# PHYSICAL CUSTODY ARRANGEMENT AND REPARTNERING:

# EVIDENCE FROM A POLICY PROMOTING JOINT CUSTODY

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ABSTRACT

Women who live after separation with their children exhibit lower repartnering rates than women without residential children. This study considered that the children's residential arrangement may be endogenously related to the repartnering intentions of the mother. Using the Divorce in Flanders study (N=1,222), the authors exploited exogenous variation in physical custody arrangement induced by a policy reform promoting joint legal custody as an instrumental variable. They found that mother-sole physical custody was significantly related to lower repartnering probabilities. This effect became stronger when allowing custody choice to be endogenous, suggesting that ordinary models underestimate the negative effect of sole physical custody on mothers' repartnering because they fail to consider endogeneity. Having the children 90 to 100 percent of the time living in the household reduces the probability to re-partner by 75 percent. The authors conclude that sole physical custody acts as an important impediment to step-family formation following divorce.

High divorce rates imply that a substantial proportion of the population experiences marriage not as a life-long institution but as a chapter in family life often followed by step family arrangements. In response to the ongoing trend of increasing divorce rates, a growing body of research has been dedicated to the nature and the dynamics of postdivorce family life. Within this line of research, there is consistent evidence that having the children from the previous union permanently in the household – in so-called sole physical custody – substantially decreases one's likelihood to form a new coresidential partnership (de Graaf & Kalmijn, 2003; Ivanova et al., 2013; Beaujouan, 2010, 2012). Lacking opportunities to meet a potential spouse are seen as the main reason behind this finding (de Graaf & Kalmijn, 2003).

So far, studies have treated the child custody arrangement after separation as exogenous to repartnering. This leaves it unclear to what extent the correlation between child custody and repartnering is due to the causal effect of the former on the latter, as unmeasured confounders could have biased the estimations. Unmeasured characteristics of mothers can lead them to be in sole physical custody and to register lower or even higher values on repartnering (DeMaris, 2014; Wooldrigde, 2003). Often unmeasured are for example attitudes – such as being family oriented. Women who highly value family life may prefer to be the primary caretakers, dedicating their time to childrearing while consciously setting aside romantic love. On the contrary, a high family orientation among mothers in sole custody may motivate them to repartner in order to restore the image of a complete family. If unmeasured characteristics make mothers more likely to opt for sole custody and less likely to repartner, the negative effect of sole custody would have been overstated in prior findings, whereas the effect of custody arrangement is likely to have been understated if unmeasured characteristics lead mothers to be in sole physical custody and register higher values on repartnering. Next to the problem of omitted variables, endogeneity can also be caused by simultaneity or reverse causality, which would be the case when people who have a new partner after separation are less inclined to take sole custody. Advanced statistical modeling can help to address endogeneity problems and to eliminate unmeasured confounds. In this article, we extend the literature on children's residential arrangement and repartnering by accounting for the potential endogeneity of the physical custody arrangement. We do this by making use of a legal policy change in custody arrangements, which we consider as an exogenous shock to individual custody decisions.

For a long time, the default arrangement after separation was that the mother received sole legal as well as sole physical custody, while the father was acquainted visitation rights (Bauserman, 2002; Bender, 1994). In recent decades, there has been a movement towards equal parental rights in legal custody decisions (Bauserman, 2002; Cancian & Meyer, 1998; Perelli-Harris & Sanchez-Gassen, 2013; Sodermans et al., 2013). Plenty of countries have adopted a legal presumption for joint legal custody, assigning both parents equal rights after a marital separation (Perelli-Harris & Sanchez-Gassen, 2013; Skinner & Davidson, 2009). Such legislation reforms have changed perceptions of the parental role after divorce and have promoted diversity in physical custody arrangements. Although mother-sole custody remains the dominant post-divorce arrangement (Cancian & Meyer, 1998; Cancian et al. 2014), shared physical custody has become more common (Juby, Marcil-Gratton, & Le Bourdais 2005; Juby, Le Bourdais, & Marcil-Gratton 2005; Skinner & Davidson, 2009).

Investigating repartnering patterns among separated mothers is important, because the transition to a new partnership can limit the negative consequences of marital dissolution and single parenthood (Jansen, Mortelmans, & Snoeckx, 2009; Kreyenfeld & Martin, 2011). Repartnering is the key for mothers' greater well-being after divorce (Symoens et al, 2013; Wang & Amato, 2000). Further, pooling the financial resources with a new partner can work as a way to increase the family's income (Amato & Maynard, 2007).The financial situation of the family often worsens after separation with especially lone mothers having a high risk of poverty (Jansen, Mortelmans, & Snoeckx, 2009; Seccombe, 2002; Smock, 1993). Among countries of the European Union, on average half (50 percent) of the single persons with dependent children was in 2013 at the risk of being poor, severely materially deprived or living in a household with low working intensity, compared to 23 percent of the couples with dependent children (Eurostat, 2014). Poverty risks are similarly accumulated among US American and Canadian single-parent families (Statistics Canada, 2012; Brown, 2010; Manning & Brown, 2006). Poverty can threaten the physical and mental health of the family and reduce the children's educational outcomes (Seccombe, 2002). Understanding the circumstances that causally affect the formation of step-families is an essential prerequisite for designing policies that aim to support families in their post-divorce life and to reduce social inequality within the population. More concretely, for formulating effective policies it is crucial to know whether the custody arrangement itself constitutes an impediment for step-family formation rather than that a correlation arises for other reasons.

Using detailed data on divorced persons from the Divorce in Flanders (DiF) study gathered in 2009, we concentrate in this study on the mother's perspective to compare the "traditional" and most common arrangement of sole physical custody to arrangements in which the mother is less involved in childrearing. This means that, although we recognize the importance of analyzing repartnering from a gender perspective, we do not concentrate on father's repartnering.

# BACKGROUND

## *Repartnering Chances – A Combination Of Needs, Attractiveness, And Opportunities*

Based on previous work by Becker (1991), Goldscheider and Waite (1986) and Oppenheimer (1988), de Graaf and Kalmijn (2003) developed a theoretical framework that aims to explain repartnering patterns with three factors: needs, attractiveness, and opportunities. These aspects also have to be considered to understand how childrearing obligations determine post-divorce repartnering chances of women with children (Ivanova et al. 2013; Jansen, Mortelmans, & Snoeckx, 2009; Turunen, 2011; Theunis et al. 2015). First, women can seek a new relationship because they have a practical, financial, emotional or social *need* for a partner. A separation, creating time constraints, makes the reconciliation of work and family life more difficult. Single parents benefit from a new partnership because a second adult living in the household can take over some of the household and child care tasks and contribute to the household income (de Graaf & Kalmijn, 2003; Turunen, 2011). Women in sole physical custody arrangements might have a higher financial need than women in alternative arrangements: fathers are less likely to pay alimonies if they are not involved in childrearing duties, turning the mothers not only into primary caretakers but also into sole providers (Juby, Le Bourdais, Marcil-Gratton, 2005). At the same time, the potential working hours (and ceteris paribus the earning capacities) of sole physical custody mothers are more restricted (Turunen, 2011). Apart from that, a new partnership also addresses the mother's emotional needs. Mothers may want to share the emotions of the day-to-day family life and have a partner on their side at social occasions. Research points out that mothers with full-time residential children might feel less lonely and less in need for a new residential partner than women in alternative living arrangements (Lampard & Peggs, 1999; Vanassche et al., forthcoming). The former therefore might prefer to stay alone or in a non-residential partnership, also in order to avoid conflicts that a residential step-family configuration might provoke (Lampard & Peggs, 1999; Beaujouan, 2010).

Second, repartnering prospects depend on how *attractive* the woman is as a potential partner. Children from a prior relationship may make the woman unattractive to a new partner, especially if they are always around. Dating a mother in sole physical custody implies the entry into step-parenthood with the entry into partnership, which might keep some potential partners away (de Graaf & Kalmijn, 2003; Ivanova, Kalmijn, & Uunk, 2012; Stewart, Manning, & Smock, 2003). Children can also be seen by the new partner as resource drain, because they reduce the availability of the woman (Vanassche et al., forthcoming; Stewart, Manning & Smock, 2003).

Third, the probability of repartnering depends on the woman's *opportunities* to meet someone (de Graaf & Kalmijn, 2003; Theunis et al. 2015). Full-time childrearing obligations limit the woman's opportunities to search on the partner market: mothers with resident children are likely to have fewer time and energy that they can spend on leisure activities and on socializing with potential mates (Ivanova et al, 2013; Koo et al. 1984; Turunen, 2011).

The determinant aspects of need, attractiveness and opportunity, developed in de Graaf and Kalimijn (2003), that relate custody arrangement to repartnering are summarized in Table 1.

Table 1: *Sole physical custody and the determinants of repartnering*

| Determinants of repartnering (de Graaf & Kalmijn, 2003) | Sole-custody mother's likelihood to repartner (relative to mothers in alternative arrangements) |
|---|---|
| Need | |
| Practical need | + |
| Financial need | + |
| Emotional need | - |
| Attractiveness | - |
| Opportunity | |
| Time | - |
| Occasions | - |

*Note*: "+" higher repartnering rate; "-" lower repartnering rate; own illustration

The empirical literature about repartnering has primarily concentrated on the role of custody obligations in explaining gender-specific differences in post-divorce partnership formation (Beaujouan, 2012; Bernhardt & Goldscheider, 2002; de Graaf & Kalmijn, 2003; Ivanova et al. 2013). The empirical findings in these studies consistently confirm that the presence of children from a previous relationship reduces the woman's (and the man's) likelihood to form a new cohabiting union (Beaujouan, 2010, 2012; Bernhardt, 2000; Bernhardt & Goldscheider, 2002; Goldscheider & Sassler, 2006; de Graaf & Kalmijn, 2003; Ivanova et al., 2013; Wu & Schimmele, 2005). Several studies further showed that a separated mother who has sole physical custody over her children is less likely to form a step-family than mothers who share childrearing with their former partners (de Graaf & Kalmijn, 2003, Ivanova et al., 2013). Women with non-residential children have similar repartnering chances than their childless peers (Beaujouan, 2010, 2012; de Graaf & Kalmijn, 2003). Beaujouan (2012) concluded from this finding that the custody arrangement is more crucial for repartnering than parenthood itself. De Graaf and Kalmijn (2003) found in their empirical analysis stronger support for the opportunity argument than for the attractiveness and needs hypotheses. Referring to a sample of Dutch women and men with and without children, the authors showed that economic theories of marriage received in general little support. Women with poor socioeconomic prospects – and thus, with great financial need – do not have higher repartnering rates than others. De Graaf and Kalmijn argued that a high degree of economic dependency may make women less attractive, with need and attractiveness working in opposite directions. According to the authors, lacking opportunities to meet and mate at work and in leisure contexts act as the main impeding factor for repartnering among women with residential children.

### *Selection Into Sole Physical Custody And Repartnering*

In prior repartnering models, custody choice was treated as exogenous. Studies did not control for selection processes that make it more or less likely that a woman is in sole physical custody. But, it is important to note that the custody arrangement is potentially related to latent attributes of the mother and further, repartnering may simultaneously influence custody decisions (de Graaf & Kalmijn, 2003). It is not uncommon to have started a new romantic partner before or at the time of marriage dissolution (Juby et al. 2005), also in our study population (Pasteels et al. 2012). Women who were already engaged in a new romantic relationship towards the end of the dissolved marriage are less inclined to obtain sole physical custody, potentially because their preferences changed with the new partner and they want to spend some child-free time with him (Beaujouan, 2010; Juby et al. 2005; Nielsen 2011).

Mothers in sole physical custody and mothers in alternative arrangements might differ in their unobserved background characteristics, while these same background characteristics may also affect the outcome of the repartnering process (Wunsch, 2007; Leridon & Toulemon, 1997). In consequence, unobserved heterogeneity between mothers in different custody arrangements and within the group of main caretakers regarding their needs, attractiveness, and opportunities, as well as the reverse effect of repartnering on custody choice might have biased the previous empirical results. For example, as mentioned beforehand, women's attitudes – although often not observed – may influence both the custody arrangement and the transition to a higher order union. Women with traditional family attitudes may prefer to be the primary caretakers in continuation of their specialization in home production during first marriage (Cancian & Meyer, 1998; Meggiolario & Ongaro, 2008). Repartnering may be a solution for these women to restore the image of a complete family (Meggiolario & Ongaro, 2008). On the other hand, they may refrain from forming a step-family because it does not conform to the image of the traditional family. They may also be disappointed in their partnership expectations and prefer to distance from men (McNamee et al., 2014). Women with more liberal attitudes may be less prone to permanently live with children and have fewer difficulties to find a new partner (Meggiolario & Ongaro, 2008). On the other hand, autonomous women, avoiding further commitments, may prefer to stay partnerless (de Graaf & Kalmijn, 2003), because they want to focus on their own needs and to take time to develop their social and human capital (McNamee et al., 2014). In consequence, the father's involvement in childcare makes the life outside a committed partnership more attractive, reducing the mother's willingness to repartner.

Next to attitudes, attractiveness of the mother is another often unmeasured characteristic that can drive custody decisions and repartnering. It is possible that mothers who are less attractive anticipate their lower chances to find a partner which is reflected in their custody decision. They might opt for sole physical custody, because they prefer to spend time with their children instead of being alone. Alternatively, one can also think that these mothers choose less often sole custody to increase their opportunities to meet a potential partner.

Previous research has found observed characteristics of the mother, such as her age at marital dissolution, her highest level of educational attainment, but also children's characteristics, such as their gender and age, to be related with the custody choice and repartnering (see Table 2). Some of these findings suggest that mothers with less need for a new partner or mothers who are less attractive might select themselves into sole-custody arrangements. For example, mothers who separate at older ages are less likely to share physical custody with the children's father and to repartner. Older mothers might have lower repartnering rates than their younger peers, because a) they have less a need to be partnered: the financial situation may be better in older ages (Le Bourdais et al., 1995) and there might be less a desire for further children (Beaujouan, 2012); b) they may be perceived as less attractive: the mechanisms of the partner market disadvantage older age in women (Beaujouan, 2012; Bumpass et al., 1990: 751; England & McClintock, 2009); c) mating opportunities are worse: there are less potential spouses available in older age groups (Ní Bhrolcháin & Sigle-Rushton, 2005; Beaujouan, 2012). Consequently, older mothers might opt for sole-custody arrangements, because they place more weight on raising their children than on a new partnership (Beaujouan, 2010; Lampard & Peggs, 1999). Furthermore, they anticipate the general mating difficulties within their age group, which make a successful partner search less likely, and render the costs of partner search too high to bear (Becker et al., 1977: pp.1155).

Educational attainment can drive negative selection, because it links sole custody to repartnering. Lower educated mothers are assumed to have fewer capacities to bargain custody arrangements and to be more open to sole custody because they are less work oriented than their higher educated peers (Juby, Le Bourdais, & Marcil-Gratton, 2005; Soderman et al. 2013). On the partner market, low educational attainments are expected to decrease a person's attractiveness (Becker 1973: 825; Oppenheimer, 1988). Some empirical studies however found no effect of educational background on women's repartnering chances (see Table 2). That might be because the effects of needs and attractiveness are counteracting or because education is influencing the repartnering chances of mothers mainly through their custody choices.

The age of the youngest child at the time of separation has been found to have a nonlinear effect (Juby et al., 2005). When children are young, mothers are more often the sole or main childcarer after separation (Cancian & Meyer, 1998; Juby et al. 2005; Soderman et al. 2013). Young children are assumed to represent an impediment for the mother's repartnering because they restrict the time available for partner search (Ivanova et al. 2013). Turunen (2011) pointed out, however, that young children may facilitate the role of the step-parent for the new partner, making women with young children more attractive on the partner market compared to women with older children. This argument was supported by his empirical findings; but, as he stated, the age of the child is correlated with the age of the parents (see also Beaujouan, 2010). Researchers who were able to control for both the age of the youngest child and the age of the mother (e.g. in Ivanova et al., 2013; Le Bourdais et al. 1995; Meggiolario & Ongaro, 2008) found more support for the first argument, showing that the mother's likelihood to enter a new union increases with children's age, net of own age. This suggests that sole-custody mothers, who more often have children in secondary school age than mothers who are not the main care takers, can be positively selected on repartnering.

Some studies suggest selection based on the child's gender. According to Juby and colleagues (2005), fathers are more involved in childrearing activities with sons; consequently, sole-mother custody is less likely if the child is a boy (Cancian & Meyer, 1998; Fox & Kelly, 1995). Mothers of girls have been assumed to be more attractive for potential spouses, because the latter can expect less potential father role conflicts (Lundberg & Rose, 2003; Turunen, 2011).

Alternatively, it can be expected that new partners favor mothers with sons, because they more easily share interests and activities with a child of the same sex. In empirical studies that investigated the effect of child's gender on their mother's likelihood of repartnering support for both arguments can be found (Lundberg & Rose, 2003; Turunen, 2011).

Finally, it seems relevant for custody decisions and repartnering who of the spouses initiated marital separation. The partner who decided to separate is more likely to start a new relationship, potentially because divorce initiators perceive favorable repartnering prospects (Sweeney, 2002). Cancian and colleagues (Cancian & Meyer, 1998; Cancian et al., 2014) argue that the parent who initiated separation should be more likely to receive the desired custody arrangement. The authors found with Court Record Data for Wisconsin that mothers who initiated the divorce are more likely to receive sole physical custody. Sodermans and colleagues (2013) found that a mutual decision to break up leads more often in joint physical custody compared to a situation in which the end of the marriage was iniated by one of the partners alone.

## *Hypotheses*

All in all, the findings from previous research lead us to suggest that the link between custody arrangement and successful repartnering may be biased by the selection into and the endogeneity of custody choice. But, the direction of the effects and the underlying mechanisms are less clear. This leads us to formulate two opposing hypotheses. On the one hand, the negative effect of sole physical custody on repartnering may be mainly due to reverse causality and negative selection processes that make it more likely that a woman with poor repartnering prospects has sole physical custody. If this is the case, the low repartnering probabilities of a woman with full-time residential children are likely to be overstated in ordinary models that do not account for selection into sole custody (*Hypothesis 1*). On the other hand, sole-custody mothers may have certain characteristics that increase their propensity to find a partner. If this effect prevails, the low repartnering probabilities of a woman with full-time residential children are likely to be understated in ordinary models that do not account for the endogeneity of custody choice (*Hypothesis 2*).

## DATA AND METHOD

### *Country setting*

In this study, we analyzed the custody arrangement and repartnering of Dutch-speaking Belgian (=Flemish) mothers, because this context – with its high divorce rates, an early implementation of joint legal custody and its small area size - seemed to us particularly suitable to answer our research question. Divorces and divorces including children are very common in Belgium. Belgian divorce rates are among the highest of wealthy countries (OECD, 2014). The Belgian period total divorce rate of 2012 implies that 53.3 percent of the marriages would end in divorce if the rates of that year remain unchanged (Statistics Belgium, 2014). Two out of three divorces involve minor children (Corijn, 2005; Statistics Belgium, 2014). The country has pioneered legal changes that promote joint physical custody, and these have effectively increased the likelihood that children live with both parents following separation (Sodermans et al., 2011). From 1995[1] onwards, parents were by default given shared legal custody after divorce, i.e., unless otherwise decided for exceptional reasons, both parents shared full parental authority jointly (Sodermans et al., 2011). With this reform, both biological parents have been targeted as ultimately responsible for the rearing of their children in terms of housing, living costs, parenting and education. Joint legal custody implies active involvement and shared decision-making in child-related matters. Nonetheless, until 2006, legislation did not set forth a default residential model (Sodermans et al., 2013).[2] Belgium allows focusing on a small geographical territory, which increases the link between legal custody and physical custody. The country is clearly divided in different quite homogenous language communities. Flanders, the northern part of the country, is Dutch speaking and covers a region of 13,552 km$^2$ with around 6 million inhabitants. The maximum distance

[1] Joint legal custody was incorporated into Belgian custody law on April 13, 1995.
[2] For recent divorces, the incidence of joint physical custody arrangements exceeded 30 percent (Sodermans et al., 2013).

from North to South is approximately100 km and from West to East 200 km, which restricts the commuting distance between the household of the mother and the father after separation to maximal two driving hours. This encourages joint physical custody, or at least renders it more feasible, and makes Flanders a particularly suitable region to study the effect of household arrangements on post-divorce repartnering.

## *Data*

The 'Divorce in Flanders' data were collected in 2009 to 2010 among first marriages of the 1971 to 2008 cohorts with an oversampling of divorced marriages (two thirds dissolved before the interview date and one third was still intact) (Mortelmans et al., 2011). The sample was selected from the Population register proportional to the marriage formation year. Partners had to be of different sex, both in their first marriage, younger than age 40 at the time of the marriage, living in Flanders both at the time of the marriage and of sampling and had to have the Belgian nationality from birth on. Only people who divorced once were allowed in the sample. From these reference marriages, both partners were approached to be interviewed.[3] The total sample, net of unit nonresponse, amounts to 6,365 persons of 4999 marriages. The year of marriage and the year of divorce that were provided during the interview was corrected in case of inconsistency using the information from the Civil Register. If the partners from the reference marriage had common children, randomly one of the children was selected as the reference child during the interview with the first partner of the reference marriage. The partners received some questions on that child during their interviews, e.g. on the time the child usually spends in their households (=physical custody). If the child alternately lived with both parents, the residential arrangement

[3] The response rates on individual level were very similar for partners from intact and non-intact marriages. The proportion of intact marriages from which both partners participated was higher than the proportion of respective non-intact marriages (34 percent versus 21 percent) (Vanassche, 2013: 25).

was recorded with the help of a calendar corresponding to a regular month without holiday periods (Sodermans et al., 2013). The parents were asked where the child lived immediately after the residential separation and whether the arrangement changed afterwards. The information applied only to the target child, but Sodermans et al. (2013) estimated that only about 6.5 percent of the families used different arrangements for their children. Focusing on the target child is therefore unlikely to bias the results.

The analytical sample consisted of mothers who experienced the dissolution of their first marriage. We limited the sample to women with children under age 18 at the time of divorce, because the presence of minor children usually led to the discussion of custody arrangements during the divorce process. We excluded women if there was no information on the custody arrangement, on education or inconsistent information on the repartnering date. The sample was limited to women who dissolved their first marriage in  the ten years before or after the custody reform of 1995. In order to avoid potential announcement effects we disregarded persons who divorced in 1995, the year of the policy reform. The final number of separated mothers considered in our analysis was 1,122. Of these, 542 (48 percent) had entered into a union within five years.

*Table 2: Sample selection procedure*

| | | N |
|---|---|---|
| Initial sample | | 6,365 |
| Excluded | | |
| - | Men | -2,965 |
| - | never divorced | -957 |
| - | childless | -594 |
| - | youngest child at time of divorce 18 years or older | -148 |
| - | information on custody missing | -25 |

| | | |
|---|---|---|
| - | information on education missing | -5 |
| - | information on who iniated separation missing | -3 |
| - | inconsistencies in repartnering date | -39 |
| - | divorced prior to 1985 | -41 |
| - | divorced after 2005 | -390 |
| - | divorced in 1995 | -76 |
| Final sample | | 1,122 |

*Note:* Data are from the “Divorce in Flanders” study, authors’ calculations.

### *Measures*

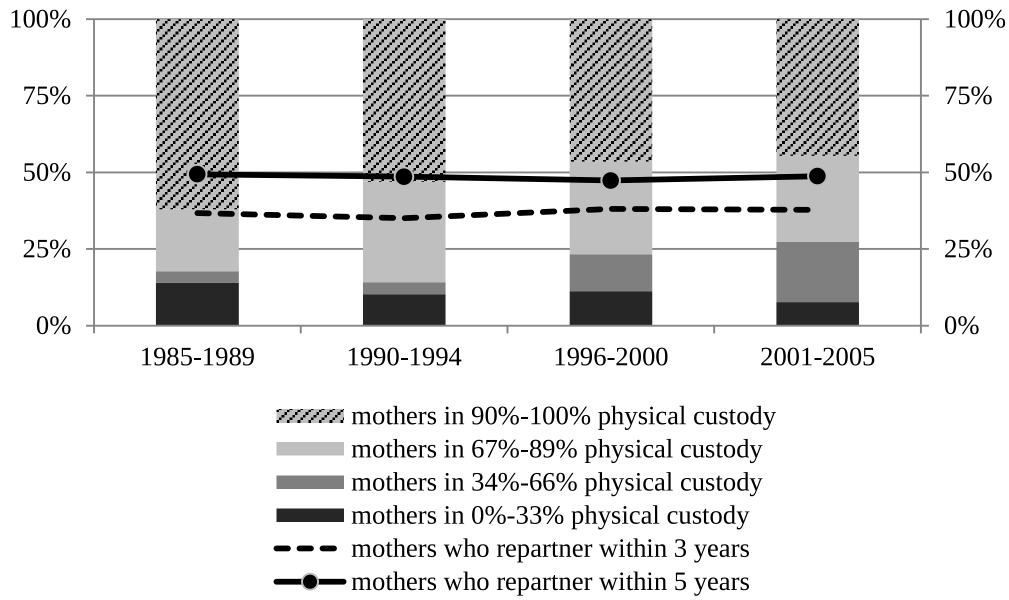
The dependent variable was defined as the formation of a household with a new partner within the first five years after the dissolution of the marital household. As a robustness check, we modeled the probability to repartner within the first three years after marital dissolution. Limiting ourselves to the first five post-separation years had several advantages. First, the restriction of the observation period to this time frame allowed a better comparison of re-partnering of older divorce cohorts and more recent divorce cohorts (which are earlier censored). Second, changes in custodial arrangements become more likely the greater the time span under study. Furthermore, a greater proportion of the children would have reached adulthood and left the maternal household (Ivanova et al., 2013). Third, this observation period was used in other studies and thus, our results can be more easily compared with other findings (Beaujouan 2010, 2012). Fourth, with the considered observation period we capture the large majority of the repartnering events in the data. 80 percent of the divorced mothers who moved together with their new partner until the interview date did so within the first five years of divorce. Considering all mothers, about 50 percent repartner within the first five years (see Figure 1), which confirms prior findings (Ivanova et al., 2013). By definition of the observation start, the mother becomes at risk of repartnering after dissolving the marital household. To avoid cases in which the dissolution of the marriage was precipitated by the presence of a new romantic partner, some studies excluded, in lack of information on the existence of a non-coresidential partnership, women who moved in with a new partner within the first months after marital dissolution. E.g., Ivanova and colleagues (2013) excluded from their sample women who experienced household formation with a new partner within the first nine months. We decided not to define any restrictions regarding the entry into observation, because this would increase the selectivity of the sample. It could be a strategy of the mother to dissolve the broken marriage only after having found a new partner in order to avoid being alone. Our data informs about the proportion of mothers that were already in a non-coresidential partnership at the time the marital household dissolved. Only as a robustness check, we withdraw these women from the sample.

The main independent variable was a binary indicator of the mother's physical custody arrangement. Berger et al. (2008) showed that sole mother's custody usually is a stable living arrangement; thus, we considered the first arrangement after marital dissolution. Only 11 percent of the mothers experienced custody changes in our sample, both into less and more equally shared custody types. Mother-sole custody has been defined in previous studies as having the children more than 66 percent of the time (Sodermans et al., 2013), more than 75 percent of the time (Cancian & Meyer, 1998) or roughly as having the children most of the time (Beaujouan 2010, 2012; Ivanova et al., 2013; de Graaf & Kalmijn, 2003). In our sample, custody arrangements varied from mothers having the children not at all living in the household (0 percent) to having the children always (100 percent). The distribution was strongly skewed towards the latter arrangement. As can be seen from Figure 1, the vast majority of the mothers who divorced between 1985 and 2005 lived more than 66 percent of the time together with their children. The most common arrangement was that the mother lived with her children 90 percent of the time or more, although this prevalence decreased over the divorce cohorts from 61 percent among women who divorced 1985-1989 to 45 percent among women who divorced in the period 2001-2005. Having the children in 90 percent or more of the time can be translated in having at maximum three days per month where the children are staying with the father. We assumed the difference in mating opportunities to be highest between mothers in this arrangement and mothers who had more child-free time at disposal. We operationalized the custody arrangement as a binary variable distinguishing between having the children almost always in the household (90 to 100 percent of the time) or less.[4] Within the latter group, mothers lived on average 60 percent of their time with their children.

Other independent variables were the woman's educational attainment level (low, medium, high), the number of the children from the divorced marriage (one child, two children or three children and more), the children's gender composition (=1 if all children are boys, 0 otherwise), as well as the age of the youngest child (continuous, in years) and of the mother at the time of separation (continuous, in years). Furthermore, we considered that whether the woman, her husband or both initiated marital dissolution had an impact on custody arrangements and repartnering outcomes. An additional variable informed about whether the woman divorced before or after the policy reform (=1 if divorced after 1995, 0 otherwise). Sample statistics are displayed in Table 3. The table also shows the difference in the proportions/means between sole physical custody mothers and mothers in other custody arrangements, including test results on the equality of proportions for binary variables, and on the equality of means for continuous variables.

[4] As a robustness check, we tested alternative categorizations, i.e. children staying in the mother's household more than 70 percent and more than 80 percent of the time.

Figure 1: *Custody type and repartnering outcome, women divorced 1985-2005*




Note: Data are from the "Divorce in Flanders" study, authors' calculations.

Table 3: *Variable definitions and sample statistics*

| Variable | Coding | N | Mean | Difference (non-sole custody(control group) minus sole custody (treatment group) ) |
|---|---|---|---|---|
| In sole physical custody (children ≥90 percent of time in household) | 0;1 | 536 | 0.48 | |
| Number of children from first marriage | | | | |
| One child | 0;1 | 405 | 0.36 | 0.01 (0.03) |
| Two children | 0;1 | 515 | 0.46 | 0.01 (0.03) |
| Three or more children | 0;1 | 202 | 0.18 | -0.02 (0.02) |
| Only boys | 0;1 | 367 | 0.33 | 0.03 (0.03) |
| Age of the youngest child at time of marital dissolution | In years 0-17 | 1,122 | 6.15 (4.53) | 0.03 (0.27) |
| Mother's age at time of marital dissolution | In years 19-52 | 1, 122 | 33.33(5.72) | 0.31 (0.34) |
| Highest educational attainment | | | | |
| Low (lower secondary school) | 0;1 | 265 | 0.24 | -0.05 (0.03)** |
| Medium (upper secondary | 0;1 | 481 | 0.43 | -0.03 (0.03) |

| | | | | |
|---|---|---|---|---|
| school) | | | | |
| High (university degree) | 0;1 | 376 | 0.34 | 0.08 (0.03)*** |
| Initiator of marital dissolution | | | | |
| Woman | 0;1 | 690 | 0.61 | 0.04 (0.03) |
| Partner | 0;1 | 238 | 0.21 | -0.07 (0.02)*** |
| Both | 0;1 | 194 | 0.17 | 0.03 (0.02) |
| Divorced after 1995 | 0;1 | 866 | 0.77 | 0.07 (0.03)*** |

*Note*: Data are from the "Divorce in Flanders" study, authors' calculations. Standard deviations of continuous variables and differences in proportions/means are in parentheses.

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

### *Model*

We used a potential outcome framework to formalize the effect of sole physical custody on repartnering. Thereby, we assumed that each mother can be exposed to two alternative states of a cause, namely being in sole physical custody or not (Wooldridge 2002: 477). Each state is characterized by a distinct set of conditions, exposure to which potentially affects an outcome of interest. Each mother has a potential outcome under each treatment state, although only one state is observed. Because it is impossible to observe both states for any mother, causal effects cannot be directly calculated at the individual level (Morgan & Winship, 2007). Exogenous variation in an instrumental variable can be used to isolate covariation in causal and outcome variables (Angrist et al., 1996; Heckman, 1997).

Our analytical strategy involved models of the probit family (Wooldridge 2002: 477). We preferred a probit model set up to an event history model, because in the latter, estimates combine tempo and level and are thus more difficult to interpret (Beaujouan, 2012; Gray et al., 2009). Furthermore, any kind of anticipatory analysis should be avoided in event history models and thus, the ordering of the decision making processes have to be clear. Using a probit model approach we focused on the repartnering outcome within the first five years, putting less emphasis on the timing of post-marital household formation. The use of probit models had further the advantage that we had to be less concerned about the ordering of the considered processes; it was sufficient to assume that probabilities are interrelated. In general, in the presence of binary endogenous and dependent variables, a recursive bivariate probit model should be preferred over linear estimation methods (Woodridge 2002: 477; Monfardini & Radice, 2008). We estimated the effect of a full-time physical custody $C_i$, on repartnering $Y_i$ in a recursive bivariate probit model, in which full-time physical custody was assumed to be a binary endogenous explanatory variable (Wooldridge 2002: 477). The model was

$$Y_i = 1[\alpha\, C_i + \gamma X_i + u_i > 0]$$

$$C_i = 1[\beta\, Z_i + \varphi X_i + \varepsilon_i > 0]$$

where the treatment indicator $C_i$ took the value one if treatment was received (the mother lived full-time with her children) and zero otherwise. The matrix $X_i$ consisted of a set of control variables. If $C_i$ was correlated with $u_{i,}$ there would be an endogeneity problem. The custody reform in 1995 was taken as an instrumental variable $Z_i$ for $C_i$. We assumed $u_i$ and $\varepsilon_i$ to have zero mean, bivariate normal distribution, $\rho_1$ = Corr($u_i$., $\varepsilon_i$) and to be independent from $Z_i$. The parameters α, γ, β, φ and ρ were estimated by maximum likelihood.

Based on the coefficients of this model we predicted treatment effects (XX). Assume a mother that did not receive treatment (C=0) so that we observe the repartnering outcome $Y_0$. What would be the repartnering outcome if the same mother would have undergone the treatment? Let $Y_1$ be her counterfactual outcome. For a mother that received treatment (C=1), we observe $Y_1$, while $Y_0$ would be her counterfactual outcome. Formally, the potential-outcome model is:

$$Y = (1 - C)Y_0 + Y_1$$

We made use of three treatment effect estimators to estimate the potential outcome under study. The potential-outcome means (POMs) are the means of $Y_1$ and $Y_0$ in our study population. The average treatment effect (ATE) is the mean of the difference ($Y_1$-$Y_0$) and the average treatment effect on the treated (ATET) is the mean of the difference ($Y_1$-$Y_0$) among the mothers that actually received the treatment (C=1). The POMs were calculated by estimating the marginal predicted repartnering probabilities of sole-physical custody mothers and mothers in other custody arrangements while keeping their other characteristics to the mean values. For the set of control variables, derivatives of the marginal predicted probabilities of repartnering success were calculated, with the estimator for custody arrangement being the ATE. The ATET was calculated based on derivatives of the predicted repartnering probability conditional on having sole physical custody. All margins and marginal effects reported were mean predictions.

Several requirements had to be met to define custody reform as an appropriate instrumental variable (DeMaris, 2014; Woodridge, 2002). First, the instrument had to be relevant; that is, it should be correlated with the treatment $C_i$. Second, the instrument should affect repartnering only through the custody choice; this is known as the "exclusion restriction". Third, the instrument should be exogenous in the equation for $Y_i$; that is, it should not share common cause with the outcome variable. Fourth, the effect of the instrument on the potentially endogenous variable should be monotone. In the following we show that these requirements were fulfilled for the chosen instrumental variable. Policy reform was found to be correlated with custody choice, but not with repartnering. Descriptive measures in Figure 1 show that while custody arrangements changed for mothers of different divorce cohorts, the proportion of divorced women who re-partner within five years remained pretty constant. Probit regressions of custody arrangement on the custody reform dummy and controls yielded significant coefficients of the reform dummy (see Appendix). Estimations using dummies for five year intervals (see Table A1 in Appendix) revealed that compared to mothers divorced in the period 2001-2005, the probability of sole physical custody was higher among mothers who divorced in the periods 1985-1989 and 1990-

1994, but not among mothers who divorced in the period 1996-2000. To check the assumption of exclusion restriction, we estimated probit regression of repartnering including the custody reform dummy in the set of control covariates (see Table A2 in Appendix). The coefficient of the reform dummy was not significant, as long as the age at marital dissolution was not controlled for. Including this information yielded a significant coefficient of the reform. The change in the coefficient could be completely attributed to a decomposition effect (an interaction effect was tested and was found to be insignificant). Repartnering probabilities were decreasing with age and women who separated after 1995 appeared to be somewhat older than women separated prior to 1995. The Smith-Blundell test tests for exogeneity of the custody arrangement in the probit model (Baum et al., 2003; Smith & Blundell, 1986). With a p-value of 0.08, this null hypothesis can be rejected. We believe the monotonicity assumption was satisfied. It would be violated if some mothers chose sole custody because of the policy reform; this is unlikely to have been the case. All in all, the tests indicate that our chosen variable presented a valid instrument.

## FINDINGS

### *Multivariate results*

The effects of physical custody and further controls on repartnering are listed as coefficients and marginal effects in Table 4. First, we estimated a univariate probit model of repartnering (Model 1). In Model 2, potential endogeneity of custody arrangement was accounted for. The coefficient results of repartnering show that the effect of being a full-time residential parent is negative and significant ($p<0.01$) in both models. The effect of the instrumental variable (divorced after the policy reform of 1995) is highly significant ($p<0.01$) in Model 2. The Wald test result shows that unobserved selection influenced repartnering ($\rho$ is significantly different from zero). This signals that there is no identification problem in the model: According to Monfardini and Radice (2008), the model may be poorly identified if the endogenous variable shows a significant estimate in the univariate probit model, whereas its coefficient and the correlation coefficient are insignificant in the recursive bivariate probit estimation.

Being in sole physical custody arrangement was associated in Model 1 with an average reduction of repartnering probability by 15 percentage points compared to mothers in other physical custody arrangements. In Model 2, we investigated whether this effect was causal by allowing for endogeneity using a recursive bivariate probit model with reform-based exclusion restriction. In this model, the effect of sole physical custody became even more pronounced: mothers with average characteristics who had the children more than 90 percent of their time after marital dissolution experienced a reduction of repartnering probability by 56 percentage points compared to their counterparts with more childfree time. This result is supporting our second hypothesis, in which we assumed that ordinary models tend to underestimate the negative effect of sole physical custody. Indeed, we found a positive effect of ρ, suggesting that unobserved characteristics make mothers more likely to be in sole physical custody and to repartner. In Table 5, the results of the counterfactual model are displayed, which allows to compare the potential outcome means of divorced mothers with and without the treatment of sole physical custody in the probit model and the recursive bivariate probit model. Ignoring the endogeneity of custody choice in the ordinar y probit model, we found that 55 percent of the mothers would repartner within the first five years after marital dissolution if they were having the children less than 90 percent of their time. In contrast, 40 percent of the mothers would repartner within the same period, if they had the children living in the household 90-100 percent of the time. This means that sole physical custody reduces the probability of repartnering by 15 percentage points or 27 percent. In other words: if the mothers had to take care of their children almost always, every fourth step-family would not be formed compared to the setting in which separated fathers share childrearing tasks with the mother. Accounting for the endogeneity of custody choice in the recursive bivariate probit model, we found that 76 percent of the mothers would repartner within the first five years after marital dissolution if the children were staying more than 10 percent of the time with the father, whereas 20 percent of the mothers would repartner within the same period, if they had the children living in the household almost always. This comparison shows that ordinary models tend to overestimate sole-custody mothers' repartnering probabilities and to underestimate the repartnering probabilities of mothers in more equally shared custody arrangements. Sole physical custody is reducing the probability of repartnering in fact by 74 percent. This means that only one quarter of the step-families would be formed if the mothers were the sole caretakers of their children, compared to the setting in which separated fathers were involved in childrearing tasks. Looking at the coefficients of the control variables of Model 1 and Model 2, we find that the number and the gender of children from first marriage did neither influence custody arrangement nor repartnering. The older the youngest child, the higher was the probability that the mother repartnered. Mother's age at the time of marital dissolution was found to negatively influence the probability of repartnering. Having low education lowered the mother's probability of repartnering by 10 percent in Model 1. Accounting in Model 2 for the fact that low educated mothers had a higher probability to obtain full-time custody, educational differences in repartnering vanished. This result shows that educational differences in repartnering can be explained with the heterogeneity in uptaking of sole physical custody among mothers from different educational groups. A similar effect can be found when looking on who initiated marital separation. The results of Model 1 shows that if the divorce was upon the initiative of the woman's partner only, her repartnering probability was significantly lower compared to the case in which the woman initiated the divorce process – alone or together with her first marriage partner. Model 2 however revealed that women who were left by their spouse had a lower probability of repartnering, because they ended more often in a sole physical custody arrangement.

Table 4 *Coefficient results and marginal effects of repartnering, univariate probit model and recursive bivariate probit model (N=1,222)*

| | Probit: Repartnering Model 1 | | Biprobit: Model 2 | | | |
|---|---|---|---|---|---|---|
| | | | Repartnering | | Sole physical custody | |
| | B | ME | B | ME | B | ME |
| Number of children form first marriage (ref= one child) | | | | | | |
| Two children | 0.04 | 0.01 | 0.05 | 0.01 | 0.03 | 0.01 |
| | (0.09) | (0.03) | (0.08) | (0.02) | (0.09) | (0.04) |
| Three or more children | 0.01 | 0.00 | 0.09 | 0.02 | 0.13 | 0.05 |
| | (0.13) | (0.05) | (0.11) | (0.03) | (0.12) | (0.05) |
| Only boys | 0.01 | 0.00 | -0.04 | -0.01 | -0.08 | -0.03 |
| | (0.09) | (0.03) | (0.08) | (0.02) | (0.08) | (0.03) |
| Age of youngest child at marital dissolution | 0.04*** | 0.01*** | 0.03** | 0.01** | 0.01 | 0.00 |
| | (0.01) | (0.01) | (0.01) | (0.00) | (0.01) | (0.01) |
| Mother’s age at marital dissolution (-33) | -0.08*** | -0.03*** | -0.07*** | -0.02*** | -0.01 | -0.00 |
| | (0.01) | (0.00) | (0.01) | (0.00) | (0.01) | (0.00) |
| Highest educational attainment (ref=High) | | | | | | |
| Low | -0.23** | -0.08** | -0.01 | -0.00 | 0.26** | 0.10** |
| | (0.11) | (0.04) | (0.11) | (0.03) | (0.11) | (0.04) |
| Medium | -0.06 | -0.02 | 0.07 | 0.02 | 0.20** | 0.08** |
| | (0.09) | (0.03) | (0.09) | (0.02) | (0.09) | (0.04) |
| Initiator of divorce (ref=Mother) | | | | | | |
| Partner | -0.42*** | -0.15*** | -0.15 | -0.04 | 0.28*** | 0.11*** |
| | (0.10) | (0.04) | (0.11) | (0.03) | (0.10) | (0.04) |
| Both | -0.15 | -0.05 | -0.14 | -0.04 | -0.07 | -0.03 |
| | (0.11) | (0.04) | (0.10) | (0.03) | (0.10) | (0.04) |
| **Sole physical custody** (ref=yes) (children ≥90 percent of time in household) | **-0.40***<br>(0.08)** | **-0.15***<br>(0.03)** | **-1.62***<br>(0.12)** | **-0.56***<br>(0.04)** | | |
| Divorced after 1995 | | | | | -0.26*** | -0.10*** |
| | | | | | (0.08) | (0.03) |
| Constant | 0.08 | | 0.56*** | | -0.09 | |
| | (0.13) | | (0.12) | | (0.14) | |
| rho | | | | | 0.90 | |
| | | | | | (0.15) | |

*Note:* Data are from the "Divorce in Flanders" study, authors' calculations. Effects in beta-coefficients (B) and Marginal effects (ME), calculated as mean over the sample. Standard errors in parentheses. Log Likelihood in Model 1: -711, Log likelihood in Model 2:-1,471. Wald test: chi2(1) = 3.70633 Prob > chi2 = 0.0542

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

Table 4 *Potential outcome means and average treatment effects (N=1,222)*

| | Probit model | Biprobit model |
|---|---|---|
| Potential outcome means (POMs) | | |
| POM of $Y_0$ | 0.55(0.02) | 0.76(0.02) |
| POM of $Y_1$ | 0.40(0.02) | 0.20(0.02) |
| Average treatment effect (ATE) | -0.15(0.03) | -0.56(0.04) |
| Average treatment effect of the treated (ATET) | -0.15(0.03) | -0.58(0.04) |
| *Change in outcome after treatment (in %)* | *-27%* | *-74%* |

*Note*: Data are from the "Divorce in Flanders" study, authors' calculations. POMs, ATE, ATET are calculated as margins and marginal effects (delta method) with mean predictions. Standard errors in parentheses.

*Sensitivity Analysis*

**Methodological robustness checks**

Even in our framework that accounts for endogeneity of custody choice, one has still to worry about the possibility that treatment is contaminated by endogeneity, because we relied on non-experimental data.[5] Endogeneity would impart selection bias to any estimates. In order to purge remaining bias we used augmented inverse probability weighting (AIPW) estimations to estimate mothers' repartnering outcomes (Stata, 2013; Tan, 2010). AIPW estimators use inverse-probability weights to account for the fact that each subject is observed in only one of the potential outcomes and an augmentation term in the outcome model to correct misspecification in treatment. The AIPW estimator has the advantage to have the double-robust property for some functional form combinations; that is, if either the outcome model or the treatment model is correctly specified, the effects are consistently estimated. ATETs could not be modeled via the AIPW estimator; we used instead the inverse-probability-weighted (IPW) estimator, which uses weighted means to disentangle the effects of treatment and confounders. A drawback of these estimators is that they are most suitable under selection-on-observables. When not all variables common to treatment assignment and outcome are observable, the outcomes are not conditionally independent of the treatment. To our knowledge, there is no probit treatment model with binary treatment and binary outcome accounting for selection-on-unobservables. With a POM($Y_0$) of 55 percent, a POM($Y_1$) of 41 percent, and both ATE and ATET of -14 percentage points, the treatment estimators appeared to be very similar to the estimators of Model 1 in Table 4. This demonstrates that the coefficients in the ordinary probit model are not strongly biased by

[5] If treatment and control units were allocated randomly, the probability density function of each of the control variables would be the same for treatment and control group. In Table 3, the equality of the proportions and means in the treatment and control group of mothers was tested. Especially educational background and whether the woman or her partner initiated the separation process were found to differ between these groups.

observed selection and that unobserved characteristics linking sole physical custody choice to successful repartnering play a central role in identifying the causal effect of physical custody on repartnering.

But, our empirical findings might hinge on the functional form assumption of the bivariate probit model. To investigate whether this could be the case, we tested the normality assumption with the Murphy's score test (Chiburis, 2009; Murphy, 2007). This goodness-of-fit score test embeds the bivariate normal distribution within a larger family of distributions by adding more parameters to the model. The model is correctly specified if the null hypothesis of normality cannot be rejected, which was the case in our estimations (Prob > $chi^2$ = 0.45).

Alternative instrumental variable models that account for selection-on-unobservables belong to the family of linear regression models. In recent literature, estimations of linear regression models in the presence of binary endogenous and dependent variables have been encouraged, especially next to a recursive bivariate probit model that relies on non-linearity. We used two-stage-least-squares (2SLS) estimators and limited information maximum likelihood (LIML) estimators in a linear regression model. The latter have the advantage to provide the same asymptotic distribution as 2SLS, whereas the finite-sample bias is reduced (Angrist et al., 2008). Monte Carlo studies have found these modelling strategies and the bivariate probit model strategy to give very similar results (Altonij et al., 2004; Angrist et al., 2008: 209). Indeed, with a treatment effect of 75 percent, the results of the linear regression model are quite identical with those of a bivariate probit (see change in outcome after treatment in Table 4).

Next to nonlinearity, we investigated the role of the exclusion restriction as the source of identification in the bivariate probit model. Estimations in which the instrumental variable was excluded reached convergence only with difficulties (after 1922 iterations) and showed an average treatment effect of sole physical custody of similar dimension and expected direction like in the original model, but with large standard errors (ATE -0.44(S.E. 0.51)).

In sum, our methodological checks point out that the causal effect of sole physical custody on repartnering is mainly identified through the exclusion restriction.

**Sample restrictions**

In Table 5, we present the results of three sets of robustness checks, where we used selected samples, a modified model set up and different instruments. For comparison, we replicate the main results in the first row of the table. As a first sample check, we restricted our sample to women who divorced up to seven years – instead of ten years – prior or after the reform year 1995 (i.e. between 1988 and 2002). As a second sample check, we included in our sample only women who divorced between 1985 and 2000. This means we analyzed the repartnering behavior of the youngest divorce cohorts up to the year 2005, the year preceding the custody reform of 2006. The results were robust to these modifications, only the standard errors increased somewhat due to the smaller sample size. As a third sample check, we excluded women who had the children less than one third of the time living in the household, because we assumed that they were potentially different from other mothers in many aspects. Excluding them did only slightly decrease our sample size; the results stayed robust to this change and the unobserved selection term ρ became even more significant. As a fourth sample check, we considered only women who had no non-coresidential partner at the time of marital dissolution. Again, model results did not change, whereas the standard errors increased somewhat.

**Modified model set up**

A test of the exclusion assumption is to use the policy dummy instead of the custody variable as the main variable of interest to model the probability of repartnering in a so-called intention-totreat (ITT) model (Angrist et al., 1996). If our assumptions hold, the effect estimated by IV analysis should have been higher in magnitude than the one obtained by the ITT analysis (Arpino et al., 2014). The last column in Table A2 shows the results of this ITT model (Model R4). Comparing the marginal effect of the policy reform (0.08; $p<0.05$) in the ITT model with the marginal effect of sole physical custody in Model 1 and Model 2 (Table 3) reveals indeed a stronger repartnering effect of the custody arrangement.

As another test of the exclusion restriction, we imposed a fake reform in 2001 on the post-reform data, following the approach of Nielsen et al. (2009). As expected, the instrument turned out to be insignificant in this model. The effect of custody changed little.

To test the importance of the length of the observation period, we defined our dependent variable as repartnering within the first three years – instead of five years – after marital split-up. The model results were robust to this modification.

Finally, we changed the our binary endogenous variable by defining those having sole physical custody as a) having the children more than 80 percent of the time, b) more than 70 percent of the time, and c) having them more than 60 percent of the time. Again, the results were robust to these changes. The average treatment effect was with -0.46 (a), -0.45 (b), and -0.44 (c) somewhat smaller than when defining sole physical custody as living with the children more than 90 percent of the time.

**Adding additional information**

In principal, other political forces could have influenced custody type and repartnering, but we are not aware of relevant reforms other than the legal custody reform in the observation period. The exclusion restriction would have been also threatened if parallel time trends affected the repartnering behaviour by e.g. increasing the mating opportunities for mothers in sole custody.

The expansion of internet for example can be seen as a potential to reduce search frictions and may be especially beneficial for the partner search outcome of divorced mothers with minor children (Bellou, 2015). Internet use in private households started in the 1990s, also in the Belgian context. We conducted a robustness check by relating the proportion of Belgian households with internet connection with data from the World Bank to the year of marital dissolution. Estimating a model equivalent to Model 1, but complemented by information on the proportion of internet users in the year of separation, we found that higher proportions of internet access were related to higher probabilities of repartnering. We found no interaction effect between the custody arrangement and the availability of internet (see Figure A1a and Figure A1b in Appendix). Furthermore, we found that – similar to the effect of the policy dummy variable in Table A2 – there was no significant effect of internet access when women's age at marital dissolution was not controlled for (Figure A1b). Considering this information in the bivariate probit model did not lead to changes in the custody coefficients, while it increased somewhat the standard errors and turned the policy reform dummy insignificant (see Table 5). The latter effect can be attributed to the substantive level of correlation (corr=0.50) between the internet and the policy reform variable.

One may also argue that the increase in female employment rates in the respective period was responsible for the shift in custody type. Working mothers might be more inclined to take up shared custody arrangements in order to be able to continue to work. On the other hand, the increase of public childcare has made childrearing obligations increasingly easy to combine with employment, suggesting a decoupling of working and custody decisions. In our sample, the proportion of full-time working mothers remained relatively stable over the observation period, whereas there was a shift from nonactivity to parttime work. Empirical studies identified a link between employment status and custody with full-time working mothers being less likely to live in sole physical custody, whereas they found no association between mother's employment and repartnering (Boeheim & Ermisch, 1998; Cancian & Meyer, 1998; de Graaf & Kalmijn 2003; Juby et al., 2005). As a robustness check, we included measures of economic activity, expressed in the mother's employment status (nonactive, parttime, fulltime) at the time of marital dissolution. This did not change the custody coefficients in the bivariate probit model. Confirming prior findings, our results showed that the employment status had no significant impact on repartnering. But, in our sample fulltime working mothers appeared to be more likely than parttime working mothers to have sole physical custody, which may underline the argument that these mothers face higher financial needs.

Through the policy reform variable we captured an exogenous shock that changed the mothers' exposure to joint legal custody regulations. We assumed this exposure to be related to the chosen physical custody arrangement. We also had information on the chosen legal custody arrangement at hand which we included into the model as a further robustness check, like proposed by Nielsen et al. (2009). Having sole legal custody increased the probabilities to have also sole physical custody, but it had no significant effect on the repartnering probability.

Overall, the treatment effects have been shown to be pretty invariant to sample restrictions, model modifications and additional covariates.

Table 5: *Robustness checks – Effects of policy reform on probability of sole physical custody and custody arrangement on repartnering in the recursive bivariate probit model*

| | N | Policy reform → sole custody (ME) | Sole custody → repartnering (ME) | rho |
|---|---|---|---|---|
| Original model (Model 2) | 1,122 | -0.10(0.03)*** | -0.56(0.04)*** | 0.90(0.15)* |
| Sample restrictions | | | | |

| | | | | |
|---|---|---|---|---|
| Only women divorced between 1988 and 2002 | 754 | -0.08(0.03)** | -0.54(0.09)*** | 0.84(0.27) |
| Only women divorced between 1985 and 2000 | 612 | -0.09(0.04)** | -0.50(0.20)*** | 0.70(0.52) |
| Only women living with the children at least 1/3 of time | 1,015 | -0.13(0.03)*** | -0.57(0.03)*** | 0.91(0.09)*** |
| Only women without a non-coresidential partner at time of marital dissolution | 937 | -0.08(0.04)** | -0.48(0.14)*** | 0.71(0.34) |
| Modified model set up | | | | |
| Fake policy reform (2001), post-reform data | 866 | -0.03(0.03) | -0.53(0.14)*** | 0.79(0.32) |
| Dependent variable: repartnering within 3 years | 1,122 | -0.11(0.03)*** | -0.50(0.08)*** | 0.76(0.20)** |
| Sole physical custody (children ≥80 percent of time in household) | 1,122 | -0.12(0.03)*** | -0.46(012)*** | 0.58(0.29) |
| Sole physical custody (children ≥70 percent of time in household) | 1,122 | -0.10(0.03)*** | -0.45(0.09)*** | 0.66(0.28)* |
| Sole physical custody (children ≥60 percent of time in household) | 1,122 | -0.08(0.03)*** | -0.44(0.10)*** | 0.62(0.24)* |
| Adding additional information | | | | |
| + information on internet use in Belgian population | 1,122 | -0.05(0.04) | -0.54(0.17)*** | 0.83(0.52) |
| + information on economic activity | 1,057 | -0.08(0.03)*** | -0.56(0.04)*** | 0.91(0.05)* |
| + information on legal custody (sole or shared) | 1,076 | -0.05(0.03)* | -0.50(0.13)*** | 0.73(0.34) |

*Note*: Data are from the "Divorce in Flanders" study, authors' calculations. Full model results are available upon request from the authors. Significance levels for marginal effects based on p-levels; significance levels for rho based on Wald test results.

## DISCUSSION

A divorce usually raises the issue of the children's residence. The decision for a certain physical custody arrangement has an impact on future private life outcomes of the parent, as several recent studies have shown. Dissolved parents who are the main child care providers have lower chances of repartnering than parents who share the physical custody with their former partner or who have non-residential children. But, so far, repartnering studies have used traditional regression approaches that did not consider that custody choice might be endogenous to repartnering. There are potentially two different effects of custody arrangement on repartnering: the effect of sole physical custody for mothers in sole physical custody and the potential effect of sole physical custody for those mothers who have an alternative arrangement. Standard regression estimates only are unbiased, if these two effects are identical. A counterfactual framework – like the one used in this study - enables to address this issue.

In this article, we discussed in detail the different channels through which custody choice and repartnering can be related. We found arguments according to which mothers that are more likely to opt for sole physical custody should be less likely to repartner. Not accounting for this selectivity might have led researchers to overestimate sole-custody mothers' low repartnering probabilities. But, we also found arguments to the contrary which linked mothers in sole physical custody to a higher likelihood of repartnering. This resulted in two opposing hypotheses, which were empirically tested by using a recursive bivariate probit model. We exploited exogenous variation in physical custody arrangement by a policy reform promoting joint legal custody as an instrumental variable. We defined sole physical custody as an residential arrangement in which the mother lives with the children 90 percent of the time or more.

Our main empirical finding is that the negative effect of sole physical custody is causal. Furthermore, low repartnering probabilities of women with full-time residential children are still understated in ordinary models that do not account for the endogeneity of custody choice. With the results of our counterfactual model, the consequences of different custody arrangements can be illustrated: We estimated that in ordinary probit models, sole physical custody reduces the probability of repartnering by 27 percent. Accounting for the endogeneity of custody choice in the recursive bivariate probit model, we found that sole physical custody is reducing the probability of repartnering in fact by 74(!) percent. Other things equal, this means that only one out of four step-families (instead of three out of four according to ordinary regression results) would be formed if the mothers had to take care of their children almost always, compared to a setting in which separated fathers share childrearing tasks with the mother.

Why does the presence of children decrease the probability of repartnering that strongly? Having the children 90 to 100 percent living in the household can be translated into at maximum 3 child-free days per month. If sole-custody mothers search a partner with the same intensity as mothers with more child-free time, the repartnering outcome is likely to be worse due the lower quantity of time available for partner search. This has been called the “opportunity” argument in previous repartnering research (de Graaf & Kalmijn, 2003). But, the search intensity of sole-custody mothers can also be lower than that of their peers, because they have less an interest in finding a new partner, which has been defined as the “need” hypothesis. The mother’s emotional needs might be more satisfied if the children are always around. Finally, the search outcome of sole-custody mothers can be worse, because the continued presence of the children keeps potential candidates away, which corresponds to the “attractiveness” hypothesis. The strong negative effect of sole physical custody on repartnering is likely to be a sum of lacking opportunities, lower needs and lower attractiveness. Although we accounted in our study for the specific selection of some women into sole physical custody, it was not possible to disentangle the different mechanisms behind.

Why had the negative impact of sole physical custody on repartnering been underestimated in the ordinary model? Ordinary models fail to appropriately account for the fact that women with and without sole physical custody are not alike. Custody arrangement is not randomly distributed among divorced mothers. Sole-custody mothers possess some characteristics related to needs, opportunities and attractiveness that are related to higher repartnering probabilities, whereas mothers not having sole physical custody have some characteristics that make them less likely to repartner. Sole custodians may have a high family orientation which makes them prone to accept the first suitable candidate in order to provide the children a new father figure and to restore the image of a complete family. After divorce, some mothers prefer to focus on their own needs and activities. Prioritizing self-actualization, they may want to commit their daily life to neither children nor the partner. Mothers who are less physically attractive might opt against sole physical custody to increase their opportunities to meet a partner. Mothers might decide against a shared physical custody arrangement because it obliges them to stay geographically close to the father. By opting for sole physical custody, the mother has more flexibility in choosing the place of living, which also facilitates moving in with a new partner.

What are the implications of our findings? The risk of being poor is doubled in single-parent families. Step-family formation can work as a way to increase the income of the family. Our study has shown that the physical custody arrangement constitutes a strong impediment for step-family formation. Policies that aim to support divorced families should especially focus on single-custody families, as their living situation is likely to be there to stay.

We believe the sample used for this study to be very well suited for our research purpose, as it includes detailed information on the custody arrangement in the period of interest. The data nevertheless has several limitations that need to be mentioned (Sodermans et al., 2013). An important drawback is the special sampling design of the “Divorce in Flanders” study. First, it excluded mothers that divorced for a second time. Leaving these women out of the sample might have led to an underestimation of mothers' repartnering probabilities. Second, information on custody arrangement referred to the selected target child, which did not represent the population of Flemish children with divorced parents, as it was on average somewhat older (Sodermans et al., 2013). Finally, as any other retrospective data, our sample could have been subject to a recall bias. We relied on individual retrospective information, as remembered and reported by the mothers during the interview. Mothers who divorced before 1995 might recall the characteristics of the first postmarital period differently than mothers who divorced more recently, e.g. they might remember not all changes over time in the custody arrangement, but only the most stable arrangements. We only took the first residential arrangement after divorce into account which might be differently remembered by mothers of different divorce cohorts.

There is a trend towards more equally shared childrearing arrangements after divorce that has been promoted by recent policy reforms. This means that in the future, sole physical custody will become less common and sole-custodians will belong to a more and more selective group. Characteristics that have been argued in this article to promote the repartnering of sole-custody mothers may not be found among the sole-custody mothers of the future. If this is the case, the negative effect of sole physical custody is likely to become more evident in future divorce cohorts.

REFERENCES

Altonji, J. G., Elder, T. E., & Taber, C. R. (2005). An evaluation of instrumental variable strategies for estimating the effects of catholic schooling. *Journal of Human Resources*, 40(4), 791-821.

Angrist, J. D., & Pischke, J. S. (2008). *Mostly harmless econometrics: An empiricist's companion*. Princeton university press.

Amato, P. R., & Maynard, R. A. (2007). Decreasing nonmarital births and strengthening marriage to reduce poverty. *The Future of Children*, 17(2), 117-141.

Baum, C.F., Schaffer, M.E., &Stillman, S. (2003). Instrumental Variables and GMM: Estimation and Testing. *The Stata Journal*, 3(1), 1-31. Working paper version: Boston College Department of Economics Working Paper No 545. http://ideas.repec.org/p/boc/bocoec/545.html

Bauserman, R. (2002). Child adjustment in joint-custody versus sole-custody arrangements: A meta-analytic review. *Journal of Family Psychology*, 16(1), 91–102.

Bellou, A. (2015). The impact of Internet diffusion on marriage rates: evidence from the broadband market. *Journal of Population Economics*, 28(2), 265-297

Bender, W. N. (1994). Joint custody: The option of choice. *Journal of Divorce and Remarriage*, 21, 115 – 131.

Beaujouan, É. (2010). Children at home, staying alone? Paths towards repartnering for men and women in France. *Centre for Population Change Working Paper*, 1-24.

Beaujouan, E. (2012). Repartnering in France: The role of gender, age and past fertility. *Advances in Life Course Research*, 17(2), 69-80.

Becker, G. S. (1991). *A treatise on the family*. Cambridge: Harvard University Press.

Bernhardt, E. (2000). Repartnering among Swedish men and women: A case study of emerging patterns in the second demographic transition. Paper presented at the FFS flagship conference.

Bemhardt, E. & Goldscheider F. K. (2002). Children and union formation in Sweden. *European Sociological Review*, 18(3), 289-299.

Berger, L. M., Brown, P. R., Joung, E., Melli, M. S., & Wimer, L. (2008). The stability of child physical placements following divorce: Descriptive evidence from Wisconsin. *Journal of Marriage and Family*, 70(2), 273-283.

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

Bzostek, S. H., McLanahan, S. S., & Carlson, M. J. (2012). Mothers' repartnering after a nonmarital birth. *Social forces,* 90(3), 817-841.

Chiburis, R. C. (2009). Score tests of normality in bivariate probit models: Comment and implementation. Working paper, University of Texas at Austin.

Corijn, M. (2005), Echtscheidingen in België: met of zonder kinderen. *CBGS-Sitemap, Uit het onderzoek*, 17 oktober 2005.

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

DeMaris, A. (2014). Combating unmeasured confounding in cross-sectional studies: Evaluating instrumental-variable and Heckman selection models. *Psychological methods*, 19(3), 380.

Dykstra, P. A. & Poortman, A. (2010). Economic resources and remaining single: Trends over time. *European Sociological Review*, 26(3), 277-290.

England P, McClintock EA (2009) The gendered double standard of aging in US marriage markets. *Population and Development Review*, 35, 797–816.

Fehlberg, B., Smyth, B., Maclean, M., & Roberts, C. (2011). Legislating for shared time parenting after separation: A research review. *International Journal of Law, Policy and the Family*, 25(3), 318–337.

Goldscheider, F. K. & Sassler, S. (2006). Creating stepfamilies: integrating children into the Study of Union Formation. *Journal of Marriage and Family*, 68(2), 275-291.

Goldscheider, F. K. & Waite, L. J. (1986). Sex differences in the entry into marriage. *American Journal of Sociology*, 92(1), 91-109.

Heckman, J. (1997). Instrumental variables: A study of implicit behavioral assumptions used in making program evaluations. *Journal of Human Resources*, 32(3), 441-462.

Gunnoe, M. L. & Braver, S. L. (2001). The effects of joint legal custody on mothers, fathers, and children controlling for factors that predispose a sole maternal versus joint legal award. *Law and Human Behavior*, 25(1), 25-43.

Jansen, M., Mortelmans, D., & 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.

Juby, H., Marcil-Gratton, N., & Le Bourdais, C. (2005). *When parents separate: Further findings from the National Longitudinal Survey of Children and Youth*. Department of Justice Canada.

Juby, H., Le Bourdais, C., & Marcil-Gratton, N. (2005). Sharing roles, sharing custody? Couples' characteristics and children's living arrangements at separation. *Journal of Marriage and Family*, 67(1), 157-172.

Kreyenfeld, M., & Martin, V. (2011). Economic conditions of stepfamilies from a cross-national perspective. *Journal of Family Research/ Zeitschrift für Familienforschung*, 23(2), 128-153.

Koo, H. P., Suchindran, C. M. & Griffith, J. D (1984). The effects of children on divorce and re-marriage: a multivariate analysis of life table probabilities. *Population Studies*, 38(3), 451-471.

Lampard R. & Peggs, K. (1999). Repartnering: the relevance of parenthood and gender to cohabitation and remarriage among the formerly married. *British Journal of Sociology*, 50(3), 443-465.

Leridon H. and L. Toulemon (1997). *Démographie. Approche statistique et dynamique des populations.* Paris, Economica, 440 p.

Lundberg, S., & Rose, E. (2003). Child gender and the transition to marriage. *Demography*, 40(2), 333-349.

Martin, C. (1994). Diversité des trajectoires post-désunion. Entre le risque de solitude, la défense de son autonomie et la recomposition familiale. *Population*, 49(6), 1557-1583.

Meggiolaro, S. & Ongaro, F. (2008). Repartnering after marital dissolution: does context play a role? *Demographic Research*, 19(5), 1913-1932.

Manning, W. D., & Brown, S. (2006). Children's Economic Well-Being in Married and Cohabiting Parent Families. *Journal of Marriage and Family*, 68(2), 345-362.

Monfardini, C., & Radice, R. (2008). Testing exogeneity in the bivariate probit model: A monte carlo study. *Oxford Bulletin of Economics and Statistics*,70(2), 271-282.

Mortelmans, D., Pasteels, I., Bracke, P., Matthijs, K., Van Bavel, J., & Van Peer, C. (2012). *Divorce in Flanders. Codebooks and questionnaires.* Retrieved from www.divorceinflanders.be.

Murphy, A. (2007). Score tests of normality in bivariate probit models. *Economics Letters*, 95(3), 374-379.

Ni Bhrolchain, M., & Sigle-Rushton, W. (2005). Partner Supply in Britain and the US. Estimates and Gender Contrasts. *Population*, 60(1-2), 37–64.

Pasteels, I., Bracke, P., Mortelmans, D., Matthijs, K., Van Bavel, J., & Van Peer, C. (2013). *Scheiden in meervoud: over partners, kinderen en grootouders.* Leuven: Acco.

Pasteels, I., & Mortelmans, D. (2013). Gescheiden en dan? Herpartneren in Vlaanderen anno 2010. *Relaties en Nieuwe Gezinnen*, 3(3), 1-66.

Pasteels, I., Corijn, M., & Mortelmans, D. (2012). Een nieuwe partner na een echtscheiding? Opleidingsverschillen bij mannen en vrouwen in Vlaanderen. *Tijdschrift voor Sociologie*, 33(3-4), 331–352.

Poortman, A. R. (2007). The First Cut is the Deepest? The Role of the Relationship Career for Union Formation. *European Sociological Review*, 23(5), 585-598.

OECD (2014). *Family Database.* www.oecd.org/social/familydatabase.

Seccombe, K. (2002). “Beating the odds” versus “changing the odds”: Poverty, resilience, and family policy. *Journal of Marriage and Family*, 64(2), 384-394.

Skinner, C., & Davidson, J. (2009). Recent trends in child maintenance schemes in 14 countries. *International Journal of Law, Policy and the Family*, 23(1), 25-52.

Smith, R. J., & Blundell, R. W. (1986). An exogeneity test for a simultaneous equation Tobit model with an application to labor supply. *Econometrica*, 54(4), 679-686.

Smock, P. J. (1993). The economic costs of marital disruption for young women over the past two decades. *Demography*, 30, 353 – 371.

Sodermans, A. K., Matthijs, K. & Swicegood, G. (2013). Co-parenting over time: the incidence and characteristics of joint physical custody families in Flanders. *Demographic Research*, 28, 821-848.

Sodermans, A. K., Vanassche, S., & Matthijs, K. (2011). Gedeelde kinderen en plusouders: De verblijfsregeling en de gezinssituatie na scheiding. *Scheiding in Vlaanderen*, 135-151.

Statistics Belgium (2014). Population - Divorces en 2012 dossier. http://statbel.fgov.be/fr/modules/publications/statistiques/population/downloads/population_-_divorces_en_2012.jsp

Sodermans, A.K.; Matthijs, K. & Swicegood, G. (2013). Characteristics of joint physical custody families in Flanders. *Demographic Research*, 28(29), 821-848.

Sweeney, M. M. (2002). Remarriage and the Nature of Divorce Does it Matter Which Spouse Chose to Leave? *Journal of Family Issues*, 23(3), 410-440.

Symoens, S., Bastaits, K., Mortelmans, D., & Bracke, P. (2013). Breaking up, breaking hearts? Characteristics of the divorce process and well-being after divorce. *Journal of Divorce & Remarriage*, 54(3), 177-196.

Turunen, J. (2011). Entering a stepfamily: Children's experience of family reconstitution in Sweden 1970-2000. *Journal of Family Research/ Zeitschrift für Familienforschung*, 23(2), 154-172.

Vanassche, S. (2013). *Stepfamily configurations and trajectories following parental divorce. A quantitative study on stepfamily situations, stepfamily relationships and the wellbeing of children.* Dissertation, KU Leuven.

Wang, H., & Amato, P. R. (2000). Predictors of divorce adjustment: Stressors, resources, and definitions. *Journal of Marriage and Family*, 62(3), 655-668. DOI: 10.1111/j.1741-3737.2000.00655.x

Wooldridge, J. M. (2002). *Econometric analysis of cross section and panel data.* The MIT press.

Statistics Canada (2012). *Income in Canada 2010: Analysis.* Ottawa: Minister of Industry. Retrieved fromhttp://www.statcan.gc.ca/pub/75-202-x/2010000/analysis-analyses-eng.htm

# APPENDIX

Table 2: *Literature overview on the determinants of custody choice and repartnering*

| Determinants in empirical analyses | Influencing sole physical custody | | Influencing repartnering | | Selectivity of sole-custody mothers |
|---|---|---|---|---|---|
| Woman's educational attainment | - | Cancian & Meyer (1998);<br>Fox & Kelley (1995);<br>Gunnoe & Braver (2001);<br>Juby et al. (2005);<br>Nielsen (2011) | +<br><br><br>/ | de Graaf & Kalmijn (2003);<br>Dykstra & Poortman (2009);<br>Le Bourdais et al. (1995)<br>Ivanova et al. (2013);<br>Bauserman (2002);<br>Turunen (2011);<br>Wu & Schimmele (2005) | (Negative) |
| Woman's age at marital dissolution | + | Cancian & Meyer (1998);<br>Cancian et al. (2014);<br>Juby et al. (2005); | - | Beaujouan (2010, 2012);<br>Ermisch & Wright (1991);<br>Le Bourdais et al.(1995) | Negative |
| Number of children | +<br><br><br>/ | Fehlberg et al. (2011);<br>Bausermann, (2002);<br>Juby et al., (2005);<br>Cancian & Meyer (1998);<br>Sodermans et al. (2013) | / | Turunen (2011) | / |
| Age of youngest child | (+) | Cancian & Meyer (1998);<br>Cancian et al. (2014) | +<br><br><br><br><br><br>- | Fehlberg et al. (2011);<br>Bausermann (2002);<br>Le Bourdais et al. (1995);<br>Meggiolaro and Ongaro (2008);<br>Poortman (2007);<br>Wu & Schimmele (2005)<br>Turunen (2011) | Positive |
| Child is a boy | -<br><br><br>/ | Cancian & Meyer (1998);<br>Cancian et al. (2014);<br><br>Sodermans et al. (2013) | -<br><br><br>+ | Turunen (2011)<br><br><br>Lundberg & Rose (2003) | Positive/Negative |

*Note*: "+" positive effect; "-" negative effect; "/" no significant effect

Table A1: Effects of the policy reform on physical custody (N=1,222)

| | Probit: Sole physical custody Model C1 | | Probit: Sole physical custody Model C2 | |
|---|---|---|---|---|
| | B | ME | B | ME |
| Number of children form first marriage (ref= one child) | | | | |
| Two children | 0.03 | 0.01 | 0.03 | 0.01 |
| | (0.09) | (0.04) | (0.09) | (0.04) |
| Three or more children | 0.14 | 0.05 | 0.14 | 0.05 |
| | (0.12) | (0.05) | (0.12) | (0.05) |
| Only boys | -0.07 | -0.03 | -0.07 | -0.03 |
| | (0.09) | (0.03) | (0.09) | (0.03) |
| Age of youngest child at marital dissolution | 0.01 | 0.00 | 0.01 | 0.00 |
| | (0.01) | (0.01) | (0.01) | (0.01) |
| Mother's age at marital dissolution (-33) | -0.01 | -0.00 | -0.00 | -0.00 |
| | (0.01) | (0.00) | (0.01) | (0.00) |
| Highest educational attainment (ref=high) | 0.26** | 0.10** | 0.25** | 0.10** |
| Low | (0.11) | (0.04) | (0.11) | (0.04) |
| | 0.20** | 0.08** | 0.20** | 0.08** |
| Medium | (0.09) | (0.04) | (0.09) | (0.04) |
| Initiator of divorce (ref=both) | | | | |
| Mother | 0.28*** | 0.11*** | 0.28*** | 0.11*** |
| | (0.10) | (0.04) | (0.10) | (0.04) |
| Partner | -0.07 | -0.03 | -0.07 | -0.03 |
| | (0.10) | (0.04) | (0.10) | (0.04) |
| Divorced after 1995 | -0.26*** | -0.10*** | | |
| | (0.09) | (0.04) | | |
| Divorce year (ref=1996-2000) | | | | |
| 1985-1989 | | | 0.47*** | 0.18*** |
| | | | (0.16) | (0.06) |
| 1990-1994 | | | 0.20* | 0.08* |
| | | | (0.11) | (0.04) |
| 2001-2005 | | | 0.04 | 0.02 |
| | | | (0.09) | (0.03) |
| Constant | -0.09 | | -0.38*** | |
| | (0.15) | | (0.13) | |

*Note:* Data are from the "Divorce in Flanders" study, authors' calculations. Model C1 and Model C2 are ordinary probit models. Effects in beta-coefficients (B) and Marginal effects (ME), calculated as mean over the sample. Standard errors in parentheses.

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

Table A2: Effects of the policy reform on repartnering (N=1,222)

| | Probit: repartnering Model R1 | | Probit: repartnering Model R2 | | Probit: repartnering Model R3 | | Probit: repartnering Model R4 | |
|---|---|---|---|---|---|---|---|---|
| | B | ME | B | ME | B | ME | B | ME |
| Number of children form first marriage (ref= one child) | | | | | | | | |
| Two children | | | -0.13 | -0.05 | 0.02 | 0.01 | 0.01 | 0.00 |
| | | | (0.09) | (0.03) | (0.09) | (0.03) | (0.09) | (0.03) |
| Three or more children | | | -0.33*** | -0.12*** | -0.02 | -0.01 | -0.05 | -0.02 |
| | | | (0.12) | (0.04) | (0.13) | (0.05) | (0.13) | (0.05) |
| Only boys | | | 0.08 | 0.03 | 0.06 | 0.02 | 0.08 | 0.03 |
| | | | (0.09) | (0.03) | (0.09) | (0.03) | (0.09) | (0.03) |
| Age of youngest child at marital dissolution | | | -0.04*** | -0.01*** | 0.04*** | 0.02*** | 0.04*** | 0.02*** |
| | | | (0.01) | (0.00) | (0.01) | (0.01) | (0.01) | (0.01) |
| Mother's age at marital dissolution (-33) | | | | | -0.09*** | -0.03*** | -0.09*** | -0.03*** |
| | | | | | (0.01) | (0.00) | (0.01) | (0.00) |
| Highest educational attainment (ref=high) | | | | | | | | |
| Low | | | -0.02 | -0.01 | -0.22** | -0.08** | -0.25** | -0.09** |
| | | | (0.10) | (0.04) | (0.11) | (0.04) | (0.11) | (0.04) |
| Medium | | | 0.11 | 0.04 | -0.06 | -0.02 | -0.09 | -0.03 |
| | | | (0.09) | (0.03) | (0.09) | (0.03) | (0.09) | (0.03) |
| Initiator of divorce (ref=both) | | | | | | | | |
| Mother | | | -0.48*** | -0.18*** | -0.43*** | -0.16*** | -0.47*** | -0.17*** |
| | | | (0.10) | (0.04) | (0.10) | (0.04) | (0.10) | (0.04) |
| Partner | | | -0.18* | -0.07* | -0.14 | -0.05 | -0.13 | -0.05 |
| | | | (0.10) | (0.04) | (0.11) | (0.04) | (0.11) | (0.04) |
| Full-time custody (90-100%) | | | -0.36*** | -0.14*** | -0.39*** | -0.14*** | | |
| | | | (0.08) | (0.03) | (0.08) | (0.03) | | |
| Divorced after 1995 | -0.02 | -0.01 | 0.04 | 0.01 | 0.17* | 0.06* | 0.21** | 0.08** |
| | (0.09) | (0.04) | (0.09) | (0.04) | (0.10) | (0.03) | (0.10) | (0.03) |
| Constant | -0.03 | | 0.47*** | | -0.09 | | -0.26* | |
| | (0.08) | | (0.13) | | (0.16) | | (0.15) | |

*Note:* Data are from the "Divorce in Flanders" study, authors' calculations. Model R1, Model R2 and Model R3 are ordinary probit models. Effects in beta-coefficients (B) and Marginal effects (ME), calculated as mean over the sample. Standard errors in parentheses.

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

Figure A1a: Interaction of physical custody and internet use at time of separation

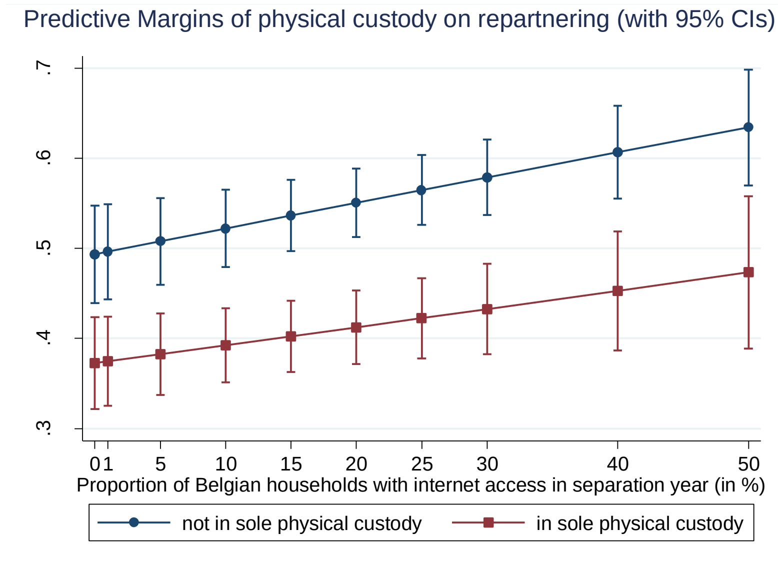
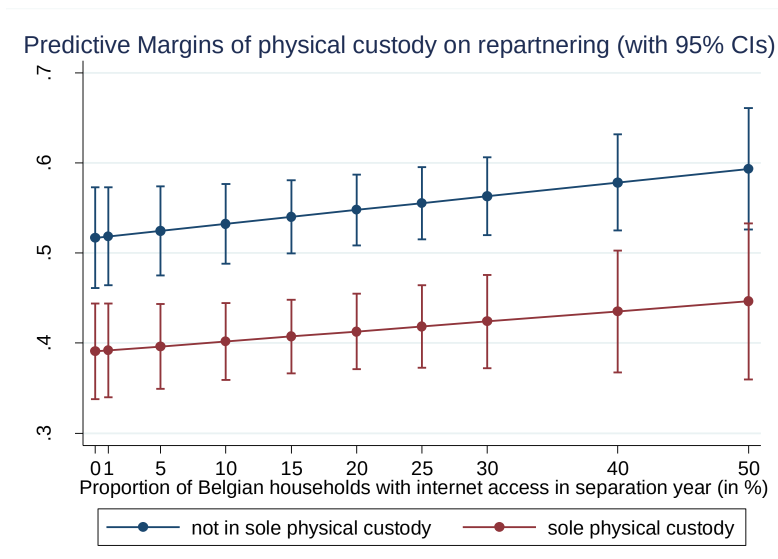



Figure A1b: Interaction of physical custody and internet use at time of separation, model disregarding age at marital separation




Note: Data are from the "Divorce in Flanders" study and the World Bank (internet access), authors' calculations. Predicted margins calculated based on coefficients of ordinary probit models on repartnering. Models of Figure A1a and Figure A1b include as control covariates: number, age, and gender of children, mother's education, initiator of divorce. Figure A1a includes also the mother's age at marital dissolution.