Technical Appendix:

Childhood Family Structure and Schooling Outcomes: Evidence for Germany

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Identification Issues and Estimation Methods

A central concern of this study is that an estimated effect of childhood family structure on education may be spurious due to the mutual association between family structure and children's schooling achievements, and some unmeasured true causal factor. For example, the association between having experience of life in a non-intact family and lower educational attainment may not be the result of family structure during childhood; differences in attainment may simply reflect the characteristics of families in which children of lone mothers are brought up.

Our econometric strategy is to apply a number of different techniques which place different assumptions on the data in order to identify the effects of experiencing life in a non-intact family during childhood. Each of these techniques has advantages and disadvantages. Although the methods differ, they share a number of common elements. For this reason, in what follows we strip down the corresponding statistical models in order to better highlight the difference in their identifying restrictions.

Sibling Difference Model

Let *j* index families and *i* index young adults (or children). For convenience, assume that the relationship we estimate is

(A.1) $$S_{ij} = \beta F_{ij} + u_{ij},$$

where S_{ij} is education, F_{ij} is a variable that indicates childhood family structure (e.g., ever lived in a non-intact family in the first 10 years of life), and u_{ij} is random shock with zero mean. In equation (A.1), which for the moment excludes other determinants of schooling,[1] β is the parameter of interest.[2] Consistent estimation of β requires that F_{ij} be uncorrelated with the disturbance term u_{ij}. We investigate this issue using a framework suggested by Behrman et al. (1994) and Rosenzweig and Wolpin (1995). Consider a two-child family. For the *i*-th child in family *j* with sibling *k*, u_{ij} can be decomposed as follows:

(A.2) $$u_{ij} = \delta_1 \varepsilon_{ij} + \delta_2 \varepsilon_{kj} + \varepsilon_j + \eta_{ij},$$

[1] In our empirical analysis, the relationship (A.1) is expanded to include a set of child- and family-specific variables that may be fixed or time-varying.

[2] In this formulation, β is assumed to be the same for all individuals. Arguably, the effect of family structure is heterogeneous (i.e., some children might be better off in a non-intact family, while others might be worse off). The sibling difference approach would apply even if one specifies a random-coefficients model in which $\beta_j = \beta + \zeta_j$, and $E(\zeta_j u_{ij}) = E(\zeta_j X_{ij}) = 0$. However, it may not be feasible to estimate such a model, because repeated observations with each family *j* are needed, whereas most of the families in our sample consist of only two or three siblings.

where ε_{ij} and ε_{kj} are the endowments or 'ability' of each sibling, ε_j denotes the genetic endowments that are common to both children of family *j*, and η_{ij} is a random shock that is specific to i in *j*, inclusive of measurement error in schooling. Endowments of both siblings are likely to be transmitted across generations in a Galton-type law of heritability (Becker and Tomes 1986):

(A.3) $$\varepsilon_{ij} = \rho\varepsilon_j + \psi_{ij},$$

where ψ_{ij} is a child-specific idiosyncratic disturbance with zero mean and uncorrelated with other unobservables (including ψ_{ik}, the corresponding random term for sibling *k*). Assuming that $0 \le \rho < 1$ implies that endowments regress towards the mean across generations. Finally, the family structure variable F_{ij} is itself a function of unobserved variables that pertain to the family (ϕ_j) and to the two siblings $(\mu_{ij}$ and $\mu_{kj})$:

(A.4) $$F_{ij} = \gamma_1\mu_{ij} + \gamma_2\mu_{kj} + \pi\phi_j + \theta_{ij},$$

where θ_{ij} is a disturbance that affects F_{ij} but does not affect S_{ij} except indirectly through F_{ij}. It is well known that the parameter β is not identified with equations (A.1)–(A.4) if π is not zero, ρ is not zero, and if either δ_1 and γ_1 or δ_2 and γ_2 are not zero,[3] even if orthogonality restrictions on the moments involving η_{ij}, ψ_{ij} and θ_{ij} are imposed. That is, β is estimated with bias if equation (A.1) is estimated across individuals with different values of family and children's endowments.[4]

The sibling difference model estimates β by comparing educational outcomes among siblings according to whether they experienced life in a non-intact family during childhood. In our two-child family case, the sibling difference estimator is computed from

(A.5) $$\Delta S = \beta\Delta F + \Delta u,$$

where $\Delta r = r_{ij} - r_{kj}$, for any term r in equation (A.5). The within-family covariance between family structure differences and the disturbance term in (A.5) is thus given by

(A.6) $$\text{cov}(\Delta F, \Delta u) = (\gamma_1 - \gamma_2)(\delta_1 - \delta_2)E(\Delta\mu\Delta\theta) + (\gamma_1 - \gamma_2)E(\Delta\mu\Delta\eta).$$

Therefore a sufficient condition for β to be identified is that $\gamma_1 = \gamma_2$; that is, parents respond to their children's idiosyncratic endowments equally. A stronger condition would be to

[3] These would be the selection-on-observables assumptions, which are relevant for all cross-sectional estimators, including those based on propensity score matching methods.

[4] This conclusion applies to propensity score matching estimates, since these too rely on the selection on observables assumption of the cross-sectional model (A.1)–(A.4).

assume that children's endowments do not affect parents' behaviour (family structure), or, alternatively, there are no intrafamily responses (i.e., $\gamma_1 = \gamma_2 = 0$).[5]

Before-After Comparisons and Quasi-Experiments

To see the assumptions needed for identification in this case, we modify our empirical framework slightly and explicitly allow parents' behaviour and child unobservables to differ over time, so that:

(A.7) $$F_{ij} \equiv F_{ijb} + F_{ija} \quad \text{and} \quad u_{ij} \equiv u_{ijb} + u_{ija},$$

with the subscripts b and a indicating some time period 'before' and 'after' a specific event occurs (e.g., death of a parent and the introduction of a divorce law reform). Equations (A.2) and (A.4) respectively become:

(A.2') $$u_{ij\tau} = \delta_\tau \varepsilon_{ij\tau} + \varepsilon_{ij} + \varepsilon_j + \eta_{ij\tau}$$

and

(A.4') $$F_{ij\tau} = \gamma_\tau \mu_{ij\tau} + \mu_{ij} + \pi\phi_j + \theta_{ij\tau},$$

for $\tau = b, a.$ A straightforward before-after comparison of outcomes for the same individual requires repeated information on $S,$ which is problematic when schooling is measured in term of highest educational attainment. Bearing this in mind, and imposing orthogonality restrictions on all moments involving $\eta_{ij\tau}$, ψ_{ij} and $\theta_{ij\tau}$, a fixed-effects estimator based on (A.1), (A.2'), (A.3) and (A.4') will imply

(A.8) $$\text{cov}(\Delta F, \Delta u) = \gamma_a \delta_a \sigma_{\varepsilon_a \mu_a} - \gamma_a \delta_b \sigma_{\varepsilon_b \mu_a} + \gamma_b \delta_b \sigma_{\varepsilon_b \mu_b} - \gamma_b \delta_a \sigma_{\varepsilon_a \mu_b},$$

where $\sigma_{pq} = \text{cov}(p,q),$ for $p,q = \varepsilon_a, \varepsilon_b, \mu_a,$ and $\mu_b.$ Notice that β cannot be identified even if we assume that the correlations between ε s and μ s are the same before and after the change of interest. Identification instead can be guaranteed if there are no intrafamily responses (i.e., $\gamma_a = \gamma_b = 0$), *or* if child endowments are not "regime" specific (i.e., $\delta_a = \delta_b = 0$).

The assumption that $\delta_a = \delta_b = 0$ is perhaps more credible when some 'exogenous' events are taken as instruments (e.g., the passage of divorce law regulations) rather than others, such as remarriage or parental death, since specific realisations of ε_{ijb} (and

[5] If the family allocates schooling so as to reinforce endowment differences between siblings, then $\gamma_1 > 0$ and $\gamma_2 < 0;$ if instead the family compensates for child-specific endowment differentials, then $\gamma_1 < 0$ and $\gamma_2 > 0$ (see Berhman et al. 1994). The assumption of no intrafamily responses was imposed by Rosenzweig and Wolpin (1991) and Currie and Cole (1993) to estimate the determinants of birth outcomes.

expectations about ε_{ija}) may ultimately lead to such events. Even when those more credible exogenous circumstances apply, one concern is that the regulations may be endogenous, in the sense that there may be trends in educational attainments of children of divorced parents that are correlated with the introduction of a specific divorce law (Gruber 2004).

To illustrate this point, let D_i denote a dummy variable that is equal to 1 if individual *i*'s parents divorced during his/her childhood and 0 otherwise, and let *q* be the time period in which the divorce law reform occurred. Suppose our outcome of interest S_{it} is determined by the following specification[6]

(A.9) $$S_{it} = \alpha_0 + \alpha_1 D_i + (\alpha_2 + \alpha_3 D_i)t + \alpha_4 \mathrm{I}(t \geq q) + \beta D_i \mathrm{I}(t \geq q) + \xi_{it},$$

with $E(\xi_{it} \mid D, t) = 0,$ where $E(\cdot)$ is the expectation operator, and the term $\mathrm{I}(w)$ is a function indicating that the event *w* occurs (i.e., the post-reform period), so that in equation (A.9) β is our parameter of interest. The parameters α_2 and α_3 reflect two different time trends in educational achievement for children in intact families and children in divorced families, respectively, while α_4 reflects a common jump in S_{it} from the time the reform is introduced onwards. From (A.9), a difference-in-difference estimate of β is given by $\beta + \alpha_3(k + k'),$ where $k + k'$ represents the average number of calendar periods (say, years) between the post-reform and pre-reform period observations in the sample.[7] Unless $\alpha_3 = 0,$ this is clearly a biased estimate of β. The bias arises precisely because the time evolution of S_{it} differs between children in intact families and children of divorced parents. With this approach, therefore, modelling group-specific trends will be crucial.

[6] The family subscript *j* has been dropped for convenience.

[7] With the family structure variable and the residuals in (A.9) following the same structures as those specified in (A.2') and (A.4'), a necessary condition for identification of β is, as before, that child endowments are independent of the reform (or that there are no intrafamily responses).