only West males older than 46 have lower odds of leaving unemployment compared to the youngest group. Thus, these findings suggest a slight divergence in the role of the age in unemployment duration and emergence of age differential between East and West males in the age group 36–45.

Furthermore, I find that education for East males is a statistically an unimportant determinant of unemployment duration during spells that originate in the 1990–1995 period . This changes over time, and education become an important determinant for spells originating in 2003–2012. For example, East males with a vocational degree, general secondary education, and higher education have approximately 2.5, 3.3, and 2.8 folds higher odds of leaving unemployment than East males with only primary education, respectively. The educational level is also more important in East male unemployment hazards for the period 2003–2012 than for West male unemployment hazard. For example, East males with vocational or general secondary education are 1.1 and 1.7 times more likely to leave unemployment than West males. Thus, as with females, I find that there is convergence in the role of education in unemployment length, but at the same time, there is a rise in educational differential between East and West German males.

I find that the role of children have no significant role in determining unemployment duration for both East and West males. On the other hand, I find that relationship status is an important determinant of unemployment hazards. In particular, I find that married East males have 1.4 higher odds of leaving unemployment than single East males during the 1990–1995 period. Further, I find that marriage becomes even more important in the period 2003-2012. There is a similar relationship for married ad single West males, but West married males have 0.3 times lower odds of leaving unemployment]than East married males. I also find that the effect of cohabitation on the unemployment duration is similar for East males between the 1990–1995 and 2003–2012 periods, as well as that the effect of cohabitation is not significantly different than a comparable effect for West males. Thus, these findings suggest that the role of a partner in unemployment hazards of East and West males is similar, and that it was not different during the transition. Finally, I find no significant pattern in the role of the residence on unemployment hazards of East and West Germans.

3.5 Conclusion

In this paper, I investigate whether there are differences in duration of unemployment and its determinants related to age, education, children, marriage between East and West Germans, and how the role of these determinants changes over time. The German Socio-Economic Panel data allows me to distinguish between East and West Germans based on the residence before unification. I use data from 1990 until 2012 to assess developments in unemployment duration and its determinants.

The results show the existence of differences in the duration of unemployment between the Easterns and Westerns regardless of gender immediately after unification. During the early days of transition, unemployment lengths are shorter in the East than the West, especially for longer spells of unemployment, but these differences disappear for males and significantly decrease for females. I find that the female unemployment length is falling over time in both the East and West, indicating higher attachment to employment. The interesting finding about durations is that the Western female unemployment duration is approaching the Eastern female's duration. This result might come as a surprise at first, but it can be explained by changes in the unemployment benefits.

In regards to determinants of unemployment, I find that their role is changing over time for East German and West German females. Further, I find initial differences in the role of age and education in unemployment duration for both males and females. For both males and females in East, I find that the role of education in the hazard function experiences the largest change over time, from being unimportant determinant to being the most important followed by the role of age. Furthermore, I find the emergence of the educational differentials between East and West Germans with education having stronger effects on unemployment duration for East Germans. I do not find any significant differences in the roles of children and the relationship with a partner in determining the length of unemployment between East and West, regardless of gender. Therefore, the findings in this paper suggest that there is convergence in the duration of unemployment and its determinants between East and West Germans. Furthermore, this research can be extended to examine if East Germans are more likely to switch occupations when leaving unemployment spells.

Table 3.1: Summary of Unemployment Spell Types

	East		West	
	Male	Female	Male	Female
Single spell				
1990–1995	310	431	320	512
1996–2002	337	268	414	667
2003–2012	377	454	662	1016
Multiple spell				
1990–1995 and 1996–2002	111	176	72	211
1996–2002 and 2003–2012	197	236	187	415
1990–1995 and 2003–2012	19	42	24	58
All three periods	68	90	30	109
Total	1419	1697	1709	2988

Note: The first three rows summarize the number of individuals who only had one spell of unemployment between 1990–2012. The second three rows summarize the number of individuals who had multiple spells across periods.

Table 3.2: Summary of Unemployment Durations by Percentiles for Easterners and Westerners

	East			West			East	West
	1990-1995	1996-2002	2003-2012	1990-1995	1996-2002	2003-2012	1990-2012	1990-2012
Female								
25th percentile duration	6	5	4	5	4	3	5	4
Median duration	15	13	10	16	11	9	13	11
75th percentile duration	33	28	20	42	31	24	27	30
Spells [exits]	1067 [942]	1152[1007]	1353[1079]	1231[1016]	1998 [1725]	2569[1972]	3572[3020]	5798[4713]
Individuals	739	770	822	890	1402	1598	1697	2988
Male								
25th percentile duration	3	3	3	3	3	3	4	3
Median duration	6	6	6	7	6	6	6	6
75th percentile duration	14	14	13	20	14	13	14	15
Spells [exits]	777[700]	1180[1031]	1219[981]	641[527]	1014[875]	1446[1122]	3176[2712]	3101[2524]
Individuals	508	713	661	446	703	903	1419	1709

Note:Durations are in months.

Table 3.3: Summary of Determinants of Unemployment Duration, Females

	East			West			East	West
	1990-1995	1996-2002	2003-2012	1990-1995	1996-2002	2003-2012	1990-2012	1990-2012
Age								
≤25	0.235	0.157	0.187	0.220	0.145	0.099	0.195	0.155
26–35	0.331	0.389	0.388	0.467	0.443	0.368	0.366	0.428
36–045	0.241	0.237	0.219	0.161	0.228	0.321	0.234	0.235
46–55	0.193	0.218	0.206	0.152	0.184	0.211	0.205	0.182
Education								
Incomplete degree	0.003	0.012	0.021	0.025	0.031	0.010	0.011	0.022
Apprenticeship	0.292	0.207	0.135	0.491	0.381	0.323	0.222	0.398
Vocational edu.	0.538	0.591	0.541	0.331	0.349	0.375	0.557	0.351
General Secondary edu.	0.023	0.048	0.095	0.087	0.124	0.129	0.050	0.114
University	0.144	0.142	0.208	0.066	0.115	0.163	0.160	0.114
Relationship Status								
Single	0.146	0.219	0.262	0.095	0.132	0.179	0.202	0.135
Married	0.722	0.566	0.471	0.836	0.756	0.709	0.602	0.767
Cohabitation	0.132	0.215	0.267	0.069	0.112	0.112	0.196	0.098
Child's age								
0–1	0.148	0.198	0.321	0.271	0.277	0.294	0.210	0.280
2–6	0.283	0.236	0.270	0.381	0.375	0.309	0.263	0.357
7–11	0.289	0.207	0.145	0.218	0.225	0.193	0.223	0.213
Distance to city								
0km	0.121	0.102	0.133	0.087	0.101	0.096	0.117	0.095
0–10km	0.210	0.182	0.215	0.250	0.244	0.260	0.202	0.251
10–25km	0.265	0.255	0.225	0.315	0.296	0.304	0.251	0.304
25–40km	0.170	0.188	0.168	0.133	0.166	0.144	0.176	0.149
40–60km	0.166	0.165	0.119	0.105	0.086	0.099	0.154	0.096
≥60km	0.068	0.108	0.140	0.110	0.107	0.096	0.101	0.105
Living in west	0.061	0.134	0.170	0.995	0.991	0.982	0.115	0.989
Observations	25389	22977	16984	36787	42793	35413	65350	114993
Spells [exits]	1067 [942]	1152[1007]	1353[1079]	1231[1016]	1998 [1725]	2569[1972]	3572[3020]	5798[4713]
Individuals	739	770	822	890	1402	1598	1697	2988

Table 3.4: Summary of Determinants of Unemployment Duration, Males

	East			West			East	West
	1990-1995	1996-2002	2003-2012	1990-1995	1996-2002	2003-2012	1990-2012	1990-2012
Age								
$\leq$ 25	0.229	0.196	0.242	0.403	0.248	0.233	0.220	0.279
26–35	0.306	0.254	0.233	0.315	0.311	0.247	0.259	0.286
36–45	0.230	0.283	0.206	0.138	0.214	0.285	0.243	0.225
46–55	0.235	0.267	0.319	0.144	0.227	0.235	0.278	0.211
Education								
Incomplete degree	0.026	0.017	0.023	0.070	0.064	0.044	0.021	0.057
Apprenticeship	0.291	0.326	0.374	0.600	0.572	0.543	0.335	0.567
Vocational edu.	0.474	0.516	0.496	0.164	0.178	0.215	0.499	0.190
General Secondary edu.	0.046	0.044	0.048	0.115	0.089	0.084	0.046	0.093
University	0.163	0.097	0.058	0.052	0.097	0.114	0.099	0.093
Relationship Status								
Single	0.318	0.386	0.430	0.505	0.472	0.432	0.385	0.463
Married	0.572	0.402	0.350	0.357	0.390	0.403	0.424	0.388
Cohabitation	0.109	0.212	0.220	0.138	0.138	0.165	0.191	0.149
Child's age								
0–1	0.056	0.036	0.055	0.065	0.055	0.059	0.047	0.059
2–6	0.172	0.074	0.102	0.136	0.103	0.109	0.107	0.113
7–11	0.190	0.114	0.066	0.113	0.106	0.102	0.115	0.106
Distance to city								
0km	0.183	0.112	0.086	0.151	0.140	0.107	0.120	0.129
0–10km	0.140	0.218	0.176	0.298	0.251	0.265	0.185	0.268
10–25km	0.182	0.216	0.183	0.273	0.276	0.260	0.196	0.269
25–40km	0.212	0.178	0.213	0.093	0.112	0.141	0.198	0.119
40–60km	0.205	0.148	0.154	0.103	0.102	0.103	0.164	0.103
$\geq$ 60km	0.077	0.129	0.188	0.082	0.119	0.123	0.137	0.112
Living in west	0.033	0.049	0.083	0.982	0.975	0.967	0.057	0.973
Observations	8133	14307	12076	7448	11269	12758	37692	34576
Spells [exits]	777[700]	1180[1031]	1219[981]	641[527]	1014[875]	1446[1122]	3176[2712]	3101[2524]
Individuals	508	713	661	446	703	903	1419	1709

Table 3.5: Estimates of the Effect of Unemployment Determinants on Female Unemployment Hazard

	East			West		
	1990–1995	1996–2002	2003–2012	1990–1995	1996–2002	2003–2012
Age						
26–35	0.87	0.67*	1.00	0.87	1.01	0.79*
	(0.71)	(0.20)	(47.71)	(0.55)	(8.35)	(0.30)
36–45	0.80***	0.56*	0.58*	0.64*	0.94	0.75*
	(0.48)	(0.13)	(0.15)	(0.18)	(1.46)	(0.24)
46–55	0.60*	0.38*	0.54*	0.51*	0.52*	0.63*
	(0.20)	(0.07)	(0.15)	(0.12)	(0.10)	(0.16)
Education						
Apprenticeship	0.82	1.27	1.46	1.34	2.10*	1.44
	(0.67)	(2.69)	(1.11)	(1.38)	(0.48)	(0.96)
Vocational Education	1.03	1.51	2.41*	1.66***	2.40*	1.84**
	(5.64)	(1.81)	(0.74)	(1.00)	(0.49)	(0.76)
General Secondary Edu.	1.00	2.14	2.83*	1.90**	2.52*	1.92**
	(48.30)	(1.41)	(0.78)	(0.96)	(0.51)	(0.75)
Higher Education	1.64**	2.14	2.96*	2.71**	3.13*	2.54*
	(0.66)	(1.41)	(0.77)	(1.24)	(0.53)	(0.68)
Children's Age						
0–1	0.30*	0.27*	0.22*	0.30*	0.36*	0.30*
	(0.03)	(0.03)	(0.02)	(0.04)	(0.03)	(0.02)
2–6	0.87	0.82**	0.89	1.04	0.96	0.79*
	(0.63)	(0.49)	(0.72)	(4.77)	(1.63)	(0.25)
7–11	0.85*	0.81*	0.89	1.28*	1.05	1.04
	(0.49)	(0.43)	(0.94)	(0.49)	(1.93)	(2.07)
Partner's Status						
Married	1.21***	1.06	1.06	0.99	0.87***	1.01
	(0.69)	(2.22)	(2.01)	(8.22)	(0.48)	(10.64)
Cohabitation	0.98	1.00	1.12	1.12	0.92	1.31*
	(5.24)	(735.29)	(1.19)	(1.24)	(1.10)	(0.46)
Distance to City						
$\leq$ 10km	1.03	0.93	1.08	1.16	0.84	0.99
	(3.52)	(1.96)	(2.08)	(1.06)	(0.52)	(11.47)
10–25km	0.98	0.89	1.06	1.07	0.90	0.85
	(5.20)	(1.06)	(2.93)	(2.17)	(0.95)	(0.60)
25–40km	1.11	0.79	1.10	0.92	0.83	0.94
	(1.48)	(0.52)	(1.98)	(1.72)	(0.55)	(2.10)
40–60km	1.05	0.70**	1.18	0.98	0.87	1.07
	(2.89)	(0.33)	(1.23)	(7.07)	(0.84)	(2.35)
$\geq$60km	1.30***	0.79	0.93	0.85	0.90	1.09
	(0.73)	(0.55)	(2.07)	(0.91)	(1.24)	(1.63)
ln(L)	−3659.6429	−3890.5203	−3829.0465	−4183.9119	−6646.7389	−7161.1392
Observations	24,736	22,954	16,976	35,569	42,615	35,396
Individuals	739	770	822	889	1,401	1,598

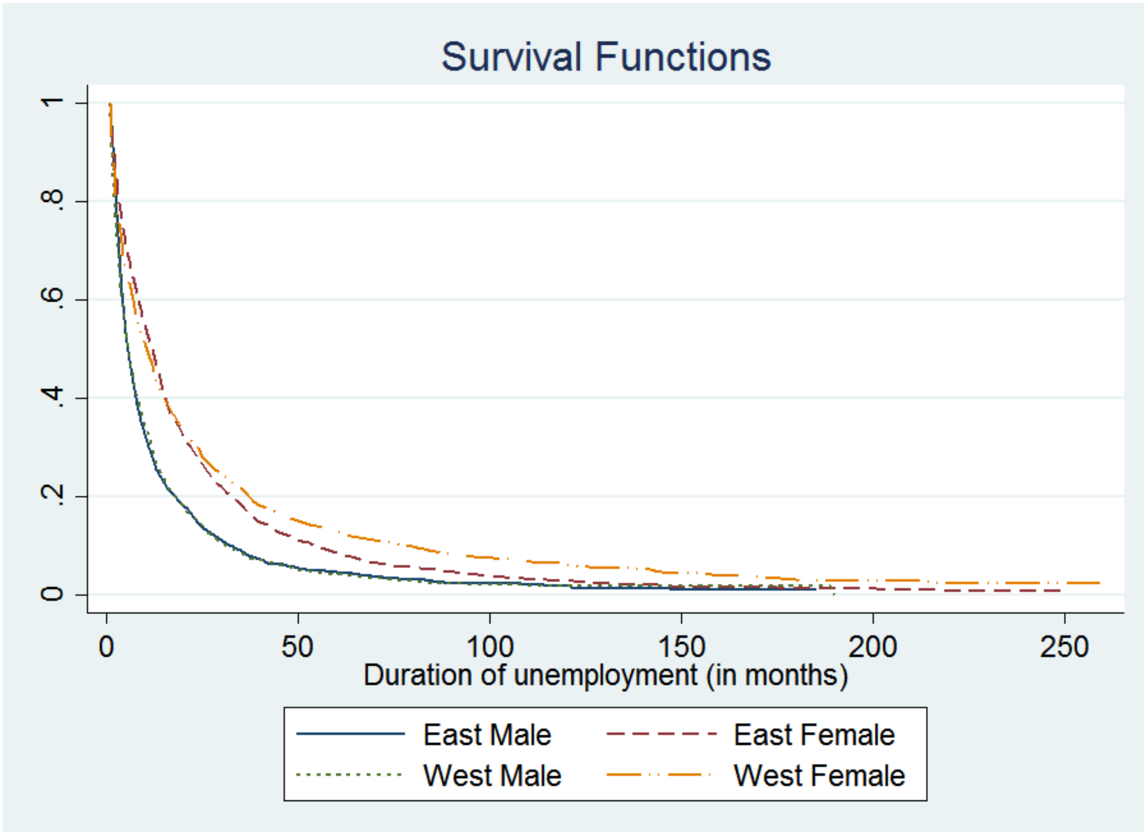
Note: Estimates are obtained using the panel logit estimator. Odds ratios are reported with standard errors clustered at individual levels due to multiple spells per individual. All regressions include year and state dummies. Determinants are time-invariant within spells, except for childrens' variables, which are time-variant. Significance levels: 1% *, 5% ** ,and 10% ***.

Table 3.6: Estimates of the Effect of Unemployment Determinants on Male Unemployment Hazard

	East			West
	1990–1995	1996–2002	2003–2012	1990–2012
Age	0.90	0.64*	1.00	1.03
26–35	(1.19)	(0.15)	(75.68)	(2.20)
	0.96	0.57*	0.77**	0.90
36–45	(3.41)	(0.11)	(0.38)	(0.70)
	0.73**	0.43*	0.56*	0.51*
46–55	(0.42)	(0.06)	(0.13)	(0.07)
Education				
Apprenticeship	1.07	1.64***	1.41	1.34**
	(3.59)	(0.86)	(1.56)	(0.64)
Vocational Education	1.07	1.98*	2.47**	1.37**
	(3.56)	(0.75)	(1.02)	(0.61)
General Secondary Edu.	0.78	2.02**	3.31*	1.61***
	(1.13)	(0.82)	(1.06)	(0.51)
Higher Education	0.91	2.20*	2.81*	2.15***
	(2.36)	(0.77)	(1.08)	(0.44)
Children's Age				
0–1	1.03	0.86	0.90	1.05
	(5.19)	(0.97)	(1.35)	(2.24)
2–6	1.03	1.04	0.82	0.78***
	(4.21)	(2.98)	(0.54)	(0.28)
7–11	0.99	0.95	0.86	0.92
	(8.81)	(1.94)	(0.92)	(1.06)
Partner's Status				
Married	1.43*	1.66*	1.62*	1.40*
	(0.55)	(0.34)	(0.38)	(0.31)
Cohabitation	1.35	1.22	1.22	1.30*
	(0.73)	(0.61)	(0.70)	(0.38)
Distance to City				
≤10km	1.45**	0.91	1.32	0.99
	(0.64)	(1.34)	(0.85)	(6.96)
10–25km	1.59*	1.09	1.61**	1.20**
	(0.53)	(1.66)	(0.62)	(0.57)
25–40km	1.19	1.23	1.43**	1.20*
	(1.03)	(0.81)	(0.71)	(0.63)
40–60km	1.29	1.01	1.22	1.16
	(0.83)	(11.33)	(1.20)	(0.93)
≥60km	1.34	1.02	1.12	1.20
	(0.88)	(7.60)	(1.82)	(0.74)
ln(L)	-2300.24	-3447.92	-3183.25	-8344.79
Observations	8,075	14,084	12,064	31,433
Individuals	508	712	661	1,709

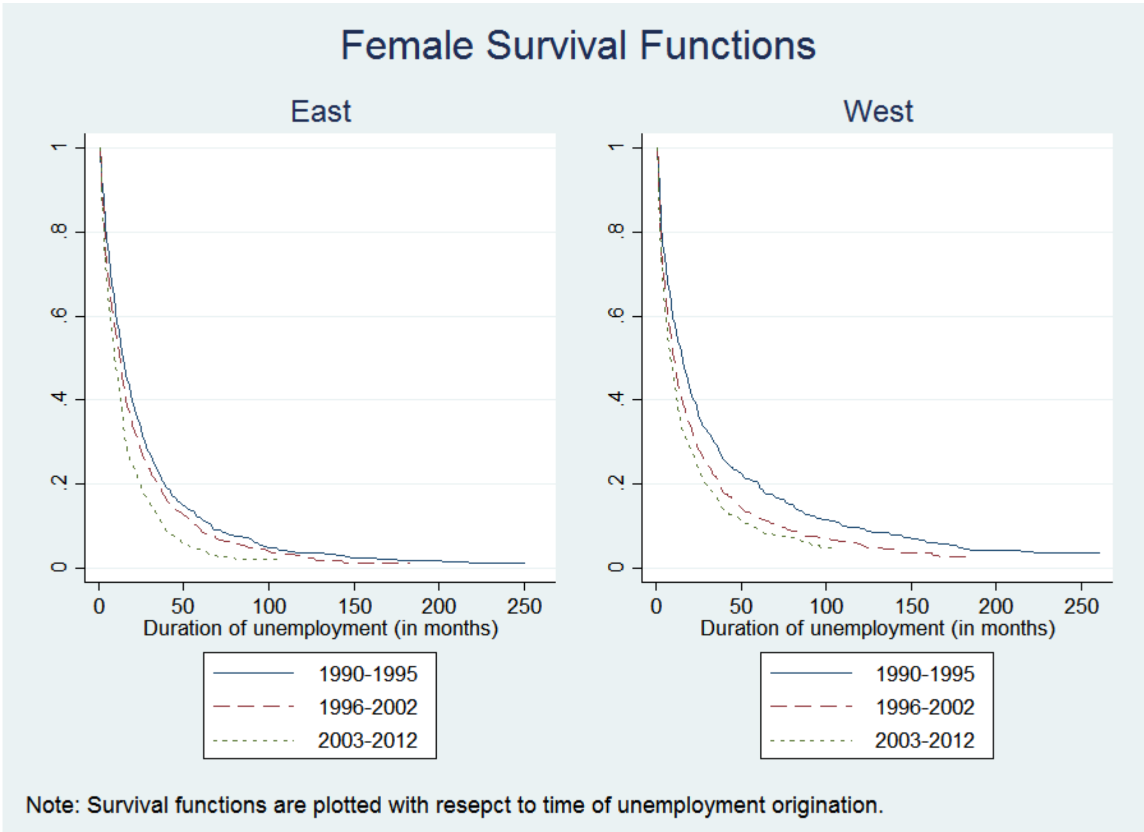
Note: Estimates are obtained using the panel logit estimator. Odds ratios are reported with standard errors clustered at individual levels due to multiple spells per individual. All regressions include year and state dummies. Determinants are time-invariant within spells, except for childrens' variables, which are time-variant. Significance levels: 1% *, 5% ** ,and 10% ***.

Figure 3.1: Survival functions, 1990-2010



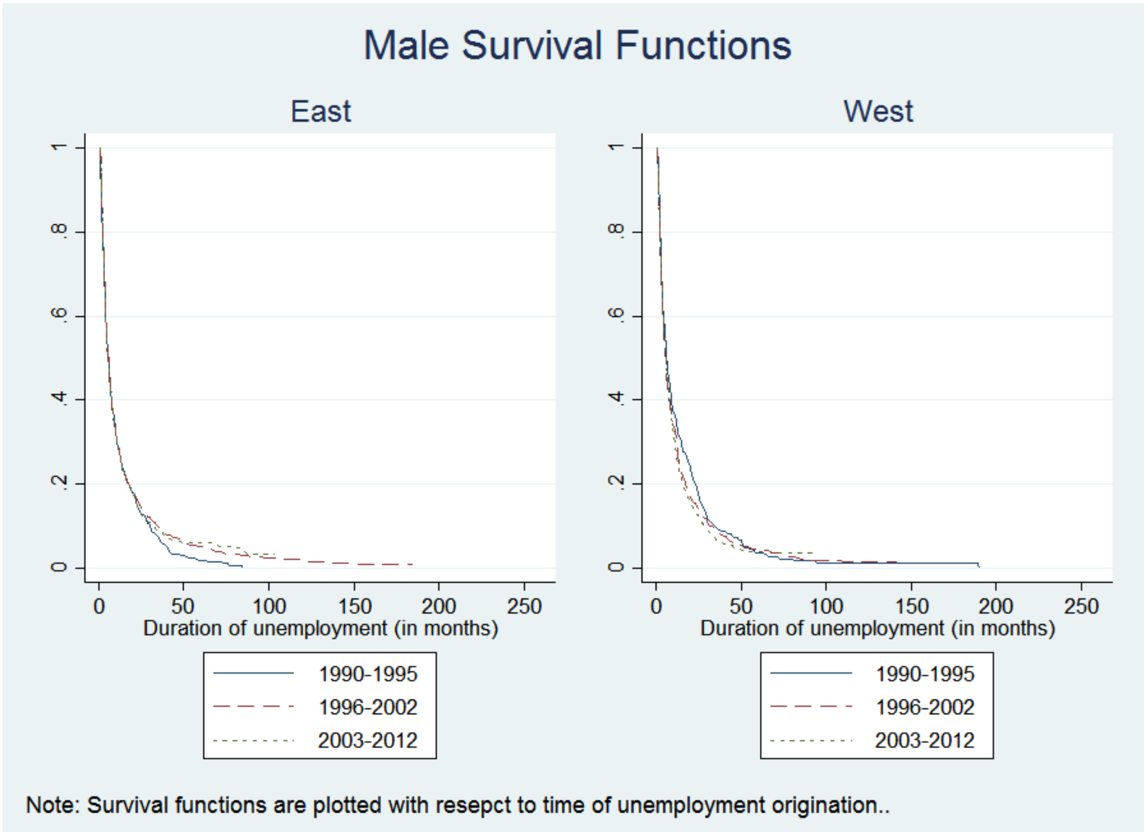
Note: The survivor functions are calculated using the Kaplan–Meier estimator. The Log-rank and the Wilcoxon test of equality between East and West female survivor functions yields p=0.05 while I find no significant differences between East and West male survivor functions.

Figure 3.2: Female unemployment survival functions by location



Note: The survivor functions are calculated using the Kaplan–Meier estimator. Testing equality of survivor functions within East and West over time shows that they are significantly different at 1 percent level. Comparing East to West female survivor functions for each period shows that largest difference was during the 1990–1995 period when the Log-rank and the Tarone-Ware tests reject the null hypothesis at 1 percent confidence level. The differences between East and West female survivor function fall over time while renaming significantly different.

Figure 3.3: Male unemployment survival functions by location



Note: The survivor functions are calculated using the Kaplan–Meier estimator. The Tarone-Ware test fails to reject the hypothesis that these survivor functions are changing over time within East and West. The Log-rank test and the Tarone-Ware suggest that there are significant differences between East and West male survivor functions at 7 percent confidence level for spells originating during 1990–1995, and the tests fail to reject the null hypothesis for other periods.

Appendices

Table 7: Robustness: Estimation Results without Exclusion Restriction

Variable	Coeff.	95% CI	
Homeowner	.275	$-.048$	1.034
Unobserved Effects			
ν_u	3.054	-1.804	4.971
ν_h	1.765	.627	1.977
$Corr(\nu_u, \nu_h)$	-.001	$-.0454$	.078
Support Points			
π_1	.048	.003	.04
π_2	0.79	.002	.125
π_3	.289	.274	.723
π_4	.583	.473	.75
Log		3251.2	
Individuals		1898	
Observations (person-spell)		3077	

Note: Estimation is performed using simulated annealing, global optimization technique, in R with package GenSA. I reported 95% CI are biased-corrected intervals that are calculated as described in Efron and Tibshirani (1986) because standard inference that relays on asymptotic properties is not appropriate due to skewness in empirical distribution of the estimate. The 95% CI for the correlation coefficient is calculated based on 1000 draws from the multivariate normal distribution with mean and covariance matrix set to the estimated parameter vector and covariance matrix.

Appendix A Model without exclusion restrictions

Alternative estimation methodology to exclusion variable, primary methodology in this paper is to exploit availability of multiple unemployment spells per person in which there is intra-person variation in homeownership status. Munch et al. (2006) use this methodology in their paper. Unfortunately, I only observe four percent of individuals who were homeowners and renters. Given the small variation in the intra-person homeownership status, the identification strategy with exclusion restriction is preferred, but I also estimate the model using intra-person variation of homeownership status as an identification strategy for the completeness. In Table 7, I present the estimates of the model that corrects for the presence of self-selection into homeownership by using intra-person variation in homeownership status. The coefficient on homeownership status indicates that homeowners have 31 percent (=exp(0.275)$-$1) higher rates of exit from unemployment than renters but the 95% biased-corrected confidence interval is between $-.048$ and 1.304 . The difference between the coefficients on homeownership in the model without correction and with correction for self-selection indicates the presence of negative correlation between unobserved effects such that unobserved characteristics that make men more likely to be homeowners also have negative impacts on the transition out of unemployment. The estimate of correlation between the unobserved effects is -0.001 and this is statistically significant. Although the intra-person variation of homeownership status is not a viable identification strategy in this situation, the estimates of correlation between unobserved effects indicate the need to correct for self-selection when estimating the effects of homeownership on exits out of unemployment.

Table 8: Main Estimation Results with Exclusion Restriction: Full Results

Variable	Unemployment Hazard			Homeownership Choice		
	Coefficient	95% CI		Coefficient	95% CI	
Homeowner	−.832	−2.460	−.464	–	–	–
Homeownership Rate	–	–	–	1.923	−.096	2.7
Age	.038	−.034	.06	−.199	−.251	−.145
Age^2	−.001	−.001	0	.003	.002	.004
Education						
Apprenticeship	.130	−.165	.639	.349	−.296	1.303
High school	−.011	−.326	1.02	.700	.052	1.788
College or Higher	.065	−.315	.726	.551	−.123	1.541
Partner						
Employed Partner	.250	−.29	.701	.321	−.379	.751
Unemployed Partner	.250	−.453	.547	−.321	−.904	.044
Non−labor Income						
Public Transfer	−.061	−.119	−.019	.016	−.041	.067
Partner's income	.035	−.005	.087	−.023	−.059	.025
Labor Market Tightness	−3.774	−4.95	−1.462	−1.389	−2.55	.894
Unobserved Effects						
ν_u	2.775	.331	4.35			
ν_h	1.960	.457	4.5			
$Corr(\nu_u, \nu_h)$	.069	.007	.13			
Support Points						
π_1	.300	.133	.718			
π_2	.115	.002	.302			
π_3	.056	.001	.295			
π_4	.529	.329	.789			
Log			3258.145			
Individuals			1898			
Observations			3077			

Note: Estimation is performed using simulated annealing, global optimization technique, in R with package GenSA. I reported 95% CI are biased-corrected intervals as described in Efron and Tibshirani (1986) because standard inferences that rely on asymptotic properties is not appropriate due to skewness in the empirical distribution of the estimators. The standard error for the correlation coefficient is calculated based on 1000 draws from the multivariate normal distribution with mean and covariance matrix set to the estimated parameter vector and covariance matrix.

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