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Union Dissolution Decisions and Childbearing in Subsequent Unions: A Study of Australian Panel Data

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NON-TECHNICAL SUMMARY

Recent decades have witnessed dramatic changes in partnership behaviour, with cohabitation, union dissolution and re-partnering (beyond marriage) on the rise in most industrialized countries, including Australia. As a consequence, childbearing within marital unions has decreased dramatically. While the increase in childbearing within cohabitation is well documented, less attention has been paid to the implications of increasing diversity in the number of partners and union statuses for contemporary childbearing. As childbearing years are the time when much of the increasing turnover in partnership occurs, these new partnerships provide opportunities for (further) childbearing. In fact, some initial evidence indicates that childbearing with new partners may constitute a large number of births in contemporary societies. The bulk of the studies, however, focused on continued childbearing with subsequent partners or childbearing in the context of step families. The dissolution of first marital unions and its relation to post-marital fertility has received little attention.

Our study contributes to understanding childbearing patterns after marital dissolution in four ways. First, we examine the associations between union dissolution decisions and post-marital first-time parenthood as well as parity progressions. Second, we elaborate and test an explanation to post-marital childbearing that builds on the thesis that relationships are instrumental for childbearing. We argue that individuals initiate union dissolutions to leave union contexts that are not deemed appropriate for parenthood or for a rewarding family life. Third, we address selective processes in post-marital fertility. Although union and childbearing decisions are highly inter-related, the specific mechanisms that lead individuals to dissolve unions, re-partner, and build or grow their families are still largely unknown and so we account for these unobserved factors in our modelling strategy. Fourth, while previous scholarship on childbearing after union dissolution has come from North America and Europe, we focus on a different context- Australia. We use data from the Household, Income and Labour Dynamics in Australia (HILDA) survey.

Our results show that while union dissolution in Australia is associated with lower rates of first-time parenthood, rates of parity progression are similar across stable first marital and subsequent unions. This suggests that in Australia re-partnering serves as a driver for continued childbearing and compensates for lost births (from dissolution) to some degree. Against our arguments of dissolution initiation as an instrument for achieving parenthood, we find that initiating the dissolution of the first marital union does not significantly reduce the time to first or higher-order conceptions. This result supports the idea that post-marital childbearing rationales are heterogeneous and suggests that union dissolution decision making is complex and includes a range of conditions that might or might not be associated with fertility plans. Finally, we find that childbearing and union dissolutions are associated on individual-specific unobserved factors. Thus, we conclude that further research is needed linking the causes and context of union dissolution with post-marital fertility behavior.

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Abstract

The extent to which childbearing occurs within marital unions has decreased dramatically over recent decades. While a wealth of studies examined recent patterns of childbearing out-of-wedlock and premarital childbearing, research has been less systematic on deciphering childbearing patterns after marital dissolution. Our study contributes to understanding of the latter by examining the associations between union dissolution decisions and post-marital first-time parenthood and parity progressions. We argue that individuals initiate union dissolutions to leave union contexts that are not deemed appropriate for parenthood or for a rewarding family life. We test this using hazard regression models for first-to-fourth order conceptions leading to live births. The analyses are done in the context of multi-process modelling to address selectivity due to individual-specific unobserved factors that lead individuals to dissolve unions, re-partner, and build or grow their families. The sample is restricted to women aged 16 to 40, who were observed since their first marriage, from the panel study Household, Income and Labour Dynamics in Australia (HILDA) survey. Our results show that while union dissolution is associated with lower rates of first-time parenthood, rates of parity progression are similar across stable first marital and subsequent unions. Initiating the dissolution of the first marital union does not significantly reduce the time to first or higher-order conceptions. Since we find conception episodes and union dissolutions to be positively associated on individual-specific unobserved factors, we conclude that further research is needed linking the causes and context of union dissolution with post-marital fertility behavior.

Keywords: childbearing; union dissolution; multi-process event-history analysis; Australia

Introduction

Family trajectories in contemporary societies are increasingly plural (Widmer and Ritschard 2009; Beaujouan 2012). Recent decades have witnessed dramatic changes in partnership behaviour, with cohabitation, union dissolution and repartnering (beyond marriage) on the rise in most industrialized countries, including Australia (Gray 2015). A consequence of this is that the universality of marriage as the first and only union type over the life course has indeed declined. A number of studies have reported the increase in childbearing within cohabitation (Kiernan 2004; Kennedy and Bumpass 2008; Perelli-Harris 2014), with some arguing that cohabitation has taken on many of the meanings of marriage (Smock 2000; Raley 2001) to the degree that marriage as an institution has almost lost its salience, including marriage as an optimal context for childbearing and childrearing (Cherlin 2004).

Less attention has been paid to the implications of increasing diversity in the number of partners and union statuses for contemporary childbearing. If childbearing is no longer exclusively taking place within the confines of the first marital union, it is not only due to increasing pre-marital fertility in temporary or in permanent cohabiting unions, but also because childbearing is increasingly common in second and subsequent unions. The small, but growing literature that have looked into childbearing after marital dissolution and their correlates coincide in that childbearing with new partners may constitute a large number of births in contemporary societies (Beaujouan 2010; Megiolaro and Ongaro 2010; Thomson et al 2014, Vanassche et al 2015). The bulk of the studies, however, focused on continued childbearing with subsequent partners (Sweeney 2010; Holland and Thomson 2011) or childbearing in the context of step families (Carlson and Furstenberg 2006). The dissolution of first marital unions and its relation to post-marital fertility has received little attention.

In this paper, we extend the scope of prior research in three ways. First, we study a wider set of childbearing episodes after the first marital union in Australia, including first-time parenthood and further parity progression regardless of partner's parental status. Second, we link previous union dissolution decisions with post-marital childbearing by looking into union dissolution initiation as a potential mechanism. Third, as union and childbearing decisions are highly inter-related with specific drivers largely unknown, we explicitly account for this in our modelling strategy. To this end, we use data from the Household, Income and Labour Dynamics in Australia (HILDA) Survey and deploy a simultaneous modelling framework of childbearing and union dissolution transition rates.

Existing Evidence

Much of our knowledge on childbearing after union dissolution (Thomson et al 2014) comes from two growing, interrelated areas of research. First, a body of research has been devoted to improve our understanding of patterns and mechanisms of fertility among couples that are also step families, where one partner co-resides with children from the other partner's prior relationship (Allen-Li 2006; Brown 2000; Heintz-Martin, Le Bourdais, and Hamplová 2014; Holland and Thomson 2011; Juby, Marcil-Gratton, and Le Boudais 2001, Vanassche et al 2015). Consistent findings of this research has shown that stepfamilies display higher birth rates than intact families with the same number of biological children. In close connection to these findings, a second growing body of evidence has revealed that multipartner fertility –childbearing with various partners over the life course– has rapidly increased in recent decades (Carlson and Furstenberg 2006, Guzzo and Furstenberg 2007, Thomson et al. 2014).

Despite the existing overlap between these two strands of research (Sweeney 2010), they cover distinct study populations. Stepfamilies do not always include a shared child in which case there is no multipartnered fertility (Guzzo 2014). In contrast, not all children in multipartnered families are coresidential -a defining feature of stepfamilies- and are henceforth outside the scope of step families literature (Kreyefeld and Heintz-Martin 2015; Guzzo 2014). Nevertheless, both literatures offer similar explanation by engaging with two main mechanisms-parenthood conferral and partnership commitment.

Parenthood conferral emphasizes individual considerations and suggests that individuals want to have children as a way to achieve the adult status of parenthood (Ivanova et al 2014; Murinko and Szalma 2016). Hence, individuals who have not yet had children from their pervious partnership are more likely to have children in their subsequent union compared to individuals that have already become parents, with childless new couples having the highest likelihood of childbearing (Vanassche et al 2015). Alternatively, partnership commitment suggests that a shared biological child can cement and bring social confirmation to a new partnership (Vanascche 2015; Ivanova et al 2014). It can therefore signal partners' commitment to each other by creating relational capital between all members of the union including children from previous partnerships (Vanassche 2015; Kreyenfeld and Heintz-Martin 2015). It follows that childbearing in higher order unions is not dependent on having children from a previous union (Vikat et al 1999) and the likelihood is in fact higher in subsequent unions compared to intact ones (Murinko and Szalma 2016; Beaujouan 2011).

Despite the two major explanations, fertility decisions after partnering are worth studying as couple considerations might indeed diverge. Unlike first unions where the birth of a first child might both confer parenthood status and solidify the relationship of both partners, these considerations might be disconnected from each other in higher order unions especially when the new union includes children from a previous relationship (Kalmijn and Gelissen 2007; Ivanova et al 2014). This disconnect has drawn a lot of scholarly attention to the role of prior children in fertility decisions after repartnering with no clear support for either mechanism (Murinko and Szalma 2016). Support for the partnership commitment hypothesis was first reported in the United States and confirmed that a repartnered woman's number of prior children did not affect her fertility after repartnering (Griffith et al 1985). Similar findings were reported for men in Sweden (Vikat et al 1999; Hollannd and Thomson 2011), France (Beaujouan 2011), the United States (Thomson and Li 2002; Stewart 2002) and Canada (Heintz-Martin et al 2014). In contrast, some studies from the Netherlands (Ivanova et al 2014; Kalmijn and Gelissen 2007), the United Kingdom (Jefferies et al 2000) and Austria (Vikat et al 2004) found that entering a new union with prior children is negatively associated with continued childbearing (i.e. partners are less likely to have a union-specific first child) lending support for the parenthood conferral hypothesis.

To this end, much of the research focus has been on understanding parenthood and parity progression in contexts where children already exist. The conditions that lead to dissolution of prior unions have received little attention to explain childbearing over the life course. In what follows, we propose an explanation for post-marital fertility based on how potential mismatches within the union context and the pathway to union dissolution can have implications for fertility life courses.

The Current Study

Our study complements and broadens the scope of previous research in several ways. First, we study a wider set of childbearing episodes including post-marital first-time parenthood and continued fertility, regardless of partner's parental status. This is important as individuals and particularly women are increasingly delaying their childbearing to prioritize the consolidation of a career (Blossfeld and Huinink 1991), and are more likely to have had multiple partners before the birth of their first child (Wu and Schimmele 2005).

Second, we provide a more nuanced examination of the parenthood conferral and partnership commitment hypotheses by building on life course theory and the notion that partnerships and childbearing trajectories are highly interdependent. To do so, we examine a potential mechanism that links both processes- union dissolution initiation. We argue that, among other reasons, individuals would initiate the dissolution of a marital union in order to achieve parenthood or a rewarding family life within the context of a new union when it is not realizable in an existing one. We build on the premise that along with prioritizing career achievements, difficulties in finding the 'right' partner might be a decisive factor in the delay in childbearing or unwanted childlessness.

Third, we address selective processes in post-marital fertility. Although union and childbearing decisions are highly inter-related, the specific mechanisms that lead individuals to dissolve unions, re-partner, and build or grow their families are still largely unknown and so we account for these unobserved factors in our modelling strategy. While previous fertility scholarship has addressed selectivity, our study is the first to examine this in the context of union dissolution and childbearing in higher order unions.

Fourth, while previous scholarship on childbearing after union dissolution has come from North America and Europe, we focus on a different context- Australia. Recent evidence from Thomson and colleagues (2014) report that in 2008 in Australia, 12 percent of women with two children, and 16 percent of women with three children, had their children with two fathers or more. Exploring post-marital childbearing and its potential link with previous union dissolution decisions is important to advance knowledge on this issue in Australia.

Life course interdependence: Partnership and Childbearing

We adopt a life course perspective which conceives individual life paths as sequences of purposive biographical transitions in different life domains (Lindenberg and Frey 1993; Huinink & Feldhaus 2009; Huinink and Kohli, 2014). We are particularly interested in the internal dynamics of life course development. That is, how decisions and expectations in one life domain (i.e. childbearing) impact behavior in another domain (i.e. union dissolution). We build on the principle of life course interdependence, and envisage progresses in partnership and childbearing trajectories to be highly intertwined. This idea resonates with recent studies showing how future parenthood prospects affect other domains of life (e.g. place of residence, occupational career) to make conditions appropriate for childrearing (Vidal et al 2017). In a similar fashion, we propose that individuals will also make adjustments to their union situation if they perceive that it is not conducive for childbearing.

Union situations requiring adjustments that may lead couples to split can be due to conflict in one or multiple core dimensions of partnerships, including family formation. On average, partners' desires relating to fertility will often differ unless there is perfect assortative mating with regard to preferences (Voas 2003; Yeatmen et al 2013). It follows, that individuals may alter their fertility preferences in the context of a relationship. If (childless) individuals enter a relationship eager to embark on their childbearing careers but later discover that this is not shared with their partner, they might either re-adjust their preferences about family –including family size and the timing of childbearing– or re-adjust the existing union context to one that matches their preferences.

Spouses might share similar preferences for family but may evaluate the prospective family life within a union in light of other family arrangements such as gender relations and division of household labour. In contemporary societies, the increasing involvement of women in the labour market has not been paralleled by an increasing involvement of men in childcare and housework (Goldscheider et al 2015). Lack of gender equity within the household limits fertility and boosts separation rates, particularly among couples where women and men have different ideologies about gender relations. Therefore, it can be expected that women will initiate union dissolutions if they expect that prospective family childrearing arrangements will not be satisfying.

In light of the above, we propose that potential mismatches within the union context with regards to family preferences and arrangements can have implications for fertility life courses. With regards to fertility, individuals will face disruptions in their fertility careers in the form of (in-) voluntary childlessness, or delayed childbearing till they find the right partner that shares their family life prospects or preferences for children.

Therefore, we expect that *childbearing rates will be lower in unions that follow the first marital union* (Hypothesis 1).

Indeed, finding a new partner plays an important role in the fertility careers of individuals following a martial dissolution (Van Bavel et al 2012). Building on the parenthood conferral hypothesis, we further address marital union dissolution as a purposive act or an instrument for finding the right partner for childbearing. We propose that one way to examine this would be to look at who initiates dissolution of a marital union in light of diverging expectations and preferences about parenthood and family living arrangements. Initiating a dissolution can be beneficial for childbearing, particularly for women, because it reduces the time spent in a partnership that would not lead to childbearing, allowing for a wider window of reproductive time to find a suitable union context for childbearing. Above and beyond parenthood conferral, we expect that initiating a separation can also be beneficial to those who already are parents yet desire leaving a union context due to unsatisfactory family life or arrangements.[1] In such cases, re-partnering and further childbearing in a more satisfactory union context may be gained earlier.

Thus, we expect that *childbearing rates are higher among the spouses who initiate the marital union dissolution than among those who did not initiate the marital union dissolution* (Hypothesis 2).

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[1] Our argument builds on the premise that along with prioritizing career achievements, difficulties in finding the 'right' partner might be a decisive factor in the delay in childbearing or in unwanted childlessness. In fact, having a supportive partner was the second most important factor (after women's health) in the decision to have a child among childless men and women in a cross-national European Study (Testa 2007). This means that individuals might have experienced a union dissolution before having the opportunity to give birth and form a new union which usually takes time (Mills et al. 2011).

Prior research finds strong interdependence between childbearing and union transitions on factors that are often unmeasured in the data that favor childbearing (i.e. shorter time to conception) and stable unions (i.e. longer time to dissolution), and vice versa. Hence, the association might not only be partly spurious, but the direction of causality is also contested. Indeed, research on childbearing within non-marital unions suggests that the choice to get married is not independent from the choice to have children (Baizan et al 2003, 2004). More recent research on childbearing and marital dissolution shows that women in higher order unions *and* women with a higher propensity for union instability *both* have higher conception rates once selectivity is controlled for, suggesting a double direction of causality is at play (Leone and Hinde 2007). Although a direct test for selectivity is not new to the fertility literature (Baizan et al 2003; Lillard 1993), it has never been applied to the study of marital dissolution and childbearing in higher order unions before. Leone and Hinde (2007) is a notable exception which we build on but extend further by testing the motivations and contexts in which women have children in higher order unions.

We hence expect that *the commented associations between childbearing and union dissolution are non-causal, and level off after adjusting for common unobserved predictors* (Hypothesis 3).

Method

Data and sample

For the empirical analyses, we use fourteen waves of the panel survey Household, Labor and Income Dynamics in Australia (HILDA, 2001-2014). The HILDA survey is a nationally representative study that allows tracking individuals over time, including the multiple union contexts in which respondents may engage over the life course (Watson et al. 2015).

To address childbearing since the first marriage, we retrieve relevant information on women's marital and fertility histories using retrospective records of information that predate the survey and contemporaneous information collected in each survey wave. We examine first-to-fourth-order childbearing episodes, because of the salience of higher-order parities in multi-partner fertility episodes. Information about cohabitating unions is available in each survey wave, and is used in our study of childbearing in cohabitating as well as marital unions following the first marital union. We use separation (or end of spouses' co-residence) to measure union dissolution. Separations usually precede divorces, and thus, can be considered an initial indicator of union dissolution. Additionally, not all separations end in divorce. We note that re-partnering with the previous spouse within a short period of time (e.g. under 1 year) is considered a transitory separation, which is not accounted for in our calculations.

The initial sample consists of 7,607 women of childbearing age, between age 16 and 40, who provide information on marital and childbearing histories. Due to the study focus, we exclude women who had children from relationships prior to the first marital union (n=32),[2] women who already had four children in their first survey participation (n=23), never married women throughout the observation window (n=3,400), and observations of women before the first marriage and following the experience of widowhood. We further restrict our sample to women who have been observed sometime in their first marital union throughout the life of the panel (2001-2014), because available information on spousal separation initiation makes reference to union dissolutions within the observation window (since 2001). This required to exclude women who separated (or became widows) from 1st marriage before 2001, or thereafter before their first survey participation (n=1,387). We also exclude survey drop-outs before 2005 (n=415), because the question on spousal separation initiation is collected since 2005. Our sample is biased towards respondents in recent marriages and respondents in lasting marriages due to these exclusions. To limit the potential effects of this bias, we further restrict the sample to women who married more recently, since 1990 (n=138).[3] After exclusions, our sample consists of 2,212 women.

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[2] Respondents reporting childbearing in pre-marital cohabitation that turns into a first marriage are not excluded from the sample. These cases contribute to analyses since the first childbearing episode observed after marriage.

[3] Our sample selection is partly driven by lack of information on marital separation initiation for unions that were dissolved before 2001 and respondents who dropped the survey before 2005. These (left-censored) cases are potentially problematic if they are not independent from our analysis outcome (e.g. early divorces leading higher post-divorce fertility). One way to address the degree of bias introduced is to analyse only individuals with complete observations (i.e. exclude those who married before they enter the

Table 1. Sample summary: episodes and events

	Frequency	Percentage
Childbearing		
Episodes		
Parity 0	1,858	47.0
Parity 1	1,128	28.5
Parity 2	743	18.8
Parity 3	228	5.7
Overall	3,957	100.0
Events (childbirth)		
First	860	47.4
Second	668	36.8
Third	209	11.5
Fourth	77	4.2
Overall	1,814	100.0
Union dissolution		
Episodes		
1st marital union	2,212	93.9
Subsequent unions	144	6.1
Overall	2,356	100.0
Events		
First marital separation	263	84.6
Further separations	48	15.4
Overall	311	100.0
N - individuals	2,212	100.0

HILDA Survey data (2001-2014).

The units of analysis are childbearing episodes in the first marital union and subsequent unions. We define a childbearing episode as the lapse of time in months since the first marital union until the first conception leading to live birth, or since the delivery until the second conception leading to live birth. We measure conceptions by subtracting nine months to the time of childbirth. Since only the year of birth is available in the dataset, we assume childbirths occur (i) one month before the interview if the respondent is partnered at the time of the interview, (ii) the last month the individual was partnered if the respondent was not partnered at the time of the interview, or (iii) in July if the individual was not partnered since the last interview. In our sample (see Table 1), we

survey). Results of an analysis applying such restriction do not differ substantively from an analysis to the larger sample, and thus, we conclude that the degree of bias introduced is small.

observe 1,858 episodes leading to 860 first-order conceptions, 1,128 episodes leading to 668 second-order conceptions, 743 episodes leading to 209 third-order conceptions, and 288 episodes leading to 77 fourth-order conceptions. We note that higher numbers of second-order episodes than first-order conception are due to some women entering the study when they already have had the first child (within the first marital union).

Key Predictors

We distinguish several partnership episodes and statuses that are relevant to our study. First, individuals are initially observed in the *first marriage* (reference category), but these can move to a post-marital episode in which we distinguish those (i) *separated* and not in a (co-residential) relationship, from those (ii) in a *post-marital union* with a new partner. Partnership episodes and statuses are specified in hazard regression models as time-dependent variables. In our sample 293 respondents experience the dissolution of the first marital union. We also observe 144 re-partnering episodes during the study window (see Table 1). Due to low numbers, we do not further distinguish those in post-marital unions across those cohabiting with a new partner and the re-married with a new spouse.

Another predictor of interest is the decision-making status of the first marital union. To those women who during the course of survey participation separated or divorced from a union, the following question was asked: *Whose decision was it to finally separate?* There are three possible response categories: (i) *Mostly mine*, (ii) *mostly partner's*, and (iii) *joint*. We note that most respondents initiate the dissolution of the union, and a much smaller proportion are initiated by partners or are joint decisions (See Table 2). This is consistent with previous research that shows that women are more likely to initiate a separation or divorce (Hewitt, Western and Baxter 2006; Hewitt 2009). We find that almost one in five separating individuals do not provide an answer to the question. We use different strategies for handling these missing cases –including list-wise deletion, dummy variable, or predictive model imputation– with no substantive differences in results across them.

Table 2. Decision maker of first marital union dissolution (women's responses)

	Frequency	Percentage
Whose decision was it to finally separate?		
Mostly mine [female partner]	115	43.7
Mostly partner's [male partner]	43	16.6
Joint	55	20.9
No response	50	19.0
Union dissolutions	263	100

HILDA Survey data (2001-2014).

Control Variables

We control for a number of theoretically-relevant predictors of union dissolution that could act as confounders to the associations under study. We consider whether the respondent lives with non-biological children by including an indicator for *step-children* in the household (ref. no step children present). Additional time-varying covariates include age categories (16-25, ref.; 25-30; 30-35; 35-40), educational attainment in three categories (high: university qualifications; medium: grade 12 and higher vocational qualifications; and low: grade 11 or below, ref.), employment status (employed; non-employed, ref.), and owner-occupancy of primary residence (ref. tenant). This is because home ownership has been shown to have an effect on post-marital childbearing, being at its peak during first marital unions and diminishing in subsequent ones (Lersch and Vidal 2014). Additional variables include calendar period (2001-2004, ref.; 2005-2009; 2010-2014), territory or state of residence (six categories) and a time-constant foreign-born indicator (ref. Australian born). We also include a measure of religious affiliation (ref. not affiliated in any religion) as it has been shown to be positively associated with fertility and marital stability (Hubert 2015) suggesting a depressed effect on marital separation and childbearing in higher order unions. Summary statistics of model covariates can be consulted in Table A1 in the appendices.

Analytical strategy

We examine first (parity 0) - and higher-order (parities 1, 2 and 3) childbearing episodes of women using hazard rate modelling. The model can be written as

$$\ln h_{ij}^{C}(t) = y_i^{C} T(t) + \sum\nolimits_k \beta_i^{C} x_{ij}^{C}(t)$$

where $\ln\, h_{ij}^{C}$ is the log-hazard of a conception c of i order $(i = 1,2,3,4)$ of woman j. The hazard rate of childbirth depends on the duration of the episode, or the time since the individual is at risk of event occurrence until the parity-specific conception of a life birth. The duration function is specified in $y_i^{C}T(t)$. Since we examine childbearing that occur in the first marital union and afterwards, we define the duration process as the time elapsed in months until conception event leading to the first life birth (or censoring) since the formation of the union. For second-to-fourth order-childbearing, we define the duration process as the time elapsed in months until conception event (or censoring) since the birth of the preceding child. We model the duration function as piecewise-linear splines (or piecewise-linear Gompertz), with two nodes at the 12th and 32th month for parity 0 and with one node at the 24th month for parity 1 and 2. Due to lower event numbers, no nodes are designated for parity 3; hence, the duration function is linear. Piecewise-linear splines enable to estimate slope-coefficients of the duration function for the designated duration segments. In general, in the context of a marital union most first births occur within the initial years since union formation, and most higher-order births occur within the first years after childbirth. Thus we expect a positive slope-coefficient for earlier duration intervals. The slope might level off for later duration intervals. All other time-varying and time-constant covariates presented in the previous section are specified in $\beta_i^{C} x_{ij}^{C}(t)$.

To test our hypothesis that childbearing rates are lower after marital separation, we will assess the size and significance of the coefficients on indicators of post-marital episodes as *separated* and in a *post-martial union* adjusting for the duration process, and additional covariates that confound the association. To test our hypothesis that initiating the dissolution of the marital union increases post-marital fertility, we will assess the size and significance of the coefficients on an indicator that combines respondent and joint initiator status, against partner's initiator status (reference category) adjusting for the duration process, the union status and episode, and additional covariates that confound the association. We justify the combination of respondent and joint initiator status in that the respondent was active in the decision making process. We have also performed analysis with separate measures for respondent and joint initiator status and no significant differences in results across indicators were found. We assess these association across first-time parenthood and continued fertility episodes through interactions of parity 0 episodes and higher-order parity episodes with union status and union dissolution decision maker indicators.

Our estimates of post-marital childbearing could be biased due to reverse causation if childbearing and union episodes are commonly predicted by factors not specified in the model. This is because the unobserved variation remains in the model error term (Kulu and Steele 2013). One solution is to model the unobserved heterogeneity using Lillard's (1993) multi-process estimation strategy (see also Upchurch et al 2002). To this end, we simultaneously estimate two equations: one for the hazard of conception together with the hazard of first marital and subsequent union dissolution. The jointly estimated equations can be written as follows

$$\ln h_{ij}^{C}(t) = y_i^{C} T(t) + \sum\nolimits_k \beta_i^{C} x_{ij}^{C}(t) + u_j^{C}$$
$$\ln h_{ij}^{S}(t) = y_i^{S} T(t) + \sum\nolimits_k \beta_i^{S} x_{ij}^{S}(t) + u_j^{S}$$

The first equation refers to the hazard model for parity-specific conceptions as specified before. It also contains a random term u_j^C that captures individual-specific unobserved heterogeneity fixed across childbearing episodes of the same women. The second equation refers to the hazard model for union dissolution, where $\ln h_{ij}^{S}$ is the log-hazard of a separation s of i order $(i = 1, \dots N)$ of women j. The first separation refers to the dissolution of the first marital union. The duration function specified in $y_i^S T(t)$ is a piece-wise linear transformation of months of the union episode for the following intervals: (i) union entry until second year, (ii) second year until seventh year, and (iii) seventh year onwards since the start of the union episode for the interval. In Australia, union dissolution of the first marital union increases during the first few years, and decreases steadily afterwards (Hewitt, Baxter and Western 2005). Time-varying and time-constant covariates are specified in $\beta_i^C x_{ij}^C(t)$, which include the predictors of union dissolution commented in the previous section. As we expect a reciprocal relationship across fertility and union dissolution outcomes, we also include indicators for the number of biological children (ref. no children; 1 child; 2 or more children) as time-varying covariates. Children are expected to be a protective factor to union dissolution, at least while they are young. We also include a time-varying covariate for union order (ref. first marital union; higher-order union), as some women are observed in unions after the dissolution of their first marriage. Higher order unions are often found to be less stable than first-order unions. Finally, a random term u_j^S captures individual-specific unobserved heterogeneity fixed across union episodes of the same women.

The equation-specific random terms are assumed to follow a joint bivariate normal distribution with mean 0 and variances and covariance to be estimated (see equation below). The correlation term ρ captures the common unobserved heterogeneity affecting the conception and union dissolution hazard equations. A significant non-zero correlation indicates that there are individual-specific unobserved factors that commonly affect both, conception episodes and union dissolutions. Hence, we can test selection as a mechanism assessing the direction and significance of the correlation term ρ. Modelling ρ also minimizes bias due to reverse causation in estimates of other model coefficients.[4]

$$\begin{pmatrix} u^C \\ u^S \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma^2_{u^C} & \rho_{u^C u^S} \\ \rho_{u^S u^C} & \sigma^2_{u^S} \end{pmatrix} \right)$$

[4] Model identification is achieved through repeated childbearing episodes and union dissolutions for some women in the sample (Steele 2008).

Results

Descriptive survival curves of first and higher-order childbearing episodes by separation status (See Figure A1 in appendices) indicate that separation associates with delayed and lower rates of first-order conceptions than remaining married. Differences across groups on timing and occurrence rates also exist but are smaller for higher-order conceptions. Breaking up separations by decision-maker status shows that those who initiate the dissolution of the union have more and faster transitions to first child than those who do not initiate. Despite this, differences are not statistically significant. These results do not account for differences across groups in age and other confounding factors; hence, we move now to the multivariate results.

Table 3 displays estimated log-hazard rates of selected model coefficients for conceptions leading to live births. Model 1 shows results for process-specific estimations, where correlation across fertility and union dissolution processes is assumed to be zero. All conceptions from any parity have been estimated together, in the same equation. Parity-specific duration parameters are included in the model to acknowledge differences in conception occurrence and timing across parities.

With respect to union episode, results in Table 3 show that the rate of conception leading to a live birth is lower after the dissolution of the first marital union, for first-time parenthood as well as for continued fertility. Log-hazard rates are negative and statistically significant for the separated. Coefficient for the separated have larger size and statistical significance for first childbearing episodes (β=-1.86, $p<0.01$) than for higher-order episodes (β=-.85, $p<0.1$). For those in a union after the first marriage, the associations are also negative, but non-statistically significant. The associated coefficients are somewhat less negative than those of the separated for first childbearing episodes (β=-.89, $p>0.1$), and particularly for higher-order episodes (β=-.08, $p>0.1$).

Turning to the union dissolution initiation, we find that the directions of the associations between union dissolution and childbearing vary if the respondents initiate the dissolution of the marital union. As expected, the association between initiating a union dissolution and first-order conception leading to live birth is positive, but non-statistically significant (β=0.48, $p>0.1$). Further, the association between initiating a union dissolution and higher-order conception leading to live birth is negative, but non-statistically significant (β=-0.28, p>0.1). Statistically insignificant associations indicate that the propensity to have a first or higher-order child after the first marital is not different across those who initiated the union dissolution and those who did not. This suggests that initiating the dissolution of the union is neither likely associated with preferences for childlessness or smaller families nor related to adverse couple experiences that undermine preferences for (further) childbearing.

Results of other model covariates in the hazard rate for first and second-order conceptions are presented as appendices in Table A2. We find that our results are mostly in line with those of previous research. The hazard rate of conception leading to first and higher-order conceptions increases during the first years after union formation and childbirth (respectively) and levels off thereafter. The coefficients for education, religious affiliation, and homeownership are all statistically significant and positive, while the coefficients for older age groups, employment status and foreign-born group are statistically significant and negative.

Table 3. Hazards of first-to-fourth-order conceptions leading to life birth (selected coefficients).

	Model 1 Process-specific estimation	**Model 2** Multiprocess estimation
First-time parenthood or Parity 0 *		
1st marriage	ref.	ref.
After 1st marriage		
Separated	-1.86 ***	-2.13 ***
	(0.55)	(0.58)
In a union	-0.89	-1.16 *
	(0.65)	(0.66)
Respondent initiates marriage dissolution		
No	ref.	ref.
Yes	0.48	0.53
	(0.61)	(0.62)
Continued fertility or Higher-order parities *		
1st marriage	ref.	ref.
After 1st marriage		

Separated	-0.85 *	-1.00 *
	(0.51)	(0.55)
In a union	-0.08	-0.20
	(0.50)	(0.53)
Respondent initiates marriage dissolution		
No	ref.	ref.
Yes	-0.28	-0.31
	(0.53)	(0.56)
Random variation (unobserved heterogeneity)		
$\sigma^2_{u^c}$	1.27 ***	1.27 ***
	(0.05)	(0.05)
$\sigma^2_{u^s}$		1.86 ***
		(0.23)
Correlation: $\rho_{u^s u^c}$		0.27 ***
		(0.07)
Episodes	3,957	3,957
Events	1,773	1,773
ln-L	-17394.84	-17386.54
P(Chi2)	<.001	<.001

HILDA Survey data (2001-2014). Significance levels: * 0.1 ** 0.05 *** 0.01
Coefficients are log-hazards. Standard errors in parentheses under coefficients. Multiprocess model is a joint estimation of childbirth hazard equation and union dissolution hazard equation. Covariates included in all models: duration function, age groups, calendar period groups, education level groups, employment status, foreign born, religious affiliation, homeownership, State or Territory of residence, and a dummy for non-response on the union dissolution decision maker.

To assess selectivity in the study association due to unobserved factors commonly affecting both processes, and sensitivity to selection in our model estimate, we estimate conception hazards and union dissolution hazards simultaneously. Results of the multiprocess estimation for the hazard of first-to-fourth order conceptions are presented in Model 2 in Table 3, and the union dissolution hazard is presented in Table A3. A non-zero correlation term ρ will be taken as evidence for common unobserved heterogeneity. In fact, we find that the correlation term is different from 0, with a statistically significant positive coefficient ($\rho = 0.27; \mathrm{p} < .01$). In our model, unobserved factors among women lead them to both dissolve a (marital) union and to achieve children earlier in the life course. This result is consistent with those of similar models based on samples of Brazilian women (Leone and Hinde 2007).

After adjusting for common unobserved heterogeneity, we observe a few changes in the size and statistical significance of model coefficients of the conception equation in Model 2. First, the negative associations between union episodes following the first marital union and first-order conceptions become larger. The association even turns statistically significant at the margin for those observed in a union after the first marital union ($\beta = -1.16$, $p<0.1$). The change in coefficients after adjusting for selection to post-marital childbearing, indicates that the lower propensity to childbearing after union dissolution observed in the previous model is accentuated by the selection into earlier childbearing by those individuals who separate. Second, the negative association for childbearing in subsequent unions also becomes larger for second-order conceptions. However, the association remains small and statistically insignificant for those observed in a union after the first marital union ($\beta = -0.20$, $p<0.1$). This result aligns well with finding of the literature, where in some instances rates of continued fertility after the dissolution of the union are not significantly lower than those in first unions. Third, the coefficients of union dissolution initiation status in Model 2 remain unchanged and statistically insignificant for first and higher-order conceptions.

Though tangential to our research in this paper, we present results of the union dissolution equation in Table A3 in the appendices. In line with previous research, we find that union dissolution is less likely among the older, the highly educated, for those in owner-occupied homes and when biological children are present in the household. We also find that step-children have no relevant impacts on the dissolution of the union. Higher order unions are associated with lower union dissolution rates, but partly level off if the higher-order unions are unmarried. One should take into account that these results are based on multiprocess estimation. That is, once controlling for the actual high propensity to separate among those who enter higher-order unions according to crude rates, subsequent unions do actually protect from union dissolution. This is likely as individuals learn from previous union experiences.

Discussion and Conclusion

In this study we assess the associations between union dissolutions and fertility decisions relating to first parenthood and continued fertility in Australia. We address the question of whether initiating a marital dissolution is associated with post-marital childbearing. We test this using long-running Australian panel data and fitting piecewise linear log-hazard models of first-to-fourth-order conceptions leading to live births. We also address issues of selectivity in the study associations by estimating conception and union dissolution hazard models simultaneously, in the framework of multilevel multiprocess estimation.

Some main findings arise from our research. First, childbearing is less likely after the dissolution of the first marital union, though the association is not statistically significant for repartnered women with children from the first marital union. This lends partial support for Hypothesis 1, as fertility rates are lower after the dissolution of the first marriage only in some cases. Second, initiating the dissolution of the marital union is not statistically associated with childbearing. This lends support against Hypothesis 2, as initiating the dissolution of the union does *not* directly affect the timing to conception leading to childbirth in unions that follow the first marriage. Third, we find childbearing episodes and union dissolutions to be positively associated on individual-specific unobserved factors. This indicates selective union dissolution among individuals with higher propensity for early childbearing, but only lends partial support for Hypothesis 3 because the associated selection bias only affects somewhat the size and statistical significance of the associations between fertility and union dissolution.

Our study is embedded in a larger literature that devotes itself to the changing intersections between nuptiality patterns and fertility dynamics. The extent to which childbearing occurs within marital unions has decreased dramatically over recent decades. This changing relationship between partnership and parenthood has been a central feature of family life transformation recently (Holland and Thomson 2011). Certain dynamics such as the increase in divorce, the diffusion of cohabitation, or the changing meaning of marriage and cohabitation-collectively referred to as the second demographic transition have certainly contributed to this trend (Lesthaeghe 1995; Thomson et al 2014). Nonetheless, people *still* desire and have children and this is not confined to first marital unions (Ivanova et al 2014; Murinko and Szalma 2016). As childbearing years are the time when much of the turnover in partnership occurs, these new partnerships provide opportunities for (further) childbearing (Holland and Thomson 2011). This suggests that repartnering serves as a driver for childbearing and compensates for lost births (from dissolution) to some degree (Megiolaro and Ongaro 2010; Thomson et al 2012; Beuajouan and Solaz 2013). Along these lines, our findings suggest that in the Australian context this is true for continued childbearing, as parity progression rates are similar across those who repartner with children from the first marriage and those who remain in the first marriage. Our research, thus, contributes to understanding of contemporary fertility patterns, by showing that the apparent negative association between levels of marital separation and fertility might be driven by childless women missing or abandoning first-time parenthood.

Conceptually, our study innovates by offering an explanation to post-marital childbearing that builds on principles of life course interdependence, with feedback processes in individual's partnership and fertility career progression. We argue that marital separation leads to lower fertility not only due to the decrease in the proportion of time that women spent outside largely stable marital unions, as some research suggested (Hobcraft and Kiernan 1995; Thornton and Young-DeMarco 2001), but also due to the conditions that it takes individuals to find the right partner, or the optimal context for childbearing and childrearing. We focus on the internal dynamics of the life course, linking decision-making processes in one life domain (i.e. initiating a union dissolution) with outcomes in other life domains (i.e. childbearing). In line with arguments of partnerships as instruments for achieving parenthood, we develop a nuanced approach signaling the role of partner match, through dissolution initiation as a key indicator. Results from our empirical analysis do not lend support for this thesis, which suggests that union dissolution decision making is complex and includes a range of conditions that might or might not be associated with fertility plans. In close relation to this, we acknowledge the relevance of selectivity on marital and childbearing decisions. Empirically, ours is one of the few studies that examines how conditions of marital union dissolution associate with childbearing in subsequent unions. We also test for the direction of causality and assess the degree of unobserved heterogeneity by deploying the multi-process estimation of the hazard rates of childbearing and union dissolution.

Despite the contributions, our research has some limitations. First, our data is left-truncated and, thus, results have to be read with some caution. Due to incomplete cohabitation histories and key variables only available at the time of survey, our sample is restricted to childbearing episodes during the study window. This results in a study sample likely biased towards recently married women, or women who postponed childbearing within the marital union. Additionally, despite efforts to reduce bias in estimates by modelling common unobserved heterogeneity -through the estimation of the correlation term ρ- we note that other sources of heterogeneity (e.g. time-varying unobserved heterogeneity) may still prevent us from making definitive causal statements about the associations. Ideally, one would add relevant common time-varying and time-dependent predictors to reduce the weight of common unobserved heterogeneity. However, the restricted sample prevents us from including a broader set of predictors. Theoretical development would also be needed to assess what predictors should be chosen.

Subsequent research can look into more nuanced model specifications, including moderation effects of women resources through the examination of interactions between union dissolution (decision-making) and education or employment that we were not able to test due to small case numbers. Our analysis should be replicated for men as well, as differences in gender roles within couples are still pervasive, where women remain responsible for household functioning and childcare. According to our thesis, union dissolution decision making could reflect fertility intentions mismatch across partners or dissatisfaction with the marital union. Although these do not necessary lead to a union dissolution if there is a prospect to embark on childbearing with the partner in a later point in time, it would be interesting to further test associations based on such indicators. The role of ex-partners' characteristics, the division of household labour and gender role attitudes could also be explored in further studies as predictors of childbearing in subsequent unions. These align well with ideas and evidence that union dissolution associates with greater heterogeneity in childbearing behaviour, and with our finding that childbearing and union dissolutions are associated on individual-specific unobserved factors. Thus, we conclude that further research is needed linking the causes and context of union dissolution with post-marital fertility behavior.

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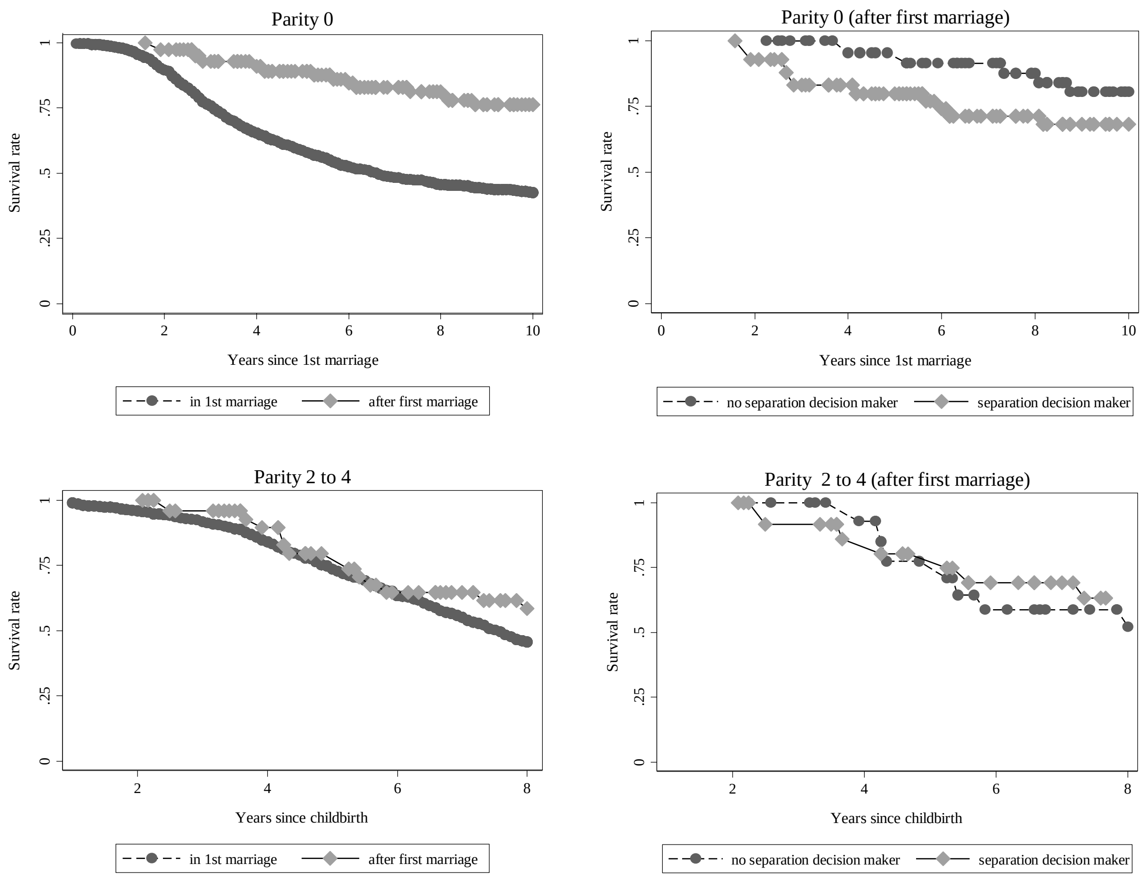
APPENDICES

Table A1. Summary statistics of model covariates (percentage of sample)

Educational level	
Low	16.0
Medium	45.3
High	41.7
Employed	80.4
Age groups	
16-25	20.5
25-30	44.2
30-35	46.7
35-40	30.3
Calendar period	
2001-2004	39.6
2005-2009	42.6
2010-2014	54.4
Step children	3.0
Foreign-born	23.0
Religious affiliation	62.5
Homeownership	51.0
State or Territory of residence	
New South Wales	32.4
Victoria	26.3
Queensland	21.4
South Australia	9.1
Western Australia	9.5
Other States and Territories	6.8
	100.0
Total (N)	(2,210)

HILDA Survey data (2001-2014). Percentage of women observed in each category throughout the observation window.

Figure A1. Parity-specific survival curves of childbearing episodes by separation status and decision maker.



HILDA Survey data (2001-2014). Survival curves based on Cox regression estimates adjusted by time varying covariates of separation of first marriage status (left panel) or separation decision-maker status (right panel). Differences across curves are statistically significant at the 95 percent for the left panel only.

Table A2. Hazard rates of first-to-fourth-order conceptions leading to life-birth.

	Model 1 Process-specific estimation	**Model 2** Multiprocess estimation
Parity 0 *		
1st marriage	ref.	ref.
After 1st marriage		
Separated	-1.86 ***	-2.13 ***
	(0.55)	(0.58)
In a union	-0.89	-1.16 *
	(0.65)	(0.66)
Respondent initiates marriage dissolution		
No	ref.	ref.
Yes	0.48	0.53
	(0.61)	(0.62)
Parity 1 to 3 *		
1st marriage	ref.	ref.
After 1st marriage		
Separated	-0.85 *	-1.00 *
	(0.51)	(0.55)
In a union	-0.08	-0.20
	(0.50)	(0.53)
Respondent initiates marriage dissolution		
No	ref.	ref.
Yes	-0.28	-0.31
	(0.53)	(0.56)
No response on marriage initiation	-0.34	-0.28
	(0.55)	(0.59)
Step children	-0.27	-0.31
	(0.28)	(0.28)
Religious affiliation	0.09	0.10
	(0.07)	(0.08)
Homeownership	0.41 ***	0.39 ***
	(0.07)	(0.07)
Educational level (ref: low)		
Medium	0.80 ***	0.81 ***
	(0.11)	(0.11)
High	1.09 ***	1.10 ***
	(0.11)	(0.11)
In employment (ref: not employed)	-1.50 ***	-1.50 ***
	(0.06)	(0.06)
Foreign-born	-0.80 ***	-0.79 ***
	(0.08)	(0.08)
Calendar period (ref. 2001-2004)		
2005-2009	0.52 ***	0.52 ***
	(0.09)	(0.09)
2010-2014	0.40 ***	0.42 ***

	(0.09)	(0.09)
State of residence (ref. New South Wales)		
Victoria	0.02	0.02
	(0.09)	(0.09)
Queensland	0.04	0.06
	(0.09)	(0.09)
South Australia	-0.23	-0.21
	(0.14)	(0.14)
Western Australia	0.13	0.14
	(0.12)	(0.12)
Others	-0.28 *	-0.28 *
	(0.15)	(0.15)
Age groups		
16-25	ref.	ref.
25-30	-0.16	-0.15
	(0.10)	(0.10)
30-35	-0.42 ***	-0.40 ***
	(0.12)	(0.12)
35-40	-1.01 ***	-0.99 ***
	(0.14)	(0.14)
Parity 0 - Piece-wise linear function of time since marriage		
0-1 years	0.34 ***	0.34 ***
	(0.06)	(0.06)
1-3 years	0.09 ***	0.09 ***
	(0.01)	(0.01)
>3 years	-0.01 ***	-0.01 ***
	(0.00)	(0.00)
Intercept	-9.98 ***	-10.04 ***
	(0.74)	(0.75)
Parity 1 - Piece-wise linear function of time since first child		
0-2 years	0.15 ***	0.15 ***
	(0.02)	(0.02)
>2 years	0.03 ***	0.03 ***
	(0.00)	(0.00)
Intercept	-9.92 ***	-9.96 ***
	(0.46)	(0.47)
Parity 2 - Piece-wise linear function of time since second child		
0-2 years	0.11 **	0.11 **
	(0.05)	(0.05)
>2 years	0.03 ***	0.03 ***
	(0.00)	(0.00)
Intercept	-11.26 ***	-11.31 ***
	(1.07)	(1.09)
Parity 3 - Linear function of time since third child		
Slope	0.02 ***	0.02 ***
	(0.00)	(0.00)
Intercept	-9.93 ***	-9.98 ***
	(0.50)	(0.50)

Random variation unobserved heterogeneity)		
$\sigma^2_{u^c}$	1.27 ***	1.27 ***
	(0.05)	(0.05)
$\sigma^2_{u^s}$		1.86 ***
		(0.23)
Correlation: $\rho_{u^s u^c}$		0.27 ***
		(0.07)
Episodes	3,957	3,957
Events	1,773	1,773
ln-L	-17394.84	-17386.54
P(Chi2)	<.001	<.001

HILDA Survey data (2001-2014). Significance levels: * 0.1 ** 0.05 *** 0.01
Coefficients are log-hazards. Standard errors in parentheses under coefficients. Multiprocess model is a joint estimation of childbirth hazard equation and union dissolution hazard equation.

Table A3. Hazard rates of union dissolution.

	Model 2 Multiprocess estimation
Union Dissolution	
Type of union	
Marital union	ref.
Cohabitation union	1.59 ***
	(0.28)
Union episode	
1st marriage	ref.
Post-marital union	-3.29 ***
	(0.48)
Number of biological children	
None	ref.
1 child	-0.81 ***
	(0.25)
2 or more children	-1.19 ***
	(0.26)
Step children	-0.40
	(0.54)
Religious affiliation	-0.21
	(0.19)
Homeownership	-1.10 ***
	(0.16)
Educational level (ref: low)	
Medium	-0.06
	(0.23)
High	-0.83 ***
	(0.27)
In employment (ref: not employed)	-0.26
	(0.17)
Foreign-born	-0.47 **
	(0.23)
Calendar period (ref. 2001-2004)	
2005-2009	1.46 ***
	(0.21)
2010-2014	1.43 ***
	(0.23)
State of residence (ref. New South Wales)	
Victoria	0.01
	(0.24)
Queensland	0.12
	(0.25)
South Australia	0.56 *
	(0.31)
Western Australia	-0.25

	(0.34)
Others	0.60 *
	(0.35)
Age groups	
16-25	ref.
25-30	-0.01
	(0.27)
30-35	-0.38
	(0.30)
35-40	-0.65 *
	(0.35)
Piece-wise linear function of time since union formation	
0-2 years	0.07 ***
	(0.02)
2-7 years	0.03 ***
	(0.01)
>7 years	0.01 ***
	(0.00)
Intercept	-9.53 ***
	(0.70)
Random variation unobserved heterogeneity)	
$\sigma_{u^c}^2$	1.27 ***
	(0.05)
$\sigma_{u^s}^2$	1.86 ***
	(0.23)
Correlation: $\rho_{u^s u^c}$	0.27 ***
	(0.07)
Episodes	3,957
Events	1,773
ln-L	-17386.54
P(Chi2)	<.001

HILDA Survey data (2001-2014). Significance levels: * 0.1 ** 0.05 *** 0.01
Coefficients are log-hazards. Standard errors in parentheses under coefficients. Multiprocess model is a joint estimation of childbirth hazard equations and union dissolution hazard equation.