Single Motherhood and (Un)Equal Educational Opportunities: Evidence for Germany*

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Abstract

We examine the effect of single motherhood on children's secondary school track choice using 12-year-old children drawn from the German Socio-Economic Panel. In line with previous studies for the U.S., the U.K. and Sweden, we find a negative correlation between single motherhood and children's educational attainment. Looking for alternative explanations for this correlation, we use probit regression models to control for factors related to single motherhood such as higher educational background, lower household income and higher labor supply of the mother. Our evidence suggests that single motherhood reduces school attainment mainly because it is associated with lower resources (household income) available for the child.

JEL Classification: I21, J12

Keywords: school choice, educational attainment, binary response model, German Socio-Economic Panel.

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1 Introduction

Equal opportunity education is one of the main underpinnings of a just society. However, a look at reality shows that the inter-generational transmission of educational attainment is high and that the ideal of a meritocracy seems quite elusive even in advanced economies. A child's education is highly correlated with income and schooling of parents and therefore educational opportunities may not be so equal after all.

For Germany the link between parental background and schooling outcomes can be expected to be particularly strong because of the very early tracking of students. In the German school system only the first four years in primary school are shared by a cohort of students. After that, students are sorted into different tracks in a three-tiered secondary school system. Only the highest secondary school track – Gymnasium – allows to enter university directly. The other two tracks, Hauptschule and Realschule, mainly prepare students for entering the labor market through the apprenticeship system. The enrollment decision for secondary school is typically made relatively early between the age of 10 and 12. This distinguishes Germany from the U.K., where the decision for or against a university entrance qualification is made at the age of 16, or the U.S., where a large majority of a birth cohort completes high school. In fact, no other OECD country tracks students earlier than Germany (Woessmann and Hanushek, 2005).

Previous empirical research on the German experience has highlighted the importance of parental income and education on children's secondary school choice and subsequent educational attainment (Dustmann, 2004, Schneider, 2004). A more detailed analysis of the income effects reveals that permanent income, measured between the ages of six and 13, is more important than transitory income at age 13 (Buechel et al, 2001). Splitting up childhood in early and late childhood periods shows that household income in late childhood is more important for secondary school choice than household income in early childhood (Jenkins and Schluter, 2002).

A further suspected contributing factor to inequality in educational attainment, apart from parental income and education, is the family structure during childhood. A detailed analysis of the effects of single parenthood on children's educational outcomes in Germany based on data from a larger nationally representative household survey is, to the best of our knowledge, still missing. We consider such an analysis timely and relevant, as a growing number of children is raised in single-parent households. In 2003, 15 percent of all children under the age of 18 lived with a single parent (13 percent in West Germany and 21 percent in former East Germany), up from 12 percent in 1996 (Statistisches Bundesamt, 2004).

Our research adds to a large amount of international evidence on the topic. Haveman and Wolfe (1995) conclude in their comprehensive review of methods and findings that children who grew up in a single-parent or step-parent family, or experienced a parental separation or divorce, have worse educational outcomes than children who grew up with both biological parents. High school drop out rates in the U.S. are higher for children living in single parent households than for children living in two parent families. Single parenthood during preschool age has a more pronounced adverse effect on high school completion rates than single parenthood at a later age (Gransky, 1995, Fronstin et al, 2001). A similar finding is obtained by Ermisch and Francesconi (2001a, 2001b, 2002) for the U.K. and by Dronkers (1994) for the Netherlands.

These papers address to a varying degree the problem of identifying the causal effect of parental separation from observational data. If single parent families differ systematically from intact families in their other characteristics – and assuming that these matter for educational decisions – then a comparison of schooling outcomes between the two family types mixes causality and selection. Essentially, it could be the case that separation is more prevalent in families with disadvantageous socio-economic characteristics, such as low income, low education levels, unemployment, alcoholism and the like. In this case, the observed educational outcome of children from intact families is not a valid counterfactual outcome for children from separated families had they not separated.

Most recently, Björklund and Sundström (2005) find for Sweden using siblings data that the negative association between separation and educational outcomes might be entirely due to selection. The extent to which their result can be transferred to other countries is an open question. Country specific differences in schooling systems and child support regulations may modify the effect of single parenthood on a child's education. The final verdict on the causality issue is still out.

Our own approach to the selection issue rests on two arguments. First, we can analyse whether separation is associated with unfavorable *observed* characteristics, and control for these in a regression analysis. In the case of income – certainly a main factor – there is the problem that it tends to decrease as a result of separation. In this case, the pre-separation income (observed for those children who lived with a single parent in late childhood or between the age of 13-16) is a better indicator of potential selection effects. Secondly, with regard to selection on *unobservables*, we provide a comparison of children in permanently intact families with those living in a two-parent family during the first twelve years of a child's life but in a single parent family thereafter, when the child is between the age of 13 and 16. Under the selection hypothesis, living in such an instable family should be associated with lower schooling outcomes even if the actual event, the separation and single parenthood, did not occur prior to the school track choice.

In this paper we extend the existing literature in three ways. First, we investigate for Germany the effects of single parenthood on children's educational outcomes, measured in terms of school track at the age of 12. The second novelty is that we test the hypothesis of childhood stage dependence, i.e., whether the effect of living in single parenthood in early childhood is stronger relative to that of living in single parenthood during late childhood. The lasting importance of early childhood development has been emphasized in the recent literature on human capital formation (Heckman, 2000). Thirdly, we identify the channels through which single parenthood affects children's secondary school choice. Is this an effect *per se* – the "emotional upheaval hypothesis" reviewed in Jonsson and Gähler (1997) – or are the reduced resources related to single parenthood – less income and less time – responsible for the children's lower educational attainment?

The data source is the German Socio-Economic Panel (GSOEP), a large German longitudinal household survey started in 1984. For our analysis, we select all 12 year old children from the birth cohorts 1992–2003. The longitudinal data structure permits to obtain current information on income, household size, family structure, and maternal work participation for each year of a child's life. We average this information over two childhood periods, early childhood from birth until the age of six, and late childhood from seven to 12, respectively, and investigate the contribution of these developmental factors to the visited school track at age 12. A further advantage of the data is that we can identify a sample of children in intact families until the age of 12 who will move into single parenthood within the following few years. As mentioned before, this comparison sample is useful for assessing selection effects.

2 Theoretical Framework

In order to model the relationship between educational attainment and single parenthood we assume the existence of an education production function

$$edu = f(p, r) + u \tag{1}$$

where edu is the child's educational attainment measured by the secondary school level at age 12, p is the child's psychological well-being, r is the amount of household resources spent in the upbringing of the child, be it money or time, and u is an i.i.d. error term. Education is an increasing function of household resources and psychological well-being, i.e., $\partial edu/\partial r > 0$ and $\partial edu/\partial p > 0$.

In this framework, the effect of parental separation is mediated through the psychological and material well-being variables. On the one hand, we can write

$$p = g(s, x) \tag{2}$$

where s is an indicator for single parenthood, and x is a vector of socio-economic characteristics, excluding resource variables. We take it as evident that $\partial p/\partial s < 0$, based on a large literature that links a child's psychological well-being to the interaction between parents and their children, and hence to the family structure (e.g., Coleman, 1988, Seltzer, 1994, Boggess, 1998). On the other hand, single parenthood clearly has also an adverse effect on resources

$$r = h(s, x) \tag{3}$$

with $\partial r/\partial s < 0$, since single parenthood can be expected to reduce household income as well as the time available for the child if the single parent needs to start working – or increase working hours – in order to support the family. After substitution, the education production function can therefore be written as

$$edu = f(g(s, x), h(s, x)) + u = \tilde{f}(s, x) + u \tag{4}$$

The reduced-form equation (4) reveals the crucial dependence of a child's educational attainment on single parenthood. Under the assumptions made above, we hypothesize a negative effect of separation on education, both because resources are diminished and because the psychological well-being is compromised. However, based on equation (4), we cannot decompose the overall effect of separation $\partial\tilde{f}(s, x)/\partial s$ into its two constituent parts. Therefore, in order to identify the relative contributions of the resources and psychological effects, respectively, we consider the alternative model, where we control for resources:

$$edu = f(g(s, x), r) + u = \hat{f}(s, x, r) + u \tag{5}$$

The partial derivative $\partial\hat{f}(s, x, r)/\partial s$ gives now the pure psychological effect, and a comparison of $\partial\tilde{f}(s, x)/\partial s$ and $\partial\hat{f}(s, x, r)/\partial s$ indicates the relative importance of the two channels.

The two equations (4) and (5) capture the essence of our theoretical framework in a static world. In addition, we are interested in the dynamic properties of the linkage between separation and educational outcome. In order to be able to address a question such as "Does it matter whether separation occurred during early or late childhood?" we generalize the static equations by formulating the relevant psychological well-being p and resources r as accumulated stock variables.

In this interpretation, p is the stock of psychological capital a child is endowed with at time T. The accumulation process can be expressed as follows

$$p = \int_0^T p(t) w_p(t) \mathrm{d}t = \int_0^T g(s(t), x) w_p(t) \mathrm{d}t$$

and similarly for r

$$r = \int_0^T r(t) w_r(t) \mathrm{d}t = \int_0^T h(s(t), x) w_r(t) \mathrm{d}t$$

The relative importance of early and late childhood events is then captured by the two weighting functions $w_p(t)$ and $w_r(t)$.

3 Empirical Methods

We give a short summary of our estimation strategy shortly. Before, we start with some general remarks on the selection problem and the potentially non-random assignment of separation. According to this hypothesis, the incidence of single parenthood is systematically related to other family specific factors that diminish educational outcomes. We can distinguish between selection on observables that arises if s and x are correlated, and selection on unobservables that arises if s and u are correlated. In this paper we allow for the selection on observables by including as many relevant variables in the regression as possible. Selection on unobservables such as the "quality" of the partnership, i.e., whether it is a stable or an unstable one, will tend to cause an overestimation of the effect of single motherhood on a child's educational attainment if not accounted for.

A first possible approach for addressing selection on unobservables is given by the method of instrumental variables. Gruber (2005), for example, argues that state level variation in divorce laws has the properties required of an instrument. For us, this is not an option as Germany has no regional variation in the divorce laws - and thus in the probability of single parenthood - nor do we have any other convincing instrument.

An alternative approach is to compare the children's educational attainment before and after parental separation. Such a differences-in-differences estimator is implemented by Piketty (2003), who shows for France that children from divorced parents tend to have lower educational attainment already before the separation. De Galdeano and Vuri (2004) provide similar results for the U.S. Since we have for each child only a single measure of educational outcome – school track at age 12 – this identification strategy is not an option either. The problem can be overcome by differencing over siblings rather than time. In this case, one compares the school track of two siblings, provided that a separation took place before the younger sibling reached the relevant age. This method was used by Ginther and Pollak (2000) and Björklund and Sundström (2005). Unfortunately, there are not sufficiently many siblings observed in the GSOEP data to allow for any reasonable analysis. What we can do, however, is to compare the educational attainment of children whose parents separate *after* the schooling decision is made at the age of 12, with those living in intact families. A significant disadvantage of the former group of children would support the selection hypothesis.

We employ in this paper a very simple measure of educational outcome, whether a child attends Gymnasium at age 12. After primary school, which lasts for four years, Germany has – as a rule – a three-tiered schooling system, with Gymnasium being the academically most demanding school track, and the requirement for entering university. The other two secondary-school tracks, Realschule and Hauptschule, are similar in that they prepare students for further specialised vocational schools or for apprenticeship training. There are exceptions to the rule. Some federal states (Schleswig-Holstein, Rheinland-Pfalz, Saarland and Hessen) have a so called Orientierungsstufe in grades five and six. This means essentially that the final school track decision is postponed by a year or two. By choosing the age of 12 for our analysis, we can be confident that the final decision about the secondary school track has effectively been made for all children in the sample.

The variable "Gymnasium" is a measure of both educational achievement and educational opportunity. It is an achievement measure, because the allocation to the three secondary school tracks at the age of 10 to 11 is to a substantial degree determined by the previous academic performance of the child. It is also an opportunity measure, since non-academic factors play a role in the decision as well, and because only Gymnasium leads directly to a university education.

Because the dependent variable is binary, we model attendance of a Gymnasium by means of a standard probit model

$$y^* = \eta + u, \quad u|\eta \sim Normal(0, 1)$$

$$y = \begin{cases} 1 & \text{if } y^* > 0 \quad \text{“Gymnasium”} \\ 0 & \text{if } y^* \leq 0 \quad \text{“Hauptschule or Realschule”} \end{cases}$$

where y^* describes a latent variable dependent on a linear index function of the form $\eta = x'\beta$ and u is an i.i.d. error with a standard normal distribution. As motivated by our discussion above, we want to decompose the overall effect into a resources affect and a psychological effect, controlling for selection as best as possible with the data at hand. Therefore, we consider three alternative models that differ in the assumptions on the index function:

Model 1: $\eta = x_1'\beta_1$

Model 2: $\eta = x_1'\beta_1 + x_2'\beta_2$

Model 3: $\eta = x_1'\beta_1 + x_2'\beta_2 + x_3'\beta_3$

In Model 1 the vector x_1 includes indicators for living in a single parent household, separated by when the event occurred. These are incidence indicators that are one if any episode of single parenthood is recorded in the data, regardless of its length, and zero else. We distinguish between two childhood periods, early childhood from age zero to six, and late childhood from age seven to 12. A third indicator is one if at least one single parent episode was recorded during both childhood periods. These three dummy variables are therefore mutually exclusive, in the sense that a child either lives in a single parent household in early, late or in both childhood periods.

The additional regressors in x_2 in Model 2 control for a potential selection or family effect. These include mother's schooling, an indicator for a foreign household head, and mother's age at birth. To control for the resources effect we include in Model 3 information on child's gender, birth order, families average per capita equivalent household income, mother's labor supply, and average number of children in the household separately for the child's early and late childhood.

The empirical reasoning is as follows: if we compare the educational attainment of children from intact and non-intact families, the difference gives us a combination of the psychological, selection and resources effect (Model 1). In order to decompose the overall effect into its constituent parts, we need to include the vectors containing the controls for the selection and resources effect, x_2 and x_3, respectively, in addition to the vector containing the single parenthood indicator, x_1. The coefficient of the latter measures then the psychological effect, i.e., the specific effect of single parenthood keeping resources constant and controlling for a selection on observables. If the parameter related to the psychological effect becomes insignificant after controlling for selection and resources, whereas the resource effect is significant, we can conclude that single parenthood causally affects children's educational attainments, and that the reasons for this effect are diminished economic resources rather than adverse psychological effects.

4 Data

The data used for this study are drawn from waves 1984-2003 of the German Socio-Economic Panel (GSOEP) (see Burkhauser et al., 2000). The GSOEP contains a great amount of information about household and personal characteristics of their members. Each member older than 16 years answers an own personal questionnaire. For younger children some basic information such as current schooling is provided by the household head in a separate questionnaire. This information is essential for the following analysis.

We select the sample we analyse as follows. First, we identify all children who are 12 years old in one of the years between 1992 and 2003. For these children, we observe the living circumstances during all or most of their childhood years "in real time", i.e., in the current responses by their parents to the annual questionnaires of the GSOEP. This means, for example, that we can use the household and personal questionnaires in 1990 to determine the household size, presence of parents, and mother's labor force participation when a child aged 12 in 2000 – and thus born in

1988 – was two years old. Since income is asked retrospectively for the previous year, we would use information from the 1991 wave to establish the income situation in 1990. Clearly, then, a major selection criterion is that the parents took part in the survey regularly during the first 12 years of a child's life. Children from households, which have gaps in their sample presence or joined the sample only later on, are typically excluded.

For the sample of 12 year old children thus defined, we have separate information on the school track visited at the time of the interview. In a few cases, this is a so-called Gesamtschule (comprehensive school). Since the classification of this school type relative to Gymnasium is ambiguous, we excluded these observations. In order to analyze specific childhood period effects, childhood is divided into two periods: early childhood from zero to six years before children enter school, and late childhood from seven to 12 years after schooling has started.

The information on household income, household size, number of children and mother's labor force participation is obtained as arithmetic mean over the respective childhood period, i.e., early childhood or late childhood. The basis for the income measure is annual household income after taxes and government transfers provided in each wave, deflated to 1995 and on an adjusted per-capita basis, where the following equivalence scale was used: The first adult in a household has a weight of 1, each additional adult 0.7 and each child in the household 0.5. The mother's labor force participation history is measured as average working hours per weekday, again averaged over the two childhood periods.

The information on family structure is included as a "incidence" indicator: it is one in the respective childhood period, if the child lived with a single mother in at least one of the years, and zero otherwise. Single father households, while a theoretical possibility, are empirically irrelevant. Further time invariant variables include the mother's educational background, mother's age at birth, the birth order and the nationality (German or foreign household head). It was not possible to include the educational background of the father or partner. Because of the large number of missing data on this variable, the sample size would have been reduced too much.

Finally, the 12 subsamples for the years 1992 to 2003 were pooled together. Accounting for missing values, the final data set consists of 985 children. Note that we exploit the panel structure of the GSOEP to construct variables from the particular current information rather than from retrospective answers. We consider this a great strength of our analysis as it should allow for better insights into the link between separation and educational attainment.

5 Results

Recall the two basic questions: Do children who grew up with a single mother have a lower probability of attending Gymnasium at age 12 than those who spent the entire childhood with both parents? And if so, why? In Table 1, we offer some first descriptive evidence regarding the association between schooling, single motherhood, and other socio-economics characteristics. The first two columns split the sample by single motherhood, whereas the second two columns contrast the group of students in Gymnasium with the rest.

From the case numbers that are provided in the bottom row, we see that the incidence of single motherhood is quite substantial. One in five, or 195, of the 985 children included in our sample have experienced an episode of single motherhood by the age of 12. This is in line with the aforementioned estimates from the Mikrozensus (Statistische Bundesamt, 2004). Of those 195 cases, 70 involve single motherhood during early childhood only, 54 cases involve single motherhood during late childhood only, and the remaining 71 cases involve single motherhood during both early and late childhood.

At first glance, the bivariate association between single motherhood and school track choice appears not very strong: 33 percent of children with any single motherhood episode attend Gymnasium, compared to 35 percent of children without. The two percentage point gap is both statistically and economically insignificant. The picture changes somewhat, however, if we distinguish between three groups of children, those exposed to single motherhood during early childhood only, those exposed to single motherhood during late childhood only, and those exposed to single motherhood during both childhood episodes. Among the latter group, only 26 percent attend Gymnasium. Hence, there is a 9 percentage point gap in the Gymnasium attendance rate between those who never encountered single motherhood and those who did in both childhood periods.

While we cannot yet tell from this analysis what explains this effect, i.e., whether it is selection or separation, and if separation, whether it is because of resources or because of psychological reasons, the potential size of a causal effect is certainly of a magnitude that warrants some closer analysis. At the same time, we have to point out that there are other factors, income and mother's education in particular, that have a much more pronounced effect on the probability of attending Gymnasium. From Table 1, we see that mothers of children in Gymnasium have 2.3 additional years of schooling. Conversely, only 11 percent of children of mothers with the minimal schooling in the data (7 years) attend Gymnasium, compared to 97 percent of mothers with the highest schooling level (18 years, which corresponds to a university degree). Similarly, a move in the income distribution from the lowest quartile in both childhood periods to the highest quartile in both childhood periods increases the fraction of children attending Gymnasium from 10 to 68 percent.

The regression results are displayed in Tables 2 and 3. The dependent variable is in all cases a dummy variable that is equal to one if the child is enrolled in Gymnasium at age 12 and zero else. A positive regression coefficient means that an increase in the corresponding regressor increases the probability of attending Gymnasium. Marginal effects vary with the individual. For an average person, the probability of attending Gymnasium is 35 percent (from Table 1). The value of the standard normal density at the 0.35 percentile is 0.37. Hence, marginal effects of an "average person" are about 0.37 times the regression coefficient.

Before we discuss the results in detail, we should mention that we considered – but abandoned – a number of alternative specifications. First, we tested whether or not to include as regressors the proportion of time spent in single motherhood during the respective childhood periods, in addition to the mere incidence, a 0/1 indicator variable. A formal test could not reject the hypothesis that these interactive terms are jointly zero. Second, we also split up the single parent variable by reason, distinguishing the three categories "never married", "living separate" and "divorced". Again, a formal test for the homogeneity of these three categories could not reject the null. Hence, we impose it in our model, also in order to conserve on degrees of freedom. Third, we included in Model 2 a measure of mother's general life satisfaction as additional selection control. We eventually dropped this variable for two reasons: it is not clear if parents with lower life satisfaction separate more often, or if separation causes lower life satisfaction; moreover, a likelihood ratio test (p-value 0.43) of a model with life satisfaction against a model without life satisfaction did not reject the null hypothesis that the life satisfaction is irrelevant.

Model 1 gives us the overall stage-dependent effect of single parenthood on the probability of attending Gymnasium at age 12. The regression results show that children who spent both childhood periods with a single mother are significantly less likely to attend Gymnasium than children from intact families. If we translate the coefficients into probability differences (evaluated at the sample means of the other variables) we obtain approximately a 10 percentage points gap. However, the effect of single motherhood during one childhood period only, be it early or late childhood, is insignificant. Thus, we find an effect of single parenthood on educational outcomes, although no significant time pattern can be found.

As we move to Model 2, we see that there is indeed a very strong transmission of educational attainment from mother to child. The coefficient of "mother's education" is positive and significant. An increase of mothers education by one year increases the child's probability of attending Gymnasium by 11 percentage points. Interestingly, these selection variables cannot explain away the single motherhood effect. To the contrary, the effect of having lived in single mother household during both childhood periods on the probability of attending Gymnasium is now larger in absolute value (-13 percentage points), and it is significant at a five percent level. In other words, families with an incidence of single parenthood are, if anything, *positively* selected. This conclusion is confirmed by an analysis of pre-separation income. In particular, we can compare the early childhood income of families who separate in late childhood (a log income of 9.43) with the early childhood income of families who never separate (a log income of 9.20). Hence, separation is actually more likely in high income families than in low income families.

Further evidence on the selection issue is provided in Table 3. Here, we consider a smaller sample of children, those who are observed in the GSOEP data until the age of 16. The main variable of interest is whether a parental separation took place between the age of 13 and 16 (but not before), i.e., after the schooling placement is observed. The hypothesis is that if these instable families had unfavorable unobserved characteristics, then growing up in such a family should be associated with inferior schooling outcomes even before the actual separation takes place. Hence, the dummy, indicating if a child lived in single motherhood between age 13 and 16 should have a negative coefficient. In our data, this is not the case, with or without additional control variables. In summary of the evidence, we conclude that selection is not such a major issue, and thus, that living in a single parent family appears to have a genuine causal effect on the probability of attending Gymnasium.

This leads us to the next question, why such an effect should exist. What are the channels through which single motherhood affects children's educational attainment? In order to address this question, we now return to Table 2 and consider the results for Model 3. From a statistical point of view, this is our preferred specification, since F-tests reject Model 1 against Model 2, and Model 2 against Model 3. The main additional variables of interest are the resource variables, i.e., the childhood period specific average household income and the mother's working hours. The effects are as expected: the probability of attending Gymnasium depends positively on income. A doubling of early and late childhood income is predicted to increase the probability of attending Gymnasium by 30 and 17 percentage points, respectively (again evaluated at the means of the other variables). The effect is significant for both periods but, in contrast to Jenkins and Schluter (2002), larger for the earlier childhood period. However, a child's educational attainment is not affected by the mother's working hours during childhood. While the point estimates are negative, both parameters are statistically insignificant.

The estimated coefficients for the number of children in the household during early and late childhood and for birth order are all individually insignificant as well. Thus, we find no evidence for sibling rivalry, although that obviously does not prove that there is none. Finally and importantly, all three coefficients of the family structure variables – lone motherhood in early childhood, late childhood, or both periods – are insignificant in this extended model. We find – as conjectured – that the observed correlation between single motherhood and secondary school track is mostly attributable to the resources effect. According to the evidence in our data, the psychological effect plays a subordinate role only.

6 Conclusions

This paper examined the relationship between spending the childhood with a single mother and secondary school track assignment of 12-years olds in Germany, using data from the German Socio-Economic Panel for the period 1984 - 2003. As in other industrialised countries, single motherhood is on the rise in Germany. One in five children in our data has experienced an episode of single motherhood by the age of 12. Moreover, the rate at which such children attend the highest secondary school track, Gymnasium, trails the rate of children from two-parent families by nine percentage points (for children who experience single motherhod during both early and late childhood).

A significant gap remains once we control in a probit regression model for family background variables related to the mother's educational attainment, age and nationality. Moreover, we cannot find a gap comparing children from intact families with those from intact families – up to the age of 12 – who will separate shortly after. Both results can be interpreted as evidence against a pure selection explanation according to which separation and single motherhood is associated with third unfavorable observed or unobserved socio-economic factors that by themselves diminish a child's chances of entering the most demanding school track.

In a further analysis, we establish that the observed adverse effect of single motherhood on secondary school track attainment is mostly attributable to diminished family resources of single mother families. When controlling for household income and mother's labor force participation, the estimated coefficients for the single-motherhood indicator variables become insignificant. The lower educational attainment of children growing up in single mother households appear therefore not to be due to psychological factors.

Our empirical strategy has exploited the long duration of the Germany Socio-Economic Panel. With 20 years of data available to us, we were in the position to extract information on household characteristics for each year of a child's life from the annual survey information provided by the parents in the past. This approach has the great advantage of providing accurate information on past income, family structure and labor force participation. The disadvantage is that even after pooling data over 12 years, we only obtain a final sample of 985 youngsters. As a consequence, all the effects are estimated only with low precision.

An alternative approach is to use retrospective information instead. In the GSOEP, such information is available for each adult person and for each year of one's adult life. It provides information on marital status and labor force status. A major drawback is that working hours and in particular income is not available. Still, it might be worthwhile to pursue such an analysis in future research. A much larger sample size might make it even feasible to implement a siblings estimator. A further extension of our research – also with such an alternative sample – could go beyond the short-term outcome "school attainment at age 12" and investigate the long-run consequences of single parenthood on labor market outcomes of children.

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Table 1. Sample Means by Single Motherhood and Gymnasium

	Single Motherhood		Gymnasium	
	no	yes	no	yes
Gymnasium	0.354	0.333	0	1
	(0.017)	(0.034)		
Single motherhood in early childhood	0	0.359	0.072	0.070
		(0.034)	(0.010)	(0.0145)
Single motherhood in late childhood	0	0.277	0.050	0.064
		(0.032)	(0.009)	(0.013)
Single motherhood in both	0	0.364	0.081	0.055
early and late childhood		(0.035)	(0.011)	(0.012)
Mother's years of schooling	10.88	11.33	10.15	12.50
	(0.085)	(0.173)	(0.068)	(0.147)
Mother's age at birth	27.59	25.91	26.55	28.56
	(0.176)	(0.377)	(0.206)	(0.241)
Foreign household head	0.313	0.169	0.359	0.145
	(0.017)	(0.027)	(0.019)	(0.019)
Log per capita income 0-6	9.203	9.184	9.095	9.392
	(0.013)	(0.027)	(0.013)	(0.018)
Log per capita income 7-12	9.284	9.208	9.170	9.453
	(0.013)	(0.027)	(0.013)	(0.021)
Weekly hours of work 0-6	2.017	3.021	2.240	2.171
	(0.089)	(0.188)	(0.103)	(0.133)
Weekly hours of work 7-12	2.577	3.975	2.816	2.923
	(0.090)	(0.212)	(0.107)	(0.141)
Number of observations	790	195	640	345

Data: German Socio-Economic Panel, own calculations.
Standard errors in parentheses.

Table 2. Probit Regression Results; Full Sample (N=985)

	(1)	(2)	(3)
Single mother 0-6	0.018	0.060	0.160
	(0.168)	(0.188)	(0.198)
Single mother 7-12	0.163	0.054	0.181
	(0.184)	(0.205)	(0.218)
Single mother 0-12	-0.300+	-0.406*	-0.057
	(0.175)	(0.202)	(0.217)
Mother's education		0.309**	0.251**
		(0.027)	(0.029)
Mother's age at birth		0.029**	0.039**
		(0.010)	(0.013)
Foreign household head		-0.094	0.154
		(0.125)	(0.136)
Log per capita income 0-6			0.806**
			(0.242)
Log per capita income 7-12			0.558*
			(0.219)
Work 0-6			-0.030
			(0.027)
Work 7-12			-0.021
			(0.026)
Log # children in HH 0-6			-0.009
			(0.241)
Log # children in HH 7-12			0.223
			(0.196)
Female			0.028
			(0.098)
Birth order			-0.146
			(0.098)
Log-likelihood	-607.1	-482.6	-449.4

Notes:

Dependent variable: Gymnasium at age 12 (yes/no).

Standard errors in parentheses.

Significance levels: + 10 percent, $*$ 5 percent, $**$ 1 percent.

The models include in addition a constant, 12 time dummies and seven federal state dummies.

Table 3. Probit Regression Results; Selection Sample (N=692)

	(1)	(2)
Single mother 13-16	0.063	-0.233
	(0.306)	(0.340)
Mother's education		0.277**
		(0.032)
Mother's age at birth		0.021+
		(0.012)
Foreign household head		-0.244+
		(0.147)
Log-likelihood	-427.7	-353.8

Notes: see Table 2.