Union fertility and stability in different family settings

A multiprocess analysis of Britain

Alessandro Di Nallo

Université de Lausanne

I study the risks of a birth and separation in step-families and families with no prior children. I test whether a birth transition is determined by the 'parenthood motive' or by the 'commitment effect'. Also, I test whether common children stabilize a partnership and whether the pre-union children are associated with higher risk of dissolution. Elements of family complexity, such as presence, number and parentage of step-children, are considered. Using multiprocess models, I model partnership transitions with fertility, allowing for the correlation between the unobserved individual characteristics affecting each process. The analysis is based on Wave 1 of UKHLS of men and women aged 16-45. The findings indicate that the parenthood and the commitment motives influence the transitions to a birth, under different family configurations. Further, the risk of separation is reduced by shared children, while the children from prior unions does not generally increase the risk of dissolution.

1. Introduction

A wealth of studies on stepfamily fertility has examined fertility behavior following a union dissolution (Beaujouan & Solaz, 2012; Jefferies et al., 2000) and, more specifically, parity progression in stepfamilies (Henz, 2002; Li, 2006; Thomson et al., 2002; Thomson & Li, 2002; Vikat et al., 1999; Vikat, et al., 2004). To study the childbearing within these unions is relevant to understand the stepfamily dynamics as well as population fertility in a context of increasing partnership churning rates (M. M. Sweeney, 2010): stepfamily fertility might lead to fertility levels higher than would be registered in uninterrupted unions (Thomson et al., 2012; Winkler-Dworak et al., 2017).

This body of research has mainly concentrated on the existence of different factors motivating childbearing in second or higher order union as opposed to childbearing in general (childless couples), and specifically whether the presence of children from previous unions impacts on partners' risk of having additional children together. In spite of this extensive interest in the study of fertility in complex families, the evidence from previous studies remains rather mixed, due to the inclusion of different indicators of family arrangements and the exclusion of cohabitating stepfamilies and/or in higher order unions.

Further, previous research has devoted little attention to study the risk of union dissolution of stepfamilies as opposed to couples without stepchildren, and it has rarely taken men's fertility into account (with some exception such as (Ivanova, et al., 2014; Stewart, 2002). Also, the presence of past-union children of at least one partner introduces complexity in the relationship between the family members, might affect children's wellbeing (Fomby & Cherlin, 2007), and has been indicated among one of the drivers of union instability (Beaujouan, 2016; Teachman, 2008).

Childbearing and partnership transitions are intertwined processes whose inputs can be the output in another process (Lillard & Waite, 1993; Steele et al., 2005). Fertility in a couple is determined by many factors, including partners' parental status and commitment to their current union. Having children could either bolster partners' mutual attachment or introduce some complication to partners' dynamics, accelerating the process of dissolution. Failing to capture the underlying relations across these processes would not accurately estimate the influence of family configurations on the couple's childbearing and dissolution risks. Thus, in this article I use multi-process models, which simultaneously account for union formation, fertility and partnership dissolution, to estimate the influence of different family configuration of step-families and nonstepfamilies on these life-course transitions.

2. Background

2.1. Childbearing progression in couples: parenthood and commitment hypotheses

Childbearing progression in childless couples could diverge from that of stepfamilies, in which at least one partner brings children to the family. In the former, the number of children depends on partners' desired fertility; in the latter, the fertility intentions are also conditioned by the children born in previous unions. For instance, partners who want two children altogether, but had only one in total, would have a second in their stepfamily; but those couples in which one partner brought two children from a dissolved partnership, would they go on to have their first shared child (and third total)?

Stepfamilies' fertility progression is complicated by the mismatch between individual and couple-level fertility. This misalignment prompts the question: does partners' previous fertility influence childbearing in current union or does the new partnership produce new motives for having children, regardless of previous fertility? In the first scenario, it is the achievement of parenthood *per se* that justifies the first shared birth, if either one partner or both partners haven't had any children yet. This 'parenthood hypothesis' implicitly implies that, if both partners in a stepfamily are already parents, the odds of a first shared child are lower than those of a couple with combined parity of at most one child. In the second scenario, it is the 'commitment hypothesis' that prevails: having a child – regardless of the children born in previous relationships – cements a union and expresses partners' pledge to each other.

Previous research has reported mixed findings, which do not clearly support one hypothesis over the other. The studies backing the parenthood hypothesis found that the birth risk is lower or at least no higher in compared to childless couples, although the effects differs by children's parentage, number and residence. Stewart (2002) in the USA and (Beaujouan, 2012) in France highlighted that couples are least likely to have a child after repartnering if both partners already have children, while Ivanova et al. (2013) in the Netherlands proved that the presence of pre-union children reduces the odds of a union-specific birth only for men but not for women. Wineberg (1990) showed that remarried women with two or more children have less chances of a shared birth in a new union compared to counterparts with one child. Likewise, (Buber & Furnkranz-Prskawetz, 2000) demonstrated for Austria that women have lower odds of a child in a second union if either partner has more than two children. In other words, in one-child stepfamilies, a first shared birth would provide a (half-)sibling to the pre-union child, while in two-children stepfamilies this 'sibling motive' would be attenuated. This would explain why some studies found a different risk of a first shared birth between couples with one vs. two or more pre-union children. However, Vikat et al. (2004) and Buber & Prskawetz (2000), both in Austria, argued that it is the presence of co-resident pre-union children influencing the odds of a shared birth, irrespective of their number. According to this argument, the intention of a couple to have a (further) child would be negatively associated with the number of children for whom a couple is responsible, either through coresidence or financial support (Hohmann-Marriott, 2015). The argument of (step)children's residence has not gained uncontroversial stand. Vikat and colleagues (1999; 2004) found little or no difference between the effects of coresident and nonresident children on stepfamily fertility in Sweden and Finland.

In contrast to these findings, other studies hold the birth of a shared child as a symbol of union commitment, also for partners who had already children from prior unions, in spite of the greater childrearing burden stepfamily couples may incur (Vikat et al., 2004). This commitment hypothesis was first formalized by Griffith et al. (1985), who found empirical evidence that a woman's number of children in previous unions does not have a significant effect on her fertility in a new one. A number of studies have lent support to this theory finding that births appear relatively unaffected by the number of past-union children Sweden (Vikat et al., 1999; Jefferies, et al., 2000; Meggiolaro & Ongaro, 2008; Thomson et al., 2014). Interestingly, Ivanova et al. (2013) confirmed the commitment hypothesis, but only for women and not for men, paving the way to new research investigating possible gender-specific patterns of childbearing in different family arrangements.

While the value of the first common child is a symbol of achieved parenthood or mutual commitment, or both, the value of a second child is primarily characterized as a full sibling to the first child. Existing evidence for this value is more mixed especially when it comes to stepfamilies. Although couples with no prior-union children seem to interiorize a two-child norm (Joshua, Lutz, & Testa, 2003), parents with past-union children and one shared child seem to show lower intentions for another birth compared to parents without shared children (Thomson & Li, 2002; Stewart, 2002). Vikat et al. (1999) hypothesized that a half sibling might act as a full sibling if stepfamilies with one shared and one prior-union child were less likely to have a further child compared to couples with a common child in the current union and no past-union children. However, they did not find congruent evidence as other factors such as the number of prior union children might exert a stronger influence. An intriguing finding about the value of the sibling value was provided by (Holland & Thomson, 2011)who linked the sibling value of step-sibling to the age distance between the shared firstborn and the youngest pre-union child. In this view, the stepsibling might act as a *de facto* full sibling if the birth spacing between the youngest pre-union child and the first shared child is fairly short (Bernstein, 1997; Ivanova et al., 2014)

2.2. Union dissolution

An extensive body of research investigated the effect of fertility on marital stability. Most evidence, including the UK, concluded that children have a protective effect on marriage, especially when they are young (Bellido, Molina, Solaz, & Stancanelli, 2016; Berrington & Diamond, 1999; Lillard & Waite, 1993; Waite & Lillard, 1991). These findings mirror the links between childbearing and divorce that have been suggested by the theory. Children are a source of union-specific capital (Becker, Landes, & Michael, 1977), as parents' utility is higher if they live together than if they do not. Also, children are a symbol of partners' commitment, which cements the union of their parents (Brines & Joyner, 1999).

However, there is mixed empirical evidence on the effects of fertility on dissolution in higher-order unions and in non-marital settings (Lyngstad & Jalovaara, 2010), and in different family configurations with regard to number, age and children's residence (Bellido et al., 2016;

Coppola & Di Cesare, 2008; Ermisch J. and Francesconi, 2000; Lillard & Waite, 1993; Manlove et al., 2012; Steele et al., 2005; Svarer & Verner, 2008). This may reflect either genuine cross-country differences, diverse specifications of family configurations (Steele et al., 2006), or factors that are supposed to affect stability in higher order unions, such as unmarried cohabitation and the presence of stepchildren (Teachman, 2008).

Cohabiting parents tend have lower risk of union dissolution than childless cohabitors, although this association is less consistent for married childless couples vs. married couples with children (Andersson, 2002; Heuveline, Timberlake, & Furstenberg, 2003; Poortman & Lyngstad, 2007; Steele et al., 2005; Wu & Musick, 2008). Further, stepfamilies tend to have a higher dissolution risk than couples without pre-union children (Erlangsen, et al., 2001; Teachman 2008; Beaujouan, 2016) and, among stepfamilies, stepmother families experience a lower risk of separation than stepfather families (Heintz-Martin, et al., 2011; Teachman, 2008).

It is not clear whether this finding is explained by the micro-dynamics of family functioning. Research in psychology has not reached uncontroversial evidence about the differential levels of quality in either family step-family arrangement. In fact, there is no evidence that children systematically report better quality relationships with stepfathers or stepmothers (King, 2006, 2007; King, Thorsen, & Amato, 2014; Vogt Yuan & Hamilton, 2006). Stepchildren are inclined either to accumulate parental figures, namely have close relationships with both stepfathers and fathers and with both stepmothers and mothers (King, 2006; White & Gilbreth, 2001), or replace nonresidential biological parents with step-parents (Ganong, et al., 2011). A key role to facilitate the development of positive ties between children and their new stepparent is played by the cooperation of the biological non-resident parent (Amato & Booth, 1996; Cummings & Davies, 1994) and by the quality of the relationship with the resident parent (Carlson et al., 2008; Weaver & Coleman, 2010).

2.3. Selectivity and interplay of childbearing and partnership processes

To the extent that lower levels of relationship quality keep childless couples from having children, and cohabiters from marrying, these results can be explained in part by self-selection of more stable unions into parenthood, into first-order unions, and out of step-family arrangements. For instance, the decision to have children is selective of the most stable couples (Lillard & Waite, 1993). Conversely, childless partners might be largely selected from persons with poorer relationship skills and/or lesser commitment to family life, which might result in higher predisposition for partnership dissolution (Lillard et al., 1995; Steele et al., 2005). These couples might have hastily formed a relationship in response to an emergency and, thus, have more likely mismatched (Lichter, et al., 2016). Likewise, stepfamilies may have different a risk of separation compared to couples without stepchildren due to unobserved characteristics. For instance, having children in prior relationship could reduce the risk of separation of the current union. Individuals with dependent children, who experience a relationship breakdown and have suffered a trauma, might be warier of entering a new union and be more selective in their choice of a new partner (Steele et al., 2005). This form of selection would lead to a negative effect of stepchildren on the risk of separation.

A few studies have tried to address the mechanisms of selection of individuals into specific partnerships patterns using sophisticated statistical techniques. Using multi-process methodology, Lillard & Waite (1993) found that the first shared child lowers the risk of divorce whereas subsequent children have opposite effect. Coppola & Di Cesare (2008) showed that second or third birth stabilize couples in Italy and Spain, while Svarer & Verner (2008) found opposite results in Denmark. In the United States, Vuri, (2002) and Bellido et al. (2016) found a deterrent effect of children on marital disruption, especially when they are younger. Using British data, Steele et al. (2005) found that women's pre-school children lower the risk of dissolution, whereas Aassve et al. (2006) showed that first and second shared child have a stabilizing influence on unions for women and men. However, none of these studies took into account the influence of stepfamily configurations on the risk of union dissolution. Henz & Thomson (2005) analyzed simultaneously the processes of childbearing and union dissolution of stepfamilies and found those without shared children more at risk of separating as opposed to stepfamilies with at least one common child. Nevertheless, their analysis ruled out non-stepfamilies from the analysis. Therefore, my study will try to fill this gap by accounting for the simultaneous processes of childbearing and partnership transition of individuals entering in stepfamilies and non-stepfamily arrangements.

3. Mechanisms and hypotheses

In this article, I identify which family arrangements have a stronger association to parity progression and partnership dissolution within stepfamilies and non-stepfamilies configurations.

3.1. Parity progression

A1. I test the *parenthood hypothesis* by examining the progression to a shared birth for childless couples without pre-union children (childless non-stepfamilies) and for stepfamilies without shared children in which only one partner has children. If the latter are less likely to have a shared birth, I can conclude that the desire to achieve the parenthood role is a valid mechanism that justifies childbearing in stepfamilies.

A2. I test whether couples in which only one of the partners is a parent have similar odds of a common birth in the current union as the couples in which both partners bring children to the family. If this hypothesis is confirmed, the value of a shared child implies that having a birth is a deliberate behavior of commitment, regardless of partners' parental status (Griffith, et al., 1985).

A3. I test the influence of a family configuration on the couple's transition to a second shared child. I hypothesize that (co-resident) step-children do not have any positive influence on a couple's first shared birth and have a negative effect on the second birth. In non-stepfamilies, the value of second born is to provide the firstborn with a full sibling (Bernstein, 1997). In stepfamilies with a shared child, the second common child would not serve either as the expression of partners' commitment to the union (*commitment effect*) or as the achievement of parenthood for one of the partners (*parenthood effect*). In addition, when the couple has co-resident pre-union children, a second shared birth might pose additional issues on the sustainability of the family budget and on the management of family life. Therefore, I argue that co-resident step-sibling(s), especially if relatively young (Bernstein, 1997; Holland & Thomson, 2011), might act as a *full sibling(s)* and, thus, deter stepparents from having a second shared birth.

3.2. Dissolution

B1. I assess whether having common children stabilizes a partnership. Although theoretical arguments prospect lower chances of dissolution for couples who have children (Gary S. Becker, 1981; Thornton, 1977; Coppola & Di Cesare, 2008), empirical evidence reported mixed results. On the one hand, the presence of young children raises the costs of a separation and lowers the risk of separation. A union dissolution with toddlers and/or infants might force parents either to raise the children alone or reduce the contact with them. Further, awareness of the severe impact of a separation when children are young (de Graaf & Kalmijn, 2003), and the normative and social pressure against dissolution (Coleman, 1988) might further raise separation costs for couples with young children.

On the other hand, (shared) children could hinder partnership stability. They might trigger a conflict within the couple by altering parents' pre-birth habits, such as time and money allocation (Del Bono, Ermisch, & Francesconi, 2012) as they impose additional obligations and reduce their parents' romantic time (Kluwer, 2010).

B2. I test whether the influence of pre-union children increases the risk of union dissolution.

Children born in previous partnerships can hinder a partnership stability in multiple ways. To begin with, stepchildren are not a 'union-specific capital' in the new relationship (Becker et al., 1977) as shared children are, and can have a different value for their step-parents, who may invest less resources and time than do biological parents (Aquilino, 2005; Hofferth & Anderson, 2003; Teachman, 2008). In fact, prior research finds that, at least among fathers, involvement with prior children is a primary source of mistrust among cohabiting parents (Reed, 2006). This is especially the case if step-parents' obligations span multiple households (e.g., Weaver & Coleman, 2005).

Stepchildren could be a source of conflict in the new family as they could manifest their adjustment problems after experiencing the emotional and psychological distress of parents' separation (Hetherington & Kelley, 2002). Also, the lack of institutionalization in a newly formed step-family implies that parents and step-parents need to negotiate complex relationships with their partners, their stepchildren, and even their stepchildren's biological parents (King et al., 2014; Manlove et al., 2012; Marsiglio, 2004). This might lead to confusion and stress that heightens partnership conflict and the risk of union disruption (Coleman et al., 2000). However, to the extent that stepchildren constitute a shared interest to the couple (in particular to the step-parent who intends to legitimize her role through active childrearing), their presence might not hamper the stability of a union. A shared child might also stabilize a stepfamily where children of both partners live, in spite of adding further complexity to the family. The newborn could strengthen the emotional bond between the partners and legitimize their parental role in the eyes of their stepchildren (Juby, Marcil-Gratton, & Le Bourdais, 2001; King et al., 2014)

Many of the previous hypotheses marginally incorporate the influence of children's age, both for the risk of (further) childbearing and union dissolution. For instance, Holland & Thomson (2011) showed that the decision of birth spacing in step-families depend also on the age of the youngest child: the risks of a second and third birth (in terms of combined parity) in a stepfamily drop substantially if the youngest child is at least 10 years old because youngest child would be too large for them to be thought of together as (half-)siblings; the risk of a second and third birth is also relatively low when the youngest child is 6 or younger, possibly because the prospect of two very young children might too demanding in terms of energy. When it comes to partnership dissolution, most research agrees that young children delay separation (Waite & Lillard, 1991; Vuri, 2002; Steele et al., 2005). This 'protective' effect of children might last until preschool age (e.g., Steele et al., 2005) or longer (Waite & Lillard, 1991). In the current study, the age of the (youngest) child is not considered in order not to increase the model complexity but future versions the analysis will seek to include this variable.

4. Data

4.1. The analytical sample

The analysis uses the first wave of the UK Household Longitudinal Study (UKHLS), 'Understanding Society', a panel study of over 30,000 households in the UK (Mcfall, 2013). Respondents were asked at the age of the first interview (between 2009 and 2010) to recall the start and end dates of all the unions (marriages or cohabiting unions) they had experienced and had lasted for at least one month. Also, they were asked to record the history of all the dates of birth of all the children they had given birth to, adopted and/or raised.

The retrospective data are matched to construct detailed union and fertility histories from age 15 for all respondents born between 1940 and 1980. This age is motivated by the early pathways of union formation and childbearing during teenage and the fact that marriage is legal from age 16. For each individual, I created three event histories, for union formation and dissolution and childbirth in a union. The censoring age for childbearing for women is set at 45 and 55 years for women and men, respectively, while the highest censoring age for the union processes is 60 years since the oldest cohort of the analytical sample was born in 1940. The final sample consists of 12751 women and 9402 men.

4.2. Partnership histories

In the union formation histories, a variable indicates for each month whether the individual entered a union. Likewise, in the union dissolution histories, a variable indicates whether the union was ongoing or ended. It is important to stress that I define a partnership as a co-residential union, regardless of the marital status. The proportion of cohabitations out of the total unions raised from 10% in the late 1980s to almost 30% in the late 2000s in Britain (Sanchez Gassen & Perelli-Harris, 2015). Although still far from being normative, cohabitation has gradually gained the role of "testing-ground" for stable unions and as an alternative to marriage (e.g., Murphy, 2000; Perelli-Harris, 2014), with 30% of first births occurring in non-marital unions in the late 2000s compared to 4% in the early 1980s (Perelli-Harris et al., 2012). Therefore, the persons living in a couple as a cohabiter or married are considered "in a union", while all the other individuals, including those "living apart together", are not considered in a union. Strictly, a new episode of union might immediately follow a separation, leading to a transition to co-residential relationship rather than to single state. However, in keeping with previous research, I treat such transitions as the same as periods of singlehood that last one month and follow a dissolution.

4.3. Birth histories

The birth history is constructed from the household record of the respondents, who retrospectively mentioned the month and year of birth of all the children they had given birth to. Following Steele et al., (2005), I dropped women and men with adopted children because it is likely that their conception intervals were affected by the adoption, and it is possible that they or one of their partners was infertile. For each child, UKHLS collected the date of birth, as well as the date of arrival of adopted and step-children in the respondent's child; the date at which the child left the household was also recorded. The respondent was not asked to identify the partner of each child. Therefore, in order to identify the (time-varying) parental status of the respondent, I reconstructed and synchronized the fertility and the partnership histories. A birth interval may begin at union formation (for the first shared birth of a couple) or after a birth within a union (all other intervals).

Births are collected in a six-month dataset. In line with other studies on fertility and unions (Heintz-Martin et al., 2011; Guzzo, 2017; Thomson et al., 2014), children born up to six months prior to the beginning of a union belong the union and are the biological children of the current partner. Conversely, if the child is born over six months prior to the union, the birth is considered born out of the union. This assumption is supported by recent evidence showing that the odds of union formation after a nonunion birth are quite low in the first six months and that the odds of forming a union with a new partner, rather than with the biological parent, increase over time (Guzzo, 2016). One of the advantages of UKHLS is that in each co-residential union, the respondent is asked whether the partner had any co-residential children from prior union and, for each of them, age and permanence in the household were available. This could represent a more serious problem for older people as it has been shown that remote events mainly suffer from recall bias in retrospective survey. This problem was limited for all the cases in which the respondent was in a union at the time of the interview. The UKHLS design allowed me to compare the respondent fertility history with the current partner's: every childbirth date being omitted was amended with the one reported by the partner and every mismatch between partners' reported dates was corrected in favour of the female's version. These solutions could not be applied for all prior unions – for which no partner's interview was available. Strategies to cope with measurement error could be made but these were not implemented given the complexity of the analysis (See Chesher, 2001; Chesher, et al., 2002; Aassve et al., 2006) for further details on the effect of measurement error in duration models).

The information collected in conjugal and parental histories allowed me to identify stepfamily episodes. The moment in which a respondent starts living with a partner and at least one child who is not a biological child marks the beginning of the stepfamily. For each child living in the respondent's house – whether born in the respondent's prior unions or stepchild– UKHLS traced the permanence in the household. It was therefore possible to determine the period in which a family was a step-mother or a step-father family or a complex stepfamily (with both partners bringing children to the union). The respondent was asked the date each (own and partner's) co-resident child left home, which enabled me to create time-varying counts of the number of children living with the respondent, distinguishing between shared children with the current partner and those born in a prior union, and those who left the household of the respondents.

The data did not include information on the number of children born to partners in previous unions. For example, unlike Thomson et al., (2002) who controlled for the combined number of children (respondent's, partner's, shared) at the beginning of a union or birth spell, I could identify only partner's children who spend a spell in the respondent's household: 'co-resident' step-children only. Therefore, I was not able to accurately specify theoretically relevant configurations of step-families, such as stepchildren of the respondent, of the partner, of both). Lack of information on partners' children means that some couples identified as not having step-children did have (non-resident) step-children through respondent's partner, and some of the families had more step-children than I was able to identify through the respondent. Nevertheless, repartnering is *not* random with respect to partners' children (see Chapter 3): childless individuals are more likely to repartner with other childless, and parents are more likely to repartner with other partners. Therefore, I would underestimate the negative effect of combined parity on the risk of stepfamily births because the larger combined parity of partners in a stepfamily reduces the potential for further births (Buber & Prskawetz, 2000; Holland & Thomson, 2011).

4.4. Explanatory variables

The richness of the data allowed me to elaborate different specifications of the stepfamily status (i.e., whether the union is a stepfamily) and configuration (whose children are whose). The family composition at the beginning of the episode is measured in three ways. The first specification distinguishes individuals' partnership status according to the stepfamily status and the number of shared children in the ongoing partnership (0 = 'the union is *not* a stepfamily, no shared children'; 1 = 'the union is *not* a stepfamily, 1 shared child'; 2 = 'the union is *not* a stepfamily, 2+ children'; 3 = 'the union is a stepfamily, no shared children'; 4 = 'the union is a stepfamily, 1 shared child'; 5 = 'the union is a stepfamily, 2+ children').

Table 1A. Family configuration, by step-family status and number of shared children

	Stepfamily or non-stepfamily	Number of shared children
0	Not a stepfamily	No shared children
1	Not a stepfamily	1 shared child
2	Not a stepfamily	2+ shared children
3	A step-family	No shared children
4	A step-family	1 shared child
5	A step-family	2+ shared children

The second reflects also the number of pre-union children in the family: (0 = 'the union is *not* a stepfamily, no shared children'; 1 = 'the union is *not* a stepfamily, 1 shared child'; 2 = 'the union is *not* a stepfamily, 2+ children'; 3 = 'the union is a stepfamily with one step-child, no shared children'; 4 = 'the union is a stepfamily with one step-child, 1 shared child'; 5 = 'the union is a stepfamily with 1 step-child, 2+ children'; 6 = 'the union is a stepfamily with 2+ step-children, no shared children'; 7 = 'the union is a stepfamily with two or more step-children, 1 shared child'; 8 = 'the union is a stepfamily with two or more step-children, 2+ children').

Table 1B. Family configuration, by step-family status, number of step-children and number of shared children

	Stepfamily or non-stepfamily	Number of shared children
0	Not a stepfamily	No shared children

1	Not a stepfamily	1 shared child
2	Not a stepfamily	2+ shared children
3	A step-family, 1 step-child	No shared children
4	A step-family, 1 step-child	1 shared child
5	A step-family, 1 step-child	2+ shared children
6	A step-family, 2+ step-children	No shared children
7	A step-family, 2+ step-children	1 shared child
8	A step-family, 2+ step-children	2+ shared children

The third specification accounts for the parentage of pre-union children: (0 = 'the union is *not* a stepfamily, no shared children'; 1 = 'the union is *not* a stepfamily, 1 shared child'; 2 = 'the union is *not* a stepfamily, 2+ children'; 3 = 'the union is a stepfamily with respondent's step-children (only), no shared children'; 4 = 'the union is a stepfamily with respondent's children only, no shared children'; 5 = 'the union is a stepfamily with partner's children only, no shared children'; 6 = 'the union is a stepfamily with children of both partners, no shared children'; 7 = 'the union is a stepfamily with respondent's children only, 1 shared child'; 8 = 'the union is a stepfamily with partner's children only, 1 shared child'; 9 = 'the union is a stepfamily with children of both partners, 1 shared child'; 10 = 'the union is a stepfamily (residual), 2+ shared children').

Table 1C. Family configuration, by step-family status, parentage of step-children and number of shared children

	Stepfamily or non-stepfamily	**Number of shared children**
0	Not a stepfamily	No shared children
1	Not a stepfamily	1 shared child
2	Not a stepfamily	2+ shared children
3	A step-family, respondent's children	No shared children
4	A step-family, respondent's children	1 shared child
5	A step-family, partner's children	No shared children
6	A step-family, partner's children	1 shared child
7	A step-family, children of both	No shared children
8	A step-family, children of both	1 shared child
9	A step-family, residual	2+ shared children

4.5. Controls

I adjust for the effects of a range of other factors that have been previously linked to partnership transitions. I control for characteristics of the current cohabiting partnership – such as duration of the current union or singlehood (in case of the process of union formation) and respondent's age –and previous partnership experiences, including previous marriage or cohabitation. A history of previous unions might be conducive of some personal attributes linked to relationship dynamics that predict partnership dissolution, particularly when children are involved (Steele et al., 2005).

The durations of partnerships and singlehood spells are derived from the partnership histories. The length of the first episode of singlehood is calculated from age 15. In general, the length of relationships is determined from the time the respondents moved in with the partner. The duration of previous marriage episodes (in the partnership formation process) includes also the time spent in premarital cohabitation. The duration of the current union (in childbearing and dissolution processes) and length of the singlehood (in formation process) are the explanatory variables of the baseline hazard function which are proxied by duration and duration-squared terms in the model. The functional form for each of the three transitions is quadratic as the hazard function was found similar to an inverse U-shaped relationship with duration. In keeping with Aassve et al., 2006, I fitted distinct base-line hazard for formation of first and subsequent partnerships and time-varying duration variables. Age is treated as a time-varying covariate, as either linear or quadratic term, while another variable controls for a given time period and is expressed by a continuous variable between 1955 and 2010 and its squared term to control for nonlinear effect.

I consider a range of background characteristics that were found to be relevant predictors in earlier studies of partnership transitions and childbearing (Steele et al., 2005; Ermisch & Francesconi, 2000). These characteristics are: the highest level of educational attainment (treated as time varying), measures of the respondent's family background, and the region of origin. Education is operationalized as a dichotomous variable ~~is~~ equal to 1 when the respondent is enrolled in full-time education, whereas another time-varying variable captures the highest educational qualification attained in five categories: (1) 'lower education': CSE and other school certificates; (2) ‘GCSE and equivalents’: standard/ordinary (O) grade, lower (in Scotland), GCSE/O level; (3) ‘A-level and equivalents’: certificate of sixth year studies, higher grade, advanced higher (in Scotland), AS level, International Baccalaureate, A level, other schools, leaving exam certificate and other schools; (4) ‘Other higher qualification’: nursing or other medical qualifications, teaching qualification (except PGCE), diploma in higher education; (5) ‘Degree’: first degree level qualification including Postgraduate education, which is the omitted category for the set of dummy variables; (6) none of the above.

Social class is found associated with the risk of union formation (Barber, 2001), break-up (Steele, et al., 2006) and childbearing (Carlson et al., 2013; Hobcraft, 2008; for a complete review, see Chapter 2). It is based on the highest occupation of the respondent’s parents as coded to the UK NS-SeC Class classification (Office for National Statistics 2005), at the age of 14. The resulting indicator is coded as an 8-category variable: (1) ‘Higher managers and professionals’, (2) ‘Lower managers and professionals’, (3) ‘Intermediate occupations’, (4) ‘Small employers and own account workers’, (5) ‘Lower supervisory and technical occupations’, (6) ‘Semi-routine occupations’, (7) ‘Routine occupations’, (8) ‘Unemployed’. I created an extra category ‘missing’ for those respondents with incomplete information about their parents’ social class. Although this approach may introduce bias into the estimates of the transitions, this bias should be small if these background characteristics are weakly correlated with the outcomes at issue, as in this case. The family stability in childhood is significantly associated to later establishment a partnership (Aassve et al., 2006; Steele et al., 2006), risk of family dissolution (Steele et al., 2006) and risk of childbearing (Liefbroer & Elzinga, 2012; Fomby & Bosick, 2013). The indicator of family disruption by the age of 16 reflects parents’ divorce and any other alternative arrangement in which the parental or maternal figure was not one of the natural parents. Ethnicity is coded as a five-category variable with *White*, *Black African and Caribbean*, *Indian/Pakistani/Bangladeshi*, *Other Asian*, and *Other*. Finally, region of residence at birth – initially represented by 12 categories, is grouped into five categories following Steele et al. (2005). Descriptive information on the outcome variables and background characteristics are displayed in Table 2.

Table 2. Descriptive statistics for the partnership and background characteristics included in the final models.

Variables	*Women*	*Men*
	N	
Number of individuals	12751	9402
Number of singlehood episodes	23521	17865
Number of partnerships episodes	21148	16150
	Percent (%)[a]	
Age at union		
< 20 years	5,0	2,4
20-24	18,7	13,1
25-29	23,9	23,6
30-34	20,3	22,7
35-39	14,9	17,3
40+	17,2	20,9
Number of previous unions		
0	74,9	73,5
1	20,2	19,7
2+	4,9	6,8
Number of children in prior unions		
0	81,5	85,1
1	8,3	7,2
2+	10,2	7,7
Number of children in current unions		
0	38,2	56,8
1	20,9	18,7
2+	40,9	24,5
Family arrangement		
No stepfamily	80,1	79,9
Stepfamily	19,9	20,1
Parentage of step-children		
No stepfamily	80,1	79,9
Respondents' children	17,2	11,8
Partner's children	1,4	5,2
Children of both	1,4	3,0
Number of step-children		

0	80,1	79,9
1	8,4	8,9
2+	11,5	11,1
Number of shared children by family status		
No stepfamily, 0 shared	38,4	43,5
No stepfamily, 1 shared	16,1	14,7
No stepfamily, 2+ shared	26,1	22,0
Stepfamily, 0 shared	13,1	13,6
Stepfamily, 1 shared	4,1	4,2
Stepfamily, 2+ shared	2,2	2,1
Cohort of birth		
1940-49	7,0	7,4
1950-59	16,0	16,1
1960-69	37,9	38,5
1970-80	39,1	38,0
Parents' higher social class		
Higher managers & professionals	14,5	14,9
Lower managers & professionals	10,5	10,7
Intermediate occupations	11,2	11,3
Small employers and self-employed	9,1	9,4
Lower supervisory and technical occupations	15,5	15,7
Semi-routine occupations	16,4	16,0
Routine occupations	9,9	9,8
Unemployed	9,1	8,9
Missing	3,9	3,2
Parents' higher education		
Degree	9,9	9,6
Some qualification	22,4	20,5
Left with some qualification	20,3	21,7
Left with no qualification	23,7	21,5
No school	1,7	2,0
Missing	22,1	24,8
Own education		
Degree	24,9	28,8
Other higher	14,3	11,8
A level etc	17,3	19,7
GCSE etc	22,6	19,1
Other qual	9,2	10,6
No qualification	11,6	9,9
Parents separated at 14	4,1	3,8

Mother is absent/dead at 14	2,2	2,1
Father is absent/dead at 14	7,9	7,2
Number of siblings		
0	11,7	11,7
1	28,9	28,6
2 or more	59,4	59,7
Ethnicity		
European	80,8	78,6
African - Caribbean	5,7	5,3
Indian, Pakistani, Bangladeshi	9,5	12,1
Other Asian	2,2	2,1
Other	1,7	1,8

[a] It calculated as percentage of episodes*individuals in the models of partnership dissolution and childbearing

5. Methods

5.1. Joint modelling approach

Besides estimating the processes of partnership formation, childbearing and partnership dissolution separately, I will estimate them simultaneously to account for factors that otherwise would produce biased results. There are several reasons why acknowledging the interplay of these processes in a 'multi-process' approach performs better than a traditional method.

In first place, the decision to end a union or to have a (further) child with a partner can be driven by individuals' unobserved traits along with the observed characteristics included in the models (e.g., Steele et al., 2005). Therefore, the processes of partnership formation, dissolution, and childbearing could be jointly influenced by 'hidden' time-invariants traits, such as family values, attitudes towards relationship hopping, or propensity to betray the partner. For instance, individuals who have children in multiple partnerships might have different personal characteristics with respect to those who have never had any child in each relationship they entered, when it comes to family values (Guzzo, 2016). Likewise, people in second or higher order unions are more prone to another break-up because they more at risk of dissolving unions compared to "average" individuals (or partners in their first union), hence they are *selected* (Licther & Quian, 2008). Therefore, unobserved time-invariant characteristics affect the probability of event occurrence and cause the durations of repeated events in the same process to be correlated. A good wealth of previous studies (Aassve et al., 2006; Steele et al., 2005) stress the importance to control for unobserved heterogeneity when modelling repeated events. If this selection on unobserved characteristics was not taken into account, the influence of a prior event (such childbearing or dissolution) on a subsequent transition would result biased by the disproportionate presence of individuals whose 'unobserved propensities' would put the at increased risk. It is the case, for instance, of parents in newly established unions. In contrast to childless people, they might be either warier of experiencing, together with their children, another traumatic separation, or less concerned with the negative consequence of a separation and, hence, more likely to go through it.

Another form of selection arises when unobserved characteristics in one process are correlated with those of related events, such as fertility and union dissolution. For instance, fertility in a previous union might have an influence not only on fertility in successive unions but also on other events, such as the dissolution or the formation of co-residential unions. Prior evidence highlighted that individuals with above-average risk of having children also display higher propensity to form further unions (Rutigliano & Esping-Andersen, 2017). Therefore, to examine the influence of fertility (or union formation or dissolution) on other processes, it is necessary to jointly model the three processes at stake and allow for the residuals in the equations to be correlated among the transitions. If did not acknowledge that key explanatory variables in my model, such as the presence of stepchildren in the household, are not independent of the residuals in the equations of union formation and dissolution, the estimates of those variables on the partnership outcomes would result biased again.

In this article, I adjust for such selection effects by estimating simultaneously the processes of partnership formation, childbearing and partnership dissolution, by introducing correlation terms between the residual components of each process in the system. Childbearing and family formation and dissolution are specified in separate equations but are estimated in a joint maximum likelihood procedure. This allows me to account for the potential endogeneity of each transition with respect to the others. Although the outcomes of the model partnership formation are not relevant for the research questions, the inclusion of this process is motivated by the endogeneity of the other two equations with respect to it (Henz, 2002; Aassve et al., 2006). Therefore, the 'partnership formation' equation controls for the selection mechanisms into the patterns of childbearing (in a union) and partnership dissolution. Further, I treat all other variables as exogenous. This assumption may be questionable for outcomes of processes which are contemporaneous to partnership and 'within-union' childbearing, such as the non-union childbearing and educational status. Although it is technically possible to extend the model to allow for the determination of the other transitions, this would severely increase the complexity and the elaboration times.

The approach of joint modelling partnership and fertility histories has an additional advantage. The estimate of the residual correlations across the processes provides insights into the latent characteristics of the individuals. For instance, if the unobserved heterogeneity terms between fertility and separation are positively correlated, I can conclude that the unobserved traits that tend to increase the risk of union dissolution are responsible for a lower risk of birth. Steele et al. (2005), using data from the NCDS, found a positive correlation between the residuals of the processes of fertility and partnership dissolution, which means that individuals with an above average risk of having children might have a latent propensity to union dissolution. An underlying correlation also exists between the processes of union formation and childbearing, as the presence of women with strong latent desire to be in a partnership who also have a tendency towards having children (Rutigliano & Esping-Andersen (2017).

5.2. Definition of state transitions and episodes

I have argued that union and fertility decisions are simultaneously influenced by some common determinants and influence each other directly. Each process is represented by a discrete-time hazard event history equation that measures the duration of the exposition to the risk of (1) establishing a new union, (2) having a (further) children within a union, (3) dissolving a union. Such durations between events of the same individual are correlated because of the presence of unobserved individual-specific characteristics that influence that occurrence of each event. Further, the three equations are estimated simultaneously in a system (hence the definition of multi-process model) to explicitly control for their mutual effect. As I assume that individuals can experience these transitions multiple times during their life time, repeated events are modelled with a two-level hierarchical structure with events nested within individuals. Following in the vein as Steele et al. (2005), I estimate the model using Monte Carlo Markov chain methods, as implemented in the software *MLwiN*.

Formally, the models can be presented as follow

$$h_{ijt}^{(U)} = \alpha^{(U)}\boldsymbol{D}_{ijt}^{(U)} + \boldsymbol{\beta}^{(U)}\boldsymbol{F}_{ijt}^{(U)} + \boldsymbol{\gamma}\boldsymbol{X}_{ijt}^{(U)} + \boldsymbol{\delta}\boldsymbol{Z}_{i}^{(U)} + u_{j}^{(U)} \qquad (1)$$

$$h_{ijt}^{(F)} = \alpha^{(F)}\boldsymbol{D}_{ijt}^{(F)} + \beta^{(F)}\left\{\boldsymbol{U}_{ijt}^{(F)}, \boldsymbol{D}_{ijt}^{(F)}\right\} + \boldsymbol{\gamma}\boldsymbol{X}_{ijt}^{(F)} + \boldsymbol{\delta}\boldsymbol{Z}_{i}^{(F)} + u_{j}^{(F)} \qquad (2)$$

$$h_{ijt}^{(D)} = \alpha^{(D)}\boldsymbol{D}_{ijt}^{(D)} + \beta^{(D)}\boldsymbol{F}_{ijt}^{(D)} + \boldsymbol{\gamma}\boldsymbol{X}_{\boldsymbol{ijt}}^{(\boldsymbol{D})} + \boldsymbol{\delta}\boldsymbol{Z}_{i}^{(D)} + u_{j}^{(D)} \qquad (3)$$

In each equation, I model the hazard of the transition to union formation (Eq. 1), childbearing (Eq. 2) and union dissolution (Eq. 3) as a function of durations, outcomes of the other processes (endogenous covariates), observed background characteristics (exogenous covariates), and (potentially correlated) unobserved (time-invariant) heterogeneity components. In Equation (1) $h_{ij}^{(U)}$ denotes the hazard of a partnership formation during the time interval t of episode i for individual j. Eqs. (1), (2) and (3) define a multi-process model. The processes of union formation and union dissolution are similar in structure, except for the fact that they are mutually exclusive and for the measurement of exposure time (see next paragraph).

The baseline clocks for each hazard consist of linear splines with a quadratic term for all the three processes. In the childbearing process, the exposure to the risk starts at union formation (for first parity) or since previous childbirth (for the following parities). Once the first child has born, individuals become at risk of having a second birth, once the second child has born, they become at risk of a third conception, and so on. In the process of union formation, the exposure is set to begin from age 15 (for first union formation), and at the dissolution of the previous union (for subsequent partnerships). In the union dissolution equation, individuals are at risk after the start of the co-residential union.

The endogenous variable of fertility $\boldsymbol{F}_{ijt}^{(p)}$, with p={U, F, D}, accounts for different family configurations regarding the childbearing in the current partnership, the existence of children from past unions and stepchildren. $\boldsymbol{X}_{ijt}^{(p)}$ is a set of time varying explanatory variables, such as the proxy of family configurations and respondent's age. In the fertility process it includes the key categorical variables capturing family configurations are updated any time a birth occurs. For instance, a couple whose partners have not experienced parenthood before moving in, are categorized as a 'nonstepfamily without shared children'. If they have a child together, the family configuration turns into 'nonstepfamily with one shared child', in the following episode *i*. $\boldsymbol{\delta Z}_{i}^{(p)}$educational level, historical period, parents' separation, family social class, region of origin, ethnicity. $u_j^{(p)}$ is the heterogeneity term representing the effect of unobserved characteristics on each process, which is not captured by the observed covariates. This means that, if the processes are related, these components are correlated across the three processes: correlated random effects (a byword for heterogeneity term) would arise if the unobserved characteristics that influence the partnership transitions are correlated with those that affect childbearing within partnerships. In essence, this means that the processes are endogenous. The error terms are assumed to be constant over time for each respondent. Thereby, they capture some kind of "life-time average characteristics". Technically, this does not imply that individuals' personal traits remain stable over time. Instead, I assume that any shift in the risk of childbearing or partnership transition is caused by observed factors – such as age and number of children (Steele et al., 2005). The model assumes that random effects components have a joint normal distribution:

$$\begin{pmatrix} u^U \\ u^F \\ u^D \end{pmatrix} \sim N \left\{ \begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_U^2 \; \rho_{FU} \rho_{DU} \\ \rho_{UF} \; \sigma_F^2 \; \rho_{DF} \\ \rho_{UD} \rho_{FD} \; \sigma_D^2 \end{pmatrix} \right\}$$

Non-zero elements of the diagonal suggest the existence of unobserved heterogeneity in a specific process, while non-zero correlation in the elements of the sub-diagonal highlights that any or all processes are endogenous. Therefore, not controlling for the potential endogeneity of the processes would bias the estimates of the key variables of family configuration.

In the section Results, I will first show the estimates from single-process models, in which the random effects u^p are assumed to be mutually uncorrelated. This is the simplest way to fit the model, which is equivalent to estimate the three processes independently. In a second specification, I will add the estimates allowing for non-zero correlation of heterogeneity components between any pair of the three random effects. The comparison of the coefficient estimates between the two methods will arguably highlights the entity of the endogeneity removed by the estimates performed with multi-process models.

The length of the time interval is grouped into six-month intervals and weighted by exposure time, following Steele et al. (2005). Although the data granularity would allow me to perform analyses to the nearest month, I opted for broader interval to reduce the length of the records and ease the computability of the models. In each six-month interval, a weight is defined as the number of months in which the respondent is exposed to the transition-specific risk. These weights are also the denominators of the dichotomous outcomes. Therefore, I assume that hazard functions and covariates are constant within the six-month period without any loss of information.

The specification of the childbearing process is not common in the demographic literature. In previous studies of stepfamilies, parity-specific models have been formulated as the baseline hazard and the influence of explanatory variables are likely to differ depending on the birth order. Therefore, the estimated parameters are not specific to each order of birth or union, as the influence of (say) social class is the same for the first and all the subsequent transitions. However, the detailed operationalization of prior fertility and shared childbearing with the current partner is motivated by the need for diversifying individuals with prior childbearing and (possibly) relationship experiences from those with no children and/or prior unions.

The focus on three simultaneous processes comes at a cost also when it comes to the identification of cohabitation and marriage, which may have different influences on fertility and partnerships. Although technically possible, I chose not to treat marriage and cohabitation as distinct events (such as in Aassve et al., 2006) not to further complicate the statistical model and the computational feasibility. Another relevant issue of identification of simultaneous models concerns the exclusion restriction to be placed on the covariates. For equations to be identified in a system, a set of covariates included in a specific equation should be ruled out in the others. For instance, some factors involved in the childbearing process should not have any influence on the partnership transition, although it is hardly sustainable on theoretical grounds. However, the identification of the model is ensured by the observation of repeated events, whereas the unobserved heterogeneity is fixed over time (Lillard, Brien, & Waite, 1995; Steele et al., 2005).

6. Results

Table 3 presents the number of valid cases of women and men and selected descriptive statistics. The analysis features 12,751 (9,402) women (men) and spans over 21,148 (16,150) episodes of unions, which results in less than two spells of union per women (men). However, the distribution of partnership per individual is rather spread out as the majority experienced only one union. The proportion of episodes leading to a birth approaches 63% (57%) for women (men) with one union, and ranges between 20 and 30% in at least one union for individuals with two unions. Of all episodes observed, roughly 20% occurred in a stepfamily (Table 2). The prevalence of stepfamilies among families is comparable to other studies (Ermisch & Francesconi, 2000). These numbers seem adequate for the models and provide the statistical power to produce statistically meaningful estimates for the different family configurations.

Table 3. Sequencing of partnerships and within-union childbearing.

Variables	*Women*	*Men*
	Percent (%)[a]	
One union		
No shared children	36,3	42,4
At least one shared child	63,7	57,6
Total	100	100
Episodes*individuals (*n*)	11538	8714
Two unions		
No children in 1st and 2nd union	38,8	46,9
At least child in 1st union; no child in 2nd union	27,6	22,8
No children in 1st union; at least one child in 2nd union	23,6	22,1
At least one child in 1st and 2nd union	10,0	8,2
Total	100	100
Episodes*individuals (*n*)	6880	4713
Three unions or more		
No shared children	42,6	48,1
At least child in one union	57,4	51,9
Total	100	100
Episodes*individuals (*n*)	2730	2723

[a] It calculated as percentage of episodes*individuals in the models of partnership dissolution and childbearing

Before the illustration of the risks of childbearing and dissolution, I comment on the estimates of the parameters of the unobserved heterogeneity reported in Table 4. The diagonal displays the estimates of the standard deviations for each of the three equations, whereas the other cells report the estimated correlations between the error terms. All standard deviations are significant, although the magnitude varies among women and men. Positive correlation reflects that individuals that have frequents transition in one process tend to do the same in the other.

No causal effect is assumed. The correlations between the processes are significant, which provides evidence that fertility outcomes and partnership processes are endogenous, in which case the estimated effects from the single process model will be biased. The direction of this bias depends on the sign of the correlation and on the effect of the endogenous variables on the process outcomes.

The estimated random-effect covariances suggests that individuals who are less likely to have a child are also more prone to rapid transitions in forming and dissolving unions, in contrast with Aassve et al. (2006) and Upchurch et. al. (2002). This is an interesting result because it reveals that those who an above average risk of childbearing tend to find a partner more slowly (or do not repartner at all) and also have a predisposition to change partner less frequently (or not at all), conditioning on the observable variables that are used in the model. Conversely, the covariance between the processes of partnership dissolution and formation is positive and significant. This is interpreted as though women and men who are prone to unstable partnerships tend to have high chance of finding a (new) partner. The covariance terms for men are generally smaller.

Table 4. Estimated random-effects covariance matrix from the multi-process model

		Women	
	Union formation	Union dissolution	Childbearing
Union formation	0,456***		
	0,042		
Union dissolution	0,203***	0,382***	
	0,097	0,102	
Childbearing	0,254	-0,170*	0,293***
	0,165	0,087	0,069
		Men	
	Union formation	Union dissolution	Childbearing
Union formation	0,299***		
	0,063		
Union dissolution	0,113	0,292***	
	0,067	0,108	

Childbearing	0,154	-0,092	0,236***
	0,138	0,087	0,094

Notes : the values in each cell are the point estimate (mean of a MCMC sample). The results are based on 10,000 MCMC sample with a burn-in of 1,000.

*** < 0,01; ** < 0,05; *<0,1

Figures 1A and 1B show conceptions risks (hazard coefficients) by combined parity and step-family configurations for women and men. Within each graph, model 1 includes the risk of separation in a single equation model, while model 2 includes the risk of separation in a multi-process model, thus estimated jointly with the other processes of union formation and childbearing. The single-process estimates are from a model that assumes that the processes of partnership transitions and childbearing are independent. The multiprocess estimates are drawn from a model which assumes dependence between processes through correlated random effects. Of primary interest in this study are the influence of current and part union outcomes of the fertility process on the odds of union dissolution. In figures 1A and 1B are displayed the estimates from the single-process and multiprocess models. Their comparison highlights the effect of allowing for the endogeneity of the existence of children with respect to partnership transitions, and viceversa (Steele et al., 2005). Appendix includes complete estimates for the key variables and the controls for models 1 and 2.

In general, in the single-process models, births are less likely at higher combined parity, regardless of the status of the family configuration. Further, estimates of models for men and women are similar and are jointly commented except when explicitly mentioned. Net of combined parity, couples without common children have higher birth risk than of those with common children, with a few exceptions. More in detail, among non-stepfamily partnerships, those with one shared child, have a higher risk of a new birth. In fact, I expect that the transition rate to first birth for childless couples is higher than the transition rate from first to second birth for the other couples. However, many of the childless couples are selected among those short-lived unions with a high dissolution risk, as a heterogeneity term is not controlled for, and might dissolve before any birth planning. Interestingly, the stepfamilies with no shared children, which should be at least as likely as childless nonstepfamilies to progress to first birth according to the commitment hypothesis, are less likely do so with respect to the baseline category.

Moving across the figures 1A and 1B, I gain new insights on different combinations of family configurations. In the specifications addressing the number of stepchildren (the second group), I find that stepfamily unions with 1 stepchild only are more likely to have a first shared birth with their partner than nonstepfamily unions with non-shared children, in line with the evidence from Guzzo (2017). This finding does not necessarily mean that stepfamilies with one shared child end up with a higher completed *couple-level* fertility than non-stepfamilies. In fact, Thomson & Holland (2011) showed that the hazard of fertility is similar for stepfamilies with one pre-union child and for nonstepfamilies who have experienced one birth. After the first shared birth, the birth risk for stepfamilies is half the risk for nonstepfamily couples (Thomson & Holland, 2011). In other words, the couples with one pre-union child might tend to concentrate the first shared birth in the first 24 months from the establishment of the union (Thomson & Holland, 2011), possibly because their higher age impose a faster pace of childbearing not to forgo a childbirth altogether. In fact, the magnitude of the coefficient at issue decreases, when age is not controlled for (not shown in this analysis). This result – which is only reported for women – would have fully supported the commitment hypothesis if the family with 2+ stepchildren and no shared birth had displayed higher risk of a common birth. Finally, stepfamily couples with two or more shared children are significantly less likely to have another child than couples with one shared child only.

Figure 1A. Fertility process: risk of a birth. Estimated odds ratios by family configuration. Women.

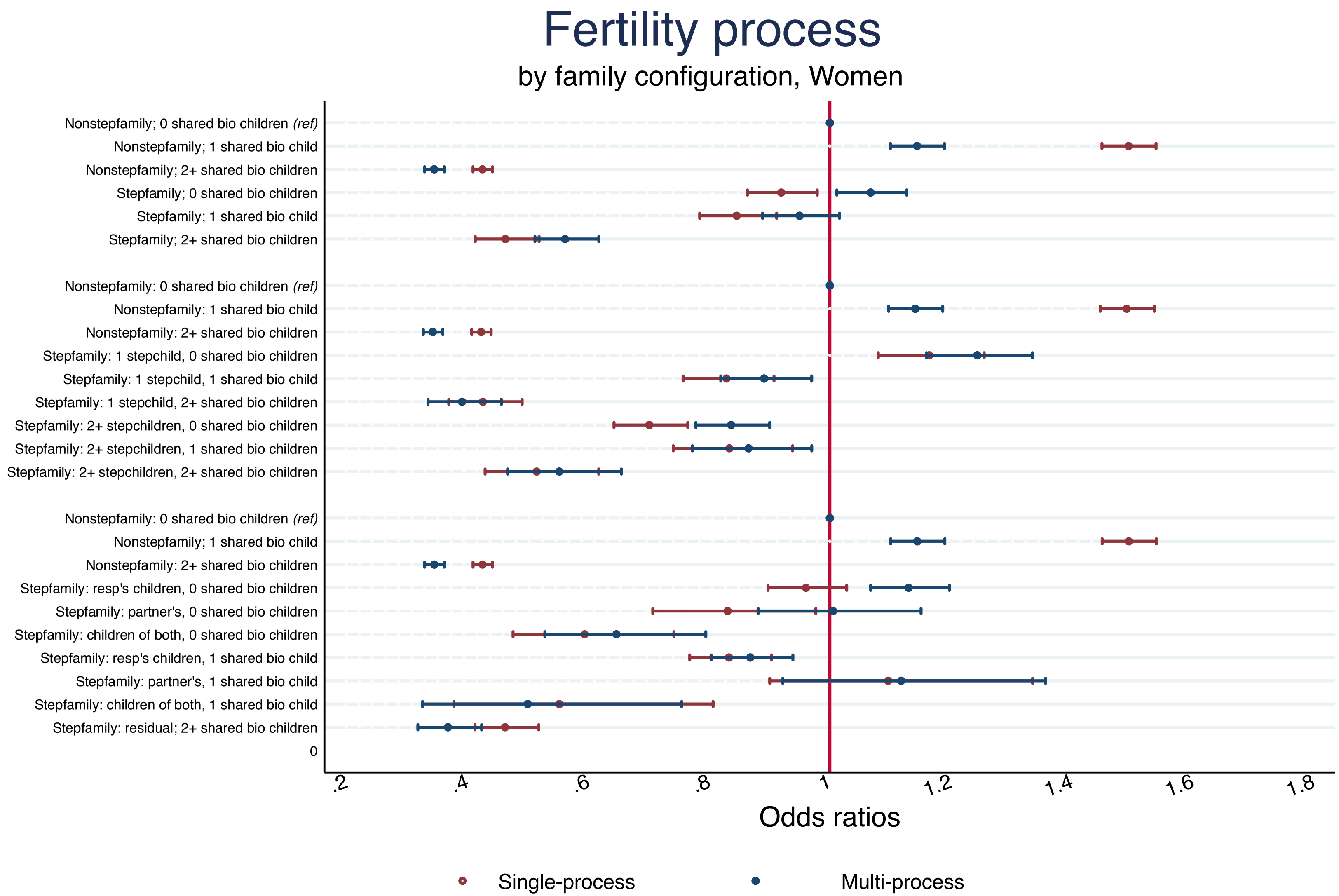
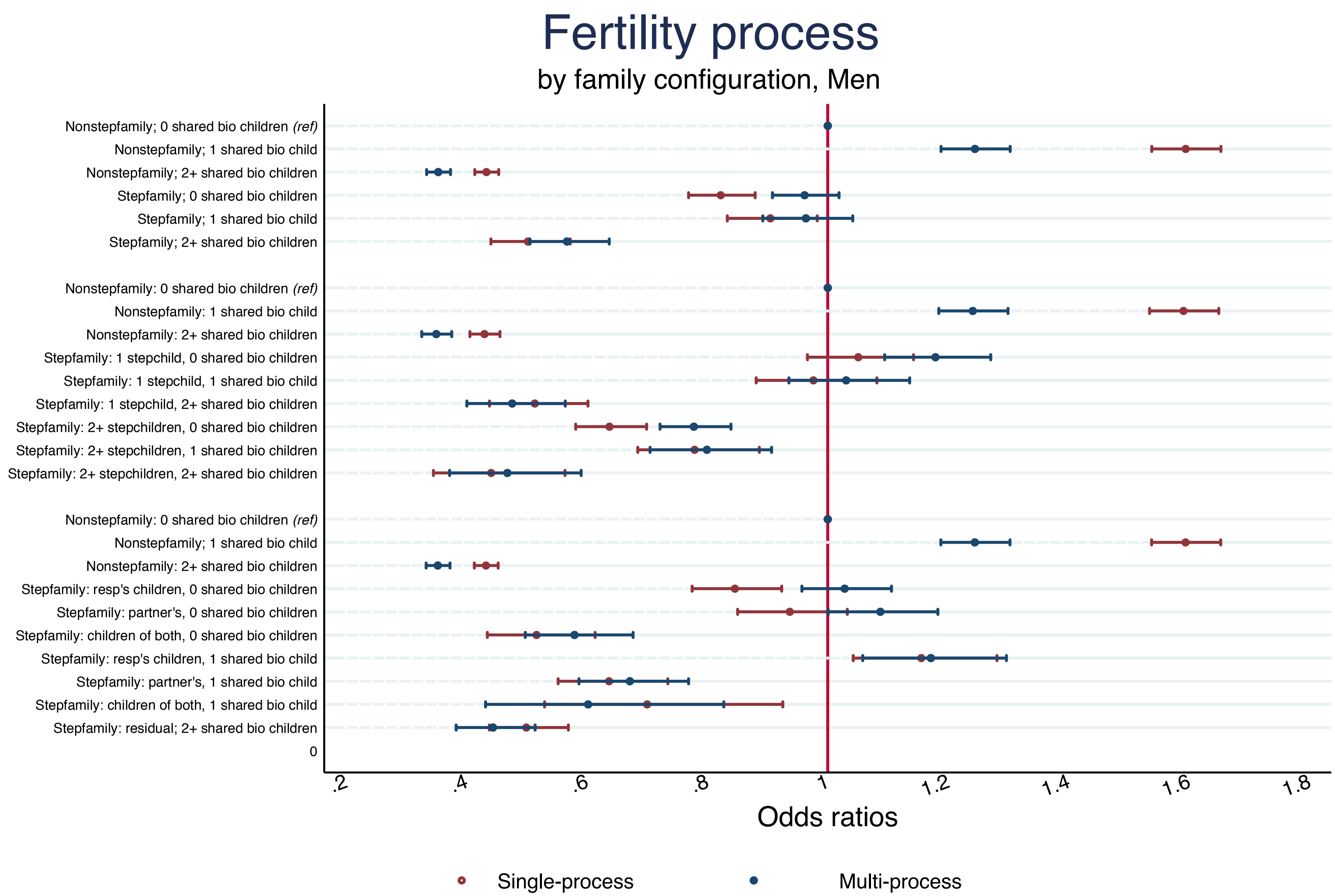


Figure 1B. Fertility process: risk of a birth. Estimated odds ratios by family configuration. Men.



When it comes to parentage of children, only stepfamilies in which the woman has one child – and the man none – have comparable risk of new childbirth relative to nonstepfamily childless unions. Interestingly, this result holds in the analyses of both men and women. Among stepfamily couples with no shared children, the risk of a first common birth is higher among the partnerships in which either partner has children vis-à-vis the 'blended' couples (those in which both partners brought children to the union), albeit not always significantly, for both genders. Among stepfamily unions with one shared child, the partnership with a father with a prior child displays a comparable risk of another shared birth to the baseline category. This result represents a clear puzzle if the parenthood or the commitment motives held true. These findings seem to suggest that *either* the parenthood *or* the commitment hypotheses are verified *under different family arrangements*.

In the multi-process model (in blue), in which unobserved association between the risk of partnership formation, separation and childbearing are taken into account, the estimated odds ratios change by a non-negligible size. The negative association between family dissolution and childbearing implies that individuals with greater propensity to stay together also tend to have more children, because partnership stability leads to higher fertility. If the childbearing process 'incorporates' the unobserved influence of 'partnership stability', a part of the negative effect of parity on further childbearing emerges so that the estimated odds are significantly different from those drawn from the single-process specification. Thus, the odds of a birth decrease for nonstepfamily couples with 1 child and 2+ children because the parity effects, which represent higher childrearing costs, are deprived of the unobserved propensity to partnership stability. When unobserved heterogeneity is accounted for, the odds of nonstepfamily couples with one shared child are 15% and 23% higher than the baseline unions, in the models of women and men respectively. Couple with two and more children are about two thirds less likely to experience a new birth than their childless peers.

Likewise, the lower stability of stepfamilies should counteract the union-binding effect exerted by one or more shared children. If stepfamilies were as stable as partnerships with no stepchildren taking out the extra risk of union dissolution of stepfamilies, I would find a greater risk of childbearing for couples without shared children or only one shared child. These effects become statistically significant in stepfamilies who have no shared children. Among male respondents, relative rates of birth increase to 20%, relative to the baseline groups; for women the relative almost doubles to 23%. Less pronounced increases are found in the relative risk of a second or in some cases a third shared birth to stepfamily couples although they do not differ much from the single-process estimates.

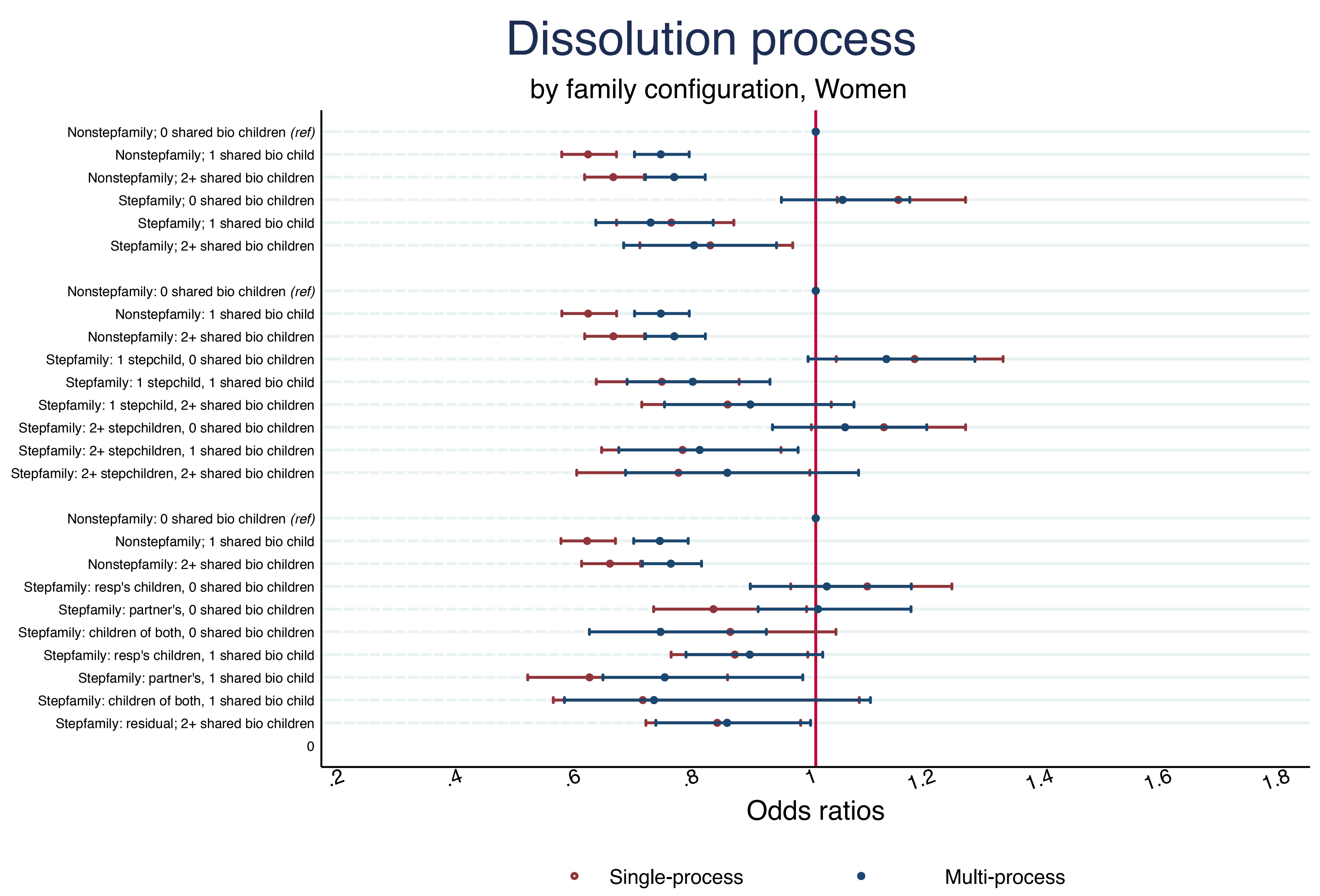
Couples in which one partner or the other has no children have higher chances of a shared birth in multi-process than in single process models. These couples – who don't share any child – are more likely to have a child than other couples with same (or similar) combined parity but no stepchild, and these couples still do have higher birth rates of couples in which each partner has one or more children (in contrast to Henz & Thomson, 2005). In other words, in line with the parenthood hypothesis, blended stepfamilies with no shared children are still less likely to have a child than stepfather or stepmother families.

The odds ratios of separation risks are shown in Graphs 2A and 2B for women and men, respectively. In single-process models, separation risks of stepfamilies are generally higher than separation risks of nonstepfamilies with some shared children, although the significance and the magnitude vary considerably across women and men. Among the former, the dissolution risk of stepfamilies with no shared children are higher than that of couples with no prior and common children. Among men, this finding is not confirmed. The birth of a first shared child in a stepfamily, which in my hypothesis could cement the partnership, arguably reduces the odds of separation vis-à-vis a stepfamily without common children, in both models in which a woman or a man are the main respondent. A second shared child does not further significantly decrease the chances of separation of all family configurations (either in a stepfamily or in a nonstepfamily). This suggests that it is not the number of shared children to affect the risk of dissolution but the very existence of common offspring. The other specifications do not add relevant evidence to these results, as significant difference by number and parentage of children are not found. The only exception emerges when the parentage of children in couples with no shared children is accounted for. Here, only the family setting in which the main respondent brings children to the family is significantly more at risk of partnership dissolution than the baseline category. In any case, I do not find evidence of a *gender* effect of the stepparent so that it is not possible to conclude that stepmother or stepfather families without shared children are more prone to dissolution.

The estimated odds ratios differ to some extent between the single and the multi-process models (in blue). Allowing for the endogeneity of fertility within partnerships has some impact on the effects of fertility outcomes on the risk that a union dissolves. In both models, the main results are confirmed as having one or more children reduces the risk of a separation for a couple. Interestingly, in the multi-process models the effects are slightly less strong, possibly owing to the negative residual correlation between the risk of dissolution and the odds of having children in a union. In the single-process model, the negative effects of shared births and the presence of stepchildren are probably biased by selection into partnership. Women and men with a higher risk of separation (those living in a union without shared children) are less likely to conceive during a partnership. These individuals decrease the risk of separation for women who are pregnant or who have children with a partner, leading to an overestimation of the "true" negative effect of the fertility outcomes on the risk of dissolution.

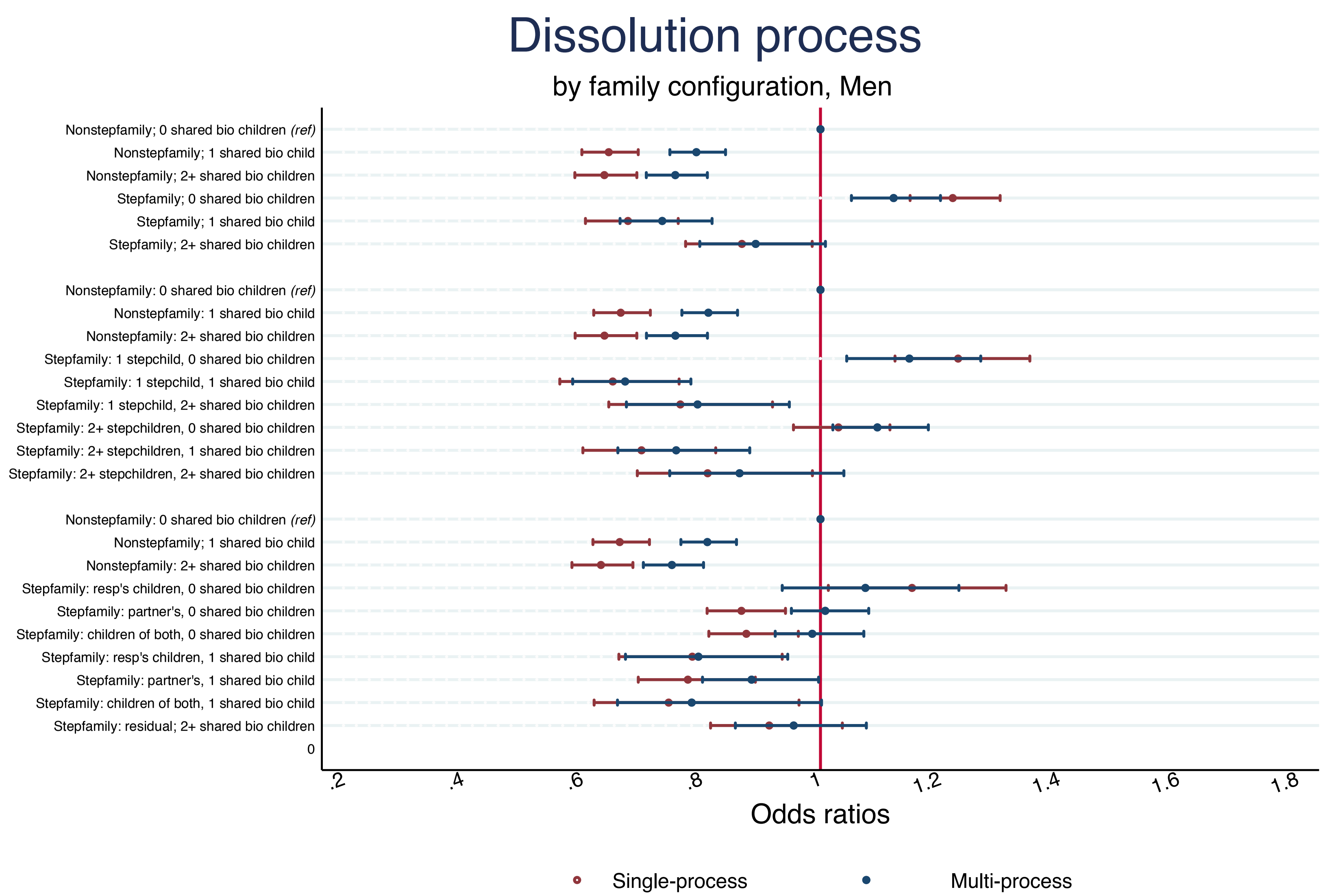
In the multi-process models, the gap in risk of separation between nonstepfamilies with children and the stepfamilies tends to close up. Nevertheless, the differences remain statistically significant between the nonstepfamilies with some shared children and the couples without common children, which confirm the binding effect of the shared offspring. In all the specification, and in both models of men and women, the difference between the estimated odds of separation between the stepfamilies and the nonstepfamilies without shared children approaches zero. This result highlights that the number of children from past union does not generally increase the risk of dissolution in couples who have not had any child.

Figure 2A. Dissolution process: risk of a separation. Estimated odds ratios by family configuration. Women.



Dissolution process
by family configuration, Women
Nonstepfamily; 0 shared bio children (ref)
Nonstepfamily; 1 shared bio child
Nonstepfamily; 2+ shared bio children
Stepfamily; 0 shared bio children
Stepfamily; 1 shared bio child
Stepfamily; 2+ shared bio children
Nonstepfamily: 0 shared bio children (ref)
Nonstepfamily: 1 shared bio child
Nonstepfamily: 2+ shared bio children
Stepfamily: 1 stepchild, 0 shared bio children
Stepfamily: 1 stepchild, 1 shared bio child
Stepfamily: 1 stepchild, 2+ shared bio children
Stepfamily: 2+ stepchildren, 0 shared bio children
Stepfamily: 2+ stepchildren, 1 shared bio children
Stepfamily: 2+ stepchildren, 2+ shared bio children
Nonstepfamily: 0 shared bio children (ref)
Nonstepfamily; 1 shared bio child
Nonstepfamily: 2+ shared bio children
Stepfamily: resp's children, 0 shared bio children
Stepfamily: partner's, 0 shared bio children
Stepfamily: children of both, 0 shared bio children
Stepfamily: resp's children, 1 shared bio child
Stepfamily: partner's, 1 shared bio child
Stepfamily: children of both, 1 shared bio child
Stepfamily: residual; 2+ shared bio children
0
.2
.4
.6
.8
1
1.2
1.4
1.6
1.8
Odds ratios
Single-process
Multi-process

Figure 2B. Dissolution process: risk of a separation. Estimated odds ratios by family configuration. Men.



7. Conclusions

In the few past decades, relevant changes in fertility behaviours as well as in union formation and stability have emerged. As adults increasingly spend longer time in the marriage market, many individuals end up forming unions with parents, thus establishing simple (with only one partner's children) or complex (with both partner's children) stepfamilies. Because of the growth of the stepfamily phenomenon, fertility, and to less extent dissolution, have become an area of research (Thomson et al., 2014; Henz & Thomson, 2005; Vikat et al. 1999, 2004), but their dynamics remain unclear. Stepfamilies feature a specific condition whereby one or both individuals have children. In nonstepfamilies at higher parities have generally lower risk of childbearing. Stepfamilies seem to have a different predisposition to fertility relative to couples in which partners are childless at the start. Further, nonstepfamilies generally display lower risk of dissolution after the birth of a first child which does not necessarily decrease at higher parity. Stepfamilies are not univocally shown to have lower or higher risk of separation because the binding effect of a shared birth is arguably compensated for by the adjustment of the stepchildren to the new family unit or by the presence of children with different parentage.

This study expands existing literature on the link between fertility and union dissolution in several ways. Rather than focusing on the outcomes of an individual's second co-residential union, I analyse all partnerships between 15 and 45 years using multilevel models that control for partnership history and personal background. I model transition from and into partnership along with childbirths within unions jointly, thus explicitly modelling the endogeneity of the presence of children with respect to the union and fertility pathways. Drawing on the literatures of family formation in stepfamilies as well as dissolution research, this study examines *five hypotheses* (three about childbearing and two about partnership separation) to assess the risk of childbearing and dissolution across different family configurations.

With respect to childbearing, the first hypothesis holds that the transition to a first shared child is affected by partners' pre-union parental status as parity progression could be dictated by partners' decision to have a child in order to achieve the status of parenthood ('parenthood hypothesis'). The second hypothesis argues that having a shared child serves to cement a couple's relationship, so that the partners signal each other commitment to the union ('commitment hypothesis'; Griffith et al., 1985; Vikat et al., 1999) and establish themselves as a family unit (Stewart, 2002). The third hypothesis holds that the subsequent shared (the second shared) child could provide a full sibling to the new-born ('sibling hypothesis') but this motive does not hold true in stepfamilies, in which a step-sibling can arguably act a full-sibling.

Overall, the results show that stepfamilies are not less likely than other couples to bear children, although childbearing within in a stepfamily is more sensitive to the configuration. Also, there is no conclusive evidence in favour of the parenthood hypothesis against the commitment hypothesis. On the one hand, stepfamilies with no shared children have higher chances to have a first common child than childless nonstepfamilies, even controlling for the unobserved heterogeneity. Further, in the analyses of women, the risk of first birth is marginally higher for the stepfamilies without a child in common than for those with a shared child. Taken together, these findings confirm the unique value of shared first birth and arguably back the commitment hypothesis, as prior children do not necessarily affect the transition to a first birth in a union. In other words, in the analyses of women's sample, the birth of a common child is important to confirm the establishment of a new couple. On the other hand, the risk of childbearing for stepfamilies with 2 or more pre-union children and no shared child is no higher than for stepfamilies with a comparable number of pre-union children and one shared child. Further, zooming in by the parentage of children, the findings reveal that couples in which both partners have already experienced parenthood are less likely to have a child, whether they already share one or they do not share any. To date, the accomplishment of parenthood by both partners proves negatively associated to the transition to a new birth either when stepfamilies have a shared child or when they do not. The latter finding is in line with previous evidence supporting the parenthood hypothesis, as a stepfamily childbearing risk tends to be lower when *both* partners achieved parenthood. When it comes to the third hypothesis, the results show that stepfamilies tend to ascribe the existing stepchildren to full-siblings of the firstborn in the couple. Indeed, nonstepfamilies with one shared child only are generally more likely to continue to a second birth compared to most of stepfamilies arrangements with the exceptions of the stepfamilies in which women enter union childless. This piece of evidence, which is mirrored in the samples of women and men, is arguably due to a gender misalignment in childbearing preferences. Further analyses should shed light on possible 'gender' differences within stepfamilies fertility behaviours.

To sum up, these results do not explicitly resolve the dichotomy between the 'parenthood' and 'commitment' motives, which has animated the debate in stepfamily fertility in the last two decades. The fairly large sample size and the innovative methods legitimize me to conclude that the two alternative strands are to be accepted as a result of the complexity of the dynamics involved in the childbearing decisions. A complex family, which is used as a synonym of stepfamily, implies difficult and, at times, contradictory decisions which mediate between parents' legitimate aspirations (e.g., the achievement of parenthood for the childless and the need for proving mutual commitment) and constraints (the presence of children within the household and consequential childbearing burden).

Jointly with the childbearing process, I also studied the influence of the presence, parentage and number of biological children and stepchildren on transition into and out of partnerships. I tested the hypothesis that having common children, regardless of the presence of stepchildren, reduces the risk of a partnership dissolution against the competing idea that shared children do impose additional obligations that arguably hinder the beneficial effects of the 'union-specific capital'. Moreover, I advanced the idea that, *ceteris paribus*, the presence of pre-union children makes a stepfamily more prone to dissolve in contrast to a nonstepfamily. I thereby tested different contexts in which the presence of stepchildren might exert a more negative influence on a partnership's stability. The results from the multi-process models confirm the hypothesis that *shared* children strengthen a relationship. Specifically, the risk of dissolution is higher in couples without shared biological children, although the influence is weaker in step-families as opposed to couples without pre-union children, in the sample of men. Also, I tested whether the stabilizing influence of shared children vanishes in more complex family settings, such as stepfamilies, and depending on the number or the parentage of the stepchildren themselves. No systematic difference between nonstepfamilies and stepfamilies with at least a shared child emerges, so the hypothesis that stepchildren represent a source of conflict within the couple, or, to some extent, hinder the family functioning has no empirical support.

I have argued that fertility, union formation and dissolution are three intertwined processes (Aassve et al., 2006) and that individuals' trajectories are simultaneously affected by some unobserved characteristics (Lillard & Waite, 1993). I proposed the interpretation that these unmeasured attributes encompass individuals' values, personal traits, or attitudes towards the specific relationship (Coppola & Di Cesare, 2008), or reflect long-term dynamics, such as propensity to infidelity or mistrust, which act as strong determinants of union dissolution and relationship hopping (Steele et al., 2005; Manlove et al., 2012). Ultimately, recent research also suggested that some individuals are more prone to repartner within stepfamilies rather with childless partners as they are not perceived as dependable partners in the marriage market (Lappegård & Rønsen, 2013; Schnor et al., 2017). In general, this intuition is confirmed as union dissolution is negatively associated with the risk of having further births, and positively associated with the risk of partnership entry, after controlling for parity, partnership history and background characteristics. This finding implies that *men and women* who are more prone to partnership dissolution have a below-average risk of having children within a union, in contrast to what Upchurch et al. (2002) and Steele et al., (2005) found in cohabiting unions, but in line with the evidence provided by Henz & Thomson (2005). The results of the single and multi-process models are surprisingly similar when I controlled for unobserved heterogeneity, especially when it comes to the dissolution process. The most plausible explanation lies on the relatively low level of significance of cross-equation correlation, in particular for men. Even though it is not the primary goal of my analysis, I also report a positive residual correlation of the hazards of the partnership transitions. There is strong evidence that *men/women* with a high chance of experiencing partnership break-up also have higher odds of forming a new partnership. Therefore, there are unobserved time-invariant attributes that make *men/women* more prone to relationship hopping.

This article has several limitations that will be be taken into account and possibly amended in the next stages. First, although UKHLS provides a wide array of past fertility and partnership histories across cohorts and over time, no information on socio-economic characteristics, such as income or employment history, is collected retrospectively. For instance, decisions concerning the start of a co-residential union or a separation – not to mention having a baby – are deeply affected by personal economic independence. Further, the economic conditions are time-varying factors which could differ at different stages of the life course and exert a different impact in first, second or higher order unions. Therefore, the methodological approach, which includes repetitive transitions, does not account for the impact of contextual changes family dynamics, resulting in a remarkable (and unfortunately unamendable) flaw.

Second, the format does not allow for the collection of detailed information of respondents' previous partners, such as full fertility history, age and socio-economic background. As previously mentioned, available data neither provide information on partners' non-resident children nor on their childcare involvement. Thus, the closest factor framing parents' engagement with pre-union children is children's residence, which is not sufficient to accurately test the influence of pre-union children on a new couple's childbearing under the childrearing responsibility theories. The absence of information of partners' non-resident children (if any) thus only allows the analysis to weigh up the role only respondents' co-resident children in family configuration. However, it must be acknowledged that previous evidence suggests that co-resident children may influence fertility (Hohmann-Marriott, 2015) and dissolution decisions (Steele et al., 2005), which makes this omission tolerable.

In line with this reasoning, I expect that the estimates of a shared birth to stepfamily couples to be conservative, especially in the sample in which women are the main respondents. It is in these couples, in which (male) partners are not explicitly asked to mention non-resident children, that they are more likely to forgo pre-union (non-resident) children. Therefore, some stepfamilies might be wrongly categorized as they might have more stepchildren than they actually had. The couple's combined number of co-resident children is negatively associated with childbearing and thus suppresses the stepfamily effect. If I had been able to account for non-resident partner's children, I could have gained stronger inference on the meaning of first and second shared birth for stepfamilies.

Finally, this paper does not address the different underlying clocks of childbearing across family types. Previous research (Henz, 20002; Holland & Thomson, 2011; Li, 2006) found a faster pace of childbearing for stepfamilies after family formation than after a shared birth. The 'stepfamily differential' is caused by the pace of childbearing, the age of the youngest child and the age of parents. For instance, previous research found that the transition to the first common child in a stepfamily is partly driven by the age of the youngest child (Henz & Thomson, 2011) and that the influence of younger children is greater than that of older children for the stability of a partnership (Steele et al., 2005). Of the aforementioned factors, only the respondent' age was eventually included in the models, while the age of the partner was missing and the age of the youngest child discarded, not to compromise the readability of the results. Further developments of this project will necessarily take the latter on board in order to challenge or confirm the existing empirical evidence.

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