Time for children and career in Germany

Time constraints, childcare and career paths

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Name: Katharina Sutter

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Zweitgutachter:	Prof. Dr. Erwin Amann

Preface

I wrote this dissertation while working as a research assistant at the Chair of Public Economics (Prof. Dr. Reinhold Schnabel) at the University of Duisburg-Essen. Writing this thesis has been the most challenging step in my academic and occupational development. During these years, I not only had the chance to further develop my skills but also to develop in a personal capacity. My path has led me to meet people whom I greatly esteem. I thank Udo Schneider for supervising my diploma thesis, since this was my motivation for writing a dissertation. The chair of Reinhold Schnabel was the perfect place to start this project.

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Part A

A Introduction

A.1 Motivation

The compatibility of family and career in Germany is a prevalent issue in political discourse as well as in economic literature (e.g., Spieß (2011), Gerlach (2008)).
One reason for this is the demographic change the country is undergoing. Germany is facing one of the lowest fertility rates in Europe (e.g., European Commission (2014)). The total fertility rate in Germany was 1.38 children per woman in 2012, compared to 1.58 in the entire European Union and 2.01 in France (European Commission (2014)). Highly educated women particularly show lower fertility compared to less educated women (Federal Statistical Office of Germany (2012b)). In addition, life expectancy is increasing (Federal Statistical Office of Germany (2009)). These two important factors largely influence the demographic structure of a country (European Commission (2012)). For instance, the old-age dependency ratio of Germany is expected to increase from 34 in 2008 to 63 resp. 67 (depending on migration) by 2060 (Federal Statistical Office of Germany (2009)). This ratio is defined as the number of people aged 65 and over divided by the number of people aged 15 to 64 ((Federal Statistical Office of Germany (2009)). This will bear remarkable consequences for the social security system, on both the revenue and expenditure sides, as well as for economic growth (for a deeper discussion, see, for instance OECD (2014), European Commission (2012)). Therefore, the utilization of present labor force potentials is considered to be important for economic growth (e.g., OECD (2014)). Women and older employees are viewed as important labor reserves (apart from migration, e.g., Reinberg and Hummel (2003), OECD (2014)).

In recent decades, the proportion of women among university graduates has been increasing (CEWS (2015)) and the share of women participating in the labor market also shows an upward trend (European Commission (2014)). Even the proportion of women in managerial positions is increasing, although it is still at a low level (Holst and Schimenta (2013)). Notwithstanding these trends, home-related work, such as housework and childcare, is still within the female domain of responsibility (e.g., Federal Statistical Office of Germany (2003)). In addition, a large share of German mothers only work part-time, indicating that the male breadwinner model still is prevalent (e.g., European Commission (2014)). While only about 9% of employed men work part-time, about 45% of employed women do so (e.g., European Commission (2014)). Part-time work often is considered to be incompatible with managerial positions (Holst et al. (2012)). There is also a remarkable difference between mothers and childless women. Childless women are more often employed in full-time positions than are mothers, and family-related and personal obligations are most often cited as reasons for their part-time employment (e.g., Keller and Haustein (2013), Federal Statistical Office of Germany (2012a)).

The compatibility of a working career and family life is a vital component for tackling the aforementioned challenges (for a detailed discussion, see OECD (2007, 2014)). The government as well as employers try to facilitate the combinability of family and working life for mothers, for example by offering (institutional) childcare (e.g., Spieß (2011), §24 SGB VIII, Gerlach (2010)). The implementation of the *Elterngeld* in 2007 as well as the right to daycare for children aged at least one year from 2013 onwards are further examples of policy instruments designed to strengthen maternal labor market attachment (KiFöG, German Bundestag (2006, 2008)). But still, while the proportion of women participating has increased, a large number of mothers that participate only works part-time (European Commission (2014)). Household chores and childcare are said to be the most prevalent reasons for part-time employment (Keller and Haustein (2013), Federal Statistical Office of Germany (2012a)).

This background informs the motivation of this thesis and the issues investigated herein. The time used for home-related and market work seems to be one reason for the prevalent part-time employment of mothers (Keller and Haustein (2013), Federal Statistical Office of Germany (2012a)). As Wang and Bianchi (2009) state, male involvement in home-related tasks might be crucial to labor market outcomes of mothers. Hence, in Chapter B, we investigate how the time allocation within couple households changes if one partner faces an employment shock. The question in focus refers specifically to the change in male time devoted to home-related tasks and the female time devoted to market work. Since the time allocation of both men and women in couple households is strongly interdependent, we make use of a difference-in-differences approach using an exogenous employment shock. The time allocation within households appears to be rather rigid in the short term.

The low representation of women in managerial positions leads to the question Chapter C focuses on. I investigate whether the duration of birth-related time outs from the labor market indicates a career penalty in terms of occupational prestige. The analysis takes account of the selection into employment by applying a Heckman selection correction. It investigates not only the prestige level per se but also the occupational mobility compared to the prestige level held before the birth of the first child. Mothers' careers are destabilized after very long time outs.

Finally, in Chapter D, the role of informal childcare for maternal labor market involvement is comprehensively analyzed. Again, time constraints might play an important role in the decision about labor market participation and working hours. Informal childcare arrangements are expected to relax these constraints to some extent and thereby to support maternal labor market participation as well as their working hours. This includes two decisions. The analysis therefore faces two major econometric challenges. The first is the potential endogeneity of informal childcare arrangements, and the second is the selection into employment when investigating work hours. Both are taken into account.

A.2 Overview and summary

I begin with the question of rigidity of intra-family allocation of home-related work. Therefore, section B focuses on the intra-household allocation of time among couples. Using the German Socio-Economic Panel, we investigate how the time allocation changes if one of the partners is forced out of market work by an involuntary lay-off. We expect that he or she will spend some of the additional time on household chores and/or childcare. Thus, our results will give some indication as to whether couples allocate their time due to time constraints or due to individual preferences. We apply a difference-in-differences approach as well as panel methods to identify the effects of negative employment shocks on childcare and housework of both partners. Our results indicate that there is only a moderate reaction with respect to domestic work, namely childcare and housework. This evidence suggests that preferences determine the division of home-related work within families and that the intra-household time allocation is relatively rigid in the short term. Public policies that loosen the time constraints (e.g., public provision of childcare) will not have large effects on the intra-family division of household work.

In section C, I analyze whether there is a penalty for birth-related leaves of absence in terms of a loss in occupational prestige upon return to the labor market. There is a broad body of research on the effect of motherhood on wages. Beyond wages, occupational prestige is also an important characteristic of a job. In this section, I ask whether there is a penalty for this as well. I use the SIOPS information of the German Socio-Economic Panel (GSOEP) to investigate this for the case of Germany. Since the SIOPS information is only observed for those who are working, a two-step model according to Heckman (1979) is used to correct for this selection. In addition, a strategy suggested by Wooldridge (1995) is used to account for the panel structure of the GSOEP. The first step estimates the probability of participation in the labor market after the first birth, i.e., of observing prestige information. The second step is performed for the level of the occupational prestige as well as for the probability of an upward or downward occupational move, conditional on selection. The descriptive analysis offers a preliminary indication that occupational mobility is higher for mothers as compared to childless women, and that it is higher if the career interruption is long. This is true for upward as well as for downward mobility. The results reveal a prestige penalty for very long career interruptions.

Section D analyzes the relationship between the use of informal childcare arrangements and the employment of mothers. It focuses on informal childcare outside the household provided by relatives, friends or neighbors. The aim is to investigate whether the presence of this kind of childcare arrangements facilitates maternal labor market involvement. I suspect that a crucial property of this kind of childcare might be its flexibility in contrast to formal childcare arrangements (also stated by Posadas and Vidal-Fernández (2012)). To analyze this question, the survey years 1999 through 2012 of the German Socio-Economic Panel (GSOEP) are used. The presence of informal childcare arrangements is most likely not exogenous to employment status. To take this into account, an instrumental variable approach is applied. The instrument for informal childcare is the information about whether the grandmother is still alive. This is considered to have a direct effect on the probability of having an informal childcare arrangement, but it is not considered to influence the labor market participation of mothers directly. Using two-stage least squares estimation, the findings reveal that women who receive the support through informal childcare are more likely to participate in the labor market. Since working hours are only observed for those who are participating in the work force, I apply a selection correction to analyze the employment hours. Informal childcare arrangements also have a positive impact on the working hours, even though the magnitude is small.

The last section summarizes the results and draws a conclusion.

Part B

B Intra-household time allocation: The effects of an employment shock

Reinhold Schnabel Katharina Sutter

B.1 Introduction

Time use in couple households remains unevenly distributed between men and women. This is not only true with regard to labor force participation at the intensive and extensive margins but also with regard to other activities such as childcare and housework. For instance, in households with children below age 17, men only spend 1.5 hours a weekday on childcare whereas women spend nearly 6 hours. Even when controlling for labor force participation, enormous differences remain between men and women. Understanding the reasons for this pattern of time allocation is interesting in itself. Moreover, the effects of economic policies intended to improve work-life balance may heavily depend on the intra-household mechanisms. More specifically, the effectiveness of public childcare programs that are intended to weaken the time constraints of families may depend on the role of those constraints on the one side and on the role of preferences on the other side. If the uneven distribution of market and non-market work within families is a result of individual preferences, the effect of loosening time constraints might be limited.

However, causal inference on the time allocation within families is difficult due to the potential simultaneity of the decisions, both between partners and between different time uses. In order to tackle this problem, we investigate how men and women respond to an exogenous variation in market work time. More specifically, we ask whether and to what extent men and women living as couples alter their time allocation in response to an involuntary reduction in working hours. There are many dimensions along which time uses could be adjusted. First, the person with “excess” time could increase his or her non-market work (childcare, housework, repairs, etc.). Second, the other spouse may also alter his or her market and non-market work. The reaction to such a shock is compelling to study because it may indicate whether the time allocation prior to the shock displays couples' preferences or whether it is mainly a result of time constraints.

We contribute to the literature on time use by employing a time variation due to a plant closure or a dismissal in order to identify responses in time use of both partners. We prefer using plant closures since those should be exogenous with respect to intra-household time allocation. We also estimate models using dismissals. Dismissals, however, may not be fully exogenous with respect to other time uses if the worker has provoked a dismissal, e.g., in order to have more time for childcare. Using an exogenous variation is important in order to identify the determinants of own non-market activities and of the other partner's time uses. We use a difference-in-differences approach as our main estimation strategy. Thereby, we identify the causal effect of a variation of one partner's time use (namely being laid off) on his or her own non-market activities as well as on the partner's decisions. The resulting estimates shed some light on the motives of time allocation within partnerships. Our treatment group consists of persons subject to a plant closure (or a dismissal); the control group consists of those persons who are not hit by such a shock. Thus, we do not have to rely on questionable instruments like some part of the literature.

Our analysis is based on the German Socio-Economic Panel (GSOEP) using the survey years 1992 to 2010. We can follow persons over time from employment to dismissal and to the adjustment of home-related work. The estimations are performed separately for men and women in order to identify gender differences in market and non-market time use. In addition, we run pooled estimations.

We find that gender differences in time use seem to be very stable. The increase of housework and childcare of men, although (weakly) significant, falls short of closing the enormous gender gap in time use. This hints to stable gender roles that are barely altered in response to employment shocks and suggests that preferences rather than time constraints play an important role in determining time allocation within families.

In the remainder of this chapter, we first turn to a discussion of the existing literature. In the third section, we describe the dataset as well as the construction of our samples. Subsequently, we present some descriptive statistics. Section five contains an outline of the econometric methods and the main results. Section six concludes.

B.2 Literature Review

Many studies investigate couples' time allocation (e.g., Álvarez and Miles (2006), Connelly and Kimmel (2009), Deding and Lausten (2006), Duguet and Simonnet (2007), García et al. (2009), Mancini and Pasqua (2011), Solaz (2005), Hallberg and Klevmarken (2003), Bloemen et al. (2010)). The studies focus on various questions, such as the influence of wages on time allocation (e.g., Kalenkoski et al. (2009), Bloemen and Stancanelli (2008)), the influence of children on parental leisure time synchronization (e.g., Barnet-Verzat et al. (2011)) or spousal influences on time allocation (e.g., Connelly and Kimmel (2009)). We will focus on the effect of a variation in one time use on the other time uses, both of oneself and of the spouse.

Nock and Kingston (1988) explore the interdependencies between working time and time with children. They distinguish between different activities, e.g., playing/educating and housekeeping as well as between single- and dual-earner households. They conclude that working hours influence the time each individual spends with children, but that the extent of this influence seems larger in the case of activities where childcare is a secondary activity. The time men spend with children is more responsive to work hours than it is for women. Álvarez and Miles (2006) focus on the relationship between female market work and male housework in Spain when the husband has a paid job. They find that female labor force participation increases both the male share of housework and the number of hours men spend on housework. The effect seems to be stronger if they consider the wife's participation to be endogenous. Anxo and Carlin (2004) study French time diary data and find that higher female working hours lower absolute male housework time but raise the male *share*. The results of Daunfeldt and Hellström (2007) indicate that more working hours result in a lower probability to do domestic work. Leeds and von Allmen (2004), using the year 1991 of the Panel Study of Income Dynamics, find a clear difference in the behavior of working and non-working women in couples. The latter do not view their partners as substitutes for home production, whereas working women do. Husbands' reaction to the behavior of their spouses depends on the presence of children.

Connelly and Kimmel (2009) examine spousal influences on non-market time uses in married couples with children. These are the wife's relative wage, which is often used as a proxy for bargaining power (e.g., Connelly and Kimmel (2009) but also Anxo and Carlin (2004), García et al. (2009)), spousal working time and spousal non-market time. They use the American Time Use Survey (ATUS, 2003-2006) and consider leisure, childcare and home production as non-market time. Due to possible endogeneity, wages, working time and spousal non-market time are instrumented. They find small effects. An increased market work time of the husband seems to result in less home production time for the wife. On the other hand, if the wife devotes more time to market work, the husband devotes a greater amount of time to care giving. They find some indication that leisure time is complementary for parents.

A study by Hallberg and Klevmarken (2003) analyzes economic incentives' effect on parental childcare time and also includes spousal time use variables. They use a Swedish panel dataset (HUS) of the years 1984 and 1993 and focus on couples with children. Their study faces the issue of the need to instrument time use variables. Their results indicate that parental time with children is mainly influenced by a change in working hours of the male parent rather than the female parent.

One of the questions considered by Hamermesh (2002) concerns how an extra hour is used (this hour occurs every autumn when people change to standard time). If people (especially mothers with young children) get an "extra hour", they mostly spend it on sleep.

Sousa-Poza et al. (2001) focus on housework and childcare in Switzerland. They investigate the determinants of those two time uses as well as their monetary value by using the 1997 Swiss Labour Force Survey. Male childcare and housework seems to be relatively independent of changes in socio-economic factors. Women, on the other hand, react to the presence of children, marital status and wage rate.

Mancini and Pasqua (2011) use data from the Italian Time Use Survey (1988/99 and 2002/03) and focus on couples with children. They explicitly discriminate between basic childcare and quality time with children. They find that female working time has a strong effect on childcare time of both spouses and that childcare is substituted within the household if the working time of women increases. Women exhibit a stronger reaction to the characteristics of the family than men do, but mothers' work hours are less responsive in 2002 than in 1988.

Bredtmann (2010) uses the German Time Use Survey. She focuses on factors influencing the way couples share their work. Since all time uses are potentially endogenous, she has to draw on instrumental variables. Her findings suggest a reaction of women on the time allocation of the partner, but not vice versa. She writes about a "first mover advantage" of men.

Deding and Lausten (2006) test several theories concerning intra-household division of market and non-market work. Individual time use is estimated, including the partners' time use as an explanatory factor. They also rely on instruments. Using the Danish Time Use Survey, they find some substitution between own and partners' time uses. But it is significant for women only; that is, if the male does more paid work, the female does more unpaid work, but not vice versa. Their results show some evidence for the so-called *assortative mating* and for *doing gender*[1].

[1] These and the other theories examined are described in detail in Deding and Lausten (2006).

Intra-household time allocation on a weekly basis is investigated by Ettema and van der Lippe (2009). According to the authors, role expectations of women play an important role for the division of paid and unpaid work between the spouses. The presence of children results in a specialization of men in market work and women in unpaid work.

We contribute to the existing literature in the following way. We take account of the possible endogeneity of working hours by defining the occurrence of a plant closure as an exogenous variation of market work time.[2] In a second set of estimations, we also include dismissals as a treatment. In contrast to using an instrumental variable approach, we thereby do not face the issue of finding reasonable instruments. In addition, we use a very rich panel dataset (GSOEP). This offers the opportunity to examine individual reactions to such an employment shock. By applying a difference-in-differences approach, we get the following information: To what extent do men and women in the short term change their non-market work hours if their market work hours are exogenously reduced? If, for example, the male partner, who is affected by a plant closure, increases his housework or childcare significantly after the treatment occurred, this indicates that the reason for the time allocation prior to the treatment have been time constraints rather than preferences. This will yield some indication as to whether time constraints or preferences are the main motive for intra-household time allocation. To the best of our knowledge, the only study that seems to be similar to the study at hand is that of Solaz (2005). She analyzes whether there is an adjustment in domestic work if one partner is unemployed. The analysis is based on the French time use survey including one wave. Therefore, it is not possible to identify an adjustment process. Moreover, being in the state of unemployment is not the same as being laid off in a given period. Instead, by using plant closures or dismissals as an exogenous treatment at time t, we do not condition on still being unemployed in the next period, t+1.

B.3 Data

We use the German Socio-Economic Panel (GSOEP)[3] including the survey years 1992 to 2010. We exclude persons who are in formal education, retired, self-employed, on maternity leave, in short-time work, in a sheltered workshop, civil servants, and persons doing military or civil service, because these groups are not at risk of being laid off. The spouse is allowed to be in any labor status, though. One may ask why we do not exclude those who are unemployed, because they cannot be treated. But this is not necessarily true. The status "not employed" refers

[2] Other papers using this exogenous variation include Schmitz (2011) and Del Bono et al. (2012).
[3] For a description of this very rich panel dataset, see Wagner et al. (2007).

to the time the interview took place. So, if a person is interviewed in March, and then starts working in May and is subject to a closure in November, we would exclude a treated individual from the treatment group. We exclude multigenerational households and other non-standard households. The sample is restricted to cohabiting couples at working age (25-59 years). This seems reasonable since couples that are not living together face different time constraints. In addition, at this age, it is unlikely that people are still attending school or are already retired. Since this work focuses on gender differences, we also exclude couples of the same sex. For families with children, we use the information about the number of children in different age groups living in the household. The composition of couples might change during the observed period, so we define a couple as a unique combination of a man and a woman. One person can be part of different couples over time and can switch from one household to another. This is necessary and plausible since living with another partner may lead to different decisions about time uses or different time constraints.

We only consider time use on weekdays. This is for two reasons. First, in the GSOEP, time use on weekends is only included every second year. Second, time use on weekends is expected to be quite different from that on weekdays, because most market work is done on weekdays. In addition, there are expected to be differences between Saturday and Sunday.

To perform the difference-in-differences estimation, which will be presented in detail in section B.5.1, a treatment group, a control group and pre- and post-treatment periods must be defined. Since we are using a panel dataset, in which a treatment can occur in every year from 1992 to 2010, this raises the question of how to define these variables. First, we clean the control group. We drop every household from the control group that was affected by a treatment at some point in time. People in the control group can have lost their job due to other reasons, for example due to the expiration of a fixed-term contract. This is reasonable because in those cases, the working time is not *exogenously* set to zero. When investigating closures separately, we also exclude from the control group those who are affected by a dismissal. In addition, we do not include a couple if both partners are affected by a treatment at the same time. For each treatment definition (closures in particular or closures and dismissals), we save two distinct datasets, one for men and one for women at risk.

For the definition of pre- and post-treatment period, we define the period-dummy separately for each two-year set and append the resulting two-year-datasets afterwards. As a result, each year, except for the first and the last, is included twice for each couple, as pre- and as post-treatment observation. This almost doubles the number of observations. Since these are only duplicates (except for the period-dummy), this may result in misleading standard errors. We use cluster-robust standard errors to take this into account.

After restricting the data in the aforementioned way, there remain 407 (326) couples with the male (female) partner being affected by a plant closure. If we include dismissals as a treatment, there are 1,167 and 902 affected couples. The control groups have the same size in both cases, namely 6,839 for men and 7,709 for women. The control group is substantially larger since the selected treatments are rare events. Note that one couple might be treated more than once. If affected twice in a row, we only consider the first treatment, since we need the time use prior to any treatment. The couples in the control group are observed one to nineteen years, but the latter case is rare. The treatment group couples are in the sample for two, four, six or eight years.

This work will also consider the case where there is no distinction between men and women being at risk. In this case, we observe 716 (1,946) couples affected by a plant closure (or also a dismissal). We append the two regular panel datasets for treated men and women and distinguish between the treated person and the partner of the treated person, which in both cases can be a man or a woman. The control group contains 8,189 couples. Since we appended the two datasets, the control group contains many observations twice, once with the man being the "potentially treated person", once with the woman being the "potentially treated person". We account for this via cluster-robust standard errors.

B.4 Descriptive Analysis

We first report some basic information about the time use of couples in our sample, which has been described in the previous section. Table B-1 contains the average hours per day (Monday through Friday) spent on different activities by men and women in the sample. Hours of market work reach about 8.5 per weekday for men and 4.6 for women. A regular finding is that women spend more time on the other reported activities than men, except for repairs. Since the numbers in table B-1 apply to all couple households irrespective of the presence of children, the hours spent on childcare are relatively low at only 3.3 for women. It is also interesting to note that men as well as women use slightly more than one and a half hours for their hobbies, i.e., leisure. The total sum of the reported activities differs between men and women. While men report an approximate sum of 13 and a half hours, women report 14 and a half hours, noting that sleep is not included.[4]

[4] This is due to the fact that sleep is included as a separate question in the GSOEP questionnaire only from the year 2008 onwards.

Table B-1: average time spent on activities – with or without children

	men			women		
	mean	sd	se mean	mean	sd	se mean
paid work	8.541	3.500	0.018	4.599	3.988	0.021
housework	0.667	0.871	0.004	2.987	1.871	0.010
shopping etc.	0.716	0.748	0.004	1.301	0.774	0.004
repairs, garden	0.951	1.070	0.005	0.598	0.829	0.004
hobby	1.687	1.658	0.008	1.659	1.528	0.008
childcare	0.834	1.638	0.008	3.255	4.988	0.025
total	13.632	3.330	0.017	14.570	4.936	0.026

Note: the numbers of observations vary around 38,500 depending on the variable under consideration due to missing values. Moreover, the number of persons differs slightly between males and females. Source: GSOEP, 1992-2010, own calculations.

In table B-2, we condition on the presence of children under age 17. Men with children work slightly more than the general population (9 compared to 8.5 hours). Women with children spend considerably less time on market work (3.5 compared to 4.6). For childcare and housework, the converse is true. While women with children spend 9.3 hours on childcare and housework, men only spend 2.1 hours on these two activities. Thus, in the presence of children, the division of labor clearly is more pronounced.

Table B-2: average time spent on activities – couples with children below age 17

	men			women		
	mean	sd	se mean	mean	sd	se mean
paid work	9.007	3.009	0.021	3.488	3.623	0.026
housework	0.624	0.864	0.006	3.385	1.927	0.013
shopping etc.	0.656	0.728	0.005	1.360	0.772	0.005
repairs, garden	0.874	0.986	0.007	0.566	0.777	0.005
hobby	1.464	1.459	0.010	1.451	1.371	0.010
childcare	1.511	1.969	0.014	5.946	5.499	0.038
total	14.333	3.199	0.023	16.322	5.433	0.038

Note: the numbers of observations range vary around 20,500 depending on the variable under consideration due to missing values. Moreover, the number of persons differs slightly between males and females. Source: GSOEP, 1992-2010, own calculations.

The large differences between men and women on one side and between women with and without children on the other side may be due to differences in opportunities (e.g., time constraints) or to preferences. One way to improve opportunities to work is the provision of public or subsidized childcare. If lack of opportunities was the main driver of the observed differences, a policy to increase access to childcare would increase labor force participation of women, reduce private childcare, and diminish the large gender differences. Using differences in access to public childcare to identify the causal effects of changes in opportunities (or time constraints) would lead to endogeneity problems if public policies were not exogenous. Thus, we want to look at changes in working hours that are imposed on workers in some exogenous manner. We first use plant closures that we deem exogenous – at least with respect to the variables of interest in our study. Thus, we next report changes in time use following a plant closure.

Table B-3 demonstrates how little the different dimensions of time use change in response to such a shock. For illustrative reasons, we present the changes for persons who lost their job due to a plant closure and who are not employed in the next period. In the following econometric analysis, we will not condition on being unemployed, since this again might be endogenous. While the reduction in hours of paid work is large, other activities, such as childcare, change little. However, the total time spent on reported activities drops considerably, i.e., time spent on other activities that are not reported increases considerably. The reaction of the partner is negligible. In particular, the partner does not make up for the reduction in paid work by increasing his or her hours of market work. The laid-off person (male or female) does not increase childcare time in a substantial way. Interestingly, the patterns of changes are akin across gender. This first descriptive analysis indicates a strong stability of behavior even after a negative employment shock.

When not conditioning on being unemployed in the period following the plant closure, the effects are somewhat weaker. The reason is that many persons leave unemployment (or more specifically, become reemployed) within a year after the plant closure.

When it comes to the changes due to a plant closure in the pooled dataset, the values are the following (if not employed in the subsequent period). The treated person, on average, reduces market work by about seven hours. Some additional time is spent on the other time uses. The change in leisure time amounts to almost one extra hour. Shopping, repairs and childcare increase by about half an hour, and housework by about 45 minutes. The total time is reduced by almost three hours. The partner's reaction is quite close to zero for all time uses. Even when it comes to childcare, the reaction of the partner only amounts to about 6 minutes less childcare time. All in all, the extent of the changes lies, as expected, in between the male and the female reaction presented in table B-3.

Table B-3: change in time spent on activities after firm closure, not employed

	men laid off			female partner		
	mean	sd	se mean	mean	sd	se mean
paid work	-8.533	3.365	0.260	-0.084	2.287	0.171
housework	0.787	1.348	0.100	-0.005	1.556	0.115
shopping etc.	0.672	1.039	0.077	-0.076	0.961	0.071
repairs, garden	0.885	1.727	0.128	-0.033	0.809	0.060
hobby	1.295	2.617	0.193	0.038	1.693	0.125
childcare	0.565	2.795	0.206	-0.197	2.808	0.208
total	-3.488	4.563	0.359	-0.402	4.223	0.320
	women laid off			male partner		
	mean	sd	se mean	mean	sd	se mean
paid work	-5.340	3.825	0.319	0.000	2.714	0.222
housework	0.766	2.483	0.200	0.006	1.072	0.086
shopping etc.	0.271	1.397	0.112	0.079	0.895	0.073
repairs, garden	0.155	1.344	0.108	0.039	1.108	0.089
hobby	0.510	1.825	0.148	0.026	2.010	0.162
childcare	0.532	3.036	0.245	0.000	1.069	0.086
total	-2.128	4.742	0.399	-0.090	3.162	0.264

Note: the numbers of observations vary around 160 depending on the variable under consideration due to missing values. Moreover, the number of persons differs slightly between males and females. Source: GSOEP, 1992-2010, own calculations.

B.5 Empirical methods and results

B.5.1 Methods

This work seeks to shed light on the effect of a treatment (i.e., an exogenous variation of the time devoted to market work) on the time use decisions of men and women living as couples. So basically, we ask what people use their time for if they are facing an employment shock. This shock, as already mentioned, can be a plant closure or a dismissal. We distinguish two definitions of a treatment, namely *closure* and *closure or dismissal*. We expect people affected by such a shock to spend their "additional" time on other liabilities, such as childcare and housework. This analysis concentrates on childcare and housework, as it is specifically concerned with time devoted to domestic work.

We apply a difference-in-differences approach (DID). This seems promising in ruling out the risk that the observed treatment effect is not due to the treatment, but due to other variable changes occurring during the period under consideration, as would be the case in a simple before-after-design (for a description, see Lee (2005)). In the DID framework, the before and after difference in the outcomes of a group that is not subject to the treatment – the control group – is compared to the time difference of the treatment group. The two differences are subtracted from each other to yield the *difference in differences*.

In the simplest case, it is made use of two differences: the difference between one year and another and the difference between treatment and control group. In the basic DID, one considers two groups and two time periods (Cameron and Trivedi (2005)). What is computed in this case is the following difference (the notation is adapted to be consistent here):

$$\widehat{\delta_1} = (\bar{y}_2^B - \bar{y}_1^B) - (\bar{y}_2^A - \bar{y}_1^A) \qquad \text{(B-1)}$$

1 and 2 are the time periods, *B* indicates the treatment group and *A* indicates the control group. Term (B-1) can be described as the *difference between the time differences of the treatment and the control group*.

In this case, it is not so simple because there is no clear treatment period for the entire sample. This is due to the fact that we use 19 years of the GSOEP and a treatment can occur in every single year. The solution to this is explained in section B.3.

In a first step, we estimate this regular DID model using OLS and including control variables such as age, number of children in different age groups, living in East Germany, years of education, but also year dummies. The model for repeated cross-sections is as follows (Wooldridge (2010)):

$$y = \beta_0 + \beta_1 dB + \delta_0 d2 + \delta_1 d2 \cdot dB + u \qquad \text{(B-2)}$$

In a second step, we apply this to more than two time periods (Wooldridge (2010)). In this case, we can use our panel datasets (in which each year is only included once) and run a pooled OLS estimation.

Since we use a rich panel dataset, we also apply a panel approach as a third step. The estimation of the following is implemented using either first differences estimated with pooled OLS or fixed effects (Imbens and Wooldridge (2009)):

$$y_{it} = \lambda_t + \tau D_{it} + x_{it}\gamma + c_i + u_{it}, \qquad t = 1, \dots, T \qquad \text{(B-3)}$$

Here, τ represents the treatment effect.

The estimation of a treatment effect is subject to several assumptions (Cameron and Trivedi (2005) as well as Lee (2005)). The first is the *conditional independence assumption* (Cameron and Trivedi (2005)). This is valid if the outcome is independent of treatment assignment, conditional on covariates. Random assignment will be sufficient here. This should not be a problem in this case since a closure (or a dismissal) can occur to anyone who is working. It is not an individual choice. In addition, we include several covariates to control for differences between the two groups. The second assumption is the *matching (or overlap) assumption* (Cameron and Trivedi (2005)). This assumption implies that for all values of the covariates, there are both treated and untreated individuals, but it can be relaxed under certain circumstances. If one is interested in the average treatment effect on the treated (ATET), the assumption can be relaxed in the sense that, for each treated household, there is an equivalent household that is not treated. According to the *conditional mean assumption*, the outcome does not define the selection into treatment (Cameron and Trivedi (2005)). This is not violated as well. A further assumption is the *common trends assumption* (Cameron and Trivedi (2005)). It assumes that time effects are the same for the treatment and the control group. Since the two groups are drawn from the same population and the groups for women and men are defined separately, we do not consider this problematic.

This work investigates the treatment effect on childcare and housework separately as well as on the sum of those two (referred to as *domestic work*). Note that we always refer to the time use on a regular weekday. Separate estimations are run for the time use of those who were laid off and for their partners. Each model is run eight times since we define four treatment groups and we estimate the reaction of each spouse. We additionally run separate estimations for couples with and without children up to age 16, although in those cases, the treatment group only amounts to about half of its original size.

As a robustness check, we "trim" the control group so as not to exceed the maximum or deceed the minimum (+ / - one standard deviation) of the treatment group covariates. In addition, we pool the datasets of treated men and women. In this case, we will distinguish between the reaction of the treated person and the reaction of the partner, irrespective of whether a man or a woman is laid off.

B.5.2 Results

B.5.2.1 Difference-in-differences estimation

Childcare

The results obtained by the regular DID estimation are presented in tables B-4 to B-6.

Table B-4: Regular DID, plant closure, childcare

	Closure (men)		Closure (women)	
	1	2	3	4
childcare time of:	men	women	men	women
treatment variable	0.193 *	-0.123	0.035	0.188
	(0.115)	(0.242)	(0.094)	(0.219)
household income (1000 €)	-0.064 ***	-0.142 ***	-0.039 ***	-0.087 ***
	(0.005)	(0.016)	(0.007)	(0.015)
# children 0-3	0.907 ***	6.000 ***	0.933 ***	6.006 ***
	(0.031)	(0.095)	(0.032)	(0.100)
# children 4-6	0.721 ***	3.093 ***	0.697 ***	3.116 ***
	(0.031)	(0.082)	(0.028)	(0.079)
# children 7-12	0.519 ***	1.871 ***	0.484 ***	1.930 ***
	(0.018)	(0.048)	(0.017)	(0.046)
# children 13-16	0.211 ***	0.617 ***	0.206 ***	0.630 ***
	(0.018)	(0.045)	(0.017)	(0.042)
years of education (woman)	0.017 ***	-0.057 ***	0.012 ***	-0.059 ***
	(0.004)	(0.010)	(0.004)	(0.010)
years of education (man)	-0.033 ***	0.042 ***	-0.032 ***	0.029 ***
	(0.004)	(0.010)	(0.004)	(0.009)
East-Germany	0.051 **	-0.650 ***	0.095 ***	-0.614 ***
	(0.024)	(0.049)	(0.025)	(0.048)
observations	53,696	53,677	59,345	59,332
R-squared	0.218	0.513	0.202	0.5

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

The treatment effect of a plant closure affecting men is positive but only weakly significant for men (table B-4). The magnitude of the reaction is also small (about 12 minutes). The sign of the female change in childcare time is negative, but the effect is insignificant. If women are at risk, the effect on male childcare is not significant. The effect on women is positive but also insignificant. The covariates referring to the number of children in different age groups all have a positive effect on the childcare time of both men and women. The age group 0-3 shows the strongest impact, and the age group 13-16 shows the weakest. This is not surprising, since the younger the child, the more care is necessary. One's own years of education seem to be related to less childcare time, whereas the opposite is true for spousal education. Living in East Germany is significantly associated with less female childcare time both if men or women are at risk. This is expected, since in East Germany the use of institutional care, even for very young children, is much more common (and available) than in West Germany (e.g., Federal Statistical Office of Germany (2012)). For men, the opposite is true. They spend more time on childcare than do men in West Germany.

If dismissals are included as a treatment, the effect of a shock affecting men significantly raises their childcare time. They spend about 22 additional minutes on childcare on weekdays. The effect on female childcare is negative but small and insignificant. Treatments affecting women do not have an impact on male childcare, but women (significantly) undertake almost an additional half-hour of childcare. Again, the coefficient of children in different age groups is smaller the older the children are. The effects of "living in East Germany" and education are comparable to the case of plant closures.[5]

If only couples with children are included, the sample is reduced by roughly half. As expected, the effect of a plant closure on childcare time almost doubles to about 23 additional minutes a day (treatment group: men). Again, the effect is only weakly significant. It seems interesting that the female partners' reaction now amounts to about 14 minutes less childcare time, but the effect still remains insignificant. The female years of education, in this case, raise female childcare time, as does male education. For men, there is a similar pattern as in the estimations described earlier. Women living in the east now spend more than two and a half hours less time on childcare, men only a few minutes. If women are potentially affected by a plant closure, their childcare time now increases by about 25 minutes, but this is not significant. The reaction of the male partners is positive and stronger than when childless couples are included, but it is not significant. The results concerning the covariates are comparable to those obtained if men are the treatment group.

In summation, it is found that people adjust their childcare time to some extent when facing an employment shock. However, the change is significant and of a considerable extent only when dismissals are included as a treatment or if only couples with children are considered. As already mentioned, dismissals might be endogenous. The results gleaned if only couples with children are included should be interpreted with caution since, in this case, the treatment group is quite small.

[5] We do not present the detailed results of those estimations since the inclusion of dismissals might cause an endogeneity problem.

What can be derived from these results? It seems that people show little change in their childcare time as a result of plant closures. There is some indication that people do not just allocate their time the way they do due to time constraints. The allocation of childcare within couples seems to be quite stable even if one partner has additional hours to allocate.

Housework

When housework is the dependent variable, the following results are found (table B-5). When men are the treatment group, closures significantly raise the time spent on housework. The same is true when including dismissals. The extent is, again, small. Men spend about 13 additional minutes on housework if closure is the treatment and about 22 more minutes if we include dismissals. In case of closures affecting men, there is almost no reaction from women; the effect is slightly positive but insignificant. If dismissals are included, the coefficient is small, negative and weakly significant. If women are the treatment group, closures do not have a significant effect on housework time, neither for men nor for women. If dismissals are included, the treatment does have an effect. Women then spend about an extra 30 minutes on housework. This coefficient is highly significant. Men do not show a reaction. The effect of children in different age groups is not as distinct as in the case of childcare. This is not surprising since the amount of housework does not depend on children's needs in the same way childcare time does. When investigating male housework time, children seem to lower the time allocated to this activity. The descending order of the age groups remains when investigating the housework time of women though. In all cases, i.e., closures or also dismissals affecting men or women, female housework time is greater the younger the children. Children seem to lower male housework time but the extent only amounts to a maximum of about three minutes. The results also reveal a difference in housework time between those living in East and West Germany. Women who live in the east significantly spend about half an hour less on housework than do their counterparts in the west. The opposite is true for men, but here the coefficient is quite small (about 6 minutes a day).

If we only consider couples with children below age 17, plant closures affecting men raise their housework time by about 9 minutes, though the effect is not significant. Men without children do about 19 extra minutes of housework in response to a plant closure (highly significant). The female partners do not show any considerable reaction to the employment shock affecting their male partners. If women are the treatment group, they increase their housework by approximately 8 minutes, but again, this is not significant. The reaction is somewhat stronger when only considering couples without children. At this point, it seems worth mentioning that in the GSOEP, there is no distinction between primary and secondary activities. It is therefore possible that couples with children do some extra housework but report it as childcare time because they consider this to be the primary activity.[6]

Table B-5: DID estimation, plant closure, housework

	Closure (men)				Closure (women)			
	1		2		3		4	
housework time of …	men		women		men		women	
treatment variable	0.220	***	0.035		0.057		0.195	
	(0.066)		(0.108)		(0.069)		(0.141)	
household income (1000 €)	-0.057	***	-0.112	***	-0.039	***	-0.053	***
	(0.004)		(0.009)		(0.005)		(0.013)	
# children 0-3	-0.034	***	1.05	***	-0.018		1.034	***
	(0.013)		(0.028)		(0.012)		(0.029)	
# children 4-6	-0.020		0.73	***	-0.027	**	0.749	***
	(0.012)		(0.027)		(0.011)		(0.027)	
# children 7-12	-0.033	***	0.635	***	-0.047	***	0.636	***
	(0.007)		(0.017)		(0.007)		(0.016)	
# children 13-16	-0.047	***	0.529	***	-0.05	***	0.529	***
	(0.010)		(0.020)		(0.009)		(0.019)	
years of education (woman)	0.032	***	-0.12	***	0.029	***	-0.132	***
	(0.002)		(0.005)		(0.002)		(0.005)	
years of education (man)	-0.014	***	-0.006		-0.018	***	0.006	
	(0.002)		(0.004)		(0.002)		(0.004)	
East-Germany	0.104	***	-0.586	***	0.117	***	-0.609	***
	(0.014)		(0.026)		(0.013)		(0.026)	
observations	53,691		53,578		59,340		59,218	
R-squared	0.025		0.228		0.026		0.209	

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

Domestic Work

Domestic work is defined as the sum of childcare time and housework time. Investigating the sum of these two seems prudent since, as discussed in the previous paragraph, both time uses are non-market work and one might do household chores while simultaneously supervising children. Designating an activity as childcare or housework is not always clear. The treatment effect approximately amounts to the sum of the two specific ones if men are affected by a closure (table B-6). They spend additional 25 minutes on domestic work. This is highly significant. Female partners' reaction is still not significant and slightly negative. If women are

[6] We do not discuss the results we get when we include dismissals as a potential treatment since those are not necessarily exogenous. This is discussed above.

at risk of facing a plant closure, the increase in domestic work by about 23 minutes is still not significant. The male reaction remains negligible. The other covariates limn a similar picture as already presented above. The numbers of children in different age groups again reveal, in a descending order, a positive coefficient in respect to housework time of both partners.

Table B-6: Regular DID, plant closure, domestic work

	Closure (men)				Closure (women)			
	1		2		3		4	
domestic time of …	men		women		men		women	
treatment variable	0.410	***	-0.105		0.092		0.382	
	(0.149)		(0.284)		(0.132)		(0.286)	
household income (1000 €)	-0.121	***	-0.252	***	-0.078	***	-0.139	***
	(0.007)		(0.020)		(0.012)		(0.026)	
# children 0-3	0.871	***	7.062	***	0.913	***	7.054	***
	(0.037)		(0.103)		(0.039)		(0.109)	
# children 4-6	0.703	***	3.814	***	0.671	***	3.859	***
	(0.037)		(0.091)		(0.034)		(0.088)	
# children 7-12	0.487	***	2.509	***	0.438	***	2.565	***
	(0.021)		(0.054)		(0.020)		(0.051)	
# children 13-16	0.164	***	1.147	***	0.156	***	1.158	***
	(0.022)		(0.053)		(0.021)		(0.050)	
years of education (woman)	0.049	***	-0.176	***	0.042	***	-0.191	***
	(0.005)		(0.012)		(0.005)		(0.012)	
years of education (man)	-0.047	***	0.035	***	-0.049	***	0.034	***
	(0.005)		(0.011)		(0.005)		(0.011)	
East-Germany	0.156	***	-1.237	***	0.213	***	-1.223	***
	(0.031)		(0.057)		(0.032)		(0.058)	
observations	53,684		53,558		59,334		59,202	
R-squared	0.147		0.503		0.142		0.483	

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

The additional time spent on domestic work as an aggregate measure of home production amounts to almost half an hour a day. So while there is not a strong reaction to a plant closure in respect to the specific non-market time uses, there is a considerable reaction of men concerning non-market work. Since the reported effects relate to average weekdays, a reaction of 25 minutes a day approximately amounts to 2 hours a week, excluding weekends. But still, this might indicate that a main driver of intra-household time allocation are preferences. On average, irrespective of the current labor status, the male market work is reduced by slightly more than three hours in response to a plant closure. Note that those who are already working again are also included in this measure. Leisure time is, on average, increased by almost half an hour. The total time reported is reduced by about one and a half hours, whereas those not affected by a treatment only show an average reduction of total time of about three minutes within one year. This reveals that time constraints are markedly loosened by a plant closure in the short term, but the additional time spent on domestic work only makes up a small part of the newfound time. What men use this time for remains unclear. Maybe they sleep more or use it for job search. The partner, however, does not react by changing his or her non-market work time to a noticeable extent.

Repeated Cross-sections

Using the repeated cross-sections model does not reveal large differences. The extent of the reactions is comparable to those obtained before. When dismissals are included as a treatment for men, the effect on female housework is highly significant. The extent is about the same. The effect on female domestic time is somewhat weaker but significant at the 5% level. Female domestic time is significantly (5%) raised in response to being laid off due to a plant closure, but the extent remains unchanged.

B.5.2.2 Panel

Childcare

For men, closures seem to have a weak, increasing effect on childcare time on weekdays (table B-7). When using fixed effects and cluster-robust standard errors, the impact is weakly significant.

For women, this effect is comparable but not significant. The spousal reactions are, again, weak and insignificant. The coefficients concerning children in different age groups also show similar effects as in regular DID. There is no significant effect of years of education. Still, the sign of the coefficients is as expected and the same as shown in the previous section. Living in the east also is insignificant. It is not surprising that those effects are not significant because we use fixed effects and those variables are not expected to vary greatly over time.

For the model including dismissals, the results are as follows. The childcare of both men and women increases when affected by a treatment. The spousal reaction is not significant and very small in both cases. The rest of the results are similar to those of the consideration of closures only. Only the dummy for *living in East Germany* shows slightly different impacts in the case of men, but it is insignificant anyway. Note, again, that including dismissals might create a problem of endogeneity.

Table B-7: Fixed Effects, plant closure, childcare[7]

	Closure (men)		Closure (women)	
	1	2	3	4
childcare time of …	men	women	men	women
treatment variable	0.197 *	-0.053	0.045	0.185
	(0.105)	(0.148)	(0.064)	(0.136)
observations	33,985	33,971	37,862	37,851
R-squared	0.077	0.306	0.075	0.294
# of couples	6,721	6,719	7,452	7,451

*** p<0.01, ** p<0.05, * p<0.1
robust standard errors in parentheses

Source: GSOEP, 1992-2010, own calculations.

If we only include those with children up to age 16, the results differ markedly. As before, this halves the sample size. If men are subject to plant closures, the treatment effect amounts to about 19 minutes (compared to 12 minutes as shown in table B-7) but it is no longer significant – not even weakly. Female partners insignificantly lower their childcare time by slightly more than nine minutes. This amount is also markedly higher than when including childless couples. If women are at risk of being affected by a plant closure, they do 21 additional minutes of childcare, but again, this is insignificant. The male reaction is also about twice as high as shown in table B-7, but there still is no significance. The reaction of both partners is approximately zero if only investigating those without young children.

When not distinguishing between men and women being affected by a treatment, the results are as follows. Again, fixed effects are used. The results do not differ to a great extent from those obtained from the separate estimations. The treated person is the one whose reaction is somewhat stronger, and the partner almost does not change his or her time schedule at all. The effect of a closure is now significant at the 5% level, which can simply be ascribed to the larger treatment group. The extent of the effect is comparable, however. The coefficients of the included covariates show some differences. The children in different age groups reveal a stronger coefficient for male childcare time, but a weaker one for female childcare time. This is expected. For both treatment definitions, more years of education of the potentially treated person result in more childcare time of the partner. Solely investigating couples with children reveals stronger reactions, as expected. The partner affected by a closure spends about 21 additional minutes on childcare. This effect is significant at the 5% level. The partner who is not treated somehow lowers his or her childcare time, but only by about 4 minutes, and the result is insignificant.

[7] The complete results are presented in the appendix.

Housework

Closures affecting men result in a significant rise in housework time of men, but again, they only spend about 14 additional minutes a day (table B-8). The female spouses do not noticeably change their housework time. If women are at risk of a plant closure, the treatment effect amounts to about 12 additional minutes of housework, but this is only weakly significant. Male spouses do not react noticeably by changing their housework time. If dismissals are also included as a treatment, male housework time is significantly raised by about 22 minutes. The reaction of female spouses is slightly negative and significant at the 5% level. If women are at risk, they significantly spend an extra half-hour on housework. Their male spouses do not vary their housework time. The covariates draw a similar picture as in the case of regular DID.

Table B-8: Fixed Effects, plant closure, housework[8]

	Closure (men)		Closure (women)	
	1	2	3	4
housework time of :	men	women	men	women
treatment variable	0.229 ***	0.047	0.060	0.201 *
	(0.063)	(0.073)	(0.053)	(0.116)
observations	33,977	33,910	37,856	37,778
R-squared	0.007	0.056	0.006	0.049
# of couples	6,720	6,715	7,453	7,446

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

Solely including couples living with children up to age 16 reveals some differences. If men are subject to plant closures, they do an additional ten minutes of housework. This effect shows a weak significance. The change in housework of their female partners is negligible. If women are at risk, they raise their housework by about nine minutes, but again, this is not significant. The same is true for their partner's reaction, which is even weaker. Again it is observed that the reactions are stronger when only including couples without children. As already mentioned, this could be due to the fact that for people with children, housework might be a secondary activity while they are watching the children.

Plant closures in the pooled version result in a rise of the housework time of the treated person of about 13 minutes (highly significant), whereas the reaction of the partner is close to zero and also insignificant. Again, children in different age groups do raise the housework time of the treated person as well as that of the partner, but the size of the coefficient is not as large as that

[8] The complete results are presented in the appendix.

in the case of childcare. Still, the younger the children, the greater the extent of the effect. One's own years of education lower housework time while the partner's education raises it, but this is not significant. The region in which the couple lives, as shown earlier in the case of fixed effects estimation, does not reveal any significant coefficient. The change in housework time if only couples with children are investigated is, again, smaller. An additional 10 minutes are spent on housework by the treated partner (weakly significant), whereas the change of the partner is negligible.

Domestic work

Adding the two time uses reveals a highly significant treatment effect of closures on male domestic work (table B-9). They spend additional 26 minutes on these activities. Again, this effect is about the same as the sum of the two separate effects. The women's reaction is negligible. The effect on female domestic work if they are at risk of being laid off due a plant closure amounts to about 23 additional minutes. In contrast to the regular DID approach, the change in domestic work is weakly significant. The male partners do not display notable reactions.

Table B-9: Fixed Effects, plant closure, domestic work[9]

	Closure (men)		Closure (women)	
	1	2	3	4
domestic time of:	men	women	men	women
treatment variable	0.427 ***	-0.015	0.105	0.385 *
	(0.140)	(0.178)	(0.083)	(0.202)
observations	33,973	33,897	37,851	37,766
R-squared	0.048	0.31	0.047	0.293
# of couples	6,720	6,714	7,452	7,445

*** p<0.01, ** p<0.05, * p<0.1
robust standard errors in parentheses

Source: GSOEP, 1992-2010, own calculations.

Pooling the dataset reveals a treatment effect that lies between that for men and for women, which again is not surprising. The covariates do not reveal large differences to the aforementioned results.

Robustness Checks

The results concerning the treatment effect on childcare time do not essentially differ from those obtained by the fixed effects approach if we instead estimate in first differences. This is true for

[9] The complete results are, again, presented in the appendix.

plant closures as well as for the "comprehensive" treatment. The extents slightly differ, but the picture drawn remains unchanged. The same is true for the estimations concerning housework time and domestic time. The control group is also trimmed so as not to exceed the maximum and not fall below the minimum (+/- one standard deviation) of the treatment group in respect to the covariates. Then fixed effects estimations are run. The results are comparable to those of the fixed effects estimation for the full sample.

B.6 Conclusions

This work investigates the question of how couples use their time when facing a negative employment shock. This shock can be a plant closure or a dismissal, noting that dismissals might be endogenous. The underlying question is whether people's time allocation is due to time constraints or to preferences. This work concentrates on childcare and housework time as domestic work. Using the survey years 1992-2010 of the German Socio-Economic Panel (GSOEP), it considers cohabiting couples at working age (25-59). We perform a difference-in-differences approach. We also exploit the panel structure of the GSOEP and apply fixed effects estimations. All our results reveal that there is some reaction in the considered time uses of men and women, but the extent is low and the changes often are insignificant. This indicates that the time allocation of men and women represents their preferences to some extent. Even if men are facing a plant closure, they do not spend substantially more time on housework or childcare. If adding the two non-market time uses and analyzing the change in domestic work as an aggregate measure, male time devoted to non-market tasks is raised considerably. But relative to the total time that needs to be reallocated, this still is a small share. The female reaction almost always is not significant.

The results support that male and female time allocation to domestic tasks is relatively stable if an employment shock occurs. On average, men have about three hours a day to reallocate if they have been facing a plant closure, from which they spend 25 minutes on additional domestic work as a reaction to this shock. Interestingly, partners do not react to a mentionable extent. There does not appear to be substantial substitution of domestic tasks between partners if one needs to reallocate his or her time. Overall, the time allocation within couples seems, in large part, to be the result of preferences rather than of time constraints.

As robustness checks, we trimmed the sample and also performed first differences estimations instead of fixed effects. Both did not change the baseline of our results. We also pooled the samples of men and women. Doing this also did not result in new findings.

The distinction between families and childless couples reveals a stronger reaction of the treated person with respect to childcare, which is not surprising. The implications do not change, though. Childcare is not raised by more than 45 minutes a day. The effect on housework is weaker than when investigating families and childless couples altogether. This might be due to the fact that the GSOEP does not contain the information about primary and secondary activities. It is important to note that conditioning on the presence of children halves the sample sizes so that the results should be interpreted with caution.

This analysis explores reactions in the short term. We only consider the time use in the year immediately after a person has been laid off. To investigate the effects in the long term, we would need more observations over a long time period for the treated group. We also do not distinguish between those already working again and those who are still unemployed. This is also due to data constraints since a plant closure is an infrequent event and we only observe 407 (resp. 326) affected households in our sample. Further restrictions according to labor status would result in too few observations. This would also result in an endogeneity problem since the occupational status the year after the treatment cannot be considered exogenous. Over all, it is likely that we somehow underestimate the treatment effects.

B.7 Appendix

Table A B-1: Complete results – Basic Difference-in-differences, domestic work

	Closure (men)				Closure (women)			
	1		2		3		4	
domestic time of …	men		women		men		women	
treatment variable	0.410	***	-0.105		0.092		0.382	
	(0.149)		(0.284)		(0.132)		(0.286)	
household income (1000 €)	-0.121	***	-0.252	***	-0.078	***	-0.139	***
	(0.007)		(0.020)		(0.012)		(0.026)	
# children 0-3	0.871	***	7.062	***	0.913	***	7.054	***
	(0.037)		(0.103)		(0.039)		(0.109)	
# children 4-6	0.703	***	3.814	***	0.671	***	3.859	***
	(0.037)		(0.091)		(0.034)		(0.088)	
# children 7-12	0.487	***	2.509	***	0.438	***	2.565	***
	(0.021)		(0.054)		(0.020)		(0.051)	
# children 13-16	0.164	***	1.147	***	0.156	***	1.158	***
	(0.022)		(0.053)		(0.021)		(0.050)	
years of education (woman)	0.049	***	-0.176	***	0.042	***	-0.191	***
	(0.005)		(0.012)		(0.005)		(0.012)	
years of education (man)	-0.047	***	0.035	***	-0.049	***	0.034	***
	(0.005)		(0.011)		(0.005)		(0.011)	
East-Germany	0.156	***	-1.237	***	0.213	***	-1.223	***
	(0.031)		(0.057)		(0.032)		(0.058)	
observations	53,682		53,556		59,332		59,200	
R-squared	0.147		0.503		0.142		0.483	

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.

robust standard errors in parentheses

Table A B-2: Results if only couples with children up to age 16 are included – effects on childcare time

	Closure (men)		Closure (women)	
	1	2	3	4
childcare time of:	men	women	men	women
treatment variable	0.380 *	-0.231	0.139	0.411
	(0.217)	(0.458)	(0.180)	(0.441)
household income (1000 €)	-0.102 ***	-0.226 ***	-0.109 ***	-0.211 ***
	(0.011)	(0.032)	(0.010)	(0.031)
years of education (woman)	0.056 ***	0.087 ***	0.056 ***	0.093 ***
	(0.007)	(0.018)	(0.007)	(0.018)
years of education (man)	-0.035 ***	0.190 ***	-0.031 ***	0.158 ***
	(0.006)	(0.018)	(0.006)	(0.016)
East-Germany	-0.114 **	-2.660 ***	-0.066	-2.61 ***
	(0.046)	(0.105)	(0.048)	(0.100)
observations	29,449	29,434	32,084	32,075
R-squared	0.065	0.211	0.061	0.201

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

Table A B-3: Results if only couples with children up to age 16 are included – effects on housework time

	Closure (men)		Closure (women)	
	1	2	3	4
housework time of …	men	women	men	women
treatment variable	0.144	0.015	0.075	0.137
	(0.095)	(0.163)	(0.097)	(0.213)
household income (1000 €)	-0.048 ***	-0.106 ***	-0.053 ***	-0.096 ***
	(0.005)	(0.014)	(0.004)	(0.011)
years of education (woman)	0.035 ***	-0.116 ***	0.031 ***	-0.115 ***
	(0.003)	(0.007)	(0.003)	(0.007)
years of education (man)	-0.017 ***	0.016 **	-0.013 ***	0.026 ***
	(0.003)	(0.006)	(0.003)	(0.006)
East-Germany	0.173 ***	-0.89 ***	0.15 ***	-0.967 ***
	(0.019)	(0.042)	(0.018)	(0.039)
observations	29,444	29,355	32,079	31,988
R-squared	0.027	0.092	0.029	0.090

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

Table A B-4: Results if only couples with children up to age 16 are included – effects on domestic work

	Closure (men)		Closure (women)	
	1	2	3	4
domestic time of …	men	women	men	women
treatment variable	0.519 *	-0.237	0.213	0.547
	(0.270)	(0.531)	(0.238)	(0.553)
household income (1000 €)	-0.150 ***	-0.332 ***	-0.161 ***	-0.306 ***
	(0.014)	(0.040)	(0.012)	(0.035)
years of education (woman)	0.091 ***	-0.028	0.087 ***	-0.020
	(0.008)	(0.021)	(0.008)	(0.021)
years of education (man)	-0.053 ***	0.206 ***	-0.044 ***	0.184 ***
	(0.008)	(0.020)	(0.007)	(0.019)
East-Germany	0.062	-3.549 ***	0.086	-3.571 ***
	(0.055)	(0.123)	(0.055)	(0.118)
observations	29,437	29,339	32,073	31,977
R-squared	0.058	0.181	0.059	0.168

*** p<0.01, ** p<0.05, * p<0.1
robust standard errors in parentheses

Source: GSOEP, 1992-2010, own calculations.

Table A B-5: Results Fixed Effect – closure, childcare time

	Closure (men)		Closure (women)	
	1	2	3	4
childcare time of …	men	women	men	women
treatment variable	0.197 *	-0.053	0.045	0.185
	(0.105)	(0.148)	(0.064)	(0.136)
household income (1000 €)	-0.033 ***	-0.128 ***	-0.011 **	-0.034 **
	(0.009)	(0.026)	(0.005)	(0.014)
# children 0-3	0.839 ***	5.634 ***	0.882 ***	5.642 ***
	(0.041)	(0.130)	(0.043)	(0.130)
# children 4-6	0.653 ***	2.921 ***	0.662 ***	2.946 ***
	(0.035)	(0.100)	(0.034)	(0.098)
# children 7-12	0.524 ***	1.968 ***	0.512 ***	2.049 ***
	(0.028)	(0.071)	(0.025)	(0.068)
# children 13-16	0.286 ***	0.964 ***	0.294 ***	1.009 ***
	(0.025)	(0.057)	(0.024)	(0.054)
years of education (woman)	0.019	-0.014	0.016	-0.009
	(0.022)	(0.058)	(0.023)	(0.056)
years of education (man)	-0.006	0.076 *	-0.022	0.033
	(0.015)	(0.044)	(0.016)	(0.045)
East-Germany	0.029	-0.000	0.235	-0.035
	(0.162)	(0.671)	(0.351)	(0.664)
observations	33,985	33,971	37,862	37,851
R-squared	0.077	0.306	0.075	0.294
# of couples	6,721	6,719	7,452	7,451

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

Table A B-6: Results Fixed Effect – closure, housework time

	Closure (men)		Closure (women)	
	1	2	3	4
housework time of :	men	women	men	women
treatment variable	0.229 ***	0.047	0.060	0.201 *
	(0.063)	(0.073)	(0.053)	(0.116)
household income (1000 €)	-0.022 ***	-0.034 ***	-0.011 ***	-0.011 *
	(0.006)	(0.012)	(0.004)	(0.006)
# children 0-3	-0.035 **	0.847 ***	-0.026	0.8 ***
	(0.016)	(0.037)	(0.016)	(0.035)
# children 4-6	-0.019	0.511 ***	-0.021	0.499 ***
	(0.015)	(0.034)	(0.014)	(0.032)
# children 7-12	-0.011	0.375 ***	-0.013	0.359 ***
	(0.012)	(0.026)	(0.011)	(0.024)
# children 13-16	-0.003	0.247 ***	-0.001	0.236 ***
	(0.014)	(0.026)	(0.012)	(0.024)
years of education (woman)	-0.017	0.013	-0.008	0.002
	(0.012)	(0.022)	(0.010)	(0.021)
years of education (man)	-0.030 ***	0.023	-0.026 ***	0.025
	(0.010)	(0.020)	(0.009)	(0.019)
East-Germany	0.175 ***	-0.137	0.212 ***	-0.227
	(0.059)	(0.173)	(0.067)	(0.156)
observations	33,977	33,910	37,856	37,778
R-squared	0.007	0.056	0.006	0.049
# of couples	6,720	6,715	7,453	7,446

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

Table A B-7: Results Fixed Effect – closure, domestic work

	Closure (men)		Closure (women)	
	1	2	3	4
domestic time of:	men	women	men	women
treatment variable	0.427 ***	-0.015	0.105	0.385 *
	(0.140)	(0.178)	(0.083)	(0.202)
household income (1000 €)	-0.055 ***	-0.161 ***	-0.022 ***	-0.043 **
	(0.011)	(0.031)	(0.008)	(0.017)
# children 0-3	0.804 ***	6.492 ***	0.857 ***	6.446 ***
	(0.048)	(0.144)	(0.050)	(0.143)
# children 4-6	0.635 ***	3.427 ***	0.641 ***	3.434 ***
	(0.043)	(0.112)	(0.040)	(0.109)
# children 7-12	0.514 ***	2.34 ***	0.499 ***	2.399 ***
	(0.034)	(0.081)	(0.030)	(0.077)
# children 13-16	0.284 ***	1.211 ***	0.293 ***	1.24 ***
	(0.032)	(0.066)	(0.030)	(0.064)
years of education (woman)	0.002	-0.004	0.007	-0.007
	(0.027)	(0.066)	(0.028)	(0.065)
years of education (man)	-0.036 *	0.095 *	-0.049 **	0.056
	(0.019)	(0.051)	(0.019)	(0.053)
East-Germany	0.204	-0.141	0.449	-0.263
	(0.176)	(0.762)	(0.360)	(0.732)
observations	33,973	33,897	37,851	37,766
R-squared	0.048	0.31	0.047	0.293
# of couples	6,720	6,714	7,452	7,445

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2010, own calculations.
robust standard errors in parentheses

Part C

C Occupational prestige: Is there a career penalty for birth-related career interruptions in Germany?

Katharina Sutter

C.1 Introduction

The reconciliation of work and family life in Germany is a topic that is often discussed in economic literature and political discourse (e.g., Spieß (2011)). One reason for this is the demographic change Germany is experiencing. The proportion of the elderly in the population is increasing (e.g., European Commission (2014)). This has far reaching consequences for the social security system as well as for economic growth (e.g., OECD (2014), European Commission (2012)). Women are therefore an important source of human capital (e.g., Reinberg and Hummel (2003), OECD (2014)). However, a large share of mothers who are in the labor force only work part-time (e.g., Federal Statistical Office of Germany (2012)) and women only make up a minor part of executive positions while they account for approximately 50% of university graduates today (e.g., Holst and Schimenta (2013), Holst and Wiemer (2010)). A female quota for those positions is often brought up in political discourse (e.g., Holst and Schimenta (2013)).

Several instruments have been implemented in the recent decades to provide opportunities for mothers to stay attached to the labor market to a greater extent. One example is the implementation of the “Elterngeld” in 2007 (e.g., German Bundestag (2006)), another is the legal right to daycare for children aged at least one year implemented in 2013 (e.g., German Bundestag (2008)). All of these measures are intended to make work and family obligations compatible. But aside from simply participating in paid labor, there is another important aspect, namely occupational prestige (i.e., powerful positions) and upward occupational mobility (also see Mandel and Semyonov (2006)).

The present study focuses on the following question: Is there a career penalty for employment interruptions linked to childbirth? In this context, the wage penalty for motherhood is often discussed (e.g., Budig and England (2001)), but this is only part of the story. One might ask whether this wage penalty is, in part, the result of a prestige penalty. In this paper, I ask whether this prestige penalty exists in Germany. I investigate whether the duration of the first career interruption related to childbirth has an effect on occupational prestige as well as on upward and downward occupational mobility. The prestige is measured using the Treiman prestige scale (SIOPS). The study focuses on mothers at working age (20-59 years) who give birth to their first child during the observation period and who have been working the year prior to the first birth. The effect of leave length is identified using a selection model according to Heckman. The selection equation displays the likelihood of observing the prestige information after birth, i.e., the re-entry behavior, while the second step displays occupational prestige. I consider two versions of the second step. The first version uses the level of the occupational prestige as a dependent variable, the second version uses a binary variable indicating an upward/downward occupational move of at least 10% compared to the SIOPS level prior to the first birth or subsequent increases resp. declines. I use regular Heckman models as well as an approach suggested by Wooldridge (1995) correcting for the potential selection bias.

I contribute to the existing literature by using the German Socio-Economic Panel (1992-2012)[1] to provide further evidence for the case of Germany and by explicitly taking selection into account. The remainder of this chapter is organized as follows. Section C.2 contains an overview of the existing literature. In the third chapter, the data and the sample are described. Descriptive statistics are presented in section C.4. Subsequently, the empirical strategy is explained and the results are discussed. The last chapter offers conclusions derived herein.

C.2 Literature Review

There is a broad body of literature on the wage penalty for motherhood (e.g., Kühhirt and Ludwig (2012), Napari (2010), Gangl and Ziefle (2009) or Budig and England (2001)) as well as on the re-entry behavior of mothers (e.g., Lalive et al. (2011), Burgess et al. (2008), Pylkkänen and Smith (2004) or Ondrich et al. (2003)). The research question of this study induces to focus on the literature about the career penalty in terms of occupational prestige.

One of the most often cited papers explicitly investigating the prestige penalty for motherhood is that of Aisenbrey et al. (2009). They analyze Sweden, the United States and Germany using Hazard Rate Models (piecewise constant exponential models as well as Cox proportional hazard models). In the case of the United States, they find a lower upward mobility but an increased downward mobility following time outs. In Sweden, their results reveal a lower upward

[1] For a description, see Wagner et al. (2007).

mobility if the career interruption is longer than 15 months. If career interruptions are long in Germany, mothers' careers are destabilized in general. The probability of a downward or upward move increases with longer leaves. Focusing on Sweden, Evertsson and Duvander (2011) investigate the effect of the duration of maternity leave on upward occupational mobility (Swedish Level of Living Survey (LNU), 1991-2000). First, they estimate the probability of taking a leave of more than 15 months using a probit model. As a second step, they apply a hazard model to obtain the rate of an upward occupational move after re-entering the labor market. In addition, a joint model controlling for unobserved time-constant heterogeneity is estimated to account for the selectivity into long or short leaves. Women taking longer leaves are found to be less likely to experience an upward occupational move.

In an event history framework, Jonsson and Mills (2001) examine the occupational mobility of Swedish mothers after they return to the labor market, but they only investigate the first "postbirth job". They discuss the role of the duration of absence from work in the occupational career of mothers, mentioning several reasons for career penalties for long time outs, namely *human capital depreciation*, *availability* and *social networks*.[2] The authors ask whether the duration of the absence has an influence on the occupational prestige upon return to the labor market. It apparently does for those who choose to take parental leave but not for those who choose to exit the labor market. Highly educated mothers have a higher chance of an upward move. The authors also mention that a causal effect might not be identified due to selection into longer or shorter time outs.

Kubis et al. (2009) ask whether the wages and the occupational prestige differ between mothers and childless women. They use the survey years 1992-2007 of the German Socio-Economic Panel and apply fixed effects models. They find differences between mothers and childless women for both indicators. Malo and Muñoz-Bullón (2008) use the British Household Panel Survey and investigate whether work interruptions affect mothers' subsequent careers using standard OLS regressions. In a first step, the authors do not control for endogeneity and find a negative effect of family-related time outs on occupational prestige. They argue that the jobs held prior to family-related interruptions might be chosen according to the expectations of those interruptions. This means that women who anticipate an interruption choose a lower-prestige job. Their results confirm this argument. A somewhat older paper considering data about British women (Social Change and Economic Life Initiative (SCELI)) is Jacobs (1997). She uses logistic regressions and finds that shorter career interruptions related to birth combined with

[2] For a description, see Jonsson and Mills (2001).

returning to full-time work are associated with a higher probability of keeping one's pre-birth occupational status. High education also is advantageous for maintaining the pre-birth status.

Several other papers also investigate questions similar to my research question, even though their focus often differs. One paper that is important to mention is Evertsson and Grunow (2012). They investigate career interruptions not only due to childbearing but also due to homemaking, unemployment or other reasons using Cox proportional hazard models. The authors compare women with continuous careers to those with discontinuous careers in Germany (German Life History Study (GLHS), West) and Sweden (LNU), two countries with rather different family policies. They do not find an effect of career interruptions on upward occupational mobility in Germany, whereas in Sweden, longer family breaks (accumulated) decrease the likelihood of an upward move. It should be noted that overall occupational mobility is higher in Sweden and almost all women work, so that a longer interruption might have a more negative effect than in Germany. In Germany, unemployment turns out to be a driver of downward occupational mobility. Judiesch and Jyness (1999) investigate the connection between leaves of absence and the career success of managers. Their data is from a financial services organization (1990–1995). They run hierarchical logistic regressions with promotion being the dependent variable. They find that leaves, irrespective of their reason and of gender, are linked to fewer promotions.

Grunow et al. (2011) investigate the reconciliation of career and family in Germany (GLHS West), USA (National Longitudinal Survey of Youth (NLSY)) and Sweden (LNU) distinguishing between different levels of education. They investigate the duration of career interruptions, whether mothers re-enter on the same prestige level as they had prior to the birth and whether they are able to keep this level in subsequent years. The authors use multivariate Cox models. In the case of Germany, they find that highly educated mothers perform better in respect to keeping their prestige level as compared to low or middle educated women. Ochsenfeld (2012) specifically considers management positions using the HIS Graduate Panel of 1997. He gives two explanations for women's underrepresentation in those occupations. The first is self-selection into specific paths of study and the second reason is that family obligations differ between genders. The author estimates Logit models as well as linear probability models with "having a managerial position" as the binary dependent variable. The estimations reveal that for women motherhood is associated with about half of the likelihood of obtaining such a position ten years after graduation. There is no such effect of fatherhood found for men. McIntosh et al. (2012) investigate the career progression of nurses in particular, a rather femaledominated occupation. They find career interruptions of more than two years to be detrimental for mothers' career progress.

Dex et al. (2008) analyze whether the likelihood of a downward occupational move has changed over time. The authors use logistic regressions. The results concerning the job after a career interruption indicate that downward occupational mobility has decreased over the cohorts that are subject to the analysis if mothers return to full-time employment but has increased if mothers return to part-time work. Taking longer career breaks is associated with an increasing downward occupational mobility over time as well. The authors also look at the occupational mobility after mothers have already returned to the labor market. The youngest cohort under examination experienced the highest probability of an upward move upon return. However, downward mobility did not decline over the cohorts.

Kahn et al. (2014) examine wages, labor force participation and occupational status. They ask whether motherhood penalties change over mothers' life-course. They use the National Longitudinal Survey of Young Women and take into account the age range of the 20s through the 50s. Using fixed effects models, they find that motherhood is penalized in respect to the three mentioned outcomes but that this penalty diminishes in the 30s and 40s.

Downward occupational mobility in Spain is investigated in a study written by Gutierrez-Domenech (2002) who focuses on the re-entry behavior of Spanish mothers. The lack of downward mobility in Spain is explained by the fact that part-time jobs are rare in Spain and therefore no significant downward mobility can be observed for those mothers who return to the labor market rather than remaining non-employed. It is worth mentioning that the author uses a rather rudimentary definition of occupational mobility.

Granqvist and Persson (2005) explicitly ask about the different occupational mobility of men and women in Sweden (LNU). They also conduct event history analyses. The authors make use of a piecewise constant exponential hazard model. They do find differences between men and women. Men are about twice as likely to get a better job as women. The effect is not as clear if the employees have more than 12 years of education. Granqvist and Persson (2005) also argue that the difference can partly to be attributed to home-related obligations, which lower women's chances on the labor market. Interestingly, parental leave does not seem to have any effect. The authors explain this by the fact that leave taking is common in Sweden and therefore it does not signal anything to a (potential) employer. They conclude that the negative effect of children reflects the negative effect of part-time work.

Mandel and Semyonov (2006) ask about the role of the welfare state when it comes to women's employment participation and occupational positions. They analyze 22 industrialized countries.

Their results show that mothers in countries with progressive welfare systems are underrepresented in managerial positions.

This review reveals that, overall, birth-related career interruptions seem to destabilize mothers' careers. But negative effects of time outs probably diminish with age, exist irrespective of whether the interruption is family-related and depend on the level of education of the women. The selectivity into certain occupations and into long or short leaves also seems to be relevant.

The contribution of this paper is that it provides further evidence for possible career penalties for career interruptions due to childbirth in Germany. To the best of my knowledge, there has not been another study investigating this exact question using this data and this method. Most studies concerning Germany concentrate on the German Life History Study (e.g., Aisenbrey et al. (2009), Evertsson and Grunow (2012)). The study by Ochsenfeld (2012) uses the HIS Graduate Panel of 1997. I first investigate the probability of observing positive SIOPS information for mothers, i.e., the probability of participating in the labor market after the first childbirth. In the second step, I investigate the effect of leave duration on the prestige level as well as on the probability of experiencing an upward or downward occupational move compared to the SIOPS level prior the first birth.

C.3 Data

I use the German Socio-Economic Panel.[3] For the purpose of this paper, this dataset is more favorable than other datasets, e.g., the German Life History Study (GLHS[4]). The main reason is that the GSOEP is not restricted to several birth cohorts. Therefore I do not only have information until the birth cohort of 1971 as in the GLHS, but rather until to the late 1980s. The HIS Graduate Panel only contains people with a degree qualifying for a profession (e.g., Ochsenfeld (2012)). In addition, the graduates are only interviewed three times tops up to ten years after graduation. The GSOEP does not have these restrictions.

I include the survey years 1992 - 2012. The sample is restricted to women at age 20 to 59. Even in those cases where men take parental leave, they only take a small share of the maximum of 14 months (e.g., Federal Statistical Office of Germany (2013)). Teenage mothers are excluded since they are considered special cases. New mothers are included as long as they are at working age. If they are older than 59 years, they are dropped. The sample contains women between the ages of 20 and 59 between 1992 and 2012 and who gave birth to their first child within this

[3] For a description of this very rich panel dataset, see Wagner et al. (2007).
[4] E.g., Mayer (2008).

observation period. I do not observe all births for all mothers and not all returns to the labor market. The sample is further restricted to those for whom I observe the first return to the labor market.

I want to show effects of the duration of the first time out from the labor market mothers take in connection with their first childbirth. For time out duration, I use the calendar information of the GSOEP. These provide information about leave behavior as well as about housewife and unemployment periods. The information is on a monthly basis. The leave variable can be defined in two different ways. One definition contains maternity leave periods only. The other also considers housewife periods, unemployment and simply “not being employed” as time out. I focus on this more comprehensive definition. It is important to note that the time out has to have started in connection with the first childbirth. I later control for the presence of subsequent career interruptions that might occur, e.g., due to a second childbirth.

In the estimations, I condition on whether a mother has reported to be employed at an interview that took place within 12 months prior to the first birth. I include single mothers as well as mothers living in couples. I exclude mothers if they are in short-time work, in vocational training, in a sheltered workshop or near retirement. For some mothers, no interview is available for the year they gave birth to their first child. I have no information about any leave or its duration for this group, so they are not included in the sample. I also drop those for whom I have no information about the year of birth of the first child in the biography data. I do not keep mothers for whom there is a gap during the period of the time out because I cannot identify changes in labor status.

The SIOPS scale is based on the ISCO88 categories of occupations (Ganzeboom and Treiman (1996)). A first version, developed by Treiman (1977), was based on ISCO68. He transformed national prestige scales into one standardized scale, the Standard International Occupational Prestige Scale (SIOPS) (Ganzeboom and Treiman (1996)). This scale only contains information about the social prestige of an occupation, not about the socio-economic factors education and income as does the ISEI index (e.g., Ganzeboom and Treiman (1996)).

The sample consists of 1,588 women who became mothers during the years 1992 through 2012 and for whom I observe the information for the year of the first birth; 1,088 of them have been working the year prior to the first birth. Of those, 1,087 took some leave and 838 re-entered the labor market at some point during the observation period. This is the estimation sample. Although this is not an extremely large sample, the number still is high enough to perform reasonable estimations.[5]

[5] If childless women are included, 806 of 2,004 interrupt their career and for 294 the return is observed.

C.4 Descriptive Analysis

1,376 of the 1,588 observed new mothers were living with a partner in the year of the first birth. Only a small share did not live with a partner. Some mothers had a partner who was not living in the same household. It is worth mentioning that I do not control for whether the partner living in the same household is the father or not. In the estimations, I control for the number of children in different age groups living in the household. This is reasonable since further children may yield further employment interruptions or postponement of re-entry into the labor market. The women in the sample have up to six children.
An upward occupational move is considered to be an upward change of at least 10% in respect to occupational prestige prior to first birth. Accordingly, a downward move is considered a downward change of at least 10%. A 10% change in occupational prestige does not necessarily mean the same change in terms of "scale points" for all mothers. This is done for the same reason as mentioned by Aisenbrey et al. (2009): A change of one point does not really imply a change in prestige on a scale that is so strongly subdivided. Using a 10% change is common in the literature (e.g., Aisenbrey et al. (2009), Evertsson and Grunow (2012)). An alternative is to use a change of five scale points, as done in Jonsson and Mills (2001). One can use different points in time as a reference to define occupational mobility. The first definition is the mobility within one year, given the mother participates again. The second definition takes the year prior to the first birth as a reference point, and the third definition takes the SIOPS value of the first job after the time out as a reference. Another interesting point is whether mothers manage to keep a once gained upward move or not. I focus on the reference year "prior to first birth" and investigate the probability of achieving a more resp. less prestigious job upon return to the labor market and of further mobility of the same direction.

The mean SIOPS of mothers amounts to 46 in the year prior to their first birth for those who work the year before (table C-1). If the total time out after first birth does not exceed 12 months, the SIOPS in the year of the return on average amounts to 47. This group consists of 196 mothers. This is an initial indicator that those mothers returning quickly manage to keep their previous position. This result is not surprising since one would expect that those mothers who return soon after childbirth are those with a high labor market commitment and that those are the women who might have focused on building a career in the first place. Those who return in the second year after childbirth (time out of 13 – 24 months, 297 mothers) re-enter on average to a job with 45 points on the SIOPS scale. The 112 mothers returning between 25 and 36 months after birth reach an average of 44 and those re-entering even later reach an average of 42 points. If the first leave is longer than 72 months, the SIOPS at re-entry amounts to 39. When looking at the SIOPS information prior to the first birth, it turns out that this value tends to be lower for those who take longer time outs. This indicates self-selection into certain occupations, which is also mentioned in some part of the literature (e.g., Evertsson and Duvander (2011), Jonsson and Mills (2001)).

Table C-1: Basic SIOPS information for mothers working the year before first birth

		SIOPS scale	
group	**# mothers**	**prior birth**	**at return**
working, year of birth – 1	1088	46	-
time out ≤ 12 months	196	49	47
12 < time out ≤ 24 months	297	46	45
24 < time out ≤ 36 months	112	46	45
time out > 36	233	44	42
time out > 72	62	42	39
for comparison:	**# women**	**SIOPS scale**	
childless, no interruption, mean	3182	46	
…, interruption, now working, mean	1672	44	

Source: GSOEP, 1992-2012, own calculations

Table C-2 shows the occupational mobility of mothers and childless women within the timeframe between two interviews, which is approximately one year, as well as the occupational mobility in respect to the SIOPS value reported in the year prior to the first birth. It is worth mentioning that for childless women with an employment interruption, only the years they reported being employed are used in Table C-2. It turns out that about 17% of mothers re-entering within one year after first birth experience an upward move compared to their SIOPS before the first birth. About 22% experience a downward move. The proportion of those experiencing an upward move is smaller than the share of those experiencing a downward move in the group that re-enters within the first year. The upward as well as the downward mobility of childless women who interrupt their employment is somewhat higher than that of new mothers who return to work within one year, noting that the difference is more distinct for upward mobility. It is interesting to see that among childless women who do not interrupt their working career, only 14% experience an upward move, while 13% experience a downward move within one year. Twenty-five percent (23%) of the childless women who do interrupt their career experience an upward (downward) move. Taking these two groups together reveals that only about 18% (16%) experience an upward (downward) move during the observed period. If mothers re-enter within the second year after first birth, about 19% manage to make an upward occupational move, 23% undergo a downward move compared to the prestige prior to first birth. Of those re-entering in the third year, 23% (26%) move up (down) compared to their pre-birth level. Twenty-one percent of the mothers returning after the oldest child is in school, namely after more than six years, experience an upward move, while 36% of this group move down the prestige scale. At first glance, it appears that mobility is higher for mothers, irrespective of time out duration, than for childless women without any interruptions. The same is true for women who interrupt due to other reasons. Further, mothers whose time out exceeds six years show a higher downward mobility and a lower upward mobility than childless women with interruptions. It seems that interruptions, irrespective of their reason, seem to destabilize women's careers. But those who became mothers experience an upward move less often than do those who have interrupted their career due to other reasons.

Table C-2: Occupational mobility and duration of time out

		mobility within one year				**mobility compared to SIOPS prior first birth**			
group	**mothers**	**↑**	**%**	**↓**	**%**	**↑**	**%**	**↓**	**%**
timeout ≤ 12	196	36	18.4	40	20.4	34	17.3	43	21.9
12 < timeout ≤ 24	297	50	16.8	52	17.5	57	19.2	67	22.6
24 < timeout ≤ 36	112	28	25.0	21	18.8	26	23.2	29	25.9
timeout > 72	62	13	21.0	11	17.7	13	21.0	22	35.5
for comparison:	**women**	**↑**	**%**	**↓**	**%**				
childless, no break	3182	451	14.2	406	12.8	n.a.		n.a.	
childless, break	1672	419	25.1	384	23.0	n.a.		n.a.	

Source: GSOEP, 1992-2012, own calculations

When investigating occupational prestige it is relevant to consider whether mothers return to part-time or full-time employment. Therefore, table C-3 shows some information about the covariates used in the estimations. Mothers are, on average, 28 years old when their first child is born. When they return, they are on average 31 years old. Almost 75% of all new mothers return to the labor market before their second child is born. Another 15% return after their second child is born. Only approximately 1% returns after the third or higher order birth. Approximately 23% live in East Germany. About 83% live with a partner the year before birth; when returning, this share amounts to almost 94%. In the year prior to their first birth, about 87% of new mothers are working full-time, 8% are working part-time and only 2% are marginally employed. When returning, only 21% are working full-time while 55% have returned to part-time work. Another 24% return to marginal employment. This offers a preliminary insight to the re-entry behavior of German mothers since the largest share of mothers return to part-time work instead of full-time work.

Table C-3: sample means – year before first birth, year of the first return to the labor market

	Year before first birth[6]			**Year of reentry**		
variable	**mean**	**sd**	**N**	**mean**	**sd**	**N**
age (year of first birth)	28.601	4.031	741	31.449	4.368	741
years of education	12.987	2.580	707	13.082	2.601	717
return after first birth	0.743	0.437	731	0.742	0.438	741
… second birth	0.150	0.358	731	0.150	0.357	741
… third or higher order birth	0.008	0.090	731	0.009	0.097	741
East Germany	0.233	0.423	731	0.235	0.424	741
partner	0.825	0.381	713	0.936	0.244	730
working full time	0.874	0.332	731	0.212	0.409	741
working part time	0.077	0.266	731	0.548	0.498	741
marginal employment	0.019	0.137	731	0.240	0.428	741
children 0-3	0.000	0.000	727	0.819	0.444	737
children 4-6	0.003	0.052	727	0.273	0.503	737
children 7-12	0.004	0.064	727	0.080	0.304	737
children 13-16	0.001	0.037	727	0.011	0.103	737
experience full time	6.207	4.109	731	7.264	4.235	741
experience part time	0.707	1.653	731	1.320	1.846	741
multigenerational household	0.003	0.052	731	0.009	0.097	741

Source: GSOEP, 1992-2012, own calculations

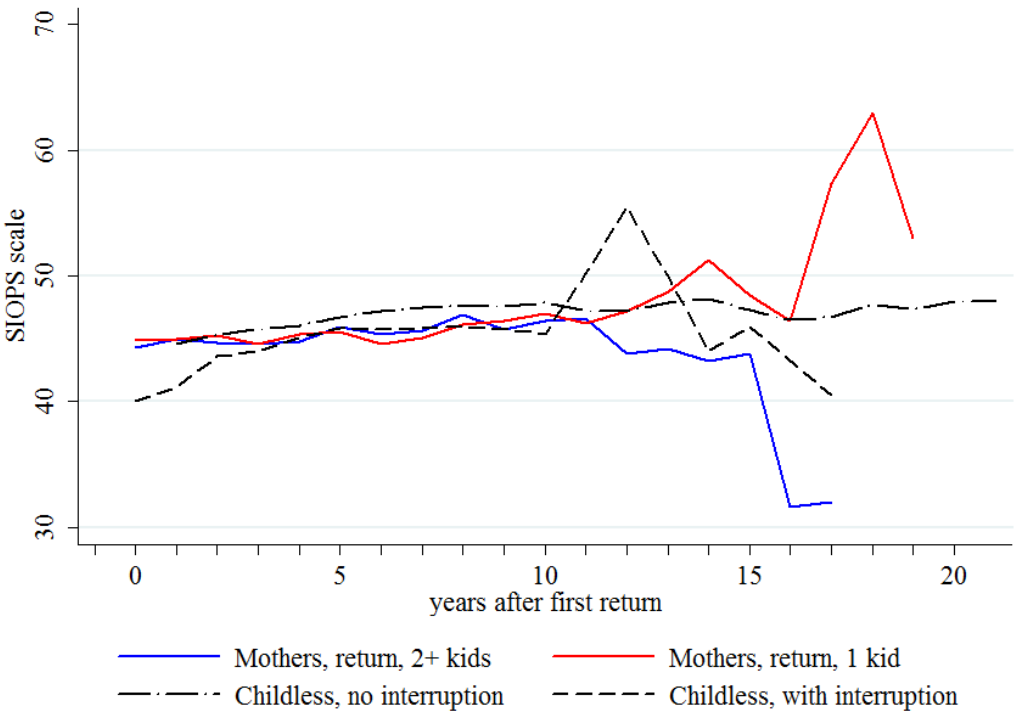
There are no children in the youngest age group (age 0-3) living in the household the year before first birth. But for older age groups, the means are not zero. These children therefore are not biological children of the interviewed woman[7]. Only 0.3% of women are living in a multigenerational household the year prior first birth. Though this share has tripled at return, it still only amounts to approximately 1%. Before birth, women on average have gained about 6 years of full-time labor market experience and about 0.7 years of part-time experience. Since

[6] This is the calendar year prior the first birth, not the last 12 months.
[7] The variable that identifies the mother-child relationship refers to "social mothers", who, in rare cases might not the biological mother (e.g., SOEP documentation).

these values refer to the year prior first birth, women might gain a view more months of experience until they start their time out due to childbirth. In addition, they might already have returned several months before being interviewed.

Figure C-1 shows the SIOPS of mothers during the years after the first re-entry. I am able to analyze the occupational mobility of mothers after they have returned to the labor market for the first time. It seems that when mothers return to the labor market, they do not show a considerably lower SIOPS level than childless women without any interruptions. However, up to 10 years after return, mothers show a slightly lower level than do childless women.[8] From this time on, mothers with one child show an increase while mothers of two or more children fail to keep pace with their childless counterparts.

Figure C-1: SIOPS after return to the labor market[9]



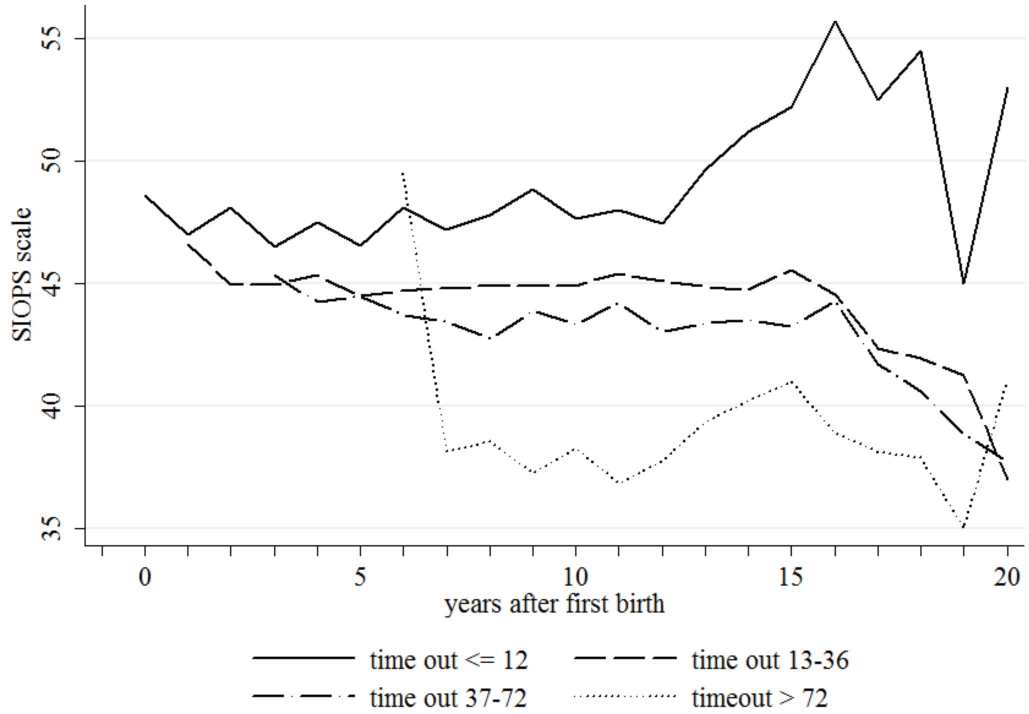
Source: GSOEP, 1992-2012, own calculations

Overall, mothers show more occupational mobility, upward as well as downward, while childless women without any interruptions show virtually no change over time. It is important to note that the group sizes for mothers are quite small after about 10 years after return to the labor market due to such factors as panel attrition. For example, the peak that shows up 18 years after first return for mothers of one child only represents two mothers.[10] For childless women

[8] Note that fertility is not completed in all cases. The biography information is only available up to the latest update.
[9] The proportion of mothers who are working can be found in the appendix (figure A C-1).
[10] The number of women represented in the graphs is shown in the appendix (table A C-2).

who interrupt their working career, there appears a greater occupational mobility after return as well.

Figure C-2: SIOPS according to maternity leave length



Source: GSOEP, 1992-2012, own calculations

A comparison of the SIOPS path according to different leave lengths is shown in Figure C-2. There is an obvious difference between those returning within one year and those taking a longer time out. In addition, a further difference becomes apparent between 13-36 months or 37-72 months. Time outs lasting longer are associated with an even lower SIOPS value. Of course these are selective groups since longer leave is associated with lower labor market commitment and these women are expected to be less career-oriented than women who quickly return to the labor market. They therefore are expected to already show a lower SIOPS level prior to first birth. After about 16 years following the first birth, the SIOPS value of the groups 13-36 months and 37-72 months declines further. Again, the group sizes are small, so that this figure should not be overly interpreted as revealing a career penalty due to leave taking.[11] But is there a prestige penalty for longer time outs or are the displayed differences attributed to other factors, such as self-selection into lower prestigious jobs with the intention of taking a long time out after childbirth?

[11] The numbers are again presented in the appendix (table A C-3).

Overall, this first analysis reveals what is expected. Interruptions destabilize women's careers. It seems as if the career interruption itself, not the reason for the break (i.e., leave due to childbirth or any other reason), is the main driver of some sort of career penalty in terms of lower prestige.

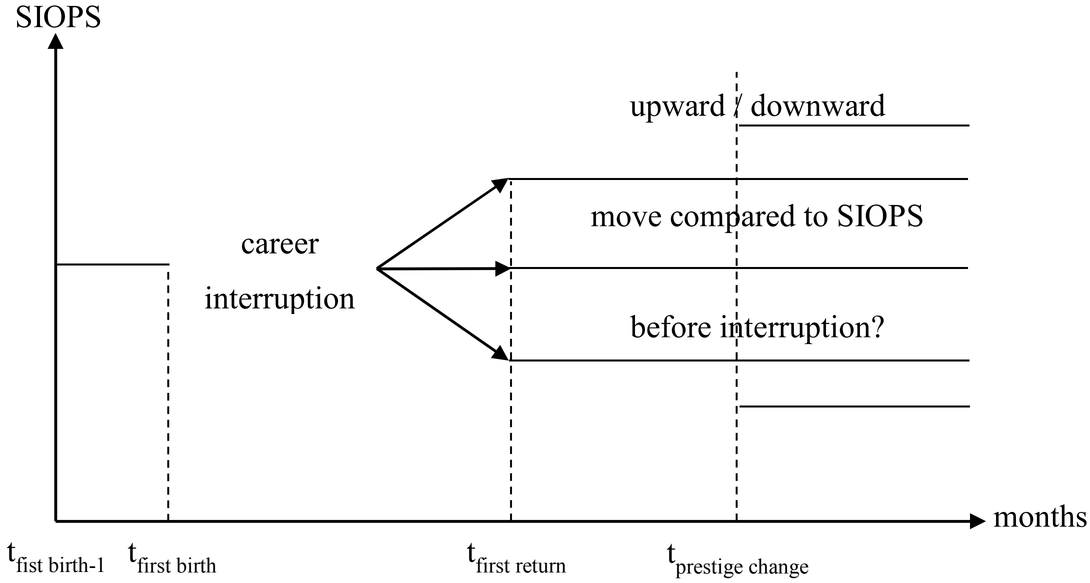
C.5 Empirical methods and results

C.5.1 Methods

The present study seeks to identify the effect of the first birth-related career interruption, and especially of its length, on the later career of women in respect to occupational prestige. This prestige penalty might be expected due to several reasons. Employers might interpret a longer career interruption as a lower commitment to the labor market (Evertsson and Duvander (2011)). Furthermore, during "inactive" times, no human capital accumulation takes place. It most likely even depreciates (e.g., Jonsson and Mills (2001), Evertsson and Duvander (2011)). Therefore, I expect a negative effect of time out duration on occupational prestige. In case of mobility, an increasing probability of a downward move is expected for those taking longer time outs.

The research question raises several methodological issues. First, the prestige information, measured by the SIOPS scale, is only observed for those who are working. The decision of whether and when to re-enter the labor market after childbirth inherits the problem of selection. I address this by using a two-step model according to Heckman (Heckman (1976), Heckman (1979)). Figure C-3 describes the relevant points during the observation period of each mother. Upward resp. downward mobility is defined as follows. The dummy takes on the value 1 if the mother reaches a prestige level that is at least 10% above the level of the position she held before her first birth for the first time upon re-entry. Afterwards it is zero, unless she reaches a further upward move of at least 10%. The same definition is used for downward moves. I therefore not only investigate the prestige level according to time out duration but also the mobility compared to the level before the first birth as well as subsequent rises resp. declines. By defining mobility in this way, I investigate one dimension of occupational mobility after childbirth. There are many further dimensions that are not in the focus of this chapter. Another dimension is, for instance, mobility upon return, taking the SIOPS at re-entry as a reference. A further dimension would investigate whether mothers achieve and keep a higher prestige or whether they drop to and stay at a lower prestige level relative to their situation prior the birth of their first child. In this case, the aforementioned dummy would take on the value 1 in each year the actual level is at least 10% higher resp. lower than the level before the first birth. The last dimension to be mentioned is the mobility within one year unrelated to the SIOPS held before first birth. A detailed investigation of all of the different kinds of occupational mobility in this context is beyond the scope of this analysis.

Figure C-3: relevant points in time



Source: Evertsson and Grunow (2012) and Evertsson and Duvander (2011) use a similar illustration.

I am interested in women who have been working prior to the birth of their first child. After childbirth, they take leave. In some cases this might exceed the duration of maternity leave entitlement. In such cases, they are not truly in maternity leave any longer but are considered housewives. The duration until mothers re-enter the labor market is measured in months. The participation decision is represented by the selection equation. I only observe SIOPS information for those mothers who actually re-entered. Of course then, an upward/downward move can only be observed for this group. Since this selection is on observables as well as unobservables, a selection model is needed to estimate the probability of participation. Given that a mother is working, I estimate the second step. This uses as a dependent variable the occupational prestige while the duration of the time out is used as explanatory. The duration is included as several dummies, each representing an interval of one year. The reference category is an interruption of no more than 12 months. Thereby I get the effect of leave duration on the level of occupational prestige compared to a very short interruption. In addition, the second step is also estimated as a binary outcome model. I estimate the effect of leave duration on the probability of achieving an upward occupational move resp. the risk of facing a downward move.

The basic method according to Heckman (1979) can be described as follows (see also Greene (2008), Cameron and Trivedi (2005)).[12] The first step contains the selection mechanism:

$$z_i^* = \omega_i{}'\gamma + u_i \text{ with } z_i = \begin{cases} 1 \; if \; z_i^* > 0 \\ 0 \; if \; z_i^* \leq 0 \end{cases} \tag{C-1}$$

Thus: $Prob(z_i = 1|\omega_i) = \Phi(\omega_i{}'\gamma)$.

The regression model of interest is

$$y_i = x_i{}'\beta + \varepsilon_i \tag{C-2}$$

noting that y_i is observed only if $z_i = 1$. It is assumed that the errors u_i and ε_i follow a bivariate normal distribution (Greene (2008), Cameron and Trivedi (2005)):

$$\begin{bmatrix} u_i \\ \varepsilon_i \end{bmatrix} \sim N \left[\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} 1 & \sigma_{12} \\ \sigma_{12} & \sigma_2^2 \end{bmatrix} \right] \tag{C-3}$$

Simply estimating the model of interest (C-2) via OLS leads to inconsistent estimates if the errors are not uncorrelated. Therefore, the estimation equation has to be corrected for selection, since

$$E(y_i|z_i = 1, x_i, \omega_i) = x_i{}'\beta + E(\varepsilon_i|u_i \geq -\omega_i{}'\gamma) \tag{C-4}$$

$$E(y_i|z_i = 1, x_i, \omega_i) = x_i{}'\beta + \sigma_{12}\lambda(\omega_i{}'\gamma) \tag{C-5}$$

which only is equal to $x_i{}'\beta$ if the errors are uncorrelated. To correct the equation of interest for selection, the following procedure is performed (e.g., Cameron and Trivedi (2005)). In the first step, a probit model is estimated where z_i is regressed on w_i to obtain $\hat{\gamma}$. $\lambda(\omega_i{}'\hat{\gamma})$ is the estimated inverse Mill's ratio (IMR) which is included as an additional regressor in the second step equation. The IMR can be written as $\lambda(\omega_i{}'\hat{\gamma}) = \phi(\omega_i{}'\hat{\gamma})/\Phi(\omega_i{}'\hat{\gamma})$, with $\phi(.)$ being the standard normal density function (pdf) and $\Phi(.)$ being the standard normal distribution function (cdf) (e.g., Cameron and Trivedi (2005)). The resulting equation is estimated using OLS (Cameron and Trivedi (2005)):

$$y_i = x_i{}'\beta + \sigma_{12}\lambda(\omega_i'\hat{\gamma}) + v_i \tag{C-6}$$

The arising problem is that the standard errors are not correct (e.g., Cameron and Trivedi (2005)). First, the errors from equation (C-6) are heteroskedastic. Second, $\hat{\gamma}$ is an estimate. Therefore the standard errors must be corrected. When correcting for sample selection manually, I use bootstrapped standard errors as suggested in Cameron and Trivedi (2005). The

[12] The notation is adapted to be consistent here.

present study not only uses the level of occupational prestige as the dependent variable of the second step, it also uses a dummy indicating whether there has been an upward/downward move or not. Therefore, the second step needs to be a probit model as well. The first step is represented by the selection equation presented above (van den Ven and van Praag (1981), Heckman (1979), Greene (2008)). The second step no longer is as simple as in equation (C-6) since the dependent variable in this case is either equal to one (if an upward/downward move occurred) or equal to zero (if there was no change in prestige or a move in the other direction). This is considered by the use of the Stata-implemented model of Van den Ven and van Praag (1981), which applies Maximum Likelihood estimation.[13]

There should be at least one variable that is included in the selection equation but not in the outcome equation (exclusion restriction) (e.g., Cameron and Trivedi (2005)). The GSOEP provides a promising variable. It is whether starting a new job is possible immediately. If it is possible to start a new job immediately, one is more likely to return to the labor market. I include this variable lagged by one year. This variable is included in the selection equation but not in the outcome equation. I also include as an exclusion restriction the female unemployment rate of the federal state of residence. These values are available from the Federal Employment Agency (2013), but only from the year 1995 onwards, so I lose three waves of my observation period in the estimation. In addition, a dummy indicating whether the mother has been working the last year is included in the first but not the second step.

The estimated equations for the SIOPS level are described as follows. The selection equation estimates the probability that a positive SIOPS value is measured after the first birth. Of course, this is the case only if the mother is working. The vector w_i from equation (C-1) contains the variables according to Table A C-1 in the appendix. The second step of the Heckman model estimates the effect of time out duration on the SIOPS level that mothers achieve during their employment career after their first birth. The variables used in this outcome equation are listed in table A C-1 as well.[14]

To account for the panel structure, an approach suggested by Wooldridge (1995) is applied. First, for each period t, a standard probit model is estimated:[15]

$$P(s_{it} = 1|\omega_i) = \Phi(\omega_i \gamma_t) \qquad \text{(C-7)}$$

[13] For a further description and the detailed derivation of the model see van den Ven and van Praag (1981).

[14] I apply Stata-implemented routines and also replicate the results by manually correcting for sample selection.

[15] As above, the notation is adapted to be consistent.

Using this, the inverse Mill's ratio is obtained for each period. Then a pooled OLS regression is run using the selected sample (Wooldridge (1995, 2010))[16]:

$$y_{it} = x_{it}\beta + \rho_1 d1_t \hat{\lambda}_{it} + \cdots + \rho_t dT_t \hat{\lambda}_{it} \text{ for all } s_{it} = 1 \quad \text{(C-8)}$$

with $d1_t$ through dT_t being year dummies. I bootstrap the standard errors here as well. The vector ω_i contains variables that are also included in x_{it}. I also include the exclusion restrictions discussed above in ω_i but not in x_{it}.

C.5.2 Results

The results are presented in tables C-4 through C-8. The results are obtained using the comprehensive definition of a career interruption. If only considering maternity leave duration, it is possible that mothers still do not work after leave has ended. This is a rather restrictive definition of an interruption, which is why I focus on the results derived using the comprehensive definition. Since the probit coefficients only contain information about sign and significance (e.g., Wooldridge (2010)), the average marginal effects are presented in the appendix.

C.5.2.1 SIOPS level

It appears that the time out duration is relevant in respect to the SIOPS level. The coefficients of time outs of 13-24, 25-36, 37-48 and 49-60 months are negative but insignificant. However, leaves of 61-72 months are associated with almost 7 points less on the SIOPS scale and those longer than 72 months are associated with more than 10 points less compared to the reference of up to 12 months (table C-4). The two categories are highly significant. Of course one must bear in mind that the groups exhibiting short resp. long interruptions are certainly selective. Women who take long interruptions probably select into less prestigious jobs in the first place since they anticipate their long interruption, e.g., because they have distinct preferences for staying with their children.[17]

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[16] I also estimate an alternative specification with added time averages of appropriate explanatory variables to specify individual effects (Wooldridge (2010), Dustman and Rochina-Barrachina (2007)). The results are comparable.

[17] However, if controlling for the SIOPS level held before the first birth, the results still reveal remarkable differences between different time out durations, even though the effects are not this strong. Since the mobility relative to the SIOPS score before the time out is investigated in particular later on, the results are not presented here.

Table C-4: Heckman estimation – SIOPS level

	SIOPS			selection		
time out 13 - 24	-0.561		(0.916)	-		
… 25 - 36	-0.333		(1.271)	-		
… 37 - 48	-1.925		(1.238)	-		
… 49 - 60	-2.150		(1.735)	-		
… 61 - 72	-6.903	***	(1.990)	-		
… > 72	-10.284	***	(2.102)	-		
age	1.555	***	(0.545)	0.100	*	(0.054)
age squared	-0.015	**	(0.007)	-0.002	***	(0.001)
years of education	2.197	***	(0.206)	0.044	***	(0.012)
return after second birth	1.848		(1.315)	-0.499	***	(0.062)
return after third+ birth	1.037		(3.114)	-0.530	***	(0.195)
East-Germany	0.654		(0.861)	0.045		(0.102)
partner	2.108	*	(1.092)	0.078		(0.085)
# children 0-3	-0.210		(0.882)	-0.834	***	(0.061)
# children 4-6	0.039		(0.641)	0.150	***	(0.052)
# children 7-12	0.493		(0.628)	0.178	***	(0.060)
# children 13-16	2.073	***	(0.751)	0.189	**	(0.088)
experience full time	-0.322	**	(0.146)	0.051	***	(0.009)
experience part time	-0.402	**	(0.183)	0.144	***	(0.018)
full time	3.314	***	(0.913)	-		
part time	4.664	***	(0.739)	-		
further interruptions	-2.217	***	(0.755)	-0.572	***	(0.050)
multigenerational	0.670		(1.563)	-0.306	*	(0.182)
not working (t-1)	-		-	-0.755	***	(0.056)
start new job now (t-1)	-		-	0.172	**	(0.081)
unemployment rate (f)	-		-	0.015		(0.010)
constant	-17.676	*	(9.306)	-0.941		(0.893)
λ			0.462			
se (λ)			0.823			
χ^2 (for comparison)			0.313			
Prob > χ^2			0.576			
observations			6,999			
clusters			817			
Wald χ^2			534.66			
Prob > χ^2			0.000			

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

cluster-robust standard errors in parantheses

Age is associated with a higher SIOPS level. Years of education show a clear and positive coefficient. Living in East Germany is not significant but positive. The same is true for the dummy *living with a partner* (weakly significant). The SIOPS level seems to be higher if there are more children at age 13-16 living in the household. The other age groups do not show significant coefficients. Both full-time and part-time experience have a negative coefficient, both being significant at the 5% level. Both full-time and part-time employment are related to a higher SIOPS. This is in line with the expectations. The dummy indicating the occurrence of further employment interruptions is highly significant and negative. Living in a multigenerational household shows a positive but insignificant coefficient. Correcting for selection does not seem to be necessary in this case ($\chi^2 = 0.313$). Therefore, I also perform a simple OLS estimation, which results in the same implications (table A C-7).

Table C-5: Results Panel estimation – SIOPS level

	SIOPS		
time out 13 - 24	-0.511		(0.926)
… 25 - 36	-0.267		(1.272)
… 37 - 48	-1.882		(1.276)
… 49 - 60	-2.106		(1.677)
… 61 - 72	-6.976	***	(1.986)
… > 72	-10.316	***	(2.209)
Inverse Mill's ratio …1995	-5.795		(6.429)
1996	-5.130		(3.836)
1997	2.830		(2.512)
1998	-1.569		(2.236)
1999	3.195		(2.480)
2000	-3.122	*	(1.772)
2001	-0.316		(1.693)
2002	-1.270		(1.665)
2003	-1.717		(1.336)
2004	-1.359		(1.631)
2005	-0.662		(1.385)
2006	0.408		(1.322)
2007	0.661		(1.482)
2008	0.596		(1.297)
2009	0.336		(1.702)
2010	-0.534		(1.554)
2011	-3.498	*	(2.073)
2012	-0.765		(1.628)
observations	4,459		
adjusted R squared	0.353		
χ^2 (IMR's)	17.62		
Prob > χ^2	0.4811		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

bootstrapped standard errors in parantheses

Applying the strategy suggested by Wooldridge (1995) results in comparable implications concerning the influence of the duration of the first birth-related career interruption on the SIOPS level. The results are presented in table C-5. A duration of 61-72 months or more than six years shows about the same influence on the SIOPS value as the basic Heckman procedure. The results concerning the control variables also are the same.

The inverse Mill's ratios (interacted with the year dummies as described in the last section) do not show any obvious pattern. Some are positive, some are negative, but all except two are insignificant. Testing the joint significance of the inverse Mill's ratio does not reject the null of no significance.

C.5.2.2 Occupational mobility

Not only the level itself but also the probability of an upward move is associated with time outs (table C-6). The categories 13-24, 25-36, 61-72 and >72 months are not relevant for the probability of an upward move. A time out of 37-48 or 49-60 months is (weakly) significantly increasing the probability of an upward move, compared to a time out of up to 12 months. In table A C-4, the average marginal effects are presented. They amount to 1.4 resp. 2.5 percentage points, compared to time outs of up to 12 months. Since the overall occurrence of an event is scarce, this is a "large effect" when comparing the two groups. However, the effect is not considered to be large overall. Therefore, longer time outs do not reduce the probability of an upward change in comparison to short interruptions. The probability of a downward move is more distinctly associated with time out duration, only 13-24 and 25-36 months are not related to downward moves. All other categories are significant and positive. The strongest influence is found for time outs lasting more than six years. The difference between those interrupting their career for more than 72 months and those who interrupt up to 12 months amounts to 7.5 percentage points (table A C-4). It must be mentioned that the extent of these results should be interpreted with caution due to the fact that upward resp. downward moves are rather rare events if defined as described above, resulting in only few positive values for the dummy of interest. Since the exclusions restriction of being able to start a new job immediately includes some missing values, some of the positive outcomes are dropped in the estimation. If this exclusion restriction is not included in the selection equation, the coefficient of time outs longer than 72 months is weakly significant and positive in the model of upward moves as well, which is more in line with the basic probit estimation for the selected sample. The results concerning downward moves are not influenced by this to a noteworthy extent. Overall, long leaves are associated with higher occupational mobility, both upward and downward, as is in line with the results found by Aisenbrey et al. (2009).

Education does not show any significant coefficient. The coefficient of the number of children in the youngest age group is negative and weakly significant. The opposite is true for children at age 4-6. For downward mobility, only the age group 7-12 years shows a weak negative coefficient. Living with a partner is weakly significantly associated with upward moves but is not associated with downward moves. Full- and part-time experience are not significant in respect to upward or downward mobility. Full-time employment is related to less occupational mobility, with only being highly significant for downward moves. The same is true for part-time work. Further interruptions and living in a multigenerational household are not associated with occupational mobility. Examining the need of the selection correction reveals that this correction seems to be needed in the case of occupational mobility. For comparison, the results of simple probit estimations are presented in the appendix.

Table C-6: Heckman Specification – mobility

	upward			downward		
time out 13 - 24	0.021		(0.096)	0.002		(0.095)
… 25 - 36	0.096		(0.117)	0.117		(0.114)
… 37 - 48	0.209	*	(0.126)	0.415	***	(0.131)
… 49 - 60	0.366	**	(0.173)	0.428	**	(0.178)
… 61 - 72	0.222		(0.196)	0.596	***	(0.189)
… > 72	0.355		(0.227)	0.751	***	(0.196)
age	-0.007		(0.068)	0.021		(0.066)
age squared	-0.000		(0.001)	-0.001		(0.001)
years of education	-0.006		(0.020)	0.020		(0.018)
return after second birth	-0.210		(0.132)	-0.361	***	(0.128)
return after third+ birth	-0.584		(0.417)	-0.289		(0.243)
East-Germany	0.089		(0.088)	-0.128		(0.093)
partner	-0.209	*	(0.108)	0.011		(0.112)
# children 0-3	-0.183	*	(0.109)	-0.029		(0.095)
# children 4-6	0.144	*	(0.086)	0.073		(0.074)
# children 7-12	-0.109		(0.088)	-0.139	*	(0.078)
# children 13-16	-0.036		(0.143)	-0.164		(0.134)
experience full time	0.014		(0.014)	-0.000		(0.013)
experience part time	-0.021		(0.018)	-0.029		(0.021)
full time	-0.205	*	(0.114)	-0.444	***	(0.103)
part time	-0.198	**	(0.094)	-0.462	***	(0.076)
further interruptions	-0.020		(0.074)	-0.075		(0.068)
multigenerational	-0.053		(0.285)	-0.478		(0.396)
constant	-1.161		(1.205)	-1.441		(1.127)
rho	0.579			0.689		
χ^2 ($\rho = 0$)	13.13			34.32		
Prob > χ^2	0.000			0.000		
observations			6,999			
number of women			817			
Wald χ^2	6243.39			148.88		
Prob > χ^2	0.000			0.000		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

cluster-robust standard errors in parantheses

Taking into account the panel structure results in comparable implications (table C-7).

Table C-7: Results panel estimation – occupational mobility

	Upward			downward		
time out 13 – 24	0.021		(0.104)	0.024		(0.106)
… 25 – 36	0.106		(0.136)	0.190		(0.135)
… 37 – 48	0.215	*	(0.128)	0.470	***	(0.157)
… 49 – 60	0.367	*	(0.195)	0.483	**	(0.211)
… 61 – 72	0.253		(0.216)	0.647	***	(0.227)
… > 72	0.361		(0.255)	0.859	***	(0.231)
Inverse Mill's ratio …1995	1.371		(1.096)	1.617		(1.607)
1996	n.a.[18]		n.a.	0.997		(2.374)
1997	-0.807		(0.996)	0.263		(0.495)
1998	-0.081		(0.438)	0.683		(0.605)
1999	0.394		(598.529)	0.655	*	(0.362)
2000	-0.282		(1.607)	0.361		(0.352)
2001	0.701		(0.712)	0.947	**	(0.440)
2002	0.371		(0.319)	0.828	**	(0.333)
2003	0.524		(0.555)	0.698	**	(0.355)
2004	0.510		(0.416)	0.765	**	(0.300)
2005	0.814	***	(0.225)	0.731	***	(0.275)
2006	1.080	***	(0.399)	0.670	**	(0.286)
2007	0.337		(0.332)	0.542	*	(0.298)
2008	0.565		(0.561)	0.849	***	(0.307)
2009	0.388		(0.369)	0.824	**	(0.395)
2010	0.708	**	(0.312)	0.721	**	(0.316)
2011	1.004	**	(0.418)	0.860	*	(0.440)
2012	0.335		(322.547)	0.111		(0.586)
observations	4,414			4,459		
number of women	786			787		
χ^2 (IMRs)	29.77			28.48		
Prob > χ^2	0.028			0.055		
Wald χ^2	183.88			248.88		
Prob > χ^2	0.000			0.000		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations
bootstrapped standard errors in parantheses

The probability of an upward move is (weakly) significantly higher for time outs of 37-48 months and 49-60 months than for time outs of up to 12 months.[19] The probability of downward moves is significantly higher for leaves longer than 3 years compared to very short interruptions. Again, the marginal effects are presented in the appendix. The results concerning covariates also are comparable to those obtained via the Heckman procedure, which is why they

[18] IMR1996 ≠ 0 predicts failure perfectly. It is dropped and 45 observations are not used.
[19] Since there are no upward moves in 1996 and the probit selection model for 1995 is based on view observations, I also estimated the upward model without these two years. The results do not change. If the year 1995 is not included in the downward model, the results do not change either.

are presented in the appendix. For some covariates, the significance levels have changed. As for prestige level, the inverse Mill's ratios do not show a distinct pattern. The test of the inverse Mill's ratios hints to joint significance. Since the event of an upward move is rather rare, i.e., there are only few positive values contained in the dependent variable, the estimates should be interpreted with caution. This is less of a problem concerning downward moves. Nevertheless, I present linear probability models since in both cases there are few 1's compared to zeros which may lead to failures in the bootstrap-replications. The results are presented in the appendix. They essentially deliver the same implications as the probit models presented in table C-7. In addition, standard random effects models are estimated and presented in the appendix.

C.5.2.3 Selection

Concerning the SIOPS level estimations, the exclusion restriction of the female unemployment rate is insignificant (table C-4). The possibility of starting a new job immediately is significant. It is associated with a higher probability of observing a positive SIOPS. Not working the year prior to the interview highly and significantly reduces the probability of observing a positive SIOPS value. Education is associated with a higher probability of observing a SIOPS value. Returning after the second, resp. third or higher order birth is associated with a lower probability of selection compared to returning before the second birth. Living in East Germany and living with a partner are not associated with participation in this sample of newly mothers. The coefficient of the number of children in the youngest age group is highly significant and negative, as expected. The other age groups also are significant but positive. Full-time and part-time experience are positive and significant while further interruptions are associated with a lower probability. These results are also expected. Living in a multigenerational household has a negative sign but is only weakly significant.

The first-stage results concerning mobility are presented in table C-8. The results are essentially the same as in the case of the standard Heckman model.

Examining whether the sample selection is relevant for the results reveals the following. If investigating the SIOPS level in the standard Heckman case, the two equations might be independent from each other ($\chi^2 = 0.31$). But for the binary outcomes, they do not seem to be independent from each other ($\chi^2 = 13.13$ resp. $\chi^2 = 34.32$). If the two equations are independent, the model can be simplified to standard methods, which in case of occupational mobility is a standard probit model (e.g., Baum (2006), Miranda and Rabe-Hesketh (2006)). This is done for comparison in all cases. The results obtained by these estimations are presented in the appendix (tables A C-7, A C-8, A C-12 and A C-13). The results are slightly more distinct than those obtained using the selection correction but are essentially the same for prestige level. In case of mobility, the simple models show more distinct effects of time out durations. The implications are unchanged, which is why the results are not discussed in detail.

Table C-8: pooled Heckman estimation – selection equations – mobility

	Selection – Upward			Selection – Downward		
age	0.100	*	(0.053)	0.102	*	(0.053)
age squared	-0.002	***	(0.001)	-0.002	***	(0.001)
years of education	0.042	***	(0.012)	0.044	***	(0.012)
return after second birth	-0.490	***	(0.062)	-0.495	***	(0.062)
return after third+ birth	-0.505	***	(0.193)	-0.546	***	(0.188)
East-Germany	0.023		(0.101)	0.028		(0.099)
partner	0.075		(0.084)	0.085		(0.081)
# children 0-3	-0.829	***	(0.060)	-0.826	***	(0.061)
# children 4-6	0.150	***	(0.051)	0.154	***	(0.051)
# children 7-12	0.176	***	(0.059)	0.175	***	(0.058)
# children 13-16	0.192	**	(0.087)	0.195	**	(0.087)
experience full time	0.051	***	(0.009)	0.050	***	(0.009)
experience part time	0.143	***	(0.018)	0.142	***	(0.018)
further interruptions	-0.562	***	(0.049)	-0.560	***	(0.049)
multigenerational household	-0.307	*	(0.182)	-0.293		(0.181)
not working (t-1)	-0.772	***	(0.055)	-0.780	***	(0.054)
start new job now (t-1)	0.158	**	(0.080)	0.146	*	(0.080)
unemployment rate (f)	0.017		(0.010)	0.017		(0.010)
constant	-0.949		(0.880)	-1.009		(0.875)
Observations			6,999			
Clusters			817			

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

cluster-robust standard errors in parentheses

It must be mentioned that the inverse Mill's ratio shows a remarkable time trend. The likelihood of measuring a positive SIOPS value increases over time. On average, the time period between the first birth and the observation period increases with the observed wave and therefore it is more likely that a mother is working. Therefore, the inverse Mill's ratio decreases over time since it basically is the "nonselection hazard" (e.g., StataCorp (2009)).

C.5.2.4 Robustness and summary of results

The results reported in the previous sections are essentially robust over the models that are implemented. The estimations result in the same implications, irrespective of whether one inverse Mill's ratio or period-specific ones are calculated. In addition, I performed the same estimation on a sample also containing childless women who turn 40 during the observation period. They are not expected to become mothers in the future. Mothers whose first child is born before the observation period are not included. Their first birth-related interruption and their SIOPS value prior to it are not observed. Including them would dilute the results since a possibly experienced prestige penalty could not be observed. The estimated effects are somewhat weaker as concerns the prestige level but more distinct in regard to mobility. This might be explained by the fact that birth-related interruptions are more often anticipated or planned. As already shown in the descriptive analysis, mothers generally re-enter the labor market at the same prestige level as childless women. However, after 2-3 years, their career paths show a slightly lower level than those of childless women who do not interrupt their careers. This difference is also found for childless women who interrupt their career. This indicates that it is the interruption itself that harms the career, not the fact that it is birth-related. But the descriptive analysis also reveals that mothers who only interrupt their employment for up to 12 months show considerably higher prestige levels during the years following the first birth compared to other categories. The lowest prestige level is found for mothers who take more than six years off.

As discussed in the previous section, I also used standard methods to estimate the effect of interest. The effects turn out to be a little more distinct if selection is not corrected.

The results reveal that, irrespective of the model used, there is a link between occupational prestige and very long leaves of absence. While career interruptions of up to three years do not significantly result in a lower prestige level than very short time outs, the opposite is true for very long interruptions. Occupational mobility is higher for mothers taking long leaves compared to those taking only up to 12 months off. This is especially true for downward mobility, indicating that longer career interruptions destabilize mothers' careers. However, upward mobility is more likely if the time out lasted three to five years. The effect sizes of observing actual changes are indeed distinct in proportion to the overall probability, but since the events under investigation are rare ones in the estimation sample, the overall effect size is not considered large. One exception to this is the effect of very long time outs on downward moves.

Over all, the results of a destabilized career are in line with the results found by Aisenbrey et al. (2009). Further investigation of other dimensions of mobility is a topic for future research.

As discussed earlier, apart from the selection into long or short career interruptions, another selection mechanism might be very important, namely the self-selection into certain occupations that are more or less prestigious. It is likely that women with a high preference for family life self-select into occupations with lower prestige, anticipating that they will take long time outs later on (e.g., Malo and Muñoz-Bullón (2008), for a discussion considering graduates see e.g., Ochsenfeld (2012)). Therefore, they might not invest in their career the same way more career-driven women do. This selection takes place long before the first child is born. If controlling for the SIOPS level prior to the first birth, the results concerning the level as well as upward mobility are somewhat weaker, while they are somewhat stronger for downward mobility. However, the implication that long rather than short leaves destabilize mothers' careers remains unchanged.

C.6 Conclusions

This paper explores the question of whether there is a penalty in terms of occupational prestige and occupational mobility for maternity related career interruptions. The present study investigates the following question by applying selection correction models according to Heckman (1979): Is there a prestige penalty for child-related career interruptions in Germany? The GSOEP provides comprehensive panel information about households, persons, couples and children. I investigate women who became mothers during 1992 through 2012 and who have been working the year before they gave birth to their first child. The selection model estimates the probability of observing any SIOPS information for new mothers after the birth of their first child, which is the case if they have already returned to the labor market, if they are still participating after their first return or if they have returned after subsequent interruptions. Occupational mobility is defined as a 10% change on the basis of the SIOPS value prior to first birth as well as further increases resp. decreases of at least 10%. The descriptive analysis reveals that mothers show a slightly lower prestige score from after 2-3 years after their re-entry. Compared to childless women who report an interruption during the observation period, they do not face any prestige penalty. It appears that the overall occupational mobility seems to be slightly higher for mothers than for childless women without any interruptions. The estimations reveal a significant influence of the duration of the first time out connected to first birth. Long career interruptions result in a lower prestige score. The prestige level is negatively influenced by leaves longer than three years. The extent of this negative effect is, as expected, higher for longer breaks. If examining occupational mobility, this pattern is different. The probability of upward mobility is only slightly higher for time outs of three to five years than for those of up to one year. Downward mobility is also more likely if the time out is long.

I used pooled versions of the standard Heckman sample selection model as well as a strategy suggested by Wooldridge (1995) to account for the panel structure. I tested the need of selection correction and additionally estimated standard models without the correction. The results obtained are quite similar irrespective of which method is used and of whether I correct for selection or not. The results are in line with those of Aisenbrey et al. (2009) and Evertsson and

Grunow (2012). Aisenbrey et al. (2009) find destabilized careers linked to longer leaves in Germany. The present work's results point in the same direction using the GSOEP and different methodology. Evertsson and Grunow (2012) do not find any association between interruptions and upward mobility in Germany. However, they only compare leaves of less than 15 months and those of more than 15 months. The present study uses more categories, which makes it possible to understand in greater depth the effects of time out duration on occupational prestige.

C.7 Appendix

Table A C-1: Variables used to estimate the selection and the outcome equation

variable	selection	outcome
time out duration	-	✓
age	✓	✓
age squared	✓	✓
years of education	✓	✓
first return after second birth	✓	✓
first return after third or higher order birth	✓	✓
living in East-Germany	✓	✓
partner in the same household	✓	✓
number of children 0-3	✓	✓
number of children 4-6	✓	✓
number of children 7-12	✓	✓
number of children 13-16	✓	✓
labor market experience full time	✓	✓
labor market experience part time	✓	✓
subsequent employment interruptions	✓	✓
working full time	-	✓
working part time	-	✓
year dummies	✓	✓
multigenerational household	✓	✓
not working (t-1)	✓	-
female unemployment rate	✓	-
possible to start a new job immediately (t-1)	✓	-

Data source: GSOEP, 1992-2012

Table A C-2: group sizes of figure C-1

years after first return	mothers, 2+ children	mothers, 1 child	childless, no interruption	childless with interruption
0	467	371	0	725
1	334	288	3182	461
2	226	210	1925	331
3	169	173	1312	268
4	127	144	1090	207
5	104	119	924	160
6	79	106	790	128
7	62	85	686	100
8	55	62	570	44
9	44	49	502	16
10	37	37	439	5
11	34	29	390	10
12	21	21	318	2
13	13	12	276	2
14	9	9	180	1
15	5	9	133	1
16	3	7	114	0
17	1	3	96	2
18	0	2	85	0
19	0	1	74	0
20	0	0	60	0
21	0	0	47	0

Source: GSOEP, 1992-2012, own calculations

Table A C-3: group sizes of figure C-2

years after first birth	<12	13-36	37-72	>72
0	239	0	0	0
1	153	49	0	0
2	160	278	0	0
3	121	306	9	0
4	112	233	85	0
5	97	231	114	0
6	96	210	127	2
7	81	189	105	23
8	73	171	96	30
9	60	153	83	36
10	50	137	79	35
11	37	114	71	31
12	35	94	62	27
13	23	74	55	28
14	16	58	48	25
15	13	45	39	23
16	10	33	33	19
17	8	23	23	19
18	4	14	15	8
19	2	11	7	7
20	1	2	4	3
21	0	0	0	0

Source: GSOEP, 1992-2012, own calculations

Table A C-4: average marginal effects – pooled Heckman, occupational mobility

average marginal effects	upward			downward		
time out 13 – 24	0.001		(0.007)	0.000		(0.009)
… 25 – 36	0.007		(0.008)	0.012		(0.012)
… 37 – 48	0.014		(0.009)	0.042	**	(0.014)
… 49 – 60	0.025	*	(0.012)	0.043	*	(0.019)
… 61 – 72	0.015		(0.014)	0.060	**	(0.020)
… > 72	0.024		(0.016)	0.075	***	(0.022)
age	-0.000		(0.005)	0.002		(0.007)
age squared	-0.000		(0.000)	-0.000		(0.000)
years of education	-0.000		(0.001)	0.002		(0.002)
return after second birth	-0.014		(0.009)	-0.036	**	(0.013)
return after third+ birth	-0.040		(0.028)	-0.029		(0.024)
East-Germany	0.006		(0.006)	-0.013		(0.009)
partner	-0.014		(0.008)	0.001		(0.011)
# children 0-3	-0.013		(0.007)	-0.003		(0.009)
# children 4-6	0.010		(0.006)	0.007		(0.007)
# children 7-12	-0.007		(0.006)	-0.014		(0.008)
# children 13-16	-0.002		(0.010)	-0.016		(0.014)
experience full time	0.001		(0.001)	-0.000		(0.001)
experience part time	-0.001		(0.001)	-0.003		(0.002)
full time	-0.014		(0.008)	-0.044	***	(0.012)
part time	-0.014		(0.007)	-0.046	***	(0.009)
further interruptions	-0.001		(0.005)	-0.008		(0.007)
multigenerational household	-0.004		(0.019)	-0.048		(0.040)
Observations			6,999			

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

standard errors in parantheses

Table A C-5: Results of manual Heckman twostep estimation (SIOPS level)

	SIOPS			selection		
time out 13 – 24	-0.563		(0.895)	-		-
… 25 – 36	-0.335		(1.233)	-		-
… 37 – 48	-1.927		(1.256)	-		-
… 49 – 60	-2.151		(1.776)	-		-
… 61 – 72	-6.906	***	(1.999)	-		-
… > 72	-10.287	***	(2.197)	-		-
age	1.557	***	(0.531)	0.102	*	(0.053)
age squared	-0.015	**	(0.007)	-0.002	***	(0.001)
years of education	2.197	***	(0.208)	0.044	***	(0.012)
return after second birth	1.845		(1.348)	-0.495	***	(0.062)
return after third+ birth	1.034		(3.202)	-0.524	***	(0.194)
East-Germany	0.655		(0.898)	0.044		(0.101)
partner	2.108	**	(1.065)	0.075		(0.085)
# children 0-3	-0.220		(0.942)	-0.837	***	(0.061)
# children 4-6	0.039		(0.636)	0.145	***	(0.051)
# children 7-12	0.494		(0.670)	0.175	***	(0.059)
# children 13-16	2.075	***	(0.763)	0.186	**	(0.088)
experience full time	-0.321	**	(0.152)	0.051	***	(0.009)
experience part time	-0.402	**	(0.190)	0.144	***	(0.018)
full time	3.315	***	(0.921)	-		-
part time	4.665	***	(0.734)	-		-
further interruptions	-2.222	***	(0.763)	-0.569	***	(0.049)
multigenerational household	0.666		(1.904)	-0.314	*	(0.182)
Inverse Mill‘s ratio	0.483		(0.908)	-		-
not working (t-1)	-		-	-0.761	***	(0.055)
start new job now (t-1)	-		-	0.172	**	(0.080)
female unemployment rate	-		-	0.015		(0.010)
constant	-17.715	*	(9.152)	-0.960		(0.885)
rho	0.050					
χ^2 (for comparison)	0.283					
Prob > χ^2	0.595					
adjusted /Pseudo R squared	0.344			0.355		
observations	4,478			7,011		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

bootstrapped standard errors in parantheses

Table A C-6: Results of manual Heckman twostep estimation (mobility)

	upward			downward		
time out 13 – 24	0.001		(0.102)	-0.033		(0.110)
… 25 – 36	0.070		(0.118)	0.086		(0.138)
… 37 – 48	0.183		(0.139)	0.387	***	(0.146)
… 49 – 60	0.357	*	(0.193)	0.397	*	(0.212)
… 61 – 72	0.195		(0.205)	0.563	***	(0.195)
… > 72	0.328		(0.242)	0.751	***	(0.210)
age	0.018		(0.074)	0.067		(0.079)
age squared	-0.001		(0.001)	-0.001		(0.001)
years of education	0.001		(0.022)	0.039	**	(0.018)
return after second birth	-0.283	*	(0.148)	-0.552	***	(0.161)
return after third+ birth	-0.641	**	(0.290)	-0.494		(0.339)
East-Germany	0.113		(0.100)	-0.087		(0.098)
partner	-0.212	*	(0.122)	0.046		(0.135)
# children 0-3	-0.387	***	(0.148)	-0.383	***	(0.149)
# children 4-6	0.166	*	(0.100)	0.127		(0.091)
# children 7-12	-0.088		(0.097)	-0.086		(0.086)
# children 13-16	-0.007		(0.155)	-0.098		(0.154)
experience full time	0.025	*	(0.014)	0.019		(0.016)
experience part time	0.003		(0.021)	0.019		(0.023)
full time	-0.205	*	(0.124)	-0.453	***	(0.123)
part time	-0.205	*	(0.106)	-0.504	***	(0.084)
further interruptions	-0.120		(0.093)	-0.288	***	(0.089)
multigenerational household	-0.099		(0.281)	-0.623	**	(0.284)
Inverse Mill‘s ratio	0.835	***	(0.183)	1.254	***	(0.174)
constant	-1.787		(1.855)	-2.651	*	(1.405)
Wald χ^2	155.46			397.23		
Prob > χ^2	0.000			0.000		
Observations	4,433			4,478		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

bootstrapped standard errors in parantheses

Table A C-7: Results of simple OLS estimation on selected sample

	SIOPS		
time out 13 – 24	-0.554		(0.916)
… 25 – 36	-0.391		(1.279)
… 37 – 48	-1.837		(1.240)
… 49 – 60	-1.973		(1.745)
… 61 – 72	-7.007	***	(1.985)
… > 72	-10.247	***	(2.086)
age	1.556	***	(0.542)
age squared	-0.015	**	(0.007)
years of education	2.195	***	(0.203)
return after second birth	1.839		(1.290)
return after third+ birth	1.272		(3.118)
East-Germany	0.756		(0.866)
partner	2.014	*	(1.087)
# children 0-3	0.101		(0.765)
# children 4-6	-0.040		(0.638)
# children 7-12	0.433		(0.628)
# children 13-16	2.056	***	(0.757)
experience full time	-0.334	**	(0.144)
experience part time	-0.421	**	(0.174)
full time	3.111	***	(0.915)
part time	4.583	***	(0.733)
further interruptions	-2.114	***	(0.712)
multigenerational household	0.756		(1.574)
constant	-17.380	**	(9.142)
F statistic	13.61		
Prob > F	0.000		
observations	4,596		
number of women	800		
R-squared	0.350		

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

cluster-robust standard errors in parantheses

Source: GSOEP, 1992-2012, own calculations

Table A C-8: results of pooled standard Probit estimations for binary outcome equations

coefficients	upward move			downward move		
time out 13 – 24	0.053		(0.098)	0.072		(0.099)
… 25 – 36	0.161		(0.122)	0.232	*	(0.120)
… 37 – 48	0.271	**	(0.129)	0.482	***	(0.138)
… 49 – 60	0.453	***	(0.173)	0.557	***	(0.178)
… 61 – 72	0.256		(0.204)	0.631	***	(0.196)
… > 72	0.490	**	(0.229)	0.756	***	(0.200)
age	-0.047		(0.071)	-0.029		(0.071)
age squared	0.001		(0.001)	0.000		(0.001)
years of education	-0.006		(0.020)	0.005		(0.019)
return after second birth	-0.087		(0.129)	-0.206		(0.131)
return after third+ birth	-0.520		(0.431)	-0.167		(0.239)
East-Germany	0.041		(0.089)	-0.215	**	(0.101)
partner	-0.217	*	(0.111)	-0.002		(0.118)
# children 0-3	0.052		(0.108)	0.247	***	(0.092)
# children 4-6	0.128		(0.087)	0.063		(0.078)
# children 7-12	-0.139		(0.089)	-0.164	**	(0.081)
# children 13-16	-0.075		(0.147)	-0.202		(0.139)
experience full time	0.006		(0.014)	-0.014		(0.014)
experience part time	-0.047	**	(0.020)	-0.068	***	(0.023)
full time	-0.239	**	(0.118)	-0.522	***	(0.110)
part time	-0.201	**	(0.095)	-0.539	***	(0.081)
further interruptions	0.099		(0.072)	0.106		(0.070)
multigenerational household	0.024		(0.301)	-0.446		(0.396)
constant	-0.383		(1.272)	-0.273		(1.230)
Wald χ^2	115.06			223.83		
average marginal effects						
time out 13 – 24	0.004		(0.008)	0.007		(0.009)
… 25 – 36	0.013		(0.010)	0.022		(0.011)
… 37 – 48	0.023	*	(0.011)	0.046	***	(0.013)
… 49 – 60	0.038	**	(0.015)	0.053	**	(0.017)
… 61 – 72	0.021		(0.017)	0.060	**	(0.019)
… > 72	0.041	*	(0.019)	0.071	***	(0.019)
age	-0.004		(0.006)	-0.003		(0.007)
age squared	0.000		(0.000)	0.000		(0.000)
years of education	-0.001		(0.002)	0.001		(0.002)
return after second birth	-0.007		(0.011)	-0.019		(0.012)
return after third+ birth	-0.043		(0.036)	-0.016		(0.023)
East-Germany	0.003		(0.007)	-0.020	*	(0.009)
partner	-0.018		(0.009)	-0.000		(0.011)
# children 0-3	0.004		(0.009)	0.023	**	(0.009)
# children 4-6	0.011		(0.007)	0.006		(0.007)
# children 7-12	-0.012		(0.007)	-0.016	*	(0.008)
# children 13-16	-0.006		(0.012)	-0.019		(0.013)
experience full time	0.001		(0.001)	-0.001		(0.001)
experience part time	-0.004	*	(0.002)	-0.006	**	(0.002)
full time	-0.020	*	(0.010)	-0.049	***	(0.010)
part time	-0.017	*	(0.008)	-0.051	***	(0.008)
further interruptions	0.008		(0.006)	0.010		(0.007)
multigenerational household	0.002		(0.025)	-0.042		(0.038)
Observations	4,551			4,596		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2012, own calculations

(cluster-robust) standard errors in parantheses

Table A C-9: complete results panel estimation – occupational mobility

	upward			downward		
time out 13 – 24	0.021		(0.104)	0.024		(0.106)
… 25 – 36	0.106		(0.136)	0.190		(0.135)
… 37 – 48	0.215	*	(0.128)	0.470	***	(0.157)
… 49 – 60	0.367	*	(0.195)	0.483	**	(0.211)
… 61 – 72	0.253		(0.216)	0.647	***	(0.227)
… > 72	0.361		(0.255)	0.859	***	(0.231)
age	-0.019		(0.074)	0.006		(0.084)
age squared	0.000		(0.001)	-0.000		(0.001)
years of education	-0.008		(0.022)	0.032		(0.021)
return after second birth	-0.121		(0.149)	-0.366	**	(0.157)
return after third+ birth	-0.367		(0.310)	-0.201		(0.316)
East-Germany	0.083		(0.100)	-0.155		(0.113)
partner	-0.208	*	(0.120)	-0.018		(0.134)
# children 0-3	-0.219		(0.152)	-0.065		(0.132)
# children 4-6	0.112		(0.100)	0.091		(0.090)
# children 7-12	-0.144		(0.105)	-0.145		(0.091)
# children 13-16	-0.067		(0.155)	-0.165		(0.156)
experience full time	0.022		(0.017)	0.008		(0.017)
experience part time	-0.012		(0.022)	-0.014		(0.025)
full time	-0.207		(0.130)	-0.480	***	(0.125)
part time	-0.218	**	(0.104)	-0.506	***	(0.090)
further interruptions	0.024		(0.087)	-0.084		(0.083)
multigenerational household	0.057		(0.323)	-0.443		(0.283)
Inverse Mill's ratio …1995	1.371		(1.096)	1.617		(1.607)
1996	n.a.[20]		n.a.	0.997		(2.374)
1997	-0.807		(0.996)	0.263		(0.495)
1998	-0.081		(0.438)	0.683		(0.605)
1999	0.394		(598.529)	0.655	*	(0.362)
2000	-0.282		(1.607)	0.361		(0.352)
2001	0.701		(0.712)	0.947	**	(0.440)
2002	0.371		(0.319)	0.828	**	(0.333)
2003	0.524		(0.555)	0.698	**	(0.355)
2004	0.510		(0.416)	0.765	**	(0.300)
2005	0.814	***	(0.225)	0.731	***	(0.275)
2006	1.080	***	(0.399)	0.670	**	(0.286)
2007	0.337		(0.332)	0.542	*	(0.298)
2008	0.565		(0.561)	0.849	***	(0.307)
2009	0.388		(0.369)	0.824	**	(0.395)
2010	0.708	**	(0.312)	0.721	**	(0.316)
2011	1.004	**	(0.418)	0.860	*	(0.440)
2012	0.335		(322.547)	0.111		(0.586)
constant	-1.932		(1.582)	-2.086		(1.952)
Observations	4,414			4,459		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

bootstrapped standard errors in parantheses

[20] IMR1996 ≠ 0 predicts failure perfectly. It is dropped and 45 observations are not used.

Table A C-10: average marginal effects, panel estimation with selection correction

	upward move			downward move		
time out 13 - 24	0.002		(0.026)	0.002		(0.010)
… 25 - 36	0.009		(0.123)	0.017		(0.012)
… 37 - 48	0.017		(0.250)	0.043	**	(0.014)
… 49 - 60	0.029		(0.426)	0.044	*	(0.019)
… 61 - 72	0.020		(0.293)	0.058	**	(0.021)
… > 72	0.029		(0.418)	0.078	***	(0.021)
age	-0.002		(0.023)	0.001		(0.008)
age squared	0.000		(0.000)	-0.000		(0.000)
years of education	-0.001		(0.009)	0.003		(0.002)
return after second birth	-0.010		(0.141)	-0.033	*	(0.014)
return after third+ birth	-0.029		(0.425)	-0.018		(0.029)
East-Germany	0.007		(0.096)	-0.014		(0.010)
partner	-0.017		(0.240)	-0.002		(0.012)
# children 0-3	-0.018		(0.254)	-0.006		(0.012)
# children 4-6	0.009		(0.129)	0.008		(0.008)
# children 7-12	-0.012		(0.167)	-0.013		(0.008)
# children 13-16	-0.005		(0.078)	-0.015		(0.014)
experience full time	0.002		(0.026)	0.001		(0.001)
experience part time	-0.001		(0.014)	-0.001		(0.002)
full time	-0.017		(0.239)	-0.043	***	(0.011)
part time	-0.018		(0.252)	-0.046	***	(0.008)
further interruptions	0.002		(0.028)	-0.008		(0.008)
multigenerational household	0.005		(0.071)	-0.040		(0.026)
Inverse Mill's ratio …1995	0.110		(1.591)	0.146		(0.146)
1996	n.a.		n.a.	0.090		(0.211)
1997	-0.065		(0.945)	0.024		(0.045)
1998	-0.006		(0.102)	0.062		(0.055)
1999	0.032		(48.515)	0.059		(0.033)
2000	-0.023		(0.353)	0.033		(0.032)
2001	0.056		(0.817)	0.086	*	(0.041)
2002	0.030		(0.433)	0.075	*	(0.031)
2003	0.042		(0.608)	0.063	*	(0.032)
2004	0.041		(0.591)	0.069	*	(0.027)
2005	0.065		(0.942)	0.066	*	(0.026)
2006	0.087		(1.253)	0.061	*	(0.026)
2007	0.027		(0.391)	0.049		(0.027)
2008	0.045		(0.657)	0.077	**	(0.028)
2009	0.031		(0.449)	0.074	*	(0.036)
2010	0.057		(0.820)	0.065	*	(0.029)
2011	0.081		(1.165)	0.078		(0.040)
2012	0.027		(26.043)	0.010		(0.053)
Observations	4,414			4,459		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

standard errors in parantheses

Table A C-11: linear regression with period-specific inverse Mill's ratios

linear regression	upward move		downward move	
time out 13 - 24	0.001	(0.008)	0.001	(0.009)
… 25 - 36	0.008	(0.011)	0.012	(0.011)
… 37 - 48	0.021 *	(0.013)	0.042 ***	(0.015)
… 49 - 60	0.037 *	(0.021)	0.044 **	(0.021)
… 61 - 72	0.022	(0.019)	0.059 **	(0.024)
… > 72	0.033	(0.022)	0.071 ***	(0.022)
age	-0.003	(0.006)	-0.001	(0.008)
age squared	0.000	(0.000)	-0.000	(0.000)
years of education	-0.000	(0.002)	0.003 *	(0.002)
return after second birth	-0.017	(0.013)	-0.041 ***	(0.015)
return after third+ birth	-0.036	(0.024)	-0.021	(0.033)
East-Germany	0.008	(0.008)	-0.010	(0.009)
partner	-0.019 *	(0.011)	-0.006	(0.010)
# children 0-3	-0.027 **	(0.011)	-0.026 *	(0.014)
# children 4-6	0.013 *	(0.007)	0.013 *	(0.007)
# children 7-12	-0.007	(0.006)	-0.005	(0.006)
# children 13-16	-0.005	(0.009)	-0.005	(0.008)
experience full time	0.002	(0.001)	0.002	(0.001)
experience part time	0.000	(0.001)	0.001	(0.002)
full time	-0.019	(0.012)	-0.063 ***	(0.015)
part time	-0.020 **	(0.009)	-0.063 ***	(0.012)
further interruptions	-0.003	(0.008)	-0.017 **	(0.008)
multigenerational household	0.007	(0.051)	-0.057	(0.043)
Inverse Mill's ratio …1995	0.064	(0.089)	0.276	(0.188)
1996	0.016	(0.013)	0.194	(0.119)
1997	-0.052	(0.062)	0.084	(0.109)
1998	-0.002	(0.043)	0.124 *	(0.070)
1999	0.067	(0.072)	0.147 **	(0.074)
2000	-0.034	(0.077)	0.106 *	(0.062)
2001	0.073 *	(0.040)	0.125 **	(0.053)
2002	0.061	(0.040)	0.188 ***	(0.066)
2003	0.041 *	(0.024)	0.093 **	(0.042)
2004	0.060	(0.037)	0.164 ***	(0.054)
2005	0.089 **	(0.037)	0.105 **	(0.046)
2006	0.125 **	(0.050)	0.095 **	(0.039)
2007	0.071	(0.045)	0.130 ***	(0.049)
2008	0.042 *	(0.025)	0.124 ***	(0.046)
2009	0.071	(0.048)	0.149 **	(0.058)
2010	0.071 **	(0.035)	0.100 **	(0.045)
2011	0.155 **	(0.065)	0.157 **	(0.067)
2012	0.033	(0.042)	0.026	(0.022)
constant	0.106	(0.117)	0.035	(0.181)

Table A C-11 continued

Observations	4,459	4,459
number of women	787	787
χ^2 (IMRs)	24.69	40.48
Prob > χ^2	0.134	0.002
Wald χ^2	164.95	202.85
Prob > χ^2	0.000	0.000

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1995-2012, own calculations

bootstrapped standard errors in parantheses

Table A C-12: Results of simple Random Effects Estimation – SIOPS level

	SIOPS		
time out 13 – 24	-0.671		(0.870)
… 25 – 36	-1.579		(1.158)
… 37 – 48	-2.682	**	(1.352)
… 49 – 60	-3.919	**	(1.613)
… 61 – 72	-7.690	***	(2.288)
… > 72	-9.533	***	(2.097)
age	0.255		(0.412)
age squared	-0.000		(0.005)
years of education	2.080	***	(0.225)
return after second birth	3.180	**	(1.288)
return after third+ birth	1.829		(3.114)
East-Germany	-0.710		(0.757)
partner	1.135	*	(0.679)
# children 0-3	-0.848	*	(0.468)
# children 4-6	-0.272		(0.412)
# children 7-12	0.010		(0.394)
# children 13-16	0.643		(0.498)
experience full time	-0.095		(0.140)
experience part time	-0.162		(0.150)
full time	2.378	***	(0.756)
part time	2.384	***	(0.590)
further interruptions	-1.578	**	(0.686)
multigenerational household	-1.083		(1.727)
constant	11.969		(7.335)
observations	4,596		
number of women	800		
Wald χ^2	350.36		
Prob > χ^2	0.000		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1992-2012, own calculations

cluster-robust standard errors in parantheses

Table A C-13: Results of simple Random Effects Probit[21]

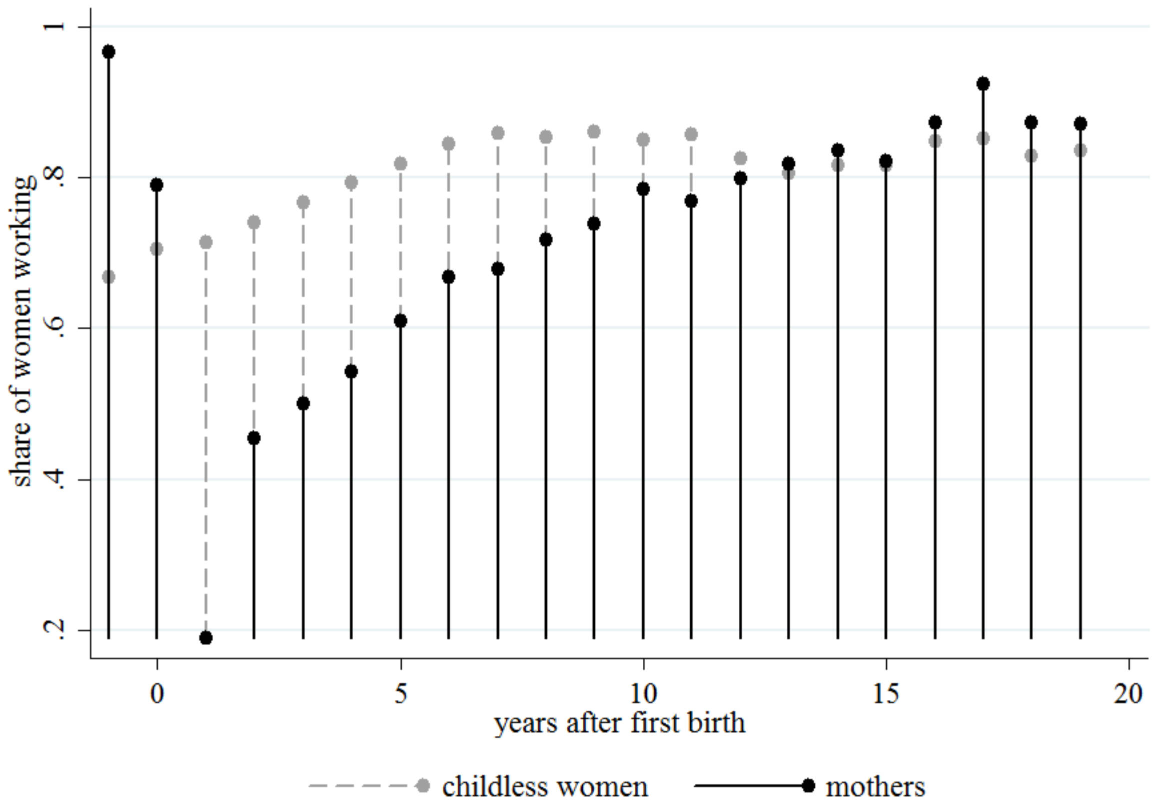
	upward move			downward move		
time out 13 - 24	0.053		(0.103)	0.072		(0.099)
… 25 - 36	0.161		(0.130)	0.232	*	(0.126)
… 37 - 48	0.271	*	(0.143)	0.482	***	(0.138)
… 49 - 60	0.453	**	(0.188)	0.557	***	(0.185)
… 61 - 72	0.256		(0.236)	0.631	***	(0.215)
… > 72	0.490	**	(0.229)	0.756	***	(0.217)
age	-0.047		(0.075)	-0.029		(0.072)
age squared	0.001		(0.001)	0.000		(0.001)
years of education	-0.006		(0.020)	0.005		(0.019)
return after second birth	-0.087		(0.143)	-0.206		(0.136)
return after third+ birth	-0.520		(0.426)	-0.167		(0.308)
East-Germany	0.041		(0.093)	-0.215	**	(0.097)
partner	-0.217	*	(0.112)	-0.002		(0.123)
# children 0-3	0.052		(0.108)	0.247	**	(0.099)
# children 4-6	0.128		(0.089)	0.063		(0.084)
# children 7-12	-0.139		(0.089)	-0.164	*	(0.087)
# children 13-16	-0.075		(0.138)	-0.202		(0.149)
experience full time	0.006		(0.015)	-0.014		(0.015)
experience part time	-0.047	**	(0.020)	-0.068	***	(0.019)
full time	-0.239	**	(0.119)	-0.522	***	(0.109)
part time	-0.201	**	(0.094)	-0.539	***	(0.082)
further interruptions	0.099		(0.078)	0.106		(0.074)
multigenerational household	0.024		(0.387)	-0.446		(0.489)
constant	-0.383		(1.293)	-0.273		(1.225)
observations	4,569			4,596		
groups	800			800		
Wald χ^2	112.39			207.41		
Prob > χ^2	0.000			0.000		

*** p<0.01, ** p<0.05, * p<0.1
standard errors in parantheses

Source: GSOEP, 1995-2012, own calculations

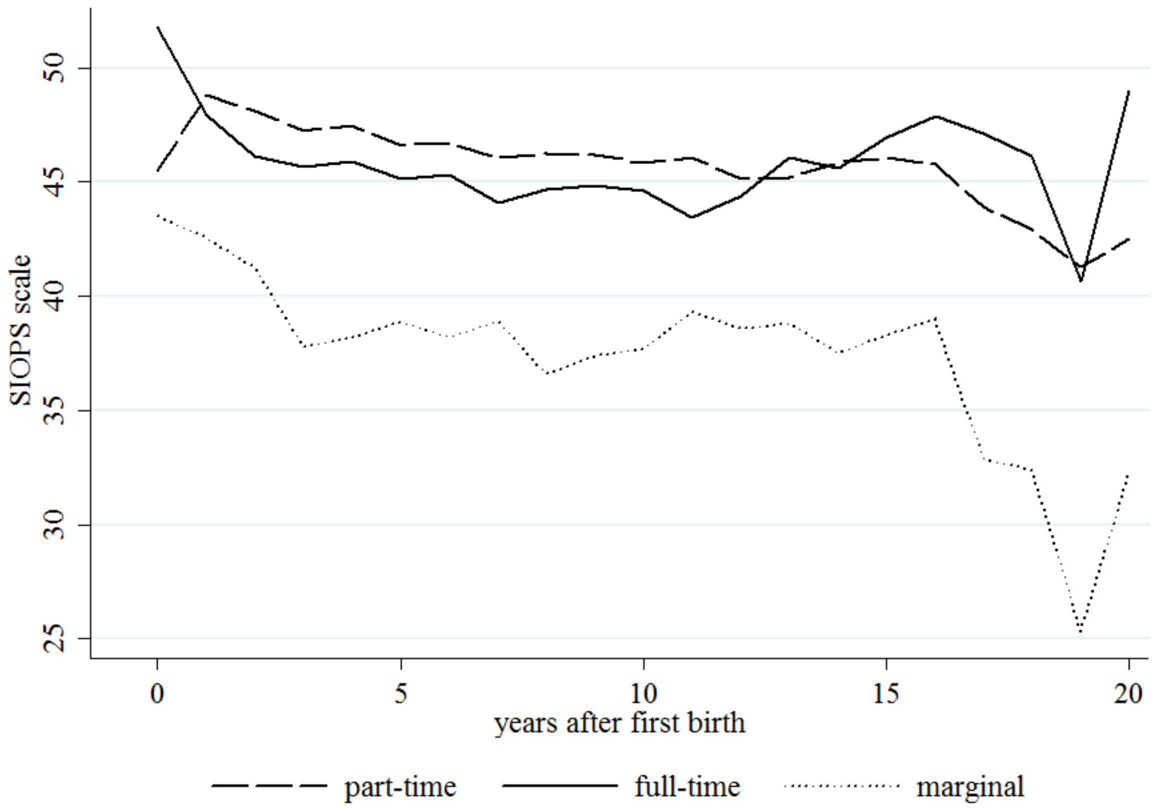
[21] Performing a linear random effects model in both cases leads to comparable implications.

Figure A C-1: proportion of mothers and childless women participating in the labor market



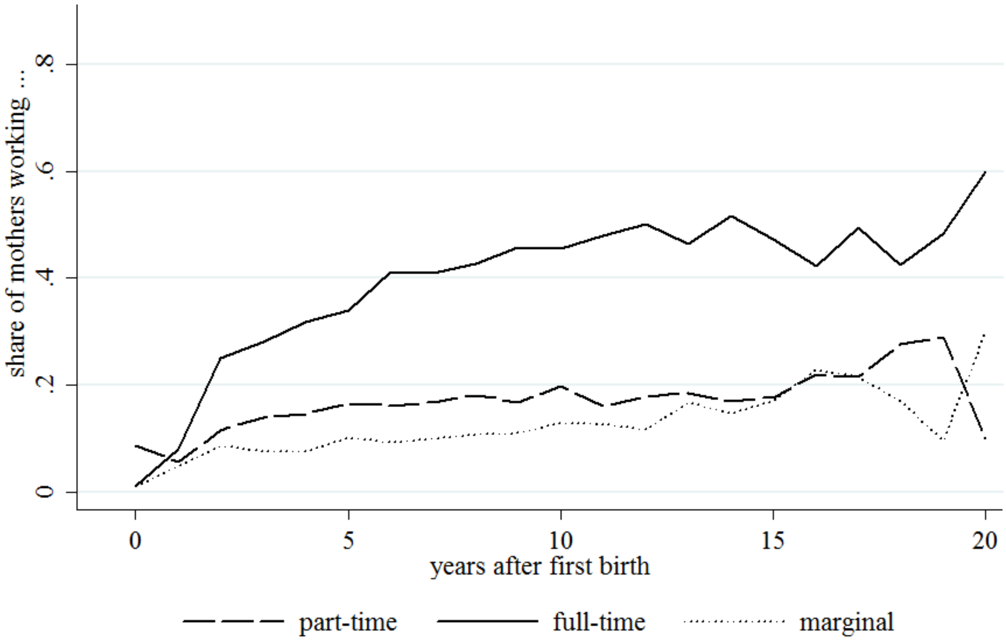
Source: GSOEP, 1992-2012, own calculations

Figure A C-2: SIOPS value of mothers being full-time, part-time or marginally employed



Source: GSOEP, 1992-2012, own calculations

Figure A C-3: share of mothers employed to different extents



Source: GSOEP, 1992-2012, own calculations

Table A C-14: group sizes of figures A C-2 and A C-3

years after first birth	full-time	part-time	marginal	working	total
0	16	2	2	20	133
1	62	87	53	202	838
2	112	242	84	438	822
3	123	246	67	436	766
4	117	252	61	430	705
5	121	247	74	442	655
6	106	268	61	435	597
7	99	240	59	398	541
8	94	220	56	370	480
9	76	206	50	332	420
10	76	175	50	301	361
11	53	158	42	253	309
12	49	137	32	218	257
13	41	102	37	180	207
14	30	91	26	147	168
15	69	26	25	120	138
16	24	46	25	95	106
17	17	39	17	73	77
18	13	20	8	41	45
19	9	15	3	27	29
20	1	6	3	10	10
21	0	0	0	0	0

Source: GSOEP, 1992-2012, own calculations

Part D

D Informal childcare and maternal employment

Katharina Sutter

D.1 Introduction

The labor market participation and the extent of the employment of mothers are of great interest in German politics (e.g., Spieß (2011), Ristau (2005)). A greater extent of maternal employment is a widely discussed topic in German politics (e.g., BMFSFJ (2014), BMI (2011)). This often is considered to be desirable due to the demographic change Germany is facing and the associated challenges for the economy as a whole (e.g., European Commission (2012), OECD (2014)). Several instruments have been implemented in recent years to make market work and family obligations more compatible (see for instance Gerlach (2008)). One of the instruments is the expansion of publicly provided childcare (e.g., 1996 - KJHG, 2005 - TAG, 2008 - KiföG[1]). The participation rate of women has been increasing in recent decades (e.g., European Commission (2014)), but the share of German mothers working full-time remains low (e.g., Keller and Haustein (2013)). The proportion of women working part-time is among the highest in Europe (e.g., European Commission (2014)). One reason for this behavior might be a lack of flexibility, which is not offered by public childcare institutions. In many cases, the office hours of such institutions do not meet the requirements, e.g., they are too short or not flexible enough, especially in West Germany (e.g., Federal Statistical Office of Germany (2012), Stöbe-Blossey (2010), Krone and Stöbe-Blossey (2010)). Therefore, dual-earner families or single mothers might use additional childcare arrangements, such as grandparents, friends or neighbors. These people are more likely to cover the hours that are not covered by public childcare arrangements, e.g., early evenings. This chapter investigates the importance of such informal childcare arrangements for the labor market behavior of mothers, both at the extensive and the intensive margin.

This study asks whether informal childcare arrangements increase maternal labor market participation and hours worked. It contributes to the existing literature by using the survey years

[1] For a description of these policy instruments see for instance Gerlach (2008). KJHG: “Kinder- und Jugendhilfegesetz“ (Child and Youth Services Act), TAG: “Tagesbetreuungsausbaugesetz“ (Day Care Expansion Act), KiFöG: „Kinderförderungsgesetz“ (Law for the support of children).

1999-2012 of the German Socio-Economic Panel (GSOEP)[2] and thereby providing further evidence for the case of Germany. I take into account the potential endogeneity of informal childcare with respect to maternal labor market participation and explicitly distinguish between the decision about labor market participation and the decision about the working hours by allowing both decisions to depend on different processes. A Heckman-type selection correction is performed when investigating the working hours.

The remainder of this chapter is organized as follows. Section D.2 summarizes the existing literature. The subsequent section contains an overview of the dataset and in section D.4 some descriptive information about the sample is presented. After discussing the empirical strategy, the results are presented and discussed. Section D.6 concludes.

D.2 Literature Review

The literature on the relationship between childcare and maternal labor supply is broad (e.g., Felfe et al. (2013), Chiuri (2000), Connelly (1992), Dimova and Wolff (2008), Duncan et al. (2004), Hansen et al. (2006), Heckman (1974), Hofferth and Collins (2000), Kreyenfeld and Hank (2000)). Some studies focus on the costs of formal childcare (e.g., Connelly (1992), Wrohlich (2004)), others on the availability (e.g., Kreyenfeld and Hank (2000)). Some literature also focuses on the role of fathers (e.g., Kitterød and Pettersen (2006), Wang and Bianchi (2009)). The present study focuses on the relationship between informal childcare and maternal employment. Informal childcare can be provided within the household (partner, cohabiting relatives), within the family, (e.g., not cohabiting grandparents), or outside the family (e.g., friends or neighbors). This kind of childcare arrangement is considered to provide flexibility to the mother (e.g., Posadas and Vidal-Fernández (2012)). Another argument for the importance of informal childcare might be that mothers trust their parents even more than they trust childcare institutions and view this form of childcare as the best substitute for their own childcare (e.g., Posadas and Vidal-Fernández (2012), Wheelock and Jones (2002)).

The role of grandparents' childcare in maternal employment is investigated by Aassve et al. (2011). They control for unobserved preferences by using a simultaneous equation approach. To model the probability of childcare by the grandmother, they use the information about whether the grandmother is alive as well as the number of siblings of the mother. Grandparental care is included in the equation modelling the probability of mothers' labor market participation. For Germany, they find a positive effect on the labor market participation of

[2] For a description of this very rich panel dataset, see Wagner et al. (2007).

mothers. The same is true for France, Bulgaria and Hungary. Albuquerque and Passos (2010) focus on southern European countries and estimate a switching probit model. In addition, the authors estimate bivariate probit models and univariate probit models. They find a positive effect of grandparental childcare, but this depends on the model used. A recent study by Posadas and Vidal-Fernández (2012) analyzes the U.S. (NLSY79). Childcare provided by grandparents is considered to be flexible, affordable and a good substitute for parental childcare. They apply an instrumental variable approach, taking grandmother's death as an instrument for grandparents' childcare. The results reveal a positive impact of access to grandparental care on maternal labor force participation, with the effect being mostly driven by mothers from disadvantaged socio-economic background. Zamarro (2011) uses SHARE data. She takes into account the simultaneity of the decisions about participation and childcare activities as well as unobserved heterogeneity in care-giving decisions. For some countries, she finds a positive and significant effect of grandparental childcare on participation of mothers, with Germany not being one of those countries. In the case of sons being investigated she finds no effect.

A study focusing on French Immigrants is written by Dimova and Wolff (2008). Apart from the effect of grandchild care on the labor market participation of the mother, they also consider the distribution of this transfer across one's own children. As a first step, they model the supply of grandchild care. In case of labor market participation they also take care of the potential endogeneity of grandparental childcare. This is done using a bivariate probit model as well as 2SLS, generalized 2SLS random effects IV and fixed effects IV estimation. They find a positive and significant effect of grandparental care on the participation of the mother. Gray (2005) also investigates the role of grandparental childcare in connection with maternal employment. She sheds some light on trends in the provision of this kind of childcare and on how it interacts with formal childcare arrangements. She concludes that informal care provided by grandparents helps mothers to enter the labor market and also to work longer hours. It must be mentioned that this study focuses on a descriptive view on the data. Hofferth and Collins (2000) analyze the influence of cost, quality as well as availability of childcare not provided by the mother on the probability of exiting the labor market. The authors estimate discrete time Logit models. They find a negative link between non-parental care arrangements and job exits. Having multiple arrangements turns out to be a possible strategy to combat the breakdown of one arrangement. They conclude that availability, cost and stability of non-parental childcare is strongly associated with the stability of maternal employment. Compton (2011) investigates family proximity and co-residence in Canada and its effect on the participation of women. She argues that this not only measures the effect of transferred care but also an insurance effect of

– as Compton (2011) calls it – “back-up” childcare. She estimates probit models for the participation decision and Tobit models for usual hours of work. She also instruments proximity in a further model. The results indicate a positive effect of proximity to one’s own mother (-in-law) on labor force participation but a negative effect of co-residence. A similar paper is written by Compton and Pollak (2014). Again, the availability of childcare and its insurance effect are mentioned. They discuss the flexibility of grandparental care compared to market-based childcare arrangements, which most often include regular and anticipated childcare hours. This study also uses another sample (“military wives”), and still finds the association between grandmothers childcare and labor force participation. They also investigate the working hours focusing on Tobit estimations and conclude that proximity is positively associated with participation, but not with hours worked. Selection correction is only mentioned as a side note.
Using SHARE data, Dimova and Wolff (2011) investigate downward time and money transfers and their influences on the labor market behavior of young mothers receiving those transfers. Ten European countries are included in the analysis. Overall, the authors find a positive effect of grandparental childcare on labor market participation as well as on the extent of employment while money transfers do not seem to matter in respect to these decisions.

Wagner (2012) investigates whether the social networks, namely the presence of kinship, friends and spouses, increase maternal employment. She uses the GSOEP and applies piecewise constant event history models for competing risks. The author analyzes six years after a birth and the transitions to full-time or part-time employment. The presence of relatives raises transitions to both full-time as well as part-time employment of West German and migrant mothers. Social support seems not to be relevant for the employment of East German mothers.

A rather early paper by Maume and Mullin (1993) investigates the influence of childcare arrangements in 1985 on female employment turnover by 1986. They focus on the role of fathers and ask whether mothers who rely on paternal childcare as the primary source of childcare are more likely to quit their job than women who have access to different childcare arrangements. They find that mothers, especially low-wage-earning mothers who rely on their husband’s childcare, are more likely to quit work.

Hansen et al. (2006) investigate the use of formal and informal childcare arrangements and the connection of maternal employment to child outcomes. They conclude that informal care is important for mothers in respect to the balancing of work and family obligations for very young children. An article of Wheelock and Jones (2002) shows that parents often rely on informal care as a complementary form of childcare to formal arrangements. The authors draw descriptive conclusions of a study focusing on informal childcare used by employed parents.

Informal care is defined as care provided by relatives, friends or neighbors, but not by the spouse. The authors find evidence that working parents depend on complementary childcare, and that this type of childcare is used to a greater extent than formal childcare. The authors also state that not only is childcare a mainly female task, but that its organization is as well. Kreyenfeld and Hank (2000) explicitly look at the availability of public childcare. Their results do not reveal any significant influence of regional availability of day care slots on the participation of mothers. Kreyenfeld and Hank (2000) argue that this indicates that the West German day care regime is inadequate since the office hours of the facilities are limited. They further argue that mothers who want to participate have to rely on additional care arrangements (social network).

My contribution to the literature is as follows. First, by using the German Socio-Economic Panel (GSOEP) I am able to use a relatively long observation period (1999-2012). In addition, this dataset contains very rich information about all members of the household, including children and the partner. For each child living in a household, it contains information about childcare arrangements. Second, as also stated in several of the aforementioned studies (e.g., Dimova and Wolff (2008), Albuquerque and Passos (2010), Kalb (2009)), there is scarce evidence about the research question in focus. This study provides further evidence for the case of Germany, using the data from as recently as 2012, resulting in very up-to-date findings. I include children up to the age children start secondary school. I use the same instrument as Aassve et al. (2011) or Posadas and Vidal-Fernández (2012) in a two-stage least-squares framework. But Aassve et al. (2011) use the Generations and Gender Survey of 2005, which only contains one wave. Posadas and Vidal-Fernández (2012) analyze the U.S. by using the NLSY79. As the authors mention, they are not able to exploit the panel structure due to the lack of data availability. I employ the panel structure of the GSOEP. Panel estimators (EC2SLS, G2SLS) suggested by the literature (e.g., Dimova and Wolff (2008), Baltagi (2008)) are applied to estimate the participation decision. Apart from participation itself, I also investigate the extent of employment measured by actual working hours. I take into account that this decision is not driven by the same process as the decision to participate and explicitly correct for this selection by using a Heckman selection correction (Heckman (1979)) in the model focusing on the working hours. I also apply a strategy suggested by Wooldridge (1995, 2010) to again consider the panel structure. Dimova and Wolff (2011) look at labor market involvement by distinguishing between no work, part-time work and full-time work. The papers of Compton (2011) as well as Compton and Pollak (2014) also investigate working hours, but with a focus on Tobit estimation. Compton and Pollak (2014) focus on proximity rather than childcare itself.

In addition, they only mention the selection correction as a side note. To the best of my knowledge, the research question of this paper has not been investigated for Germany in the way I do using this exact empirical strategy.

D.3 Data

I use the survey years 1999 through 2012[3] of the German Socio-Economic Panel (GSOEP).[4] Before 1999, the information about childcare arrangements cannot be separated with sufficient clarity, and for some years, there is no information about informal childcare at all. The sample is restricted to women. The GSOEP contains detailed information about the childcare arrangements of children living in the household. I only consider women between the ages of 20 and 59. I do not include women who are near retirement with zero working hours, who are in vocational training, in short-time work or in a sheltered workshop since these are special cases. I only consider women who live with at least one child up to the age of 12. This is considered a crucial age in respect to the need to being cared for by adults.[5] I include single mothers as well as mothers who live with a partner. Childcare arrangements are distinguished in the following way. There are (public) institutions, e.g., kindergartens or nannies, and informal arrangements, e.g., family, friends or neighbors. The questionnaire explicitly focuses on arrangements on a regular basis, meaning that cases in which grandmothers might fill-in as an exception are not included. However, it does not distinguish between grandparents and other relatives. It must be mentioned that the question about informal arrangements only refers to persons outside the household. This means childcare provided by cohabiting grandparents is not included. Multigenerational households are a rather rare household type in the considered sample (about 3% of the mothers in the sample live in such a household). Nevertheless, this might result in an underestimation of the effect later on. Another issue when investigating the importance of informal childcare might be the care that the grandparents need. This is not considered to be an issue here, since only approximately 7% of the women in the sample spend time on this task. But these 7% do not all refer to grandparents who are in need, but to care in general. This means it is not necessarily the grandparent who is in need.

Overall this results in a sample containing 5,995 women with at least one child up to age 12 living in the household. 2,995 of these women use childcare provided by grandparents, friends or neighbors at least at some point in time, while 3,787 women use formal childcare.

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[3] The year 2003 is excluded since there is no information available about informal childcare arrangements.

[4] For a description of this very rich panel dataset, see Wagner et al. (2007).

[5] E.g., from age 13 on, German children are allowed to start working to a small extent (§ 5 JArbSchG).

D.4 Descriptive Analysis

The share of women using informal childcare is somewhat smaller than the share using formal childcare (2,995 vs. 3,787 of 5,995). Table D-1 contains further information, and it is to be noted that these values are averaged over all included survey years.[6] Each year, on average, contains 2,140 women. For the descriptive analysis, I define three age groups for the children living in a household. The first group includes children at the age range 0-2 years, the second group includes children between 3 and 5 years, and the oldest age group represents age 6 through 12. This makes it possible to distinguish pre-school age and school age. Three years is the age at which children have a legal entitlement to a place in a kindergarten in Germany (Spieß (2011), § 24 SGB VIII).

As presented in table D-1, on average, 29.7% of those living with a partner use some form of informal childcare while 44.7% use some form of formal arrangement. Almost 16% even use both forms. 41.3% do not report to use any of the two childcare sources; 13.8% only rely on informal arrangements while 29% only rely on formal arrangements. Table D-1 also reports the values separately according to the age of the youngest child. In the youngest age group, the share of those using no childcare amounts to 39.9%, in the age group 3-5 this amounts to only 11.5% while in the oldest age group it amounts to 57.5%. This is expected since the oldest age group is in school, which is not included in *formal childcare*. For the youngest age group, the value is lower, but still relatively high. This might be explained by the fact that mothers take leave from the labor market, e.g., maternity leave, in connection with childbirth. Therefore, during the first months of a child's life there is no need for regular childcare arrangement. 32.4% (42.4%) of the cohabiting women whose youngest child is 2 years tops use some form of informal (formal) childcare. If the youngest child is between 3 and 5 years old the value amounts to 37.2% (82.1%) and if the youngest child is at least six years old, it is 24.5% (26.8%).

Examining women living without a partner reveals a slightly different pattern: 40.1% use informal childcare, 46% make use of formal childcare and 20.6% use both forms while 34.9% use no childcare. The difference to the women living with a partner is most distinct for informal childcare (more than 10% points). This indicates that informal arrangements are an important source of childcare. It is likely that in couple households, the partner is an important source of childcare that is lacking in single households. This might be one reason for the higher share of single women using informal childcare compared to the share of those living with a partner.

[6] The values by survey year are presented in the appendix.

Table D-1: basic sample information (averaged over the observed years)[7]

youngest child	group	partner		no partner	
all	all	1888		252	
	using informal childcare	561	29.7 %	101	40.1 %
	using formal childcare	849	44.7 %	116	46.0 %
	using both forms	301	15.9 %	52	20.6 %
	using none	779	41.3 %	88	34.9 %
	only informal	260	13.8 %	49	19.4 %
	only formal	548	29.0 %	64	25.4 %
0-2	all	469		35	
	using informal childcare	152	32.4 %	15	42.9 %
	using formal childcare	199	42.4 %	14	40.0 %
	using both forms	69	14.7 %	6	17.1 %
	using none	187	39.9 %	14	40.0 %
3-5	all	487		61	
	using informal childcare	181	37.2 %	28	45.9 %
	using formal childcare	400	82.1 %	51	83.6 %
	using both forms	150	30.8 %	25	41.0 %
	using none	56	11.5 %	7	11.5 %
6-12	all	932		156	
	using informal childcare	228	24.5 %	59	37.8 %
	using formal childcare	250	26.8 %	52	33.3 %
	using both forms	82	8.8 %	22	14.1 %
	using none	536	57.5 %	67	42.9 %

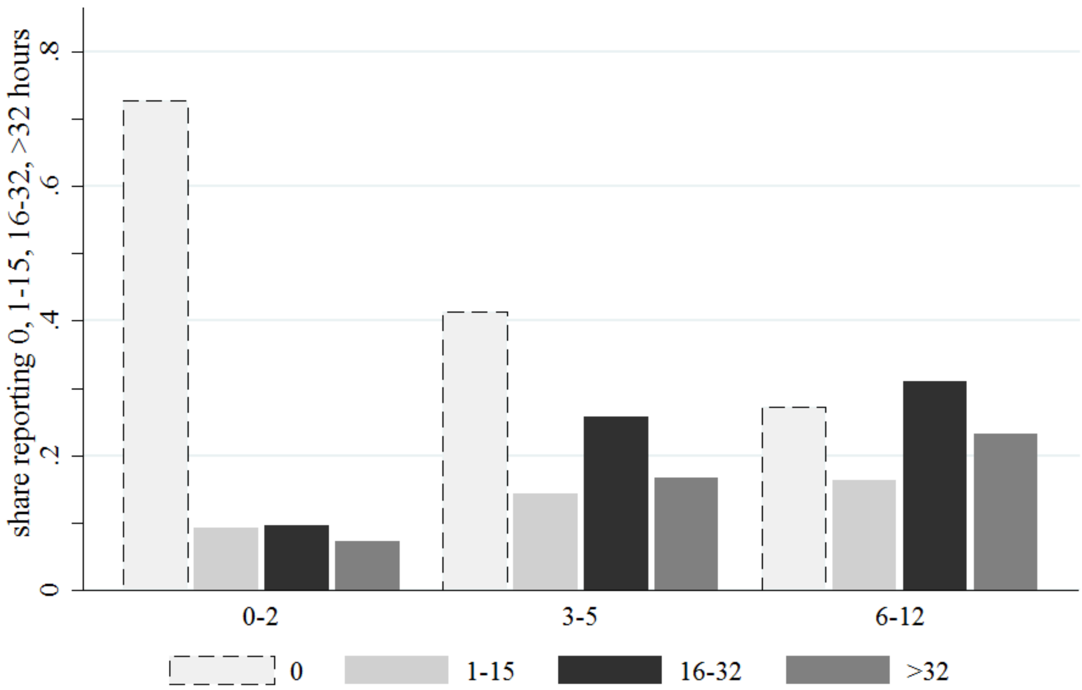
Source: GSOEP 1999-2012, excluding 2003[8], own calculations

If the youngest child is no older than 2 years, 42.9% use informal childcare, 40% use a formal arrangement. The values for the age group 3-5 amount to 45.9%, resp. 83.6%. In the age group 6-12 the shares are 37.8% and 33.3%. Even though the group sizes (averaged over time) in the latter cases are quite small, it must be mentioned that the vast majority of pre-school aged children older than 2 years are in formal childcare. This is the case for both single and cohabiting women and also is in line with the rates reported in official statistics (e.g., Federal Statistical Office of Germany (2014)). Since the youngest child can be considered the "bottleneck", figure D-1 represents the extent of employment according to the age of the youngest child. The extent is measured by the actual working hours per week. The following categories are defined: not employed, up to fifteen hours a week, 16-32 hours a week and more than 32 hours a week. To define this, the generated information of the GSOEP about actual

[7] Of course the values are rounded to integral numbers.
[8] The year 2003 is not included since for this year I have no information about informal childcare.

working hours is used. I do not use the contractual working hours since these often are not fixed (e.g., for the self-employed). The figure includes all mothers in the sample in all waves they are observed. This means that one mother can be represented in more than one age group of the youngest child over time, so the figure gives information about the average situation during the observation period. The findings are as expected. The share of mothers not working amounts to more than 70% if the youngest child is in the age group 0-2. If the youngest child is aged 3-5, part-time employment becomes common (categories 1-15, 16-32) and is even higher when the youngest child is at the oldest age group. Non-participation remains at a remarkable level over all groups, it never drops far below 30%. Full-time employment is rather low for group 1, amounting to somewhat more than 15% for group 2 and reaching 23% for group 3. While it exceeds marginal employment once the youngest child is at least three years old, it is less common than part-time employment in all groups.

Figure D-1: share of mothers working to different extents (hour categories) by age group of youngest child)[9]



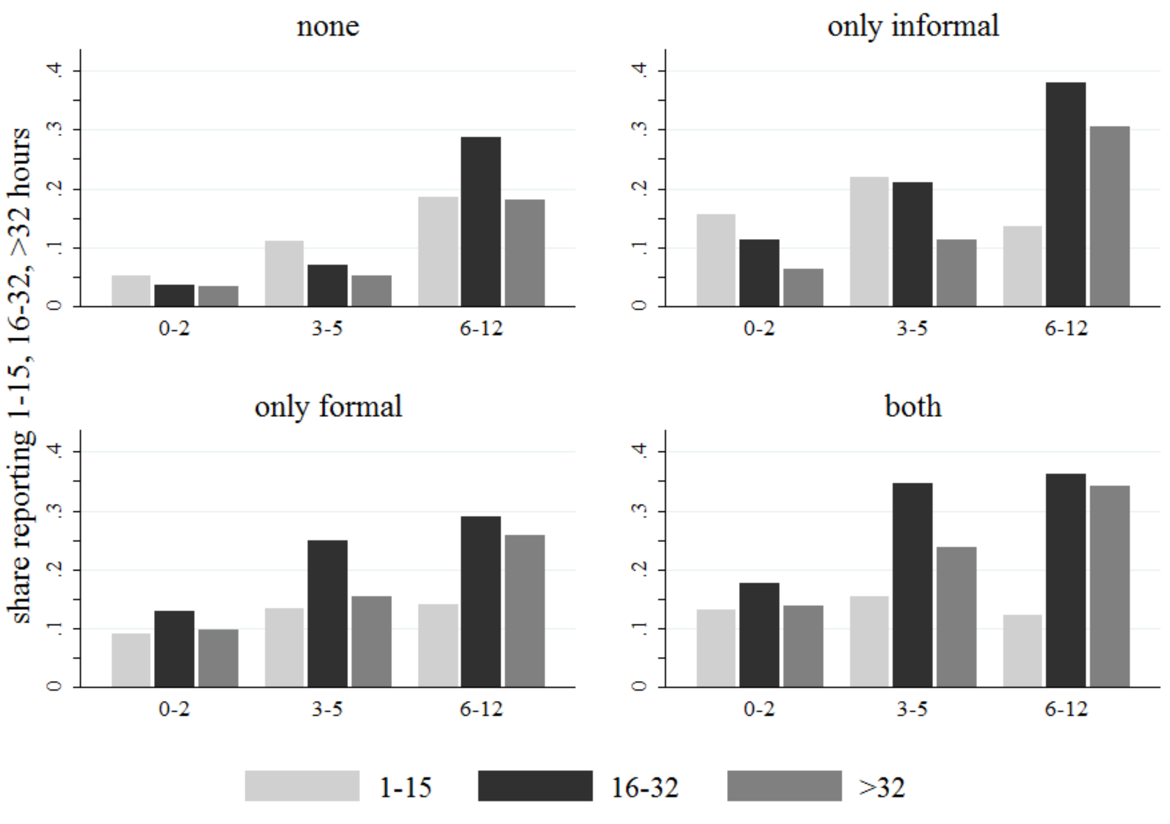
Source: GSOEP, 1999-2012, excluding 2003, own calculations

Figure D-2 distinguishes between women who receive help through informal childcare and those who do not. Apart from informal childcare, formal childcare is a major source of childcare. For this reason, the figure further distinguishes between those who receive informal support in addition to formal childcare and those who solely rely on informal arrangements. For

[9] The values contained in the figure are presented in the appendix.

reasons of clarity, only working women are presented in figure D-2.[10] First take a look at the group with the youngest child aged 0-2. All categories are more strongly represented among the women who use informal childcare arrangements in addition to formal arrangements than among women who solely rely on formal arrangements. 16-32 hours is the most prevalent employment category, both if solely formal childcare or if additional informal childcare is used. If informal childcare is the only source, “16-32” and “>32” hours are less prevalent than if only formal childcare is used while “up to 15 hours” is more prevalent than if formal childcare is the only form used. Of course, if no out-of-home childcare is used, the share of mothers working is very small.

Figure D-2: share of mothers working to different extents (hour categories, by age group of youngest child and combination of childcare arrangements)[11]



Source: GSOEP, 1999-2012, excluding 2003, own calculations

Considering the age group 3-5 reveals the following. These children are still not attending school, but there is a legal right to a kindergarten spot (§ 24 SGB VIII). For those using no childcare arrangement, the picture is slightly different. Those who are participating mainly work up to 15 hours. Comparing those who only rely on informal childcare to those only using formal

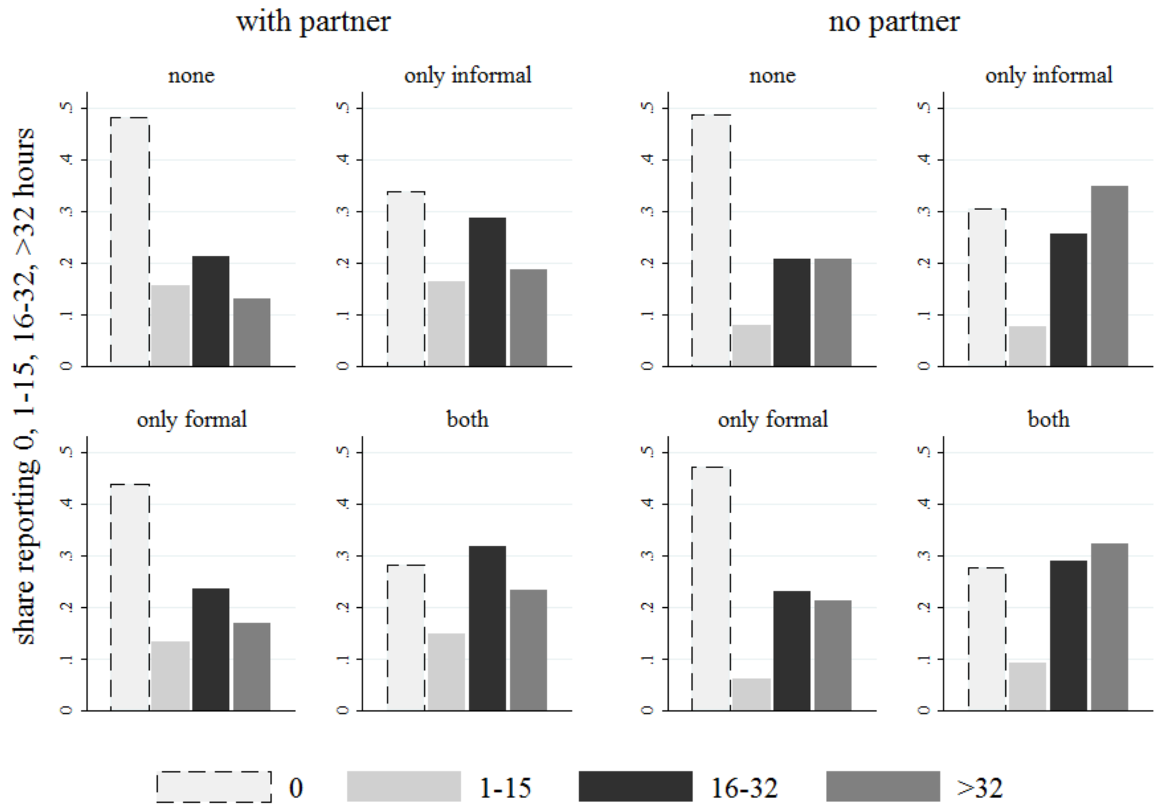
[10] A version with the women who are not working is presented in the appendix.
[11] The values contained in the figure are presented in the appendix.

arrangements reveals that the category 1-15 is more common among women who only use informal childcare while the categories 16-32 and more than 32 hours are more common among those only using formal arrangements. If both forms are used, all three categories are more prevalent. If the youngest child is at the oldest considered age group, the employment pattern of mothers is considerably different. Even if no childcare arrangement is used, all categories are more strongly represented than if the youngest child is at a younger age group. But still, (quasi) full-time employment (>32 hours) is low at 18.1%. The category 16-32 is much more strongly represented if both forms are used. The share working between 16 and 32 hours is 7.3% points higher if informal childcare is used additionally. The difference concerning (quasi) full-time employment is also notable at 8.4% points. This indicates that informal childcare might foster participation and increase the extent of employment.

Figure D-3 distinguishes between women living with a partner and those living without a partner. This does not necessarily mean they do not have one, but if they do, they are not cohabiting. If both forms of childcare arrangements are used, the non-participation rate of cohabiting and of single women are comparable. Marginal employment (1-15 hours) is remarkably higher for the cohabiting women, while the opposite is true for (quasi) full-time employment. If only informal childcare is used, the categories 1-15 and 16-32 hours are more prevalent among cohabiting women while (quasi) full-time employment is much more prevalent among singles. If the only source of childcare is the formal type, (quasi) full-time employment is less common among cohabiting women while marginal employment (1-15 hours) is more common among this group. The share of women not participating is slightly higher among single mothers. In case women do not use any of the two sources of childcare, non-participation is comparable for both groups. Again, 1-15 hours are more prevalent among cohabiting women. The category ">32 hours" is more common if no partner is living in the same household. Overall, this might indicate that cohabiting women participate to a comparable extent as their single counterparts but that they work fewer hours. Full-time employment (or quasi full-time) is less common for cohabiting mothers compared to single mothers irrespective of the forms of childcare they do or do not use.

Overall, the descriptive analysis indicates that informal arrangements may be an important source of childcare, either as the only source or as a supplement to formal childcare that might facilitate maternal labor market involvement. In general, single mothers and cohabiting mothers participate to a comparable extent, while single mothers seem to work more hours than do cohabiting women.

Figure D-3: extent of employment of women living with / without a partner[12]



Source: GSOEP, 1999-2012, excluding 2003, own calculations

D.5 Empirical methods and results

D.5.1 Methods

Participating in the labor market resp. extending the working hours is a decision that underlies clear time constraints. The weekly time a mother can devote to market work is strictly related to other time uses, such as other home-related chores, leisure and childcare. Overall, the total time devoted to all different time uses strictly sums up to 168 hours. If the mother benefits from either formal childcare, childcare of the father or cohabiting relatives or informal childcare arrangements outside the household, her own time devoted to childcare might decrease. The time devoted to other chores or leisure increases as a result. The time can be used for individual leisure, other home-related work, or for employment.[13] The data only contains information about whether or not informal arrangements are used, but not about its amount. Therefore, I investigate whether the presence of a regular informal childcare arrangement outside the

[12] The values contained in the figure are presented in the appendix.
[13] When investigating formal childcare, its costs often are considered to raise the mothers reservation wage (e.g., Wrohlich (2004)). This is less likely when it comes to informal childcare. I do not consider babysitters to be a problem since in the GSOEP, people are explicitly asked for “carer: paid care“.

household results in a higher participation and in more working hours. Informal childcare is expected to raise maternal employment as well as working hours.

The use of informal childcare arrangements is most likely not exogenous in respect to participation. Mothers might first decide to work, and because of this decision, arrange some informal childcare on a regular basis. This is not considered to be a severe problem when it comes to working hours. This would be more likely if the amount of informal childcare were under investigation, but not if its presence per se is the subject of investigation. Another challenge is that working hours are only observed for those who are actually working. The decision about working hours is unlikely to depend on the same process as the participation decision. This results in a sample that is likely characterized by selection into employment. While it is unlikely that both decisions are based on the same process, it also is unlikely that they are completely independent from each other (e.g., Cameron and Trivedi (2005)). To account for these characteristics adequately, I divide the analysis into two models.

Model 1 - participation

The first model investigates the role of informal childcare arrangements in respect to participation. To account for the potential problem of endogeneity, I use an instrument that has also been mentioned in the literature (e.g., Aassve et al. (2011)), namely the information about whether the grandmother is still alive.[14] This fact is considered to have a distinct, positive effect on the likelihood of receiving support through informal childcare but to have no direct effect on the employment decision. One might argue that care for elderly is a potential issue, but since I focus on families with rather young children, this is not considered to be an issue. Of the whole sample, only 6.8% spend time on care, 6.5% spend time on care and have a living mother. Overall, 2,860 women have a child in the youngest age group at some point in time, and only 3.3% of them spend some time on care. The values amount to 4.8% for those with the youngest child at age 3-5 and to 8.2% for those whose youngest child is in the age group 6-12.[15] In addition, this includes all types of care, not only the care for the own mother.

The dummy *"mother alive"* almost never varies over time. And since I investigate mothers of rather young children, most grandmothers are still alive. This indicates that the instrument is to be handled with care. Therefore, I also estimate the model using an additional instrument. This is the lagged information about informal childcare. Since the use of such instruments is not undisputed (e.g., Angrist and Krueger (2001)), I do not focus on these results.

[14] This is drawn from biography information of the GSOEP as well as from information about changes of the family situation.

[15] Note that due to the data being panel each mother might be contained in all of the three groups over time.

Participation is a binary variable (y_{1i}). The variable that has to be instrumented, namely informal childcare, is binary itself (y_{2i}). A general model for a binary outcome with an endogenous regressor can be described as follows (Wooldridge (2010), Cameron and Trivedi (2005))[16]:

$$y_{1i} = 1[x_i'\beta_1 + \gamma_1 y_{2i} + u_{1i} > 0] \tag{D-1}$$

$$y_{2i} = 1[x_i'\beta_2 + \delta z_i + u_{2i} > 0] \tag{D-2}$$

Both y_{1i} and y_{2i} are observed outcomes of latent variables (Cameron and Trivedi (2005)):

$$y_{1i}^* = x_i'\beta_1 + \gamma_1 y_{2i} + u_{1i} \tag{D-3}$$

$$y_{2i}^* = x_i'\beta_2 + \delta z_i + u_{2i} \tag{D-4}$$

Applying a two-stage least squares procedure (2SLS) to this is widely discussed in Angrist and Pischke (2009). An application is found in, e.g., Dimova and Wolff (2008). The advantage of using 2SLS is that there is no need for distributional assumptions about the error terms (Angrist and Pischke (2009)). They argue that a just-identified 2SLS is approximately unbiased. At least one variable (z_i) must be included in (D-4) that is unequal to zero and not included in (D-3) (just-identified). If there are two or more such variables and one endogenous variable, the model is over-identified (e.g., Angrist and Pischke (2009), Wooldridge (2010)). As stated above, I use the information about whether the grandmother is still alive. According to Angrist and Pischke (2009), the results of 2SLS reveal what is called the *local average treatment effect (LATE)*. I use cluster-robust standard errors for the pooled estimation. In addition, EC2SLS and G2SLS (Baltagi (2008)) results are reported.

Model 2 – working hours

The second step is to investigate the working hours for those who are working. As mentioned previously, this process neither is the same as the participation decision nor is it expected to be completely independent from it. This indicates that a Heckman selection correction is to be used (e.g., Cameron and Trivedi (2005), Heckman (1979)). As an exclusion restriction, I use the female unemployment rate of the federal state of residence. This variable is included in the *participation-model* but not included in the *hours-model*.[17] The information is available from the Federal Employment Agency (2013). Only working women are included, which results in a selected sample. This even reduces the variety in the instrumental variable. Given that the

[16] The notation is adapted to be consistent.
[17] A list of the variables included in each equation can be found in the appendix.

only instrumental variable is a dummy with little variation and only one exclusion restriction for the selection model, I apply the following procedure. For the selection correction, I estimate the participation equation using a basic probit model $P(y_{1i} = 1|x_i, z_i) = \Phi(x_i'\beta_1 + \delta_1 z_i)$ (e.g. Wooldridge (2010)) to obtain the inverse Mill's ratio. In this model, the dummy *"mother alive"* is included instead of the dummy *"informal childcare"* (reduced form). The inverse Mill's ratio then is included as an additional regressor in the estimation of the working hours for those who are working (Heckman selection model: e.g., Cameron and Trivedi (2005), Heckman (1979)). Informal childcare is treated as exogenous in this model. It is reasonable to assume that the endogeneity is mainly a problem in the first model. The decision to work might influence the probability of using some form of informal childcare. But once this decision is made, the number of hours likely does not influence the use of informal childcare per se, which is the only available information. This results in (e.g., Wooldridge (2010)):

$$w_{1i} = x_{1i}'\beta_3 + \gamma_3 y_{2i} + \rho\hat{\lambda}_i + u_{3i} \tag{D-5}$$

w_{1i} represents the actual (positive) weekly working hours and $\hat{\lambda}_i$ is defined as:

$$\lambda\left(x_i'\hat{\beta}_1, \hat{\delta}_1 z_i\right) = \frac{\phi(x_i'\hat{\beta}_1 + \hat{\delta}_1 z_i)}{\Phi(x_i'\hat{\beta}_1 + \hat{\delta}_1 z_i)} \tag{D-6}$$

with $\phi(x_i'\hat{\beta}_1 + \hat{\delta}_1 z_i)$ being the standard normal density function (pdf) and $\Phi(x_i'\hat{\beta}_1 + \hat{\delta}_1 z_i)$ being the standard normal distribution function (cdf) (e.g., Cameron and Trivedi (2005)). The inverse Mill's ratio is derived from a regular probit model of y_{1i} on x_i and z_i. x_i from the probit model and x_{1i} in equation (D-5) should not be identical. Therefore, I include the female unemployment rate of the federal state of residence as an exclusion restriction in the probit model but not in (D-5) (e.g. Cameron and Trivedi (2005)). This is expected to influence the likelihood of employment but not the amount of hours worked. Additionally, as discussed by Wooldridge (1995, 2010), separate probit models for each year are estimated so that ρ may vary over time to take into account the panel structure.[18]

Much like informal childcare, formal childcare may also be endogenous. This is not the focus of this analysis. However, I also estimate a pooled two-stage least squares model with both informal and formal childcare being considered as endogenous. The lagged existence of formal childcare is used as an instrument for contemporary formal childcare. Informal childcare is

[18] I also estimate an alternative specification with added time averages of appropriate explanatory variables to specify individual effects (Wooldridge (2010), Dustman and Rochina-Barrachina (2007)). The results are comparable.

instrumented using the lagged information and the information of whether the mother is still alive.

D.5.2 Results

D.5.2.1 Participation

The results concerning participation are presented in tables D-2 and D-3. They correspond with expectations. The use of informal childcare is expected to have a positive effect on the labor force participation of women. In addition to the covariates presented in the tables, I include year dummies, taking 1999 as the reference year.[19] The estimation reveals a weakly significant and clearly positive effect of informal childcare on the likelihood of participation. The standard errors are cluster-robust. The effect resulting from the 2SLS model (0.372) is distinctly higher than that found using OLS estimation (0.129).[20] As discussed above, the instrument has a weakness, namely its lack of comprehensive variation over time as well as between individuals. Therefore, the effect size revealed by the 2SLS should be interpreted with caution. I estimated the same model including another instrument: the lagged endogenous variable. The results are presented in the appendix. This reveals an effect of 0.184. The use of this instrument is not undisputed (e.g., Angrist and Krueger (2001)). Nonetheless, the 2SLS results reveal that OLS most likely underestimates the effect of informal childcare on the probability of maternal labor market participation. The effect clearly is positive and also distinct, since all estimations reveal an effect that is larger than 12 percentage points. This result is in line with the literature. An underestimation of the effect by the use of OLS is also found in other studies taking into account the endogeneity of informal childcare (e.g., Aassve et al. (2011), Posadas and Vidal-Fernández (2012)).

The coefficients for formal childcare are positive and highly significant, both in the outcome equation and in the first stage equation. As mentioned earlier, formal childcare might be endogenous as well.[21] Age is significant and positive in the 2SLS model. Years of education are positively associated with the probability of being employed. Living with a partner is negatively related to the probability of receiving out-of-home informal help, which is expected since the partner might be an important source of childcare, superseding other informal sources.

[19] I do not include school attendance since it is compulsory and therefore highly correlated with the age of the children. However, including it does not change the results and conclusions. As expected, its coefficient is positive.
[20] The OLS results can be found in the appendix.
[21] If formal childcare is modeled to be endogenous as well, the estimated effect of informal childcare on participation is of the same size as if formal childcare is treated as exogenous, while the effect of formal childcare is slightly stronger.

But it shows a positive coefficient for participation. This also is expected since the partner might provide childcare and hence more flexibility. Living in a multigenerational household is insignificant in both stages.

Table D-2: results of 2SLS estimation[22]

	participation			informal cc		
informal childcare	0.372	*	(0.192)	-		
formal childcare	0.169	***	(0.014)	0.064	***	(0.007)
age	0.020	***	(0.006)	0.008		(0.006)
age squared	-0.001	***	(0.000)	0.000	***	(0.000)
years of education	0.027	***	(0.002)	0.007	***	(0.002)
living with a partner	0.061	**	(0.024)	-0.111	***	(0.015)
multigenerational household	-0.005		(0.032)	-0.039		(0.030)
age of youngest child	0.056	***	(0.002)	-0.006	***	(0.002)
number of children 0-12	-0.019	**	(0.008)	-0.029	***	(0.007)
experience part-time	0.048	***	(0.003)	0.014	***	(0.001)
experience full-time	0.028	***	(0.002)	0.011	***	(0.001)
East-Germany	0.046	**	(0.020)	0.033		(0.023)
German citizenship	-0.026		(0.021)	0.079	***	(0.015)
unemployment rate (f)	-0.006	***	(0.002)	-0.002		(0.002)
mother alive	-			0.086	***	(0.015)
constant	-0.557	***	(0.121)	0.286	**	(0.107)
observations	24,527					
number of women	5,010					
R-squared	0.327			0.068		
Cragg-Donald Wald F	61.405					

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations

cluster-robust standard errors in parantheses

The age of the youngest child in the household shows a positive coefficient in the participation equation, indicating that participation is more prevalent among mothers of older children. This is in line with the descriptive analysis. The coefficient is negative in the first stage, which is expected since older children need less care. The number of children has a negative coefficient in both stages. Experience, either part-time or full-time, is positively associated with informal childcare as well as with participation. Living in East Germany shows a positive coefficient in the second stage. German citizenship is associated with a higher probability of using informal childcare but seems not to be related to the participation probability. This is in line with expectations because, for those women, it is more likely that their mothers live nearby. The female unemployment rate shows what is expected: a higher rate is associated with a lower

[22] The results of the instrumental variable estimations are obtained using the Stata command *ivreg2*, provided by Baum et al. (2010).

likelihood of participation. The dummy indicating whether the mother is still alive is highly significantly associated with a higher likelihood of using informal childcare.

In addition, I estimate a bivariate probit model. The bivariate probit accounts for both indicators, namely participation and the use of informal childcare, being binary. The bivariate probit model is discussed in Wooldridge (2010) and Angrist and Pischke (2009). The results obtained via 2SLS and bivariate probit cannot be directly compared in their magnitude, since 2SLS gives the *LATE,* while with the bivariate probit model, it is easy to calculate the *average treatment effect (ATE)* or the *average treatment effect on the treated (TOT)*, assuming a bivariate normal distribution (Angrist and Pischke (2009)). If a bivariate probit is applied, the baseline remains unchanged. I focus on the 2SLS estimation for the reason of weaker assumptions (e.g., Angrist and Pischke (2009)). The results of the bivariate probit model are presented in the appendix.

Table D-3: results EC2SLS and G2SLS – participation [23]

	EC2SLS				G2SLS			
	participation		informal cc		participation		informal cc	
informal childcare	0.380	***	-		0.394	***	-	
	(0.054)				(0.125)			
formal childcare	0.165	***	0.016		0.163	***	0.066	***
	(0.007)		(0.012)		(0.010)		(0.006)	
mum alive	-		0.076	***	-		0.084	***
			(0.011)				(0.011)	
observations	24,527							
number of clusters	5,010							
Wald χ^2	11,977.19				11,829.99			
Cragg-Donald Wald F[24]	5.48				53.89			

*** p<0.01, ** p<0.05, * p<0.1
standard errors in parantheses

Source: GSOEP, 1999-2012, excluding 2003, own calculations

Employing the panel methods reveals the results presented in table D-3. The results of the two estimators are essentially the same. I focus on the G2SLS results. The presence of informal childcare arrangements leads to a significantly higher probability of being in the labor force. The coefficient of the G2SLS is slightly higher than that obtained using the EC2SLS. The EC2SLS is a weighted version of the within estimator and the between estimator (Baltagi (2008)). The results are comparable to those obtained by the pooled 2SLS estimation, which is why they are not discussed in detail.

[23] The complete results are presented in the appendix. Note that the EC2SLS transforms each variable twice in the first stage. The complete results of the first stage are also presented in the appendix.
[24] Statistic obtained using *xtoverid*, provided by Schaffer and Stillman (2010).

To sum up, the estimations, irrespective of the choice of the empirical model, reveal a positive impact of informal childcare arrangements on the likelihood of female labor force participation. However, the precise size of this effect requires further research, e.g. by the use of stronger resp. less controversially discussed instruments.

D.5.2.2 Working hours

Deciding whether to participate in the labor market is only one facet of maternal employment. This section investigates the effect of informal childcare arrangements on the working hours of mothers. Therefore, only working mothers are included. As discussed above, I use a Heckman selection correction to correct for the selection into employment. Since I use the inverse Mill's ratio (IMR) obtained from the selection equation as an additional regressor, I use bootstrapped standard errors (e.g., Cameron and Trivedi (2005)). The IMR is derived from a probit model in which the endogenous regressor *informal childcare* is replaced by its instrument, namely *mother alive* (reduced form). The estimation is implemented using the Stata-implemented Heckman's two-step procedure (Heckman (1979)). Again, a set of year dummies is also included. The results are presented in tables D-4 and D-5.

The coefficient of informal childcare is positive and highly significant. However, the extent of the effect is quite small, amounting to slightly more than one hour each week. This result indicates that informal childcare does not facilitate maternal full-time instead of part-time employment. Since I refer to actual rather than contractual working hours, the results imply that informal childcare makes it easier for mothers to work overtime or to stay at work longer than they could otherwise.[25] For example, if mothers must pick up their children at some childcare institution, the grandmother might instead collect the children so that the mother can stay at work. This result is in line with the findings of Compton and Pollack (2014) regarding proximity. They also find a positive impact on the participation decision, but not on the hours-decision. Formal childcare also shows a highly significant and positive coefficient, which is considerably higher than that of informal childcare. This is expected, since informal childcare is often an additional source of childcare beyond formal arrangements. Age is slightly negatively related to working hours. Years of education are positively associated with working hours, which is clearly expected. Women living with a partner seem to work less hours than their single counterparts, which is in line with the descriptive analysis. The age of the youngest child and the number of children show expected signs, age is positively associated with the

[25] If instead the agreed working hours are investigated, this leads to an even smaller effect.

working hours while the number of children shows a negative coefficient. Labor market experience, both full-time and part-time, is related to slightly more hours. East German women work more hours. This is expected since full-time employment with even young children in the household is much more common in East Germany than in West Germany (e.g., Keller and Haustein (2013)). German citizenship seems to be negatively correlated with working hours. The inverse Mill's ratio is significant and positive.

Table D-4: results of Heckman two-step estimation – working hours[26]

	hours			working		
informal childcare	1.178	***	(0.201)	-		
formal childcare	3.696	***	(0.319)	0.644	***	(0.022)
age	-0.717	***	(0.167)	0.039	***	(0.015)
agesq	0.000		(0.002)	-0.002	***	(0.000)
years of education	1.211	***	(0.051)	0.113	***	(0.004)
partner	-3.471	***	(0.298)	0.096	***	(0.031)
multigenerational	0.828		(0.786)	-0.050		(0.071)
age of youngest child	1.260	***	(0.072)	0.196	***	(0.004)
number of children 0-12	-1.818	***	(0.172)	-0.116	***	(0.016)
experience part-time	0.139	***	(0.060)	0.223	***	(0.004)
experience full-time	0.963	***	(0.041)	0.117	***	(0.003)
East-Germany	8.219	***	(0.245)	0.242	***	(0.046)
German citizenship	-2.905	***	(0.366)	0.019		(0.033)
inverse Mill's ratio	4.074	***	(0.665)	-		
unemployment rate (f)	-			-0.026	***	(0.005)
mother alive	-			0.149	***	(0.038)
constant	22.278	***	(3.347)	-2.670	***	(0.264)
observations	24,084					
Wald χ^2	5,251.07					
χ^2 (λ)	37.50					
Prob > χ^2	0.000					

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations
standard errors in parantheses

Using probit models for each year reveals the results presented in table D-5. For each period, a probit model is estimated to derive the inverse Mill's ratio. These estimates therefore are not presented. According to Wooldridge (1995, 2010), a pooled OLS is then run on the selected sample. One can draw the same conclusion from these results as from those obtained by the

[26] The Stata-implemented Heckman two-step procedure only allows for cluster-robust inference if using the bootstrap. If this is done, or the Maximum Likelihood estimation or manual selection correction is applied, the results remain unchanged.

standard Heckman model. The covariates also show comparable coefficients. All inverse Mill's ratios are positive and also (highly) significant.[27]

Table D-5: working hours – selection with Probit model for each year[28]

	hours		
informal childcare	1.164	***	(0.266)
inverse Mill's ratio…1999	6.392	***	(1.332)
2000	2.735	**	(1.177)
2001	3.447	***	(1.103)
2002	2.821	**	(1.151)
2004	4.045	***	(1.200)
2005	4.290	***	(1.302)
2006	3.626	***	(1.109)
2007	3.289	***	(1.130)
2008	3.482	***	(1.319)
2009	3.484	**	(1.535)
2010	4.360	***	(1.422)
2011	3.596	***	(1.350)
2012	4.513	***	(1.475)
constant	21.886	***	(5.684)
observations	14,083		
R-squared	0.293		
χ^2 (IMR's)	31.32		
Prob > χ^2	0.003		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations
bootstrapped standard errors in parantheses

Overall, the analysis of working hours reveals a positive impact of informal arrangements. However, the extent is rather small. The analysis gives no proof that informal childcare might facilitate maternal full-time employment rather than part-time employment.

D.5.2.3 Robustness and summary of results

There are several requirements for a variable to be considered a valid instrument. First, an instrument must be uncorrelated with the residuals of the equation of interest, i.e., $Cov(z_i, u_{1i}) = 0$ (e.g., Wooldridge (2010)). Assuming that whether the mother is still alive directly influences the likelihood to receive help in the form of informal childcare without influencing the likelihood of being employed through any other channel than childcare availability is reasonable. Second, the instrument must be relevant, meaning the coefficient of z_i is unequal to zero ($Cov(z_i, y_{2i}) \neq 0$, e.g., Wooldridge (2010)). The first-stage estimates

[27] Including time averages as discussed in Wooldridge (1995, 2010) doesn't change the results either.
[28] The complete results are presented in the appendix.

indicate that this requirement is met by the chosen instrument as well. Nonetheless, the instrument has a weakness. It is a dummy and it does not vary greatly over time and between individuals. Most grandmothers are still alive since I investigate a sample of rather young women with children in the age group 0-12. Considering only working women even reduces the extent to which the instrumental variable varies. Therefore, I do not overwork the instrument and the analysis of working hours focuses on the results derived by the probit model(s) and the selection model, especially since endogeneity is most likely a severe problem in the participation model rather than in the hours model. This is reasonable since employment hours might be relevant for the amount of childcare used, not necessarily for the use itself. I also estimate the participation model using the lagged endogenous dummy as an additional instrument. Since this practice is discussed controversially in the literature (e.g., Angrist and Krueger (2001)), I do not focus on this specification. The literature uses several instruments. However, as Compton and Pollack (2014) discuss in respect to their instrument for informal childcare, namely proximity, these may have weaknesses as well.

However, the results concerning the participation decision indicate a positive effect of the presence of informal childcare arrangements on the probability of participating in the labor market. This is true irrespective of what specification is used. The 2SLS estimates show a remarkably higher coefficient, indicating that OLS underestimates the effect. When including the lagged dummy about informal childcare as an additional instrument, the coefficient is closer to that of the OLS estimation and is highly significant, but the effect is still is of a greater extent than that obtained via OLS.[29] The true effect most likely amounts to an extent somewhere in between the OLS result and the 2SLS result presented above. However, it is reasonable to conclude that informal childcare seems to clearly facilitate maternal employment. This is in line with previous literature (e.g., Aassve et al. (2011)).

Concerning working hours, I account for the selection into employment since the hours are observed only for those who are working. I use a Heckman approach as well as an approach suggested by Wooldridge (1995, 2010), which accounts for the panel structure. The results are comparable for both models. If informal childcare of the last year is included in the selection equation of the standard Heckman model as well, the coefficient of informal childcare in the hours equation is slightly higher at 1.408 hours. If it is included in the period-specific probit models, the coefficient of informal childcare amounts to 1.378 hours and is still highly significant. Nevertheless, the effect is small in all applied estimations. The presence of informal childcare is found to raise work time by about 1.3 hours a week, which may indicate that

[29] The results are presented in detail in the appendix.

informal childcare does not facilitate maternal full-time employment instead of part-time employment. Since I refer to actual rather than contractual working hours, the results imply that informal childcare makes it easier for mothers to work overtime or to stay at work slightly longer.[30] For example, if mothers must pick up their children at some childcare institution, this is a fixed time that cannot be extended easily. Grandmothers or other forms of informal childcare might assume this task so that the mother can stay at work longer. But, as already mentioned, the results do not imply that informal childcare facilitates full-time rather than part-time employment of mothers.

D.6 Conclusions

Maternal labor market participation as well as employment hours are of great interest in German politics (e.g., Spieß (2011), Ristau (2005)). The demographic change Germany is facing and the challenges for the economy wrought by this change (e.g., European Commission (2012), OECD (2014)) have led to the extent of market work of mothers being a widely discussed topic in German political discourse (e.g., BMFSFJ (2014), BMI (2011)). This analysis investigates whether maternal employment and working hours are increased by the presence of informal childcare arrangements. Two problems when solely relying on formal arrangements might be the lack of flexibility and the inability to cover irregular time slots (Posadas and Vidal-Fernández (2012), Compton (2011), Compton and Pollak (2014)). A grandmother might assume tasks spontaneously, i.e., if the child is sick, or they might take on other obligations on a regular basis, such as collecting the children at some institution or covering certain time slots that are not covered by formal institutions. This is much more complicated if one only relies on formal arrangements.

One empirical challenge when investigating this question is the potential endogeneity of informal childcare arrangements with respect to participation. Informal childcare is instrumented using information about whether the mother of the woman is still alive. This is expected to have a direct influence on the likelihood of having access to informal childcare, but it is not expected to influence the participation of the woman through any other channel. A challenge in the estimation of the effect on working hours is the selection into employment in the first place. The decision about whether to participate might depend on another process than the decision about hours (e.g., Cameron and Trivedi (2005)). This selection is taken into account

[30] If the agreed working hours are investigated, this leads to an even smaller effect.

when investigating the working hours using a Heckman-type correction (Heckman (1979), Wooldridge (1995, 2010)).

Overall, informal childcare seems to be an important factor that supports mothers in combining market work and family life. The results concerning labor market participation are in line with expectations and with previous literature. Maternal employment is more likely when mothers have support through social networks outside the household.

Informal childcare also seems to have a positive influence on working hours once a mother has decided to participate in the labor market. The extent of this effect is small. Using a Heckman selection correction reveals an effect of slightly more than an hour. Therefore, the results do not indicate that informal childcare arrangements are crucial for full-time employment instead of part-time employment. The presence of informal childcare arrangement rather seems to provide more flexibility within a certain work arrangement, for instance, the flexibility to stay at work slightly longer because someone else can pick up the children from a childcare institution.

Since only out-of-home childcare is investigated, this may lead to an underestimation of the importance of informal childcare. The contribution of fathers as well as cohabiting relatives is not investigated. This is done for two reasons. First, including fathers' contribution to childcare would lead to many further methodological challenges, since the time use of partners is most likely interdependent. Second, when it comes to other cohabiting relatives, e.g., grandparents, these are not only a possible source of childcare but might also be in need of care themselves. This would not relax the mothers' time constraints, it would instead further tighten them. To include all of these other kinds of informal childcare is an issue for further research. The present study is a further step towards the investigation of this source of childcare dealing with endogeneity as well as with selection. It further illustrates the positive effects of informal childcare on maternal participation in the labor market.

D.7 Appendix

Table A D-1: Sample by observation year

	group	with partner	without partner
1999		1,771	
	all	1,582	189
percentage reporting…	using informal childcare	31.7	39.7
	using formal childcare	38.5	39.2
	using both forms	14.3	15.3
	using none	44.1	36.5
	only informal	17.4	24.3
	only formal	24.2	23.8
2000		2,968	
	all	2,657	311
percentage reporting…	using informal childcare	27.9	40.2
	using formal childcare	40.1	38.9
	using both forms	13.4	19.3
	using none	45.4	40.2
	only informal	14.5	20.9
	only formal	26.6	19.6
2001		2,635	
	all	2,361	274
percentage reporting…	using informal childcare	32.4	40.5
	using formal childcare	42.1	39.8
	using both forms	15.9	19.7
	using none	41.3	39.4
	only informal	16.6	20.8
	only formal	26.3	20.1
2002		2,663	
	all	2,394	269
percentage reporting…	using informal childcare	35.4	41.3
	using formal childcare	43.1	42.8
	using both forms	17.5	21.9
	using none	39.0	37.9
	only informal	17.9	19.3
	only formal	25.6	20.8
2003		2,450	
	all	2,189	261
percentage reporting…	using informal childcare	n.a.[31]	n.a.
	using formal childcare	43.5	42.5
	using both forms	n.a.	n.a.
	using none	n.a.	n.a.
	only informal	n.a.	n.a.
	only formal	n.a.	n.a.

[31] This type of childcare was not included in the 2003 questionnaire.

Table A D-1 continued:

2004		2,313	
	all	2,059	254
percentage reporting…	using informal childcare	32.2	44.9
	using formal childcare	44.7	49.2
	using both forms	17.5	24.4
	using none	40.6	30.3
	only informal	14.7	20.5
	only formal	27.2	24.8
2005		2,174	
	all	1,899	275
percentage reporting…	using informal childcare	25.6	34.2
	using formal childcare	44.5	46.9
	using both forms	15.2	18.2
	using none	45.0	37.1
	only informal	10.4	16.0
	only formal	29.4	28.7
2006		2,329	
	all	2,025	304
percentage reporting…	using informal childcare	26.9	40.1
	using formal childcare	43.3	49.0
	using both forms	14.4	21.7
	using none	44.2	32.6
	only informal	12.5	18.4
	only formal	28.9	27.3
2007		2,150	
	all	1,870	280
percentage reporting…	using informal childcare	26.6	33.9
	using formal childcare	44.2	44.6
	using both forms	14.9	17.9
	using none	44.0	39.3
	only informal	11.8	16.1
	only formal	29.4	26.8
2008		1,936	
	all	1,668	268
percentage reporting…	using informal childcare	28.1	38.8
	using formal childcare	44.9	42.2
	using both forms	15.0	18.3
	using none	42.0	37.3
	only informal	13.1	20.5
	only formal	29.9	23.9
2009		1,959	
	all	1,690	269
percentage reporting…	using informal childcare	29.9	43.9
	using formal childcare	50.2	53.9
	using both forms	17.9	23.4
	using none	37.8	25.7
	only informal	12.0	20.4
	only formal	32.2	30.5

Table A D-1 continued:

2010		1,735	
	all	1,499	236
percentage reporting…	using informal childcare	28.2	40.3
	using formal childcare	52.2	48.7
	using both forms	17.7	18.6
	using none	37.3	29.7
	only informal	10.5	21.6
	only formal	34.6	30.1
2011		1,654	
	all	1,465	189
percentage reporting…	using informal childcare	31.9	41.3
	using formal childcare	52.3	52.9
	using both forms	18.8	27.0
	using none	34.5	32.8
	only informal	13.2	14.3
	only formal	33.5	25.9
2012	2000	1,530	
	all	1,369	161
percentage reporting…	using informal childcare	27.6	44.7
	using formal childcare	52.9	54.7
	using both forms	16.8	28.0
	using none	36.3	28.6
	only informal	10.8	16.8
	only formal	36.1	26.7

Source: GSOEP 1999-2012, own calculations

Table A D-2: mean of different employment statuses, cohabiting or single

age of youngest child	extent of employment	all	with partner	without partner
0-2	zero hours	**0.728**	0.722	0.809
	1-15 hours	**0.093**	0.096	0.046
	16-32 hours	**0.096**	0.098	0.069
	>32 hours	**0.072**	0.072	0.069
3-5	0	**0.414**	0.408	0.466
	1-15	**0.143**	0.150	0.095
	16-32	**0.257**	0.260	0.233
	>32	**0.166**	0.163	0.192
6-12	0	**0.273**	0.270	0.289
	1-15	**0.163**	0.177	0.078
	16-32	**0.310**	0.315	0.283
	>32	**0.231**	0.215	0.329

Bold values: figure D-1 Source: GSOEP 1999-2012, excluding 2003, own calculations

Table A D-3: number of observations (tables A D-4 – A D-6)

youngest child	all	formal	informal	none	both
all	27,817	7,948	4,007	11,263	4,599
0-2	**6,553**	1,780	1,181	2,608	984
3-5	**7,124**	3,591	453	815	2,265
6-12	**14,140**	2,577	2,373	7,840	1,350
cohabiting					
all	24,538	7,122	3,375	10,124	3,917
0-2	6,092	1,684	1,073	2,432	903
3-5	6,334	3,252	408	728	1,946
6-12	12,112	2,186	1,894	6,964	1,068
single					
all	3,279	826	632	1,139	682
0-2	461	96	108	176	81
3-5	790	339	45	87	319
6-12	2,028	391	479	876	282

Bold values: figure D-1 Source: GSOEP 1999-2012, excluding 2003, own calculations

Table A D-4: mean of different employment statuses

youngest child	extent	all	formal	informal	none	both
all	zero hours	0.416	0.443	0.333	0.482	0.281
	1-15 hours	0.141	0.127	0.151	0.149	0.140
	15-32 hours	0.246	0.236	0.282	0.213	0.314
	>32 hours	0.177	0.175	0.213	0.138	0.247
0-2	0	0.728	0.670	0.656	0.871	0.542
	1-15	0.093	**0.090**	**0.155**	**0.051**	**0.132**
	16-32	0.096	**0.130**	**0.113**	**0.035**	**0.177**
	>32	0.072	**0.097**	**0.063**	**0.035**	**0.138**
3-5	0	0.414	0.441	0.428	0.757	0.246
	1-15	0.143	**0.135**	**0.219**	**0.110**	**0.154**
	16-32	0.257	**0.250**	**0.210**	**0.070**	**0.346**
	>32	0.166	**0.154**	**0.113**	**0.053**	**0.237**
6-12	0	0.273	0.290	0.154	0.324	0.150
	1-15	0.163	**0.141**	**0.136**	**0.185**	**0.123**
	16-32	0.310	**0.289**	**0.380**	**0.287**	**0.362**
	>32	0.231	**0.258**	**0.306**	**0.181**	**0.342**

Bold values: figure D-2

Source: GSOEP 1999-2012, excluding 2003, own calculations

Table A D-5: mean of different employment statuses– cohabiting

youngest child	extent	all	formal	informal	none	both
all	zero hours	0.418	**0.440**	**0.338**	**0.481**	**0.282**
	1-15 hours	0.150	**0.134**	**0.164**	**0.157**	**0.148**
	16-32 hours	0.247	**0.236**	**0.287**	**0.214**	**0.319**
	>32 hours	0.166	**0.170**	**0.187**	**0.130**	**0.234**
0-2	0	0.722	0.669	0.639	0.866	0.534
	1-15	0.096	0.092	0.166	0.053	0.138
	16-32	0.098	0.131	0.118	0.036	0.183
	>32	0.072	0.096	0.063	0.037	0.134
3-5	0	0.408	0.431	0.424	0.745	0.240
	1-15	0.150	0.142	0.225	0.120	0.157
	16-32	0.260	0.254	0.211	0.071	0.350
	>32	0.163	0.152	0.108	0.054	0.234
6-12	0	0.270	0.277	0.149	0.320	0.144
	1-15	0.177	0.155	0.150	0.197	0.140
	16-32	0.315	0.290	0.399	0.291	0.375
	>32	0.215	0.256	0.275	0.170	0.316

Bold values: figure D-3

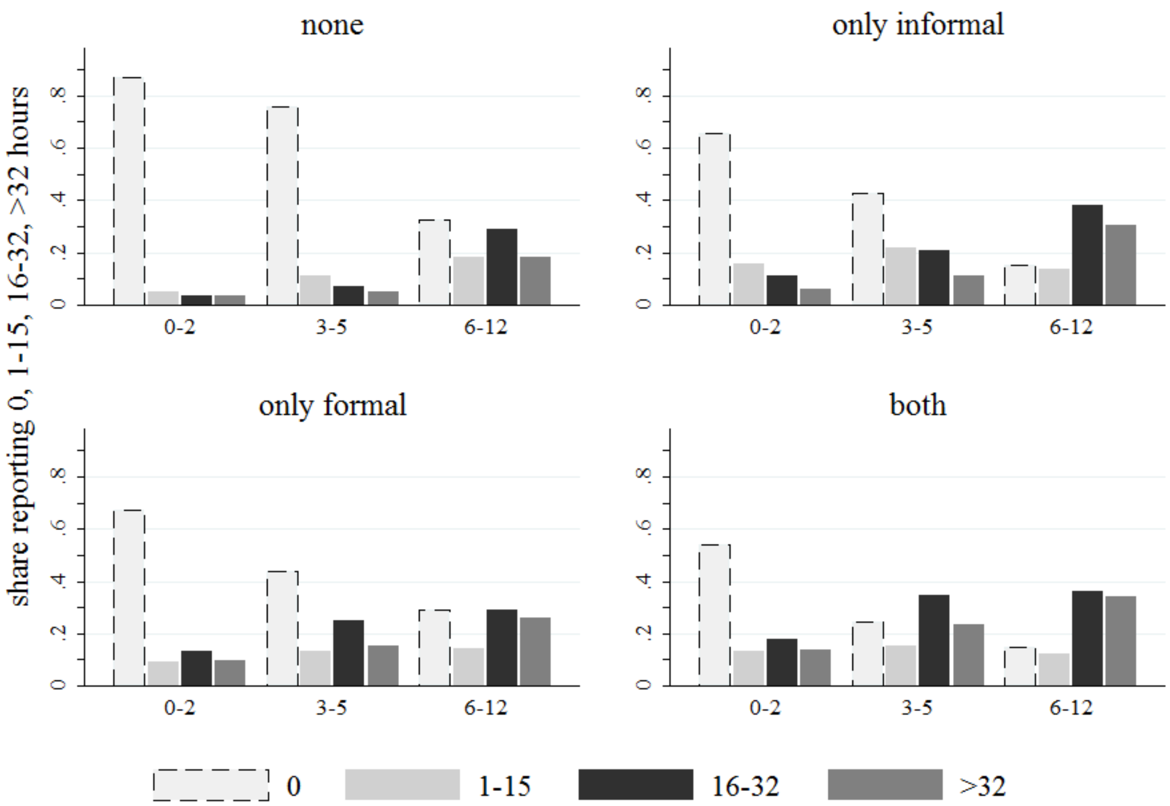
Source: GSOEP 1999-2012, excluding 2003, own calculations

Table A D-6: mean of different employment statuses– single

youngest child	extent	all	formal	informal	none	both
all	zero hours	0.405	**0.472**	**0.305**	**0.487**	**0.279**
	1-15 hours	0.078	**0.063**	**0.078**	**0.079**	**0.094**
	16-32 hours	0.240	**0.231**	**0.256**	**0.208**	**0.290**
	>32 hours	0.260	**0.213**	**0.348**	**0.207**	**0.323**
0-2	0	0.809	0.698	0.824	0.943	0.630
	1-15	0.046	0.052	0.046	0.034	0.062
	16-32	0.067	0.115	0.065	0.023	0.111
	>32	0.069	0.115	0.056	0.000	0.185
3-5	0	0.466	0.537	0.467	0.862	0.282
	1-15	0.095	0.065	0.156	0.034	0.135
	16-32	0.233	0.204	0.200	0.057	0.317
	>32	0.192	0.177	0.156	0.046	0.254
6-12	0	0.289	0.361	0.173	0.358	0.174
	1-15	0.078	0.064	0.077	0.092	0.057
	16-32	0.283	0.284	0.305	0.260	0.312
	>32	0.329	0.269	0.432	0.265	0.440

Bold values: figure D-3 Source: GSOEP 1999-2012, excluding 2003, own calculations

Figure A D-1: figure D-2 including the category "zero hours"



Source: GSOEP, 1999-2012, excluding 2003, own calculations

Table A D-7: list of variables used in each equation

	Model 1		*Model 2*	
variable	working	informal cc	hours	working
informal childcare	✓	-	✓	-
formal childcare	✓	✓	✓	✓
age	✓	✓	✓	✓
age squared	✓	✓	✓	✓
years of education	✓	✓	✓	✓
partner in the same household	✓	✓	✓	✓
multigenerational household	✓	✓	✓	✓
age of youngest child	✓	✓	✓	✓
# of children 0-12	✓	✓	✓	✓
experience part-time	✓	✓	✓	✓
experience full-time	✓	✓	✓	✓
living in East Germany	✓	✓	✓	✓
German citizenship	✓	✓	✓	✓
year dummies	✓	✓	✓	✓
inverse Mill's ratio(s)	-	-	✓	✓
female unemployment rate	✓	✓	-	✓
mother alive	-	✓	-	✓
informal childcare (t-1)	-	-/✓	-	-/✓

Data source: GSOEP, 1999-2012

Table A D-8: results simple OLS

	participation			hours		
informal childcare	0.129	***	(0.007)	1.027	***	(0.273)
formal childcare	0.180	***	(0.006)	2.486	***	(0.276)
age	0.023	***	(0.005)	-1.004	***	(0.267)
agesq	-0.001	***	(0.000)	0.007	*	(0.004)
years of education	0.028	***	(0.002)	1.011	***	(0.075)
partner	0.034	***	(0.011)	-3.529	***	(0.479)
multigenerational	-0.010		(0.027)	0.558		(1.074)
age of youngest child	0.054	***	(0.001)	0.922	***	(0.063)
number of children 0-12	-0.027	***	(0.006)	-1.566	***	(0.263)
experience part-time	0.052	***	(0.001)	-0.192	***	(0.056)
experience full-time	0.030	***	(0.001)	0.756	***	(0.043)
East-Germany	0.056	***	(0.017)	8.205	***	(0.418)
German citizenship	-0.003		(0.013)	-3.085	***	(0.691)
unemployment rate (f)	-0.006		(0.002)	-		
constant	-0.490	***	(0.094)	34.245	***	(4.827)
observations	26,767			15,252		
R-squared	0.371			0.286		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations
cluster-robust standard errors in parantheses

Table A D-9: results bivariate Probit estimation

	participation			informal cc		
informal childcare	1.423	***	(0.093)			
formal childcare	0.489	***	(0.033)	0.190	***	(0.022)
age	0.024		(0.022)	0.050	**	(0.020)
agesq	-0.001	***	(0.000)	-0.002	***	(0.000)
years of education	0.092	***	(0.008)	0.026	***	(0.007)
partner	0.241	***	(0.043)	-0.325	***	(0.041)
multigenerational	0.016		(0.107)	-0.113		(0.092)
age of youngest child	0.186	***	(0.007)	-0.017	***	(0.005)
number of children 0-12	-0.057	***	(0.022)	-0.075	***	(0.023)
experience part-time	0.178	***	(0.011)	0.048	***	(0.005)
experience full-time	0.091	***	(0.006)	0.039	***	(0.004)
East-Germany	0.175	***	(0.067)	0.082		(0.070)
German citizenship	-0.104	**	(0.048)	0.264	***	(0.050)
unemployment rate (f)	-0.021	***	(0.007)	-0.005		(0.007)
mother alive	-			0.321	***	(0.056)
constant	-2.802	***	(0.383)	-1.134	***	(0.346)
observations	24,527					
Wald χ^2	6936.53					
Wald test $\rho = 0$	46.25					
Prob > χ^2	0.000					

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations
cluster-robust standard errors in parantheses

Table A D-10: results EC2SLS and G2SLS estimators – participation

	EC2SLS				G2SLS			
	participation		informal cc		participation		informal cc	
informal childcare	0.380	***	-		0.394	***	-	
	(0.054)				(0.125)			
formal childcare	0.165	***	0.016		0.163	***	0.066	***
	(0.007)		(0.012)		(0.010)		(0.006)	
age	0.019	***	0.010	**	0.022	***	0.006	
	(0.004)		(0.005)		(0.004)		(0.004)	
agesq	-0.001	***	-0.000	***	-0.001	***	0.000	***
	(0.000)		(0.000)		(0.000)		(0.000)	
years of education	0.027	***	0.008	***	0.027	***	0.008	***
	(0.001)		(0.001)		(0.002)		(0.001)	
partner	0.061	***	-0.100	***	0.063	***	-0.111	***
	(0.010)		(0.010)		(0.016)		(0.009)	
multigenerational	-0.004		-0.001		-0.002		-0.049	**
	(0.020)		(0.023)		(0.021)		(0.022)	
age of youngest child	0.056	***	-0.012	***	0.056	***	-0.005	***
	(0.001)		(0.002)		(0.001)		(0.001)	
# children 0-12	-0.019	***	-0.044	***	-0.019	***	-0.026	***
	(0.005)		(0.006)		(0.005)		(0.005)	
experience part-time	0.048	***	0.013	***	0.047	***	0.014	***
	(0.001)		(0.001)		(0.002)		(0.001)	
experience full-time	0.027	***	0.009	***	0.027	***	0.010	***
	(0.001)		(0.001)		(0.001)		(0.001)	
East-Germany	0.047	***	0.024	*	0.045	***	0.033	**
	(0.013)		(0.014)		(0.014)		(0.014)	
German citizenship	-0.026	**	0.074	***	-0.025	*	0.079	***
	(0.010)		(0.010)		(0.014)		(0.010)	
unemployment rate (f)	-0.006	***	-0.001		-0.006	***	-0.002	
	(0.001)		(0.001)		(0.001)		(0.001)	
mum alive	-		0.076	***	-		0.084	***
			(0.011)				(0.011)	
constant	-0.533	***	-0.201	***	-0.600	***	0.318	***
	(0.077)		(0.076)		(0.084)		(0.078)	
observations	24,527							
number of clusters	5,010							
Cragg-Donald Wald F[32]	5.48				53.89			

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations
standard errors in parantheses

[32] Statistic obtained using *xtoverid*, provided by Schaffer and Stillman (2010).

Table A D-11: Complete list of instruments used for the EC2SLS estimator – participation

	EC2SLS			
	within transformation		between transformation	
formal childcare	0.063	***	0.016	
	(0.007)		(0.012)	
age	0.017	**	0.010	**
	(0.007)		(0.005)	
agesq	-0.000		-0.000	***
	(0.000)		(0.000)	
years of education	0.007		0.008	***
	(0.008)		(0.001)	
partner	-0.092	***	-0.100	***
	(0.016)		(0.010)	
multigenerational	-0.142	***	-0.001	
	(0.042)		(0.023)	
age of youngest child	0.006	***	-0.012	***
	(0.002)		(0.002)	
# children 0-12	0.006		-0.044	***
	(0.008)		(0.006)	
experience part-time	-0.005		0.013	***
	(0.003)		(0.001)	
experience full-time	0.004		0.009	***
	(0.004)		(0.001)	
East-Germany	0.131	**	0.024	*
	(0.052)		(0.014)	
German citizenship	0.048		0.074	***
	(0.043)		(0.010)	
unemployment rate (f)	-0.004		-0.001	
	(0.003)		(0.001)	
mum alive	0.049		0.076	***
	(0.034)		(0.011)	
constant	-0.201 ***			
	(0.076)			
observations	24,527			

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations

standard errors in parantheses

Table A D-12: selection correction using Probit models for each year – hours

	hours		
informal childcare	1.164	***	(0.266)
formal childcare	3.568	***	(0.390)
age	-0.727	**	(0.292)
agesq	0.001		(0.004)
years of education	1.189	***	(0.080)
partner	-3.497	***	(0.483)
multigenerational	0.887		(1.132)
age of youngest child	1.224	***	(0.094)
number of children 0-12	-1.797	***	(0.283)
experience part-time	0.107		(0.089)
experience full-time	0.942	***	(0.062)
East-Germany	8.238	***	(0.427)
German citizenship	-2.897	***	(0.704)
inverse Mill's ratio…1999	6.392	***	(1.332)
2000	2.735	**	(1.177)
2001	3.447	***	(1.103)
2002	2.821	**	(1.151)
2004	4.045	***	(1.200)
2005	4.290	***	(1.302)
2006	3.626	***	(1.109)
2007	3.289	***	(1.130)
2008	3.482	***	(1.319)
2009	3.484	**	(1.535)
2010	4.360	***	(1.422)
2011	3.596	***	(1.350)
2012	4.513	***	(1.475)
constant	21.886	***	(5.684)
observations	14,083		
adjusted R-squared	0,291		
χ^2 (IMR's)	31.32		
Prob > χ^2	0.003		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations

bootstrapped standard errors in parantheses

Results if *“informal childcare in t-1”* is included as an instrument

Table A D-13: pooled 2SLS – participation

	participation			informal cc		
informal childcare	0.184	***	(0.017)	-		
formal childcare	0.171	***	(0.008)	0.014		(0.007)
age	0.026	***	(0.007)	0.003		(0.005)
agesq	-0.001	***	(0.000)	-0.000	**	(0.000)
years of education	0.030	***	(0.002)	0.005	***	(0.001)
partner	0.046	***	(0.013)	-0.062	***	(0.011)
multigenerational	-0.019		(0.036)	-0.004		(0.030)
age of youngest child	0.052	***	(0.002)	-0.010	***	(0.001)
number of children 0-12	-0.033	***	(0.007)	-0.038	***	(0.005)
experience part-time	0.051	***	(0.002)	0.007	***	(0.001)
experience full-time	0.030	***	(0.001)	0.006	***	(0.001)
East-Germany	0.066	***	(0.019)	0.022		(0.015)
German citizenship	-0.013		(0.016)	0.052	***	(0.010)
unemployment rate (f)	-0.007	***	(0.002)	-0.002		(0.002)
mother alive	-			0.041	***	(0.011)
informal childcare (t-1)	-			0.454	***	(0.009)
constant	-0.537	***	(0.121)	0.251	**	(0.089)
observations	17,614					
number of women	4.216					
R-squared	0.363			0.273		
Cragg-Donald Wald F	2,370.83					
Hansen J statistic (p-value)	0.489					

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations

cluster-robust standard errors in parantheses

Table A D-14: EC2SLS / G2SLS – participation

	EC2SLS				G2SLS			
	participation		informal cc		participation		informal cc	
informal childcare	0.154	***	-		0.191	***	-	
	(0.011)				(0.015)			
formal childcare	0.169	***	-0.028	**	0.167	***	0.016	**
	(0.007)		(0.012)		(0.007)		(0.007)	
age	0.022	***	0.000		0.027	***	0.003	
	(0.005)		(0.005)		(0.005)		(0.005)	
agesq	-0.001	***	-0.000		-0.001	***	-0.000	***
	(0.000)		(0.000)		(0.000)		(0.000)	
years of education	0.030	***	0.002		0.030	***	0.005	***
	(0.001)		(0.001)		(0.001)		(0.001)	
partner	0.042	***	-0.017	*	0.047	***	-0.065	***
	(0.009)		(0.009)		(0.009)		(0.010)	
multigenerational	-0.021		0.024		-0.019		-0.009	
	(0.024)		(0.024)		(0.024)		(0.025)	
age of youngest child	0.051	***	-0.009	***	0.052	***	-0.010	***
	(0.001)		(0.002)		(0.001)		(0.001)	
# children 0-12	-0.035	***	-0.025	***	-0.033	***	-0.038	***
	(0.005)		(0.005)		(0.005)		(0.005)	
experience part-time	0.051	***	0.001		0.050	***	0.007	***
	(0.001)		(0.001)		(0.001)		(0.001)	
experience full-time	0.03	***	0.002	**	0.030	***	0.007	***
	(0.001)		(0.001)		(0.001)		(0.001)	
East-Germany	0.068	***	0.015		0.065	***	0.022	
	(0.014)		(0.013)		(0.014)		(0.014)	
German citizenship	-0.011		0.013		-0.012		0.055	***
	(0.010)		(0.009)		(0.011)		(0.012)	
unemployment rate (f)	-0.007	***	-0.001		-0.007	***	-0.002	
	(0.001)		(0.001)		(0.001)		(0.002)	
mum alive	-		0.009		-		0.043	***
			(0.011)				(0.012)	
informal childcare (t-1)	-		0.812	***	-		0.424	***
			(0.008)				(0.007)	
constant	-0.448	***	0.511	***	-0.558	***	0.277	***
	(0.089)		(0.113)		(0.084)		(0.088)	
observations	17,641							
number of women	4,216							
Wald χ^2	9,429,36				9,357.98			
Cragg-Donald Wald F[33]	202.88				2,001.38			
Sargan Hansen (p-value)	0.000				0.235			

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations

standard errors in parantheses

[33] Both the Cragg-Donald Wald F and the Sargan-Hansen statistic are obtained using the Stata command *xtoverid*, provided by Schaffer and Stillman (2010).

Table A D-15: complete first stage of EC2SLS – participation

	EC2SLS			
	within transformation		between transformation	
formal childcare	0.041	***	-0.028	**
	(0.007)		(0.012)	
age	-0.005		0.000	
	(0.008)		(0.005)	
agesq	-0.000		-0.000	
	(0.000)		(0.000)	
years of education	-0.009		0.002	
	(0.012)		(0.001)	
partner	-0.084	***	-0.017	*
	(0.016)		(0.009)	
multigenerational	-0.137	***	0.024	
	(0.046)		(0.024)	
age of youngest child	0.003		-0.009	***
	(0.002)		(0.002)	
# children 0-12	-0.008		-0.025	***
	(0.008)		(0.005)	
experience part-time	-0.003		0.001	
	(0.003)		(0.001)	
experience full-time	0.004		0.002	**
	(0.004)		(0.001)	
East-Germany	0.162	***	0.015	
	(0.052)		(0.013)	
German citizenship	0.087	**	0.013	
	(0.043)		(0.009)	
unemployment rate (f)	-0.005		-0.001	
	(0.003)		(0.001)	
mum alive	0.019		0.009	
	(0.032)		(0.011)	
informal childcare (t-1)	0.031	***	0.812	***
	(0.008)		(0.008)	
constant	0.511 ***			
	(0.112)			
observations	17,614			

*** p<0.01, ** p<0.05, * p<0.1
standard errors in parantheses

Source: GSOEP, 1999-2012, excluding 2003, own calculations

Table A D-16: pooled Heckman – both instruments in the selection equation (reduced form) – hours[34]

	hours			working		
informal childcare	1.408	***	(0.231)	-		
formal childcare	3.270	***	(0.339)	0.601	***	(0.028)
age	-0.795	***	(0.194)	0.046	**	(0.019)
agesq	0.003		(0.003)	-0.002	***	(0.000)
years of education	1.143	***	(0.056)	0.116	***	(0.005)
partner	-3.521	***	(0.331)	0.156	***	(0.036)
multigenerational	0.268		(0.947)	-0.042		(0.091)
age of youngest child	1.134	***	(0.074)	0.191	***	(0.005)
number of children 0-12	-1.822	***	(0.196)	-0.150	***	(0.019)
experience part-time	0.016		(0.063)	0.224	***	(0.005)
experience full-time	0.865	***	(0.044)	0.119	***	(0.003)
East-Germany	8.513	***	(0.272)	0.296	***	(0.053)
German citizenship	-2.916	***	(0.414)	-0.018		(0.039)
inverse Mill's ratio	2.174	***	(0.699)	-		
unemployment rate (f)	-			-0.030		(0.006)
mother alive	-			0.124	***	(0.046)
informal childcare (t-1)	-			0.300	***	(0.026)
constant	25.621	***	(3.916)	-2.880	***	(0.335)
observations			17,302			
Wald χ^2			4131.13			
χ^2 (λ)			9.68			
Prob> χ^2			0.002			

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations
standard errors in parantheses

[34] The Stata implemented Heckman two-step procedure only allows for cluster-robust inference if using the bootstrap. If this is done the results do not change.

Table A D-17: selection with Probit model for each year – working hours

	hours		
informal childcare	1.378	***	(0.320)
formal childcare	3.206	***	(0.429)
age	-0.803	**	(0.330)
agesq	0.003		(0.005)
years of education	1.131	***	(0.100)
partner	-3.543	***	(0.569)
multigenerational	0.312		(1.413)
age of youngest child	1.116	***	(0.107)
number of children 0-12	-1.801	***	(0.288)
experience part-time	0.000		(0.097)
experience full-time	0.854	***	(0.073)
East-Germany	8.512	***	(0.495)
German citizenship	-2.940	***	(0.854)
inverse Mill's ratio…2000	1.997		(1.398)
2001	2.150	*	(1.157)
2002	1.599		(1.178)
2005	2.883	**	(1.231)
2006	2.097	**	(1.066)
2007	1.648		(1.139)
2008	2.005		(1.499)
2009	2.294		(1.804)
2010	2.788	*	(1.635)
2011	-0.048		(1.549)
2012	2.121		(1.538)
constant	26.153	***	(6.658)
observations	10,738		
number of women	3,159		
R-squared	0.298		
χ^2 (IMR's)	11.43		
Prob > χ^2	0.408		

*** p<0.01, ** p<0.05, * p<0.1 Source: GSOEP, 1999-2012, excluding 2003, own calculations

bootstrapped standard errors in parantheses

Part E

E Conclusions

Germany is facing demographic changes in the form of an ageing society. A low fertility rate together with an increasing life expectancy influence the demographic structure (European Commission (2012, 2014), Federal Statistical Office of Germany (2009)). The old-age dependency ratio of Germany is increasing (Federal Statistical Office of Germany (2009)). This has remarkable consequences for the social security system, on both the revenue and expenditure sides, as well as for economic growth (e.g. OECD (2014), European Commission (2012)). In this context, a broader female labor market involvement is a desirable goal of politics (e.g., OECD (2014), Reinberg and Hummel (2003)). One important requirement for this goal to be achieved is the compatibility of family life and working career.

This dissertation investigates three questions regarding this compatibility. The first relates to intra-household time allocation. To examine male involvement in home-related work, this dissertation investigates how men react if they have time to do engage in this work. The second question regards the career penalty of birth-related career interruptions of women. This goes beyond wages by explicitly taking into account the occupational prestige. The third question relates to the role of informal childcare in respect to maternal employment.

Chapter B asks how couples allocate their time to home-related work if they are facing an employment shock. This question is informed by the unequal division of these tasks in couple households and the question about whether this is due to time constraints or to preferences. We use the survey years 1992 – 2010 of the German Socio-Economic Panel. Since the time allocation within couples is strongly interdependent, we use a difference-in-differences (DID) approach to identify the effect of an employment shock in form of a plant closure. We also exploit the panel structure of the GSOEP and apply fixed effects estimations. The results reveal some reaction in the considered time uses of men and women. The effects are small and mostly insignificant. This indicates that preferences are an important reason for prevalent time allocation. The time men devote to housework or childcare is not increased considerably after facing an employment shock. Even the sum of both home-related chores is not increased remarkably relative to the total time that needs to be reallocated. The female reaction is almost always insignificant. The robustness checks did not alter the results. It must be borne in mind that this analysis displays the reaction in the short term. We only consider the time uses the year immediately after a person has been laid off. To investigate the effects in the long term, we would need more observations over a long time period for the treated group.

Chapter C investigates whether the duration of the career interruption after the first birth is detrimental to a woman's career. Since the prestige level is only observed if a woman is working, a Heckman selection model is applied to correct for the selection into employment. Regardless of the model used, the results reveal a link between prestige and leaves longer than three years. The analysis of occupational mobility compared to the prestige level prior the first birth indicates destabilized careers in case of long time outs instead of short ones. If childless women are included in the analysis, the effects become somewhat weaker in respect to the SIOPS level, but stronger for mobility. Apart from the selection into employment after childbirth, the self-selection into certain occupations that are more or less prestigious might be another source of selection in this context. This selection takes place long before the first child is born. If controlling for the SIOPS level prior to the first birth, the results concerning the level are less distinct while those concerning downward mobility are slightly more distinct. However, the analysis only investigates one dimension of occupational mobility, namely achieving at least one upward resp. facing at least one downward move of at least 10% on the SIOPS scale.

Chapter D investigates the role of informal childcare arrangements in maternal labor market involvement. The analysis takes into account the endogeneity of informal childcare arrangements. When investigating working hours, I correct for the selection into employment. Informal childcare is instrumented using the information about whether the mother of the woman is still alive. Informal childcare seems to facilitate the combination of labor market participation and family life. I find a positive influence of the use of informal childcare on the likelihood of women to participate, even though the effect size depends on the instruments used. Using a Heckman selection correction when investigating the effect of the presence of informal childcare arrangements on the working hours reveals a positive effect of slightly more than an hour. Therefore, the results do not indicate that informal childcare arrangements are crucial for full-time employment instead of part-time employment. The presence of informal childcare arrangement seems to provide some more flexibility within a certain work arrangement.

Overall, this thesis gives some insights into the prevalent situation of mothers concerning the compatibility of a working career and family life. The intra-household allocation of time seems to be pretty stable in Germany, at least in the short term, indicating that preferences are an important determinant of the prevalent intra-household time allocation. This thesis also offers indication that mothers' careers are not per se destabilized by childbirth, but the length of career interruptions is probably an important factor in this respect. The analysis in chapter C reveals that short time outs of up to three years show no distinct destabilizing effect. In addition, interruptions due to other reasons seem to be at least as important for the career path as birth-related interruptions. One important factor in labor market participation seems to be the presence of informal childcare arrangements. The most prevalent argument for this is that informal childcare, e.g., grandparental supervision, provides flexibility to the mother and is associated with more trust than is institutional childcare (e.g., Posadas and Vidal-Fernández (2012), Compton (2011), Compton and Pollack (2014)). Even though this dissertation could not reveal the precise size of this effect, it clearly indicates a distinct, positive effect, since all estimations reveal a remarkable effect size. Therefore, as also mentioned in some part of the literature, policies that provide the aforementioned flexibility might facilitate maternal employment. While the thesis reveals a positive impact on participation, the effect on working hours is small. Informal childcare arrangements may therefore promote maternal labor market participation, but are probably not the crucial determinant in the choice of full-time over part-time employment. It is to be noted, however, that this thesis only investigates the presence of these arrangements per se and not the amount of informal childcare.

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