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The Diversity of Union and Fertility Trajectories in the UK

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TESI DOCTORAL UPF 2019

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Declaration of Authorship

I certify that the thesis I have presented for examination for the PhD degree of the Pompeu Fabra University is solely my own work other than where I have clearly indicated that it is the work of others (in which case the extent of any work carried out jointly by me and any other person is clearly identified in it).

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Alessandro Di Nallo

Alla mia famiglia

Acknowledgements

I thank Gøsta Esping-Andersen for selecting me as a PhD student and for his special mentoring throughout the four years. I am also grateful to the professors who contributed to my research ideas and provided insightful feedback: Bruno, Clara, Aïda, Pau, Maria José, Luis, Jorge, Visnja.

I am very grateful to Roberta for being more than a special colleague.

I am really thankful to Léa for her outstanding contribution to my personal and academic growth. I am filled with gratitude to my colleagues and fellow PhD students: Luīze, Daniela, Lesia, Jorge, Alina, My Lan, Danilo, Sergio, Giulia, Francesco, Olatz, Juan Ramon, João, Francesco, Wan Li, Natalia, Stefano.

I owe a heartful "thank you!" to all the people I met in Barcelona: Alessandro, Meriem, Gabriel, Daria, Juan, Lara, Cesare, Veronica, Enrico, Aina, Cristian, Barbara, Jeroni, Diederik, Toni, Roger, Maricia.

I mention with pleasure all the fantastic people I met in Milan during my research stay: Veronica, Sissi, Luca, Valentina, Emanuele, Alessia, Marco, Filippo, Rebecca, Francesco; and all the friends I have regularly met at the conferences: Danilo, Ludovica, Chiara, Maria, Anna, Alice, Nicola, Cecilia, Benedetta, Agnese, Valeria.

During these years, I have seen many friends of mine become accomplished people in different countries and through diverse experiences. Overall, I am happy that distance and work did not set us apart: Giuseppe and Lisa, Simone and Daniela, Valeria, Marco and Teresa, Andrea, Rodolfo and Susanna, Alessandro, Filippo, Luca, Predrag, Davide, Andrea, Pierluca, Antonio and Marta, Emanuele, Roberto, Giorgio, Chiara, Barbara, Ettore, Sara, Riccardo, Barbara, Sara, Chiara.

Devo un ringraziamento speciale ai miei genitori che mi hanno seguito con affetto ed entusiasmo nel corso di questi anni difficili. Mi hanno incitato e mi hanno discretamente spronato, nei giusti momenti. Spero che, quando maneggeranno questo spesso volume, sentano i propri sacrifici ripagati.

Abstract

This thesis explores the increasing complexity in the life course in the United Kingdom. The first article illustrates the association between parents' socio-economic position and children's transition to first union and parenthood. The analytical sample (N=35,880) is drawn from the UK Household Longitudinal Study (Understanding Society) with respondents being born between 1930 and 1980, to examine not only the effect of children's background on union formation and parenthood but also whether this effect changes over birth cohorts, periods, and the life course, and varies by gender. The results lend support to the hypothesis of a negative relation between socio-economic family background and timing of first union – whether in a marriage or in a cohabitation – and a first non-marital birth.

In the second article, I study the risk of a birth and the risk of separation in different union settings, such as step-families and families with no prior children. I test countervailing hypotheses on childbearing transition. I also test the association between (biological and step) children influence the risk of union dissolution. Using multilevel multiprocess models with simultaneous equations, I model partnership transitions jointly with fertility. The analysis is based on the partnership and birth histories of the Wave 1 of UKHLS (Understanding Society) of men and women aged 16-45. The findings indicate that both the parenthood and the commitment motives influence the transitions to a birth, under different family configurations. Further, the risk of separation is reduced by the presence of shared children, while the existence of children from prior unions does not generally increase the risk of dissolution.

The third article assesses whether parenthood influences repartnering for women and men and explores how repartnering is associated with parental status of the prospective partners. The results, based on the Wave 1 of UKHLS, suggest that mothers, and to a lesser extent fathers, are less likely to repartner than their childless counterparts. Among parents who have child custody, there emerges a distinct gender gap because mothers exhibit a significantly lower rate of repartnering than fathers. Finally, coresident single parents are relatively less likely to repartner with child- less individuals, and single fathers more frequently form two-parent stepfamilies than do mothers.

Keywords: Fertility; Stepfamilies; Cohabitation; Marriage; Repartnering; Multilevel multistate model; Understanding Society

Resum

Aquesta tesi explora la creixent complexitat del curs de vida al Regne Unit. El primer article il·lustra l'associació entre la posició socioeconòmica dels pares i la transició dels infants a la primera unió i paternitat. La mostra analítica (N = 35,880) es dibuixa a partir de UKHLS (Understanding Society) amb els enquestats nascuts entre els anys 1930 i 1980, per examinar no només l'efecte dels antecedents infantils sobre la formació i la paternitat dels sindicats, sinó també si aquest efecte canvia cohorts de naixement, períodes i el curs de vida, i varia segons el sexe. Els resultats donen suport a la hipòtesi d'una relació negativa entre els antecedents socioeconòmics de la família i el moment de la primera unió-ja sigui en un matrimoni o en una convivència- i un primer naixement no marital.

En el segon article, estudio el risc d'un naixement i el risc de separació en diferents entorns sindicals, com ara les famílies de passos i les famílies sense fills anteriors. Puc provar hipòtesis compensatòries sobre la transició infantil. També prova l'associació entre els nens (biològics i de pas) que influeixen en el risc de la dissolució de la unió. Mitjançant models multicanal de múltiples processos amb equacions simultànies, modelo transicions d'associació conjuntament amb la fertilitat. L'anàlisi es basa en les històries d'associació i naixement de la Wave 1 de UKHLS d'homes i dones de 16 a 45 anys. Els resultats indiquen que tant la paternitat com els motius de compromís influeixen en les transicions a un naixement, en diferents configuracions familiars. A més, el risc de separació es redueix per la presència de nens compartits, mentre que l'existència de fills de sindicats previs no sol augmentar el risc de dissolució.

El tercer article valora si la paternitat influeix en la repartició de dones i homes i explora com el repartiment està associat amb l'estat parental dels socis potencials. Els resultats suggereixen que les mares i, en menor mesura, els pares tenen menys probabilitats de repartiment que els seus homòlegs sense fills. Entre els pares que tenen custòdia de menors, sorgeix una diferència de gènere diferent perquè les mares presenten una taxa significativa de repartiment que els pares. Finalment, els pares monoparentals són relativament menys propensos a distribuir-se amb persones sense fills, i els pares solters freqüentment formen famílies parentals que les mares.

Paraules clau: Fertilitat; Step-family; Cohabitació; Matrimoni; Model de multi-nivell i multi-estat; repartnering; UK

Table of contents

List of Tables

List of Figures

Chapter 1

Introduction

All Western European Countries have experienced major changes in partnership and fertility trends, such as increased divorce, delayed or forgone marriage, formation of new partnerships, multiple-partner fertility (Beaujouan & Solaz, 2012; Perelli-Harris & Lyons-Amos, 2013; Thomson, et al., 2014). These components of what is commonly deemed the "Second Demographic Transition" (van De Kaa, 2002) have re-shaped the family landscape over the past half century, so that much of family life now unfolds outside the bounds of marriage.

The increase in divorce rates over the last half a century has been considered to be one of the main factors undermining people's partnership histories and a major threat to fertility (Thomson, et al., 2012; van Bavel, et al., 2012). However, the raising divorce trends have not necessarily meant a retreat from partnership in general. Empirical evidence suggests that younger generations have increasingly opted for alternative union arrangements, such as cohabitation (Beaujouan & Ní Bhrolcháin, 2011) and that the majority of divorcees repartner (Coleman et al., 2000; Sweeney, 2002) with a stronger penchant for cohabitation over remarriage (Z. Wu & Schimmele, 2005). The fact that an increasing number of people separate (or divorce) and repartner at childbearing ages means that fertility and partnership decisions are frequently made over a longer age span and besides the first-order union (Sweeney, 2010). This thesis addresses demographic transitions– entry into first union and parenthood; repartnering; childbearing and separation over the life course – in the United Kingdom, a country that epitomizes the growing variety of patterns in adult life.

The UK is an interesting case study due to the existence of many of these demographic changes: a sharp rise in divorce rates in the 1970s followed by a generalized increase in cohabitation (Kiernan, 1998; Berrington & Diamond, 2000), relatively high overall levels of completed fertility with significant levels of childlessness (Berrington, et al., 2015), a late average age at entry into motherhood, high rates out-of-marriage births (Winkler-Dworak, et al., 2017), one of the most heterogeneous patterns of union formation in Europe Europe (Perelli-Harris & Lyons-Amos, 2015).

Fertility outside marriage is socially (Basten, et al., 2014). Over the last two decades, the proportion of women of childbearing age who are married has fallen from 53 per cent in 1991 to 34 per cent in 2008 (O'Leary, et al., 2010). Changing age patterns of fertility are linked to social, cultural and economic factors and have occurred simultaneously with changes in partnership formation. During the same period, the mean age at first marriage for women has risen from 25.5 years in 1991 to 29.9 years in 2008, and for men from 27 to over 32 in the same years (O' Leary et al., 2010). These changes in marriage behaviour have been accompanied by increases in the proportion of people cohabiting, and the proportion living alone: union dissolution has become a common experience, especially for cohorts born after 1960 (Basten et al., 2014). Marriage rates have also declined and only been partly offset by higher rates of cohabitation (Berrington & Diamond, 2000). Not accidentally, the majority of the British population see little difference between marriage and cohabitation (Duncan & Phillips, 2008).

The establishment of a partnership has been recognized as one of the milestones of the transition to adulthood (Shanahan, 2000). In the past, this typically meant the start of the first and only marriage. However, the choice of the first partnership type is now no longer limited to traditional life-long marriage, as cohabitation has become increasingly common and replaced marriage as the most frequent arrangement of first (Mike Murphy, 2000a; Ní Bhrolcháin & Beaujouan, 2012), and has become an integral part of family life in Western countries (Kennedy & Bumpass, 2008; Kiernan, 2001). Cohabitation has therefore gradually gained the role of test-bed for marriage and also alternative to it (e.g., Murphy, 2000; Perelli-Harris, 2014). In the United Kingdom, between 1980 and 2007, the proportion of men and women entering a marital union before the 25th birthday has fallen from 51% to 10.4% and from 71% to 23%, respectively, while the prevalence of those opting for cohabitation by the same age has increased from 16% to 38.5% and from 18% to 50.8%, respectively (Beaujouan & Ní Bhrolcháin, 2011).

The surge of cohabitation is evidenced also in the partnership context of childbearing for the cohorts born in the 1940s through the 1970s. A dramatic increase in out-of-wedlock childbearing, most notably in cohabiting unions, had occurred by the early 1990s. The marriage rate among women experiencing their transition to motherhood within cohabitation was lower in Britain than in any other country in Western Europe (Kiernan, 2001). In 1991, 30 per cent of births took place outside marriage. By 2008, 45 per cent of babies were born to unmarried parents (O'Leary et al., 2010). Also "solo" parenthood (having a child outside of a union) appears to have increased mainly in the United Kingdom, besides the United States (Kiernan, 2004; Hobcraft, 2008).

The transition to adulthood is a period of the life course characterized by the density of transitions into new roles and responsibilities such as movement out of school and into the labour market, a union, parenting, and independent residence. These events are so crucial to the life course that deviations from their 'normative' timing and sequential order have supposedly profound effects on later life: most research considers a stable partnership to be the optimal context for childbearing and childrearing (Kiernan & Hobcraft, 1997; Thornton & Young-DeMarco, 2001). A stable partnership lowers childrearing costs for each parent and may enhance the benefits of children through mutual enjoyment and caring. Children also benefit from stable partnerships (Amato, 2001), providing additional motivations to avoid or postpone childbearing when one has no partner or union stability is in doubt. Conversely, off-time or out-of-order transitions, such as becoming a parent early or in a cohabitation are shown to be negatively consequential to the children's life course (Chase-Lansdale, et al., 1995; Kiernan & Hobcraft, 2001; Waite & Lillard, 1991; Hardy, et al., 1998; Kiernan & Hobcraft, 2001; Krohn, et al., 1997). Therefore, to the extent that the establishment of first union and parenthood have consequences for individuals' future socioeconomic trajectories and well-being (Steven L. Nock, 1998) and is linked to outcomes for children (Sarah McLanahan, 2011), this is an important topic that has implications for both research and public policy.

A wide array of previous studies showed that intergenerational associations of entry into union and fertility are primarily a consequence of socioeconomic commonalities across generations. In other words, if socioeconomic traits are associated with the timing of transition to adulthood and parenthood, any observed association in these outcomes could be due to similarities in socioeconomic characteristics across generations. While the existence of intergenerational transmission of timing of fertility and union formation is well documented, the family arrangements in which these transitions occur (whether in marriage or in a cohabitation) are largely unexplored, and so are the mechanisms behind these transmissions.

Following in the vein as Michael & Tuma (1985) and Barber (2001), the purpose of the Chapter 2 is to identify the factors of family background that might affect the timing and the type of union formation and transition to parenthood, adopting retrospective information on adulthood transition of individuals interviewed in the Wave 1 of

Understanding Society. Furthermore, I will also contribute to the literature by considering different dimensions of the stratification order – namely social class and parental education – which arguably have independent influence on children's (Elo, 2009) and help the understanding of the mechanisms underlying these intergenerational processes. Finally, I will examine whether the effects of parental socioeconomic resources on the timing of first marriage vary both historically and over the young adult life course. Drawing on cultural theories of individuation (e.g., Lesthaeghe, 1983) and the life-course approach (e.g., Hogan & Astone, 1986), I assess whether these family background characteristics may have weakened over historical time and also weaken as young adults age.

First unions such as cohabitations have been consistently found to be less stable than first unions as marriages (Poortman & Lyngstad, 2007). In the United Kingdom increases in cohabitation underlie a concomitant rise in the prevalence of repartnering across cohorts (Perelli-Harris & Lyons-Amos, 2015; Gałęzewska, et al., 2014). The first union has progressively abandoned the long-term commitment that used to be for the generations born up to the 1940s-1950s (Murphy, 2000) and has shifted toward a 'trial stage' for a future more stable union (Perelli-Harris, 2014). Partnership trajectories have become diverse according to the type and number of unions formed during the life course. It is likely that second and higher-order unions differ from the first union as they often involve individuals with more complex life trajectories, with multiple spells of partnerships, children from previous relationships, and the continuing influence of previous partners and their family members (Poortman & Lyngstad, 2007; Beaujouan, 2016; Teachman, 2008). Also, post-separation singlehood, in contrast to the first singlehood, which precedes the first partnership, involves individuals who have learned about the process of breakup (Poortman, 2007). Going through this often-painful process may cause people to be more cautious the next time, be less committed to and invest less mental resources in the second or higher-order union compared to the first (Furstenberg & Spanier, 1984; Poortman, 2007). Furthermore, marriage market conditions have also changed because people are older when they search for a partner for the second time, and therefore the pool of potential partners is more restricted (Teachman, 2008). Thus, it is likely that the factors linked to the formation of second and higher-order unions are not the same as those linked to the formation of the first union.

The existing research has devoted little attention to a key determinant of union formation: the relationship career. Most evidence on union formation has focused on social-demographic and socioeconomic determinants and, with the exception of Poortman (2007), on the formation of either first or second unions. Chapter 3 aims to fill this gap in empirical knowledge and addresses the interrelationships between first and higher-order unions jointly with the childbearing history of both partners involved in repartnering, which is another innovation in the literature. Most of the research on how children affect the probability of a union has generally examined the question only for those with children. Further, the focus has been almost entirely on women and co-resident children, thus generally ignoring men and fathers with non-resident children. With the rise in separation and out-of-union childbearing, men have increasingly not lived with the children they have fathered (Henz, 2014). The issue of repartnering takes on particular salience in this context, because most men and women who enter new relationships do so before the children born of prior relationships are grown, and possibly receive care from different persons: a co-resident parent, his/her new partner, and a non-resident parent (Cancian, et al., 2014; Carlson, et al., 2008; Haux, et al., 2015; Kiernan, 2006). Ultimately, the high level of union disruption creates a large pool of individuals and children at risk of forming ‘simple’ stepfamilies (in which one partner brings children born to another partner to the new couple) or ‘complex’ stepfamilies (in which both partners live with their children born in a previous union). Whether this family complexity is neutral for the adults and children’s well-being has been debated with conclusions that hint at family malfunctioning (Brown & Manning, 2009; Stewart, 2005; Balbo & Ivanova, forthcoming).

Chapter 3 seeks tries to establish whether systematic patterns of repartnering exist according to an individual’s gender, parental status, children residence and his/her prospective partner’s parental status. Using retrospective partnership and fertility histories of individuals born in the decades 1950s through 1970s drawn from the Wave 1 of Understanding Society, I assess whether there exists a significant difference in repartnering chances between men and women and whether, and to what extent this gap is explained by the parental status (having children or being childless) and by the residence status of children (having co-resident children or non-resident children), their age and number. The second objective of the analysis is to clarify which type of next partner is systematically associated to an individual’s parental status. Are the childless more or less likely to repartner with other childless or with custodial parents? Or with non-resident parents? Do these patterns vary according to gender? Existing research knows very little about the processes involved in this fundamental restructuring of family relationships. These questions are addressed for the first time in the United Kingdom, where the fairly high levels of female labour-force participation, the relatively low male involvement in daily childcare (Lewis, 2002)and post-separation children's custody (Trinder, 2010) reinforce the unequal burden of childcare responsibilities of women.

Repartnering dramatically increases the complexity of household arrangements as a substantial proportion of separated and divorced parents (Sweeney, 2010), and previously unpartnered single parents, mainly mothers, (Ermisch & Francesconi, 2000) enter new unions. It is estimated that in the 1990s about 30% of mothers in the United Kingdom had formed stepfamilies, and 86% of children living in a stepfamily were residing with their mother and a stepfather (Ermisch & Francesconi, 2000). Among divorced mothers with at least one child, 32% went on to have another child within 5 years in a new relationship, and 45% had another child within 10 years (Hohmann-Marriott, 2015; Jefferies, Berrington, & Diamond, 2000).

When a stepfamily is formed, a child might have two or three adults taking on a parental role. When a stepparent has had children, a child could also have step-siblings, who might live together and shift back and forth between the household of the parents (Elizabeth Thomson, 2014). Eventually, separated or divorced individuals who enter new unions have motives and opportunities to have (further) children (Thomson, 2014). Offspring in higher-order union could be half-siblings if parents have had prior children. These children add complexity to the family as they may not live in the same family or live further apart in age than full siblings.

Ultimately, higher union formation might multiply the number and the type of relationships in a family if at least a child is present: in this case, a partner becomes a stepparent, and a child becomes also a stepchild. The configuration complexity escalates if children from both partners are present and both are actively engaged in childrearing, as the biological children of distinct parents become stepchildren. The family complexity tops the level of complexity if partners have a shared child so that the linkages between half-siblings are set up. The implications of family complexity are sizeable and have been only marginally explored in the demographic literature. Of higher-order unions and stepfamilies occurring across the life course, those formed among women and men of childbearing age are particularly relevant for children's wellbeing (Thomson & McLanahan, 2012; Thomson, Winkler-Dworak, Spielauer, & Prskawetz, 2012; Balbo & Ivanova, forthcoming). Also, when a union is dissolved, there is likely greater claim on economic resources and assistance, and partners tend to experience extra stress also during the formation of new unions (Amato & Kane, 2011; Ferri & Smith, 1998; Steele, Kallis, & Joshi, 2006).

Chapter 4 addresses the consequences of different family configurations on the risks of childbearing and separation. Few prior studies have estimated the transition to a new birth and to union breakup across different family settings, including the stepfamilies. The last chapter will collect elements of the lines of research of stepfamilies and multiple-partner fertility and elements of research in partnership dissolution to estimate the risk of transition to a further child and the risk of in different family arrangements (e.g., childless couples vs. stepfamilies; couples with biological shared children only vs. couples with some stepchildren, etc.).

Although any union is at risk of new childbearing, stepfamilies feature specific conditions whereby one or both individuals have children. In 'nonstepfamilies' (families who do not have children from previous unions), higher parities are associated with lower risk of further childbearing. Stepfamilies seem to have a different predisposition for fertility but no clear conclusion has emerged, due to the complexity of the possible family combinations. Two lines of research have examined fertility across partnerships. The stepfamily literature has predominantly examined childbearing by comparing how likely the transition to having a common child in the new union is for individuals with and without prior (Buber & Furnkranz-Prskawetz, 2000; Griffith, et al., 1985; Prskawetz, et al., 2003; Stewart, 2002; Vikat, et al., 1999). Stepfamilies often report the plan to have children in the future (Hohmann-Marriott 2015; Stewart 2002) and one-half of stepfamilies actually have a shared child (Holland & Thomson, 2011), even at a combined parity that is higher than for nonstepfamilies. One motivation is that stepfamilies could be especially motivated to have shared child to signal the couple's status as a family and their mutual commitment (Griffith et al., 1985). Another motivation hints at one partner's willingness to experience parenthood and achieve the adult status if one partner hasn't had any child. In the existing literature, these hypotheses have alternatively obtained empirical support or raised doubts, probably as a consequence of distinct (and complex) mechanisms underlying the childbearing process.

Another line of research has looked at the recently recognized phenomenon of multiple-partner fertility, which is a greater domain of stepfamily fertility. This field has mainly grown separately from the literature about stepfamilies, perhaps because multiple-partner fertility has been largely investigated as a deleterious behavior originating from problematic relationship behavior and persisting in low socio-economic strata (Carlson

& Furstenberg, 2006; Guzzo, 2017; Klerman, 2007). On the one hand, multiple-partner fertility is a condition of the individuals (focusing on whether a person has children with two or more partners), whereas stepfamily fertility is a characteristic of the union (concerning on whether a union has children from either partner's prior relationships, regardless of residence, and on whether the partners have a shared child). While the strand of stepfamily literature concentrates on couple-level fertility, the multiple-partner fertility research addresses individuals' lifetime fertility, which spans across multiple unions. Chapter 4 puts the two frameworks together to focus on childbearing in couples (a typical approach in the stepfamily research) in a life course perspective (which borrows mostly from the multiple-partner fertility). Thus, I will take into consideration all episodes in which individuals are at risk of childbearing – the unions – in models that explicitly account for repeating events.

A critical component of stepfamily life that has received much less attention is the risk of dissolution as opposed to couples who do not have stepchildren or do not have children altogether. Incomplete institutionalization has been often deemed as one explanation for the instability of higher-order unions and/or stepfamilies. Certainly, selectivity in background characteristics and inclination for partnership dissolution do play a role (Bumpass, Sweet, & Martin, 1990). Also, the parental responsibilities and/or the higher union order for at least one partner are plausible predictors of separation as well. However, evidence on the association between partnership dissolution and having children is somewhat mixed. Earlier research has found different, even opposite, effects of having children on partnership dissolution across countries and in different family situations with regard to, for example, the number, age, and residence of children (Coppola & Di Cesare, 2008; Lillard & Waite, 1993; Lyngstad & Jalovaara, 2010; Steele, Kallis, Goldstein, & Joshi, 2005; Svarer & Verner, 2008; J. Teachman, 2008). In case of second and higher-order unions, the picture is also incomplete and varies either when only marriages are taken into account (Kulu, 2014) or when second unions (including married and unmarried relationships) are analysed (Poortman & Lyngstad, 2007; Beaujouan, 2016).

Chapter 4 aims to fill this gap by analyzing the risk of union dissolution of different family configurations. I seek to innovate the literature by studying the fertility behaviours jointly with partnership events, using the partnership and fertility histories of women and men interviewed in Wave 1 of Understanding Society. More in detail, I examine the relationship between the risk of a couple to have a (further) child and the risk of the formation and dissolution of the ongoing union, in different family settings (e.g., nonstepfamily vs. stepfamily; couple with shared children vs. couple without shared children). I assume that the decision to have a child with a partner is likely to be jointly determined with the decision to form that union or to move from a partnership to singlehood. If decisions about partnerships and childbearing are jointly determined, there are unobserved components underlying these processes that are correlated. Further, unobserved factors are likely to influence individuals to (i.e.) self-select into multiple partnerships, or to explain their inclination for serial childbearing in different unions. In other words, performing analyses on demographic transitions without accounting for individuals' heterogeneity, which is possibly unobserved and reflects one's beliefs and values (Lillard & Waite, 1993; Thornton, 1977), only reveals part of the story. I will apply simultaneous hazard models (Lillard & Waite, 1993; Steele et al., 2005) to study fertility decisions and union transitions as intertwined processes, telling direct reciprocal effects from common determinants. Along with this strategy, I will also a 'simple-process' models, which disregards these unobserved determinants, to assess to what extent a simultaneous modeling improves the comprehension of union dissolution and fertility decisions.

Each of the following three chapters of the thesis will thus investigate diverse transitions at different stages of the life course in the United Kingdom. Chapter 2 assesses the transition to first union and parenthood in cohabitation and marriage in association with the socio-economic background of the young adults. Chapter 3 explores the repartnering chances of the childless as opposed to parents for women and men, and whether these individuals systematically form a new union with a childless partner or a parent. Chapter 4 estimates jointly the risk of childbearing and separation in couples in which individuals enter during the course of their life, with a special attention for complex family configurations. Chapter 5 summarizes the contributions and main findings of chapters 2 through 4, discusses their limitations, and traces outlines for future research.

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Chapter 2

The influence of parents’ social class and education on children’s transition to first union and parenthood in the United Kingdom.

Abstract

Family background shapes young adults’ decisions in their transition to adulthood, and the outcomes of these decisions lay the foundation for their subsequent life course. This study examines how parents’ education and social class influence their children’s union formation and parenthood formation. I examine the timing of entry into a first union (a married or a cohabiting union), the choice between marriage and cohabitation, the timing of entry into parenthood and the choice between a non-marital and a marital first birth. The analytical sample (N=35,880) is drawn from the UK Household Longitudinal Study (Understanding Society) with respondents being born between 1930 and 1980, to examine not only the effect of children’s background on union formation and parenthood but also whether this effect changes over birth cohorts, periods, and the life course, and varies by gender. The results lend support to the hypothesis of a negative relation between socio-economic family background and timing of first union – whether in a marriage or in a cohabitation – and a first non-marital birth.

1. Introduction

In the field of intergenerational transmission of partnership behavior and childbearing, the influence of parental socio-economic background has been extensively investigated (Axinn & Thornton, 1992; Mooyaart & Liefbroer, 2016; Thornton, et al., 2008) Most research has concentrated on the timing of first union and first parenthood, which are generally found to occur earlier for children whose parents are low-educated (Axinn & Thornton, 1992; Goldscheider, et al., 2009; South, 2001). Further, the literature has highlighted that low socio-economic background of origin is linked with disadvantaged family demographic behavior, such as early and non-marital childbirth (Duncan, et al., 1998; Haveman & Wolfe, 1993), early cohabitation (J. Teachman, 2003) as well as divorce (Teachman, 2002), and with socio-economic disadvantage later in life (e.g, Furstenberg, 2008).

However, little research has explicitly addressed the influence of socio-economic background on union and parenthood timing among adults entering their first cohabitation. Including cohabitation in the study of union formation and transition to parenthood is very relevant. In the United Kingdom, such as in most Western Countries, cohabitation has become increasingly common and replaced marriage as the most frequent arrangement of first relationships (Beaujouan & Ni Bhrolcháin, 2011)

This study contributes to the literature also by addressing explicitly the role of two distinct measures of social background (parents' class and education) onto the entry into union and parenthood. Recent stratification research has acknowledged that distinct measures of social background capture different dimensions of the stratification order (e.g., Chan & Goldthorpe, 2007) and have differential effects on children's outcomes (Elo, 2009; Torssander & Erikson, 2009) and demographic behavior (Dahlberg, 2015).

The purpose of this article is to identify the factors of family background that affect the timing and the type of union formation and transition to parenthood, drawing on economic theories of individual utility (Becker, 1981), socialization (Mclanahan & Bumpass, 1988), and life course approach (Elder, 1975; Hogan & Astone, 1986). Partnership and fertility decisions are distinct, albeit closely connected (Barber, 2001; Michael & Tuma, 1985), although the previous research in sociology and demography tend to separate these subjects, with few exceptions (e.g., Barber, 2001; Michael & Tuma, 1985), and marriage and childbearing have become less linked over time (Pagnini & Rindfuss, 1993). However, "becoming a parent, like becoming a spouse, reflects a decision to adopt an adult role" (Micheal & Tuma, 1985, p. 517) and childbearing preferences play a role in determining marriage timing (Jennifer S. Barber & Axinn, 1998).

In stratification research, the influence of parents on children is considered an indicator of opportunity in the society (Barber, 2001) because individual's socioeconomic status tends to mirror that of the previous generation (Kolk, 2014). The transition to adulthood has significant implications for transmission of parental socioeconomic disadvantage (Barber, 2001) with consequences for individuals' well-being and further life course (Barber, 2001; Wiik, 2009). Therefore, the analysis of the influence of parents' socio-economic background on union formation and transition to parenthood ultimately improves the understanding on persistence of social inequality. Unions formed earlier have higher chances of disruption (Berrington & Diamond, 1999; Lyngstad, 2006), cohabiting unions tend to be less stable than marital unions (Beck, et al., 2010; Graefe & Lichter, 1999; McLanahan, 2011), and children born in cohabiting unions are more likely to end up living with a single mother compared to children born in marital unions (Manning & Brown, 2006). Further, the timing of childbearing has serious implications for the material and psychological wellbeing of parents (Hoffman, 1998; Upchurch & Mccarthy, 1990). Early age at first birth and out-of-wedlock childbearing are linked to subsequent adult social disadvantage (Chase-Lansdale et al., 1995; Duncan et al., 1998; Hardy et al., 1998; Hobcraft, 2008; Kiernan & Hobcraft, 2001).

Benefitting from a large sample size from the UK Household Longitudinal Study, which amounts to almost 35,000 individuals, I assess the link between socio-economic background on the transition to adulthood of children born between 1930 and 1980. In addition to the main hypotheses, I assess whether the influence of parents' education and social class has varied across the cohorts, and over the children's life course.

2. Background

2.1 Background and transition to union

Previous research extensively addressed the influence of parents' background on timing and occurrence of a number of life events, including cohabitation (e.g., Wiik, 2009; Moyaart & Liefbroer, 2016), marriage (e.g., Micheal & Tuma, 1985; Axinn & Thornton, 1992; Wiik, 2009), first birth (e.g., Micheal & Tuma, 1985; Barber, 2001).

Overall, the existing evidence points out that individuals from advantaged social background (in terms of parents' high social class, occupations, or education) tend to experience trajectories of transition to first union at a slower pace relative to individuals with poorer social (Axinn & Thornton, 1992; Barber, 2001; Blossfeld & Huinink, 1991; Clarkberg, 1999; Dahlberg, 2015; Michael & Tuma, 1985; Mulder, et al., 2006; Rijklen & Liefbroer, 2009; Sassler & Goldscheider, 2004; South, 2001; Wiik, 2009). However, the evidence about the association between parents' background and first cohabitation is smaller and more mixed. Studies on the United States found the transition to cohabitation (vs. direct marriage) associated with parents' more advantaged background (e.g., Clarkberg, 1999; Cohen & Manning, 2010; Lichter & Qian, 2008) with more disadvantaged background (e.g., Bumpass & Lu, 2000; Manning & Cohen, 2015) or reported no significant effect (Sassler & Goldscheider, 2004; Xie, et al., 2003).

A few studies in Europe have simultaneously compared the effects of background on timing of first marriage versus first cohabitation for men and women. In Norway, Wiik (2009) found that only respondents entering cohabitation were (negatively) affected by parental education (especially mothers). In the Netherlands, Moyaart & Liefbroer (2016) found that father's and mother's education is associated with earlier entry into cohabitation as opposed to marriage. In Italy, (Schröder, 2006) found parents' education positively associated with women's transition to cohabitation and negatively associated with marriage. In Bulgaria, Hoem & Kostova (2008) found the opposite association between mother's education and daughters' risk of experiencing cohabitation rather than marriage.

These conflicting results might be primarily explained by the diversity of the country context and by the use of different indicators of parents' socio-economic background. In fact, Hoem & Kostova (2008), Moyaart & Liefbroer (2016) and Schröder (2006) adopted parents' highest educational attainment, while Wiik (2009) used an indicator of parents' social class. A multidimensional perspective into family social background would thus help untangle the different underlying mechanisms of parents' background influencing children's transition to first union.

2.2 Background and transition to parenthood

Most studies addressing the association between parents' education and children's parenthood focused on timing (e.g., Barber, 2001; Dahlberg, 2015; Blossfeld & Huinink, 1991; Baxter et al., 2008; Lappegård & Rønsen, 2005; Michael & Tuma, 1985; Reeder, 2013; Rijken & Liefbroer, 2009; Hofferth & Goldscheider, 2010; Steenhof & Liefbroer, 2008; Upchurch & Mccarthy, 1990). Overall, the findings highlight that the likelihood of early parenthood declines as family resources (broadly defined) increase (Aassve, 2003; Barber, 2000, 2001; Ravanera & Rajulton, 2006).

A line of research analyzed the predictors of early or non-marital parenthood and identified socioeconomic disadvantage as a particular risk factor (Aassve, 2003; Carlson, et al., 2013; Fomby & Bosick, 2013; Högnäs & Carlson, 2012; Manning & Cohen, 2015), also in the UK, where childhood disadvantage is associated with earlier entry into parenthood and with parenthood occurring in less favourable partnership contexts (Buxton et al., 2005; Chase-Lansdale et al., 1995; Hobcraft, 2008; Kiernan & Hobcraft, 1997). Among the studies concentrating on first childbearing beyond teenage and early adulthood (Barber, 2000, 2001; Dahlberg, 2015; Rijken & Liefbroer, 2009; Steenhof & Liefbroer, 2008), only Barber (2001) and Dahlberg (2015) included indicators for both parents' education and social economic status and only Dahlberg (2015) controlled for the index of person's own education attainment.

However, only few studies addressed the partnership context (marriage vs. cohabitation) of first parenthood of women (Aassve, 2003), men (Carlson et al., 2013) or both (Barber, 2000; Barber, 2001; Hobcraft, 2008). Father's higher education is associated with a lower risk of a non-marital first birth of daughters (Aassve, 2003; Upchurch, et al., 2002; Musick, 2002) and sons (Carlson et al., 2013) and, to a lesser extent, to sons' marital birth (Carlson et al., 2013). Mother's education reduces both the sons' and the daughters' rates of premarital birth (Barber, 2001) but has a marginal negative effect on daughters' marital birth and a non-significant influence on sons' marital birth (Barber, 2001). With respect to social status, Hobcraft (2008) found excess risk of becoming a parent up to age 25 for married men and women, who experienced economic deprivation in childhood, compared to their counterparts who did not.

These findings can be explained by the fact that children born in disadvantaged families are more likely to leave the parental home earlier and thus begin family formation earlier, while young children from better-off families are more likely to delay union and childbearing, and particularly less likely to experience non-marital first births (Aassve, 2003). In other words, young women who have little hope of social or economic advancement are more likely to choose childbearing at younger ages and outside of marriage (Edin & Kefalas, 2005). Nevertheless, no previous study addressed the joint influence of parents' educational background and social class on the union context in which parenthood occurs.

To date, although prior articles paid attention to social background as a predictor of the timing of union and childbearing, previous research lacks a comprehensive view of how different dimensions of social background affect union and fertility timing across the reproductive ages of men and women, and a thorough understanding of the mechanisms leading to a first marital or non-marital union and parenthood, for both women and men. These gaps limit the comprehension of which aspects of social background matter, the mechanisms through which they have an impact on those outcomes, and whether their influence varies for men and women, over the life course, and across cohorts.

3. **Hypotheses & mechanisms**

3.1 Social Background and transition to adulthood: Theoretical framework

Most sociological research addressing the influence of family background on children transition to adulthood adopts two theoretical approaches: the 'opportunity cost'/'rational agent' frameworks theorized by (Becker, 1960) and the socialization hypothesis first introduced by McLanahan & Bumpass (1988) in the socio-demographic literature. In Becker's model, individuals decide to enter unions in response to economic incentives, such as parents' material and nonmaterial resources, which influence their position in the marriage market. Likewise, the demand for children, who are viewed as 'durable good', is positively associated with income, which is severely dependent on resources inherited by the family of origin.

The socialization theory provides an alternative view about the association between parents' background and children's transition to adulthood (Hynes, et al., 2008; Mclanahan & Bumpass, 1988; Plotnick, 1992; Scaramella, et al., 1998; Wu & Martinson, 1993). By socialization of values, previous studies mean *conscious and unconscious learning* (Murphy & Wang, 2001) and *transmission of values and preferences* (Barber, 2000) from parents to children. Children's own preferences on when to form the first union and whether to become parents, is either directly (De Valk & Liefbroer, 2007) or indirectly (Wiik, 2009) a product of their parents' preferences. Of course, this does not rule out that parents have preferences that are independent of, and perhaps in conflict with, the child's own desires (Wiik, 2009). Therefore, parents' educational achievement and socioeconomic status mirror the values that are salient for shaping children's preferences (e.g., Axinn & Thornton, 1993; Barber 2000; Hofferth & Goldscheider, 2010; Wiik, 2009; Dahlberg, 2015).

As Dahlberg (2015) argues, parents' educational attainment and socioeconomic status are two *complementary* and not *mutually exclusive* dimensions of socioeconomic stratification on the intergenerational transmission of fertility and parenthood behavior. Most research has adopted a variety of indicators of socio-economic background (e.g., educational attainment, social class, income, indicators of economic deprivation) without discussing the theoretical justification of each factor. Recently a growing body of research in stratification has addressed more specifically *which* measures capture *which* aspects of parents' social background (Buis, 2013; Bukodi, 2012; Chan & Goldthorpe, 2007; Chan & Goldthorpe, 2004; Dahlberg, 2015, Geyer, et al., 2006). On the one hand, there is large consensus that, education captures cultural differences and abilities in embracing knowledge and socializing norms with the children (Dahlberg, 2015). On the other hand, class encompasses economic well-being and income prospects (Goldthorpe & McKnight, 2006), employment conditions, such as differences between manual and non-manual work (Dahlberg, 2015; Erikson & Goldthorpe, 1992). In the following paragraphs, I motivate parents' class and education as separate dimensions of social background, and I discuss the pathways of transmission to union formation and parenthood.

3.2 Parents' education and union formation

Parental education embodies the human capital that is crucial for the transmission of advantage from one generation to the next (Becker, 1993). Parents' educational attainment is a close proxy of parents' attitudes and beliefs about family life which, through socialization, influence how their children want to behave (Wiik, 2009). Children with more educated parents may be socialised differently than children with low-educated parents. First, highly educated parents should more able at persuading them to postpone early union formation (Axinn & Thornton, 1992) because they might be more aware of potential negative consequences of choices made in the early life-course (Farkas, 2003). Also, educated parents could nurture children's higher career aspirations (Dubow, et al., 2009; Mooyaart & Liefbroer, 2016; Schoon & Parsons, 2002), induce them to extend their schooling and, hence, delay their first co-residential union. Finally, highly educated parents might provide their children with better support during scholarship (Lareau & Cox, 2011) and, ultimately, ease a slower entry into the labour market (Van Winkle, 2018). On the contrary, lower educated parents might be not particularly motivated to persuade their children to invest in education and promote ambitious educational prospects.

These mechanisms particularly apply to the link between parents' education and age at marriage of their offspring (Micheal & Tuma, 1985; Blossfeld & Huinink, 1991). Conversely, the influence of parents on the timing of first cohabitation is less evident because cohabitation is more easily reversible than marriage. If cohabitation is perceived as a temporary and more informal arrangement, parents would not interfere much with their children's decision. Instead, if cohabitation is selective of more individualistic and non-traditional individuals (Smock, 2000 → Wiik, 2009), in line with the Second Demographic Transition narrative, parents' higher education could be associated with less traditional pathways – such as entry into cohabitation vs. marriage –because parents socialise with their children more liberal values (Cohen & Manning, 2010; Lichter & Qian, 2008; Moyaart & Liefbroer, 2016).

Based on these arguments, *I expect that parents' education has a different influence on children's transition to marriage and cohabitation, and that higher-educated parents might delay the transition to marriage as opposed to cohabitation* (hypothesis 1).

3.3 Parents' education and parenthood

Parents' education can have a direct influence on the fertility through the socialization of norms, values and preference for the timing of transition to parenthood (Barber, 2001). A wealth of studies also concentrated on the multifaceted indirect influence of parents' educational background on children' transition to parenthood. First, as the timing entry into parenthood is affected by young adults' educational attainment (Blossfeld, & Huinink, 1991), the intergenerational transmission of fertility (Steenhof & Liefbroer, 2008)might operate through the intergenerational transmission of education (Blau & Duncan, 1967). Second, as union formation is a strong determinant of the transition to parenthood (Perelli-Harris et al., 2012), the tendency to postpone the first union among the young adults from better-off families (e.g., Wiik, 2009) might imply the postponement of first parenthood as well. Conversely, disadvantaged young adults, deprived of the opportunity to access higher education (Osgood, Ruth, Eccles, Jacobs, & Barber, 2005), often look to family early parenthood as an attainable marker of adulthood (Cherlin, Cross-Barnet, Burton, & Garrett-Peters, 2008; Edin & Kefalas, 2005).

England, et al., (2011) argue that low-educated parents might not effectively socialize knowledge about contraception and family planning with their children, who might incur more easily in early or unintended births. Further, Barber (2001) points out that mothers tend to socialize with children their preference for birth timing within marriage, since non-marital first births tend to occur more often "by accident" (Brown & Eisenberg, 1995) as opposed to marital first births, which need more planning. In other words, parents' preferences about childbearing are more likely to affect marital rather than pre-marital childbearing behavior because parents tend to socialize their attitudes about marital rather than non-marital childbearing (J. S. Barber, 2000; Jennifer S. Barber, 2001; Barber & Axinn, 1998). Therefore*, I expect that higher-educated parents might more effectively delay children's transition to parenthood within marriage than parenthood within cohabitation* (hypothesis 2).

3.4 Parents' social class and union formation

A well-off family fosters young adults' marriageability by signaling its wellbeing through its status symbols (Becker, 1993). Indicators of family advantage, such as higher parental income, along with a limited number of siblings and an intact family structure (Axinn & Thornton, 1992), increase the absolute advantage of remaining single and lower the advantage of entering a co-residential union (whether marital or non-marital) earlier.

Children raised in a wealthy context develop the same consumption aspirations as their parents (Easterlin, 1980) and might be willing to leave their household, and hence form a union, only when their standard of living matches their consumption aspirations (Axinn & Thornton, 1992). Also, wealthy parents are more likely to have the financial wherewithal to keep their children off premature unions (Axinn and Thornton, 1992) or, conversely, to ease their children's settlement in a union by offering financial help (Mulder & Smits, 1999).

Conversely, at the low end of the income distribution, young adults with few parental resources are more likely to form co-residential unions early because they may have relatively more to gain from pooling their resources with a partner (Meier & Allen, 2008), or because they find parents' household less attractive (Michael & Tuma, 1985).

Another mechanism might predict opposite pathways to first union for low and high-class offspring. Parents' economic resources can promote early transition to first union by providing financial stability through parents' safety net (Dahlberg, 2015). Therefore, well-off parents might (unintentionally) increase their children's motivation for leaving home early. On the contrary, the children from disadvantaged families might be dissuaded from forming a new household early.

Marriage is commonly perceived as a more stable arrangement and long-term commitment as opposed to cohabitation (Lesthaeghe, 1983), and parents' ability to make it economically viable is determinant (Wiik, 2009). Further, the search process for a marital partner has higher "transaction costs", especially in periods of economic hardship or uncertainty (Oppenheimer, 1994, p. 308), and dissolving a marriage is costlier than dissolving a cohabitation (Wiik, 2009). As cohabitation provides the same benefits of marriage, such as companionship and resource pooling, young adults from disadvantaged background might find cohabitation a more convenient option for their first union. Under these assumptions, *I hypothesize that parents' high social class should encourage young adults to remain longer in the parental home if the hypotheses of "marriageability" and "higher occupational aspirations" hold true* (hypothesis 3a). *I also hypothesise that children from well-off families would be willing to their household earlier if the "safety net" assumption prove true (*hypothesis 3b).

Finally, I assume that advantaged parents tend to delay especially children's entry into marriage. Conversely, lower SES children would result in higher propensity to opt for unmarried cohabitation rather than direct marriage (hypothesis 3c).

3.5 Parents' social class and parenthood

In an economic perspective, the choice of having children in a marriage (as opposed to cohabitation) reveals that gains to marriage are larger (Willis, 1999). In other words, persons tend to have a child when their utility in a family is higher than if they remain single. Marriage reveals a preference for a long-term relationship, which nurtures stronger emotional bonds and is more efficient as far as consumption choices and children's upbringing are concerned. Conversely, entering out-of-wedlock parenthood, and possibly at early age, might be considered more attractive to individuals with limited future economic prospects or few expectations from marriage due to social exclusion (Aassve, 2003), in spite of valuing children as much as or more than (Edin & Kefalas, 2005) their counterparts with greater resources.

In a search-theoretical framework, young adults spend time looking for a possible match, and a union is formed when the potential partner guarantees an acceptable level of utility (Oppenheimer, 1988). Persons from higher socio-economic backgrounds might be particularly selective in order to preserve the socio-economic level of their family of origin. If people enter a union early, before their education is completed, the information about the mate's earning potential and perspective class is not fully disclosed (Oppenheimer, 1988). In the short term, cohabitation, whose dissolution costs are lower than those of marriage, may be an alternative to a costly search for a long-term partner for people form more advantaged families. Only when high-class individuals end schooling and have a firm foothold in the labour market, they might search for a better match and plan parenthood in a solid arrangement such as marriage. In other words, the latter might experience marital parenthood only when they meet partners with good prospects. Therefore, it is the risk aversion to 'downward mobility' that might lead advantaged young adults to delay parenthood to find a match that maximizes their chances of upward mobility or class stability. Marital parenthood, as opposed to out-of-wedlock childbearing, would be just more coherent with their view of utility maximization in the long run.

In the light of these arguments, *I expect that the pathways to parenthood of children from disadvantaged background are most likely to occur outside of marriage and those of children from less disadvantaged origin preferably occur within marriage* (hypothesis 4).

3.6 Cohort variability

Historical changes in union formation and stratification processes might have led to a decay of the influence of family background on offspring's transitions (South, 2011). The major cause could be attributable to the shift in values from "solidarity and social group adherence" in favour of "autonomy and self-realisation", as hypothesised by the SDT theory (Lesthaeghe, 2010; van De Kaa, 2002). In general, the role of parents in stirring children's life-course choices has become less prominent: parents have become less authoritative and have increasingly put more emphasis on an upbringing punctuated by stimuli and freedom rather than by discipline. In the United Kingdom, such as in many other Western Countries, traditional gender roles and norms have been increasingly questioned and challenged over time (Hobcraft, 2008).

These changes, along with the technological burst, led to increasing female employment and a narrowing of educational gap between the high educated and the low educated. Education expansion has enabled adults to provide for themselves without requiring the use of parental resource, has strengthened the social acceptance of alternative arrangements such as cohabitation, and formed the critical mass of young adults across all social who embraced it in the 1980s (Murphy, 2000). It appears legitimate to argue that *parents' influence has ultimately decreased across the cohorts* (hypothesis 5).

3.7 Life course variability

Family is expected to influence mainly adolescents and young adults, who are more exposed to their parents' socialisation and also have less resources to become independent (South, 2011). For instance, parents who deter their children from attending college, or who are unable to support their graduate studies, limit their option in later adulthood and might indirectly promote an early union. Conversely, parents who spur their child to take on graduate studies to favour a strong foothold into the labour market are potentially postponing (or limiting) their children's union and childbearing prospects (Barber, 2001). Parents' role is also crucial when their children approach early adulthood to guarantee the financial stability for a marriage. Nevertheless, as children start their professional career, gain more economic stability, and possibly settle down in distant places from their origin families, the effects of family resources, in particular education (Axinn & Thornton, 1992) should weaken and eventually disappear completely. Ultimately, *the influence of parental education and social class tend to decrease over children's life course* (hypothesis 7).

3.8 Other factors

Gender differences

Opportunity costs for family formation may differ by gender because of traditionally different role models of women and men. The typical role of men as an economic provider for the family has as a salient consequence that women enter union and have children earlier than men (Goldscheider & Waite, 1986; Uecker & Stokes, 2008; Winkler-Dworak & Toulemon, 2007; (Hynes et al., 2008)). Few studies have weighed up the influence of parental background by gender and only Michael and Tuma (1985) found stronger effects for women than for men. High educated parents may want their daughters to postpone family formation to benefit from better paid jobs and the consequent easier work-family reconciliation (Barber 2000; Wiik 2009). However, Axinn & Thornton (1992) in the United States, Hobcraft (2008) in the UK and Wiik (2009) in Norway did not find substantial gender differences. Limited research on men's fertility points out that men with fewer family resources are more likely to experience early fertility than their more advantaged counterparts (Baxter et al., 2008; Glick, Ruf, White, & Goldscheider, 2006; Pears, Pierce, Kim, Capaldi, & Owen, 2005), while evidence on union transitions in the UK showed that men with fewer opportunities are less likely to form a partnership of any kind (Bukodi, 2012).

Own education

Parents' education attainment might be indirectly conducive of children' partnering and childbearing behaviour because children are expected to pursue their parents' educational achievement (Michael & Tuma, 1985). Wide and consistent evidence showed that educational level and enrolment is one of the most influential factors that delay the timing of first union and first parenthood (Blossfeld and Huinink 1991; Michael & Tuma, 1985; Toulemon & Winkler-Dvorak, 2007). Educational attainment delays marriage because it delays the transition to a stable work and, consequently, the time when individuals enter the marriage market (Oppenheimer, 1988). Cohabitation, instead, requires less financial commitment and, possibly, fewer couple-oriented activities, it is more compatible with student life (Thornton, Axinn, & Teachman, 1995), and has been seen by some as a response to the delay in marriage, facilitated by the availability of modern contraceptives (Oppenheimer, 1988). Further, in a search model of partnering, individuals look for a suitable match in terms of socio-economic resources and form a union with a partner that is above the threshold of acceptability (Oppenheimer, 1988). Young adults from advantaged background set their cultural and economic standards on the basis of those of their family and tend to discard potential parents that do not match that benchmark. In case individuals entered a first union early, before completing their studies and having a foothold in the labour market, they would choose a partner before her socio-economic status is fully disclosed. For this reason, risk-averse individuals from advantaged strata, aiming at preserving their family background, avoid "early partnering" and find "later partner search" optimal (Oppenheimer, 1988; Wiik, 2009).

Other features of family background: parents' union dissolution

Robust evidence for many other Western countries, including the UK, highlights the intergenerational transmission of partnership behaviours (Amato, 1996; Kiernan, 1992; Liefbroer & Elzinga, 2012): individuals whose parents separated are more prone to entry into union and parenthood at a younger age, and particularly into cohabitation and non-marital parenthood (Michael & Tuma, 1986; Cherlin et al., 1995; McLanahan and Bumpass 1988; McLanahan and Sandefur 1994; Kiernan 1992; Upchurch et al. 2002; Kiernan 2004; Hofferth and Goldscheider, 2010, Fomby & Bosick, 2013; Hognas & Carlson, 2013). Family instability influences the levels of material and emotional support available in childhood and teenage, such as household financial stability (Wu, 1996), parents' mental and physical health (Osborne, Berger, & Magnuson, 2012) and the quality of parent–child relationships (Cavanagh, 2008).

Growing up in a non-intact family might shape – even permanently – children's negative attitudes toward marriage (Axinn and Thornton, 1996; South, 2001), and develop normative attitudes and beliefs that may subsequently contribute to their likelihood of having a nonmarital first birth (Hognas & Carlson, 2013). Second, parents' separation might speed up the departure of children, affected by a conflictive parent-child relationship (Raab, 2017), in search for emotional support and intimacy outside of the family of origin (Wu & Martinson, 1993). Third, as unmarried or single mothers are assumed to be less able to control their children's behaviour, the young adults are more likely to engage in risky sexual behaviour and experience early (non-marital) pregnancy (McLanahan & Sandefur, 1994; Michael & Tuma, 1985)

4. Data and methods

4.1 Sample

I use data from Understanding Society, a survey that started in 2009-2010 with a nationally representative sample of roughly 43,000 individuals. The survey collects contemporary and retrospective information of interviewees' employment, partnerships and fertility history and has a longitudinal design, with interviews collected annually. Individuals' fertility histories are drawn from individuals' reports that recall the date of birth of each child, the possible date of departure from the household and the motivation (e.g., death or a separation). This means that any transition taking place within work and family domains is recorded to the nearest month. Also, retrospective information may contain a higher level of recall error compared to the panel where every piece of information refers to the previous year. I use a six-month time scale to construct the sequence-type representation for life-course trajectories during early adulthood. The survey is based on probability sampling techniques that assure they are nationally representative. The analyses do feature sampling and non-response weights. The analytical sample includes 35,844 individuals (19959 women and 16084 men) born between 1930 and 1980.

In the survey individuals are asked to report the start and end dates (in year and months) of all their marital and cohabiting relationships, their childbirths, and age at which they attained their highest qualification and whether (and when) they had additional schooling in their life. The data were organised into two longitudinal histories: a "union dataset" containing information on their first marital or cohabiting relationship, and a "fertility dataset" with information about their childbearing, along with educational histories, parental background, and other time-varying and time-invariant characteristics.

4.2 Measures

I identify outcomes in distinct models for transition to union and to parenthood. The dependent variable are two indicators capturing when (a) the first marriage or a cohabitation occurs, and (b) the first conception occurs within a marriage or outside a marriage, in a given interval.

The principal explanatory variables are two dimensions of parents' background: parents' educational attainment and social class. The highest level of both *parents' education* – as reported by the respondent at the age of 14 – is used to determine parents' education with four categories (1 = *Degree*, 2 = *High school*, 3 = *Some schooling*, 4 = *No school*). A residual category captures missing information about both parents' education. *Parents' social class* is coded according to a version of the National Statistics Socio-Economic Classification (NS-SEC), following Bukodi & Goldthorpe's (2015) strategy. Eight levels are derived: 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*. In accordance with the dominance method (Erikson, 1984), in the current model specifications, I use social class origin indexed as the highest among father's and mother's class at respondent's age 14. Again, an indicator for missing information is included. This method of identification of social class is coherent with cross-cohort comparison, according to Bukodi & Goldthorpe (2015) who applied it to the same data source.

Education of the respondent is measured both as educational attainment and educational enrollment, both time-varying. Respondent's highest level of education is measured as the highest level of completed education through six distinct ordinal categories: 0 = *no qualification*, 1 = *qualification lower than GCSE*, 2 = *GCSE*, 3 = *A-level*, 4 = *Higher vocational qualifications (mainly nursing and teaching)*, 5 = *Degree qualifications*.

The variable *age* is constructed as the number of years and months since age 15. This variable captures the duration dependence of the estimated hazard. To allow for non-linearity, a quadratic term for age is added to the equation in addition to the linear specification. The variable *cohort of birth* addresses any temporal trends in the chance of first union formation and parenthood, as well as the changes in the choice between partnership and childbearing arrangements. This item is measured continuously in years with values ranging from 1930 (the baseline year) to 1980. I interact the indicators of socio-economic background with age and cohort of birth to study more in detail the possibly differential impact of the key explanatory variables over individuals' life course and across decades, respectively.

Ethnicity is represented with a set of five categories: 1 = *White (reference)*, 2 = *Indian/ Bangladeshi/Pakistani*, 3 = *African Black & Caribbean*, 4 = *Asian*, 5 = *Other*. According to previous evidence, ethnic group is a relevant factor in influencing the age at first union formation (e.g., Michael & Tuma, 1986), first birth (Schoen et al., 2009; Reed, 2014) and non-marital childbearing (Cherlin et al. 2008; Edin and Kefalas 2005; Hobcraft, 2008).

Family structure is proxied by an indicator capturing whether respondent's parents had separated by age 14 (0 = *intact family*, 1 = *separation*); two dichotomous variables reflecting whether the mother or the father left the household or died before individual's fourteenth birthday (0 = *mother/father present*, 1 = *mother/father died or left the household*); and a categorical variable accounting for the family size (0 = *no siblings*, 1 = *1 sibling*, 2 = *2 or more siblings*). The presence of siblings could lower family resources (Michael & Tuma, 1985), thus, moderating the impact of socio-economic background, and because children who grow up in large families with many siblings might develop the desire to enter union and/or parenthood earlier than their peers, through a socialisation process (Michael & Tuma, 1985; Barber, 2001).

The variable *economic growth* is measured as a 12-month lagged GDP variation on a yearly basis, starting from 1949, the year after the Office for National Statistics first estimated the UK GDP. Another proxy of the economic cycle used in earlier analysis is the level of the *house prices* to address one of the macro-economic factors affecting couples' decision to move in (Clark, 2012). This item is operationalised as the 12-month lagged logarithm of the average price per square yard in each specific region. For descriptive statistics of the variables, see Table 1.

4.3 Analytical strategy

I applied multinomial discrete time event history models (Hoffman and Duncan, 1988) to estimate the hazard of the timing of the following transitions for men and women. The dependent variable has three possible outcomes in both analyses: (a) marriage or cohabitation vs. remaining single; (b) marital birth or non-marital birth vs. remaining childless.

The probability modelled in the multinomial logit model is the conditional probability of marital or non-marital union (or birth) versus none in a given month, since no outcome occurred beforehand. Two equations are estimated simultaneously in each analysis: (a) the log-odds of marriage versus no union formation, the log-odds of cohabitation versus no union formation; (b) the log-odds of marital birth versus no birth, the log-odds of non-marital birth versus no birth.

In the 'union dataset', young adults are exposed to the risk of first union at the onset of their fifteenth birthday until they start a marriage or a cohabitation. The same occurs in the 'fertility dataset', where the respondents are exposed to the risk they experience a pregnancy in a marital or non-marital union, nine months before the birth. More specifically, only first births that occur less than 9 months after marriage, or before marriage, are considered conceived out-of-wedlock. For each month before an event occurs, the dependent variable is coded as '0'; for the month in which a transition to union or parenthood occurs, it is coded as '1'. If the respondents do not experience any event before age 45, or in case of death, the observations are censored, in both samples.

My approach uses a hierarchical modeling strategy, by adding groups of variables to a baseline model so that a model is nested within the subsequent models. The goal is to highlight the existence of direct and indirect effect of the two key dimensions of social background on the outcomes of interest and assess whether and to what extent it is moderated by other factors. I correct the results for the non-independence of the observations within unions by a robust cluster variance estimator.

Table 2.1. Descriptive statistics

Variables	Women	Men
	N	
Number of individuals	19959	16084
Number of first marriage	11691	8725
Number of first cohabitation	6793	5755
Number of first birth in marriage	7431	5200
Number of first birth outside marriage	3577	2073
	Percent (%)	
Cohort of birth		
1930-39	9,7	10,9
1940-49	16,8	17,9
1950-59	19,6	19,5
1960-69	25,5	24,4
1970-80	28,4	27,3
Parents' higher social class		
Higher managers & professionals	12,3	12,7
Lower managers & professionals	9,5	9,8
Intermediate occupations	10,5	9,9
Small employers and self-employed	9,6	9,8
Lower supervisory and technical occupations	15,7	15,6
Semi-routine occupations	17,4	17,2
Routine occupations	11,5	11,5
Unemployed	8,7	8,1
Missing	4,9	5,4
Parents' higher education		
Degree	7,6	7,4
Some qualification	20,2	18,0
Left with some qualification	17,0	17,6
Left with no qualification	30,5	29,2
No school	1,7	1,8
Missing	23,0	25,9
Own education		
Degree	18,9	23,7
Other higher	12,4	9,8
A level etc	14,8	19,4
GCSE etc	21,1	16,8

Other qual	11,0	12,3
No qualification	21,2	16,6
Missing	0,6	1,3
Parents separated at 14	4,1	3,6
Mother is absent/dead at 14	2,7	2,5
Father is absent/dead at 14	8,6	7,9
Number of siblings		
0	16,5	17,5
1	27,8	28,2
2 or more	55,7	54,3
Rural area	22,7	21,8
Ethnicity		
European	83,5	81,1
African - Caribbean	5,7	4,8
Indian, Pakistani, Bangladeshi	7,0	9,0
Other Asian	1,7	1,7
Other	2,2	3,4

5. Results

5.1 Descriptive results

Figures 2.1 through 2.4 show the hazards of entry into union and parenthood by parents' highest social class (red) and highest educational level (orange). Figure 2.1 shows that men and women whose parents have higher levels of education and social class enter first marriage at later age. Conversely, individuals with parents from lower social class and education have a higher hazard of early cohabitation as opposed to their counterparts from higher background (Fig. 2.2). Women whose parents have a degree have a lower hazard of experiencing first marital motherhood up to age 35 compared to all other counterparts (Fig. 2.3). For men, this crossover occurs after age 38. A postponement effect, which is not compensated for at later ages, is also visible for individuals raised in higher social classes. The distribution of hazard of entry into out-of-wedlock parenthood (Fig. 2.4) does not strictly follow any educational or social gradient and is characterized by a low kurtosis. Women with low educated parents have a higher risk of becoming parents in teenage years or slightly after coming of age. For men, parental education has a similar influence, albeit more moderated. Interestingly, women from two of the highest social classes ('high managerial positions' and 'self-employed') display the highest risk of non-marital first birth after age 32, when the hazard of the same event is much lower for the other women. These findings show that out-of-wedlock motherhood consists of two very distinct pathways: a teen-age transition that not always occur in an established union (Kiernan, 2006, check it out), and a relatively late event for those women who possibly see cohabitation as an alternative arrangement to marriage.

Figure 2.1. Entry into marriage as first union, by parents' background. Women and men.

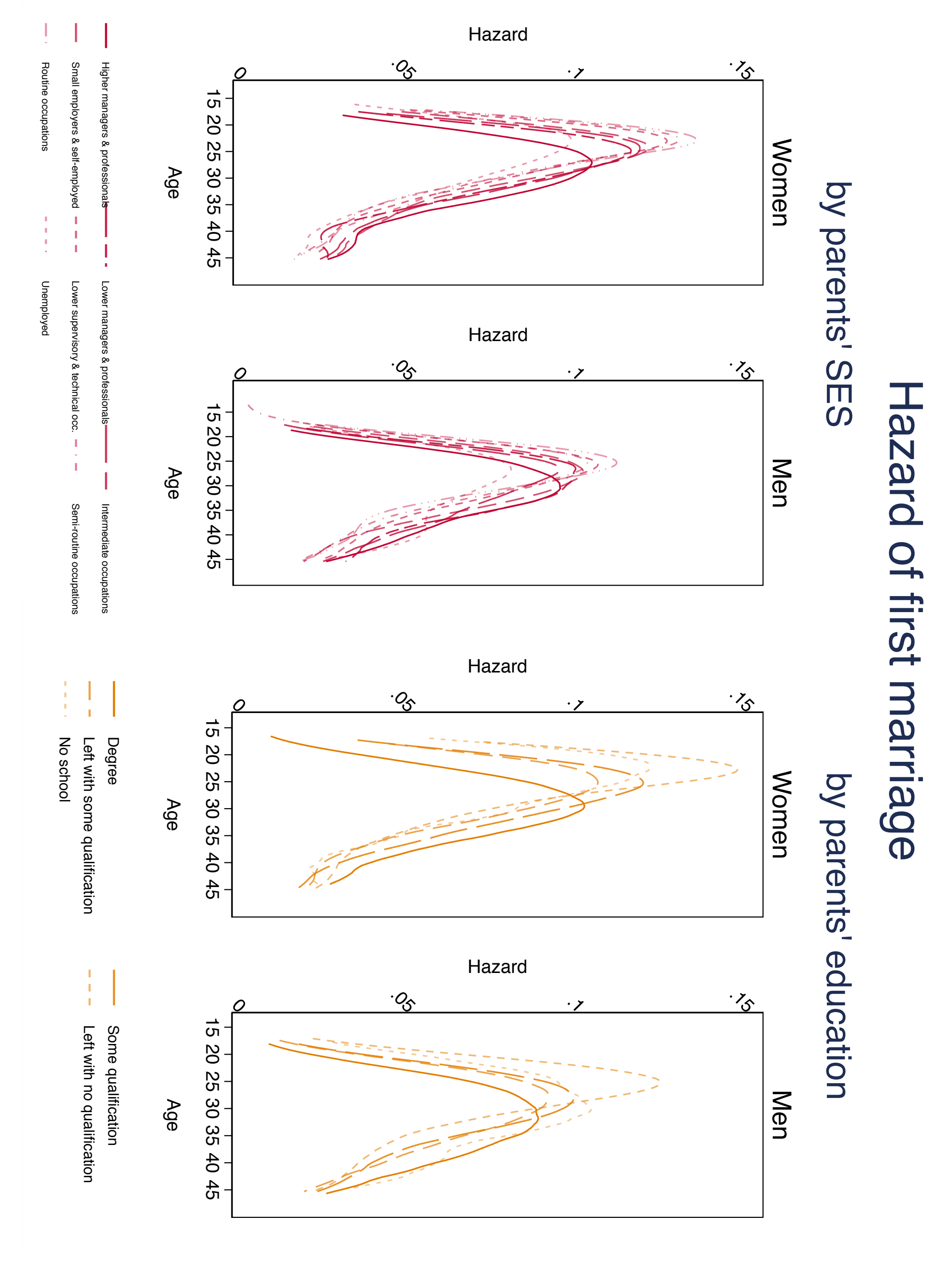


Figure 2.2. Entry into cohabitation as first union, by parents' background. Women and men.

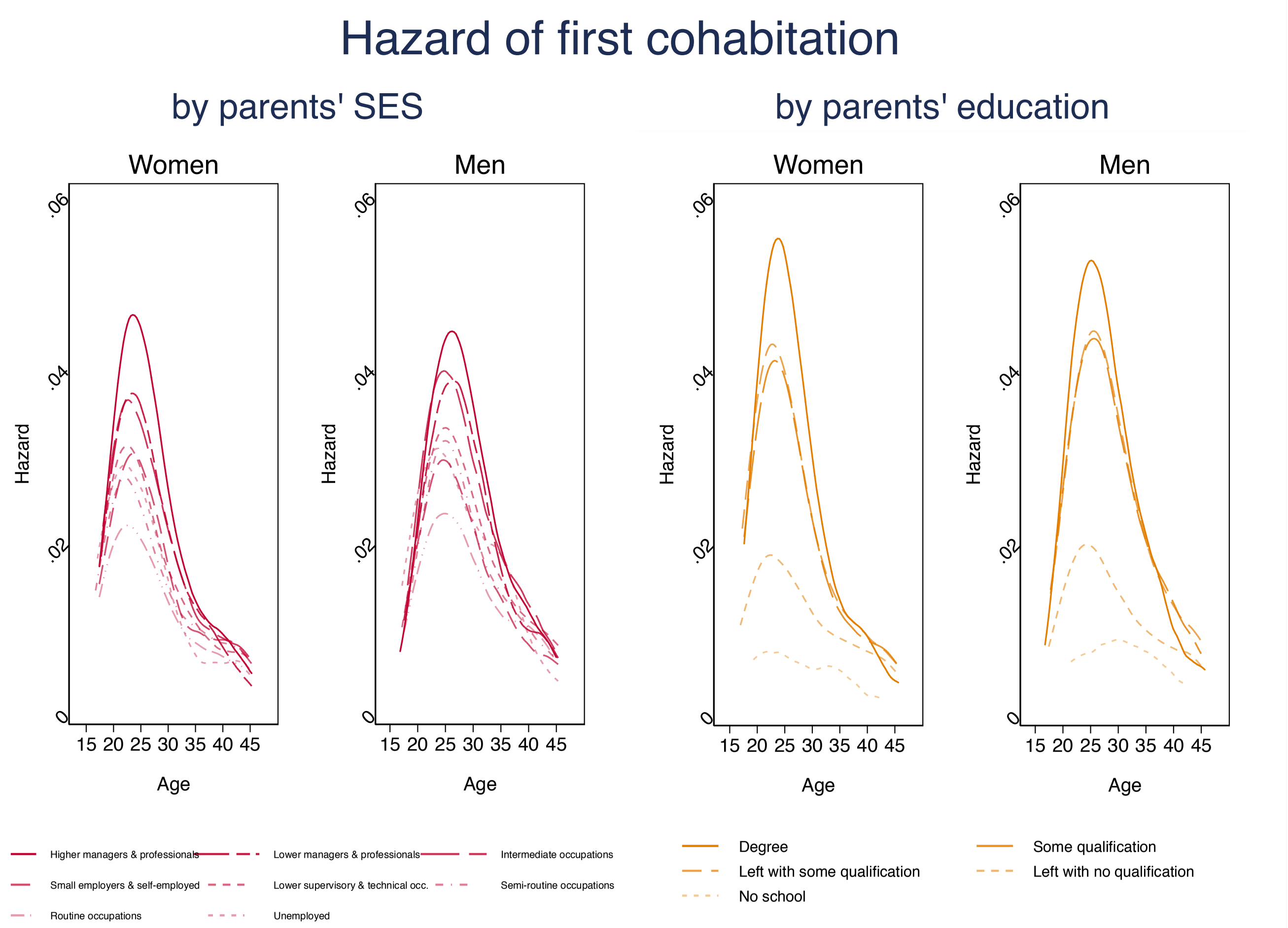


Figure 2.3. Entry into marital parenthood, by parents' background. Women and men.

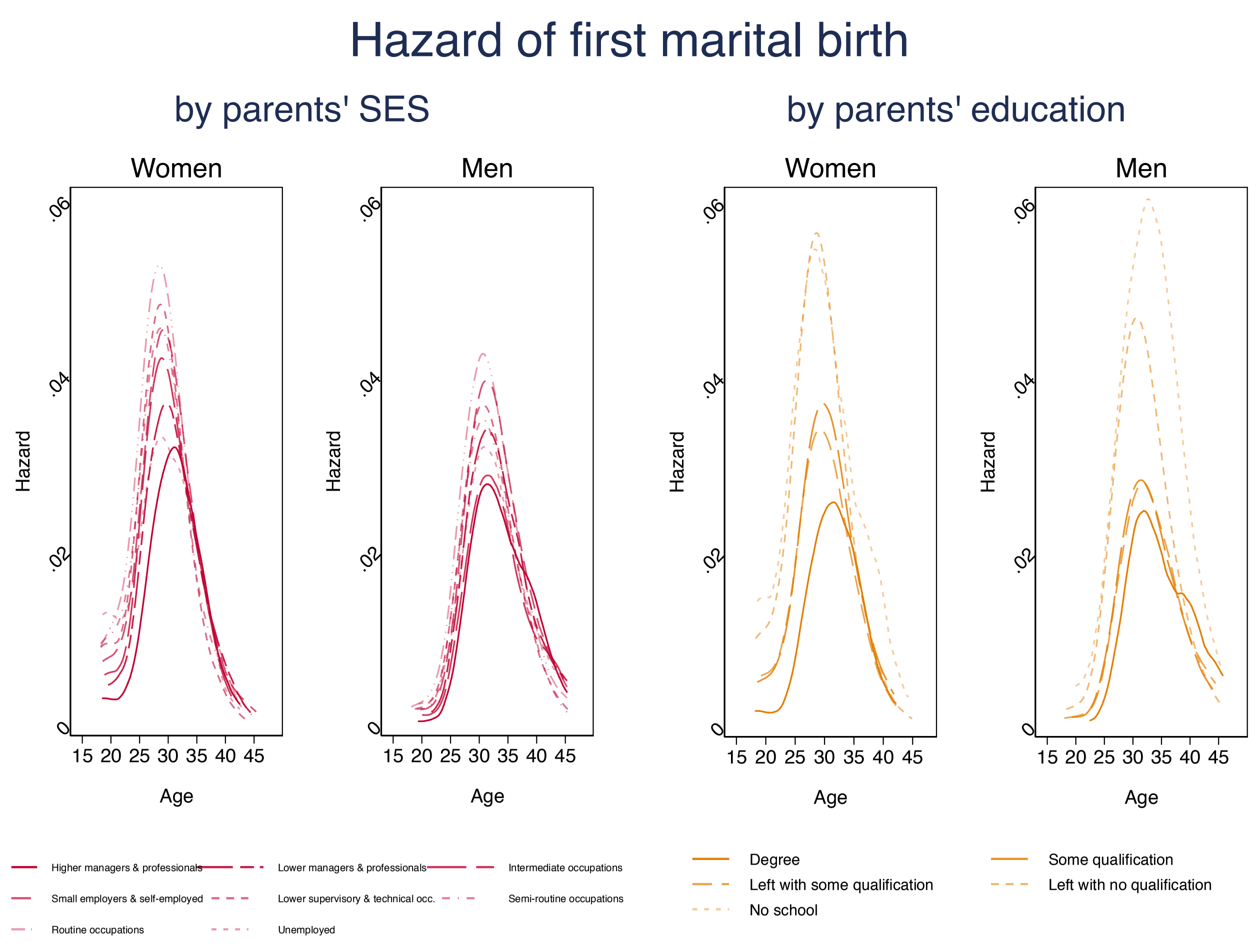
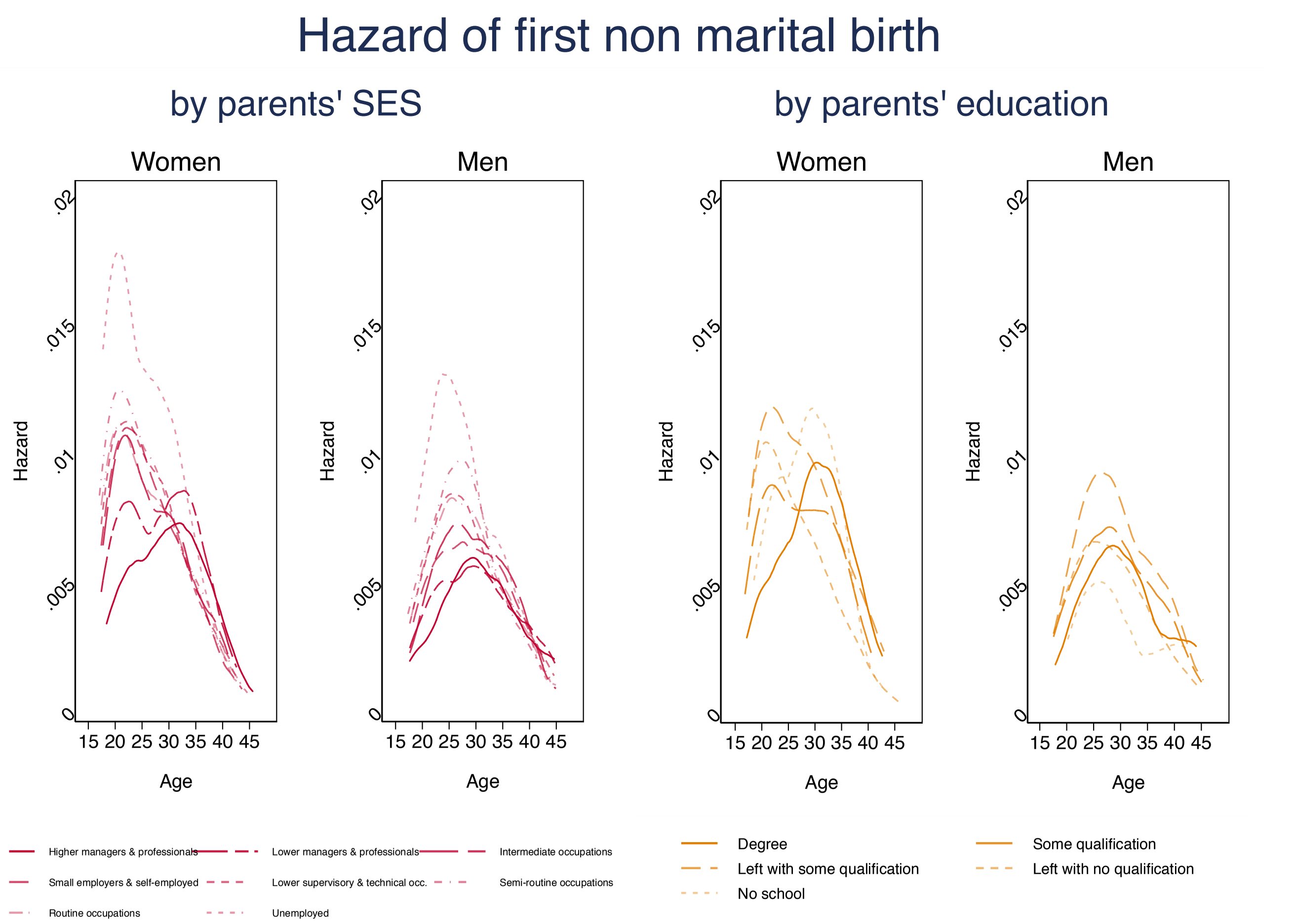


Figure 2.4. Entry into parenthood outside marriage, by parents' background. Women and men.



6. **Multivariate results**

6.1 Transition to union formation

Figures 2.5 and 2.6 display the results from two event-history multinomial logistic regression models of the likelihood of cohabiting or marrying relative to no union formation (the reference category) in a given monthly observation interval, for men (Fig. 2.5) and women (Fig. 2.6) separately. In the Appendix, Tables 2.A1-2.A4 show the full set of results as log-odds and odd ratios of all the covariates.

The multinomial logit coefficients are interpreted as the difference in odds of cohabiting or marrying in a given month for a single individual in one group compared with that of one in the reference group. I use stepwise models to highlights the correlations among the explanatory variables. Models 1 (red) and 2 (orange) display only one measure of social background at a time (parents' socio-economic status and parents' educational level, respectively), with controls for respondents' age and year of birth and gender, family stability, number of siblings, ethnicity, region of origin, metropolitan residence, and period of observation. Model 3 (blue with an empty circle) combines the two key independent variables, and Model 4 (blue with a full circle) includes two measures of respondent's educational level: a time-varying categorical variable capturing the highest educational attainment and a dichotomous variable reflecting on-ongoing full-time schooling.

Models 1 and 2, in which only one dimension of social background is included at a time, show that both dimensions of parents' background have a moderate negative association with the risk of direct marriage, while no clear pattern emerges in the risk of cohabitation. Parents' social class seems to have a quite steady gradient-like effect on marriage for men and women across the distribution (Dahlberg, 2015). However, class should not be strictly interpreted in a hierarchical order and interpretations of gradients should be made with some caution. For instance, the self-employed are difficult to place in a scale of occupational prestige. For women, the relative odds of entering a direct marriage is higher for the lower categories: lower supervisory, semi-routine occupations, and routine occupations. For men, this pattern holds with even wider difference between men whose parents were unskilled and skilled workers (except for the self-employed). When it comes to the competing risk of cohabitation, the pattern of parents' class is not clear and coefficients' magnitude is much smaller. For women and men, having parents from semi-routine and lower supervisory occupations is associated with higher risk of transition to cohabitation with respect to “higher managers” but this link does not hold for “routine occupations”, which do not differ from the reference group.

Parents’ education seems to have a steady gradient-like effect on the risk of marriage across the distribution for men and women. A different mechanism operates between parents’ highest educational attainment and children’s competing risk of a cohabitation. Females whose parents have a degree are significantly more likely to enter a cohabitation first (vs. remaining single) compared to their counterparts whose parents have (other) lower qualifications. However, males who reported parents having “left school” or reporting “no qualification” are as likely as the reference category to form a cohabitation in first place.

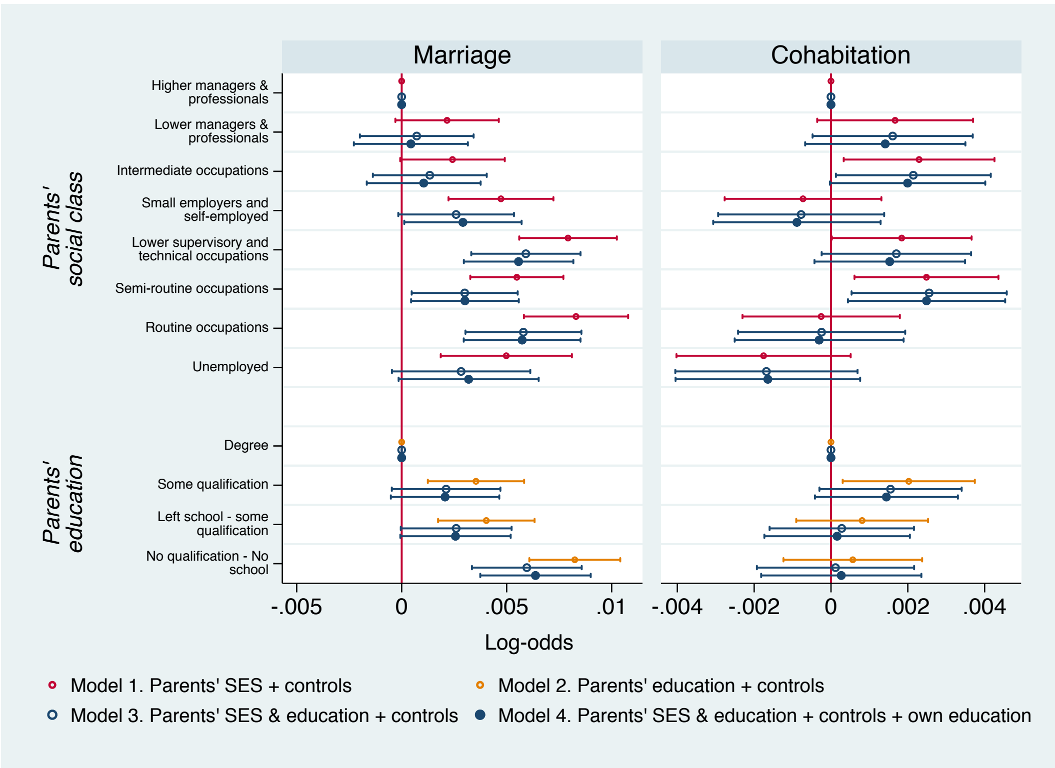
Model 3 shows that the two dimensions of social background decrease in magnitude when the measures are included simultaneously. This is not surprising given that the two measurements of social background are positively correlated with one another. Interestingly, the coefficients remain significant at 5% level with the exception most lower levels of education for both men and women with respect to the risk of cohabitation. A conclusion is that for men and, especially, for women, parental education is not a relevant dimension for the risk of cohabitation (vs. remaining single).

Model 4 adds two indicators of a person’s own educational history, which is found correlated with parents’ education in previous research (Blossfeld & Huinink, 1991). The inclusion of own education generally sterilizes the influence of parents’ attainment with respect to cohabitation, while the parents’ educational gradient on the risk of marriage remain apparent for men and, to a greater extent, for women. Previous empirical work has demonstrated that schooling postpones entry into marriage and cohabitation, whereas higher levels of education and earnings encourage union formation (Blom, 1994; Kravdal, 1999; Sassler and Goldscheider, 2004; Wiik, 2009). The analysis generally confirms these findings for men but not for women.

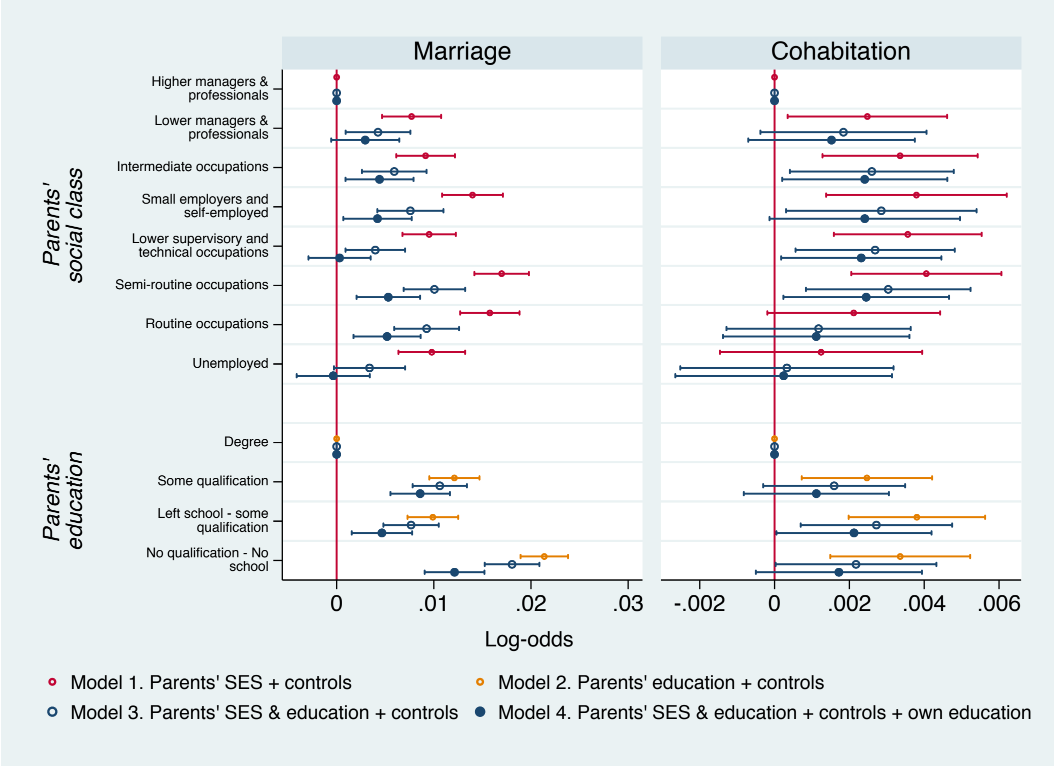
Table 2 illustrates the log-odds of the categories of children’s educational attainment in Model 4. The time-varying measures for educational level highlight that men with a degree are more likely to enter a marriage (cohabitation), vis-à-vis the A-level ($\beta = -0{,}099$; odds-ratio: $e^{\beta} = 0.906$) and GCSE holders ($\beta = -0{,}231$; e^{β} =0,794), and those with no qualification ($\beta = -0{,}329$; $e^{\beta} = 0{,}720$), respectively. A significant difference in the risk of cohabitation – but not in marriage – concerns men with “other qualification”. In contrast to the findings about men, the odds of entering first marriage versus staying single are lower for women with a degree compared to their less educated counterparts. Nevertheless, women's likelihood of entering the cohabitation follows a similar pattern as men's, with lower educated postponing this transition for longer with respect to degree holders. School enrollment decreases the likelihood of entering the first union, be it cohabitation or marriage, although the magnitude of the relationship is somewhat stronger for persons marrying directly. The reduction in the odds of entering a first marital (non-marital) union relative to no union formation in a given month by being a full-time student is 40,5% (28,6%) for men and 62,9% (40,6%) for women. However, these results should be interpreted with caution, as schooling and educational attainment are all potentially endogenous. It cannot be ruled out the individuals – and women in particular – make decisions about union formation, union type and education jointly. For instance, young adults who are 'family-oriented' may prefer to partner and leave education earlier than more career-oriented counterparts (Wiik, 2008).

Figure 2.5. Multinomial logit regression. Risk of first union: marriage and cohabitation vs. singlehood.

A. Men



B. Women



As mentioned in the theoretical section, social background can affect fertility indirectly, through the age at which the mother had her first child, economic resources, and educational careers and attainment. In contrast with previous evidence, I find that having experienced parental separation during childhood does not significantly reduce the likelihood of marrying directly and does not increase the odds of entering the first cohabitation for women. Conversely, being raised in numerous families is associated with higher chances of entering the first cohabitation for both men and women, while having a sibling reduces the risk of direct marriage for women. This result can be interpreted in the light of the theory holding that numerous families have less resources to devote to long term family arrangements – such as marriage – and privilege less financially demanding arrangements such as cohabitations. Ethnic groups display very diverse patterns to first union formation. Individuals who identify themselves as White are generally more likely to end up in a cohabitation (vs. remaining single) as opposed to all the other ethnic groups, while the Indian, Pakistani and Bangladeshi, and the "other" Asian are significantly more likely to form a direct marriage.

Table 2.2. Multinomial logit regression. Risk of first union: marriage and cohabitation vs. singlehood. Selected coefficients from Model 4. Men and women.

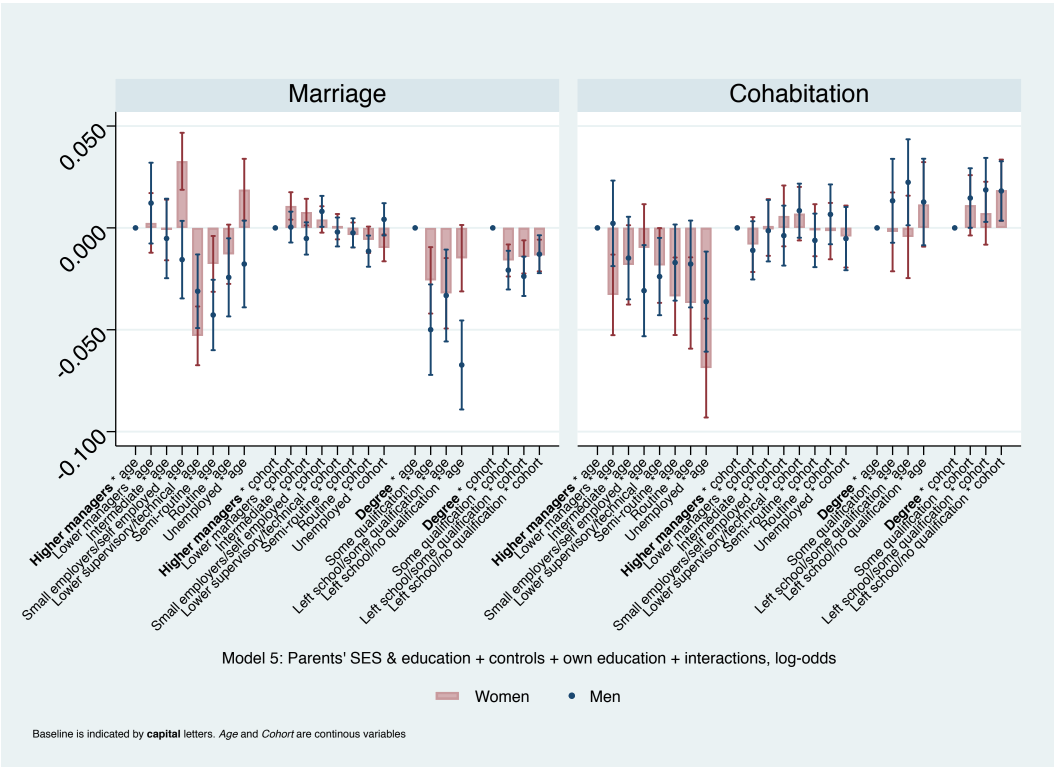
	Men		Women	
Variables	Marriage	Cohabitation	Marriage	Cohabitation
Highest education (*ref:* degree)				
Other higher	-0,001	0,025	0,089*	-0,027
A level etc	-0,099***	0,000	0,302***	-0,041
GCSE etc	-0,231***	-0,124***	0,081**	-0,174***
Other qual	-0,055	-0,155***	0,325***	-0,188***
No qualification	-0,329***	-0,268***	0,186***	-0,386***
Full-time education	-0,519***	-0,337***	-0,992***	-0,522***
N. siblings (*ref:* no siblings)				
1 siblings	0,098***	0,191***	-0,116***	0,219***
2+ siblings	0,153***	0,370***	-0,012	0,366***
Parents separated	-0,121	0,318***	-0,063	0,051
Ethnicity (*ref:* European White)				
African, Caribbean	-0,319***	-0,580***	-0,105**	-0,662***
Indian, Pakistani, Bangladeshi	0,862***	-2,291***	1,373***	-2,640***
Other Asian	0,274***	-1,593***	0,552***	-1,343***
Other	-0,067	-1,091***	0,029	-0,964***

6.2 First union. Interactions with age, cohort of birth and economic cycle

In the theoretical discussion, I hypothesized that the influence of socio-economic background of parents might change according to the age of children and the historical period. I interacted the two dimensions of family socio-economic background with these two variables expressed as continuous variables. Figure 2.6 includes the interaction terms with age and cohort of birth only in a model featuring all the covariates of Model 4 plus the interaction terms. All the interactions with the proxy of the economic cycle were not significant (see Appendix) and thus not included in Figure 2.6. The terms indicating the interaction with respondents' age and parents' socio-economic class are statistically significant for entry into cohabitation and marriage and reveal a quite clear pattern: the negative sign on the timing of first cohabitation and first marriage decreases with children's age and, in magnitude, the effect is more sizeable for the individuals from less disadvantaged background. This implies that the dependence of children from their parents' resources vanishes over time and it does more so for the less wealthy. Further, the coefficients for the interactions between cohort of birth and three parents' educational groups are negative for the entry into marriage, and positive for the entry into cohabitation (only marginally positive for women). This indicates that the effect of parents with education ranging from "no qualification" to "some qualification" on first marriage weakens and becomes less positive over historical time.

Conversely, the influence of the same educational groups on children's predisposition for cohabitation grows with respect to those from the highest educational layers.

Figure 2.6. Multinomial logit regression. Risk of first union: interaction of parent's background with children's age and year of birth. Separate models of women and men.



By the same token, the interplay between these three categorical groups and children's age is negative and statistically significant, for both men and women when it comes to the risk of first marriage (vs. remaining single): compared to their higher-educated counterparts, the influence on first marriage of mid-to-low educated parents decreases over children's life course. The other interaction factors displayed in Figure 2.6 do not provide clear evidence as displayed by the rather erratic pattern and don't lead to conclusive evidence.

6.3 Transition to parenthood

I present results of the transition to parenthood for men and women with nested models, in keeping with the analysis of union formation. Figures 2.7 A and B show the results from event-history multinomial logistic regression models of the likelihood of becoming parent in a marriage or in an alternative arrangement (be it in cohabitation or as a single) relative to not having a child in a given monthly observation interval.

Model 1 presents how the level of economic well-being in the child home correlates with the transition to parenthood. Among those who experience the first parenthood in marriage, a wealthier environment significantly decreases the odds of a first marital birth among men and, to a larger extent, women. Economic well-being in the upbringing is even more negatively related to a non-marital birth: the log-odds of experiencing a non-marital birth versus remaining childless for women and men whose parents' highest social class is "lower supervisory", "semi-routine" and "routine" occupations range from 0.443 to 0.651 for men and from 0.459 to 0.638 for women, compared to the reference group, "higher managers and professionals".

In model 2, I verify the role of socialization, parents' expectations, and other intangible aspects in childhood and teenage years by examining the role of parents' highest educational attainment. The findings highlight that increasing levels of parents' education are significantly associated with a postponement of parenthood in a marital and non-marital unions, net of other controls. In particular, having parents with 'no qualification or no schooling' is positively associated having a child in a marriage for women (men) relative to remaining childless in a specific month, given that no union was formed before than month. Similarly, having higher educated parents is significantly related to lower odds of experiencing a non-marital birth (vs. remaining childless).

The influence of each dimension of parents' background decreases when the proxies of social and educational background are jointly considered (Model 3) and further falls when indicators of individuals' educational history are accounted for (Model 4). The most striking finding is the loss of significance of parents' education on the odds of non-marital parenthood in the full model, both for women and men. Parents' education, however, continues to have an independent and generally negative influence on the prospects of a marital birth. For instance, women and men whose parents have no qualification have 20.8% ($e^{0.189} = 1.208$) and 32.5% ($e^{0.281} = 1.325$) higher odds of marital birth than those with university-educated parents, respectively. Parents' social class also reduces its influence in magnitude and significance although some findings confirm those shown in Model 1, especially when it comes to the transition to non-marital birth: economic well-being is substantially associated with lower chances of having a first non-marital birth.

Women with parents from semi-routine and routine occupation having a 41.9% ($e^{0.350} = 1.419$) and 45.5% ($e^{0.375} = 1.455$) higher odds of an out-of-wedlock birth (vs. childlessness) compared to the reference category. For men, in particular, there seems to be a cleavage among the top four categories of parents' social class (ranging from the "higher managers & professionals" to the "small employers & self-employed") and the bottom four (including the "unemployed"). When it comes to the risk of transition to marital parenthood, the findings are mixed for both genders. For men, the positive effect of parents' occupational class remains significant for "lower supervisory", "routine" occupations and for the unemployed (at 5% level), while for women the odds are significant also for those with parents from "intermediate" and "semi-routine" occupations (at 5% level).

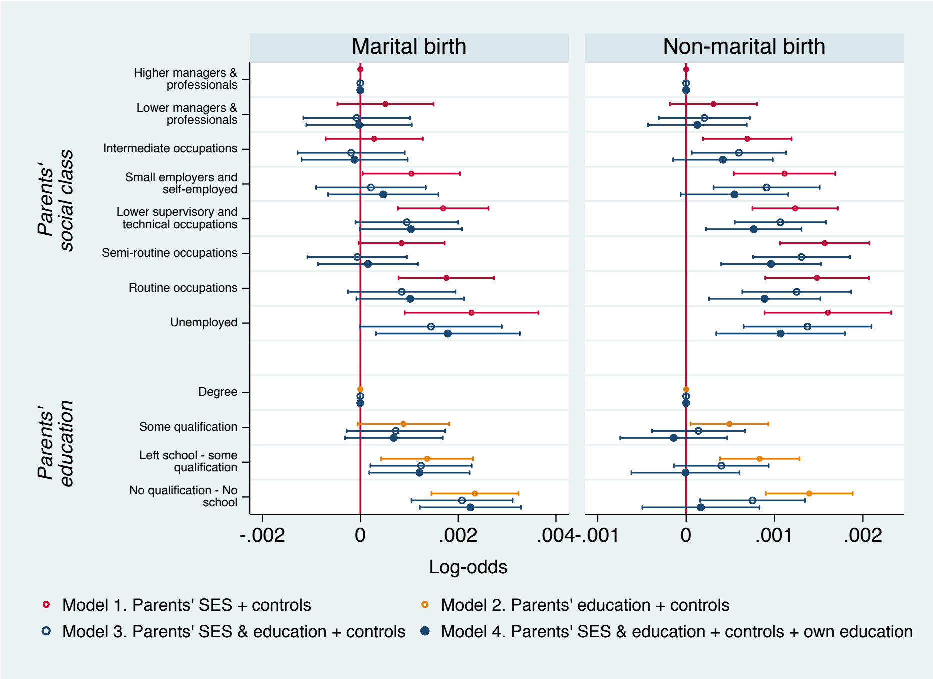
In line with the process of family formation, enrollment in education has a strong negative effect on the propensity to become a parent either in marriage or in another setting, regardless of gender: the risks of marital motherhood and fatherhood fall by 52.3% ($1 - e^{-0.740} = 1 - 0{,}477$) and 38,4%, respectively, when individuals are enrolled in education; and the odds of a non-marital motherhood and fatherhood decline by 70,2% and 63,3%, respectively, during full-time schooling. Interestingly, higher education has a stronger and negative effect on women's marital parenthood while it is generally non-significant for men. Only non-qualified men and those with an "A-level et al." have lower odds of entering a marital birth compared to their higher educated counterparts. On the contrary, every step up the educational attainment is associated with a lower relative risk of having a child in a non-marital union for women and men.

Family structure has a very marginal association with the transition to marital parenthood. Also, children from numerous families have higher odds of having a first birth out of marriage. The process of parenthood is also greatly diverse ethnicity-wise. All ethnic groups – with the exception of the African & Caribbean – are less prone to non-marital birth, and more likely to experience their first birth in a marriage vs. the White, regardless of gender.

In general, social class seems a robust predictor of entry into parenthood, particularly outside marriage, while the influence of parents' education is limited to the prediction of marital parenthood, for both men and women. I also tested the associations between father and mother's separate indicators of social background and the interaction between the latter and gender (instead of separate models). The results are shown in Appendix B1-B4.

Figure 2.7. Multinomial logit regression. Risk of first parenthood: marital and out-of-wedlock vs. no birth.

A. Men



B. Women

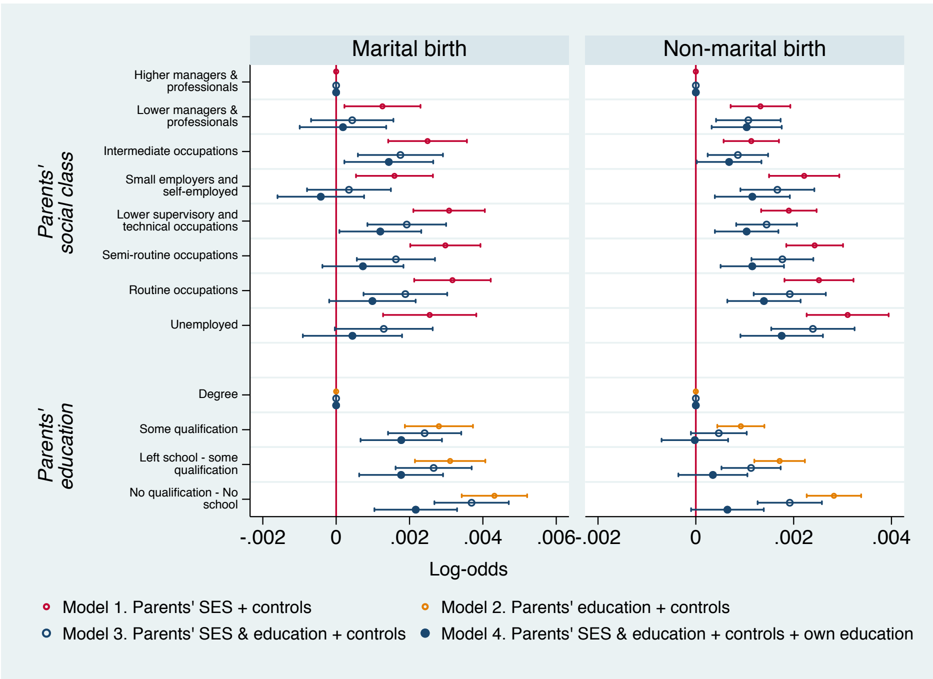


Table 2.3. Multinomial logit regression. Risk of parenthood: in a marital or non-marital setting. Selected coefficients from Model 4. Men and women.

	Men	Women

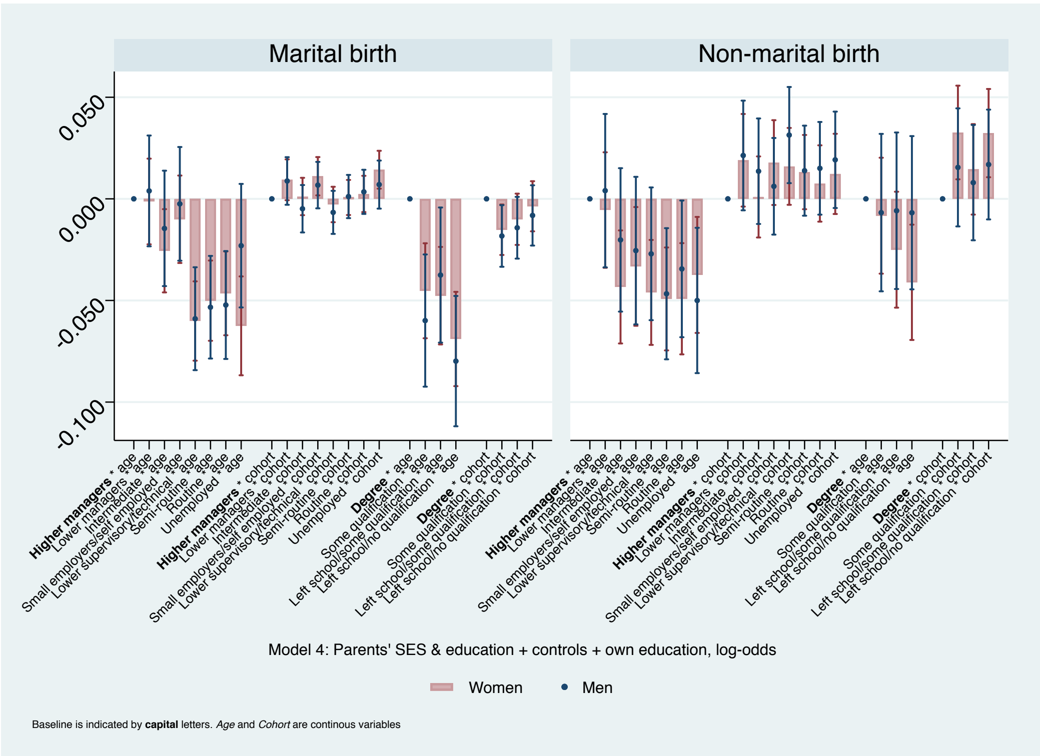
Variables	Marital birth	Non-marital birth	Marital birth	Non-marital birth
Highest education (*ref:* degree)				
Other higher	0,067	0,242**	0,247***	0,421***
A level etc	-0,100**	0,432***	0,117**	0,416***
GCSE etc	0,034	0,672***	0,362***	0,688***
Other qual	0,009	0,729***	0,373***	0,688***
No qualification	-0,205***	0,633***	0,467***	0,931***
Full-time education	-0,484***	-1,003***	-0,740***	-1,211***
N. siblings (*ref*: no siblings)				
1 siblings	-0,011	0,126	-0,030	-0,010
2+ siblings	0,077**	0,398***	0,061*	0,205***
Parents separated	-0,143	0,441***	-0,010	0,440***
Ethnicity (*ref*: European White)				
African, Caribbean	0,009	0,706***	-0,058	0,625***
Indian, Pakistani, Bangladeshi	1,263***	-1,616***	1,277***	-1,639***
Other Asian	0,628***	-1,236***	0,434***	-1,176***
Other	0,287**	-0,786***	0,505***	-0,411***

6.4 First birth. Interactions with age, cohort of birth and economic cycle

Figure 2.8 features the interaction of the dimensions of children's background with age and cohort of birth to highlight any differential influence over the life course and over time. The influence of parents' socio-economic position on the risk of marriage clearly vanishes as children age. The terms associated to the less advantaged strata of the society, both in terms of social class and education, are significantly negative. This implies that the dependence of children from their parents' material and immaterial resources tends to decrease over their life course, as far as the decision of marrying is concerned.

Also, the interaction coefficients of age and social class associated to the process of non-marital birth are generally significant for women. The effect of parents with an occupation equal to or lower than "intermediate" on first marriage weakens as daughters age. This finding can be interpreted as though women from more disadvantaged origin become increasingly less prone to an out-of-wedlock first birth, as opposed to women from the top social classes. Figure 2.8 suggests that women from lower social class are relatively more likely to enter non-marital motherhood in teenage years or as they come of age. On the contrary, women from wealthier classes tend to opt for first motherhood out of marriage after the age of 30. Interestingly, the non-significant influence of the variable "cohort" reveals that there hasn't been any significant diffusion of first out-of-wedlock birth in specific strata of the society as opposed to others, over time. This is not in contradiction with the findings from the models on union entry, which do not highlight any class-specific raise in the risk of first cohabitation risk over time either. In other words, these estimates debunk the hypothesis that non-marital birth has become increasingly selective of specific income groups in the UK in the period at issue.

Figure 2.8. Multinomial logit regression. Risk of first birth: interaction of parents' background with children's age and year of birth. Separate models of women and men



7. Discussion

This study improves the understanding on the relation between socio-economic background and transition to first union and to parenthood. Overall, the results lend support to the hypothesis of a negative relation between socio-economic family background and timing of first union – whether in a marriage or in a cohabitation, as found in many European and US studies – and a first non-marital birth.

In addition to the existing body of research, I explicitly account for the union context of first union formation and parenthood. In second place, I have introduced two different measures of social background – class and education – into model of union formation and parenthood transition and showed the influence of the explanatory variables on the two competing outcomes: marital and non-marital union, marital and non-marital birth. The study reveals that distinct measures of parents' socio-economic background have distinct effects on entry into first union and parenthood.

The first part of the analysis, which focuses on the transition to first union, shows diverging effects of parental social class and education on the timing of marriage and cohabitation. Introduced separately, both social class and education, have a clear negative association with the risk of entering union, either in a marital or in a non-marital arrangement. When included simultaneously, the evidence is less univocal. Parents' education has the most robust influence on the risk of first marriage, but has virtually none with respect to first cohabitation, when children's education is accounted for: children – particularly female – of highly educated parents had lower entry rates into marriage. The other dimension – social class – has a more attenuated influence on entry into union and, overall, not conducive of any social class gradient.

The divergent effects of parents' social background and the different influence on marriage and cohabitation confirm the hypothesis that the dimensions of background exert different effects and that the two partnership types are diverse. Interestingly, parental education clearly affects the timing of first marriage (independently of respondent's economic situation during childhood and other variables included) but not the timing of first cohabitation, in contrast to Wiik (2009) and Mooyaart & Liefbroer (2016). The effect of parental education on marriage is overall negative and generally non-significant on cohabitation for the observation period. The reasons for the "cohabitation exception" might be attributed to the self-selection of individuals into different family arrangements: people who marry (directly) and cohabit diverge on several domains ranging from the risk in dissolution (Liefbroer & Dourleijn, 2006) to leisure time engagement (Smock, 2000). This finding also confirms that parents *socialize* with children an opportune timing to enter a union during the upbringing and exert a *social control* more extensively during early during their children's adulthood. Higher-educated parents might be more aware of the cost of a long-term commitment, such as marriage, and might deter their offspring from haphazard early unions. Indeed, the negative influence of education on timing of marriage formation lowers as children age, which supports the hypothesis that parents' preferences might matter more for their children during teenage and early adulthood. Parents with a degree themselves did probably delay their entry into first union (Barber, 2001) passing onto their offspring their age preference for union formation. These pathways of parents should be transmitted to children's behavior, both in terms of educational choice and work, and, thus, their attitude towards family planning. However, the data do not provide information either on the age when parents started their own unions or on the partnership arrangement.

A wider glance at the analysis reveals a third explanation: cohabitation – originally pioneered by a restricted élite group – might have become a serious alternative to marriage over the last decades for larger social groups. Indeed, I found a negative interaction between children's year of birth and parents' education on the risk of direct marriage and a positive interaction with respect to the risk of cohabitation (especially for men). Further, the observed effect of parents' education on the timing of first marriage has decayed significantly over the cohorts under observations for the three lower educational groups. These attenuated effects of family education are consistent with the theoretical claims that individual demographic choices have become less responsive to norms and social pressure over time (Bumpass, 1990). These findings suggest that parental education – among families of non-university graduated – was a more important determinant of marriage timing when direct marriage was the standard route into a union (Wiik, 2009) and has gradually taken a toll on children's decision to cohabit. Nonetheless, this is not surprising since the proportion of young adults cohabiting prior to an eventual marriage has steadily increased (Beaujouan & Ní Bhrolchain, 2012).

First marriage and first cohabitation are generally postponed by individuals from more advantaged background, though the evidence is not compelling for all social classes. It is plausible to assume that parents from wealthier background are more able to induce their children to weigh up the partnership option. The economic theory would also claim that young adults stay home longer are less likely to start their own families if the opportunity cost of staying home is too high. Also, children from more advantaged background, used to high consumption habits, might have less incentive to leave home and possibly face some financial hardship early in their independent life. However, I cannot rule out that for people from more disadvantaged origins – and specifically women – the option of early cohabitation might be viewed as an opportunity to emancipate from family household and gain from pooling resources with their cohabiting partner.

Further, previous studies (e.g., Wiik, 2009) stressed that parents' socio-economic status might orientate children's own search behavior in the marriage market, as young adults seek a partner to maximize their long-term wellbeing (Oppenheimer, 2009). Under the assumption that children set the benefits and costs of a union according to the socio-economic standard of their family of origin, they intend to partner with someone from a similar socio-economic condition to minimize their risk of financial downfalls. Since dissolving a cohabitation is more easily remediable as opposed to marriage, individuals from advantaged background could be pursuing an optimal match even while cohabiting. Evidence highlights that this mechanism applies more to women from wealthier background, who defer marriage more than cohabitation, than men.

The second part of the analysis concentrates on the effects of social background on the type of first birth, whether in a marital or non-marital arrangement, using the same analytical strategy illustrated in the first part. The findings display quite strong evidence in favour of the negative association between the dimensions of social advantage and the timing of first birth either within or outside marriage, for men and women. In keeping with Aassve (2003), I find that women from better-off families have lower rates of pre-marital childbearing. When social class and education are jointly included in the final specification along with all the other control, the results prove rather robust, with few exceptions. First, children from wealthier background and raised by more educated parents have lower rates of entry into marital non-marital parenthood compared to their less advantaged counterparts. Second, offspring from the higher-educated delay their parenthood vis-à-vis those whose parents are less educated only within marriage while no significant difference emerges for out-wedlock parenthood. However, the direct influence of parents' education is absorbed by the two indicators of educational histories (a time-varying categorical variable accounting for the highest educational level and a dichotomous variable for ongoing schooling), which exert a largely negative influence on the risk of marital (for women only) and non-marital parenthood (also for men). The results suggest some mechanisms of transmission of parents' social background on the transition to first birth.

Parents' social class appears a robust predictor more for entry into non-marital than for marital birth for both women and men. Therefore, out-of-wedlock fertility is also function of economic resources as the classic theories of fertility would suggest. A plausible explanation could be that wealthier parents are aware that an early birth in an unstable (or no) union might compromise the economic wellbeing of their children. By the same token, children from higher social background have higher occupational ambitions (Harkonen & Bihagen, 2011; Manzoni et al., 2014) and tend to forgo early parenthood, especially whether in precarious conditions, not to compromise these goals. As far as parents' education is concerned, I hypothesized that it could either reflect preference for normative life-course sequencing and avoidance of unstable partnership for childbearing, intergenerational transmission of cultural habits, or weaker inclination for family formation. However, the fact that children raised in higher-educated families do not differ from their counterparts when it comes to the risk of a non-marital birth remains largely unexpected and contradicts previous evidence (Aassve, 2003).

A partial justification is represented by the strong role played by own education, which "drains" part of the influence of parents' education, as it is well established in the literature (e.g. Blossfeld & Huinink, 1991). Own (higher) education is a greater predictor to avoid nonmarital birth than parents' education either because higher-educated individuals have more to lose in terms of their socioeconomic attainment, or because they have greater knowledge how to avoid unintended fertility (Musick et al., 2010). As Carlson et al. (2013) point out, (unintended) non-marital fertility can be explained in part by lack of information and awareness that would otherwise prevent childbearing.

Turning to the interactions with cohort of birth and age, the results stress the role of family economic resources and education, to some extent, have a strong foothold on adolescent and young adults' plans for childbearing. Lower levels of parents' social class dissipate their effect on marital and non-marital (for women) birth at older ages.

This study supports that claim that intergenerational reproduction of inequality occurred in the UK for the cohorts born between 1930s and 1980s. Parents' social class is a robust predictor of entry into marital and non-marital union and has a sizeable effect on the risk of a non-marital birth. Parents' educational attainment has a significant effect on marital union and parenthood. Ultimately, the net effect of socioeconomic background may work against intergenerational social mobility if individuals from higher social backgrounds continue career progression longer by postponing childbearing and individuals from lower social class are more prone to engage in parenthood in less stable family arrangements.

This study has some limitations. A typical pitfall when dealing with male fertility, is that men tend to underreport births of children more than women do (Rendall et al., 1999).

In spite of this, I think that male fertility still deserves study for two reasons. First, the factors that affect male fertility might not be the same as those influencing female fertility, as the impact of parenthood on men's career might still be different from that on women', especially at young age (Michael & Tuma, 1985). Second, if underreporting of births by males is unrelated to the family background characteristics (which is not too implausible), then I still obtain a reasonable estimate of the effects of these characteristics on early entry into parenthood by males. A second limitation is the lack of variables that might be salient to union formation and childbearing. In particular, additional social factors—such as parents' attitudes, values, own family formation behavior, and parental involvement, age at union formation and childbirth (Barber, 2001)—as well as glimpses of respondent's psychological traits and time-varying information on income and social class are key factors. Future research with more fine-grained data in these domains could usefully help better weigh up the contribution of parental socio-economic background to the demographic dynamics at issue.

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Appendix

Table 2.A1. Entry into union. Log-odds. Men

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLE	Marr.	Cohab.	Marr.	Cohab.	Marr.	Cohab.	Marr.	Cohab.
Lower managers & professionals	0,189***	0,124**			0,136***	0,127**	0,130**	0,123**
	(0,048)	(0,052)			(0,050)	(0,054)	(0,050)	(0,054)
Intermediate occupations	0,089*	0,123**			0,048	0,127**	0,041	0,129**
	(0,049)	(0,050)			(0,051)	(0,051)	(0,051)	(0,051)
Small employers and self-employed	0,210***	0,016			0,136***	0,029	0,142***	0,036
	(0,047)	(0,056)			(0,050)	(0,059)	(0,051)	(0,059)
Lower supervisory and technical occupations	0,288***	0,135***			0,218***	0,145***	0,203***	0,144***
	(0,043)	(0,047)			(0,046)	(0,049)	(0,046)	(0,050)
Semi-routine occupations	0,205***	0,117**			0,120***	0,139***	0,114**	0,148***
	(0,042)	(0,047)			(0,046)	(0,050)	(0,047)	(0,051)
Routine occupations	0,342***	0,023			0,259***	0,038	0,257***	0,044
	(0,044)	(0,056)			(0,048)	(0,058)	(0,048)	(0,059)
Unemployed	0,144***	-0,106*			0,073	-0,079	0,085	-0,056
	(0,055)	(0,063)			(0,057)	(0,064)	(0,057)	(0,065)
Missing	-0,063	-0,248***			-0,125	-0,220**	-0,120	-0,196**
	(0,080)	(0,089)			(0,081)	(0,090)	(0,081)	(0,090)
Some qualification			0,161***	0,125**	0,087	0,084	0,083	0,079
			(0,053)	(0,050)	(0,056)	(0,053)	(0,056)	(0,053)

Left school - some qualification			0,188***	0,055	0,110*	0,010	0,107*	0,005
			(0,053)	(0,051)	(0,056)	(0,054)	(0,056)	(0,054)
Left school - no qualification			0,342***	0,041	0,239***	-0,007	0,244***	0,009
			(0,049)	(0,054)	(0,054)	(0,058)	(0,055)	(0,059)
No school			0,336***	-0,457**	0,236***	-0,502***	0,262***	-0,468**
			(0,083)	(0,183)	(0,087)	(0,184)	(0,087)	(0,185)
Missing			0,220***	-0,091*	0,141***	-0,126**	0,150***	-0,112**
			(0,050)	(0,051)	(0,054)	(0,053)	(0,055)	(0,054)
Age	0,839***	0,597***	0,839***	0,596***	0,841***	0,597***	0,722***	0,501***
	(0,020)	(0,018)	(0,020)	(0,018)	(0,020)	(0,018)	(0,022)	(0,020)
Age squared	-0,015***	-0,010***	-0,015***	-0,010***	-0,015***	-0,010***	-0,013***	-0,008***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Year of birth	0,058***	0,151***	0,056***	0,153***	0,057***	0,152***	0,053***	0,147***
	(0,008)	(0,014)	(0,008)	(0,015)	(0,008)	(0,015)	(0,008)	(0,015)
Year of birth squared	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (*ref.* no siblings)	0,102***	0,219***	0,110***	0,216***	0,107***	0,203***	0,098***	0,191***
	(0,033)	(0,047)	(0,033)	(0,047)	(0,033)	(0,047)	(0,033)	(0,047)
2 siblings (*ref.* no siblings)	0,157***	0,381***	0,174***	0,386***	0,157***	0,374***	0,153***	0,370***
	(0,030)	(0,044)	(0,029)	(0,044)	(0,030)	(0,044)	(0,030)	(0,044)
Parents separated	-0,115	0,307***	-0,122	0,313***	-0,116	0,317***	-0,121	0,318***
	(0,084)	(0,088)	(0,084)	(0,088)	(0,084)	(0,088)	(0,084)	(0,088)
Absent/dead mother	0,009	0,199**	-0,021	0,149*	0,018	0,212**	0,018	0,225**
	(0,077)	(0,091)	(0,077)	(0,089)	(0,077)	(0,090)	(0,077)	(0,090)
Absent/dead father	0,087	0,073	-0,028	-0,048	0,086	0,078	0,074	0,065
	(0,054)	(0,074)	(0,046)	(0,069)	(0,054)	(0,074)	(0,054)	(0,074)
African - Caribbean	-0,346***	-0,623***	-0,369***	-0,612***	-0,348***	-0,594***	-0,319***	-0,580***
	(0,061)	(0,062)	(0,060)	(0,062)	(0,061)	(0,062)	(0,061)	(0,063)

Indian, Pakistani, Bangladeshi	0,851***	-2,343***	0,840***	-2,319***	0,844***	-2,302***	0,862***	-2,291***
	(0,035)	(0,110)	(0,036)	(0,110)	(0,036)	(0,110)	(0,037)	(0,110)
Other Asian	0,270***	-1,618***	0,237***	-1,589***	0,263***	-1,592***	0,274***	-1,593***
	(0,078)	(0,141)	(0,078)	(0,141)	(0,078)	(0,141)	(0,079)	(0,142)
Other	-0,103	-1,166***	-0,154*	-1,197***	-0,101	-1,120***	-0,067	-1,091***
	(0,093)	(0,120)	(0,090)	(0,116)	(0,093)	(0,120)	(0,093)	(0,120)
Rural area	0,060**	-0,076**	0,053*	-0,081**	0,062**	-0,079**	0,062**	-0,083**
	(0,028)	(0,035)	(0,028)	(0,035)	(0,028)	(0,035)	(0,028)	(0,035)
yearly % GDP (t-12)	-0,113***	0,285***	-0,109***	0,282***	-0,111***	0,283***	-0,096***	0,290***
	(0,029)	(0,038)	(0,029)	(0,039)	(0,029)	(0,039)	(0,029)	(0,039)
Other higher							-0,001	0,025
							(0,051)	(0,056)
A level etc							-0,099***	0,000
							(0,037)	(0,042)
GCSE etc							-0,231***	-0,124***
							(0,042)	(0,046)
Other qual							-0,055	-0,155***
							(0,043)	(0,057)
No qualification							-0,329***	-0,268***
							(0,041)	(0,055)
Missing							-0,929	-0,899
							(0,588)	(0,715)
Currently in education							-0,519***	-0,337***
							(0,041)	(0,045)

Table 2.A2. Entry into union. Odds ratio. Men.

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Mar.	Cohab.	Mar.	Cohab.	Mar.	Cohab.	Mar.	Cohab.
Lower managers & professionals			1,208***	1,132**			1,145***	1,136**
			(0,058)	(0,059)			(0,057)	(0,061)
Intermediate occupations			1,093*	1,131**			1,049	1,136**
			(0,054)	(0,056)			(0,053)	(0,058)
Small employers and self-employed			1,234***	1,016			1,145***	1,029
			(0,058)	(0,057)			(0,057)	(0,060)
Lower supervisory and technical occupations			1,333***	1,145***			1,244***	1,156***
			(0,057)	(0,054)			(0,057)	(0,057)
Semi-routine occupations			1,228***	1,124**			1,128***	1,150***
			(0,052)	(0,053)			(0,052)	(0,058)
Routine occupations			1,407***	1,024			1,296***	1,039
			(0,062)	(0,057)			(0,062)	(0,061)
Unemployed			1,155***	0,899*			1,076	0,924
			(0,063)	(0,056)			(0,061)	(0,059)
Missing			0,939	0,780***			0,882	0,802**
			(0,075)	(0,069)			(0,072)	(0,072)
Some qualification					1,175***	1,133**	1,091	1,087
					(0,062)	(0,057)	(0,061)	(0,057)
Left school - some qualification					1,207***	1,056	1,117*	1,010
					(0,063)	(0,054)	(0,063)	(0,055)

Left school - no qualification					1,408***	1,042	1,270***	0,993
					(0,069)	(0,056)	(0,069)	(0,058)
No school					1,400***	0,633**	1,266***	0,605***
					(0,117)	(0,116)	(0,110)	(0,112)
Missing					1,246***	0,913*	1,151***	0,881**
					(0,063)	(0,046)	(0,062)	(0,047)
Age	0,839***	0,597***	0,839***	0,596***	0,841***	0,597***	0,722***	0,501***
	(0,020)	(0,018)	(0,020)	(0,018)	(0,020)	(0,018)	(0,022)	(0,020)
Age squared	-0,015***	-0,010***	-0,015***	-0,010***	-0,015***	-0,010***	-0,013***	-0,008***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Year of birth	0,058***	0,151***	0,056***	0,153***	0,057***	0,152***	0,053***	0,147***
	(0,008)	(0,014)	(0,008)	(0,015)	(0,008)	(0,015)	(0,008)	(0,015)
Year of birth squared	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (ref, no siblings)	2,305***	1,815***	2,315***	1,816***	2,315***	1,815***	2,320***	1,816***
	(0,047)	(0,033)	(0,047)	(0,033)	(0,047)	(0,033)	(0,047)	(0,033)
2 siblings (ref, no siblings)	0,985***	0,990***	0,985***	0,990***	0,985***	0,990***	0,985***	0,990***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Parents separated	1,059***	1,166***	1,059***	1,163***	1,058***	1,165***	1,059***	1,164***
	(0,009)	(0,017)	(0,009)	(0,017)	(0,009)	(0,017)	(0,009)	(0,017)
Absent/dead mother	0,999***	0,999***	0,999***	0,999***	0,999***	0,999***	0,999***	0,999***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Absent/dead father	1,104***	1,263***	1,107***	1,245***	1,116***	1,241***	1,113***	1,225***
	(0,037)	(0,059)	(0,037)	(0,058)	(0,037)	(0,058)	(0,037)	(0,057)
African - Caribbean	1,195***	1,483***	1,170***	1,464***	1,190***	1,471***	1,170***	1,453***
	(0,035)	(0,064)	(0,035)	(0,064)	(0,035)	(0,064)	(0,035)	(0,063)
Indian, Pakistani, Bangladeshi	0,888	1,358***	0,892	1,359***	0,885	1,368***	0,891	1,373***
	(0,075)	(0,120)	(0,075)	(0,119)	(0,075)	(0,121)	(0,075)	(0,121)

Other Asian	0,962	1,141	1,009	1,221**	0,980	1,161*	1,018	1,236**
	(0,074)	(0,102)	(0,078)	(0,111)	(0,075)	(0,103)	(0,079)	(0,112)
Other	0,985	0,937	1,091	1,076	0,972	0,954	1,089	1,081
	(0,046)	(0,064)	(0,059)	(0,079)	(0,045)	(0,066)	(0,059)	(0,080)
Rural area	0,693***	0,525***	0,707***	0,536***	0,691***	0,542***	0,706***	0,552***
	(0,042)	(0,033)	(0,043)	(0,033)	(0,042)	(0,034)	(0,043)	(0,035)
yearly % GDP (t-12)	2,326***	0,094***	2,342***	0,096***	2,316***	0,098***	2,326***	0,100***
	(0,080)	(0,010)	(0,083)	(0,011)	(0,083)	(0,011)	(0,085)	(0,011)
Missing							0,395	0,407
							(0,232)	(0,291)
No qualification							0,720***	0,765***
							(0,030)	(0,042)
Other qual							0,946	0,856***
							(0,041)	(0,049)
GCSE etc							0,794***	0,884***
							(0,033)	(0,041)
A level etc							0,906***	1,000
							(0,033)	(0,042)
Other higher							0,999	1,025
							(0,051)	(0,057)
Currently in education							0,595***	0,714***
							(0,024)	(0,032)

Table 2,A3. Entry into union. Log-odds. Women

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Mar.	Cohab.	Mar.	Cohab.	Mar.	Cohab.	Mar.	Cohab.
Lower managers & professionals	0,199***	0,099**			0,109**	0,079	0,054	0,065
	(0,043)	(0,049)			(0,044)	(0,050)	(0,044)	(0,050)
Intermediate occupations	0,306***	0,121***			0,217***	0,100**	0,153***	0,095*
	(0,042)	(0,046)			(0,042)	(0,048)	(0,043)	(0,048)
Small employers and self-employed	0,434***	0,115**			0,281***	0,100*	0,168***	0,083
	(0,040)	(0,052)			(0,042)	(0,054)	(0,042)	(0,054)
Lower supervisory and technical occupations	0,287***	0,152***			0,152***	0,130***	0,035	0,113**
	(0,038)	(0,044)			(0,040)	(0,046)	(0,040)	(0,047)
Semi-routine occupations	0,423***	0,162***			0,266***	0,141***	0,120***	0,118**
	(0,037)	(0,044)			(0,039)	(0,047)	(0,040)	(0,047)
Routine occupations	0,420***	0,062			0,269***	0,039	0,137***	0,036
	(0,039)	(0,052)			(0,041)	(0,055)	(0,042)	(0,055)
Unemployed	0,392***	0,016			0,232***	0,008	0,105**	0,012
	(0,044)	(0,059)			(0,046)	(0,062)	(0,046)	(0,062)
Missing	-0,076	0,093			-0,185***	0,092	-0,299***	0,081
	(0,065)	(0,078)			(0,066)	(0,079)	(0,066)	(0,080)
Some qualification			0,388***	0,111**	0,310***	0,074	0,242***	0,054
			(0,045)	(0,044)	(0,047)	(0,047)	(0,047)	(0,047)
Left school - some qualification			0,348***	0,167***	0,244***	0,122**	0,150***	0,097**
			(0,046)	(0,045)	(0,048)	(0,049)	(0,048)	(0,049)
Left school - no qualification			0,618***	0,141***	0,489***	0,092*	0,325***	0,072

			(0,043)	(0,047)	(0,046)	(0,051)	(0,046)	(0,052)
No school			0,847***	-0,564***	0,701***	-0,604***	0,500***	-0,571***
			(0,064)	(0,187)	(0,066)	(0,188)	(0,068)	(0,189)
Missing			0,387***	-0,020	0,285***	-0,061	0,129***	-0,077
			(0,044)	(0,045)	(0,046)	(0,049)	(0,047)	(0,050)
Age	0,461***	0,519***	0,465***	0,518***	0,467***	0,519***	0,323***	0,340***
	(0,013)	(0,020)	(0,013)	(0,020)	(0,013)	(0,020)	(0,013)	(0,022)
Age squared	-0,009***	-0,009***	-0,009***	-0,009***	-0,009***	-0,009***	-0,006***	-0,006***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Year of birth	0,022***	0,216***	0,016**	0,217***	0,019***	0,216***	0,014**	0,210***
	(0,007)	(0,017)	(0,007)	(0,017)	(0,007)	(0,017)	(0,007)	(0,017)
Year of birth squared	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,001***	-0,000***	-0,001***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (*ref*, no siblings)	-0,132***	0,236***	-0,125***	0,224***	-0,127***	0,225***	-0,116***	0,219***
	(0,028)	(0,046)	(0,028)	(0,045)	(0,028)	(0,046)	(0,028)	(0,046)
2 siblings (*ref*, no siblings)	-0,001	0,371***	0,020	0,369***	0,003	0,371***	-0,012	0,366***
	(0,024)	(0,043)	(0,024)	(0,042)	(0,024)	(0,043)	(0,024)	(0,043)
Parents separated	-0,018	0,052	-0,065	0,041	-0,038	0,045	-0,063	0,051
	(0,060)	(0,085)	(0,060)	(0,085)	(0,060)	(0,085)	(0,060)	(0,085)
Absent/dead mother	0,131**	0,129	0,044	0,143*	0,147**	0,145*	0,126**	0,139
	(0,058)	(0,086)	(0,058)	(0,083)	(0,059)	(0,085)	(0,059)	(0,085)
Absent/dead father	0,188***	0,007	0,125***	0,011	0,199***	0,015	0,205***	0,023
	(0,042)	(0,075)	(0,036)	(0,066)	(0,042)	(0,075)	(0,042)	(0,075)
African - Caribbean	-0,103**	-0,701***	-0,117***	-0,680***	-0,120***	-0,680***	-0,105**	-0,662***
	(0,043)	(0,054)	(0,044)	(0,055)	(0,044)	(0,055)	(0,044)	(0,055)
Indian, Pakistani, Bangladeshi	1,369***	-2,672***	1,356***	-2,658***	1,349***	-2,645***	1,373***	-2,640***
	(0,030)	(0,146)	(0,030)	(0,146)	(0,031)	(0,147)	(0,031)	(0,147)
Other Asian	0,517***	-1,392***	0,490***	-1,378***	0,508***	-1,374***	0,552***	-1,343***
	(0,062)	(0,122)	(0,062)	(0,122)	(0,063)	(0,122)	(0,063)	(0,122)
Other	-0,029	-1,049***	-0,101	-1,005***	-0,015	-1,004***	0,029	-0,964***
	(0,085)	(0,118)	(0,083)	(0,114)	(0,085)	(0,118)	(0,085)	(0,118)

Rural area	0,133***	0,079**	0,143***	0,076**	0,141***	0,076**	0,176***	0,080**
	(0,023)	(0,032)	(0,023)	(0,032)	(0,023)	(0,032)	(0,023)	(0,032)
yearly % GDP (t-12)	0,071***	0,163***	0,068***	0,161***	0,070***	0,163***	0,118***	0,175***
	(0,023)	(0,038)	(0,023)	(0,039)	(0,023)	(0,039)	(0,023)	(0,039)
Other higher							0,089*	-0,027
							(0,047)	(0,052)
A level etc							0,302***	-0,041
							(0,040)	(0,042)
GCSE etc							0,081**	-0,174***
							(0,040)	(0,044)
Other qual							0,325***	-0,188***
							(0,044)	(0,059)
No qualification							0,186***	-0,386***
							(0,041)	(0,054)
Missing							0,454	-0,440
							(0,573)	(0,713)
Currently in education							-0,992***	-0,522***
							(0,032)	(0,039)

Table 2.A4. Entry into union. Odds ratio. Women

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Mar.	Cohab.	Mar.	Cohab.	Mar.	Cohab.	Mar.	Cohab.
Lower managers & professionals	1,220***	1,104**			1,115**	1,082	1,056	1,067
	(0,052)	(0,054)			(0,048)	(0,054)	(0,046)	(0,053)
Intermediate occupations	1,358***	1,129***			1,243***	1,105**	1,165***	1,099*
	(0,057)	(0,052)			(0,053)	(0,053)	(0,050)	(0,053)
Small employers and self-employed	1,543***	1,121**			1,324***	1,106*	1,183***	1,087
	(0,062)	(0,058)			(0,055)	(0,060)	(0,050)	(0,059)
Lower supervisory and technical occupations	1,332***	1,164***			1,164***	1,138***	1,035	1,119**
	(0,051)	(0,051)			(0,046)	(0,053)	(0,042)	(0,052)
Semi-routine occupations	1,527***	1,176***			1,305***	1,151***	1,128***	1,125**
	(0,056)	(0,051)			(0,051)	(0,054)	(0,045)	(0,053)
Routine occupations	1,522***	1,064			1,309***	1,040	1,146***	1,037
	(0,060)	(0,056)			(0,053)	(0,057)	(0,048)	(0,057)
Unemployed	1,479***	1,016			1,261***	1,008	1,111**	1,012
	(0,065)	(0,060)			(0,058)	(0,062)	(0,052)	(0,063)
Missing	0,927	1,098			0,831***	1,096	0,741***	1,085
	(0,060)	(0,086)			(0,055)	(0,087)	(0,049)	(0,087)
Some qualification			1,475***	1,117**	1,364***	1,077	1,274***	1,055
			(0,067)	(0,050)	(0,063)	(0,050)	(0,060)	(0,050)
Left school - some qualification			1,416***	1,182***	1,276***	1,130**	1,161***	1,101**
			(0,066)	(0,054)	(0,061)	(0,055)	(0,056)	(0,054)

Left school - no qualification			1,855***	1,152***	1,631***	1,096*	1,384***	1,074
			(0,080)	(0,054)	(0,074)	(0,056)	(0,064)	(0,056)
No school			2,334***	0,569***	2,016***	0,547***	1,649***	0,565***
			(0,149)	(0,106)	(0,134)	(0,103)	(0,111)	(0,107)
Missing			1,472***	0,980	1,330***	0,941	1,138***	0,925
			(0,065)	(0,044)	(0,061)	(0,046)	(0,053)	(0,046)
Age	1,586***	1,680***	1,592***	1,678***	1,595***	1,681***	1,381***	1,405***
	(0,020)	(0,033)	(0,020)	(0,033)	(0,021)	(0,033)	(0,018)	(0,031)
Age squared	0,991***	0,991***	0,991***	0,991***	0,991***	0,991***	0,994***	0,994***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Year of birth	1,022***	1,241***	1,016**	1,243***	1,019***	1,241***	1,014**	1,233***
	(0,007)	(0,021)	(0,007)	(0,021)	(0,007)	(0,021)	(0,007)	(0,021)
Year of birth squared	0,999***	0,999***	0,999***	0,999***	0,999***	0,999***	1,000***	0,999***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (*ref*, no siblings)	0,876***	1,266***	0,882***	1,252***	0,880***	1,253***	0,890***	1,245***
	(0,024)	(0,058)	(0,024)	(0,057)	(0,024)	(0,057)	(0,025)	(0,057)
2 siblings (*ref*, no siblings)	0,999	1,450***	1,020	1,446***	1,003	1,450***	0,988	1,442***
	(0,024)	(0,062)	(0,025)	(0,061)	(0,024)	(0,062)	(0,024)	(0,062)
Parents separated	0,982	1,053	0,937	1,042	0,963	1,046	0,939	1,052
	(0,059)	(0,090)	(0,056)	(0,088)	(0,057)	(0,089)	(0,056)	(0,090)
Absent/dead mother	1,140**	1,138	1,045	1,154*	1,159**	1,156*	1,135**	1,149
	(0,067)	(0,098)	(0,061)	(0,096)	(0,068)	(0,099)	(0,067)	(0,098)
Absent/dead father	1,207***	1,007	1,134***	1,011	1,221***	1,015	1,228***	1,023
	(0,051)	(0,075)	(0,041)	(0,067)	(0,051)	(0,076)	(0,051)	(0,077)
African - Caribbean	0,902**	0,496***	0,890***	0,506***	0,887***	0,507***	0,901**	0,516***
	(0,039)	(0,027)	(0,039)	(0,028)	(0,039)	(0,028)	(0,040)	(0,028)
Indian, Pakistani, Bangladeshi	3,932***	0,069***	3,880***	0,070***	3,852***	0,071***	3,948***	0,071***
	(0,118)	(0,010)	(0,118)	(0,010)	(0,118)	(0,010)	(0,122)	(0,010)

Other Asian	1,677***	0,249***	1,632***	0,252***	1,662***	0,253***	1,737***	0,261***
	(0,104)	(0,030)	(0,102)	(0,031)	(0,104)	(0,031)	(0,109)	(0,032)
Other	0,972	0,350***	0,904	0,366***	0,985	0,367***	1,029	0,381***
	(0,082)	(0,041)	(0,075)	(0,042)	(0,084)	(0,043)	(0,088)	(0,045)
Rural area	1,142***	1,082**	1,153***	1,079**	1,151***	1,079**	1,192***	1,083**
	(0,026)	(0,034)	(0,026)	(0,034)	(0,027)	(0,034)	(0,028)	(0,034)
yearly % GDP (t-12)	1,073***	1,177***	1,071***	1,175***	1,072***	1,177***	1,125***	1,192***
	(0,025)	(0,045)	(0,025)	(0,045)	(0,025)	(0,045)	(0,026)	(0,046)
A level etc							1,352***	0,960
							(0,054)	(0,040)
GCSE etc							1,084**	0,840***
							(0,044)	(0,037)
Other qual							1,384***	0,828***
							(0,060)	(0,049)
No qualification							1,204***	0,680***
							(0,049)	(0,037)
Missing							1,574	0,644
							(0,901)	(0,459)
Currently in education							0,371***	0,594***
							(0,012)	(0,023)

Table 2.B1. Entry into parenthood. Log-odds. Men.

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock
Lower managers & professionals	0,074	0,113			0,011	0,058	0,016	0,011
	(0,063)	(0,109)			(0,065)	(0,110)	(0,066)	(0,110)
Intermediate occupations	0,029	0,306***			-0,020	0,250**	-0,013	0,144
	(0,065)	(0,101)			(0,067)	(0,103)	(0,067)	(0,104)
Small employers and self-employed	0,142**	0,400***			0,053	0,319***	0,076	0,148
	(0,061)	(0,103)			(0,066)	(0,107)	(0,067)	(0,108)
Lower supervisory and technical occupations	0,221***	0,443***			0,143**	0,368***	0,152**	0,216**
	(0,056)	(0,091)			(0,060)	(0,095)	(0,061)	(0,096)
Semi-routine occupations	0,103*	0,546***			0,007	0,450***	0,029	0,275***
	(0,056)	(0,090)			(0,061)	(0,095)	(0,062)	(0,097)
Routine occupations	0,230***	0,535***			0,134**	0,450***	0,152**	0,281***
	(0,059)	(0,099)			(0,063)	(0,104)	(0,064)	(0,105)
Unemployed	0,222***	0,651***			0,143*	0,557***	0,179**	0,373***
	(0,071)	(0,105)			(0,075)	(0,108)	(0,075)	(0,110)
Missing	-0,153	0,279*			-0,221*	0,193	-0,197*	0,044
	(0,112)	(0,158)			(0,115)	(0,160)	(0,115)	(0,160)
Some qualification			0,135*	0,223**	0,101	0,088	0,096	-0,022
			(0,072)	(0,106)	(0,075)	(0,109)	(0,075)	(0,110)
Left school - some qualification			0,191***	0,352***	0,159**	0,185*	0,156**	0,024
			(0,071)	(0,104)	(0,075)	(0,108)	(0,076)	(0,110)

Left school - no qualification			0,319***	0,506***	0,269***	0,286**	0,281***	0,068
			(0,067)	(0,105)	(0,074)	(0,113)	(0,074)	(0,115)
No school			0,388***	0,162	0,342***	-0,053	0,385***	-0,286
			(0,104)	(0,255)	(0,109)	(0,259)	(0,110)	(0,261)
Missing			0,164**	0,490***	0,125*	0,315***	0,141*	0,101
			(0,069)	(0,101)	(0,074)	(0,106)	(0,074)	(0,109)
Age	1,348***	0,341***	1,348***	0,340***	1,349***	0,341***	1,306***	0,257***
	(0,034)	(0,025)	(0,034)	(0,025)	(0,034)	(0,025)	(0,035)	(0,028)
Age squared	-0,022***	-0,006***	-0,022***	-0,006***	-0,022***	-0,006***	-0,021***	-0,005***
	(0,001)	(0,000)	(0,001)	(0,000)	(0,001)	(0,000)	(0,001)	(0,000)
Year of birth	0,034***	0,063***	0,033***	0,063***	0,034***	0,064***	0,029***	0,061***
	(0,010)	(0,021)	(0,010)	(0,021)	(0,010)	(0,021)	(0,010)	(0,021)
Year of birth squared	-0,001***	-0,000	-0,001***	-0,000	-0,001***	-0,000	-0,001***	-0,000
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (*ref*, no siblings)	-0,014	0,084	-0,010	0,073	-0,011	0,094	-0,011	0,123
	(0,043)	(0,085)	(0,043)	(0,084)	(0,043)	(0,085)	(0,043)	(0,085)
2 siblings (*ref*, no siblings)	0,070*	0,408***	0,083**	0,419***	0,071*	0,406***	0,077**	0,395***
	(0,038)	(0,075)	(0,038)	(0,075)	(0,038)	(0,075)	(0,038)	(0,075)
Parents separated	-0,160	0,528***	-0,170	0,517***	-0,151	0,515***	-0,141	0,435***
	(0,114)	(0,128)	(0,115)	(0,130)	(0,114)	(0,128)	(0,114)	(0,129)
Absent/dead mother	-0,060	0,175	-0,095	0,122	-0,047	0,169	-0,043	0,119
	(0,105)	(0,141)	(0,104)	(0,138)	(0,105)	(0,142)	(0,105)	(0,142)
Absent/dead father	-0,029	-0,087	-0,092	-0,035	-0,032	-0,091	-0,045	-0,068
	(0,072)	(0,122)	(0,061)	(0,103)	(0,073)	(0,122)	(0,073)	(0,121)
African - Caribbean	-0,024	0,642***	-0,040	0,626***	-0,028	0,642***	0,000	0,720***
	(0,089)	(0,084)	(0,088)	(0,084)	(0,089)	(0,084)	(0,089)	(0,086)
Indian, Pakistani, Bangladeshi	1,249***	-1,663***	1,242***	-1,634***	1,235***	-1,653***	1,249***	-1,587***
	(0,045)	(0,157)	(0,046)	(0,157)	(0,047)	(0,158)	(0,047)	(0,158)

Other Asian	0,605***	-1,374***	0,578***	-1,415***	0,601***	-1,380***	0,622***	-1,214***
	(0,108)	(0,292)	(0,108)	(0,292)	(0,108)	(0,293)	(0,109)	(0,293)
Other	0,263**	-0,814***	0,233*	-0,863***	0,267**	-0,842***	0,277**	-0,772***
	(0,124)	(0,215)	(0,122)	(0,214)	(0,124)	(0,216)	(0,125)	(0,218)
Rural area	0,032	-0,283***	0,027	-0,284***	0,035	-0,276***	0,029	-0,265***
	(0,036)	(0,064)	(0,036)	(0,064)	(0,036)	(0,064)	(0,036)	(0,064)
yearly % GDP (t-12)	0,038	-0,058	0,037	-0,058	0,036	-0,058	0,039	-0,034
	(0,035)	(0,071)	(0,035)	(0,071)	(0,035)	(0,071)	(0,035)	(0,071)
Other higher							-0,993	-13,390***
							(1,003)	(0,109)
A level etc							-0,208***	0,635***
							(0,051)	(0,097)
GCSE etc							0,007	0,729***
							(0,052)	(0,096)
Other qual							0,033	0,672***
							(0,050)	(0,085)
No qualification							-0,101**	0,432***
							(0,047)	(0,085)
Missing							0,066	0,242**
							(0,056)	(0,109)
Currently in education							-0,614***	-1,051***
							(0,122)	(0,106)

Table 2.B2. Entry into parenthood. Odds ratios. Men

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock
Lower managers & professionals	1,077	1,119			1,011	1,060	1,016	1,011
	(0,067)	(0,122)			(0,066)	(0,117)	(0,067)	(0,112)
Intermediate occupations	1,030	1,358***			0,980	1,283**	0,988	1,155
	(0,067)	(0,137)			(0,066)	(0,133)	(0,066)	(0,120)
Small employers and self-employed	1,153**	1,492***			1,054	1,376***	1,079	1,159
	(0,071)	(0,154)			(0,069)	(0,147)	(0,072)	(0,125)
Lower supervisory and technical occupations	1,248***	1,558***			1,154**	1,445***	1,164**	1,241**
	(0,070)	(0,142)			(0,069)	(0,137)	(0,071)	(0,119)
Semi-routine occupations	1,108*	1,726***			1,007	1,568***	1,029	1,316***
	(0,062)	(0,155)			(0,061)	(0,150)	(0,064)	(0,127)
Routine occupations	1,259***	1,707***			1,144**	1,569***	1,165**	1,325***
	(0,074)	(0,169)			(0,072)	(0,163)	(0,074)	(0,140)
Unemployed	1,249***	1,918***			1,153*	1,745***	1,196**	1,452***
	(0,089)	(0,201)			(0,086)	(0,188)	(0,090)	(0,159)
Missing	0,859	1,322*			0,802*	1,213	0,821*	1,045
	(0,096)	(0,209)			(0,092)	(0,194)	(0,094)	(0,167)
Some qualification			1,145*	1,250**	1,106	1,092	1,100	0,978
			(0,082)	(0,132)	(0,083)	(0,119)	(0,083)	(0,108)
Left school - some qualification			1,210***	1,422***	1,172**	1,204*	1,169**	1,025
			(0,086)	(0,148)	(0,089)	(0,130)	(0,088)	(0,113)

Left school - no qualification			1,376***	1,659***	1,309***	1,330**	1,325***	1,070
			(0,092)	(0,175)	(0,097)	(0,150)	(0,098)	(0,123)
No school			1,474***	1,176	1,408***	0,949	1,470***	0,752
			(0,153)	(0,300)	(0,154)	(0,245)	(0,161)	(0,196)
Missing			1,178**	1,633***	1,134*	1,370***	1,151*	1,106
			(0,081)	(0,164)	(0,084)	(0,145)	(0,085)	(0,120)
Age	3,851***	1,406***	3,851***	1,405***	3,853***	1,407***	3,692***	1,293***
	(0,130)	(0,035)	(0,130)	(0,035)	(0,130)	(0,035)	(0,128)	(0,037)
Age squared	0,979***	0,994***	0,979***	0,994***	0,979***	0,994***	0,979***	0,995***
	(0,001)	(0,000)	(0,001)	(0,000)	(0,001)	(0,000)	(0,001)	(0,000)
Year of birth	1,035***	1,066***	1,033***	1,065***	1,034***	1,067***	1,029***	1,063***
	(0,011)	(0,022)	(0,011)	(0,022)	(0,011)	(0,022)	(0,011)	(0,022)
Year of birth squared	0,999***	1,000	0,999***	1,000	0,999***	1,000	0,999***	1,000
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (*ref*, no siblings)	0,986	1,088	0,990	1,075	0,990	1,099	0,989	1,130
	(0,043)	(0,092)	(0,043)	(0,091)	(0,043)	(0,093)	(0,043)	(0,096)
2 siblings (*ref*, no siblings)	1,073*	1,504***	1,086**	1,521***	1,074*	1,501***	1,080**	1,484***
	(0,041)	(0,113)	(0,041)	(0,114)	(0,041)	(0,113)	(0,041)	(0,112)
Parents separated	0,852	1,696***	0,844	1,676***	0,860	1,674***	0,868	1,544***
	(0,097)	(0,217)	(0,097)	(0,218)	(0,098)	(0,215)	(0,099)	(0,199)
Absent/dead mother	0,942	1,191	0,909	1,130	0,954	1,184	0,958	1,126
	(0,099)	(0,168)	(0,095)	(0,156)	(0,100)	(0,168)	(0,100)	(0,160)
Absent/dead father	0,971	0,917	0,912	0,965	0,968	0,913	0,956	0,934
	(0,070)	(0,112)	(0,055)	(0,100)	(0,070)	(0,111)	(0,070)	(0,113)
African - Caribbean	0,976	1,900***	0,961	1,870***	0,972	1,900***	1,000	2,054***
	(0,087)	(0,160)	(0,085)	(0,157)	(0,086)	(0,160)	(0,089)	(0,176)
Indian, Pakistani, Bangladeshi	3,488***	0,190***	3,462***	0,195***	3,439***	0,191***	3,488***	0,205***
	(0,156)	(0,030)	(0,158)	(0,031)	(0,161)	(0,030)	(0,164)	(0,032)

Other Asian	1,832***	0,253***	1,782***	0,243***	1,824***	0,252***	1,862***	0,297***
	(0,198)	(0,074)	(0,192)	(0,071)	(0,198)	(0,074)	(0,203)	(0,087)
Other	1,301**	0,443***	1,263*	0,422***	1,306**	0,431***	1,319**	0,462***
	(0,161)	(0,095)	(0,154)	(0,090)	(0,162)	(0,093)	(0,164)	(0,101)
Rural area	1,033	0,753***	1,027	0,752***	1,035	0,759***	1,029	0,767***
	(0,037)	(0,048)	(0,037)	(0,048)	(0,037)	(0,049)	(0,037)	(0,049)
yearly % GDP (t-12)	1,038	0,944	1,038	0,944	1,037	0,944	1,040	0,967
	(0,036)	(0,067)	(0,036)	(0,067)	(0,036)	(0,067)	(0,037)	(0,069)
Other higher							1,068	1,274**
							(0,060)	(0,139)
A level etc							0,904**	1,540***
							(0,043)	(0,130)
GCSE etc							1,034	1,959***
							(0,052)	(0,166)
Other qual							1,007	2,073***
							(0,052)	(0,199)
No qualification							0,812***	1,887***
							(0,041)	(0,184)
Missing							0,370	0,000***
							(0,372)	(0,000)
Currently in education							0,541***	0,350***
							(0,066)	(0,037)

Table 2.B3. Entry into parenthood. Log-odds. Women

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock
Lower managers & professionals	0,135**	0,355***			0,058	0,285***	0,021	0,228***
	(0,055)	(0,082)			(0,056)	(0,084)	(0,056)	(0,084)
Intermediate occupations	0,284***	0,368***			0,212***	0,300***	0,164***	0,210***
	(0,053)	(0,079)			(0,054)	(0,081)	(0,055)	(0,081)
Small employers and self-employed	0,226***	0,593***			0,115**	0,471***	0,019	0,307***
	(0,054)	(0,081)			(0,056)	(0,084)	(0,057)	(0,085)
Lower supervisory and technical occupations	0,342***	0,566***			0,241***	0,459***	0,154***	0,303***
	(0,048)	(0,072)			(0,050)	(0,075)	(0,051)	(0,076)
Semi-routine occupations	0,338***	0,683***			0,222***	0,549***	0,117**	0,350***
	(0,048)	(0,070)			(0,050)	(0,074)	(0,051)	(0,075)
Routine occupations	0,362***	0,680***			0,251***	0,555***	0,146***	0,375***
	(0,050)	(0,078)			(0,053)	(0,081)	(0,053)	(0,082)
Unemployed	0,345***	0,774***			0,237***	0,638***	0,129**	0,442***
	(0,058)	(0,081)			(0,060)	(0,084)	(0,060)	(0,085)
Missing	0,009	0,551***			-0,070	0,447***	-0,159*	0,313***
	(0,088)	(0,108)			(0,089)	(0,109)	(0,089)	(0,110)
Some qualification			0,364***	0,643***	0,274***	0,437***	0,117*	0,161**
			(0,060)	(0,075)	(0,062)	(0,079)	(0,063)	(0,081)
Left school - some qualification			0,333***	0,259***	0,259***	0,097	0,179***	-0,026
			(0,060)	(0,078)	(0,062)	(0,082)	(0,062)	(0,082)

Left school - no qualification			0,382***	0,452***	0,295***	0,253***	0,189***	0,058
			(0,061)	(0,078)	(0,063)	(0,082)	(0,064)	(0,083)
No school			0,499***	0,641***	0,386***	0,382***	0,226***	0,107
			(0,058)	(0,078)	(0,061)	(0,083)	(0,062)	(0,085)
Missing			0,510***	0,593***	0,403***	0,328**	0,192*	-0,097
			(0,096)	(0,156)	(0,098)	(0,159)	(0,100)	(0,162)
Age	0,994***	0,222***	0,994***	0,222***	0,996***	0,225***	0,949***	0,122***
	(0,022)	(0,018)	(0,022)	(0,018)	(0,022)	(0,018)	(0,023)	(0,021)
Age squared	-0,018***	-0,005***	-0,018***	-0,005***	-0,018***	-0,005***	-0,017***	-0,003***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Year of birth	0,055***	0,100***	0,050***	0,098***	0,053***	0,101***	0,055***	0,108***
	(0,009)	(0,018)	(0,009)	(0,018)	(0,009)	(0,018)	(0,009)	(0,019)
Year of birth squared	-0,001***	-0,000***	-0,001***	-0,000***	-0,001***	-0,000***	-0,001***	-0,000***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (*ref*, no siblings)	-0,055	-0,101*	-0,050	-0,094	-0,050	-0,075	-0,030	-0,011
	(0,037)	(0,061)	(0,037)	(0,061)	(0,037)	(0,061)	(0,037)	(0,062)
2 siblings (*ref*, no siblings)	0,059*	0,194***	0,068**	0,204***	0,061*	0,200***	0,061*	0,204***
	(0,032)	(0,054)	(0,032)	(0,054)	(0,032)	(0,054)	(0,032)	(0,054)
Parents separated	0,017	0,511***	-0,045	0,484***	0,002	0,490***	-0,010	0,436***
	(0,087)	(0,096)	(0,087)	(0,097)	(0,087)	(0,096)	(0,087)	(0,097)
Absent/dead mother	-0,038	0,224**	-0,073	0,190*	-0,031	0,193*	-0,040	0,138
	(0,082)	(0,101)	(0,082)	(0,100)	(0,083)	(0,102)	(0,083)	(0,103)
Absent/dead father	0,009	0,024	-0,014	0,135*	0,014	0,033	0,021	0,022
	(0,057)	(0,085)	(0,050)	(0,077)	(0,057)	(0,085)	(0,057)	(0,085)
African - Caribbean	-0,070	0,618***	-0,078	0,612***	-0,076	0,608***	-0,055	0,638***
	(0,074)	(0,062)	(0,074)	(0,062)	(0,074)	(0,063)	(0,074)	(0,065)
Indian, Pakistani, Bangladeshi	1,306***	-1,561***	1,294***	-1,542***	1,302***	-1,585***	1,282***	-1,625***
	(0,040)	(0,124)	(0,041)	(0,123)	(0,041)	(0,124)	(0,042)	(0,125)

Other Asian	0,401***	-1,225***	0,374***	-1,277***	0,402***	-1,246***	0,437***	-1,166***
	(0,094)	(0,199)	(0,094)	(0,199)	(0,094)	(0,199)	(0,095)	(0,199)
Other	0,465***	-0,403***	0,434***	-0,458***	0,479***	-0,444***	0,508***	-0,397***
	(0,101)	(0,149)	(0,100)	(0,149)	(0,101)	(0,150)	(0,101)	(0,151)
Rural area	0,050*	-0,357***	0,048	-0,356***	0,055*	-0,346***	0,077***	-0,306***
	(0,029)	(0,050)	(0,029)	(0,049)	(0,029)	(0,050)	(0,029)	(0,050)
yearly % GDP (t-12)	-0,054*	-0,166***	-0,055*	-0,170***	-0,056*	-0,170***	-0,046	-0,132**
	(0,030)	(0,055)	(0,030)	(0,055)	(0,030)	(0,055)	(0,030)	(0,054)
Other higher							0,247***	0,421***
							(0,050)	(0,080)
A level etc							0,117**	0,417***
							(0,050)	(0,069)
GCSE etc							0,362***	0,689***
							(0,045)	(0,067)
Other qual							0,373***	0,690***
							(0,050)	(0,083)
No qualification							0,468***	0,937***
							(0,047)	(0,076)
Missing							1,100*	1,207*
							(0,586)	(0,722)
Currently in education							-0,741***	-1,213***
							(0,079)	(0,071)

Table 2.B4. Entry into parenthood. Odds-ratios. Women.

	Model 1		Model 2		Model 3		Model 4	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
VARIABLES	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock	Marital	Out-of-wedlock
Lower managers & professionals	1,144**	1,427***			1,060	1,330***	1,021	1,257***
	(0,063)	(0,117)			(0,059)	(0,111)	(0,057)	(0,105)
Intermediate occupations	1,328***	1,445***			1,236***	1,350***	1,178***	1,234***
	(0,071)	(0,114)			(0,067)	(0,109)	(0,064)	(0,100)
Small employers and self-employed	1,254***	1,810***			1,122**	1,601***	1,019	1,359***
	(0,067)	(0,146)			(0,063)	(0,135)	(0,058)	(0,116)
Lower supervisory and technical occupations	1,408***	1,762***			1,272***	1,583***	1,166***	1,354***
	(0,068)	(0,127)			(0,064)	(0,119)	(0,059)	(0,103)
Semi-routine occupations	1,402***	1,980***			1,248***	1,732***	1,124**	1,419***
	(0,067)	(0,139)			(0,063)	(0,129)	(0,057)	(0,107)
Routine occupations	1,436***	1,973***			1,286***	1,742***	1,158***	1,455***
	(0,072)	(0,153)			(0,068)	(0,141)	(0,062)	(0,119)
Unemployed	1,412***	2,169***			1,267***	1,893***	1,138**	1,555***
	(0,082)	(0,175)			(0,076)	(0,159)	(0,069)	(0,132)
Missing	1,009	1,735***			0,932	1,563***	0,853*	1,367***
	(0,089)	(0,187)			(0,083)	(0,171)	(0,076)	(0,151)
Some qualification			1,439***	1,903***	1,315***	1,548***	1,124*	1,174**
			(0,086)	(0,143)	(0,081)	(0,122)	(0,071)	(0,096)
Left school - some qualification			1,396***	1,296***	1,296***	1,102	1,196***	0,974
			(0,084)	(0,102)	(0,080)	(0,090)	(0,075)	(0,080)

Left school - no qualification			1,465***	1,571***	1,343***	1,288***	1,208***	1,060
			(0,090)	(0,122)	(0,085)	(0,106)	(0,077)	(0,088)
No school			1,646***	1,898***	1,471***	1,465***	1,254***	1,113
			(0,095)	(0,148)	(0,090)	(0,122)	(0,078)	(0,095)
Missing			1,665***	1,809***	1,496***	1,388**	1,212*	0,907
			(0,160)	(0,283)	(0,147)	(0,221)	(0,121)	(0,147)
Age	2,703***	1,249***	2,702***	1,249***	2,706***	1,252***	2,583***	1,130***
	(0,061)	(0,022)	(0,061)	(0,022)	(0,061)	(0,022)	(0,060)	(0,023)
Age squared	0,983***	0,995***	0,983***	0,995***	0,983***	0,995***	0,983***	0,997***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
Year of birth	1,056***	1,105***	1,051***	1,103***	1,054***	1,106***	1,056***	1,114***
	(0,009)	(0,020)	(0,009)	(0,020)	(0,009)	(0,020)	(0,010)	(0,021)
Year of birth squared	0,999***	1,000***	0,999***	1,000***	0,999***	1,000***	0,999***	1,000***
	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)	(0,000)
1 sibling (*ref*, no siblings)	0,946	0,904*	0,951	0,910	0,951	0,928	0,970	0,989
	(0,035)	(0,055)	(0,035)	(0,056)	(0,035)	(0,057)	(0,036)	(0,061)
2 siblings (*ref*, no siblings)	1,061*	1,215***	1,071**	1,226***	1,063*	1,221***	1,062*	1,226***
	(0,034)	(0,065)	(0,035)	(0,066)	(0,034)	(0,066)	(0,034)	(0,066)
Parents separated	1,017	1,667***	0,956	1,623***	1,002	1,633***	0,990	1,547***
	(0,089)	(0,160)	(0,083)	(0,158)	(0,088)	(0,157)	(0,087)	(0,150)
Absent/dead mother	0,962	1,251**	0,930	1,210*	0,970	1,213*	0,961	1,148
	(0,079)	(0,127)	(0,076)	(0,121)	(0,080)	(0,124)	(0,079)	(0,118)
Absent/dead father	1,009	1,024	0,986	1,144*	1,014	1,033	1,021	1,023
	(0,057)	(0,087)	(0,050)	(0,088)	(0,058)	(0,087)	(0,058)	(0,087)
African - Caribbean	0,933	1,855***	0,925	1,844***	0,927	1,837***	0,946	1,892***
	(0,069)	(0,115)	(0,068)	(0,115)	(0,068)	(0,115)	(0,070)	(0,123)
Indian, Pakistani, Bangladeshi	3,691***	0,210***	3,647***	0,214***	3,678***	0,205***	3,603***	0,197***
	(0,147)	(0,026)	(0,149)	(0,026)	(0,153)	(0,025)	(0,151)	(0,025)

Other Asian	1,493***	0,294***	1,454***	0,279***	1,495***	0,288***	1,547***	0,312***
	(0,140)	(0,059)	(0,137)	(0,056)	(0,141)	(0,057)	(0,146)	(0,062)
Other	1,591***	0,668***	1,543***	0,632***	1,615***	0,641***	1,663***	0,672***
	(0,160)	(0,100)	(0,155)	(0,094)	(0,163)	(0,096)	(0,169)	(0,102)
Rural area	1,051*	0,700***	1,049	0,700***	1,056*	0,707***	1,080***	0,736***
	(0,031)	(0,035)	(0,030)	(0,035)	(0,031)	(0,035)	(0,032)	(0,037)
yearly % GDP (t-12)	0,947*	0,847***	0,947*	0,843***	0,946*	0,844***	0,955	0,876**
	(0,028)	(0,046)	(0,028)	(0,046)	(0,028)	(0,046)	(0,029)	(0,048)
Other higher							1,280***	1,524***
							(0,064)	(0,122)
A level etc							1,124**	1,517***
							(0,056)	(0,105)
GCSE etc							1,436***	1,993***
							(0,065)	(0,134)
Other qual							1,452***	1,994***
							(0,073)	(0,166)
No qualification							1,596***	2,553***
							(0,075)	(0,194)
Missing							3,005*	3,342*
							(1,762)	(2,411)
Currently in education							0,477***	0,297***
							(0,038)	(0,021)

Figure 2.A1. Kaplan Meier. Transition to 1st union

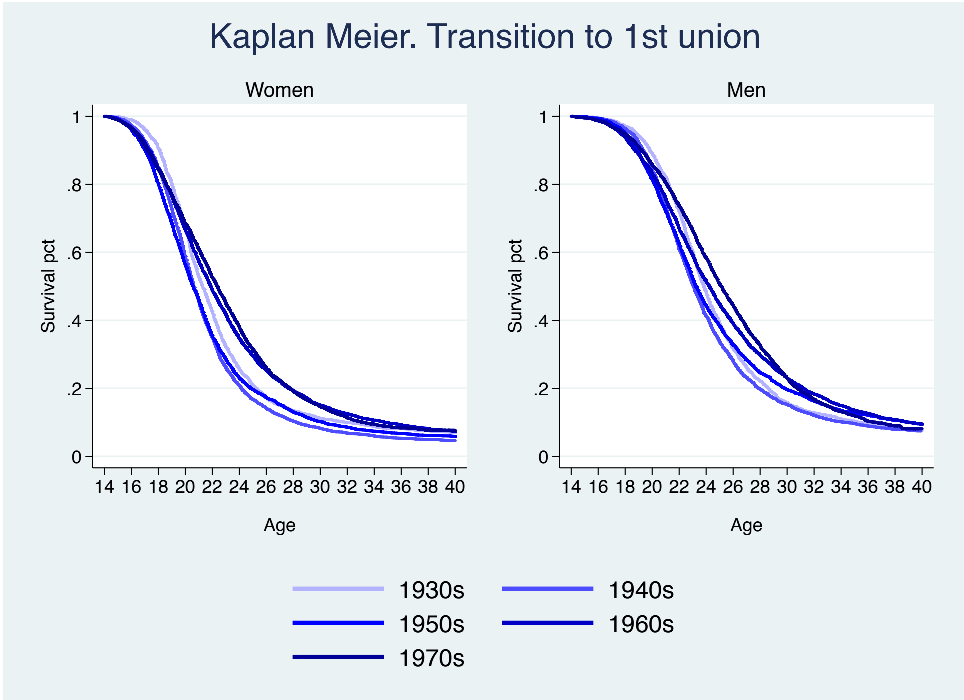


Figure 2.A2. Kaplan Meier. Transition to 1st marriage

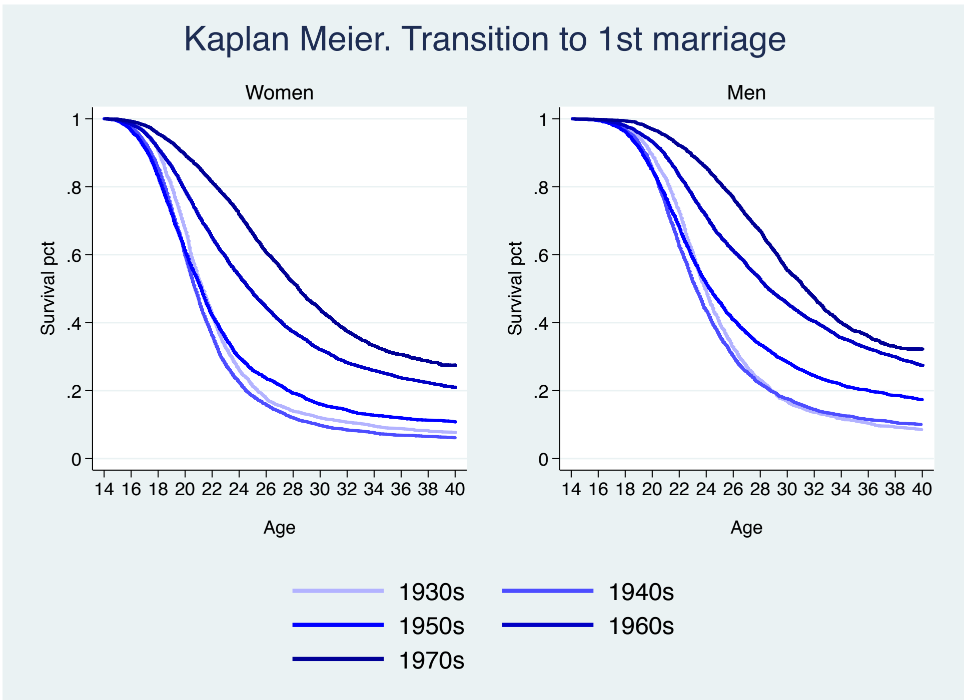


Figure 2.A3. Kaplan Meier. Transition to 1st marital birth

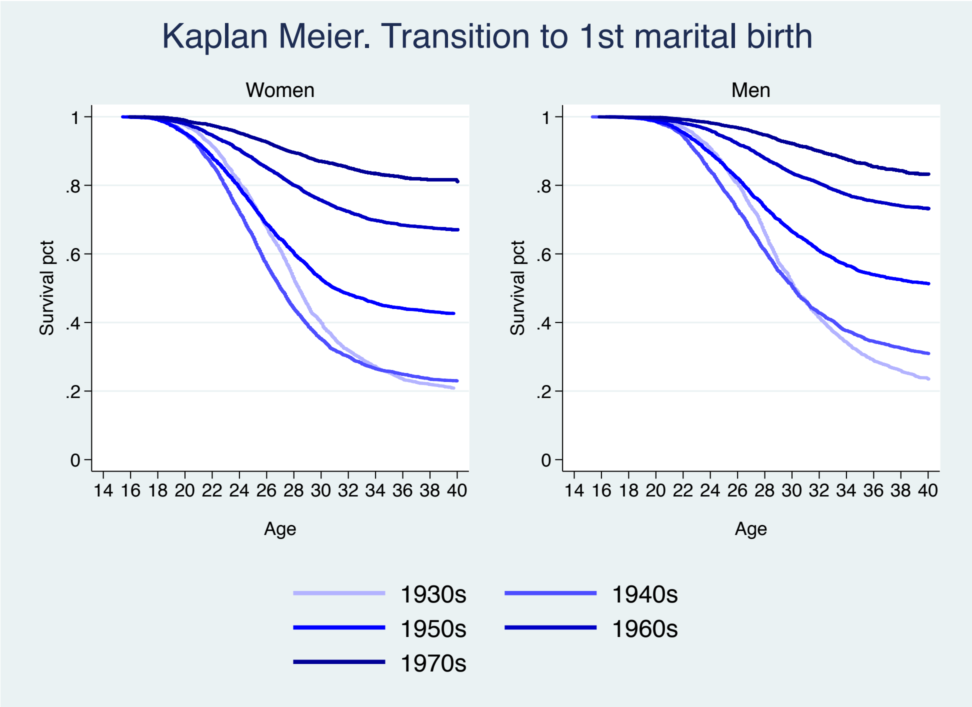
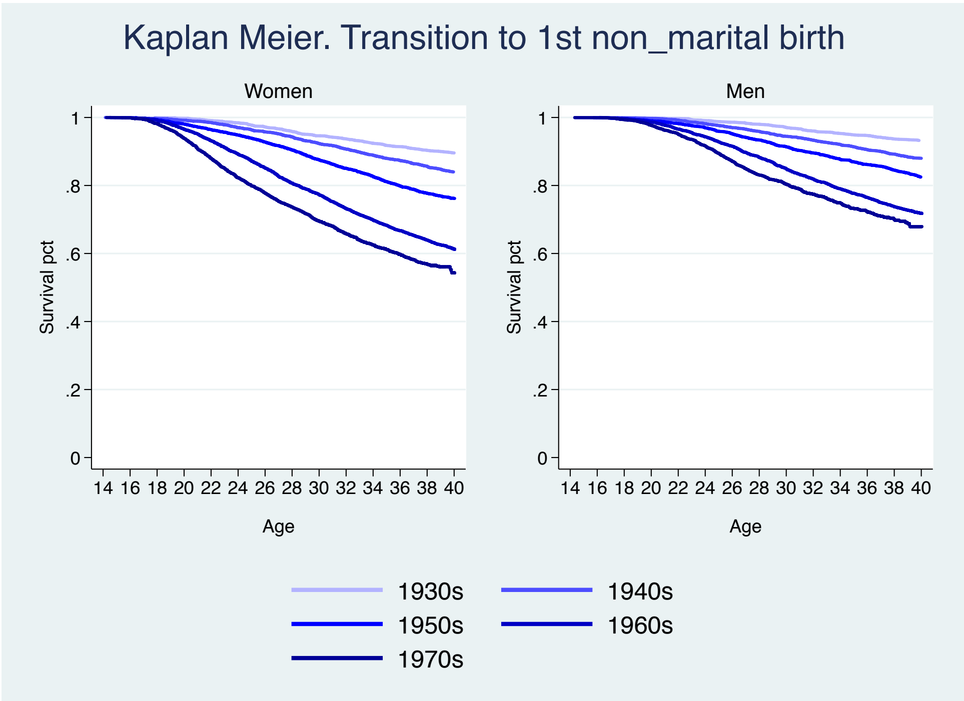


Figure 2.A4. Kaplan Meier. Transition to 1st non-marital birth



Chapter 3

Gender gap in repartnering: the role of parental status and custodial arrangements

Abstract

This study assesses whether parenthood influences repartnering for women and men and explores how repartnering is associated with parental status of the prospective partners. Previous research has not demonstrated whether gender differences in repartnering are conditional on the presence of children. This study aims to better disentangle the specific gender differentials in repartnering probabilities conditional on parenthood and child custody status. The analytical sample consists of 5,372 women and 3,375 men who reported at least one partnership dissolution in the British Understanding Society survey. Multilevel event history models with Markov Chain Monte Carlo simulations are used to estimate the probabilities of (a) finding a new partner and (b) finding a new childless partner or a new partner who has children. The results suggest that mothers, and to a lesser extent fathers, are less likely to repartner than their childless counterparts. Among parents who have child custody, there emerges a distinct gender gap because mothers exhibit a significantly lower rate of repartnering than fathers. Finally, coresident single parents are relatively less likely to repartner with childless individuals, and single fathers more frequently form two-parent stepfamilies than do mothers. This suggests the presence of a gender divide in repartnering that is especially apparent when child custody is taken into account. The presence of children also reduces the possibility of forming unions with childless individuals.

1. Introduction

The rise in cohabitation, divorce and separation, coupled with the higher frequency of repartnering, has produced greater diversity in partnering trajectories in most European countries. Over the last decades, both men and women are increasingly likely to enter into higher-order unions (Elzinga & Liefbroer, 2007), and this is also true across a wider age span (Beaujouan, 2012). The transition out of a partnership and into a new one has implications for the psychological readjustment of ex-partners (e.g. Tavares & Aassve, 2013; Wang & Amato, 2000) and for dependent children's wellbeing (Amato & Kane, 2011), but also for the formation of stepfamilies. Compared to childless couples, unions with step-children are more complex institutions and might be affected by a lack of clarity over roles within the family (Sweeney, 2010). Thus, the issue of who forms higher-order unions and what role children play in shaping their parents' behaviour is relevant for family functioning (Brown & Manning, 2009; Stewart, 2005).

Existing research on repartnering behaviour has primarily focused on women and has shown that, after a separation or a divorce, mothers are less likely to form a new partnership with respect to childless women (e.g., Wu & Schimmele, 2005; Beaujouan, 2012). Evidence for children's role in repartnering among men is more limited and inconsistent, due to the lack of detailed measures for father-child contact after union dissolution. Consequently, there is little evidence on inter-gender differences in repartnering by parental status (childless vs. parents) and by children's residential status (e.g., co-resident vs. non-resident). Furthermore, there has been little research into whether repartnering behaviour is influenced by the parental status of perspective partners, and it remains unclear how family formation varies by gender.

This study seeks to fill these gaps by assessing both how parental status is linked to entry into a new union and whether the union that is formed includes the partner's children. The first part of the analysis examines whether children's residential status, number and age are associated with lower chances of a new union for separated parents, when compared to their childless peers. The second part addresses the question of the influence of (own) children and a prospective partner's children on the likelihood of women and men forming a union, either with a childless individual or with a parent. Both analyses will contribute to the literature by assessing the existence of systematic gender differences in these repartnering patterns.

Drawing on data from the British survey Understanding Society, this study considers the repartnering process after first and higher-order marital and cohabiting union dissolutions in the 18 to 50 age range. The focus on all episodes of singlehood over a life course, in contrast to most existing research, is motivated by the increasing prevalence of individuals in multiple relationships (marriage and cohabitation) in Britain (Beaujouan & Ní Bhrolcháin, 2011; Sanchez Gassen & Perelli-Harris, 2015). Further, the data structure – with multiple episodes of singlehood nested within the same individual – requires standard multilevel event-history models combined with a multilevel approach and Markov Chain Monte-Carlo simulations. The advantage of these *frailty* models is that they explicitly address the issue of self-selection on unmeasured characteristics associated with union entry and exit (Allison, 1982).

2. Background

2.1 Repartnering, predictors and gender differences

Partnership trajectories have become increasingly complex as life-time marriages have become less common (Elzinga & Liefbroer, 2007) and as cohabitation has gradually been chosen as an alternative to marriage (e.g., Murphy, 2000; Perelli-Harris, 2014). Serial co-residential partnership has increased in European countries over the last four decades and Britain has one of the highest rates of partnership turnout (Galezewska et al., 2014; Perelli-Harris et al., 2012). Women and men who experienced at least one partnership break-up accounted for between 30 and 40 percent of those born between 1945 and 1970, and about two thirds of women and three quarters of men found a new partner (Beaujouan & Ní Bhrolcháin, 2011). It is also clear that the rise in serial co-resident partnerships has been driven by cohabitation rather than by marriages (Berrington & Stone, 2017) and has increasingly involved children born within dissolved unions (Boertien, 2016).

Empirical evidence shows that there are important differences in how men and women repartner (Wu and Schimmele 2005). Overall, men are more likely to repartner, with shorter spells between two consecutive unions (e.g., de Graaf & Kalmijn, 2003; Poortman, 2007; Wu & Schimmele, 2005). This difference widens with age (Beaujouan, 2012). In fact, marriage market mechanisms work in favour of men, who tend to find partners across a larger age range (Gelissen, 2004; Goldscheider & Sassler, 2006) and might have a later schedule of family formation or reformation (Lampard & Peggs, 1999;

Beaujouan, 2012). Past union experiences may also affect men and women differently. The demise of a union arguably hits women harder than men, particularly when it comes to the impact of prior unions and the duration of the union (Poortman, 2007). Women might be more prone to bear the emotional burden of a dissolution and might generally have less desire for a new relationship (Beaujouan, 2012).

Existing research has generally looked at second union formation after a divorce, and neglected higher-order and post-cohabitation repartnering, for both men and women. The first contribution of this study will be to fill this gap by looking at the formation of second and higher order unions for both men and women. By including the type and duration of previous unions, and the length of previous periods of singlehood for repeating events, it is possible to gain more knowledge about the cumulative influence of experiences of new union formation and to highlight gender differences.

2.2 The role of parenthood, parental custody and children's characteristics in repartnering

Parenthood is a key factor for intra-gender (mothers vs. childless women) and inter-gender (mothers vs. fathers) differentials (Goldscheider & Sassler, 2006; de Graaf and Kalmijn 2003; Poortman 2007; Wu and Schimmele 2005). Mothers are significantly less likely to repartner than childless women (e.g., Beaujouan, 2012; de Graaf & Kalmijn, 2003; Steele et al., 2005; Wu & Schimmele, 2005). This is especially the case for women with many (and young) children (Ivanova et al., 2013; Poortman, 2007), and for women with a non-marital first birth (Upchurch, et al., 2002).

A few studies have analysed how fatherhood influences men's new union prospects, but the evidence does not show any clear gap between childless men and fathers. Under certain circumstances, fathers are more likely to enter a union than childless men because fathers seem to be understood to be more reliable partners (Goldscheider & Sassler, 2006; Wu & Schimmele, 2005). Other studies have found no or a non-significant difference in partnering probabilities between fathers and childless men (e.g., de Graaf & Kalmijn, 2003; Ivanova et al., 2013; Sweeney, 1997; Skew et al., 2009). These findings do not clarify to what extent repartnering patterns are attributable to gender, parental status (parents vs. childless) and children's residence status (co-resident vs. non-resident), not to mention data limitations concerning children's characteristics (Beaujouan, 2012; Poortman, 2007; Goldscheider & Sassler, 2006; Sweeney, 1997; Wu & Schimmele, 2005). The second contribution of this study is to include a wider and more accurate array of information for children, such as their residence status, age and number, thus disentangling, in terms of repartnering possibilities, the role of parenthood, and children's residence, from gender.

There are three arguments which can explain the role of children in the repartnering process and the diverse responses of men and women to union prospects: opportunity, attractiveness, and need (e.g. Becker, 1981; de Graaf & Kalmijn, 2003; Ivanova et al., 2013; Kalmijn, 1998; Vanassche et al., 2015). Table 3.1 summarises the mechanisms leading to an increase (or a decrease) in the probability of union formation. From the combination of these mechanisms, I offer three distinct hypotheses on the role of: (1) individuals' parental status (co-resident vs. non-resident parents and childless individuals); (2) parents' gender and residence status (co-resident mothers vs. fathers); and (3) children's age and number.

The first argument affirms that financial commitments and time dedication to childcare reduce a parent's opportunities for finding a new partner (Koo et al., 1984), especially if there is more than one dependent child and even more so if all or some of these are still very young (e.g. Huerta et al., 2013). This reasoning should apply to a lesser extent to non-resident parents, though their visitation schedule and paternal duties, if any, may dissuade them from planning a new union (Lampard & Peggs, 1999). In Britain, where legislation encourages separated or divorced parents to negotiate children's arrangements through private agreements, mothers generally retain physical custody of the children (Blackwell & Dawe, 2003; Hunt & Roberts, 2004), while fathers are more likely to commit to temporary stays and financial help (Blackwell & Dawe, 2003; Trinder, 2010; Harding & Newnham, 2015). Similar numbers of fathers, ranging from 9% to 12%, are involved in children's shared (Trinder, 2010) and exclusive custody (Blackwell & Dawe, 2003; Peacey & Hunt, 2009), whereas around 10% of fathers, according to recent estimates (Poole et al., 2015), lose touch with their children altogether (as opposed to roughly 40% in the 1980s, Bradshaw & Miller, 1991). Hence, in Britain, a substantial proportion of adults at risk of entering a new co-residential union have children, either living with them or tied to them emotionally and financially, with the caretaking burden falling disproportionately on women.

The second argument holds that having a child decreases one's attractiveness to new potential partners. The presence of a child from a previous union may signal ongoing contact with the former partner (Monte, 2007), and scare away potential future step-parents (Stewart, et al., 2003). This motivation might vary by gender as previous studies on the mating process reveal asymmetric preferences: women are more inclined to form unions with partners who have children than are men (Bernhardt & Goldscheider, 2002; South, 1991). An explanation from the anthropology literature views women's childcare involvement as being less dependent on genetic inheritance than men's (Hofferth & Anderson, 2003; Waynforth, 2013). An alternative hypothesis (the "good father effect") depicts custodial fathers – as opposed to childless men and non-resident fathers – as impressing with commitment to their children's care and their dependability in a prospective family (Lappegård & Rønsen, 2013). Conversely, custody is considered normative for women and any retreat from maternal responsibilities may even hamper women's repartnering prospects.

The third argument emphasizes different factors that spur mothers and fathers to find a new partner. The need for a new partner may arise so as to compensate for financial loss (Jansen et al., 2009), the psychological distress caused by a separation (Wang & Amato, 2000), or distance from children (Tavares & Aassve, 2013). The mechanisms could be gender-specific. On the one hand, women might be more financially affected by a union dissolution than men, especially if they have dependent children (Dykstra & Poortman, 2010). On the other hand, separated mothers are less inclined to repartner compared to childless women, as they fear a new partner's interference with their established childcare routine (Beaujouan, 2012). Conversely, fathers with dependent children might purposely search for a new partner who could take on the role of stepmother and become a surrogate for a missing maternal figure (Bernhardt & Goldscheider, 2002).

Taking these three motives together, *I expect co-resident parents to be less likely to repartner than non-resident parents and childless individuals (H1)*. Parents living with dependent children might have lower chances of a new partnership relative to non-resident parents and childless individuals. This might be so either because they have fewer opportunities to find new partners (the opportunity hypothesis), or because potential partners might prove reluctant to take on the role of step-parents (the attractiveness hypothesis). *I also hypothesise that co-resident mothers are less likely to repartner than co-resident fathers (H2).* Children may be less of an obstacle to a new union for co-resident fathers, who arguably benefit from the extra bonus of "good-father" reliability (the attractiveness hypothesis) and who might be more motivated to search for a partner (the need hypothesis) than custodial mothers. Together with children's residence status, two other child characteristics should help disentangle these conflicting expectations. *I expect the number of children to be negatively associated with the chances of repartnering (H3a), while the age of children could be either positively or negatively associated with the chances of repartnering (H3b).*

Table 3.1. Summary of mechanisms of union formation with respect to three motives (needs, attractiveness, and opportunity), by parental status

Own parental status↓	**Probability of repartnering**		
	Opportunity	**Attractiveness**	**Need**
Childless person	+ no time constraints	+ no step-parenthood	+ partnership
Non-resident parent	– time constraints (children: number & young age)	– step-parenthood/part-time step-family formation (♂ mainly) – left custody of child (♀ mainly)	– step-parenthood/step-family formation (♂ mainly) + "good father" (♂ only)
Co-resident parent	– time constraints (children: number & young age)	+ partnership + economic support (♀ mainly)	+ partnership – childcare routine interference (♀ mainly) + economic support (♀ mainly) + childrearing support for toddlers & infants (♂ mainly)

Table 3.2. Summary of additional mechanisms of union formation with a childless partner or a parent, with respect to three motives (needs, attractiveness, and opportunity), by parental status.

Own parental status↓	**Probability of repartnering**		
	Opportunity	**Attractiveness**	**Need**
	Type of new partner↓		
	Childless partner		
Childless person	+ social activities	+ homogamy	+ fertility intentions

Non-resident parent	+ pool of potential partners	– no homogamy (partly)	
Co-resident parent	– social activities	– no homogamy	
	Parent		
Childless person	– no access to social networks of parents – pool of potential partners (♀ mainly)	– no homogamy – step-family formation	
Non-resident parent	+ access to social network of parents – pool of potential partners (♀ mainly)	+ homogamy (partly) – step-family formation + propensity to "social parenthood" (♀ mainly)	+ childrearing support (some)
Co-resident parent	+ access to social network of parents – pool of potential partners	+ homogamy – formation of a complex step-family + propensity to "social parenthood" (♀ mainly)	+ childrearing support

The number of children is positively associated with parents' time constraints for social activities (the opportunity hypothesis), while the age of the youngest child could either further limit time availability for custodial parents (the opportunity hypothesis) or urge fathers (more than mothers) to find childrearing support in a new partnership (the need hypothesis).

2.3 The role of partners' parental status

The second part of the analysis explores how likely different parental statuses (co-resident parent vs. non-resident parent and the childless) are to form a simple or a complex step-family, or a childless union. Two partners form a stepfamily if at least one of them brings children to the new union. If the children in a stepfamily are biologically linked to only one partner, this new union is a 'simple step-family'. If both partners bring children into the new family, they form a 'complex step-family' (Sweeney, 2010), and both assume the role of step-parent.

Table 3.2 provides a summary of the key theoretical arguments, besides those stated above, on the interplay of the parental status of both partners, along with the residential arrangements for their children. Then, I illustrate two hypotheses on the association of individual parental status (co-resident and non-resident parenthood, and childlessness) with the type of perspective partner. The first argument, which can be ascribed to 'opportunity' motives, predicts that people belonging to similar social networks are more likely to find a partner within these circles (de Graaf & Kalmijn, 2003). This might be the case with custodial and, to some extent, non-resident parents, who attend their children's social activities, such as schools and recreational clubs (Eggebeen & Knoester, 2001), while childless partners have better opportunities to meet partners without children during social activities or at work (De Graaf and Kalmijn, 2003). In line with Kalmijn (1998), these pathways are expected to enhance the matching prospects of the childless with the childless and co-resident parents with co-resident parents.

In second place, preferences also play a role in shaping the type of partnership. It is reasonable to predict a pattern of homogamy for childless individuals, reluctant to take on the role of stepparents. Similarly, separated parents, specifically those living with children, might be more inclined to form a union with other parents because the empathy between persons with the same parental status should favour the formation of a step-family where the partners both have children (Kalmijn, 1998). A competing hypothesis predicts, instead, that the two-parent stepfamily is less likely. Two single parents with established family routines may not find a new equilibrium as a stepfamily (Cherlin, 1978). For instance, partners' parenting style with the mutual step-children may not match: the reciprocal habituation of the step-siblings might prove problematic; or the step-parent's role might prove more ambiguous if a child's biological parent is in frequent contact with the family (Hofferth & Anderson, 2003). Faced with these prospects, two potential partners will perhaps renounce living together.

Third, following the need hypothesis, repartnering with a parent who has experience in raising children may provide valuable childrearing support. Women might be relatively more averse to embark on social motherhood (caring for someone's children) as their contribution to childrearing, which is generally higher than that of men, would perhaps prove to be time-consuming in a step-family (Goldscheider & Sassler, 2006). Nevertheless, for the opposite reason, women might embrace step-parenthood more readily than men would, because their childcare engagement is less dependent on genetic lineage (Waynforth, 2013) and because they are more tolerant than men of the family history of a potential partner (Goldscheider et al., 2009). Finally, the converging goal of the transition to parenthood might promote a union between two childless individuals, while the same motivation would not apply to individuals who already have children (Buber & Furnkranz-Prskawetz, 2000). The joint consideration of these three motives leads me to hypothesize that *childless individuals would be more likely to repartner with other childless individuals compared to their counterparts with dependent or non-resident children (H4); and that co-resident and non-resident parents would be as likely as childless people to repartner with other parents (H5).*

Few prior studies have explored the patterns of how parents or childless partners enter partnerships. Bernhardt & Goldscheider (2002) for Sweden, Goldscheider & Sassler (2006) for the U.S. and Vanassche et al. (2015) for Flanders found that custodial mothers are less likely to enter union formation with childless partners (versus no union formation) compared to childless women. Non-resident mothers are, meanwhile, less likely to repartner than are their childless counterparts. However, empirical findings on fathers' repartnering chances are not clear, arguably because of the different specifications of children's residence. The third innovation of this study is, thus, to shed light on the existence of a systematic gender gap in repartnering among parents and childless partners on the type of a new partnership.

3. Data and Methods

The empirical analysis is based on Wave 1 of Understanding Society (UKHLS) (https://www.understandingsociety.ac.uk) data, a British survey that started in 2009-2010 with a nationally representative sample of 43,674 individuals. It collects contemporary and retrospective information on employment, partnerships and fertility history, and has a longitudinal design with annual interviews. Fertility histories are drawn from individuals' reports that recall the date of birth of each child, the possible date of departure from the household, and the reason (e.g. death or a separation).

The following analyses concentrate on the life-course events reported by 5,372 women and 3,375 men born between 1950 and 1979 who reported at least one relationship breakup. The analytical sample results from the exclusion of (a) 9,044 individuals who did not mention any relationship or who did not accurately recall the dates of their unions; (b) 21,218 who never separated or divorced; (c) 4,118 who were born outside the cohort range; and (d) 547 who did not match further restrictions listed below. All singlehood spells begin in the month of union breakup and end when a new union begins, or are censored: (a) when the respondent turns 50; or (b) in the month of data collection, if no union is reported. The age span is set to end at 50 because the chances of living with dependent children are very low for the separated and the divorced. I account only for periods of singlehood following unions that lasted twelve months or longer, regardless of their legal status, in order to isolate the more stable co-resident unions. Further, I exclude all spells of singlehood lasting fewer than six months in order to study only individuals who spend a significant period on the repartnering market. In fact, repartnering can be endogenous because the sequence of union dissolution and partner search can be reversed, with individuals deciding to interrupt their relationships after having met another partner (e.g. Ivanova et al., 2013).

Table 3.3. Descriptive statistics, by gender and spell of singlehood

Variable	Women, mean or frequency			Men, mean or frequency		
	Spell of singlehood					
	1st	2nd	3rd	1st	2nd	3rd
Unions with a childless partner[a]	3243	534	78	2054	398	74
Unions with a parent[a]	241	50	9	479	121	19
Childless	31,3%	24,6%	24,9%	45,2%	36,1%	29,8%
Parents with non-resident children (only)	6,8%	12,1%	12,4%	27,8%	37,9%	48,1%
Parents with co-resident children	61,9%	63,8%	63,2%	27,0%	26,0%	22,1%
by number						
1	24,7%	25,7%	28,5%	15,1%	16,1%	14,4%
2+	37,2%	38,2%	34,7%	12,0%	10,0%	7,7%
by age						
0-6	22,1%	16,9%	16,6%	11,0%	11,2%	7,2%
7-12	19,2%	20,3%	19,7%	8,0%	7,3%	5,3%
13-18	13,2%	17,1%	15,5%	5,0%	4,6%	8,1%
Over 18	7,2%	9,4%	11,4%	2,8%	2,8%	1,4%
Ever married	66,6%	68,4%	74,6%	58,2%	53,4%	48,6%
Married in previous union	66,6%	40,2%	30,4%	58,2%	27,1%	18,5%
with children	53,3%	38,0%	27,6%	42,1%	23,1%	14,4%
Cohabiting in previous union	33,4%	60,4%	70,7%	41,8%	73,1%	82,2%
with children	15,6%	37,8%	48,3%	13,4%	35,1%	45,6%
Duration of previous union	8,74	6,30	4,91	7,88	5,23	4,70

Time since union dissolution	5,56	4,53	3,22	4,54	3,62	3,60
Age at union dissolution	35,21	39,43	41,7	35,34	38,85	41,60
Year of separation	1991,58	1996,10	1998,72	1991,84	1995,40	1997,90
Parents' separation	28,0%	36,3%	41,0%	26,4%	31,9%	35,6%
Birth cohort						
1950-1954	12,2%	12,0%	10,9%	13,8%	12,3%	12,5%
1955-1959	15,1%	14,6%	14,5%	16,9%	16,5%	16,8%
1960-1964	20,0%	21,4%	24,4%	20,1%	23,7%	25,5%
1965-1969	20,4%	23,2%	24,9%	22,0%	24,3%	23,1%
1970-1974	18,6%	19,5%	23,8%	16,4%	15,7%	15,9%
1975-1979	13,4%	9,2%	15,5%	19,7%	7,5%	6,3%
Ethnicity						
European	88,0%	91,9%	95,9%	89,6%	90,8%	92,3%
African-Caribbean	6,6%	5,4%	3,1%	5,3%	6,2%	5,2%
Indian, Pakistani, Bangladeshi	2,9%	1,1%	0	3,1%	1,4%	1,0%
South-East Asia	1,1%	0,7%	0,5%	0,5%	0,1%	0
Other	1,3%	0,8%	0,5%	1,3%	1,3%	1,4%
Education						
ISCED 0-1-2	35,2%	34,8%	33,7%	35,1%	33,2%	36,5%
ISCED 3-4	51,9%	53,9%	57,0%	54,4%	57,6%	55,8%
ISCED 5-6-7	12,9%	11,3%	9,3%	10,5%	9,2%	7,7%
Father has a job	84,5%	82,9%	80,8%	86,7%	84,1%	80,3%
N	5372	1196	193	3375	828	208
Person-periods	61269	11653	1398	31615	6499	1668

Note: source is Wave 1 of Understanding Society. Spells from 4th onwards are not shown. Means are calculated over persons, not person-periods, and per spell, unless otherwise stated, and are computed in the last month of each singlehood spell, except for age at union dissolution. Means for these time-varying variables are calculated in years refer to the last six-month episode of the spell. [a] indicates the number of events.

Periods of singlehood following a partner's death are not analysed. In keeping with the literature on partnership formation and dissolution (e.g., Berrington & Diamond, 1999; Wu & Schimmele, 2005), the relationship is considered ended when the partners' co-residence terminates and not when divorce is formalised. Accordingly, LAT individuals who do not co-reside are not defined as cohabiting in the survey, and their post-dissolution singlehood cannot be traced.

3.1 Variables

The measures of the variables are presented in Table 3.3, which illustrates dependent outcomes along with time-invariant and time-varying variables for partnership history, family background, year of separation and educational achievement. Two dependent variables are used in the analysis. The first represents the formation of a partnership (either cohabitation or marriage) in a self-reported month and year (1 = *a new partnership*; 0 = *singlehood*). The second defines the union start either with a childless partner (1 = *a new childless partner*; 0 = *singlehood*) or with a partner bringing children to the new partnership (1 = *a parent as a new partner*; 0 = *singlehood*). It is possible to detect whether the partner has some dependent children because the respondent is asked to give the birth date of her stepchildren, if any, and the period of co-residence with them for each union; no details about partner's non-resident children are provided. Three time-varying specifications of parental status are used to test the hypotheses. First, I account for (a) the presence of at least one co-resident child in the parent's household (1 = *some co-resident children*; 0 = *no co-resident children*); and (b) the existence of some non-resident children (1 = *non-resident children only*; 0 = *no non-resident children*). If both dichotomous variables equal zero, the respondent is thus childless. Understanding Society does not specify childcare time allocation and it is possible that separated parents, and particularly fathers, who claimed full-time residence for their children actually had joint custody or other part-time arrangements in place. For this reason, the figures concerning fathers' co-resident children might be partially inflated (and, similarly, those for non-resident children might be deflated). In the remaining two specifications, alternative indicators of the presence of co-resident children are included: a four-category variable indicating the age of the youngest co-resident child (1 = *0-6 years old*; 2 = *7-12 years old*; 3 = *13-18 years old*; 4 = *older than 18*), and a two-category variable representing the number of co-resident children (1 = *one child*; 2 = *two children or more*).

In the analyses, I control for four characteristics of previous unions. A dummy variable identifies the previous type of union (1 = *cohabitation*; 0 = *marriage*). A previous cohabitation should imply a lower level of emotional attachment to the former partner (Nock, 1995), a lower stigma in case of dissolution, and might motivate less caution about a new union, as opposed to a previous marriage (Poortman, 2007). Another dichotomous indicator captures the number of previous unions (1 = *two or more previous unions*; 0 = *one previous union*). The first breakup could be more destabilising than the following ones, and a number of separations are associated with a less risk-taking attitude (Poortman, 2007) and less commitment to partnerships (Lappegård & Rønsen, 2013). Six time-varying splines allow for us to model duration dependency of time from the break-up: 0 = *less than 1 year (reference)*; 1 = *1-2 years*; 2 = *2-4 years*; 3 = *4-6 years*; 4 = *6-10 years*; 5 = *10+ years*. Finally, a set of time-invariant spline functions indicates the duration of the previous union: 1 = *1-3 years*; 2 = *3-5 years*; 3 = *5-10 years*; 4 = *10+ years* (reference). I also account for the time-varying non-linear age of the respondent, from 1 = *18-22 years old* to 6 = *41-50 years old*, to control for the age effects caused by the variable pool of eligible partners in the repartnering market (Bumpass, et al., 1990). To examine trends over time in relationship instability, I control for the logarithm of the year of separation (continuous variable, range: 1966-2010) such as in de Graaf & Kalmijn (2003). Research on the intergenerational association of family structure implies that individuals' union instability may echo their own childhood family disruptions (Amato, 1996; Liefbroer & Elzinga, 2012), parents' interpersonal problems (Axinn & Thornton, 1996), or economic hardship (Kiernan, 1992). For this purpose, I use a dummy variable indicating whether parents were still a couple when the respondent was 16 (1 = *parents' dissolution*; 0 = *intact family*), and one indicator for whether the father was employed at the same age (1 = *father had a job*; 0 = f*ather was unemployed*). Finally, higher education positively influences either first partnership formation (e.g. Winkler-Dworak & Toulemon, 2007) or repartnering (e.g., Ivanova et al., 2013; Manning et al., 2003) due to greater attractiveness in the marriage market (Kaufman, 2000) and to better social integration (e.g. de Graaf & Kalmijn, 2003). Educational levels are measured by the cross-nationally comparable International Standard Classification of Education (ISCED; UNESCO, 2011): 1 = *ISCED 0-1-2 (less than or equal to lower secondary education)*; 2 = *ISCED 3-4 (upper and post-secondary education)*; 3 = *ISCED 5-6-7 (tertiary education and higher, reference)*. Eventually, a five-category time invariant variable captures a respondent's self-reported ethnicity: 1 = *European* (reference); 2 = *African-Caribbean*; 3 = *Indian, Pakistani, Bangladeshi*; 4 = *South-East Asian*; 5 = *Other ethnicity*.

4. **Analytical strategy**

The empirical strategy draws on the models developed by Steele (2008; 2011) for multilevel discrete-time event-history models for competing risks. This approach brings some major advantages. First, it makes possible the inclusion of repeated events rather than first-order transitions. Related to this, it can identify the influence of previous partnerships on new unions and helps disentangle the role of prior children from that of past partnerships. Further, it explicitly accounts for the individuals' unobserved heterogeneity – he or she has attractive personality traits or prefers a partnership to singlehood – which may lead to unstable relationships and multiple partnership entries. Therefore, the duration of the episodes for each respondent could be correlated with one another through the unobserved heterogeneity component (or frailty). The inclusion of this component in the analytical model corrects for episode dependency and tests the assumption made in standard methods that all durations must be independently distributed (Steele, 2011). The traditional non-frailty models are also performed as robustness checks and used to better identify the interplay of the individual-level component and its covariates.

The repartnering process is modelled as a sequence of singlehood episodes in which individuals risk entry into a new union. For each episode of singlehood, I construct a person/six-month file containing time-varying and invariant information about the individual, because the person-month dataset – although possible – is mathematically challenging for the software. Any spell starts in the semester of the last union dissolution and ends with the formation of a new union or with a censored spell for those who have not experienced the event by the end of the observation period. All episodes of singlehood are nested within each individual, which yields a two-level data structure. The multilevel (or random-effect) event history models adopted in this study were purposely developed for a hierarchical data structure (Steele, 2011).

Two models are tested. The first addresses the risk of entry into a union; the second, a competing risk model, addresses the risks of transition from singlehood to a new union with (1) a childless partner (vs. staying single); or (2) with a partner bringing children (vs. staying single), and employs two binary response models. No direct transition from state (1) to (2) occurs in the sample and is, thus, modelled in the analyses. The hazard of making a transition to the new state can be defined as a two-level random effects logistic model:

$$h_{ijt} = \log\left(\frac{p^{(r)}{}_{ijt}}{1 - p^{(r)}{}_{ijt}}\right)$$

where r $(r = 1)$ in the first model, and r $(r = 1, 2)$ in the second model and $p^{(r)}{}_{ijt}$ is the probability that a transition r occurs at time t during episode j for the individual i.

$$h_{ijt} = \alpha_{ij}{}^{(r)}(t) + \beta_{ij}X_{ijt}{}^{(r)} + \gamma_i W_i{}^{(r)} + u_i{}^{(r)}$$

where $\alpha_{ij}{}^{(r)}(t)$ is a function of time and consists of linear splines capturing the duration of the single status after union dissolution; $X_{ijt}{}^{(r)}$ and $W_i{}^{(r)}$ are vectors of time-varying and invariant covariates with, respectively, coefficients β_{ij} and γ_i; $u_i{}^{(r)}$ capture individuals' random effects and are assumed to follow a normal distribution with zero mean and variance σ^2. The model assumes that, conditional on u_i, the duration of episodes for the same individual are independent. I present models for distinct samples of men and women in Tables 4 to 6, and pooled models of men and women to highlight key differences in parental status by gender in Figures 3.1 and 3.2.

The competing risk model is computationally demanding and is estimated with Markov Chain Monte-Carlo (MCMC) methods with the MLwiN package runmlwin for STATA 13 (Leckie & Charlton, 2012b). The starting values of coefficients are derived from the IGLS (Iterative Generalized Least Squares) algorithm. The MCMC estimation includes a burn-in of 10,000 iterations, followed by a monitoring period of 100,000 iterations. In addition, I apply parameter expansion to improve convergence.

Although the multilevel approach brings clear advantages, the interpretation of the results requires that certain limitations be considered. First, the model does not allow for unobservable time-varying characteristics. Therefore, it must be assumed that, for instance, individuals' predisposition to relationship hopping and attractiveness is stable over time. Second, though the estimation of individual-level heterogeneity tackles possible self-selection on unmeasured characteristics associated with multiple union entries and exits, the coefficients should still be read as associations rather than causal effects. For instance, the presence of co-resident children is not exogenous with respect to repartnering. Indeed, it is possible that single parents selectively choose to live without children to enhance their chances on the repartnering market.

5. Results

Model 1 in Table 3.4 (women) and Table 3.5 (men) presents the effect of children's residence status on parents' chances of forming a new partnership, with childless individuals as the reference category. The results are presented as odds ratios, which represent the relative likelihood of someone with given characteristics entering a union, compared to someone who remains single.

The presence of co-resident children in the household is associated with less frequent repartnering for mothers (about 40% lower) as compared to their childless counterparts and to non-resident mothers: it does not, though, decrease the chances of new partnerships for men. These findings support the idea that the combined mechanisms of attractiveness, need and opportunities ultimately hamper the repartnering chances of custodial parents differently by gender.

Models 2 and 3 address how new union formation is conditioned by childcare burdens, which is proxied by the age of the youngest co-resident child, and by the number of dependent children. Mothers' repartnering odds are sizeably reduced in the earlier stages of a child's life, and pick up slightly as the child grows, but they remain significantly lower than for childless women and non-resident mothers. This negative "child age gradient" on new union odds does not emerge from the analysis of fathers. The odds for custodial fathers with dependent children aged 0 to 6 are comparable to those of childless men, and even decrease as the child ages. The other indicator of childcare burden confirms the hypothesis that the number of children influence parents' repartnering. Mothers' odds appear to be coherent with the hypothesis that raising two or more children is more burdensome than raising one child (the coefficients are statistically different at 95% level), while the likelihood of repartnering with co-resident fathers does not significantly change according to the number of dependent children. A number of robustness checks (i.e. modifying the age range, the time windows of the previous unions and the time elapsed since the dissolution) did not show relevant changes.

Table 3.4. Effects of characteristics of parenthood status on new union formation. Women.

	Model1		Model 2		Model 3	
	OR	SE B	OR	SE B	OR	SE B
No children (*Ref*)						
Non-resident only	1,020	0,084	1,009	0,084	1,142	0,095
Some co-resident	0,597***	0,028				
# co-resident						
1 child			0,636***	0,032		

2+ children			0,560***	0,031		
Youngest co-resident						
0-6 years					0,575***	0,028
7-12 years					0,626***	0,035
13-18 years					0,750***	0,053
Over 18 years					0,806**	0,081
Prev. cohabitating	1,045	0,046	1,039	0,046	1,032	0,045
2+ previous unions	0,994	0,063	0,993	0,063	1,002	0,063
Time since dissolution						
(Ref: 0-1 years)						
1-2 years	0,913*	0,046	0,916*	0,047	0,906*	0,046
2-4 years	0,929	0,054	0,934	0,054	0,910	0,052
4-6 years	0,888	0,070	0,895	0,071	0,856**	0,068
6-10 years	0,776**	0,077	0,785**	0,078	0,732***	0,073
Over 10 years	0,656***	0,092	0,666***	0,093	0,594***	0,084
Duration of last union						
(Ref: 10+ years)						
1-3 years	0,703***	0,047	0,686***	0,047	0,734***	0,050
3-5 years	0,817***	0,050	0,800***	0,050	0,852***	0,052
5-10 years	0,748***	0,043	0,740***	0,043	0,781***	0,046
Age (Ref: 18-22 years)						
23-26 years	1,182*	0,106	1,192**	0,107	1,182*	0,105
27-30 years	1,173	0,119	1,191*	0,122	1,167	0,118
31-35 years	0,911	0,104	0,930	0,107	0,897	0,102
36-42 years	0,679***	0,090	0,692***	0,093	0,642***	0,086
43-50 years	0,392***	0,062	0,395***	0,063	0,355***	0,056
Education						
(Ref: ISCED 5-6-7)						
ISCED 3-4	1,017	0,041	1,019	0,041	1,011	0,045
ISCED 0-1-2	0,773***	0,046	0,780***	0,047	0,768***	0,040
ρ	0,089		0,094		0,078	
χ^2	1355		1354		1417	
N	5372		5372		5372	
Persons-periods	74763		74763		74763	

Note: *OR* is "odds ratios". *SE B* is "standard error of the coefficient". Additional controls: whether parents separated by age 16, year of separation log, ethnicity and employed father when the individual was 16.

***<0.01; ** <0.05; * <0.1

The influence of a previous partnership history is measured in terms of experience of cohabitation (vs. marriage) in the previous union, the number of previous unions and the length of the previous partnership. Among these, only the last seem to matter for men and women, as the chances of repartnering increase significantly after exiting longer relationships. The time elapsed since the end of the previous union has a non-linear effect on repartnering, with men's chances falling steeply after six years and women's after one year. Repartnering probabilities are essentially flat from eighteen through 35 and decrease sharply thereafter. Those with little education are severely disadvantaged compared with those who have a degree. Among family background characteristics, parents' partnership stability is positively associated only with men's new union formation, and this holds true, too, for the father's employment.

Table 3.5. Effects of characteristics of parenthood status on new union formation. Men.

	Model1		Model 2		Model 3	
	OR	SE B	OR	SE B	OR	SE B
No children *Ref*						
Non-resident only	1,073	0,061	1,073	0,061	1,056	0,058
Some co-resident	0,932	0,053				
# co-resident						
1 child			0,937	0,061		
2+ children			0,925	0,069		
Youngest co-resident						
0-6 years					0,947	0,060
7-12 years					0,971	0,070
13-18 years					0,780*	0,074
Over 18 years					0,731*	0,111
Prev, cohabitating	1,081	0,056	1,081	0,056	1,081	0,054
2+ previous unions	1,005	0,064	1,005	0,065	1,006	0,062
Time since dissolution						
(Ref: 0-1 years)						
1-2 years	0,924	0,055	0,924	0,055	0,927	0,051
2-4 years	0,918	0,065	0,918	0,065	0,924	0,062
4-6 years	0,924	0,089	0,924	0,089	0,931	0,089
6-10 years	0,738**	0,089	0,738**	0,089	0,744**	0,089
Over 10 years	0,730*	0,123	0,730*	0,123	0,742	0,119
Duration of last union						

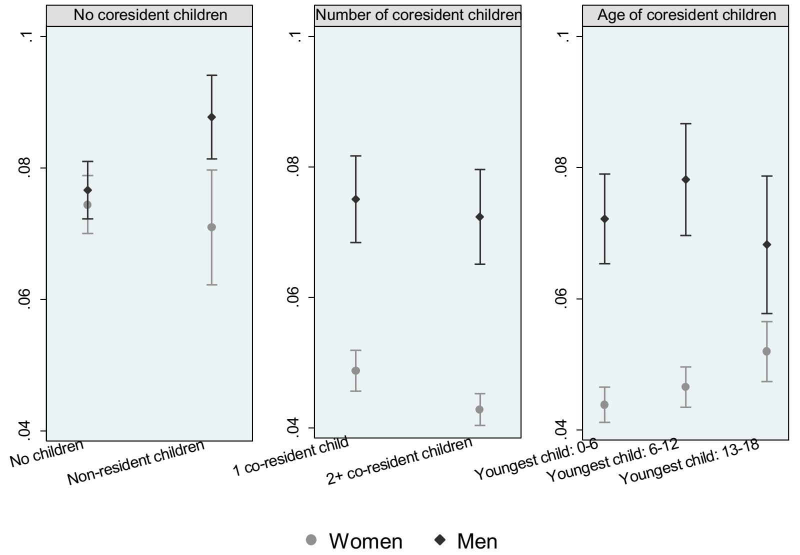
(Ref: 10+ years)						
1-3 years	0,631***	0,050	0,630***	0,050	0,620***	0,046
3-5 years	0,635***	0,047	0,634***	0,047	0,623***	0,044
5-10 years	0,677***	0,048	0,677***	0,049	0,663***	0,045
Age (Ref: 18-22 years)						
23-26 years	1,244*	0,169	1,244*	0,169	1,243*	0,158
27-30 years	1,246*	0,182	1,247*	0,183	1,244*	0,172
31-35 years	0,979	0,162	0,980	0,163	0,979	0,155
36-42 years	0,706**	0,145	0,707**	0,145	0,712**	0,141
43-50 years	0,395***	0,100	0,395***	0,100	0,407***	0,100
Education						
(Ref: ISCED 5-6-7)						
ISCED 3-4	0,935	0,042	0,935	0,042	0,939	0,049
ISCED 0-1-2	0,658***	0,056	0,658***	0,056	0,660***	0,039
ρ	0,061		0,067		0,067	
χ^2	497		496		505	
N	3375		3375		3375	
Persons-periods	40095		40095		40095	

Note: *OR* is "odds ratios". *SE B* is "standard error of the coefficient". Additional controls: whether parents separated by age 16, year of separation log, ethnicity and employed father when the individual was 16.

***<0.01; ** <0.05; * <0.1

Finally, unobserved heterogeneity plays a role. The total variance explained by individual frailty ranges from about 6 to 9 percent and is significantly different from zero. The alternative models without the individual-level component (not shown) present lower odds ratios for the splines of singlehood time and slightly higher odds ratios for the proxies of custodial parenthood. Unobserved heterogeneity is thus negatively associated with the duration dependence of repartnering and positively associated with co-resident parenthood. This implies that people with children possess traits like family-oriented values that increase the likelihood of union formation.

Figure 3.1. Probability of a new union by parental status. Number and age of co-resident children.



Note: Average predicted probability in each 6-month episode. Pooled sample of men and women

Pooling men and women in the same model and interacting gender with the other covariates helps me identify how the association between parental status and new union formation differs between men and women. At first glance, Figure 3.1 highlights a significant imbalance between genders in terms of repartnering. Fathers' chances of finding a new partner remain steadily above those of mothers, regardless of children's residence status, number and age. Only the repartnering probabilities of childless men and women do not differ significantly. Assuming that the number of children and their age are good proxies for the childcare burden and that these affect custodial mothers and fathers to a similar extent, the findings clearly reveal gender-driven mechanisms of repartnering. However, it is not possible to test exhaustively whether it is the "attractiveness hypothesis" or the "need hypothesis", or both, that explain the gap between custodial fathers and mothers.

Table 3.6 illustrates the results of the competing risk model of union formation with a childless or a parent partner. Models are presented for women and men, separately. The results are presented as odds ratios and represent the relative risk of an individual with determined characteristics entering a specific union (with a childless partner or with a person with co-resident children), compared to remaining single. Both men and women with dependent children are less likely to enter a union with partners without children (vs. remaining single), compared to childless individuals.

Table 3.6. Competing risks model. Effects of characteristics of parenthood status on new union formation by type of partner. Odds ratios.

	Women		Men	
VARIABLES	Partner without children	Partner with children	Partner without children	Partner with children
	Model 4		Model 5	
No children (*Ref=1*)				
Non-resident children (only)	0,990	1,179	1,116	1,317***
Some co-resident children	0,570***	1,059	0,845***	1,224*
N	5372		3375	
Persons-periods	74763		40095	

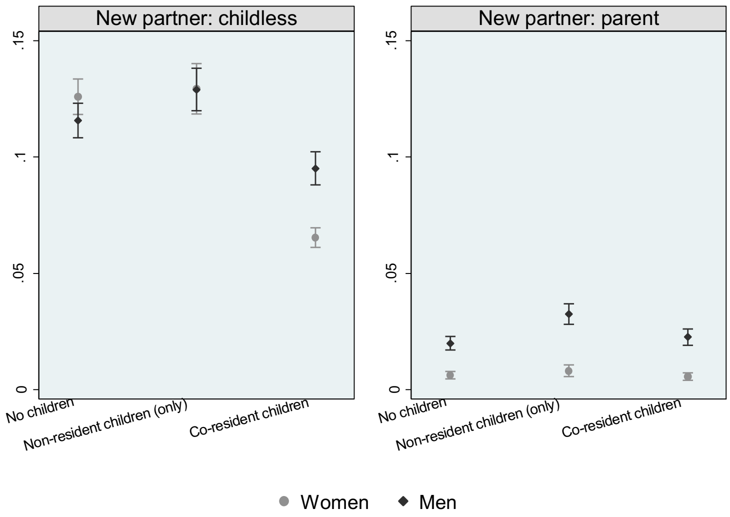
Note: Separate samples of men and women. Additional controls: any previous marriage (dummy), more than one previous union, time since union dissolution, length of previous union, family status at age 16, logarithm of year of separation, age, ethnicity, education, father employed when respondent is 16.

***<0.01; ** <0.05; * <0.10.

Conversely, custodial mothers are as likely as their childless counterparts to enter a union with other parents, and fathers are marginally more likely to do so. The combination of these results highlights a partner selection mechanism. Custodial parents' lack of attractiveness within the pool of childless people could be made up for by assortative mating mechanisms between single parents. Fathers with non-resident children are significantly more likely to repartner with a parent (vs. remaining single), but not to form unions with childless partners (vs. remaining single), at least with respect to the childless. Non-resident mothers, instead, do not systematically differ from childless women in repartnering. These outcomes suggest that child custody, rather than parenthood per se, lowers parents' risks of new union formation with a childless partner and that potential childless partners might not be motivated to embrace the role of step-parents.

Figure 3.2 shows women and men's probabilities of repartnering with childless partners and with co-resident parents in three different parental conditions: childlessness; non-resident parenthood; and co-resident parenthood. Distinct "inter-gender" behaviours of conjugal reconstruction emerge. Childless women and men have not significantly different repartnering probabilities with other childless partners. However, the condition of custodial parent implies a distinct gap between mothers and fathers in terms of repartnering with non-parents. Further, a gender imbalance shows up in repartnering probabilities with a co-resident parent for all three parental conditions. Men are significantly more likely to enter unions with parent partners than are women.

Figure 3.2. Probability of a new union with a childless partner or a partner with co-resident children by parental status.



Note: Average predicted probability in each 6-month episode. Pooled sample of men and women

These findings do not support the argument that men opt out of step-parenthood more than women. Conversely, the evidence suggests that the relatively higher probability that men will partner with parents, compared to women, could also be driven by the larger availability of separated custodial mothers. In conclusion, these results do not support either the hypothesis of women's greater willingness to care for "someone else's children" or men's supposedly greater resistance to embarking on step-parenthood, as hypothesised by Bernhardt & Goldscheider (2002).

6. Discussion

This study has examined the determinants of new union formation and the competing risks of entering a union with a childless partner and with a partner who has children. Few studies have taken together the role of children and potential partners in new relationships. This article improves upon previous work by offering, through an innovative methodology, an inter-gender analysis for Britain, including people's full partnership history.

I tested the hypothesis that the presence of co-resident children is negatively associated with new union formation, and particularly with entering a union with a childless partner. I also predicted that parental responsibilities might be associated with a gender gap in terms of men and women's new union prospects, and influenced different transitions to co-resident step-parenthood. The results partially confirm these hypotheses. In the first part of the study, I find that only co-resident mothers repartner at a lower rate than their childless counterparts do, while fathers do not. This evidence deviates from previous studies, showing that fathers with custody are not quicker to repartner than other men (in contrast with Goldscheider & Sassler, 2006), and that the gender differences in repartnering are significant even when their parenthood status is considered (in contrast with Ivanova et al., 2013). Further, the parental childcare burden – proxied by children's age and number – has a negative influence on mothers' repartnering prospects, but not on fathers'. The material limitations associated with the presence of children, which reduce parents' time for meeting potential partners and for forming new relationships, might explain the gap between custodial mothers and childless women. But these limitations do not untangle the gender gap between custodial parents.

Assuming that childrearing constraints are the same for single mothers as for fathers, this gap can be partly explained by the combination of custodial fathers' relatively greater attractiveness or by their more proactive partner search, or, indeed, by both. In this respect, the "good-father" hypothesis might be justified by the very specific socio-demographic profile of custodial fathers (Haux et al., 2016), who perhaps appear more dependable to partners (Lappegaard & Ronsen, 2013). Also, childbearing expectations for the new union may differently shape the repartnering behaviour of men and women. Fathers are able to have children at older ages and may, thus, be in less of a hurry in their quest for a new conjugal experience. By contrast, mothers who are still in their childbearing years may look for a partner to have children with, but they may also renounce childbearing as their reproductive cycle ends earlier. Finally, some studies have also stressed that women may be more susceptible to the demise of a relationship (e.g. Poortman, 2007) and that they tend to disengage from romantic life in the aftermath of a failure (Beaujouan, 2012). The presence of children could be a further reason to renounce a new partnership so as to avoid interference in an established childcare routine. Other studies have attributed this gender gap to age-dependent behaviour and to partners' availability, as women's partnership formation decreases at the beginning of their early thirties, relative to men (Beaujouan, 2012). Additional research focusing on age-specific repartnering behaviour, an issue not addressed in this study, might shed some more light on gender differences.

Similarly to Ivanova et al.'s (2013) findings, the repartnering of non-custodial parents proceeds at the same pace as that of childless individuals, regardless of gender. Two explanations seem plausible and they are not necessarily mutually exclusive. It could be that non-resident parents, freed from child-custody burdens, behave similarly to childless individuals ("constraint hypothesis"). Also, it is possible that non-resident parents face some constraints that impair their union prospects (such as duties with their absent children and potential partners' scepticism about their previous relationships), and that they compensate for this by seeking a partner to make up for their children's absence.

The pooled model also highlights a gap between non-resident fathers and mothers. However, the interpretation of this divide has to account for the different profiles of the two groups. The overwhelming majority of fathers do not retain child custody and they are heterogeneous in many domains (e.g. Haux, et al., 2015; Poole et al., 2015). Women who live apart from their children are rarer and generally represent two very distinct conditions: those who voluntarily relinquish child custody in order to guarantee their environmental stability, for instance, after leaving the household to live with a new partner; and those who are involuntarily removed from parental tasks, having lost custody rights via a court decision, or because children themselves refuse contact (Kielty, 2006). Those who voluntarily give up on child custody might be expected to trigger separation and to more rapidly set up a new union when they find a new partner (Beaujouan, 2012); while those who have involuntarily been cut off from their children might find that the pathway to repartnering could be impaired by troubled socio-economic conditions (e.g., Goldscheider & Sassler, 2006). Therefore, the sample exclusion of individuals who repartner within six months of their union dissolution – which might rule out most women from the first group – could partly explain the repartnering gap between non-resident mothers and fathers.

The second part of the analyses highlights how custodial responsibilities also shape the type of new union. In general, custodial children are a deterrent to potential unions with childless partners, but not with other parents: co-resident mothers are not less likely to repartner with other parents and custodial fathers have marginally higher odds of repartnering with other parents as compared to their childless counterparts. The following mechanisms might shed light on these results. In the first place, co-resident parents may prove more attractive to other parents than they are to childless men and women because of partnership homogamy, as also shown by Vanassche et al. (2015). Second, childless men and women might be unwilling to embrace step-parenthood, a challenging trial when it comes to role definition within the couple and relationship build-up with non-biological children.

The idea that custodial parenthood – and not parenthood per se – is unattractive in the repartnering market is also borne out by the inter-gender comparison of probabilities for a new union. Women's co-resident children reduce the odds of repartnering with childless partners to a larger extent than men, while the few women who do not retain child custody are as likely to enter new unions with childless partners as non-custodial fathers are. A gender gap in the probability of a new union with a parent emerges clearly: here, men's chances are notably greater than women's regardless of their parental status. A plausible explanation may lie in the prevalence of custodial mothers (vs. fathers) among potential partners on the remarriage market, as hypothesised by Bernhardt & Goldscheider (2002).

This study has some limitations. First, the nature of child custody is not exogenous to parental characteristics; some parents, for instance, may purposely reject their parental responsibilities in order to enhance their chances of repartnering. Further, the data have not allowed me to measure the proportion of time separated parents devote to their children. The descriptive statistics suggest that fathers with part-time custody sometimes claim to have full childcare responsibility. This measurement error could inflate the rate of repartnering of custodial fathers, some of whom may only have part-time arrangements. Finally, I did not address the influence of parental status on whether individual choose cohabitation or marriage: the sample size was insufficient. Future research addressing this repartnering mechanism might shed more light on key issues such as partnership stability and family functioning among newly established couples and step-families.

Despite these data shortcomings, the repartnering gap finds further empirical support in this analysis of British data. This divide is explained by mothers' lower repartnering rates, both with childless partners and with parents, compared to other women and men, and is further shown by the greater number of women among custodial parents. More gender-egalitarian provisions in child custody could help balance out the realities faced by separated parents and ultimately ease mothers' new union formation.

Recent British research suggests that the 50%-50% custody division represents a very small share of custody arrangements, while the 65%-35% division is more common, be it among those who separated on friendly terms or through judicial procedure (Trinder, 2010). This arrangement is not formalized by any specific legislative provision though the Government has undertaken some concrete steps to expand its use (The Ministry of Justice & Department for Education, 2012), and the Children and Families Act (2014, p. 11) has formally replaced the dichotomy "residence order" vs. "contact order", to identify, respectively, the custody holder and the other parent with a more neutral "child arrangements order" (Harding & Newnham, 2015). Growing evidence about repartnering, including this study, suggests the need for gender-equal policies in post-separation child arrangements.

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Appendix

Table 3.A1. Robustness check. Alternative version of Table 3.4

Sample including people exiting all unions (also shorter than 12 months) and experiencing at least 1-month singlehood episodes

	Women			Men		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
No children (Ref=1)						
Non-resident children only	0,997	0,984	1,127	1,079	1,079	1,065
	(0,080)		(0,091)	(0,057)	(0,057)	(0,057)
Some co-resident children	0,584***			0,913*		
	(0,025)			(0,049)		
# co-resident						
1 child		0,622***			0,922	
		(0,029)			(0,056)	
2+ children		0,545***			0,900	
		(0,028)			(0,066)	
Youngest co-resident			0,736***			0,779**
0-6 years			0,559***			0,926
			(0,025)			(0,059)
7-12 years			0,606***			0,957
			(0,032)			(0,075)
13-18 years			0,736***			0,779**
			(0,050)			(0,083)
Over 18 years			0,782**			0,712**
			(0,076)			(0,122)
Prev. cohabiting	1,057	1,051	1,046	1,151***	1,151***	1,151***
	(0,044)	(0,044)	(0,043)	(0,056)	(0,056)	(0,056)
2+ previous unions	0,965	0,963	0,973	0,946	0,945	0,947
	(0,059)	(0,059)	(0,060)	(0,054)	(0,054)	(0,054)
Time since dissolution (Ref: 0-1 years)						
1-2 years	1,038	1,041	1,031	1,108*	1,108*	1,111**
	(0,049)	(0,049)	(0,049)	(0,058)	(0,058)	(0,059)
2-4 years	1,074	1,080	1,056	1,053	1,053	1,058

	(0,056)	(0,056)	(0,055)	(0,063)	(0,063)	(0,063)
4-6 years	1,049	1,058	1,018	1,134	1,133	1,139*
	(0,073)	(0,074)	(0,071)	(0,088)	(0,088)	(0,090)
6-10 years	0,974	0,986	0,929	0,901	0,901	0,904
	(0,083)	(0,084)	(0,080)	(0,086)	(0,086)	(0,088)
Over 10 years	0,874	0,889	0,806*	0,850	0,850	0,858
Duration of last union (Ref: 0,5-1 years)						
1-3 years	1,294***	1,295***	1,292***	1,140**	1,140**	1,140**
	(0,068)	(0,068)	(0,067)	(0,061)	(0,061)	(0,061)
3-5 years	1,506***	1,512***	1,504***	1,158**	1,158**	1,156**
	(0,088)	(0,088)	(0,086)	(0,070)	(0,070)	(0,070)
5-10 years	1,377***	1,396***	1,373***	1,247***	1,248***	1,242***
	(0,089)	(0,092)	(0,088)	(0,087)	(0,087)	(0,087)
10+ years	1,862***	1,908***	1,778***	1,877***	1,881***	1,904***
	(0,136)	(0,141)	(0,130)	(0,155)	(0,156)	(0,159)
Year of separation	0,997	0,997	0,998	1,004	1,004	1,004
	(0,002)	(0,002)	(0,002)	(0,003)	(0,003)	(0,003)
Living with biological parents at 16	1,086**	1,090**	1,082**	1,104**	1,104**	1,105**
	(0,044)	(0,045)	(0,043)	(0,055)	(0,055)	(0,055)
Age						
23-26 years	1,111	1,119	1,107	1,298***	1,299***	1,298***
	(0,078)	(0,079)	(0,077)	(0,121)	(0,121)	(0,121)
27-30 years	1,067	1,081	1,056	1,333***	1,334***	1,331***
	(0,083)	(0,084)	(0,081)	(0,127)	(0,127)	(0,127)
31-35 years	0,783***	0,797***	0,765***	1,053	1,054	1,052
	(0,066)	(0,068)	(0,064)	(0,107)	(0,108)	(0,107)
36-42 years	0,521***	0,529***	0,487***	0,747***	0,748***	0,752**
	(0,049)	(0,050)	(0,046)	(0,084)	(0,084)	(0,084)
43-50 years	0,281***	0,281***	0,250***	0,424***	0,424***	0,437***
	(0,030)	(0,030)	(0,027)	(0,054)	(0,054)	(0,056)
Ethnicity (Ref: White)						
African - Caribbean	0,368***	0,366***	0,375***	0,645***	0,646***	0,643***
	(0,033)	(0,033)	(0,034)	(0,060)	(0,060)	(0,060)
Indian, Pakistani, Bangladeshi	0,381***	0,377***	0,386***	0,820*	0,820*	0,817*

	(0,051)	(0,051)	(0,051)	(0,094)	(0,094)	(0,093)
Other Asian	1,104	1,093	1,102	0,653*	0,652*	0,661*
	(0,171)	(0,171)	(0,167)	(0,161)	(0,161)	(0,163)
Other	0,506***	0,506***	0,516***	0,776	0,775	0,776
	(0,076)	(0,076)	(0,076)	(0,148)	(0,148)	(0,148)
Education						
(Ref: ISCED 5-6-7)						
ISCED 3-4	1,018	1,020	1,011	0,963	0,963	0,966
	(0,038)	(0,039)	(0,037)	(0,039)	(0,039)	(0,039)
ISCED 0-1-2	0,777***	0,784***	0,771***	0,680***	0,680***	0,682***
	(0,045)	(0,046)	(0,044)	(0,053)	(0,053)	(0,053)
Father having a job	1,106**	1,107**	1,102**	1,077	1,077	1,078
	(0,054)	(0,055)	(0,053)	(0,068)	(0,068)	(0,068)
ρ	0,0969	0,102	0,0871	0,0691	0,0622	0,0621
χ2	1507	1503	1502	563	509	517
N	5689	5689	5689	3605	3605	3605
Persons-periods	84237	84237	84237	46638	46638	46638

Table 3.A2. Robustness check. Alternative version of Table 3.5.

Sample including people exiting all unions (also shorter than 12 months) and experiencing at least 1-month singlehood episodes

	Women		Men	
	Partner		Partner	
	Without children	With children	Without children	With children
	Model 4	Model 4	Model 5	Model 5
No children *(Ref=1)*				
Non-resident children only	1,003	1,142	1,046	0,837***
	(0,080)	(0,332)	(0,059)	(0,046)
Some co-resident children	0,573***	1,010	1,381***	1,235**
	(0,025)	(0,165)	(0,148)	(0,130)
Prev. marriage	1,022	1,113	0,907*	1,573***
	(0,047)	(0,201)	(0,045)	(0,175)
2+ previous unions	0,974	1,118	0,991	1,148
	(0,058)	(0,196)	(0,061)	(0,130)
Time since dissolution (Ref: 0-1 years)				
1-2 years	0,973	0,786	0,926	0,904
	(0,048)	(0,140)	(0,054)	(0,097)
2-4 years	1,030	0,867	0,957	0,929
	(0,055)	(0,164)	(0,061)	(0,112)
4-6 years	1,028	0,709	1,025	0,821
	(0,075)	(0,178)	(0,087)	(0,136)
6-10 years	0,943	0,805	0,840	0,631**
	(0,082)	(0,228)	(0,091)	(0,128)
Over 10 years	0,845	0,901	0,752*	0,732
	(0,100)	(0,342)	(0,111)	(0,198)
Duration of last union (Ref: 0,5-1 years)				
3-6 years	1,145***	1,183	1,002	1,054
	(0,058)	(0,230)	(0,056)	(0,131)
6-10 years	1,037	1,016	1,051	1,309**

	(0,059)	(0,214)	(0,070)	(0,178)
10+ years	1,332***	1,357	1,407***	2,040***
	(0,089)	(0,321)	(0,110)	(0,314)
Year of separation	0,999	0,995	0,997	0,994
	(0,002)	(0,002)	(0,004)	(0,004)
Living with biological parents at 16	1,071	1,074	1,015	1,509***
	(0,045)	(0,171)	(0,054)	(0,151)
Age				
23-26 years	1,138	1,259	1,171	1,581
	(0,095)	(0,455)	(0,151)	(0,669)
27-30 years	1,053	1,471	1,133	2,584**
	(0,091)	(0,531)	(0,144)	(1,050)
31-35 years	0,786***	1,153	0,925	2,439**
	(0,073)	(0,433)	(0,120)	(0,998)
36-42 years	0,538***	1,000	0,674***	2,545**
	(0,054)	(0,395)	(0,092)	(1,051)
43-50 years	0,289***	0,460*	0,410***	1,505
	(0,032)	(0,194)	(0,060)	(0,640)
Ethnicity (*Ref: White*)				
African – Caribbean	0,375***	0,195***	0,696***	0,436***
	(0,038)	(0,085)	(0,070)	(0,098)
Indian, Pakistani, Bangladeshi	0,314***	0,182**	0,898	0,334***
	(0,047)	(0,127)	(0,118)	(0,122)
Other Asian	0,965	0,141	0,733	0,119*
	(0,171)	(0,183)	(0,219)	(0,152)
Other	0,638***	0,000	1,062	0,351*
	(0,106)	(0,000)	(0,201)	(0,199)
Education (Ref: ISCED 5-6-7)				
ISCED 3-4	1,015	1,173	0,884***	1,141
	(0,041)	(0,171)	(0,040)	(0,106)
ISCED 0-1-2	0,815***	0,490***	0,631***	0,815
	(0,049)	(0,109)	(0,047)	(0,118)
Father having a job	1,057	1,169	1,052	1,435***
	(0,053)	(0,226)	(0,069)	(0,177)
Var	0,358***	2,670**	0,240***	0,711***

	(0,082)	(1,200)	(0,075)	(0,237)
Covariance	0,298*	0,298*	0,165	0,165
	(0,157)	(0,157)	(0,120)	(0,120)
N	5689	5689	3605	3605
Persons-periods	84237	84237	46638	46638

Chapter 4

Union fertility and stability in different family settings
A multiprocess analysis of Britain

Abstract

I study the risk of a birth and the risk of separation in different union settings, such as step-families and families with no prior children. I test whether the transition to a child is determined by (a1) the 'parenthood motive' (the birth risk is greater if the child is the first biological child for at least one partner), or (a2) the 'commitment effect' (the risk of a birth is higher if it is the first shared birth). I also test whether (b1) having common children stabilizes a partnership and (b2) the influence of pre-union children increases the risk of union dissolution. Some elements of family complexity, such as the presence, the number and the parentage of step-children, are taken into account. Using multilevel multiprocess models with simultaneous equations, I model partnership transitions jointly with fertility, allowing for the correlation between the unobserved individual-level characteristics that affect each process. The analysis is based on the partnership and birth histories of the Wave 1 of UKHLS (Understanding Society) of men and women aged 16-45. The findings indicate that both the parenthood and the commitment motives influence the transitions to a birth, under different family configurations. Further, the risk of separation is reduced by the presence of shared children, while the existence of 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 4.1. 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 4.2. 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 4.3. 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 4.4.

Table 4.4. 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

$$h_{ijt}^{(U)} = \alpha^{(U)} D_{ijt}^{(U)} + \beta^{(U)} F_{ijt}^{(U)} + \gamma X_{ijt}^{(U)} + \delta Z_{i}^{(U)} + u_{j}^{(U)} \quad (1)$$

$$h_{ijt}^{(F)} = \alpha^{(F)} D_{ijt}^{(F)} + \beta^{(F)} \left\{ U_{ijt}^{(F)}, D_{ijt}^{(F)} \right\} + \gamma X_{ijt}^{(F)} + \delta Z_{i}^{(F)} + u_{j}^{(F)} \quad (2)$$

$$h_{ijt}^{(D)} = \alpha^{(D)} D_{ijt}^{(D)} + \beta^{(D)} F_{ijt}^{(D)} + \gamma X_{ijt}^{(D)} + \delta Z_{i}^{(D)} + u_{j}^{(D)} \quad (3)$$

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 $\boldsymbol{t}$ 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 $F_{ijt}^{(p)}$, 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. $X_{ijt}^{(p)}$ 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. $\delta Z_i^{(p)}$ 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 4.5 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 4.5. 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 4.6. 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 4.6. 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 4.1 and 4.2 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 4.1 and 4.2 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 4.1 and 4.2, 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 4.1. Fertility process: risk of a birth. Estimated odds ratios by family configuration. Women.

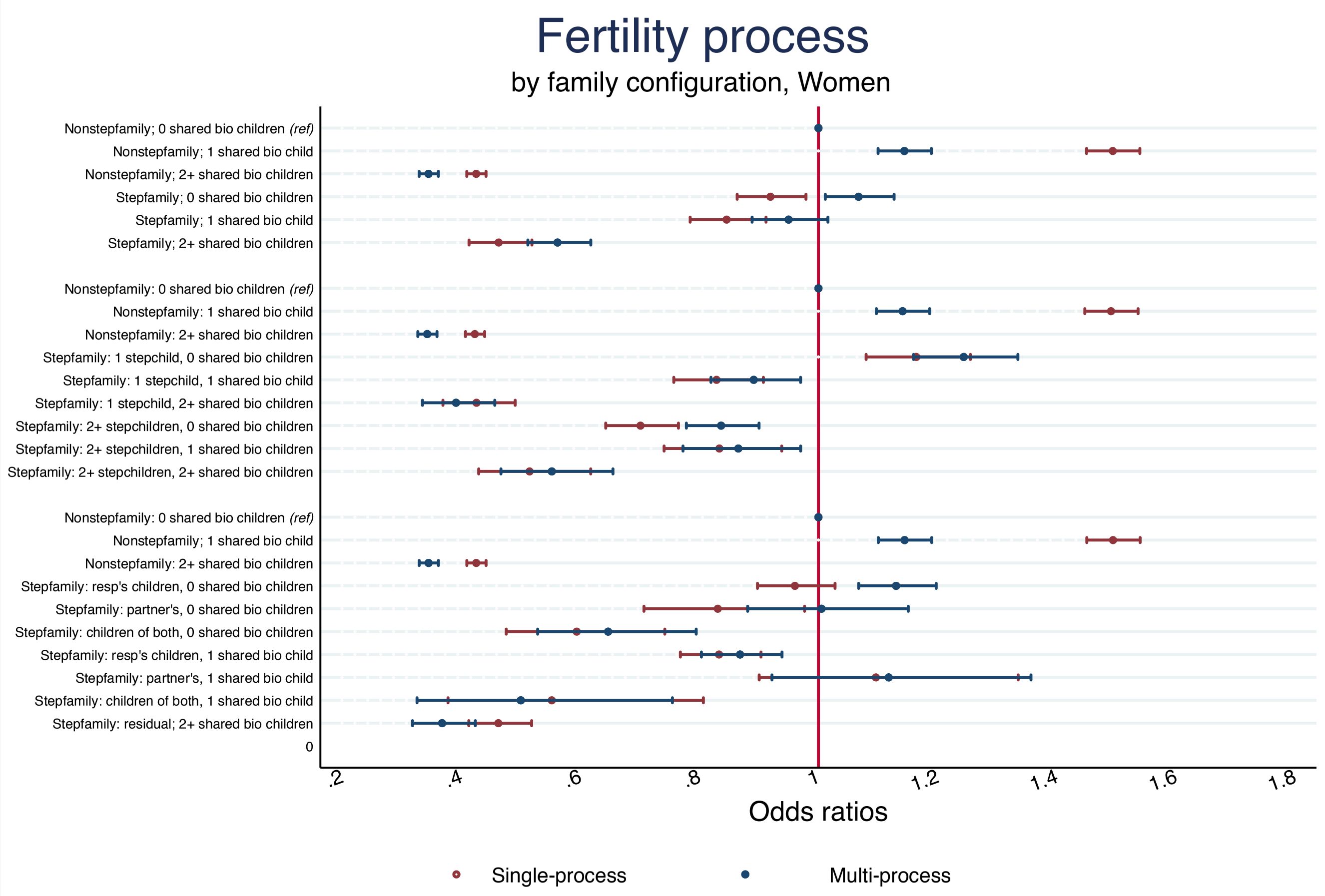
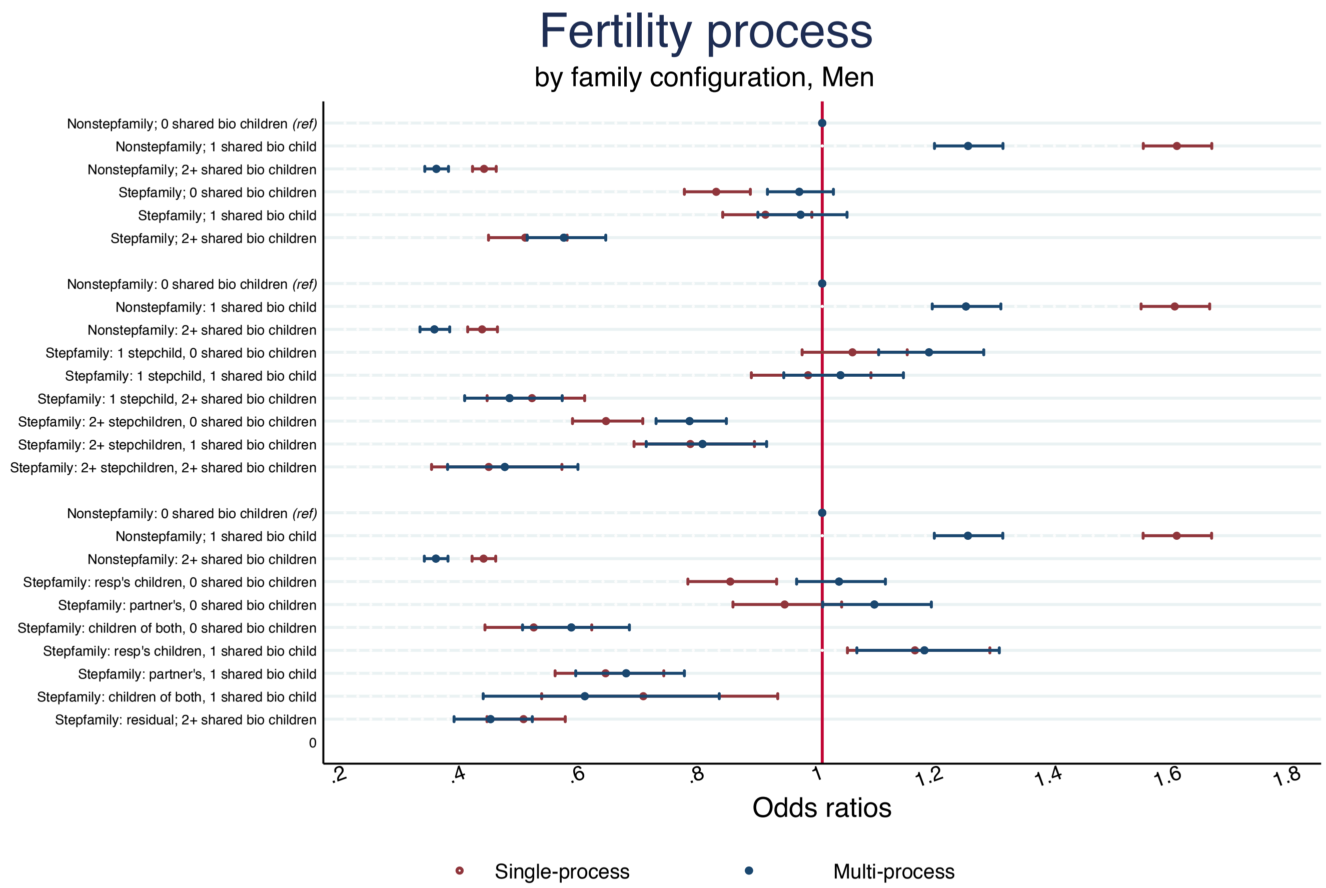


Figure 4.2 Fertility process: risk of a birth. Estimated odds ratios by family configuration. Men.



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 Fig. 4.3 and 4.4 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 4.3. Fertility process: risk of a birth. Estimated odds ratios by family configuration. Women.

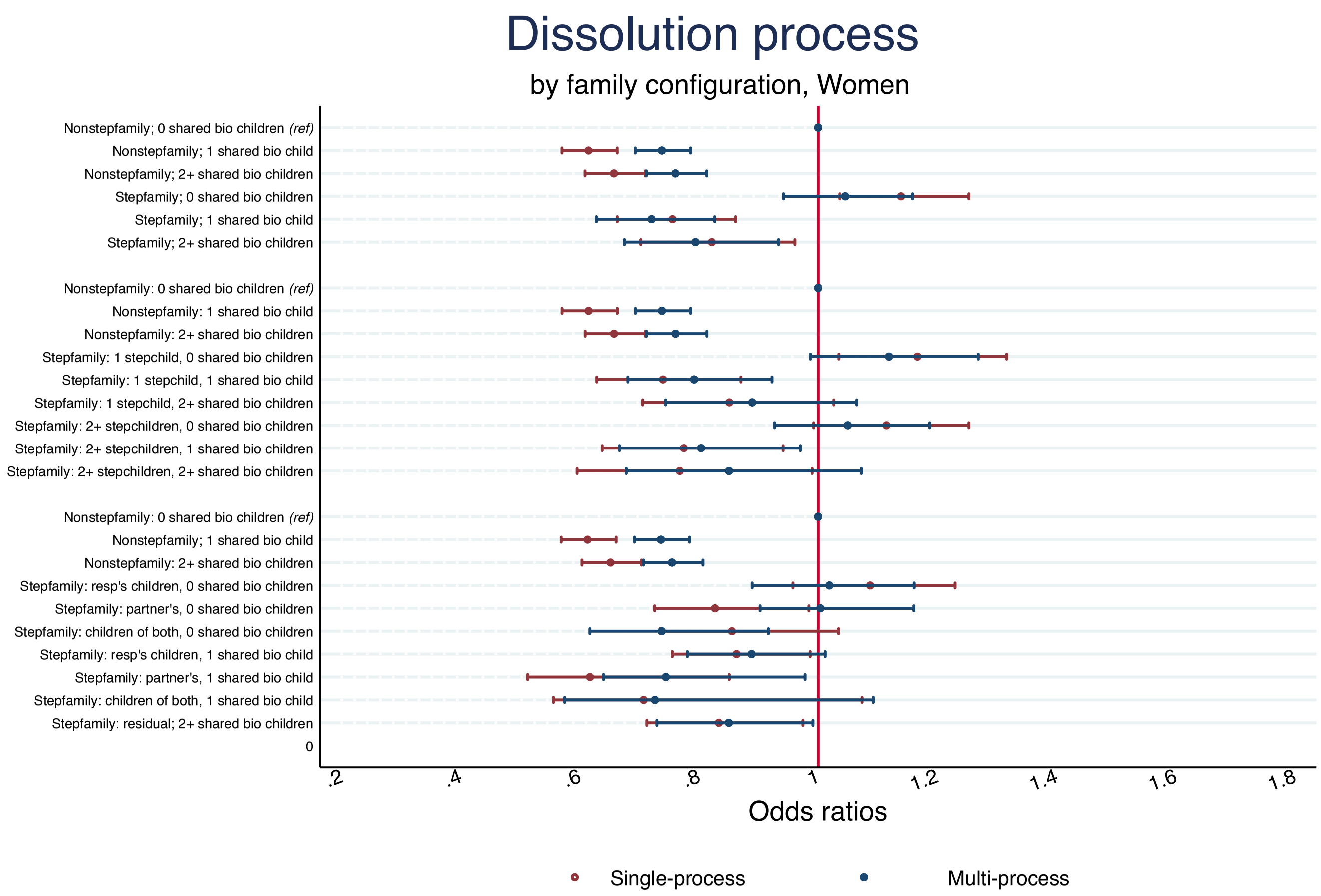
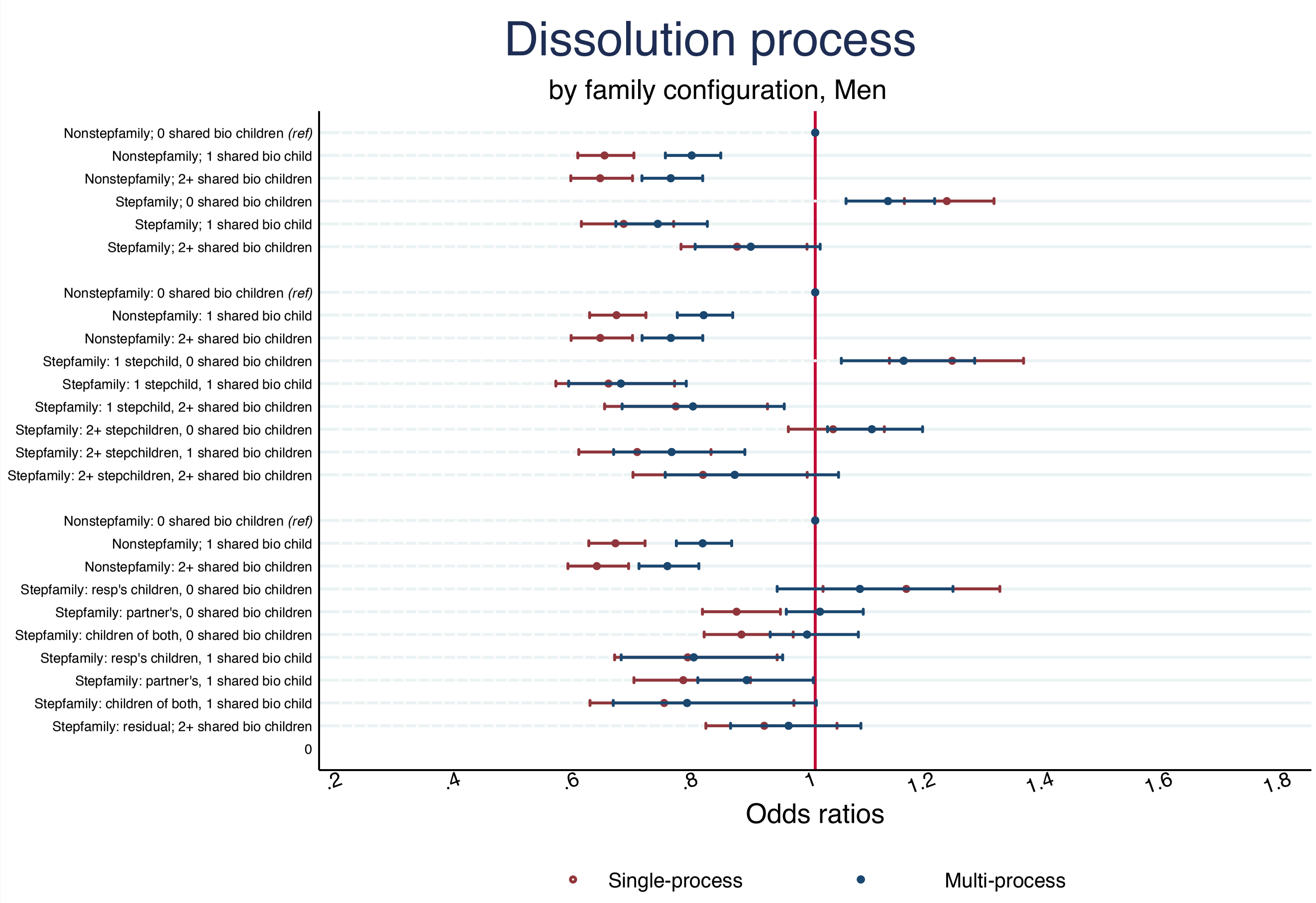


Figure 4.4. Fertility process: risk of a birth. Estimated odds ratios by family configuration. Men.



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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Appendix

Table 4.A.1. Estimated coefficients and standard errors from models of the transition from singlehood to union. Women and men.

	Women		Men	
	Single-process model	Multiprocess model	Single-process model	Multiprocess model
No children *(Ref=1)*				
Co-resident children; 1 child	0,987	0,965	1,694***	1,599***
	(0,036)	(0,039)	(0,096)	(0,098)
Co-resident children; 2+ children	0,643***	0,618***	0,948	1,082
	(0,029)	(0,031)	(0,072)	(0,083)
Non-resident children; 1 child	1,360***	1,373***	1,134*	1,179**
	(0,161)	(0,172)	(0,077)	(0,081)
Non-resident children; 2+ children	1,332**	1,348**	1,370***	1,323***
	(0,149)	(0,159)	(0,101)	(0,103)
Prev. marriage	0,528***	0,537***	0,477***	0,477***
	(0,024)	(0,028)	(0,026)	(0,026)
Number of previous unions (*Ref: none*)				
1 previous union	0,317***	0,311***	0,427***	0,432***
	(0,026)	(0,028)	(0,042)	(0,044)
2+ previous unions	0,280***	0,265***	0,326***	0,352***
	(0,032)	(0,035)	(0,044)	(0,045)
Duration (continuous)	0,993***	0,994***	0,996***	0,996***
	(0,001)	(0,001)	(0,001)	(0,001)
Duration (squared)	1,000***	1,000***	1,000***	1,000***
	(0,000)	(0,000)	(0,000)	(0,000)
Current age (continuous)	1,627***	1,648***	1,864***	1,876***
	(0,030)	(0,035)	(0,038)	(0,037)
Current age (squared)	0,993***	0,994***	0,991***	0,990***
	(0,000)	(0,000)	(0,000)	(0,000)
Cohort of birth (*Ref: 1940-1944)*				
Cohort: 1945-49	0,931**	0,950	0,969	0,991
	(0,033)	(0,037)	(0,043)	(0,052)
Cohort: 1950-54	0,877***	0,852***	0,838***	0,821***
	(0,033)	(0,036)	(0,040)	(0,042)

Cohort: 1955-59	0,730***	0,745***	0,681***	0,696***
	(0,028)	(0,028)	(0,032)	(0,033)
Cohort: 1960-64	0,595***	0,603***	0,626***	0,654***
	(0,022)	(0,023)	(0,028)	(0,032)
Cohort: 1965-69	0,585***	0,597***	0,608***	0,619***
	(0,022)	(0,024)	(0,028)	(0,031)
Cohort: 1970-74	0,583***	0,592***	0,609***	0,613***
	(0,023)	(0,025)	(0,028)	(0,029)
Cohort: 1975-79	0,626***	0,633***	0,634***	0,648***
	(0,025)	(0,027)	(0,031)	(0,033)
Parents separated	1,019	1,040	1,184***	1,121***
	(0,052)	(0,056)	(0,077)	(0,077)
Parents' education (*Ref: Degree)*				
Missing	1,009	1,013	1,073	1,062
	(0,041)	(0,043)	(0,050)	(0,053)
No school	1,169**	1,156**	1,435***	1,421***
	(0,090)	(0,095)	(0,121)	(0,127)
Left with no qualification	1,195***	1,202***	1,235***	1,252***
	(0,047)	(0,049)	(0,059)	(0,060)
Left with some qualification	1,158***	1,172***	1,081*	1,195*
	(0,046)	(0,049)	(0,051)	(0,055)
Some qualification	1,158***	1,163***	1,159***	1,162***
	(0,045)	(0,046)	(0,055)	(0,056)
Ethnicity (*Ref: White*)				
African - Caribbean	0,551***	0,579***	0,680***	0,693***
	(0,023)	(0,024)	(0,036)	(0,033)
Indian, Pakistani, Bangladeshi	1,652***	1,660***	1,491***	1,431***
	(0,055)	(0,057)	(0,050)	(0,052)
Other Asian	0,867**	0,871**	0,757***	0,745***
	(0,052)	(0,052)	(0,055)	(0,052)
Other	0,636***	0,642***	0,493***	0,502***
	(0,053)	(0,055)	(0,046)	(0,048)
Education (*Ref: Degree*)				
No qualification	1,129***	1,117***	0,839***	0,832***
	(0,044)	(0,045)	(0,036)	(0,037)
Other qual	1,164***	1,121***	1,005	1,028
	(0,047)	(0,047)	(0,044)	(0,045)
GCSE etc	1,024	1,054	0,945	0,962

	(0,035)	(0,042)	(0,036)	(0,037)
A level etc	1,140***	1,153***	1,039	1,062
	(0,037)	(0,039)	(0,036)	(0,040)
Other higher	1,077**	1,071*	1,083*	1,092*
	(0,040)	(0,042)	(0,048)	(0,049)
Respondent being in education	0,326***	0,352***	0,516***	0,504***
	(0,011)	(0,010)	(0,022)	(0,026)

Table 4.A.2. Estimated coefficients and standard errors from models of the transition to a birth. Women and men.

	Women		Men	
	Single-process model	Multiprocess model	Single-process model	Multiprocess model
Nonstepfamily; 0 shared bio children *(Ref=1)*				
Nonstepfamily; 1 shared bio child	1,493***	1,137***	1,595***	1,232***
	(0,023)	(0,027)	(0,029)	(0,032)
Nonstepfamily; 2+ shared bio children	0,425***	0,357***	0,433***	0,389***
	(0,011)	(0,015)	(0,013)	(0,026)
Stepfamily; 0 shared bio children	0,919**	1,096**	0,822***	0,952
	(0,030)	(0,031)	(0,029)	(0,033)
Stepfamily; 1 shared bio child	0,845***	0,938*	0,905**	0,964
	(0,036)	(0,037)	(0,042)	(0,049)
Stepfamily; 2+ shared bio children	0,463***	0,572***	0,502***	0,586***
	(0,032)	(0,040)	(0,039)	(0,047)
Prev. marriage	0,875***	0,875***	1,102**	1,087**
	(0,033)	(0,033)	(0,050)	(0,055)
Number of previous unions (*Ref: none*)				
1 previous union	0,967	0,982	0,861***	0,882***
	(0,028)	(0,032)	(0,030)	(0,032)
2+ previous unions	0,982	0,985	0,869***	0,896***
	(0,055)	(0,057)	(0,047)	(0,049)
Duration (continuous)	1,017***	1,018***	1,018***	1,018***
	(0,001)	(0,001)	(0,001)	(0,001)
Duration (squared)	1,000***	1,000***	1,000***	1,000***
	(0,000)	(0,000)	(0,000)	(0,000)
Current age (continuous)	1,218***	1,217***	1,159***	1,159***
	(0,011)	(0,011)	(0,014)	(0,014)
Current age (squared)	0,996***	0,997***	0,997***	0,997***
	(0,000)	(0,000)	(0,000)	(0,000)
Cohort of birth (*Ref: 1940-45*)				

Cohort: 1945-49	0,881***	0,884***	0,945*	0,950*
	(0,023)	(0,025)	(0,028)	(0,029)
Cohort: 1950-54	0,833***	0,841***	0,879***	0,881***
	(0,022)	(0,023)	(0,026)	(0,027)
Cohort: 1955-59	0,895***	0,887***	0,928**	0,932**
	(0,024)	(0,024)	(0,028)	(0,028)
Cohort: 1960-64	0,888***	0,889***	0,912***	0,915***
	(0,023)	(0,024)	(0,027)	(0,028)
Cohort: 1965-69	0,862***	0,869***	0,843***	0,852***
	(0,022)	(0,023)	(0,025)	(0,027)
Cohort: 1970-74	0,814***	0,819***	0,830***	0,834***
	(0,022)	(0,022)	(0,026)	(0,027)
Cohort: 1975-79	0,746***	0,750***	0,700***	0,707***
	(0,023)	(0,025)	(0,027)	(0,026)
Parents separated	1,033	1,058	0,985	0,992
	(0,034)	(0,036)	(0,044)	(0,048)
Parents' education (*Ref: Degree)*				
Missing	1,003	1,005	1,021	1,017
	(0,027)	(0,027)	(0,033)	(0,032)
No school	1,199***	1,206***	1,267***	1,283***
	(0,063)	(0,064)	(0,080)	(0,082)
Left with no qualification	0,985	0,991	1,030	1,032
	(0,026)	(0,022)	(0,033)	(0,034)
Left with some qualification	0,985	0,989	1,029	1,035
	(0,026)	(0,029)	(0,033)	(0,036)
Some qualification	0,978	0,984	1,000	1,001
	(0,026)	(0,023)	(0,032)	(0,035)
Ethnicity (*Ref: White*)				
African - Caribbean	1,244***	1,247***	1,339***	1,342***
	(0,043)	(0,045)	(0,058)	(0,059)
Indian, Pakistani, Bangladeshi	1,479***	1,484***	1,580***	1,587***
	(0,034)	(0,036)	(0,038)	(0,041)
Other Asian	1,066	1,080	1,170***	1,175***
	(0,055)	(0,057)	(0,064)	(0,066)
Other	1,239***	1,249***	1,386***	1,393***
	(0,065)	(0,068)	(0,083)	(0,086)

Education (*Ref: Degree*)				
No qualification	1,202***	1,195***	1,061**	1,089**
	(0,030)	(0,032)	(0,030)	(0,032)
Other qualification	1,043*	1,059*	1,005	1,026
	(0,026)	(0,026)	(0,026)	(0,029)
GCSE etc	1,042**	1,046**	0,989	0,987
	(0,022)	(0,024)	(0,024)	(0,025)
A level etc	1,004	1,014	0,955**	0,968
	(0,021)	(0,025)	(0,022)	(0,026)
Other higher	1,059***	1,066***	0,966	0,975
	(0,023)	(0,024)	(0,026)	(0,027)
Respondent being in education	0,634***	0,627***	0,740***	0,733***
	(0,047)	(0,049)	(0,047)	(0,048)

Table 4.A.3. Estimated coefficients and standard errors from models of the transition from union to singlehood. Women and men.

	Women		Men	
	Single-process model	Multiprocess model	Single-process model	Multiprocess model
Nonstepfamily; 0 shared bio children *(Ref=1)*				
Nonstepfamily; 1 shared bio child	0,613***	0,683***	0,667***	0,814***
	(0,027)	(0,030)	(0,030)	(0,029)
Nonstepfamily; 2+ shared bio children	0,671***	0,791***	0,614***	0,772***
	(0,029)	(0,031)	(0,031)	(0,029)
Stepfamily; 0 shared bio children	1,189***	1,079***	1,286***	1,122***
	(0,052)	(0,052)	(0,039)	(0,054)
Stepfamily; 1 shared bio child	0,711**	0,683**	0,651***	0,772**
	(0,059)	(0,063)	(0,051)	(0,059)
Stepfamily; 2+ shared bio children	0,828*	0,809*	0,889**	0,831
	(0,081)	(0,086)	(0,079)	(0,085)
Prev. marriage	0,783***	0,775***	0,725***	0,774***
	(0,045)	(0,048)	(0,049)	(0,047)
Number of previous unions (*Ref: none*)				
1 previous union	1,395***	1,302***	1,615***	1,585***
	(0,076)	(0,075)	(0,088)	(0,089)
2+ previous unions	1,664***	1,652***	2,115***	2,078***
	(0,155)	(0,152)	(0,167)	(0,162)
Duration (continuous)	1,000	1,000	1,000	1,000
	(0,001)	(0,001)	(0,001)	(0,001)
Duration (squared)	1,000***	1,000***	1,000	1,000*
	(0,000)	(0,000)	(0,000)	(0,000)
Current age (continuous)	0,993	0,994	0,883***	0,884***
	(0,013)	(0,014)	(0,015)	(0,015)

Current age (squared)	1,000	1,000	1,001***	1,000***
	(0,000)	(0,000)	(0,000)	(0,000)
Cohort of birth (*Ref:1940-44*)				
Cohort: 1945-49	1,177***	1,169***	1,096	1,109*
	(0,072)	(0,073)	(0,079)	(0,082)
Cohort: 1950-54	1,437***	1,429***	1,335***	1,297***
	(0,086)	(0,088)	(0,097)	(0,093)
Cohort: 1955-59	1,621***	1,623***	1,466***	1,432***
	(0,097)	(0,102)	(0,104)	(0,107)
Cohort: 1960-64	1,901***	1,934***	1,512***	1,517***
	(0,115)	(0,119)	(0,110)	(0,118)
Cohort: 1965-69	2,008***	2,075***	1,495***	1,501***
	(0,123)	(0,129)	(0,110)	(0,117)
Cohort: 1970-74	2,178***	2,203***	1,454***	1,473***
	(0,139)	(0,140)	(0,113)	(0,118)
Cohort: 1975-79	2,417***	2,442***	1,497***	1,469***
	(0,162)	(0,167)	(0,125)	(0,128)
Parents separated	1,326***	1,386***	1,272***	1,303***
	(0,069)	(0,073)	(0,089)	(0,093)
Parents' education (*Ref: Degree)*				
Missing	0,902*	0,926	0,913	0,934
	(0,048)	(0,052)	(0,057)	(0,058)
No school	0,872	0,853*	0,685**	0,664**
	(0,109)	(0,109)	(0,112)	(0,117)
Left with no qual	0,865***	0,832***	0,766***	0,748***
	(0,046)	(0,047)	(0,049)	(0,050)
Left with some qual	0,876**	0,863***	0,796***	0,775***
	(0,047)	(0,049)	(0,052)	(0,054)
Some qualification	0,926	0,937	0,912	0,956
	(0,048)	(0,048)	(0,056)	(0,057)
Ethnicity (*Ref: White*)				
African - Caribbean	1,315***	1,334***	1,250***	1,284***
	(0,066)	(0,068)	(0,084)	(0,089)
Indian, Pakistani, Bangladeshi	0,413***	0,435***	0,413***	0,428***
	(0,030)	(0,032)	(0,037)	(0,040)

Other Asian	0,666***	0,689***	0,424***	0,441***
	(0,077)	(0,079)	(0,099)	(0,103)
Other	0,983	0,991	1,037	1,032
	(0,112)	(0,123)	(0,127)	(0,128)
Education (*Ref: Degree*)				
No qualification	1,025	1,042	1,057	1,057
	(0,052)	(0,054)	(0,065)	(0,066)
Other qual	1,049	1,061	1,098	1,112*
	(0,055)	(0,059)	(0,064)	(0,065)
GCSE etc	1,071	1,075	1,200***	1,195***
	(0,046)	(0,043)	(0,061)	(0,062)
A level etc	1,141***	1,154***	1,139***	1,144***
	(0,050)	(0,055)	(0,055)	(0,056)
Other higher	0,980	0,923	1,086	1,045
	(0,047)	(0,049)	(0,062)	(0,068)
Respondent being in education	1,254***	1,243***	1,311***	1,305***
	(0,082)	(0,083)	(0,104)	(0,104)

Chapter 5

Conclusions

Each chapter of the thesis addressed crucial transitions in the life course highlighting the diversity in individuals' behaviour. Chapter 2 analysed the role of parents' background on the transition to adulthood, in terms of entry into first union and first parenthood. One of the key findings to emerge from the research illustrated in Chapter 2 is the importance of family background characteristics for explaining the risk of first cohabitation vs. first marriage, and the risk of a marital vs. a non-marital union. Interestingly, the determinants of each transition are not always significant and comparable in magnitude, and the two dimensions of parents' background do seem to play different roles on the outcomes.

To be sure, the timing and the type of entry into first union and parenthood is influenced by a myriad of other explanatory variables, such as family structure, ethnicity, culture and personal beliefs, young adults' social and economic conditions, historical conditions (e.g., Barber, 2001; Oppenheimer, 1988; South, 2001). However, I find – such as most studies in these areas – that the influence of family background persist even after other established correlates of transition outcomes are accounted for. Despite the bulk of research on transition to adulthood, few studies have considered the choice in family arrangement to have a first union or a first child as a dependent variable. Also, very few studies have considered that possibility that family background characteristics associated with transitions are contingent on either historical time or individuals' age.

Higher parents' education is associated with postponed entry into first marriage and first marital birth compared with the lower educated and lower-class parents. Higher social class is negatively linked with the risk of marriage and with the risk of non-marital birth. This result is coherent with previous research on the timing of first marriage (Axinn & Thornton, 1992; Mooyaart & Liefbroer, 2016; South, 2001), while evidence on family type of first parenthood is little and rather sparse. The influence of parents' education is stronger for first marriage than for first cohabitation. Education is a proxy of parents' social capital and aspirations. Higher educated parents should prevent children from incurring in abrupt transitions with enduring consequence. For instance, the decision to marry has an impact on a longer horizon, and higher educated parents tend discourage a rapid transition to marriage than an early cohabitation, which appears more remediable. Social class is a proxy of family resources. On the one hand, children from higher social backgrounds have higher occupational aspirations and better information on how to achieve these goals. On the other hand, they are more attractive in the marriage market and be more prone to early transitions. The findings definitely privilege the idea that higher class background predicts longer occupational career progressions, slower transition to marriage, slower or forgone transition to non-marital birth.

I used a long historical time window with cohorts born in 1930s through 1980, thus entering first union and parenthood in very different cultural and socio-economic circumstances. The results depict the influence of parents with low education decay significantly over the time of analysis with respect to marriage, in line with other studies (Sassler & Goldscheider, 2004b; South, 2001; Wiik, 2009). Conversely, the influence in low educated parents on cohabitation have increased over time. Billari & Liefbroer (2010) showed that the higher educated have developed a new set of priorities such as a period of independent living, unmarried cohabitation and the postponement of childbearing. In earlier years, when rates of cohabitation were low across different socio-economic groups, the influence of socio-economic background on marriage timing was relatively strong. Nonmarital cohabitation has dramatically increased for all demographic subgroups and has become increasingly common especially among the lower educated in Europe (Perelli-Harris et al., 2010). Children from lower socioeconomic status have possibly substituted nonmarital cohabitation for marriage later than children from higher background – and the impact of family socio-economic background on the formation of first marriages has supposedly become weaker over time. As South (2001) suggest, a historical change in the effect of family background characteristics on the transition to cohabitation is a possible mechanism that explains the relative diminution of the impact of family background on marriage timing. These attenuated effects are consistent with theories claiming that individuals' responses to structured social pressured – partly personified by parents – have become less responsive over time (Bumpass, 1990).

Chapter 2 has explored the empirical implications, among the others, of the 'individuation theory'. By 'individuation', scholars mean that the post-war demographic changes – such as the baby-boom fertility declines, the increase in nonmarital fertility ratios, and the rises in nonmarital cohabitation and divorce – must be explained not only in terms of structural forces but also as a by-product of the increasing emphasis on individual freedom (Lesthaeghe, 1983; South, 2001). For instance, in the context of the first union patterns in United Kingdom, the rising age at marriage and the diffusion of alternatives can be motivated also by a shift toward marriage as an institution (Murphy, 2000b). Also, the marked increase in cohabitation unions is considered as evidence of the individuation theories (Lesthaeghe, 1983): for instance, in Western societies, individuals embracing more liberal and secularised views tend to place less positive attitudes towards marriage (Lesthaeghe & Surkyn, 1988). One implication of the individuation thesis is that, over time, individuals' responses to group pressures – including parents' – tend to wane. In other words, demographic behaviour has increasingly escaped the influence of institutional arrangements and norms and has become more dependent on individual discretion (Lesthaeghe & Surkyn, 1988). Over time, the impact of family socio-economic background has weakened during the second half of the 20th century. It cannot be neglected that over recent decades the impact of parents' education on their children's educational attainment have waned as well (Blossfeld & Huinink, 1991). Accordingly, historical changes in the intergenerational transmission of education might also explain a decrease of the influence of parents' education on timing marriage and marital birth.

The association of family background on the transitions to first marriage and first marital birth varies over their children's life course. The influence of family resources on the transition to marriage wanes as adolescents and young adults age, in line with previous research (South 2001; Wiik, 2009; Mooyaart & Liefborer, 2016). Also, parents' higher resources are associated with postponement of marital birth at early stages but this links weakens at older ages especially for lower classes. This is coherent with the life-course perspective that claims the existence of age-specific explanatory factors on demographic transitions (Hogan & Astone, 1986). Family socio-economic background exerts its stronger influence on young adults, who have not gained their economic independence yet and, in many cases, have not put enough geographical distance from their origin families yet. The effects of some family background characteristics do weaken over time with respect to some transition but not with the respect to others. For instance, the vanishing role of parents' socio-economic role on children's age at first marriage can be motivated by the increasing number of young adults still at risk of first marriage in their late twenties and throughout their thirties (Berrington & Diamond, 2000; O'Leary et al., 2010). Most of young adults might have reached economic independence by that ages, as the influence of their parents' economic resources contextually subside.

This area of research could benefit from further developments. Future research on family origin and life-course transition should aim to disentangle the process of psychological maturation of children from the influence of parents. For instance, the 'individuation' thesis, which claims that many demographic changes are explained by ideal shifts toward greater individualism, does not identify conclusive explanations of the mechanisms linking the cultural norms to key demographic events. Future research might explore more accurately the channels of transmission from parents' background to children's transitions using information on children's values and beliefs, which may act as moderators of this pathway. To this end, longitudinal - and not retrospective – information about individuals are needed.

Finally, international comparisons are insightful to establish difference between countries in the influence of parents' background on transition to adulthood. As Mooyaart & Liefbroer (2016) note, individuals have less difficulties to settle down into long-term family arrangements, such as marriage and marriage with children, in countries with higher welfare expenditure. In many Southern European countries, where a large amount of public resources is devoted to elderly people (Thévenon, 2011), the reliance on family resources is greater, especially among the higher classes, the de-standardization of trajectories is limited and the stay in parental home is more persistent (Sironi, Barban, & Impicciatore, 2015). Wide and consistent cross-country findings on young adults' trajectories might thus serve to inform policy-makers about the relationship about social class and new generations' patterns of social mobility. The last crisis and the consequent diffusion of 'short-term' job contracts have hit hard on the new generations and intensified the linkage between young adults and parents' resources. The probable outcome is the decline in the possibilities of self-determination and social mobility. New research investigating these patterns across social classes is thus urgent.

Chapter 3 contributed to an ongoing discussion on the role of children in the union formation process and on the gender differences – and similarities – in the parent/partner roles. There has been a dramatic increase in the number of unions (marriages and cohabitations) and a relevant surge in the number of children living with only one parent. These phenomena require more understanding on how the prospect of sharing the household with children affects the union formation behaviour of their co-resident parent and his/her potential partner. Likewise, the children who would not co-reside with the newly formed couple, but remain financially and emotionally dependent on their parent, might be influential.

The results show that despite the growing ‘normalization’ of post-dissolution singlehood and the increasing supply of partners who have already children, there is clear evidence of diverging chances of repartnering between the childless and the parents, and between mothers and fathers. Co-resident parents repartner at a lower rate than childless individuals do. The custodial responsibilities can reduce parents’ time and willingness to meet potential partners and form new relationships. When these constraints do not exist – namely for non-custodial parents – repartnering proceeds at a faster pace as opposed to co-resident parents. Although non-resident mothers and fathers are a selected sample in the pool of parents, their propensity to find a partner more rapidly suggests that it is not parenthood *per se* to slow down the process of repartnering. Their proactivity on the repartner market could be motivated by the desire to compensate for children’s absence along with the partial exemption from the burden of parental duties.

The negative association between parental childcare burden and new union prospects for women partially applies to men. Fathers’ repartnering chances are higher than mothers’ and are not responsive to children’s number and only marginally to the age of the youngest child. Assuming that childrearing constraints are the same for single mothers and fathers, such a gap is supposedly explained (also) by the combination of custodial fathers’ relatively higher attractiveness and more proactive partner search. Previous studies have claimed that this gender gap has further motivations: for instance, women may be more susceptible to the demise of a relationship (e.g., Poortman, 2007) and tend to disinvest from romantic life after a break-up (Beaujouan, 2012); also, women’s partnership formation largely decreases at the beginning of their early 30s, relative to men (Beaujouan, 2012) and is severely influenced by their socio-economic status (Ivanova & Begall, 2015; Shafer & James, 2013). The variability of repartnering habits by age and socio-economic background was not the research goal of Chapter 3 but it is in my research agenda for the next years.

The second part of the analysis highlights that custodial responsibilities could also shape the type of new union. Custodial parents have lower chances of repartnering with childless individuals and with parents and mothers are systematically less likely to repartner with childless partners than are fathers. I justify these results with the idea that custodial parenthood is not attractive in the repartnering market: childless people are supposedly less willing to embrace step-parenthood, a challenging trial when it comes to role definition within the couple and relationship build-up with non-biological kids. In contrast to Vanassche et al., (2015), I do not find evidence that parents are more likely to start a union with another parent than with a childless partner. A plausible justification may reside in the greater prevalence of potential childless partners. The same motivation can explain men's higher likelihood probability of repartnering with custodial parents compared to women's. As hypothesised by Bernhardt & Goldscheider (2002), the higher availability of women with dependent children could increase the opportunity for men to enter a union with partners who live with children from previous unions.

A caveat. All these findings should not be interpreted in terms of causal inference. Custody arrangements are not randomly distributed among separated mothers and fathers. Recent evidence showed that unobserved factors make some individuals more likely than others to have sole physical custody and form a new union Schnor et al. (2017). I argued that non-resident parents are those who relinquish their parental prerogatives, as they anticipate that the chances to find a partner being a custodial parent are lower. Conversely, it is also possible the co-resident parents are disproportionally selected among those with positive attitudes toward repartnering: they might have above-average inclination for living in a family, which makes them more prone to accept the first suitable candidate to have a parental figure for their children. Future studies should address better information on custody assignment exploiting, for instance, reports of court orders or, at least, by clarifying whether the decision of sole custody was bargained between the ex-partners or imposed via court order. Ultimately, difference-in-difference designs benefitting from legislations reforms or instrumental variable approaches – like in Schnor et al. (2017) – represent the golden standard to circumvent self-selection bias and study the repartnering process from a causal point of view.

In future studies, the focus on two overlooked categories – the fathers and the non-resident children –needs much more theoretical reflection and empirical development. The decline in the likelihood that men live with children is a well-documented phenomenon in the United Kingdom (Henz, 2014; Kiernan, 2006b). Recent studies have started to depict the characteristics of non-resident fathers also in the United Kingdom (Haux et al., 2015), although there is very limited evidence on the details of father-child contact after a separation in the European countries. The nature of non-resident father-children could be more predictive of fathers' post dissolution chances of new partnership than the residence status (Stewart, 2003). Nevertheless, paternal engagement with childcare – in terms of visitation schedule, active participation in the daily care, economic support – is still relatively understudied topic, due to the large non-response rates in dedicated surveys (e.g., Millennium Cohort Study, see Aassve & Pronzato, 2017).

On the policy side, the results suggest that promoting men's co-residential responsibilities for children is not as detrimental for their repartnering probabilities as it is for women. Increasing the number of parents with active custodial responsibilities is probably good for their own personal aspirations as long as the rule of custody sharing are clear enough and adults are responsible enough to manage the conflict that emerge during coparenting. On the one hand, it seems plausible to conclude that men's involvement with children does not depress repartnering and that men with coresident children are not scared away by stepparenthood duties. On the other hand, the custody burden mothers bear seems determinant to discourage them from repartnering. The load of care and daily tasks appear the most convincing reason to justify their relatively smaller chances of new union formation. As I signaled in Chapter 3, more determined gender-equal policies in post-separation children arrangements, such as abandoning the old-fashioned dichotomy between a 'physically custodial parent' and 'visiting parent' to adopt more egalitarian arrangements, would benefit ex-partners well-being and, thus, chances of repartnering.

Studies that consider the formation and dissolution of second- and higher- order unions are rare, with noteworthy exceptions (e.g., Aassve et al., 2006; Steele et al., 2005; Steele et al., 2006). Studies that consider stepfamilies within the framework of life-course theory are even more rare. Most research on formation and outcomes of co-residential unions in the United Kingdom typically considers women and in first unions, or first marriage, or in first cohabitation, or after the first dissolutions. The greatest merit of the aforementioned studies was to incorporate separations (from marital or cohabiting unions) and new partnerships in a dynamic context was inevitable in the United Kindgom at the turn of the century. The greater diversity among young adults in their trajectories and sequencing of union formation and childbearing (Berrington, 2003) has then led many other scholars to focus on 'non-normative' transitions in the last years: in fertility (e.g. Berrington, Stone and Beaujouan, 2015; Ní Bhrolcháin and Beaujouan 2012); in partnership behaviour (e.g. Berrington, Perelli-Harris and Trevena, 2015; Ní Bhrolcháin and Beaujouan 2013), in parenthood (Berrington, 2014). In the final Chapter of this thesis, I tried to address the complexity of partnerships and childbearing dynamics in the holistic view of the life course, thus expanding the traditional focus on punctual transitions.

The reflection about stepfamilies has been generally marginalized in British research although noteworthy studies have recently outlined the context (Ermisch & Francesconi, 2000; O' Leary et al., 2010; Galezewska et al., 2014) and investigated the topic of fertility intentions in complex families (Hohmann-Marriott, 2015). In addition to innovating the literature of the United Kingdom, I also set out to "bring men back" into the study of stepfamilies and family transitions in general, as argued by Goldscheider and Kaufman (1996). In spite of the bulk of research on stepfamilies in Europe, most studies have been based on women's retrospective reproductive and conjugal histories, with notable exceptions (Ivanova et al., 2013; Thomson & Li, 2002). One of the main challenges of investigating stepfamilies has been briefly and vividly termed "incomplete institutionalization" by Andy Cherlin (1978). Originally, this expression hinted at the absence of conventional roles taken on by family members when no biological-only kinship exists. Over the decades, it has extensively labelled the complexity of alternative families formed by at least one partner who experienced a previous union and, preferably, at least child to either partner.

The complexity of stepfamilies is primarily mirrored by the childbearing, which, according to my findings, does not strictly adhere the rules advanced in the existing literature. In stepfamilies, it seems that childbearing decisions are made sequentially as if one or both partners revise their ideal number of children to accommodate the children's configuration in the new union (Iacovou & Tavares, 2011). In a childless couple, the decision of a first shared child seems motivated by the desired to manifest partners' capital and experience parenthood. In a stepfamily, the decision of a first shared child is dependent on the presence and the number of stepchildren and, possibly, on the gender of the partner who brought children to new union. On the one hand, in many stepfamilies, the first shared child is also the first born to one of the parents, providing unique values not associated with step-parenthood, such as kin ties and continuation of a family line (Homann-Marriott, 2015). On the other hand, couples with at least two stepchildren supposedly give up on a shared child more frequently than couples with one stepchild, possibly because they perceive higher costs of children, making no difference between stepchildren and biological children (Bulatao, 1981).

In a childless couple, a second shared child might be valued for the biological relationship to the first. In a stepfamily, a second shared birth does not appear as valued as for a nonstepfamily, because the cost associated to children from previous partners supposedly act as a disincentive (Griffith et al., 1985; Ganong and Coleman, 1988). Therefore, second births in complex families prove to depend strongly on the number of pre-union children, as found in Vikat et al. (1999). Ultimately, Chapter 4 reveals that step-children perceived as less costly than biological children when stepfamilies decide on a first shared birth, but only under specific conditions (for instance, when only one step-child lives in the household). When it comes to the second or higher-order shared birth, the step-children seem to be as strong deterrent on childbearing as couples' biological children.

In second place, the allegedly higher instability of stepfamilies as opposed to first-union and childless couples due to their incomplete institutionalization (Bumpass & Castro Martin, 1990) has no empirical support. I do find only that a first child cements a union, *ceteris paribus*, also controlling for the unobserved heterogeneity. No additional risk of break-up affects partnerships who share at least one child living with at least one half-sibling, vis-à-vis families who do not have stepchildren. In contrast to Henz & Thomson (2005), I do not find evidence that a partnership's stability is positively associated to the number of shared children. In contrast to Heintz-Martin & Le Bourdais (2011), stepfather families do not prove systematically more prone to dissolution than stepmother families or blended families. In general terms, the number of risk factors of dissolution – partners' degree of union, partners' parental status, type of current union – is so high and interconnected that is very hard to accurately outline a profile of the family configurations more prone to break-up.

Family psychology has sought more extensively than demographic research to establish the mechanisms linking family well-being to separation in a variety of family combinations. The empirical research has investigated: the well-being of second or higher-order marriage with respect to first order marriage (e.g., Barrett, 2000; Hughes & Waite, 2009; Williams & Umberson, 2004); the quality of emotional bonds between stepfathers and children (King et al., 2014; Stewart, 2005; White et al., 1985; Weaver & Coleman, 2010), stepmothers and children (e.g., Vogt Yuan & Hamilton, 2006, Ganong et al., 2011; Weaver & Coleman, 2005); the well-being in married or cohabiting stepfamilies (Brown, 2006; Manning et al., 2014, 2015).

Cutting-edge research on stepfamily well-being, which addresses family structure, measures of stepfamily processes and relationship quality, family members' mental health and socio-cultural variables (e.g., Ganong & Coleman, 2017; Ganong, Coleman & Russell, 2015) should represent a reference point for demographic research. As a demographer, I think that research in demography should complement the interest in transitions (e.g., to a – new – birth, to a dissolution) with a focus on processes (e.g., perception of relationship quality; see Brown et al.2015; and Balbo & Ivanova, forthcoming) and broaden the way complexity is conceptualized. To this end, a possible development is to examine the family complexity from the point of full-, half-, and step-siblings and not only from the point of view of parents with respect to their children (Brown et al., 2015). Another way to expand the knowledge about stepfamilies is to capture complexity over time: a person is not only a node of links with his/her current family members, but also a hub of relationships with his/her past family members. For instance, researchers are including non-resident parents and siblings who do not live in their stepfamily household (Carlson, McLanahan, Brooks-Gunn, 2008; King, 2007; Schenck et al., 2009) obtaining more complex perspective of family processes. The growing awareness of complexity of family is pushing scholars to implement the best practices from other disciplines to describe, with increased detail, the role of stepfamily members, and gain clearer pictures of how the dynamics of various types of stepfamily household unfold.

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