The effect of having a child with a disability on repartnering: Evidence from Swedish register data

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Abstract

We examine how childhood disability, using cerebral palsy (CP) as a case study, affects repartnering. Although several studies have examined the effects of childhood disability on family dissolution, far less is known about its impact on repartnering. In particular, existing evidence on repartnering in families of children with disabilities is restricted to mothers. To our knowledge, this is the first study to examine repartnering among both mothers and fathers of children with disabilities using nationwide data. Therefore, we develop a search-and-match model of the marriage market and apply it to comprehensive Swedish register data covering children diagnosed with CP born 1991–2014 and matched controls (5:1) based on sex, birth year, and municipality of residence at diagnosis. Using Cox survival models to study time to repartnering among parents who become single, we find that single mothers of children diagnosed with CP face a 23% lower repartnering hazard than mothers in the matched control group, while we find no statistically detectable difference for single fathers. These gendered results indicate that having a child with a disability imposes a substantially greater burden on single mothers' repartnering prospects than single fathers', consistent with gender differences in caregiving constraints emphasized by the theoretical model. These findings may inform the design of support policies for single parents of children with disabilities, particularly mothers who are often primary caregivers.

Keywords: Disability, Cerebral palsy, Repartnering, Marriage market, Gender equality

1 Introduction

Disability is a complex phenomenon that can impact families in various ways. Children with disabilities require additional care and support, which can alter intra-household dynamics and decision-making processes related to family formation and dissolution. Despite the importance of understanding the impact of disability on family composition, there are substantial gaps in the literature.

The impact of having a child with a disability on family dissolution has been a topic of interest in previous literature (see e.g. Mauldon, 1992; Risdal and Singer, 2004; Lundeby and Tøssebro, 2008; Hartley et al., 2010; Hatton et al., 2010; Bengtsson et al., 2011; Loft, 2011; Namkung et al., 2015; Tøssebro and Wendelborg, 2017; Müller et al., 2022; Doyle et al., 2023). However, research on the impact of childhood disability on *repartnering* following family dissolution is limited. Research on repartnering in the general population has found that having a child affects the likelihood of repartnering, especially for mothers (Ivanova et al., 2013; Schnor et al., 2017; Di Nallo et al., 2023). This effect could be different in single parents of children with disabilities, given that they face different circumstances than single parents of typically developing children, including worse mental and physical health, higher poverty, lower employment, and weaker financial support from non-residential parents than relevant comparison groups (Balbo and Burciu, 2026).

The limited existing evidence on repartnering in families of children with disabilities is overwhelmingly mother-centered, leaving fathers' repartnering largely unstudied (Mauldon, 1992; Bengtsson et al., 2011; Tøssebro and Wendelborg, 2017).[1] To our knowledge, only three studies have examined repartnering in this context, all focusing exclusively on mothers. An early contribution, Mauldon (1992), uses the 1981 Child Health Supplement to the U.S. National Health Interview Survey to estimate a Weibull proportional hazards model of mothers' remarriage following divorce and finds no statistically significant association between a child's disability or health limitations and the remarriage hazard. However, this evidence has limited relevance for contemporary repartnering: the remarriage model is estimated on a subsample of 844 children whose parents divorced after 1971 and who were aged nine or younger at divorce, and only about 4% of children in the underlying data are classified as "disabled" (and about 2% as "limited"), implying relatively few disabled observations in the remarriage sample and correspondingly limited

[1] Related evidence from the UK describes partnering and repartnering using changes in household composition for children with early cognitive delay. After propensity-score matching on family socioeconomic circumstances, differences in separation, partnering, and repartnering are not statistically significant (Hatton et al., 2010).

statistical power. More recent Scandinavian evidence is consistent with reduced repartnering among mothers, but it typically measures repartnering at specific points in the child's life course rather than modeling time to repartnering. Using a representative Danish survey of families with 11-year-old children, Bengtsson et al. (2011) show that, conditional on not living with the biological father, mothers of children with disabilities are more likely to still be single by the time the child is age 11. Using Norwegian longitudinal survey data linked to population registers, Tøssebro and Wendelborg (2017) finds that repartnering among separated mothers of children with disabilities does not differ statistically from other mothers when children are around ages 11–13, but is significantly lower when children are around ages 14–16. Taken together, the small existing literature suggests that mothers of children with disabilities experience lower repartnering rates, while evidence on fathers is lacking.

Theory and prior research offer mixed expectations regarding how having a child with a disability might affect a single parent's chances of repartnering. On one hand, children in general impose time and resource demands that can hinder single parents' opportunities to meet new partners, for example, by reducing social interactions (de Graaf and Kalmijn, 2003). This effect tends to be more pronounced for mothers, who are often the primary caregiver and thus face greater childcare responsibilities, as well as greater reluctance from prospective partners (Bumpass et al., 1990; Hatton et al., 2010; Ivanova et al., 2013; Schnor et al., 2017; Tøssebro and Wendelborg, 2017). A child's disability could exacerbate these challenges due to more intensive caregiving needs and potential social stigma (Hatton et al., 2010; Tøssebro and Wendelborg, 2017). On the other hand, economic theory suggests that single parents may seek a new partner to pool resources and share caregiving responsibilities to counter the financial losses experienced by single parents after divorce or separation (Jansen et al., 2009).

We add to this literature by investigating the impact of having a child with cerebral palsy (CP) on repartnering among single parents, examining both mothers and fathers. For methodological reasons, CP has previously been used to examine, for example, the effect of having a child with a disability on the likelihood of divorce, the probability of having an additional child, and labor-market outcomes (Müller et al., 2022; Asuman et al., 2023). Our analysis follows this literature and is based on a comprehensive register-based dataset that includes all children diagnosed with CP in Sweden born between 1991 and 2014, along with a matched control group from the general population. In these data, repartnering is observed through either marriage or the birth of a child with a new partner, thus capturing formalized unions and cohabiting relationships that progress to a shared birth.

From an empirical perspective, using CP as a proxy for disability is helpful for several reasons. Firstly, CP is the most common motor disability in children. It affects approximately 2-2.5 per 1,000 live births in developed countries (Oskoui et al., 2013) and has been described as representative of other congenital and early-onset disabilities (Michelsen et al., 2015). Thus, findings for CP may be informative for other congenital or early-onset disabilities that involve persistent caregiving demands. Secondly, CP is a recognized disability that is unlikely to remain undiagnosed. Thirdly, anticipatory behavior of parents (e.g., termination of pregnancy) is unlikely to introduce bias because CP itself is typically diagnosed only after birth. Fourth, although the literature reports mixed results on the association between CP and prenatal characteristics (Oskoui et al., 2013; Solaski et al., 2014; Michael-Asalu et al., 2019; Müller et al., 2022), a general result is that this association is weak (Rosenbaum et al., 2007; Müller et al., 2022). We illustrate this for our data and show that, in line with previous results for Sweden, the occurrence of CP is largely uncorrelated with prenatal characteristics. Thus, CP offers features that help mitigate identification challenges noted in previous studies. Fifth, there is substantial variation in function among persons with CP; some function independently while others require assistance to undertake daily activities (Rosenbaum et al., 2007). Hence, CP is a particularly relevant condition when the effects of having a child with a disability on parental outcomes are studied, as the variation in functional ability can be utilized to study the impact of functional limitation in itself. Finally, CP is a permanent disability with far-reaching implications, often requiring a lifelong parental commitment. Such long-term conditions are understudied compared to acute health shocks, which may be empirically exogenous but are not representative of permanent caregiving demands.

To our knowledge, this is the first study to examine repartnering among both mothers and fathers of children with disabilities using nationwide register data. We provide new evidence of a striking gender disparity in repartnering behavior when a child has a disability, in a context (Sweden) known for high gender equality and strong social support. Specifically, our findings suggest that having a child with CP has a gendered impact on repartnering, with single mothers of children with CP having, on average, a 23% lower repartnering hazard than their counterparts, while repartnering outcomes appear to be unaffected for single fathers. The results provide insight into how a child's disability can alter parents' decision-making processes and opportunities in the marriage market. These findings have important implications for policymakers designing support programs for single parents of children with disabilities.

2 Theory

We develop a search-and-matching model of the marriage market to formalize how caring for a child with CP might lower a single parent's repartnering hazard, and why mothers could be more affected than fathers. The model can easily be generalized to other long-term childhood disabilities that entail persistent caregiving demands and may affect meeting rates and partner acceptance, although we present the model specifically for CP in line with our empirical approach. The model builds on standard marriage-market theory (e.g., Becker et al., 1977; Mortensen, 1988; Burdett and Coles, 1999), but incorporates additional costs and constraints associated with raising a child with CP. In particular, we allow for (i) *caregiving costs* that reduce a single parent's utility while single and may be partially alleviated by repartnering, (ii) a *screening/availability* factor that lowers the effective rate of viable contacts, either because some potential partners are unwilling to consider parents of children with CP or because caregiving constraints reduce those parents' availability to meet new partners, and (iii) *parent-role–specific parameters* reflecting that mothers often bear greater childcare responsibilities and may face different societal expectations than fathers (Ivanova et al., 2013; Schnor et al., 2017; Hatton et al., 2010; Tøssebro and Wendelborg, 2017). The model yields predictions for how the presence and severity of a child's CP may affect repartnering hazards, and how these effects could differ for single mothers versus single fathers. To align with our empirical design, we restrict attention to single parents whose children do or do not have CP. Extending the framework to include childless individuals is conceptually straightforward but does not shed additional light on our empirical contrasts and is therefore omitted here.

2.1 Setup and Assumptions

Consider a continuum of single mothers (m) and fathers (f) on the heterosexual marriage market.[2] Parents are characterized by whether they have a child with CP as opposed to having a child without CP. Let $d_i \in \{0, 1\}$ indicate if parent i has a child with CP. We denote the utility cost associated with caring for a child with disability by $\alpha_i \geq 0$. If $d_i = 0$ (CP not present), we set $\alpha_i = 0$; if $d_i = 1$, then $\alpha_i > 0$, with larger α_i representing more severe disability (higher caregiving burden). We also allow α to differ by the parent's gender: in general, we expect $\alpha_m > \alpha_f$, as single mothers tend to incur greater daily childcare burdens than single fathers (who often

[2]For simplicity, we focus on heterosexual pairings and parents only. The model can be extended to other partnership contexts as needed.

have lower custody or more external support). This reflects the empirical reality that mothers usually bear primary responsibility for children after union dissolution (Ivanova et al., 2013; Schnor et al., 2017; Hatton et al., 2010; Tøssebro and Wendelborg, 2017).

Each single parent i derives utility U_i^S if they remain single and U_i^P if they enter a partnership with some partner j. We define:

$$U_i^S = u(y_i) - \alpha_i \,, \tag{1}$$

where $u(y_i)$ is the base utility from individual i's income or resources when single. The term α_i captures the disutility from raising a child with CP alone, as compared to a typically developing child. This includes direct monetary costs, time expenditures, stress, and any reduced leisure or earning capacity.[3] A higher severity of the child's condition increases α_i, lowering the single parent's utility. This formalizes the intuitive idea that single-handedly caring for a child with CP is taxing and can reduce a parent's well-being compared to caring for a child without CP.

When a single parent and a potential partner meet, they decide whether to form a match (repartner). We assume a *random matching process*: in each instant (continuous time), single parents encounter potential partners at a Poisson arrival rate $\lambda > 0$. Meetings are random draws from the pool of available singles. We introduce a screening/availability filter that represents social barriers to meetings and reduces effective meeting opportunities for parents of children with CP: let $0 \leq \sigma_m, \sigma_f < 1$ denote the fraction of potential opposite-sex partners not in the meeting pool of a single parent of a child with CP due to obstacles related to the disability (e.g., stigma or perceived challenges, or the parent's limited availability due to isolation or a reduced social life). Thus, a mother of a child with CP meets potential partners with the effective rate $(1 - \sigma_m)\lambda$ (less than the baseline λ), and similarly a father at rate $(1 - \sigma_f)\lambda$. This captures the idea that some potential partners may screen out parents of children with CP, while caregiving demands may also reduce those parents' own availability to meet new partners. In our notation, both channels are summarized by σ_g ($g \in \{m, f\}$) as a reduction in

[3]For a parent with a typically developing child, $\alpha_i = 0$, so that $U_i^S = u(y_i)$. We do not model utility components of parenthood that are common to both groups separately, since all individuals in our empirical setting are parents. Instead, α_i captures the additional CP-related burden relative to having a typically developing child, which may include not only direct monetary and time costs but also any extra emotional strain associated with the child's condition. Likewise, we do not model resource pooling or household public goods as separate terms; instead, common gains from partnership are absorbed into the match-surplus term ϵ_{ij}. This normalization avoids double counting and keeps the model focused on the CP-related differences of interest.

the effective rate of viable contacts (Hatton et al., 2010; Tøssebro and Wendelborg, 2017). We expect $\sigma_m \geq \sigma_f$ if single mothers face greater ex-ante screening and/or have less availability to meet (e.g., time constraints), i.e., fewer effective meeting opportunities.

Each meeting between parent i and a potential match j yields a pair-specific draw of match quality, ε_{ij}, which reflects idiosyncratic mutual compatibility, affection, and any gains from cooperation (e.g., companionship or household public goods). We treat ε_{ij} as a random variable with cumulative distribution $F(\varepsilon)$ (with full support on $\mathbb{R}$ and with 0 representing a neutral quality). A higher ε_{ij} means a better match. Importantly, if either partner has a child with CP ($d_i = 1$ or $d_j = 1$), the new partner will share the caregiving burden to some extent. Specifically, if parent i has a child with CP ($\alpha_i > 0$) and marries partner j, we assume that the partner takes on a fraction ϕ of the childcare burden, with $0 \leq \phi \leq 1$. The parameter ϕ represents the share of the *focal single parent's* childcare burden that a new partner would assume if a match forms. For clarity, we index ϕ by the focal parent: ϕ_m is the share of the childcare burden a new partner would assume when the focal parent is a mother, and ϕ_f is the analogous share when the focal parent is a father. Given prevailing caregiving norms, it is plausible that $\phi_f \geq \phi_m$, but we treat ϕ as exogenous in the baseline.

2.2 Match Utilities and Acceptance Decisions

When parent i and potential partner j meet, they observe the pair-specific match quality ε_{ij}. If they decide to partner, each will compare their expected utility in that partnership to their utility if they remain single. We assume utility in marriage is additively separable into the individual's own baseline $u(y)$ (now possibly from combined resources) plus the match quality, minus any costs related to children. To keep the analysis tractable, we suppose that marriage does not change an individual's intrinsic income utility $u(y_i)$ (partners do not fully pool income or share equally with no net change in individual utility baseline), and focus on the net new utility from the partnership. Empirically, we interpret this simplification as assuming that resource pooling upon repartnering does not systematically differ between parents with and without a child with CP (conditional on covariates). Under this condition, the estimated impact of CP on repartnering primarily captures caregiving/stigma channels. Thus, the utility for parent i if repartnered with j with match quality ε_{ij} is:

$$U_i^P = u(y_i) + \varepsilon_{ij} - C_{ij}\,, \tag{2}$$

where C_{ij} represents the cost or disutility for parent i associated with raising a child with CP together with partner j, compared to raising a typically developing child with partner j. Similarly, the utility of partner j is:

$$U_j^P = u(y_j) + \varepsilon_{ji} - C_{ji}\,, \tag{3}$$

where $\varepsilon_{ji} = \varepsilon_{ij}$. However, the cost term, C is not symmetric as it reflects the caregiving burden shared (equally or unequally) in the new household. For instance, suppose parent i is a mother of a child with CP ($\alpha_i > 0$). In that case, we model the mother's cost as $C_{ij} = (1 - \phi_m)\alpha_i$ (she still carries a portion of the burden) and the new partner's cost as $C_{ji} = \phi_m\alpha_i$. Here $\phi_m\alpha_i$ is the part of the extra childcare cost due to CP assumed by the new partner, and $(1-\phi_m)\alpha_i$ is the remaining cost borne by the mother. If instead parent i is the father of a child with CP who repartnered with partner j, the cost borne by the father is $C_{ij} = (1 - \phi_f)\alpha_i$ and the cost borne by the new partner is $C_{ji} = \phi_f\alpha_i$.

Given these utilities, the parent and the new partner will form a match (repartner) if and only if *both* individuals weakly prefer the match to staying single. That is, the following individual rationality conditions must hold simultaneously:

$$U_i^P \geq U_i^S \quad \text{and} \quad U_j^P \geq U_j^S\,. \tag{4}$$

Substituting from (1) and (2), we can express the net gains from repartnering for each side. When parent i is a mother, her gain from repartnering is:

$$\Delta_i \;\equiv\; U_i^P - U_i^S \;=\; \Big[u(y_i) + \varepsilon_{ij} - (1-\phi_m)\alpha_i\Big] - \Big[u(y_i) - \alpha_i\Big] \;=\; \varepsilon_{ij} + \phi_m\alpha_i\,. \tag{5}$$

She gains the match surplus ε_{ij} *plus* relief equal to $\phi_m\alpha_i$ because her childcare burden is partially shared in the new partnership. The prospective partner's gain (with j denoting the partner) is:

$$\Delta_j \;\equiv\; U_j^P - U_j^S \;=\; \Big[u(y_j) + \varepsilon_{ji} - \phi_m\alpha_i\Big] - u(y_j) \;=\; \varepsilon_{ji} - \phi_m\alpha_i\,, \tag{6}$$

For a match to be accepted by both, we require $\Delta_i \geq 0$ and $\Delta_j \geq 0$. The focal mother's acceptance requires $\varepsilon_{ij} + \phi_m\alpha_i \geq 0$, which will hold for any moderately positive or even slightly negative ε_{ij} as long as $\phi_m\alpha_i$ is sufficiently large. In other words, because being single is so costly for her (she carries α_i alone), she may accept a partner even if the immediate match quality is not ideal, provided the partnership offers relief in caregiving. The partner's acceptance requires $\varepsilon_{ji} - \phi_m\alpha_i \geq 0$, i.e.,

$$\varepsilon_{ji} \geq \phi_m\alpha_i\,. \tag{7}$$

The case when parent i is a father meeting a prospective partner is symmetric, with ϕ_f instead of ϕ_m.

Equation (7) is typically the binding constraint: the new partner will only enter the union if the pair-specific match quality ε_{ji} is high enough to offset the additional expected cost $\phi_m \alpha_i$ (Becker et al., 1977; Hatton et al., 2010; Tøssebro and Wendelborg, 2017). If the child has a severe condition (large α_i), the bar $\phi_m \alpha_i$ is higher, meaning only an exceptionally good match (large ε_{ji}) will induce repartnering. By contrast, if $\alpha_i = 0$ (no CP present), the condition reduces to $\varepsilon_{ji} \geq 0$ and any meeting that yields non-negative compatibility is acceptable to both sides.

If both partners *have* children with CP (say costs α_i and α_j), each would get relief from their own burden but take on part of the other's child's care. For brevity, we do not derive that case in full, as it does not shed particular light on our empirical analyses; however, it entails a similar logic, which is that higher combined child-related costs necessitate a sufficiently high ε_{ij} and ε_{ji} to make matching mutually beneficial.

Because the pair-specific match quality is randomly drawn (and assumed $\varepsilon_{ij} = \varepsilon_{ji}$), the inequality (7) will sometimes be satisfied and sometimes not. Let $\tau_d \equiv \phi_g \, \alpha_g$ denote the reservation match quality for a focal parent $g \in \{m, f\}$ when a child with CP is involved, and $\tau_0 \equiv 0$ when no CP is present. The probability that a given meeting results in a match is then

$$P(\text{match} \mid \text{CP}) = 1 - F(\tau_d), \qquad P(\text{match} \mid \text{no CP}) = 1 - F(\tau_0) = 1 - F(0). \tag{8}$$

Here $F(0) = \Pr\{\varepsilon_{ij} < 0\}$, so $1 - F(0)$ is the baseline acceptance probability without CP. Since $\tau_d > 0$ and F is strictly increasing in a neighborhood of 0 (consistent with full support), we have $1 - F(\tau_d) < 1 - F(0)$. Thus, *the presence of a child with CP strictly lowers the probability that any given meeting leads to a successful partnership*, because the match quality must clear a higher hurdle $\phi\alpha$ to be acceptable. This formalizes the intuition that having a child with CP makes an individual less "matchable" in the marriage market. In economic terms, the child's CP decreases the expected surplus of a new match. While we focus here on the additional cost of having a child with CP as compared to having a typically developing child in accordance with our empirical analysis, this is consistent with the idea that a prospective partners' children from a previous union can reduce the gains from the new partnership (see Becker et al., 1977, who discuss children as a form of marriage-specific capital primarily valued by their biological parents).

It is worth noting the role of gender in these conditions. If the single focal parent is a mother, the prospective partner is a man and the relevant cost share is ϕ_m (the fraction of her care burden a new partner would take

on). If, on the other hand, the single focal parent is a father, the prospective partner is a woman, and the cost share is ϕ_f. As discussed, it is plausible that $\phi_f > \phi_m$: on average, stepmothers might contribute more to caregiving than stepfathers. This means a stepmother would bear a larger portion of the child-related cost than a stepfather, yielding higher $\tau_d = \phi_f \alpha_f$, making the acceptance condition more stringent for prospective stepmothers of children with CP than for prospective stepfathers. However, we expect $\alpha_f < \alpha_m$ as single fathers of children with CP are less often the primary caregiver and/or can outsource some care. This makes the acceptance condition less stringent for prospective stepmothers than for prospective stepfathers. Thus, in the model, the net effect of gender on the acceptance condition is ambiguous and depends on whether gender differences in α dominate gender differences in ϕ or vice versa.

If mothers of children with CP face greater overall difficulty in repartnering than fathers of children with CP (as empirical studies often find for the population in general (Ivanova et al., 2013; Schnor et al., 2017)), it could be due to a larger care burden, (α_m) and/or larger stigma (σ_m), which reduces the effective rate of viable contacts from λ to $(1 - \sigma_m)\lambda$. Also, although we expect $\phi_m < \phi_f$, the opposite is possible and would imply a higher average caregiving share expected from a prospective partner and stepfather (ϕ_m), which raises the acceptance threshold $\phi_m \alpha_m$, reducing the rate of repartnering for mothers of children with CP. This would be expected in societies with stronger preferences for gender equality.

2.3 Equilibrium Matching and Implications

In equilibrium, singles employ the acceptance rule derived above: a partnership is formed if and only if the realized match quality $\varepsilon_{i,j} = \varepsilon_{ji}$ is equal to or exceeds the appropriate reservation value (which is 0 if no child with CP is present and $\phi_g \alpha_g, g \in \{m, f\}$ when a child with CP is involved). We assume a steady-state setting where the population shares of single mothers and single fathers are constant. A formal equilibrium would require that the flow of new matches equals the flow of relationship dissolutions for each group, and that individuals optimize based on acceptance conditions given the meeting rates and utility flows. In our context, we take the reservation thresholds as derived from individual rationality (which indeed maximize each individual's expected utility in a search model with random matching (Mortensen, 1988)).

The matching process can be summarized as follows. Given random matching, the rate at which a single parent of a child with CP finds a partner is $\lambda(1-\sigma_g)[1-F(\phi_g \alpha_g)]$, where $g \in \{m, f\}$ indexes the focal parent (mother

or father). A single parent whose child is not affected by CP finds a partner at rate $\lambda[1-F(0)]$. In our empirical analysis these expressions correspond to the hazard rates of repartnering in continuous time. The presence of σ_g and α_g acts as a multiplicative dampener on the hazard for parents of children with CP: σ_g reduces the effective rate of viable contacts, while α_g (through $\phi_g\alpha_g$) raises the acceptance threshold conditional on a meeting. The implied hazard ratio is therefore

$$\frac{(1-\sigma_g)[1-F(\phi_g\alpha_g)]}{1-F(0)} < 1, \tag{9}$$

relative to an otherwise similar parent whose child does not have CP.

The model delivers several testable implications. First, as implied by Equation (9), single parents caring for a child with CP will have a lower instantaneous hazard of repartnering at each point in time than comparable parents whose children are not affected by CP. Second, the disadvantage in repartnering is expected to be larger for more severe CP: as α increases, the pool of acceptable matches shrinks and the effective hazard rate falls. Third, there may be a gender asymmetry: given typical gender roles, single mothers of children with CP might experience a larger reduction in repartnering hazards than single fathers of children with CP. Mothers often have greater caregiving obligations that curtail their time for dating (reducing the *effective* rate of viable contacts from λ to $(1-\sigma_m)\lambda$) and may be perceived by men as less available or more financially encumbered, lowering mothers' attractiveness on the marriage market (raising the acceptance threshold) while fathers are less affected (Ivanova et al., 2013; Schnor et al., 2017). In the model, the supply-side incentive for mothers to repartner (high α_m makes Δ_m large) may be offset by the demand-side reluctance of potential partners (high $\phi_m\alpha_m$ raises the required ε), resulting in prolonged singlehood for mothers relative to fathers.

Finally, the model offers intuition for interpreting our empirical strategy. The Cox proportional hazards model in the empirical section assumes that a factor like having a child with CP multiplies the baseline hazard of repartnering by some constant factor (less than 1). Our theoretical setup is consistent with this: the derived hazard ratio, $(1-\sigma_g)[1-F(\phi\alpha)]/(1-F(0))$ represents a proportional reduction in the flow probability of partnership formation. Moreover, the model highlights sources of *heterogeneous treatment effects*: for example, the proportional impact of having a child with CP on the hazard may differ by gender g and by severity of the child's CP (captured by α_g), which justifies our exploration of heterogeneous effects in the empirical analysis. If, say, mothers of children with a severe CP face the greatest penalties (high σ_m and α_m), the theory would predict and the data show

this subgroup to have the lowest repartnering rate (the most reduced hazard relative to counterparts). On the other hand, any institutional or individual factors that mitigate these costs, such as disability support payments, supportive social networks, or a new partner's altruism, would effectively lower α_g or raise ϕ_g, thereby easing the repartnering constraint.

In summary, our model illustrates how having a child with CP can create both *search friction* (fewer opportunities to meet new partners) and a *match quality hurdle* (higher standards needed to entice a new partner) in the marriage market. These theoretical mechanisms underpin the hypotheses we take to the data: namely, that caring for a child with CP significantly reduces single parents' repartnering prospects, with stronger effects for mothers than fathers and for more severe disabilities.

Dynamic mapping to empirical time variation

Our baseline theory is static, but it naturally maps to a dynamic environment in which the caregiving burden, stigma/obstacles to social opportunities, as well as sharing norms, evolve with either the child's age or the duration of singlehood. Let t denote time since the start of the single spell, and let the primitives be allowed to vary: $\alpha_g(t)$, $\sigma_g(t)$, and $\phi_g(t)$. We maintain a common baseline meeting intensity $\lambda(t)$ for all groups (absorbed by the Cox baseline hazard).

For a single parent $g \in \{m, f\}$ (mother or father), let $d \in \{0, 1\}$ indicate whether the child has CP, with $d = 1$ for CP and $d = 0$ for a typically developing child. The instantaneous repartnering hazards are

$$h_g^1(t) = \lambda(t)\Big(1 - \sigma_g(t)\Big)\Big[1 - F\Big(\phi_g(t)\,\alpha_g(t)\Big)\Big], \qquad h_g^0(t) = \lambda(t)\Big[1 - F(0)\Big]. \tag{10}$$

Hence the model implies a *time-varying* within-gender hazard ratio

$$\mathrm{HR}_g(t) = \frac{h_g^1(t)}{h_g^0(t)} = \Big(1 - \sigma_g(t)\Big)\,\frac{1 - F\Big(\phi_g(t)\,\alpha_g(t)\Big)}{1 - F(0)}\,, \tag{11}$$

which has the same multiplicative structure as our static Cox specification (Equation 9). In words, CP reduces the hazard at each t through fewer effective meetings and stigma, $1 - \sigma_g(t)$, and a higher acceptance threshold, $\phi_g(t)\alpha_g(t)$.

The time variation in the hazard ratio can be due to two mechanisms. First, a structural time-variation can be caused by the ageing of the child, leading to a change in $\alpha_g(t)$ if care needs ease or increase at specific stages, and/or a change in stigma and barriers for socialization $\sigma_g(t)$ as routines

stabilize. The second mechanism (composition or frailty) arises because unobserved heterogeneity induces selection, even when the primitives are time-invariant. If higher-hazard individuals repartner earlier, the average hazard among survivors changes over t. Both mechanisms rationalize the time-varying hazard ratios we estimate.

3 Data

3.1 Defining the dataset

Our data comprises all children with a CP diagnosis in Sweden born between 1991 and 2014. In Appendix Section A1, we describe the symptoms and causes of CP and explain in detail how we identify the children in the treatment group. For each child with CP, a control group of five children comparable to the diagnosed child in the dimensions of sex, municipality of residence at the time of first observed CP diagnosis, and birth year, has been added to the data set. Children in the control group might have other disabilities than CP. This is deliberate, since children with CP can also have other disabilities, and removing children with disabilities from the control group might exaggerate the estimated effect of CP on repartnering.[4]

The children with CP and the controls have been matched with their parents using national registries. Further, information on family status and other background variables of the cases, controls, and parents from 1990 to 2015 was linked to the data set. Thus, we follow each family (child and parents) for up to 26 years, depending on when the child was born.

We focus on parents who become single at some point, either who were single at the birth of the child or who were in a relationship when their child was born but separated or became widowed at a later time point. However, we restrict the sample to firstborn children from single births to address identification concerns. Both multiple pregnancies and first births are associated with an increased risk of a CP diagnosis (see Section 3.2), as well as parental outcomes and experiences, including relationship dynamics. Therefore, our main analysis is limited to families where the child with CP (or control child) was the mother's first live birth from a singleton pregnancy. However, we also run our analyses on the whole sample of parents who are single at some point in time, unrestricted by first and single births.

[4]With the exception of spina bifida, which is excluded from both cases and controls. The data set used was designed to study individuals with CP or spina bifida. This is thus a technical restriction related to the data construction, but not expected to affect the analysis.

We also exclude children who were adopted after birth from the final sample, as adoptive parents of children with CP may have been aware of the child's disability at the time of adoption.[5] In addition, we exclude 122 observations with missing data on the background variables used in the analysis (see Section 3.2). Our final sample consists of 6,431 children, of whom 1,004 have a CP diagnosis, and 5,427 do not.

3.2 Variables

Besides the indicator of whether a child has CP, we have information on the severity of their condition for a subsample consisting of 862 of the 1,004 children diagnosed with CP. The most common classification of functional ability of children with CP is the Gross Motor Function Classification System (GMFCS), a five-level system based on self-initiated movement abilities (Rosenbaum et al., 2007; Palisano et al., 2008). Following Müller et al. (2022), who also use the Swedish CP register to study the effect of having a child with CP on the decisions to divorce or to get additional children, we classify children with the expanded and revised GMFCS classification 1-3 as less severely affected (ability to initiate movement with or without the support of manual devices), and children with GMFCS classifications 4 and 5 as more severely affected (unable to self-initiate movement). GMFCS has been found to correlate strongly with intellectual disability (Reid et al., 2018) and communication and manual ability (Compagnone et al., 2014) and is often used in the disability literature as an overall measure of severity. This subset is used to study whether the severity of the CP diagnosis (captured in α_g) matters for repartnering.

The Swedish register data allows us to observe parents' civil status and whether parents live together each year using Statistics Sweden's family identification numbers. Thus, we have information on whether or not parents are living together at the time of birth and subsequent years. We observe repartnering of a parent when the parent either marries or has an additional child with a new partner with whom they are cohabiting. This definition is broader than remarriage-only measures used in some prior work (e.g., Mauldon (1992)), but it still does not capture cohabiting unions that do not transition to marriage or a shared birth. Unfortunately, we do not observe new relationships when neither of the above is fulfilled. Our definition of repartnering thus captures relatively strong relationship commitments, excluding more casual cohabitations or relationship formations that do not

[5]However, we retain four same-sex female couples in which the non-biological mother adopts the child.

result in either marriage or a shared child. This also means that our analysis relies on the assumption that the propensity to transition from a cohabiting relationship (or other forms of serious relationships) without a joint child to either marriage or having a child is the same for the treatment and control groups. We discuss the credibility of this assumption in Section 4.1 and Section 5.3.

Table 1 presents summary statistics of parents' relationship status depending on the CP status of the child. Descriptively, 25.9% of single mothers in the control group repartner, compared to only 19.9% of single mothers to children with CP. Single fathers repartner at higher rates with smaller difference between groups (27.0 vs. 25.9%). This suggests a possible negative effect of child disability for mothers, which we examine formally with survival models.

Our estimations include a rich set of background variables capturing both birth characteristics and parental socioeconomic status. Summary statistics for these variables, disaggregated by the child's CP status, are presented in Table 2. In addition to the variables shown in the table, we also control for the region of birth and include birth-year dummies grouped into two-year clusters.

Previous studies have found the prevalence of CP to be higher for male children (Chounti et al., 2013), children born in multiple pregnancies (Korzeniewski et al., 2018), and children of younger and older mothers (Schneider et al., 2018; Zhou et al., 2023). Also, people of low socioeconomic status have been found more likely to have children diagnosed with CP (Oskoui et al., 2013; Solaski et al., 2014; Michael-Asalu et al., 2019), although Müller et al. (2022) found no association in Sweden using a subset of the data used here. As both the characteristics of the birth and parents' socioeconomic status are likely to play a role in parents' decisions to divorce and repartner, in Table 3 we examine the extent to which CP can be predicted by these factors. We do this for our main sample of parents of first-born children in single births (Model A), as well as for the whole sample of parents who are single at some point in time (Model B).

The results for Model B are in accordance with the concern that the variables first-born and multiple birth are strongly related to CP. This supports the decision to focus only on first-born children from single pregnancies in our analysis. Turning to Model A, the only statistically significant predictor of CP is mother's age. The adjusted R^2 of model A is very low (<0.01), and the F-statistic is not statistically significant ($p = 0.18$), suggesting that parental socioeconomic factors explain CP to a limited extent in our data. Still, because mothers' age is associated with CP and is a likely determinant of repartnering behavior, we will scrutinize its role in our robustness analysis.

Table 1: Marriage market summary statistics for parents of firstborn children in single births

	No CP	**CP**	**Mild CP**	**Severe CP**
Mothers who are single at some point	5192	951	564	257
	(37.2)	(38.3)	(37.7)	(39.9)
Non-single mothers	8783	1533	933	387
	(62.8)	(61.7)	(62.3)	(60.1)
If single:				
Mothers who repartner†	1344	189	109	51
	(25.9)	(19.9)	(19.3)	(19.8)
Average number of years until single‡	4.5	3.8	3.8	3.5
	(5.0)	(4.9)	(4.9)	(4.4)
If repartnered:				
Average number of years until repartnered§	5.6	6.2	6.3	5.5
	(3.6)	(4.1)	(4.4)	(3.4)
Fathers who are single at some point	4972	911	544	240
	(35.8)	(37.0)	(36.8)	(37.6)
Non-single fathers	8897	1554	934	398
	(64.2)	(63.0)	(63.2)	(62.4)
If single:				
Fathers who repartner†	1330	234	137	58
	(26.8)	(25.7)	(25.2)	(24.2)
Average number of years until single‡	4.8	4.2	4.1	3.8
	(5.0)	(5.0)	(4.9)	(4.5)
If repartnered:				
Average number of years until repartnered§	5.4	5.9	6.1	5.1
	(3.9)	(3.8)	(3.7)	(3.2)

Notes: Standard deviation or percentages in parentheses. †Percentage of single mothers/fathers who repartner. ‡Average number of years between the birth of the child until becoming a single parent. §Average number of years from becoming a single parent until repartnering occurs, conditional on repartnering.

Table 2: Summary statistics of single parents of firstborn children in single births

	No CP	**CP**	**Mild CP**	**Severe CP**
Child: female	42.58	42.13	39.70	39.70
Mother immigrated	13.83	13.56	14.72	11.67
Father's income: 1st quintile	28.74	30.07	30.67	29.57
Father's income: 2nd quintile	24.13	23.03	21.28	24.51
Father's income: 3rd quintile	20.07	19.14	18.44	21.01
Father's income: 4th quintile	15.33	16.19	17.73	12.45
Father's income: 5th quintile	11.73	11.57	11.88	12.45
Child $<$ 1 year when mother single	39.27	46.90	46.81	48.25
Child 2-6 years when mother single	32.34	29.34	30.14	29.57
Child 7-10 years when mother single	14.23	11.67	10.82	13.23
Child $>$ 11 years when mother single	14.16	12.09	12.23	8.95
Region of birth: Götaland	44.07	44.37	44.86	42.80
Region of birth: Svealand	40.93	39.01	39.54	38.13
Region of birth: Norrland	15.00	16.61	15.60	19.07
Mother's education: Mandatory	21.86	22.50	22.87	20.23
Mother's education: Secondary	51.93	51.21	48.76	54.47
Mother's education: Higher	26.21	26.29	28.37	25.29
Mother single at 29 years or younger	46.94	45.85	44.50	48.25
Mother single at 30-36 years	30.82	30.81	32.27	30.35
Mother single at 37 years or older	22.25	23.34	23.23	21.40
Birthplace: In or around a large city	38.12	37.12	37.59	35.41
Birthplace: In or around a medium size town	39.71	39.33	39.72	39.30
Birthplace: In or around small towns, rural	22.17	23.55	22.70	25.29
Father immigrated	16.07	15.37	15.26	14.17
Mother's income: 1st quintile	33.69	35.57	35.11	34.17
Mother's income: 2nd quintile	23.97	21.41	20.77	22.92
Mother's income: 3rd quintile	17.86	17.01	17.83	14.17
Mother's income: 4th quintile	14.24	15.92	15.07	19.58
Mother's income: 5th quintile	10.24	10.10	11.21	9.17
Child $<$ 1 year when father single	35.58	43.25	43.20	44.58
Child 2-6 years when father single	34.21	30.74	31.99	30.42
Child 7-10 years when father single	15.06	12.84	11.76	15.42
Child $>$ 11 years when father single	15.14	13.17	13.05	9.58
Father's education: Mandatory	21.68	22.83	22.06	25.42
Father's education: Secondary	56.15	55.21	53.31	54.58
Father's education: Higher	22.16	21.95	24.63	20.00
Father single at 32 years or younger	45.84	45.66	45.40	48.33
Father single at 33-40 years	32.54	32.27	33.46	31.25
Father single at 41 years and older	21.62	22.06	21.14	20.42
Region of birth: Götaland	44.17	43.80	44.49	42.50
Region of birth: Svealand	40.43	38.86	39.34	36.25
Region of birth: Norrland	15.41	17.34	16.18	21.25
Birthplace: In or around a large city	37.83	36.66	37.13	33.33
Birthplace: In or around a medium size town	39.24	38.97	39.89	38.33
Birthplace: In or around small towns, rural	22.93	24.37	22.98	28.33
Observations - children	5427	1004	592	270
Observations - mothers	5192	951	564	257
Observations - fathers	4972	911	544	240

Notes:Parents' age groups and income groups are defined as terciles and quintiles for mothers/fathers in the control group in the whole sample (including non-single parents).

Table 3: Predictors of CP

	Model A			**Model B**		
	Coeff	p	robust p	Coeff	p	robust p
Firstborn				0.028	<0.01	<0.01
Multiple birth				0.309	<0.01	<0.01
Child: female	0.000	0.97	1.00	0.002	0.72	1.00
Mother immigrated	-0.010	0.53	1.00	-0.002	0.81	1.00
Mother's education: Secondary	-0.007	0.59	1.00	-0.018	0.03	0.57
Mother's education: Higher	-0.015	0.37	1.00	-0.029	0.01	0.17
Mother: 26-32 years at birth	0.021	0.10	0.93	0.007	0.41	1.00
Mother: >33 years at birth	0.081	<0.01	<0.01	0.034	<0.01	0.06
Mother's income: 2nd quintile	-0.012	0.40	1.00	-0.001	0.89	1.00
Mother's income: 3rd quintile	-0.024	0.13	0.97	0.003	0.78	1.00
Mother's income: 4th quintile	-0.018	0.31	1.00	0.000	0.97	1.00
Mother's income: 5th quintile	-0.031	0.13	0.97	0.000	0.98	1.00
Father immigrated	-0.016	0.31	1.00	-0.006	0.58	1.00
Father's education: Secondary	-0.007	0.58	1.00	-0.005	0.57	1.00
Father's education: Higher	-0.015	0.35	1.00	-0.014	0.18	0.99
Father: 26-32 years at birth	0.007	0.61	1.00	0.009	0.38	1.00
Father: >33 years at birth	0.005	0.78	1.00	0.011	0.34	1.00
Father's income: 2nd quintile	-0.025	0.07	0.84	-0.019	0.03	0.57
Father's income: 3rd quintile	-0.017	0.25	1.00	-0.017	0.07	0.85
Father's income: 4th quintile	0.006	0.72	1.00	-0.015	0.15	0.98
Father's income: 5th quintile	-0.014	0.47	1.00	-0.007	0.56	1.00
Region of birth: Svealand	-0.008	0.60	1.00	-0.011	0.25	1.00
Region of birth: Norrland	0.017	0.26	1.00	-0.002	0.84	1.00
Region of birth: Stockholm	0.009	0.67	1.00	-0.006	0.68	1.00
Birthplace: In/around medium-sized towns	0.006	0.69	1.00	0.006	0.52	1.00
Birthplace: Rural	0.016	0.32	1.00	-0.001	0.89	1.00
Child born 2003-4	0.002	0.95	1.00	-0.012	0.42	1.00
Child born 2005-6	-0.026	0.27	1.00	-0.008	0.60	1.00
Child born 2007-8	0.008	0.74	1.00	0.023	0.14	0.98
Child born 2009-10	-0.027	0.29	1.00	0.020	0.24	1.00
Child born 2011-12	-0.011	0.69	1.00	0.015	0.41	1.00
Child born 2013-14	0.020	0.60	1.00	0.006	0.82	1.00
Child born 1991-2	-0.010	0.62	1.00	0.006	0.66	1.00
Child born 1993-4	-0.016	0.45	1.00	-0.004	0.74	1.00
Child born 1995-6	-0.016	0.48	1.00	-0.003	0.83	1.00
Child born 1997-8	-0.005	0.84	1.00	-0.011	0.44	1.00
Child born 1999-2000	-0.008	0.71	1.00	-0.002	0.89	1.00
Number of observations		5596			12835	
R-squared		<0.01			0.01	
F-test (p-value)		0.18			<0.01	

Notes: Results of OLS estimation. Robust p-values are corrected for multiple hypothesis testing using the Romano Wolf procedure

4 Empirical strategy

We estimate how the repartnering hazard, h, of single mothers and fathers is affected by their child having a CP diagnosis using the Cox proportional hazards (PH) model, as predicted by our search-and-matching framework. The hazard at time t is the instantaneous rate of repartnering among parents who are still single at time t. Formally,

$$h(t) = \lim_{\Delta t \to 0} \frac{\Pr(t \leq T < t + \Delta t \mid T \geq t)}{\Delta t}, \tag{12}$$

where T denotes the time to repartnering. The Cox proportional hazard model is then defined as

$$h(t) = h_B(t) e^{\beta CP + \gamma' x}, \tag{13}$$

where $h_B(t)$ is the baseline hazard, CP status is the main independent variable of interest and γ is a vector of background variables included as controls. For the variable CP, we estimate e^{β}, the hazard ratio of CP status. When the hazard ratio equals one, CP status implies no change in $h(t)$. In our theoretical framework, this corresponds to the case where CP does not affect either the effective meeting rate ($\sigma_g = 0$) or the acceptance threshold ($\alpha_g = 0$). A value above (below) one implies an increase (decrease) in $h(t)$. Note that the hazard ratio e^{β} maps directly to the factor $(1 - \sigma_g) \frac{1 - F(\phi_g \alpha_g)}{1 - F(0)}$ implied by our search-and-matching model (Section 2).

Throughout, we report effects in terms of hazard ratios and refer to changes in the *hazard* (instantaneous rate) rather than the probability or likelihood of repartnering, unless explicitly stated otherwise. When repartnering is not observed within the follow-up window, the observation is treated as right-censored.

A central assumption of the Cox proportional hazards model is that the effect of a unit change in an independent variable remains constant over time.[6] For the model to be correctly specified, the proportional hazards assumption must hold not only for the main covariate of interest (CP status), but also for covariates whose non-proportionality could bias the estimated effect of CP. A violation of this assumption may lead to model misspecification and biased estimates of the hazard ratio.

[6] This follows from Equation 13, where the ratio of predicted hazards for two groups (e.g., parents of children with and without CP), $h_1(t)/h_0(t)$, remains proportional over time as the baseline hazard $h_B(t)$ cancels out.

To assess whether the assumption holds, we examine whether the Schoenfeld residuals are constant over time.[7] If the residuals show systematic variation, this indicates that the proportional hazards assumption is violated, i.e., that the effect of the covariate varies over time.

To address violations of the proportional hazards assumption for such covariates, we apply stratification. This allows the model to estimate separate baseline hazard functions for different strata, thereby allowing the baseline hazard to differ flexibly across strata and avoiding the need to impose proportionality for these covariates. When the proportional hazards assumption is violated for our main covariate of interest, CP status, we implement two alternative approaches: (i) estimating time-dependent hazard ratios, and (ii) estimating weighted average hazard ratios. Time-dependent hazard ratios are obtained by extending the Cox model to include an interaction between CP status and a function of time, $\psi(t)$, which captures the temporal dynamics of the hazard ratio. When problematic background variables are also stratified, Equation (13) becomes:

$$h(t \mid S_1, \ldots, S_n) = h_B^{(s)}(t) e^{\beta_1 CP + \beta_2 CP \cdot \psi(t) + \gamma' x}, \tag{14}$$

where x is a vector of the remaining background covariates, $S_1, \ldots, S_n$ are the background variables on which we stratify the model, s indexes the resulting strata, $h_B^{(s)}(t)$ denotes the stratum-specific baseline hazard, and $\psi(t)$ is a user-specified function of time.

Since one of the primary objectives of this paper is to compare the effect of having a child with CP on the repartnering hazard between mothers and fathers, it is crucial to use identical model specifications for both groups. Therefore, we apply the methods described above to both mothers and fathers, even though the proportional hazards assumption holds for mothers and a standard Cox hazard ratio could, in principle, be estimated. For the time-dependent models, we specify the time function $\psi(t)$ using natural splines, a flexible, data-driven approach that allows for variation in time trends between groups. To prevent overfitting while still capturing relevant temporal dynamics, we use two degrees of freedom in the spline specification. This approach ensures that any differences in time trends are appropriately modeled across the estimations.

While time dynamics are important, it is also informative to summarize the effect over the observed time period in one measure. To assess the overall effect of CP on repartnering when the proportional hazards assumption does not hold, weighted average hazard ratios can be estimated, as suggested in

[7] Schoenfeld residuals measure the difference between the observed covariate value for an event and the value expected under the Cox model.

Schemper et al. (2009). Instead of assuming that the hazard ratio is constant over time, it is allowed to vary and the weighted average is calculated. To reduce the influence of a small number of late-time events on the weighted average hazard ratio, we truncated extreme weights (above the 95th percentile). This increases the robustness of the estimated average weighted hazard ratio while retaining all observations.

In our main analysis, we apply the methods described above to estimate the hazard ratio associated with having a child diagnosed with CP, separately for mothers and fathers. We estimate both the overall effect of a CP diagnosis and the effects by severity. Finally, we explore socioeconomic heterogeneity in these estimates.

4.1 Robustness analyses

First, since our main results are estimated to include only first-born children from single births, we test whether our results generalize to all parents who are single at some point in time.

Second, while the background variables included in our analysis are in line with the controls used in the bulk of literature on the topic (Hjern and Thorngren-Jerneck, 2008; Tseng et al., 2018), Müller et al. (2022) found that the physical and mental health of parents was a predictor of CP using the same data set as in this paper for the period 2004-2015. At the same time, the physical and mental health of parents can be related to the probability that they separate and potentially repartner (see e.g. Dupuy and Galichon, 2014; Guner et al., 2018; Pevalin and Ermisch, 2004). Consequently, the parents' health is a potentially omitted variable that could bias our results.

To estimate the magnitude of this potential bias, we re-estimate our main results for the years for which parental health variables are available, with and without parents' health variables included as covariates. Specifically, we include a dummy for whether the mother smoked in the three months before pregnancy (only mothers' estimation), mothers'/fathers' mental health diagnoses, and mothers'/fathers' CCI[8], both during the three calendar years before the birth of the child. Thus, even though these variables are available for 2001-2015, there is a need for a three-year lag in our estimation approach to capture pre-event health conditions. This results in estimates with parental health controls for the years 2004-2015.

[8]The Charlson Comorbidity Index (CCI) is a method for categorizing comorbidities that affect patients' mortality risk. Each defined comorbidity category is given a weight between 0 and 6, which is summed to a single comorbidity score (Charlson et al., 1987). Although created to capture mortality risk, it is often used to measure general health.

Third, we analyze if the lack of information on cohabiting relationships without common children poses a problem for our results (see Section 3.1). Müller et al. (2022) found that parents of children with CP are less likely to have additional children (in the current relationship). It could be that they are also less likely to have additional children in new relationships. If this is the case, our results could be driven by the fact that we simply do not observe repartnering amongst parents of children with CP. As a motivation that this is not the case, we analyze whether parents who have remarried without having a common child differ in their propensity to later have additional children between the treatment and the control group.

Fourth, we study whether the treatment group was more likely to divorce before the birth of the child. Since we do not have information on unmarried cohabiting relationships without common children, we focus on divorces rather than separations in general. This exercise sheds light on whether the treatment group is inherently more prone to separation than the control group, and thus if our results are driven by having a child with CP or selection.

5 Results

5.1 Testing the proportional hazards assumption and model selection

In this section, we assess the proportional hazards assumption for the baseline model defined in Equation (13). When violations of this assumption are detected, we modify the model accordingly. This ensures that we obtain valid coefficient estimates for CP status.

Appendix Table A1 illustrates the results of Schoenfeld tests for both CP status and the included background variables for mothers and fathers, overall and by severity of CP (Equation 13). First, there is evidence against the proportional hazards assumption for fathers for both overall CP status and mild CP ($p = 0.03$ and $p = 0.04$, respectively), and weaker evidence at the 10% level for mothers ($p = 0.08$ for CP and $p = 0.07$ for mild CP). Second, for both mothers and fathers, child age at the start of the single spell, parent age at the start of the single spell, education, income, and the child's birth-year indicators all show evidence of non-proportionality. Additionally, degree of urbanization fails the test for mothers. Since our main interest is the estimated effect of CP, we pay particular attention to covariates whose non-proportionality could bias that estimate.

We begin by addressing covariates that violate the proportional hazard

assumption. To test which of the variables are confounders, we study (i) the relationship between the background variables and repartnering in Appendix Table A2 and (ii) the association between the background variables and CP status presented in Appendix Table A3. Comparing these results with the Schoenfeld tests in Appendix Table A1, we find that, for both mothers and fathers, parent age and child age at the start of the single spell are the only covariates that are associated with both CP status and repartnering and also violate the proportional hazards assumption.

We therefore estimate a stratified model, as in Equation 14, in which the baseline hazard is allowed to vary across strata defined by these two variables, thereby accommodating their non-proportional effects. We then repeat the Schoenfeld test for the updated, stratified model and present the results for the main variable, CP status, only. The results, presented in Table 4, show that even after stratification, CP status continues to violate the proportional hazards assumption for fathers, suggesting that the effect of CP on fathers' repartnering varies over time.

Table 4: Schoenfeld tests of Equation 14

	Mothers		**Fathers**	
	(1)	(2)	(3)	(4)
CP	0.21		0.04	
Mild CP		0.15		0.06
Severe CP		0.56		0.90
Global	<0.01	<0.01	<0.01	<0.01

Notes: P-values of Schoenfeld tests, testing the null hypothesis that the Schoenfeld residuals are independent of time. P-values lower than 0.05 suggest time dependence of residuals and thus violation of the proportional hazard assumption.

Consequently, the time dependence of the CP effect must be addressed in the estimation. To account for the violation of the proportional hazards assumption, we estimate both a time-varying hazard ratio model (see Equation (14)), which captures how the hazard ratio evolves over time, and a weighted average hazard ratio, which summarizes the time-dependent effect in a single measure.

5.2 Main results

Table 5 reports weighted-average hazard ratios for repartnering associated with having a child with CP for both mothers and fathers. In addition to our main specification, which controls for background variables, we also present estimates that omit stratification and covariate adjustment to capture the overall effect of having a child with CP. The results show that, after adjusting for socioeconomic background and birth characteristics, mothers of children diagnosed with CP have, on average, a 23% lower repartnering hazard than comparable mothers in the control group. In our framework, this reduced-form effect reflects two channels: (i) a higher acceptance threshold, $\phi_m \alpha_m$, and (ii) ex-ante screening/limited availability, summarized by σ_m (which reduces the effective rate of viable contacts from λ to $(1-\sigma_m)\lambda$). The Cox model identifies only their combined impact. For fathers, we do not find a statistically significant effect.

Columns (3) and (6) of Table 5 present estimates disaggregated by CP severity. The results do not provide evidence that the repartnering penalty is larger for severe than for mild CP. This is not in accordance with the theoretical prediction that severity increases α_g, which should reduce repartnering rates. One possible explanation is that, in the Swedish institutional setting, support tends to increase with care needs, especially through access to personal assistance, so that measured severity may translate imperfectly into the effective parental burden captured by α_g.

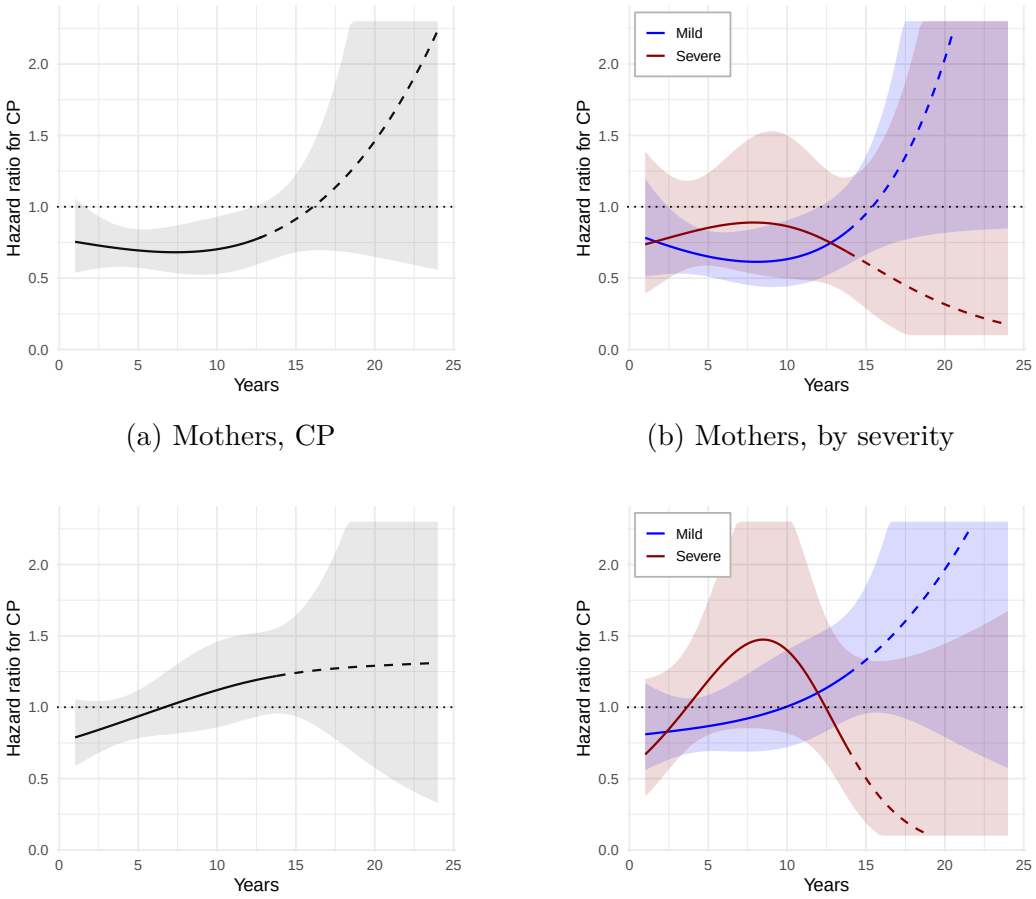
Figure 1 presents estimates from the Cox model with a time-varying hazard ratio. The figure shows that while the hazard ratios for mothers and fathers are similar one year after becoming single, they evolve differently over time: the hazard ratio for mothers remains relatively stable, whereas it increases for fathers. This pattern is consistent with the dynamic dampening factor $(1-\sigma_g(t))\,\frac{1-F(\phi_g(t)\alpha_g(t))}{1-F(0)}$ in Section 2. For both mothers and fathers, the hazard ratio associated with having a child with mild CP follows a convex pattern over time, while the estimated curve for severe CP exhibits an inverse U-shape. It is important to note that 95% of observed repartnering events occur within the first 13 years for mothers and 14 years for fathers. This is reflected in the widening confidence intervals in subfigures (a) and (c) beyond these time points. To highlight this, the estimated curves beyond this limit are represented with a dotted line in Figure 1. As such, caution is warranted when interpreting the shape of the curves over longer time horizons.

Table 6 presents a heterogeneity analysis by parents' and children's age at the start of single spell, education, immigration background, and income. For mothers, we observe that the effects are strongest for the youngest group and mothers whose child is 1 year old or younger. However, the differences be-

Table 5: Weighted average hazard ratios of repartnering

	Mothers			**Fathers**		
	(1)	(2)	(3)	(4)	(5)	(6)
CP	0.753***	0.767***		0.955	0.967	
	(0.076)	(0.075)		(0.070)	(0.069)	
Mild CP			0.757***			0.948
			(0.096)			(0.087)
Severe CP			0.771*			0.909
			(0.137)			(0.132)
Backgr. var.		✓	✓		✓	✓
Observations	6,143	6,143	6,013	5,883	5,883	5,756

Notes: Results are expressed as weighted average hazard ratios. Standard errors in parentheses. *p<0.1, **p<0.05, ***p<0.01. Background variables included as covariates are: child's gender, mother's/father's education, mother's/father's income at birth, immigration status of the mother/father, the child's year of birth (in clusters of two years), information on the place of birth (region and the degree of urbanization of the birth municipality.). Background variables stratified over are: mother's/father's age and child age when single spell begins.



(a) Mothers, CP

(b) Mothers, by severity

(c) Fathers, CP

(d) Fathers, by severity

Figure 1: Time-varying hazard ratio for repartnering associated with having a child diagnosed with CP. Dotted lines indicate portions of the estimated curves that lie beyond the 95th percentile of observed time to repartnering. Background variables included as covariates are: child's sex, mother's/father's education, mother's/father's income at birth, immigration status of the mother/father, the child's year of birth (in clusters of two years), information on the place of birth (region and the degree of urbanization of the birth municipality). Variables used for stratification are: mother's/father's age and child's age when single spell begins.

tween the estimated coefficients are not statistically significant. For fathers, we observe little or no effect among younger fathers, those with primary or secondary education, and those in the lower income quintiles. In contrast, the estimated effects tend to be larger for older, highly educated, and high-income fathers and some of these differences are statistically significant. In the table, total effects marked in red are statistically different from total effects marked in blue.

Table 6: Socioeconomic heterogeneity in the effect of having a child with CP on repartnering

		Mothers		**Fathers**	
		Total effect	SE	Total effect	SE
(1)	Mothers <30 yo/Fathers<33 yo	0.73***	0.09	0.99	0.08
	Mothers 31-36 yo/Fathers 34-40 yo	0.89	0.15	0.97	0.14
	Mothers 37+/Fathers 41+	0.74	0.31	0.8	0.25
(2)	Child 1 year or less when single	0.68***	0.12	**0.85**	0.11
	Child 2-6 years when single	0.83	0.11	1.03	0.10
	Child 7-10 years when single	1.01	0.23	**1.25**	0.20
	Child 11 year or older when single	0.56	0.39	0.92	0.28
(3)	Mandatory education	0.88	0.13	0.96	0.14
	Secondary education	0.73***	0.10	1.04	0.09
	Higher education	0.7*	0.21	0.76	0.17
(4)	Parent born in Sweden	0.77***	0.08	0.95	0.07
	Parent immigrated	0.7	0.29	1.06	0.20
(5)	Income: 1st quintile	0.75**	0.12	1.04	0.11
	Income: 2nd quintile	0.87	0.14	**1.1**	0.14
	Income: 3rd quintile	0.63**	0.18	0.96	0.16
	Income: 4th quintile	0.74	0.23	**0.73**	0.21
	Income: 5th quintile	0.98	0.33	0.66	0.31

Notes: Total effect of having a child with CP by parents' and child's age at the start of the single spell, education, income, and immigration status. Results presented as weighted average hazard ratios. *p<0.1, **p<0.05, ***p<0.01. For each variable (age, education, income, and immigration status), a weighted average proportional hazard is estimated for a model including an interaction between CP and the variable in question. The total effect for a given group is calculated as the sum of the coefficient for CP and the interaction between CP and the group in question. Total effects in red are significantly different (p values uncorrected for multiple hypothesis testing) from total effects in blue: A child that is 7-10 years old vs. a child that is 1 year or less, for fathers (p=0.095) and fathers in the 4th income quintile vs. fathers in the 2nd income quintile (p=0.097).

5.3 Robustness results

5.3.1 Robustness to using the full sample

Appendix Tables A4 and A5 report summary statistics for the full sample of parents who are single at some point in time. Relative to our main estimation sample, the full sample differs primarily in age composition and related variables such as income. Repartnering rates are also slightly lower for both mothers and fathers, which is consistent with the older age profile of the full sample.

We then re-estimate the weighted average hazard ratios from our main specification using the full sample. The results, reported in Appendix Table A6, are broadly similar to those in Table 5, suggesting that our main conclusions are not driven by the restriction to first-born children from singleton births.

5.3.2 Robustness to parental health controls

Summary statistics for the included parental health variables over the period 2004–2015, during which parental health is observed, are presented in Appendix Table A7. These statistics indicate that, on average, parents of children with CP exhibit poorer health outcomes than parents in the control group.

To assess whether omitting health variables biases our main results, we compare estimates from models with and without parental health controls, keeping all other aspects identical. The results, shown in Appendix Table A8, reveal minimal differences in the estimated effects. This suggests that, in the subsample for which parental health is observed, omitting these controls does not materially affect the estimated association between CP and repartnering.

Although the primary purpose of these estimations was not to analyze changes over time, an interesting pattern emerges. Comparing the results in Appendix Table A8 to those in Table 5, we observe that the hazard ratios, especially for fathers, are notably lower during the 2004–2015 period than over the full 1991–2015 period. This suggests that the negative effect of having a child with CP on repartnering may have become more pronounced over time, especially for fathers. This more equal distribution across gender over time could potentially be capturing a trend towards less gendered child-caring roles.

5.3.3 Robustness to unobserved repartnering

We cannot fully rule out the possibility that some of our results are influenced by a lower likelihood of having additional children in the treatment group, which could increase the chance of unobserved repartnering. Although we cannot directly observe all repartnering events, we can examine whether there is a difference in the likelihood of having an additional child among parents who remarry without having had a common child prior to the marriage. This serves as a test of whether our results might be driven by a higher propensity among repartnered but unmarried parents without a child with CP to have additional children.

The results of this analysis, presented in Appendix Table A9, show no significant association between having a child with CP and the likelihood of having an additional child within a new marriage. Assuming that fertility behavior among unmarried repartnered parents is similar to that of remarried parents, the results suggest that it is unlikely that unobserved repartnering biases our main findings.

5.3.4 Pre-birth divorce as a selection check

To assess whether parents of children with CP are inherently more likely to divorce—and thus select into singleness differently than other parents, we compare divorce rates between the treatment and control groups prior to the birth of the child. For our main data consisting of parents of firstborn children in single births, the percentage of parents in the treatment group that divorced before the birth of the child is 4.5% while the same percentage in the control group is 4.1%, the difference is not statistically significant (p=0.56 in a two-sided t-test). For the whole sample, the divorce rate prior to the child's birth is 9.1% versus 8.4% in the control group. Again, the difference is not statistically significant (p=0.27 in a two-sided t-test). Thus, we fail to find evidence of selection bias.

5.3.5 Heterogeneity by type of repartnering

In our main analysis, repartnering clusters two types of repartnering: when marriage takes place or when a common child is born to a cohabiting couple. As these two groups might differ, we estimate the main model for them separately and present the results in Appendix Table A10. The results show that the hazard ratio of CP for mothers who repartner by having a child in a new cohabiting relationship is lower than for remarried mothers, suggesting that the former group is more affected by having a child with CP. The difference between the two groups is statistically significant at the 5% level. Notably,

when we restrict the endpoint to remarriage only, the estimated effect for mothers is small and statistically indistinguishable from zero (Appendix Table A10), which is consistent with Mauldon (1992)'s finding of no association between child disability/health limitations and mothers' remarriage hazards.

A likely reason for this is that mothers who remarry before having a child in the new relationship are, on average, older at the time they become single (30 years, as compared to 26 years for those who repartner by child), and the child is, on average, older when this happens (4.4 years, as compared to 2.7 years for those who repartner by child). As observed in Table 6, higher age and older child when single spell starts is related to a hazard ratio that is closer to 1 and less statistically significant.

6 Discussion

We study how having a child with a disability (CP specifically) affects single parents' repartnering hazards using Swedish register data and Cox proportional hazards models. We estimate how the hazard of repartnering differs between parents of children diagnosed with CP and otherwise comparable parents whose child does not have CP. Because the data exhibit non-proportional hazards, we report both weighted-average hazard ratios and time-varying hazard ratios. Our results suggest that mothers of children diagnosed with CP have, on average, a 23% lower repartnering hazard than mothers in the control group, whereas for fathers we find no statistically detectable difference. We also do not find evidence that the estimated effects differ systematically by severity, although the estimates for fathers with severe CP are somewhat more negative and imprecise.

These findings are consistent with the search-friction and match-quality mechanisms outlined in Section 2, and with a disproportionate burden on single mothers. They also align with earlier work emphasizing the increased caregiving responsibilities and obstacles to repartnering faced by single mothers (Hatton et al., 2010; Tøssebro and Wendelborg, 2017). Our study extends this literature by quantifying repartnering outcomes and by providing evidence on fathers, a group that has been largely absent from prior work. The absence of a statistically detectable effect for fathers is consistent with evidence that children with disabilities are disproportionately likely to live with women, and especially with single mothers, after parental separation (Cohen and Petrescu-Prahova, 2006).

Our results also relate to earlier evidence on maternal remarriage. Using U.S. survey data, Mauldon (1992) estimates a Weibull hazard model for time to mothers' remarriage after divorce and reports no statistically

significant association between child disability/health limitations and the remarriage hazard. This is consistent with our remarriage-only specification (Appendix Table A10). In contrast, the more pronounced effect in our main endpoint arises because we also capture repartnering through cohabitation that progresses to a shared birth, a pathway not reflected in remarriage-only measures.

Having children has been suggested to reduce social interaction and therefore the likelihood of meeting a new partner (de Graaf and Kalmijn, 2003). In our framework, such reductions operate through σ_g, which lowers the effective rate of viable contacts (from λ to $(1-\sigma_g)\lambda$). This could be more pronounced in the case of disability, especially if caring for the child reduces employment opportunities and thus the likelihood of meeting a new partner through work (de Graaf and Kalmijn, 2003). Our findings also complement recent descriptive evidence from the UK showing that single parents of children with disabilities experience substantial health and economic disadvantage relative to both single parents of non-disabled children and partnered parents (Balbo and Burciu, 2026). However, prior work has also hypothesized a countervailing force: the need for support may increase a single parent's willingness to repartner (Bumpass et al., 1990). In our framework, repartnering relieves a share of the caregiving burden ($\phi_g\alpha_g$), which makes the focal parent more willing to accept a match. But because match formation requires mutual acceptance, the binding condition is the prospective partner's requirement that match quality be high enough to offset that burden (threshold $\phi_g\alpha_g$), while the rate of viable contacts may also be reduced by σ_g. The observed hazard is therefore governed by the stricter partner-side threshold together with any reduction in viable contacts.

The results of the current study must be understood in the Swedish context, where a tax-financed universal healthcare system reduces health-insurance-related incentives for repartnering. A relatively comprehensive social insurance system, including caregiver benefits, attendance allowances, transport, and care subsidies, compensates parents/families for lost income and/or increased expenses (Alriksson-Schmidt et al., 2021). This is expected to reduce the need to repartner due to financial reasons. A key support in social insurance is the user-controlled personal assistant, which is also expected to somewhat alleviate the need for repartnering for caregiving purposes. Although there are general benefits within the social insurance system that are income-based, which may thus be affected by the increased family income from repartnering, benefits connected to disability are generally need-based and thus unaffected by repartnering. That said, if the financial gains from repartnering (via resource pooling or interactions with programs) differ systematically between CP and non-CP parents, our estimates will also reflect

those differences. In that case, the reported hazard ratios should be read as a reduced-form composite of (i) the effective rate of viable contacts (screening/availability), (ii) a tougher acceptance threshold arising from caregiving demands, and (iii) any group-specific financial/resource incentives to repartner. These channels are not separately identified in our design.

As noted above, our endpoint captures higher-commitment partnerships (legal marriage or a shared child in cohabitation), and we do not observe repartnering without a marriage or a common child. This is a limitation of our study. However, we find no evidence that the likelihood of remarriage differs between parents with versus without a child with a disability in our data (Appendix Table A10), consistent with prior evidence on maternal remarriage (Mauldon, 1992). In the Swedish context, cohabitation is common and ensures mainly the same rights as those associated with marriage. The major differences are related to dissolution of the relationship and inheritance, factors that are not expected to be of special concern with respect to disability. While this does not rule out all forms of unobserved repartnering, it reduces concerns that differential transitions from cohabitation to our endpoints drive the main results.

Prior work in the general population shows that simply having children substantially lowers repartnering for mothers (e.g., Ivanova et al., 2013; Schnor et al., 2017; Di Nallo et al., 2023). Our estimates add an additional penalty specific to caring for a child with CP. Taken together, a back-of-the-envelope interpretation is that mothers of children with CP face a markedly larger overall disadvantage than childless women. If the motherhood penalty in prior studies implies roughly a 20-40% lower repartnering rate for mothers, combining it with our estimated CP-related reduction (about 23%) suggests an overall reduction on the order of 40–55% relative to childless women.[9] Because several studies find that childless women and childless men repartner at similar rates, the same back-of-the-envelope calculation conveys the gap relative to childless men as well (Ivanova et al., 2013; Schnor et al., 2017; Di Nallo et al., 2023). In short, our results imply that the motherhood penalty documented in the broader literature is amplified when the child has a disability.

Despite the strengths of the data, the study has limitations. Although robust, our reliance on Swedish national register data may limit external validity beyond Nordic and similar institutional settings. Our findings nevertheless have implications for policymakers and support services addressing the unique challenges faced by single mothers of children with disabilities.

[9]These figures are illustrative and rely on multiplying effects reported in different contexts; they should be read as approximate magnitudes rather than point estimates.

For targeted policy responses, future research could benefit from exploring the underlying mechanisms driving these gender disparities and examining how existing support structures influence the likelihood of repartnering. This is of particular interest in terms of access to personal assistance, which is expected to vary over the child's age and thereby cause variation in the parental care burden (α_g). More evidence from other institutional and cultural settings would also help clarify how general these patterns are and which policy responses are most effective.

Declarations

Competing interests

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Ethics approval

The study uses anonymized administrative register data accessed under approved research agreements. Ethical approval was granted by the Swedish Ethical Review Authority and the Regional Ethical Review Board in Lund (Dnr: 2018/1000, 2021-00164, and 2022-01150-02).

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Data availability

The data used in this study are confidential administrative register data accessed under restricted conditions for research purposes from Statistics Sweden, the Swedish National Board of Health and Welfare, and the CPUP register. The data cannot be shared publicly. The data can be accessed after standard application procedures from the register holders, subject to ethical approval and data protection regulations.

Generative AI and AI-assisted technologies

The authors used ChatGPT (OpenAI) for limited language editing and to improve phrasing, clarity, and readability. After using this tool, the authors reviewed and edited the text as needed and take full responsibility for the content of the published article.

References

Alriksson-Schmidt, A. I., Jeglinsky, I., Jonsdottir, G., Seid, A. K., Klevberg, G., Buschmann, E., Jahnsen, R., et al. (2021). Living life with cerebral palsy? a description of the social safety nets for individuals with cerebral palsy in the nordic countries. *Scandinavian Journal of Public Health*, 49(6):653–665. Published online December 15, 2020.

Asuman, D., Gerdtham, U. G., Alriksson-Schmidt, A. I., Rodby-Bousquet, E., Andersen, G. L., and Jarl, J. (2023). Pain and labor outcomes: A longitudinal study of adults with cerebral palsy in Sweden. *Disability and Health Journal*, 16(3):101479.

Balbo, N. and Burciu, R.-D. (2026). The double challenges of single parents raising children with disabilities. *European Journal of Population*. Online ahead of print.

Becker, G. S., Landes, E. M., and Michael, R. T. (1977). An economic analysis of marital instability. *Journal of Political Economy*, 85(6):1141–1187.

Bengtsson, S., Hansen, H., and Røgeskov, M. (2011). Børn med en funktionsnedsættelse og deres familier. *SFI – Det Nationale forskningscenter for velfærd*.

Bumpass, L. L., Sweet, J. A., and Castro Martin, T. (1990). Changing patterns of remarriage. *Journal of Marriage and Family*, 52(3):747–756.

Burdett, K. and Coles, M. G. (1999). Long-term partnership formation: Marriage and employment. *The Economic Journal*, 109(456):F307–F334.

Charlson, M. E., Pompei, P., Ales, K. L., and MacKenzie, C. R. (1987). A new method of classifying prognostic comorbidity in longitudinal studies: Development and validation. *Journal of Chronic Diseases*, 40(5):373–383.

Chounti, A., Hägglund, G., Wagner, P., and Westbom, L. (2013). Sex differences in cerebral palsy incidence and functional ability: a total population study. *Acta Paediatrica*, 102(7):712–717.

Cohen, P. N. and Petrescu-Prahova, M. (2006). Gendered living arrangements among children with disabilities. *Journal of Marriage and Family*, 68(3):630–638.

Compagnone, E., Maniglio, J., Camposeo, S., Vespino, T., Losito, L., De Rinaldis, M., Gennaro, L., and Trabacca, A. (2014). Functional classifications for cerebral palsy: Correlations between the gross motor function classification system (GMFCS), the manual ability classification system (MACS) and the communication function classification system (CFCS). *Research in Developmental Disabilities*, 35(11):2651–2657.

de Graaf, P. M. and Kalmijn, M. (2003). Alternative routes in the remarriage market: Competing-risk analyses of union formation after divorce. *Social Forces*, 81(4):1459–1498.

Di Nallo, A., Ivanova, K., and Balbo, N. (2023). Repartnering of women in the United States: The interplay between motherhood and socio-economic status. *Population Studies*.

Doyle, A., Craig, S., and McConkey, R. (2023). Changes over 15 years in lone parenting of Irish persons with intellectual disability. *Journal of Family Studies*, 29(2):841–852.

Dupuy, A. and Galichon, A. (2014). Personality Traits and the Marriage Market. *Journal of Political Economy*, 122(6):1271–1319.

Guner, N., Kulikova, Y., and Llull, J. (2018). Marriage and Health: Selection, Protection, and Assortative Mating. *European Economic Review*, 104:169–190.

Hartley, S. L., Barker, E. T., Seltzer, M. M., Floyd, F., Greenberg, J., Orsmond, G., and Bolt, D. (2010). The relative risk and timing of divorce in families of children with an autism spectrum disorder. *Journal of Family Psychology*, 24(4):449–457.

Hatton, C., Emerson, E., Graham, H., Blacher, J., and Llewellyn, G. (2010). Changes in Family Composition and Marital Status in Families with a Young Child with Cognitive Delay. *Journal of Applied Research in Intellectual Disabilities*, 23(1):14–26.

Himmelmann, K. and Uvebrant, P. (2018). The panorama of cerebral palsy in Sweden part XII shows that patterns changed in the birth years 2007–2010. *Acta Paediatrica*, 107(3):462–468.

Hjern, A. and Thorngren-Jerneck, K. (2008). Perinatal complications and socio-economic differences in cerebral palsy in Sweden - A national cohort study. *BMC Pediatrics*, 8(1):49.

Hollung, S. J., Vik, T., Wiik, R., Bakken, I. J., and Andersen, G. L. (2017). Completeness and correctness of cerebral palsy diagnoses in two health registers: implications for estimating prevalence. *Developmental Medicine & Child Neurology*, 59(4):402–406.

Ivanova, K., Kalmijn, M., and Uunk, W. (2013). The effect of children on men's and women's chances of re-partnering in a European context. *European Journal of Population*, 29(4):417–444.

Jansen, M., Mortelmans, D., and Snoeckx, L. (2009). Repartnering and (re)employment: Strategies to cope with the economic consequences of partnership dissolution. *Journal of Marriage and Family*, 71(5):1271–1293.

Korzeniewski, S. J., Slaughter, J., Lenski, M., Haak, P., and Paneth, N. (2018). The complex aetiology of cerebral palsy. *Nature Reviews Neurology*, 14(9):528–543.

Loft, L. T. G. (2011). Child health and parental relationships: Examining relationship termination among danish parents with and without a child with disabilities or chronic illness. *International Journal of Sociology*, 41(1):27–47.

Lundeby, H. and Tøssebro, J. (2008). Family Structure in Norwegian Families of Children with Disabilities. *Journal of Applied Research in Intellectual Disabilities*, 21(3):246–256.

Mauldon, J. (1992). Children's risks of experiencing divorce and remarriage: Do disabled children destabilize marriages? *Population Studies*, 46(2):349–362.

Michael-Asalu, A., Taylor, G., Campbell, H., Lelea, L. L., and Kirby, R. S. (2019). Cerebral Palsy: Diagnosis, Epidemiology, Genetics, and Clinical Update. *Advances in Pediatrics*, 66:189–208.

Michelsen, S. I., Flachs, E. M., Madsen, M., and Uldall, P. (2015). Parental social consequences of having a child with cerebral palsy in Denmark. *Developmental Medicine & Child Neurology*, 57(8):768–775.

Mortensen, D. T. (1988). Matching: Finding a partner for life or otherwise. *American Journal of Sociology*, 94(S1):S215–S240.

Müller, V., Gerdtham, U., Alriksson-Schmidt, A., and Jarl, J. (2022). Parental decisions to divorce and have additional children among families with children with cerebral palsy: Evidence from Swedish longitudinal and administrative data. *Health Economics*, 31(10):2170–2186.

Namkung, E. H., Song, J., Greenberg, J. S., Mailick, M. R., and Floyd, F. J. (2015). The Relative Risk of Divorce in Parents of Children With Developmental Disabilities:Impacts of Lifelong Parenting. *American Journal on Intellectual and Developmental Disabilities*, 120(6):514–526.

Oskoui, M., Coutinho, F., Dykeman, J., Jetté, N., and Pringsheim, T. (2013). An update on the prevalence of cerebral palsy: a systematic review and meta-analysis. *Developmental Medicine & Child Neurology*, 55(6):509–519.

Palisano, R. J., Rosenbaum, P., Bartlett, D., and Livingston, M. H. (2008). Content validity of the expanded and revised Gross Motor Function Classification System. *Developmental Medicine & Child Neurology*, 50(10):744–750.

Peterson, M. D., Ryan, J. M., Hurvitz, E. A., and Mahmoudi, E. (2015). Chronic Conditions in Adults With Cerebral Palsy. *JAMA*, 314(21):2303–2305.

Pevalin, D. J. and Ermisch, J. (2004). Cohabiting unions, repartnering and mental health. *Psychological Medicine*, 34(8):1553–1559.

Reid, S. M., Meehan, E. M., Arnup, S. J., and Reddihough, D. S. (2018). Intellectual disability in cerebral palsy: a population-based retrospective study. *Developmental Medicine and Child Neurology*, 60(7):687–694.

Risdal, D. and Singer, G. H. (2004). Marital Adjustment in Parents of Children with Disabilities: A Historical Review and Meta-Analysis. *Research and Practice for Persons with Severe Disabilities*, 29(2):95–103.

Rosenbaum, P., Paneth, N., Leviton, A., Goldstein, M., and Bax, M. (2007). A report: the definition and classification of cerebral palsy April 2006. *Developmental Medicine & Child Neurology*, 49(Suppl. 109):8–14.

Schemper, M., Wakounig, S., and Heinze, G. (2009). The estimation of average hazard ratios by weighted Cox regression. *Statistics in Medicine*, 28(19):2473–2489.

Schneider, R. E., Ng, P., Zhang, X., Andersen, J., Buckley, D., Fehlings, D., Kirton, A., Wood, E., van Rensburg, E., Shevell, M. I., and Oskoui, M. (2018). The Association Between Maternal Age and Cerebral Palsy Risk Factors. *Pediatric Neurology*, 82:25–28.

Schnor, C., Pasteels, I., and Van Bavel, J. (2017). Sole Physical Custody and Mother's Repartnering After Divorce. *Journal of Marriage and Family*, 79(3):879–890.

Solaski, M., Majnemer, A., and Oskoui, M. (2014). Contribution of socio-economic status on the prevalence of cerebral palsy: a systematic search and review. *Developmental Medicine & Child Neurology*, 56(11):1043–1051.

Tøssebro, J. and Wendelborg, C. (2017). Marriage, Separation and Beyond: A Longitudinal Study of Families of Children with Intellectual and Developmental Disabilities in a Norwegian Context. *Journal of Applied Research in Intellectual Disabilities*, 30(1):121–132.

Tseng, S. H., Lee, J. Y., Chou, Y. L., Sheu, M. L., and Lee, Y. W. (2018). Association between socioeconomic status and cerebral palsy. *PLOS ONE*, 13(1):e0191724.

Zhou, L., Meng, Q., von Ehrenstein, O. S., Xiao, J., Gao, Y., Wu, Y. W., Ritz, B., and Liew, Z. (2023). Parental Age and Childhood Risk for Cerebral Palsy in California. *The Journal of Pediatrics*, 255:147–153.e6.

Online Appendix A

A1 CP

A1.1 Symptoms and causes

The underlying cause of CP is brain damage that occurs either in utero, during birth, or during the first two years of the child's life (Rosenbaum et al., 2007). While the brain damage by definition is non-progressive, secondary and comorbid conditions such as pain and progressive musculoskeletal complications tend to worsen over time (Peterson et al., 2015). Persons with CP often experience sensation, perception, cognition, and communication impairments (Asuman et al., 2023). The cause of the brain damage varies and is often unknown (Korzeniewski et al., 2018), although multiple factors are often considered to contribute (Himmelmann and Uvebrant, 2018; Korzeniewski et al., 2018). Known risk factors include multiple gestation, low birth weight, low gestational age, and pregnancy and birth complications, although considered largely non-operational in developed countries (Korzeniewski et al., 2018). While some studies have noted an association with socioeconomic factors (Michael-Asalu et al., 2019; Solaski et al., 2014), other studies like Müller et al. (2022) found no association with parental education but rather with parental health status. Many identified risk factors overlap, making it hard to discern their separate contribution (Korzeniewski et al., 2018).

The most common classification of functional ability of children with CP is the Gross Motor Function Classification System (GMFCS), a five-level system based on self-initiated movement abilities (Rosenbaum et al., 2007; Palisano et al., 2008). Although other definitions of severity are possible, GMFCS has been found to correlate strongly with intellectual disability (Reid et al., 2018) and communication and manual ability (Compagnone et al., 2014) and is often used in the disability literature as an overall measure of severity.

A1.2 Identifying the treatment group

The children with CP in the treatment group were identified through the Swedish National Patient Register (SPR), the Cerebral-palsy follow-up program and National Quality Register (CPUP), and the Swedish Medical Birth Register. Hollung et al. (2017) noted that 14% of persons with a CP diagnosis in the Norwegian patient register did not have CP, as incorrect diagnoses are never removed from the register. Therefore, identifying the population based

on the Swedish patient register risks including persons incorrectly assigned a CP diagnosis. For this reason, we exclude those who were not registered with a CP diagnosis after the age of 3 (i.e., likely to be misdiagnosed) and those with a first CP diagnosis after a diagnosis of acquired brain damage (ICD10: G00, S061, S062, S063, S067, S068, S069. ICD9: 348B, 851, 852, 853, 584, 320) if that brain damage occurred after the age of two (by definition, CP cannot be caused by brain damage acquired after the age of two). We also excluded persons with diagnoses considered incompatible with CP (ICD10: G60, G61, G62, G71, G72, G834, G95, E71, E72, E74, E75, E76, E830, G114, G12, G31, G37, Q06, Q743. ICD9: 337, 356, 359, 358, 344G, 330, 270D, 270E, 270G, 270H, 272H, 277F, 275B, 334A, 334B, 335, 336, 728D, 742F), and persons written off in the CPUP register.

A2 Tables and figures

Table A1: Schoenfeld tests of Equation 13

	Mothers		**Fathers**	
	p	p	p	p
CP	0.08		0.03	
Mild CP		0.07		0.04
Severe CP		0.78		0.97
Child: female	0.67	0.52	0.98	0.89
Parent immigrated	0.51	0.46	0.24	0.25
Child 2-6 years when single	0.38	0.51	0.44	0.39
Child 7-10 years when single	<0.01	<0.01	<0.01	<0.01
Child 11+ years when single	<0.01	<0.01	<0.01	<0.01
Parent's education: Secondary	0.43	0.42	0.32	0.31
Parent's education: Higher	0.01	0.01	<0.01	<0.01
Age when single: Mothers 30-36 yo, fathers 33-40 yo	<0.01	<0.01	<0.01	<0.01
Age when single: Mothers 37 yo+, fathers 41 yo+	<0.01	<0.01	<0.01	<0.01
Parent's income: 2nd quintile	0.31	0.29	0.43	0.34
Parent's income: 3rd quintile	0.51	0.50	0.57	0.79
Parent's income: 4th quintile	0.13	0.16	0.06	0.06
Parent's income: 5th quintile	<0.01	0.01	<0.01	<0.01
Region of birth: Svealand	0.79	0.68	0.30	0.25
Region of birth: Norrland	0.27	0.29	0.68	0.64
Stockholm	0.57	0.58	0.41	0.35
Child born 1993-4	<0.01	<0.01	0.20	0.25
Child born 1995-6	0.34	0.38	<0.01	<0.01
Child born 1997-8	0.16	0.29	0.65	0.62
Child born 1999-2000	0.96	0.97	0.57	0.57
Child born 2001-2	0.25	0.26	0.11	0.17
Child born 2003-4	0.24	0.22	0.66	0.66
Child born 2005-6	<0.01	<0.01	0.08	0.09
Child born 2007-8	<0.01	<0.01	<0.01	<0.01
Child born 2009-10	0.32	0.31	0.01	0.01
Child born 2011-12	0.36	0.35	0.04	0.05
Child born 2013-14	0.25	0.26	0.17	0.17
In/around medium-sized towns	0.63	0.67	0.87	0.89
Rural	0.02	0.01	0.41	0.52
Global	<0.01	<0.01	<0.01	<0.01

Notes: P-values of Schoenfeld tests, testing the null hypothesis that the Schoenfeld residuals are independent of time. P-values lower than 0.05 suggest time dependence of residuals and thus violation of the proportional hazard assumption.

Table A2: Weighted average hazard ratios for repartnering, all covariates included.

	Mothers	**Fathers**
CP	0.77***	0.97
Child: female	1.05	1.04
Parent immigrated	0.63***	0.90
Child 2-6 years when single	1.62***	1.27***
Child 7-10 years when single	1.36***	0.66***
Child 11+ years when single	0.96	0.33***
Parent's education: Secondary	0.91	1.06
Parent's education: Higher	0.85*	1.13
Age when single: Mothers 30-36 yo, fathers 33-40 yo	0.44***	0.66***
Age when single: Mothers 37 yo+, fathers 41 yo+	0.16***	0.33***
Parent's income: 2nd quintile	1.03	0.97
Parent's income: 3rd quintile	1.02	1.08
Parent's income: 4th quintile	0.91	1.13
Parent's income: 5th quintile	0.89	1.17
Region of birth: Svealand	1.14*	1.05
Region of birth: Norrland	1.11	1.08
Stockholm	0.76**	0.90
Child born 2003-4	0.91	0.79**
Child born 2005-6	0.91	0.78*
Child born 2007-8	0.75*	0.62***
Child born 2009-10	0.63**	0.42***
Child born 2011-12	0.43***	0.45***
Child born 2013-14	0.20	0.52
Child born 1991-2	1.13	1.08
Child born 1993-4	1.07	1.00
Child born 1995-6	1.07	0.98
Child born 1997-8	1.05	1.08
Child born 1999-2000	1.21*	0.96
In/around medium-sized towns	1.05	1.13
Rural	1.15*	1.06
Number of events	1533	1564
Number of obs	6143	5883

Notes: Results are expressed as weighted average hazard ratios. *p<0.1, **p<0.05, ***p<0.01.

Table A3: CP status regressed on mothers' and fathers' background variables

	Mothers	**Fathers**
Child: female	0.00	0.00
Mother 30-36/Father 33-40 yo when single	0.03***	0.02*
Mother 37+/Father 41+ yo when single	0.07***	0.04**
Child 2-6 yo when single	-0.05***	-0.05***
Child 7-10 yo when single	-0.08***	-0.07***
Child 11+ yo when single	-0.09***	-0.07***
Parent immigrated	-0.01	-0.01
Parent's education: Secondary	0.00	0.00
Parent's education: Higher	-0.01	-0.01
Parent's income: 2nd quintile	-0.01	-0.02
Parent's income: 3rd quintile	-0.01	-0.01
Parent's income: 4th quintile	-0.01	0.01
Parent's income: 5th quintile	-0.02	-0.01
Region of birth: Svealand	-0.01	-0.01
Region of birth: Norrland	0.01	0.02
Region of birth: Stockholm	-0.01	0.01
Birthplace: In/around medium-sized towns	0.00	0.01
Birthplace: Rural	0.01	0.02
Child born 2003-4	0.01	0.00
Child born 2005-6	-0.02	-0.03
Child born 2007-8	0.01	0.00
Child born 2009-10	-0.02	-0.04
Child born 2011-12	-0.01	-0.02
Child born 2013-14	0.01	-0.01
Child born 1991-2	0.00	-0.01
Child born 1993-4	0.00	-0.01
Child born 1995-6	0.01	-0.01
Child born 1997-8	0.00	0.01
Child born 1999-2000	0.00	0.00
Number of observations	6143	5883
Adj. R-squared	0.003	0.002
F-test (p-value)	0.025	0.065

Notes: Results of OLS estimation. *p<0.1, **p<0.05, ***p<0.01.

Table A4: Summary statistics of the full sample, including parents of children born in multiple births and children with older siblings.

	No CP	**CP**	**Mild CP**	**Severe CP**
Child: female	42.62	42.99	42.06	42.06
Mother immigrated	16.07	16.56	16.81	16.30
Father's income: 1st quintile	22.46	23.64	23.66	23.34
Father's income: 2nd quintile	23.75	22.29	21.32	22.74
Father's income: 3rd quintile	21.28	21.25	21.91	20.93
Father's income: 4th quintile	18.96	18.95	19.65	17.51
Father's income: 5th quintile	13.55	13.87	13.46	15.49
Child < 1 year when mother single	32.54	41.30	40.47	43.86
Child 2-6 years when mother single	35.54	31.42	33.53	30.58
Child 7-10 years when mother single	15.70	12.57	11.37	13.68
Child > 11 years when mother single	16.23	14.71	14.63	11.87
Region of birth: Götaland	44.95	46.88	48.83	44.47
Region of birth: Svealand	39.24	37.11	35.87	37.63
Region of birth: Norrland	15.82	16.01	15.30	17.91
Mothers education: Mandatory	22.25	24.59	24.50	23.34
Mothers education: Secondary	53.64	52.47	50.92	54.93
Mothers education: Higher	24.11	22.94	24.58	21.73
Mother single at 29 years or younger	33.76	35.36	33.95	39.24
Mother single at 30-36 years	33.50	33.47	34.20	32.39
Mother single at 37 years or older	32.74	31.17	31.86	28.37
Birthplace: In or around a large city	35.27	34.51	33.44	34.61
Birthplace: In or around a medium size town	40.50	41.90	42.56	41.25
Birthplace: In or around small towns, rural	24.22	23.59	24.00	24.14
Father immigrated	18.43	18.39	17.95	18.43
Mother's income: 1st quintile	28.04	31.04	30.20	31.78
Mother's income: 2nd quintile	24.55	22.57	22.09	24.58
Mother's income: 3rd quintile	20.11	19.16	19.15	16.53
Mother's income: 4th quintile	16.23	15.75	15.53	16.53
Mother's income: 5th quintile	11.07	11.47	13.03	10.59
Child < 1 year when father single	29.88	38.22	37.36	40.89
Child 2-6 years when father single	36.97	32.80	35.20	31.36
Child 7-10 years when father single	16.28	13.48	12.17	15.04
Child > 11 years when father single	16.88	15.50	15.27	12.71
Fathers education: Mandatory	21.87	23.40	22.95	24.36
Fathers education: Secondary	56.48	56.15	55.05	56.14
Fathers education: Higher	21.65	20.45	22.00	19.49
Father single at 32 years or younger	34.31	36.11	34.94	39.41
Father single at 33-40 years	35.29	33.78	35.63	31.99
Father single at 41 years and older	30.40	30.11	29.42	28.60
Region of birth: Götaland	45.10	46.90	48.75	44.92
Region of birth: Svealand	38.81	36.47	35.38	35.38
Region of birth: Norrland	16.09	16.63	15.88	19.70
Birthplace: In or around a large city	34.91	33.68	32.70	32.20
Birthplace: In or around a medium size town	40.43	42.41	43.40	41.95
Birthplace: In or around small towns, rural	24.66	23.92	23.90	25.85
Observations - children	12,113	2,091	1,241	520
Observations - mothers	11,704	2,007	1,197	497
Observations - fathers	10,890	1,839	1,102	448

Notes: Parents' age groups and income groups are defined as terciles and quintiles for mothers/fathers in the control group in the whole sample (including non-single parents).

Table A5: Marriage market summary statistics, full sample

	No CP	**CP**	**Mild CP**	**Severe CP**
Mothers who are single at some point	11703	2005	1196	497
	(35.3)	(36.2)	(36.2)	(35.2)
Non-single mothers	21495	3537	2109	915
	(64.7)	(63.8)	(63.8)	(64.8)
If single:				
Mothers who repartner†	2685	381	210	100
	(22.9)	(19.0)	(17.6)	(20.1)
Average number of years until single‡	5.0	4.4	4.4	3.9
	(5.1)	(5.1)	(5.1)	(4.8)
If repartnered:				
Average number of years until repartnered§	5.5	5.9	5.8	6.1
	(3.6)	(3.9)	(4.0)	(3.6)
Fathers who are single at some point	11389	1936	1159	472
	(34.4)	(35.1)	(35.5)	(33.6)
Non-single fathers	21694	3569	2110	934
	(65.6)	(64.8)	(64.5)	(66.4)
If single:				
Fathers who repartner†	2828	449	257	110
	(24.8)	(23.2)	(22.2)	(23.3)
Average number of years until single‡	5.2	4.6	4.6	4.2
	(5.1)	(5.2)	(5.1)	(4.8)
If repartnered:				
Average number of years until repartnered§	5.6	5.6	5.5	5.6
	(3.9)	(4.2)	(4.3)	(3.2)

Notes: Standard deviation or percentages in parentheses. †Percentage of single mothers/fathers who repartner. ‡Average number of years between the birth of the child until becoming a single parent. §Average number of years from becoming a single parent until repartnering occurs, conditional on repartnering.

Table A6: Weighted average hazard ratios of repartnering, full sample

	Mothers			**Fathers**		
	(1)	(2)	(3)	(4)	(5)	(6)
CP	0.834***	0.830***		0.937	0.948	
	(0.054)	(0.053)		(0.052)	(0.051)	
Mild CP			0.794***			0.958
			(0.070)			(0.064)
Severe CP			0.841*			0.872
			(0.097)			(0.100)
Backgr. var.		✓	✓		✓	✓
Observations	13,708	13,708	13,396	13,325	13,325	13,020

Notes: Results are expressed as weighted average hazard ratios. Standard errors in parentheses. *p<0.1, **p<0.05, ***p<0.01. All parents who are single at some point are included in the estimation. Background variables included as covariates are: dummies for first born children and multiple births, child's gender, mother's/father's education, mother's/father's income at birth, immigration status of the mother/father, the child's year of birth (in clusters of two years), information on the place of birth (region and the degree of urbanization of the birth municipality.). Background variables stratified over are: mother's/father's age and child age when single spell begins.

Table A7: Summary statistics of health variables available 2004-2015

	No CP	**CP**	**Mild CP**	**Severe CP**
	Percent	Percent	Percent	Percent
Mother smoked	30.09	35.58	36.02	31.03
CCI mother	3.80	4.31	4.30	3.45
Mother: mental health diagnosis	9.33	13.48	14.52	12.93
CCI father	3.70	3.70	3.19	3.42
Father: mental health diagnosis	6.50	9.07	7.98	8.55
Observations - children	1,677	315	198	86
Observations - mothers	1,446	251	164	67
Observations - fathers	1,431	255	170	69

Notes: The final sample consists of children with no missing information on the background variables and the health variables. Mother smoked = mother smoked in the 3 months prior to the pregnancy. CCI = parent's CCI index is at least 1. Mental health diagnosis = parent had at least one mental health diagnosis in the three calendar years prior to childbirth.

Table A8: Weighted average hazard ratios estimated with and without parental health controls. Estimation period: 2004-2015

	Mothers		**Fathers**	
	(1)	(2)	(3)	(4)
CP	0.702*	0.701*	0.823	0.814
	(0.196)	(0.196)	(0.168)	(0.168)
Background variables	✓	✓	✓	✓
Health covariates		✓		✓
Observations	1,696	1,696	1,762	1,762

Notes: Results are expressed as hazard ratios. Standard errors in parentheses. *p<0.1, **p<0.05, ***p<0.01. The model is estimated over the period 2004-2015. Background variables included as covariates are: child's sex, mother's/father's education, mother's/father's income at birth, immigration status of the mother/father, the child's year of birth (in clusters of two years), information on the place of birth (region and the degree of urbanization of the birth municipality). Background variables stratified over are: mother's/father's and child's age when single spell begins. Parental health controls are (included as covariates): a dummy for smoking status of mother before pregnancy, mental health diagnoses of mothers/fathers before birth and mothers'/fathers' CCI index before birth.

Table A9: Hazard rates for having an additional child in a new marriage.

	Mothers	**Fathers**
CP	1.117	1.367
	(0.270)	(0.240)
Background variables	✓	✓
Observations	263	348

Notes: Results are expressed as hazard ratios. Standard errors in parentheses. *p<0.1, **p<0.05, ***p<0.01. The model is only estimated for parents who have remarried before having an additional child in the new relationship. Background variables stratified over are: mother's/father's and child's age when mother/father remarries, the child's year of birth (in clusters of two years), child's sex. Mother's/father's immigration status, mother's/father's education, income and characteristics of the place of residence when the older (focal) child was born are included as a covariates.

Table A10: Weighted average hazard ratios of repartnering by type of repartnering

	Remarriage		**Repartner by child**	
	Mothers	Fathers	Mothers	Fathers
CP	0.896	0.915	0.641***	1.090
	(0.120)	(0.097)	(0.110)	(0.106)
Background variables	✓	✓	✓	✓
Observations	5,269	5,160	5,441	5,003

Notes: Results are expressed as weighted average hazard ratios. Standard errors in parentheses. *p<0.1, **p<0.05, ***p<0.01. Background variables included as covariates are: child's sex, mother's/father's education, mother's/father's income at birth, immigration status of the mother/father, the child's year of birth (in clusters of two years), information on the place of birth (region and the degree of urbanization of the birth municipality.). Background variables stratified over are: mother's/father's and child's age at the start of single spell.