Union formation and educational differentials in micro- and macro-level economic conditions in France (1993-2008)

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Abstract

The short-term impacts of economic recession, rising economic uncertainty and its educational underlying driver on fertility are well documented in the literature. Whereas postponement of union formation has been suggested as one of the main pathways through which economic conditions affect fertility, some papers have directly addressed the micro- and macro-level economic conditions on union formation. Union formation (especially marriage) hazards are theorized to decrease due to a lack of financial and social long-term prospects. Prolonged school enrolment, postponement of transitions and flexible partnership forms are identified as coping strategies to uncertainty in literature. However, resulting from different attachments to the labour market and the divergence in meanings that educational groups attribute to living arrangements, we expect variation in timing of union formation among varying educational levels. This contribution therefore aims to examine the entry into a first unmarried co-residence union and marriage after cohabitation among different educational levels in relation to their employment status and aggregate-level economic context. The analyses use union and employment histories (1993-2008) of male and female respondents between the age of 16 and 39 from the French Harmonized Histories and the second wave of the Generations and Gender Survey. To test our research hypotheses the analyses draw on discrete-time event history methods. We find that employment is particularly an important prerequisite for union formation among more educated men. In correspondence, susceptibility to aggregate-level economic context with regard to entry into a first cohabiting union prevails amidst highly educated men. A notable result is that men and women respond differently to both types of uncertainty, implying gender-specific effects. Our analyses furthermore suggest that some of the outcomes for the transition from cohabitation to marriage are surprisingly comparable .

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1. Introduction

Economic shocks have inspired demographic researchers to investigate the relationship between economic conditions and demographic behaviour for several decades. Since the financial crisis hit continental European economies in 2007, the topic has gained a lot of renewed interest (Sobotka et al. 2011). Because a substantial amount of evidence suggests that young school-leavers in particular are a vulnerable group in adverse economic conditions (Gangl 2002; Bell et al. 2007; Verick 2009; Skirbekk et al. 2010; Danzinger and Ratner 2010), researchers have drawn attention to the hampering effects of economic insecurity on transitions to adulthood (Adsera 2005; Berkowitz King 2005; Mills and Blossfeld 2005; Bucholz et al. 2009; Sobotka et al. 2011; Kreyenfeld et al. 2012). Analyses addressing both the micro- (Kravdal 2002; Kreyenfeld 2010) as well as at the aggregate-level (Adsera 2005; Goldstein et al. 2013) have principally shown a negative impact of insecure economic conditions on family formation processes. Some evidence has shown that union formation is related to economic variables at the individual and aggregate level as well, particularly concerning the selection of a type of union (i.e. marriage or cohabitation) (Kravdal 1999; Oppenheimer 2003; Prioux 2003; Wiik 2009; Kalmijn 2011; Vergauwen and Neels submitted). Correspondingly, delayed union formation has been suggested as one of the pathways through which economic conditions affects fertility timing in countries where the tie between marriage and childbearing has remained strong (Sobotka et al. 2011). Regarding union formation findings by Kalmijn (2011) highlight the importance of both income (affordability) and the favourable prospects of being employed itself. It is not only argued that having a sufficient amount of economic resources is important as a prerequisite for living independently together (Rindfuss and VandenHeuvel 1990; Aassve et al. 2002), but also the significance of regular employment and the development of an auspicious career trajectory are often emphasized as well (Kalmijn 2011; Blossfeld et al. 2005). In terms of Oppenheimer's uncertainty hypothesis social consequences of work are equivalently important in the partnering process. Work tends to structure potential partners' lifestyles and determines a considerable part of the socialization process to an adult role. During periods of employment uncertainty assortative mating is impeded since it is more difficult to predict how partnerships will work out (Oppenheimer 1988). In this contribution we address the role of employment of both men and women for union formation. Typically the importance of women's economic position is expected to depend to a great extent on the gender context (Sweeney 2002; Liefbroer and Corijn 1999). In gender egalitarian societies it is hypothesised that the importance of the male employment status diminishes as the costs of setting up a household can be shared between men and women (Jalovaara 2012).

In fertility research education has been identified as an underlying driver of economic effects on postponement of parenthood (Kreyenfeld 2010; Neels et al. 2013). In research on union transitions, however, educational differentials in micro- and macro-level economic effects have remained unaddressed (except for a similar study on the Netherlands, see De Lange et al. 2014). In this contribution it is hypothesised that educational attainment mediates the relationship between economic conditions and prospects at both the micro- and macro-level in union formation processes. On the one hand, educational levels are in general differently related to labour market positions and careers, and on the other hand the meaning and expectations that different socio-economic groups attach to living arrangements tend to vary (Clarkberg 1999; Nazio and Blossfeld 2003; Kiernan 2004). In first instance the focus in this study is on the transition to a first unmarried cohabitation in France. The incidence of cohabitation in this country is high as well as childbearing within cohabitation is comparatively common. Moreover, French co-residences are often long-lived and a large share of cohabitations are not dissolved by a marriage (Heuveline and Timberlake 2004). In this country, Heuveline and Timberlake (2004) have cohabitation ideal-typically identified as 'an alternative to marriage'. However, Hiekel and colleagues (2012) show that within France also a substantial heterogeneity of cohabitation types exists. Not only was marriage found to be an equal alternative for marriage, but also a considerable portion cohabits as a prelude to marriage or because their financial situation is not secure enough to marry. The latter type suggests that meanings of cohabitation are to some extent emerging from varying socio-economic positions.

Applying a longitudinal approach, we aim to examine whether there are differences in timing of entering a first unmarried union between different educational credentials with regard to male and female employment states. In similar vein, we study how aggregate-level economic conditions affect timing among different educations. This contribution thus addresses both the role of economic uncertainty at the individual level and the interplay between individual characteristics and macro-economic uncertainties. Furthermore, the same individual and cross-level interactions are tested for the transition from cohabitation to marriage as well.

2. Theoretical background

2.1 France

Union formation in France

In correspondence with many other northern and western European countries France has witnessed a sharp decrease in marriage rates after the early 1970s. This downward trend has persisted since 2000 after a short revival as a consequence of changed income tax rules in 1996 (Prioux 2005). This is related to an evolving delay in the entry of a first union and a massive shift in type of first union that took place between 1960s and 1990s. Throughout this period the share of couples starting their union as unmarried cohabitants dramatically increased from 10 to 90 per cent (Toulemon 1997). Further, France has become a forerunner in non-marital childbearing (Klüsener et al. 2013). About half of all children were born outside of marriage by 2005 (Beaumel et al. 2007). The vast majority of these births occurs within cohabiting unions (Kasearu and Kutsar 2011). However, in France most of the parents still marry at some point before or after giving birth (Perelli-Harris et al. 2012). Hence, policies that regulate the relationship between partners and children might influence the decision to marry. In 1999 France introduced a new form of civil union named 'Pacte civil de solidarité' (PACS). These registered partnerships distinguish cohabitation as a specific family institution while equipping these couples with comprehensive family rights and responsibilities (Barlow 2004). Nonetheless, in this contribution we consider French unregistered partnerships when referring to unmarried cohabitation. For both Pacs and unregistered partnerships regulations between parents and children (e.g. laws on paternity and custody) remain rather restricted. Regarding rights and obligations between partners (e.g. tax advantages, financial maintenance, division of household goods after divorce, alimony, etc.), on the contrary, registered partnerships are more harmonised with marriages. By and large the French government has shown reluctance to regulate unregistered co-residence on latter domains (Perelli-Harris and Sánchez Gassen 2012).

The economic situation of young adults

In France young adults are more often in unemployment or in school enrolment compared to other European countries (Chambaz 2001). During the 1990s and 2000s the French youth unemployment has been moderately higher than the European average (Eurostat 2014c). A more elaborated look at unemployment rates furthermore suggests that unemployment is differentiated by gender and educational level. Figure A1 in appendix presents French education-specific unemployment rates (1993-2008) for men and women separately between ages 15 and 39. The level of female unemployment is considerably higher in comparison with the male unemployment level. For both sexes a decreasing unemployment tendence is shown between the mid 1990s and 2008. It is however demonstrated that low educated workers are more prone to be unemployed, whereas the highest educatonal degrees show a lower unemployment risk.

The employment protection in France is found to be rather high (compared to other corporatist welfare state regimes like Belgium) (Breen 2005). Hence, French school-leavers are found to experience somewhat more delay in entering their first significant job (Wolbers 2007). As a result, young adults might extend the period of living with their parents. Besides the support of parents, people below 30 are often eligible to receive forms of welfare support as well (Chambaz 2001). In her comparative study - taking into account job incomes, state support and parental help - Chambaz (2001) however concludes that French young people have high living standards within a European context.

2.2 The uncertainty hypothesis and postponement of partnering

Transitions to adulthood have often been associated with the affordability clause (Rindfuss and VandenHeuvel 1990). In other words, the presence of sufficient (e.g. economic) resources for young people to gain the autonomy to take up an adult role in society. However, the establishment of a stable labour market position is often hampered for many school-leavers (Brzinsky-Fay 2007; Wolbers 2007). Decreasing income and poor employment prospects hinder young adults in proceeding to economic independence (Bell et al. 2007). According to Oppenheimer's (1988; 1997) uncertainty hypothesis a difficult career-entry process is expected to promote a delayed transition to marriage. This ambiguity experienced in early adult life is often associated with her concept of 'career immaturity' (Oppenheimer 2003). This career stage is defined as a period of financial dependency and uncertainty about individuals' long-term life styles and is characterised by stopgap jobs, low-paid jobs, regular unemployment and unfavourable labour perspectives. During the course of career immaturity partnering tends to be postponed as a result of lacking information concerning the individual's financial and social potential (Oppenheimer et al. 1997). Put differently, career immaturity strongly signals uncertainty regarding financial provisions and future life styles (Kalmijn 2011). Career maturity is reached if one's occupational situation and professional perspectives are considered as sufficient and lasting.

Uncertainty at the micro-level

A straightforward indicator of micro-level career immaturity is *school enrolment*. It typically competes with union and family formation and thus diminishes the hazards of entering a union (particularly marriage) and parenthood (Oppenheimer et al. 1997; Mills and Blossfeld 2005). Previous findings (in a variety of countries) have repeatedly indicated that full-time studies are likely to be incompatible with setting up an independent household (Coppola 2004; Hango and Le Bourdais 2007; Mulder et al. 2006; Winkler-Dworak and Toulemon 2007). Some authors have reported that the effect of enrolment is stronger on direct marriage in comparison with cohabitation (Baizán et al. 2003; Wiik 2009). The lack of material resources (Upchurch et al. 2001), incompatibility with adult family role activities (Marini 1985; Blossfeld and Huinink 1991) and uncertain future prospects (Mills 2000) account for the negative correlation between union and family formation processes and school career. It is additionally suggested that a prolonged stay in the educational system is another pathway to explain postponement of other transitions during economic decline. Poor labour market prospects may thus cause an increasing opportunity cost of school-leaving. Human capital investments might pay off as economic uncertainty emerges (Bucholz et al. 2009). In summary, an extended school career accounts for an important part of the explanation of a higher amount of career immaturity at young ages (Kohler et al. 2002; Ní Bhrolcháin and Beaujouan 2013).

In line with Oppenheimer's uncertainty hypothesis several studies have particularly found negative effects of (male) *unemployment* on marriage chances in different country settings (Kalmijn 2011). Considering unmarried cohabitation the negative association with unfavourable labour market positions are found to be less strong (Kalmijn and Luijkx 2005; Liefbroer 2005; Wiik 2009; Kalmijn 2011; Jalovaara 2012). For marriage it is argued that normative expectations on completed schooling, economic independence and material standards of living are relatively high in comparison with unmarried co-residence (Mäenpää 2009). However, in Sweden and France work promotes first union formation regardless whether it is cohabitation or marriage (Bracher and Santow 1998; Winkler-Dworak and Toulemon 2007). In a multi-national research project initiated by Blossfeld and colleagues (2005) a much wider variety of labour market uncertainty measures have equivalently shown the hampering impact of employment uncertainty on the hazards of partnerships and parenthood. This leads to the conclusion that besides the longer time spent in education, deferring long-term commitment decisions (i.e. partnering and childbearing) constitutes an additional adaptive strategy for economic uncertainty (Bucholz et al. 2009). Regarding the transition from cohabitation to marriage more mixed results are suggested (Bracher and Santow 1998; Francesconi and Golsch 2005; Kalmijn 2011; Kravdal 1999). Possible explanations for latter findings relate to relationship investments and changing partnership characteristics within cohabitation. Couples who have already invested a lot in their union (i.e. financial resources or commitment) might be more independent from labour force participation when they proceed to marriage. For couples who buy a house or enter parenthood economic uncertainty might even be independent from the decision to marry. Legal considerations on the division of household goods or custody after dissolution might trigger couples to marry as well (Perelli-Harris and Sánchez Gassen 2012). A last speculation addresses a changed economic situation for the partner. As the economic prospects of a partner have surged during cohabitation, this might elevate marriage risks (Kalmijn 2011). Following the large body of empirical evidence and theoretical considerations, we hypothesise that employment is a more important prerequisite for entering a first unmarried co-residence in comparison with marriage after cohabitation (H1).

Uncertainty at the macro-level

Özcan et al. (2010) have indicated that aggregate unemployment may affect individual behaviour, even if a person does not have an individual unemployment experience. Studies on fertility behaviour combining (un)employment at the individual-level and aggregate-level show that the macro-level unemployment effects hold while controlling for individual labour market status (Hoem 2000; Kravdal 2002). Both a changing risk perception or an average wage decrease resulting from deteriorating conditions may account for this (Kravdal 2002). Economic conditions are furthermore expected to affect job characteristics, especially those of young workers (OECD 2008; Verick 2009). In addition, the attention of young workers increasingly turns to their professional career since they have to work harder and longer to obtain a fixed, full-time contract (Mills and Blossfeld 2005). Although a great deal of evidence supports the hypothesis of a negative association between economic insecurity and fertility, delayed union formation has been suggested as one of the pathways through which economic downturn entails postponement of fertility (Sobotka et al. 2011).

Oppenheimer (1988) has identified economic recession as one of the factors affecting career development and impinge on the timing of marriage. In line with studies reporting an effect of aggregate-level unemployment on first birth hazards (Neels et al. 2013), unemployment rates are found to have a delaying effect (1970-2004) on first union formation in several countries (Vergauwen and Neels submitted). For France the postponement effect appears the strongest among men. In accordance, Prioux (2003) has provided evidence for a strong correlation between youth unemployment and first union formation between 1975 and 1998. The author concludes that economic context exerts a powerful influence on the timing of union formation in France. Moreover, the results report an absence of significant outcomes for type (cohabitation or marriage) of first union. Considering French cohabitants embarking on marriage, unemployment rates only affect the youngest age groups (Vergauwen and Neels submitted). This gives rise to the hypothesis that less prosperous macro-economic conditions are expected to entail a postponement effect (regardless of individual employment status) on entry of a first cohabitation (H2a), whereas the effects on marriage after cohabitation are predicted to be weaker (H2b).

2.3 The differential impact of economic insecurity by gender

To a great extent it is expected that gender differences in the effects of micro- and macro-level economic conditions depend on the symmetry of gender roles in a society (Thomson and Bernhardt 2010). Following Becker's New Home Economics (1981) couples should take advantage from the specialization of partners in different activities within a household. Hence, men traditionally play a financial provider role whereas women specialise in domestic production. Given Becker's specialisation model marriages are thought to function the most efficient in correspondence with these gender roles. Put differently, this theory predicts a positive effect of male employment and a negative one for female employment. In male-breadwinner societies more weight is hence attached to men's economic position as a prerequisite to union formation.

France has however experienced a steady increase in female labour force participation over the last fifty years (Leridon and Toulemon 1995). Between 1990 and 2008 the ratio of female to male shares in the labour force participation rates (ages 15-64) has even risen from 0.71 to 0.82 (ILO 2012). France has evolved to a more collaborative model in a dual-breadwinner society where predominantly both partners contribute financially to the household. French women appear to be relatively more employed in health and education professions (INSEE 2012). Regarding the effects of economic adversity at the macro-level Verick (2009) highlights that service and public sectors are less sensitive to economic context. In more gender-neutral societies opportunity costs for work are theorised to be smaller for women. As a result, women's economic resources may facilitate union formation as these ease the burden on men and the costs of setting up a household are more often shared. Therefore the attractiveness of self-sufficient women increases in egalitarian countries (Oppenheimer 1988; Bracher and Santow 1998; Jalovaara 2012). Some prior articles in the United States have reported delayed union formation among women as a consequence of being out of the labour force (Goldscheider and Waite 1986; McLaughlin et al. 1993; Oppenheimer 1994). Recent empirical studies in Europe and France lend support to this hypothesis as well. Kalmijn's (2011) contribution on the role of men's income and employment on cohabitation and marriage finds that the importance of men's economic position diminishes in gender-equal settings. Research on gender differences in France observes a convergence in the female and male effects of working status on union formation between 1955 and 1998. Comparing earlier periods to more recent periods, the working status entails a more important role for female entry into a union in France (Winkler-Dworak and Toulemon 2007). From the increasing importance of female financial contributions arises the hypothesis that micro- and macro-level economic uncertainty defers entry into a first union to a similar extent for men and women (H3a).

The question however emerges whether a difference exists between couples who choose for cohabitation instead of marriage. The large number of people cohabiting prior to marriage indicates that this group has become heterogeneous while it is suggested that people refraining from cohabitation before marriage are nowadays a select group (Liefbroer and Dourleijn 2006). In countries with high cohabitation rates couples who are selected in direct marriage increasingly share a particular religious or ethnic background (Sobotka 2008). In Europe religious people tend to reject the benefits and exclusivity of marriages as a form of partnership less (Pongrácz and Spéder 2008). Moreover, Kalmijn (2011) finds that the weakening effect of men's employment on first union is at least partly attributed to the fact that cohabitation is more widespread in countries with the most symmetric gender roles. It might be the case that the group marrying without pre-marital cohabitation is specifically prone to different attitudes regarding the gendered division of labour (Xie et al. 2003; Baxter 2005). Although a vast majority chooses to cohabit as a first union, still most people eventually marry or want to marry (Andersson and Philipov 2002; Bernhardt 2004). Research has shown that the division of labour is more egalitarian in cohabitation and that these couples' pre-marital cohabitation experiences contribute to a more equal sharing of housework in subsequent marriages. Women who cohabitated prior to marriage perform less household work and are more likely to do outdoor work compared to married women without prior cohabitation (Baxter 2005; Batalova and Cohen 2002). For women patterns established during the cohabitation stage are likely to be carried over to marriage. The transition to parenthood, which is often associated with the legalisation of a non-marital union, is however found to be crucial in the development of an unequal gap in the division of labour between men and women (Baxter et al. 2008). Nevertheless, we expect for the transition from cohabitation to marriage a similarity in the effects of male and female micro- and macro-level economic uncertainty (H3b).

2.4 The differential impact of economic insecurity by education

The impact of economic insecurities on partnership transitions is likely to vary by different educational groups. The differential results are expected to depend highly on both the extent that educational groups are diversely affected by economic uncertainty (Gangl 2002; Mills and Blossfeld 2005; Verick 2009) and whether the meaning different educational levels attach to consensual unions diverges (Easterlin et al. 1990; Clarkberg 1999; Nazio and Blossfeld 2003).

Under changing labour market circumstances (e.g. economic downturn), low education groups are considered as a vulnerable group. Youth lacking human capital are typically found in the most detrimental labour market positions during periods of economic recession (Müller and Shavit 1998; Verick 2009). Hence, the economic recession effect for the highest educated might be expected to be less strongly pronounced (De Lange et al. 2014). However, different mechanisms may simultaneously be at work. On the one hand, a crowding-out process possibly replaces less educated workers from their traditional labour market positions by better-educated applicants in times of poor economic prospects. Resulting from scarcity in labour market positions, the better educated thus start to compete for low education vacancies (Dolado et al. 2000). Although a high education is no full guarantee for protection against economic recession, workers with a high education often succeed to bridge adverse economic conditions by taking up inadequate work (Berkowitz King 2005; Francesconi and Golsch 2005). On the other hand, being more educated than usually required to perform a job adequately (i.e. being overeducated) has been suggested to yield harmful job characteristics (Verhaest and Van der Velden 2012). Literature states that overeducated workers tend to have lower wages, receive less formal training and report less job satisfaction (Hartog 2000; Verhaest and Omey 2006, 2009). Furthermore, Liefbroer and Corijn (1999) have reported that the high educated generally follow career tracks where earnings are strongly related to age and experience. Among this category adverse prospects may therefore lead to deteriorations of the economic situation relative to their aspirations. In other words, a temporary break from the labour market or a limited period of overeducation potentially drive postponement of household transitions until they are sufficiently established in their career track. This leads to the fourth hypothesis which predicts that the negative effects of macro-economic adversity and individual employment insecurity predominantly pertain to the highest educational attainments (H4). The better educated are thus expected to be more prone to an uncertainty hypothesis vis-à-vis union transition hazards.

Oppenheimer and colleagues (1988; 1997) have argued that the individual's educational attainment constitutes an important factor in the process of assortative mating. Education provides information on the individual's socio-economic position in the long run, a high level of education may therefore enhance first union formation (especially marriage). However, the evidence for this shows to be mixed (Kravdal 1999; Goldscheider et al. 2006; Kalmijn 2011). The least educated are likely to have different motivations to form a union as unmarried co-residence might account for an uncertainty-reduction strategy as well. This is predominantly the case for young adults who have left the parental home since it is reported for Europe that economic needs (at both the individual- and country-level) and a low education prolong co-residence between parents and adult children (Isengard and Szydlik 2012). The precarious labour market perspectives among the least educated reach a threshold during economic adversity where living together is thought to become more beneficial (Easterlin et al. 1990). A shift from single-living to non-marital cohabitation is equivalently considered as another strategy to cope with economic uncertainty (Easterlin et al. 1990; Nazio and Blossfeld 2003; Bucholz et al. 2009). In literature unmarried co-residence unions are often identified as flexible partnerships matching volatile and uncertain labour market circumstances (Mills and Blossfeld 2005). In many cases, unmarried co-residence is characterised by low dissolution costs, resources can be easily pooled and economies of scale are provided (Oppenheimer 1988; Liefbroer and Dourleijn 2006). In similar vein, Mclanahan (2004) has suggested that emerging adverse economic conditions in the last decades have entailed changes in the relationship between education on the one hand and union and family formation on the other. This implies the hypothesis that the most disadvantaged groups on the labour market might be attracted to form consensual unions as micro- or macro-level labour market prospects become disadvantageous. Particularly for the transition from single-living to cohabitation the uncertainty-reduction hypothesis predicts a less negative effect of economic adversity among the least educated (H5).

3. Data & methods

3.1 Data

The research questions considered in this contribution require individual information on both union and employment histories complemented with macro-economic data. The analyses use data from the French Harmonized Histories (1993-2005). The Harmonized Histories is a standardised comparative database providing retrospective information on transitions between living arrangements and parental home leaving constructed from different surveys in several European countries (Perelli-Harris et al. 2010). For France Harmonized Histories data are drawn from the Generations and Gender Survey (GGS)[2] (first wave). The retrospective design of this survey raises questions on the accuracy of collected event histories, as several sources of bias have been documented in the literature (Blossfeld and Rohwer 2002; Blossfeld et al. 2007). Analyses assessing the validity of demographic data in the GGS show that female first marriage rates calculated retrospectively from the GGS approximate vital statistics quite well for more recent periods (Vergauwen et al. forthcoming). Bias may result from errors in reporting the start and dissolution of consensual unions (Hayford and Morgan 2008), but the validity of retrospective data in this respect could not be assessed given the lack of vital registration data. In the second wave of the GGS retrospective full employment histories are questioned as well. In addition, the second wave provides an update of the French union histories between 2005 and 2008. Both union and employment histories of the different waves are in this paper considered to combine information on the timing of transitions between living arrangements and individual employment status.

3.2 Model specifications and dependent variables

For each eligible respondent (aged 16-39) a person-year-file is constructed to conduct individual union and employment histories between 1993 and 2008. We use multilevel discrete-time event history models in a competing framework to consider the effect of micro- and macro-level economic conditions on two types of events: 1) the interval between 16 year and the entry into an unmarried cohabitation and 2) the entry into a marriage following cohabitation (Allison 1982). These events are investigated for men and women separately. To exclude the pre-arrival experience of first generation immigrants and to tackle a possible link between union formation and immigration we observe migrants only after the first year of residence in their destination country (Bracher and Santow 1998). Pre-arrival and arrival person-years are therefore left out of the risk set.

In the first part of the analysis 16-year old never-partnered men and women are followed until (i) a first unmarried cohabitation (living together for at least three months) takes place, or (ii) a direct marriage occurs or (iii) single individuals are censored (last observation at age 39 or in year 2008). We draw on random-intercept multinomial logit models to contrast never-partnered person-years (i.e. years ‘at risk’) with events (i.e. first cohabitation, direct marriage or censoring). Calculations on the French datafile indicates that about 90% of first unions between 1993 and 2008 are unmarried cohabitations. To exclude associations between our variables of interest and direct marriage or censoring (i.e. disturbances), the latter are considered as a separate group of outcomes in the multinomial equation. This contribution spotlights the estimated coefficients contrasting singlehood years with a cohabitation event. In addition, person-years (level 1) are nested in persons (level 2). These models thus include a random frailty-coefficient to account for unobserved heterogeneity at the individual level (Mills 2011).

The second set of models considers men and women who have entered unmarried cohabitation. Only person-years in which men and women are living together (for the first time) with a partner without being married are considered to analyse whether the cohabitation is converted into a marriage. Respondents in cohabitation are censored after the age of 39 (or after the year 2008) or at the time when the cohabiting union dissolves. Similar to the first set of models censoring and union dissolution are regarded as a separate outcome in random-intercept multinomial logit models. Hence, the comparison of interest concerns the period of cohabitation and a marriage event.

[2] Part of the Generations and Gender Project (PAU, UNECE: http://www.ggp-i.org)

Random frailty-coefficients are included to control for time-constant unobserved characteristics at the individual level.

3.3 Independent variables

The multivariate hazard models for entry into first cohabitation include following micro-level covariates (time-constant and time-varying): age, educational attainment, duration since graduation, living at the parental home, calendar time, wave and employment status. The gender- and education-specific unemployment rates (1993-2008) are included as a macro-level covariate (Eurostat 2014b). The models addressing the transition from cohabitation to marriage take duration since entry into cohabitation as the baseline hazard function. Table 1 presents the distribution of person-years by individual-level time-constant and time-varying covariates.

Age. The models analysing the entry into the first union include a cubic effect of age (time-varying) as the baseline hazard function. The age variables are centred at 16 years. The models analysing the conversion of cohabitation into marriage include age at entry in cohabitation as a time-constant variable (linear and quadratic effects).

Duration since entry into cohabitation. The model analysing the conversion of cohabitation into marriage use duration of the consensual union (time-varying) as the baseline hazard function (linear and quadratic specifications).

Educational attainment. The highest reached educational degrees (time-constant) of the second questioning round (GGS Wave 2) are considered in the analyses. Based on the ISCED-classification (1997) three levels of education are distinguished (Perelli-Harris et al. 2010). The 'high' category serves as the reference category here.

Duration since graduation. As noted earlier, school graduation increases the risk to experience several transitions to adulthood dramatically. The age of ending studies is consequently considered to be crucial in the timing of several events. In addition, Skirbekk and colleagues (2004) find that the time elapsed since leaving school has also a persistent influence on the timing of these events. This study points out that social age (i.e. age of the individual's school cohort) is of a particular relevance for demographic behaviour. Due to social interactions and peer-group influences of fellow school cohort members, people with comparable social age are expected to cluster event timings in terms of duration since graduation (rather than biological age). Notice that social age may differ from biological age due to differences in the length of the school careers or the month of birth (see Skirbekk et al. 2004). To construct this variable we used answers on the question about the year of reaching the highest educational degree. All person-years in education are represented in this variable by value 0. Subsequently, the clock starts on the first year after a person ended his or her studies. To capture the non-linear relationship between partnership transitions and elapsed time since graduation, the association is specified as a cubic effect.

Calendar time. In accordance with developments in other European countries an ongoing delayed entry into a first union and marriage is reported for France (Prioux 2003; Sobotka and Toulemon 2008). The hazards of entering a first union or marriage after cohabitation are therefore likely to vary throughout the observation period due to postponement independent from prolonged education or economic factors. Hence, a time indicator interacted with age (quadratic effect) is included that controls for unmeasured processes that render postponement effects.

Parental home leaving status. Previous studies have indicated that the co-residence of adult children with their parents is closely related to economic necessity on the children's part. Aassve and

Table 1. Distribution of person-years for time-constant and time-varying covariates, women and men aged 16-39, 1993-2008

	From single to first union		From cohabitation to marriage	
	Men	Women	Men	Women
Age				
16-21	0.4150	0.5190	0.0710	0.1279
22-27	0.3356	0.2901	0.3844	0.4416
28-33	0.1604	0.1154	0.3500	0.2790
34-39	0.0890	0.0754	0.1946	0.1515
Age at cohabitation				
16-21			0.2820	0.4534
22-27			0.5095	0.4549
28-33			0.1881	0.0722
34-39			0.0204	0.0194
Duration since cohabitation				
0-2 years			0.4193	0.4045
3-5 years			0.2596	0.2510
6-10 years			0.2297	0.2065
11-15 years			0.0767	0.1016
> 15 years			0.0147	0.0364
Educational attainment				
Low	0.1280	0.0809	0.1507	0.1312
Middle	0.4602	0.3491	0.5301	0.4116
High	0.4118	0.5700	0.3191	0.4571
Duration since graduation				
In education	0.3663	0.5171	0.0643	0.1884
1-2 years	0.1424	0.1415	0.0708	0.1187
3-5 years	0.1681	0.1222	0.1716	0.1977
6-10 years	0.1546	0.1016	0.2937	0.2457
11-15 years	0.0872	0.0614	0.2245	0.1458
> 15 years	0.0815	0.0562	0.1751	0.1037
Calendar time (year)				
1993-1996	0.3523	0.3158	0.2718	0.2586
1997-2000	0.2966	0.2885	0.2743	0.2687
2001-2004	0.2200	0.2600	0.2524	0.2501
2005-2008	0.1311	0.1357	0.2015	0.2226
Living at parental home				
No	0.4686	0.5022		
Yes	0.5314	0.4978		
Employment status				
Student	0.3690	0.5138	0.0595	0.1793
Employed	0.5641	0.4224	0.8986	0.6990
Inactive	0.0612	0.0470	0.0394	0.0622
Homemaker	0.0057	0.0168	0.0025	0.0595
N respondents	837	1,138	782	1,111
N person-years	5,978	7,493	4,014	5,913
N censored	168	188	247	362
N of cohabit.	616	861		
N of marriages	(53)	(89)	294	425
N of separations			(241)	(324)

Source: French Harmonized Histories and French GGS Wave 2, calculations by author colleagues (2002) show that employment and income are among the important determinants of the decision to leave the parental home. A recent study by Isengard and Szydlik (2012) confirms that employment status, reflecting the need structure of an adult child, is associated with prolonged co-residence with parents in Europe. One possibility is that for those living away from the parental home economic uncertainty might facilitate cohabitation (Easterlin et al. 1990), whereas for adults living at home adverse conditions postpone transitions to adulthood. To control for the relationship between employment status and co-residence of adult children with their parents a parental home leaving status is included in the models assessing the transition from singlehood to first unmarried cohabitation.

Employment status. The time-varying information GGS Wave 2 provides on employment status is collapsed to a categorical variable reflecting three states: student, employed and inactive. For cohabiting women a category 'homemaker' is retained as well. Regarding the transition from never-partnered to first union we observe that a majority of the person-years represent the employed (encompasses employed, self-employed, in military service or working as a family member in a family business) or student (or in apprenticeship) status (93.31% and 93.62% among men and women respectively). For men 4.10% of the single person-years consists of unemployment while this proportion amounts 3.86% for women. Regarding cohabiters we find that 89.86% of the male person-years depict the employed status, whereas this proportion equals 69.90% amidst women. Further we observe that a substantial part of cohabiting women experiences years (5.95%) of being homemaker. 3.24% and 5.28% of unemployed person-years are reported for men and women respectively. The inactive category consists further of minor shares of ill or disabled person-years and years in the 'other' status. To examine differential effects of economic (in)security by educational credential, interaction effects are considered between education and employment status.

Wave. In the GGS design solely the original sample of individuals surveyed within the first wave is considered within the follow-up waves. As a result the original sample is three years older in the second wave (21-82 years old) compared with the first (18-79 years old) (Simard and Franklin 2005). Given this sample design it is expected that, due to the changing age structure, an artificial increase in the proportion of transitions of interest (i.e. young adults forming co-residences) will occur between both waves (2005-2008). Hence, we estimate a dichotomous variable controlling for changing transitions hazards between the first and second wave.

Unemployment rates. The individual-level data from the Harmonized Histories are complemented with aggregate-level time-series data on unemployment rates (15-39) between 1993 and 2005 drawn from Eurostat (2014b). Figure A1 in appendix presents unemployment rates by gender and education. The annual unemployment rates are obtained by results from the European Labour Force Study (EU-LFS) and calculated as a percentage of the civilian labour force (by age, sex and education) (Eurostat 2014a). Macro-level unemployment rates have been frequently used as an indicator of economic context in research on union and family formation. A review of the literature indicates that unemployment and consumer sentiment reflect the effect of economic recession on family formation more closely than general indicators as GDP or inflation rate (Sobotka et al. 2011). Research has provided empirical support for significant positive correlations between individual perceived job insecurity and aggregate unemployment rates (Green et al. 2000; Green 2003). The labour market state is thus suggested to have clear implications for the assessment of the individual employment situation and has psychological consequences for those who are employed (Erlinghagen 2008). To examine whether the impact of macro-level unemployment on union formation hazards varies by education, interaction effects between unemployment rates and education are tested.

4. Results[3]

4.1 Transition into first cohabitation

The results of the event-history models for entering an unmarried co-residential union are presented in Table 2. The full results for Model 1 and 3 are shown in appendix (Table A1). The cubic effect of duration since graduation shows that hazards of entering a cohabiting union increase after finishing school to reach a peak and subsequently decline. Although this schedule looks similar for both genders, the results suggest less strong effects among women. Net of the duration since graduation effect the model predicts a strong association between partnering hazards and age as well. We observe that cohabitation hazards peak earlier for women in comparison with men. This indicates that women have earlier timetables for union formation in terms of age. Model 1 further demonstrates increasing first union hazards among men and women throughout the observation period (i.e. the calendar time effect). For men this trend is partly explained by macro-economic context and the second wave sample design (Model 3). The female increase, however, remains largely unexplained. The insignificant interaction terms between age and calendar time show that this trend effect is not attributed to a particular age. Controlling for the differential timing of leaving the educational system and employment status, first cohabitation hazards do not diverge substantially by educational attainment. In addition parental home leaving exerts a differential effect on the entry into first cohabitation between men and women in France. For men having left the parental home leaving facilitates union formation ($p < 0.050$), whereas for women we find an absence of significant differences between co-residence with parents and independent living. In other words, women often enter a first union directly from the parental home. Men partner more frequently after a period of single living.

Model 1 furthermore shows that employment strongly enhances the transition to first cohabitation among men. Being a student typically delays union formation. This effect is however partly explained by the duration since graduation polynomial where being in education is represented as well. Being inactive at the labour market negatively affects the likelihood to start a first cohabitation. In terms of the model estimates first cohabitation hazards for inactive men are less than half of the employed men's hazards. Contrastingly, we find that for women being inactive renders no significant differences with being employed regarding the risk of entering a first cohabitation. A strong negative effect for the student status is observed for women as well. While female education enrolment defers union formation, female (individual) employment insecurity does not postpone union formation. To verify whether employment status affects cohabiting hazards differently between diverging educational levels, interactions terms between both variables are included in Model 2 (Table 2). For men it appears that employment is particularly important for the transition to a first cohabitation among the highest and intermediate educational levels. The model predicts that the risks of entering a first cohabitation are more than two and three times higher for employed compared to inactive men among the middle and highest educated respectively. Both effects are significant at the 99% confidence level. For men with a low education the model reports that the hazards of entering a first cohabitation are not significantly different between an inactive and employment status. Put differently, for this group men in uncertain labour market positions and employed men tend to have equal partnering risks. The female equivalent of this model suggests that being employed yields no significantly different impact on first cohabitation hazards from being

[3] As all models are estimated drawing on quasi-likelihood methods, model diagnostics based on log-likelihood outcomes are absent.

inactive among all educational credentials. Although inactive women have higher partnering risks for women with a low and middle education, differences remain insignificant at the 95% confidence level. In addition, studying delays union formation significantly more than being inactive for the lowest and highest education groups ($p < 0.050$). For women with an intermediate education this difference is borderline significant ($p = 0.057$).

Model 3 shows that for men aggregate unemployment impedes the likelihood to start a first cohabitation as well. This effect is rather substantial. A 1 percentage point increase in unemployment rate reduces the male first union hazards by 12% ((1 – 0.88) x 100). Sensitivity analysis indicates that models without employment status render identical effects for unemployment rate. Hence, macro-economic uncertainties are not channeled through individual employment status. No significant relationship for unemployment rates is observed for women. Our findings thus provide no evidence for a substantial effect of macro-economic context on women's transition to first cohabitation. By allowing variation of the unemployment effect by education (Model 4), we find that the male effect is particularly articulated among the highest educations. Notwithstanding that the difference between male high and low educated is not significant at the 95% confidence level, the magnitude of

Table 2. Odds ratios for entering an unmarried co-residential union among never-partnered singles (Model 2-4), men and women aged 16-39 in France (1993-2008)

	Model 2				**Model 3**				**Model 4**			
	Men		Women		Men		Women		Men		Women	
	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig
					Micro-level covariates							
Employment status												
Student					0.65	*	0.61	***	0.64	*	0.61	***
Employed					1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*
Inactive					0.43	***	1.17	-	0.43	***	1.17	-
Low education*												
Student	1.55	-	0.29	*								
Employed	1.34	-	0.61	-								
Inactive	1.00	*Ref.*	1.00	*Ref.*								
Middle education*												
Student	1.44	-	0.63	-								
Employed	2.42	**	0.83	-								
Inactive	1.00	*Ref.*	1.00	*Ref.*								
High education*												
Student	2.21	-	0.53	*								
Employed	3.37	**	0.99	-								
Inactive	1.00	*Ref.*	1.00	*Ref.*								
					Macro-level covariates							
Unemployment rates												
UR (15-39)					0.88	**	0.99	-	0.78	***	0.95	-
Unemployment rates*												
Low education									1.17	-	1.08	-
Middle education									1.20	*	1.07	-
High education									1.00	*Ref.*	1.00	*Ref.*

Source: French Harmonized Histories and French GGS Wave 2, calculations by author
Significance levels: not significant (-), $p < 0.050$ (*), $p < 0.010$ (**), $p < 0.001$ (***)
Models controlled for duration since graduation, age, leaving of the parental home, calendar time(years)*age, educational attainment and wave (models 3 and 4)

the effect is much smaller for men with a low education (e(b) = 0.91 = 0.78 x 1.17). Between a high and middle educational level (e(b) = 0.94 = 0.78 x 1.20) the unemployment association diverges significantly ($p < 0.050$). For women the effect varies in the same direction, differences are however smaller and insignificant.

4.1 Transition into first marriage following cohabitation

A summary of the full results for the models regarding the transition from cohabitation to marriage is provided in Table A1 (Model 5 and 7). In correspondence with the results for the first cohabitation transition, duration since graduation variables indicate significant effects on post-cohabitation marriage hazards. The combined effects of the linear, quadratic and cubic specified variables indicate that marriage hazards first increase rapidly and decline the longer graduation is finished. Our findings identify a similar relationship with duration since graduation. This implies that cohabitations are often short-lived before converted into a marriage. Longstanding cohabitations experience smaller marriage risks. In addition, age at cohabitation is not significantly related to marriage hazards for either men and women. Among both genders positive educational gradients are furthermore reported in Model 5. Net of the other effects, cohabitants with a high education show higher risks to enter a marriage. However, these educational differentials are not significant. The model further displays a negative effect of calendar time and significant interaction terms between a quadratic specification of age and calendar time. Pertaining to men an elaboration of these estimations points out that interaction effects compensate for the negative effect of calendar time. In other words, we find neither a general trend effect marriage after cohabitation hazards nor changing age effects through time. Nevertheless, amidst women we find that particularly the youngest ages defer marriage more significantly throughout the observation period.

Concerning employment status we only consider two categories for the male equivalent of this set of models. Men are either regarded employed or inactive. Less than 6% of cohabiting male person-years are spent in education between the age of 16 and 39 (Table 1). Furthermore, only three marriage events materialise while being student and in cohabitation. The results of Model 5 imply that the marriage hazards of inactive co-residing men are about half of employed men's hazards (Table A1). This effect is insignificant at the 95% confidence level ($p = 0.087$). In contrast to their male counterparts, a substantial proportion of women is still enrolled while cohabiting with a partner (Table 1). In tandem with weaker effects for duration since graduation and earlier partnering timings in terms of age (cfr. first cohabitation models), this suggests that young women often combine enrolment in education with unmarried co-residence. In Model 5 the category of homemaker is retained as a separate category for women. A small 6% of female person-years in cohabitation represents the homemaker status. These women appear to have a substantial risk to experience a transition from cohabitation to marriage. The same finding is reported for inactive cohabiting women. Both experience higher transition hazards in comparison with their employed counterparts. Differences are however moderate (34% and 27% respectively) and statistically insignificant. In correspondence with findings for the transition to first union, we find some (weaker) additional evidence for the proposition that particularly for men individual-level employment security is important to make transitions between living arrangements. Model 6 allows the effects of employment status to vary over educational levels. For men with a high education employment is identified to be an important prerequisite to proceed to marriage. Their marriage risk is estimated more than four times higher compared to inactive cohabiting men with a high education ($p < 0.050$). For men with a medium education employment doubles marriage hazards. This effect remains however insignificant at the 95% confidence level. Among the least educated employment entails a relatively negative effect on marriage hazards. The difference with the inactive category is nevertheless limited and insignificant. For women we find higher marriage hazards for inactive labour market positions over all educational degrees. The negative effect of employment relative to inactivity is even the strongest for women with a high educational attainment (marriage hazards 37% lower, p = 0.087). Students are furthermore observed to have (insignificantly) diminished marriage risks among all educational levels. Particularly among men we have to be cautious with interpreting the interaction effects as interacting inactivity and education results in small numbers of cases and events in this set of models.

In line with the outcomes for individual-level economic insecurity the impact of aggregate-level economic context remains rather constrained for partners who marry after a period of unmarried co-residence. Increasing unemployment rates involve a small negative impact on the marriage hazards of cohabiting men (Model 7). A 1 percentage point increase in unemployment rate reduces the male marriage hazards by 4% ((1 – 0.96) x 100). Model 8 indicates that the context effect is the strongest

Table 3. Odds ratios for entering a marriage among cohabitants (Model 6-8), men and women aged 16-39 in France (1993-2008)

	Model 6				**Model 7**				**Model 8**			
	Men		Women		Men		Women		Men		Women	
	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig
Micro-level covariates												
Employment status												
Student							0.86	-			0.86	-
Employed					1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*
Inactive					0.53	-	1.26	-	0.54	-	1.27	-
Homemaker							1.33	-			1.35	-
Low education*												
Student			0.24	-								
Employed	0.73	-	0.69	-								
Inactive	1.00	*Ref.*	1.00	*Ref.*								
Middle education*												
Student			0.78	-								
Employed	1.97	-	0.94	-								
Inactive	1.00	*Ref.*	1.00	*Ref.*								
High education*												
Student			0.57	-								
Employed	4.42	*	0.63	-								
Inactive	1.00	*Ref.*	1.00	*Ref.*								
Macro-level covariates												
Unemployment rates												
UR (15-39)					0.96	-	1.07	-	0.88	-	1.10	-
Unemployment rates*												
Low education									1.22	-	0.84	*
Middle education									1.10	-	1.02	-
High education									1.00	*Ref.*	1.00	*Ref.*

Source: French Harmonized Histories and French GGS Wave 2, calculations by author
Significance levels: not significant (-), p < 0.050 (*), p < 0.010 (**), p < 0.001 (***)
Models controlled for duration since graduation, duration since cohabitation, age at cohabitation, calendar time(years)*age, educational attainment and wave (models 7 and 8)

for the highest educated. Nonetheless, the differential impact for the least and middle educated are statistically insignificant ($p > 0.050$). For cohabiting women an adverse economic context facilitates a marriage transition (Model 7). A 1 percentage point increase in unemployment rate increases the female marriage hazards by 7% ((1.07 – 1.00) x 100). The p-value of 0.084 is considered on the borderline of statistical significance. This positive relationship between marriage hazards and economic uncertainty is predominantly situated among intermediate and higher educational levels (Model 8). For their low educated counterparts aggregate-level unemployment yields comparatively ($p < 0.050$) a more negative effect ($e(b) = 0.92 = 1.10*0.84$).

5. Discussion

5.1 Answering our research hypotheses

Previous studies have shown that economic uncertainty at different levels affects the likelihood to make partnership transitions. This contribution aimed to examine to what extent timing of transitions between living arrangements are differently related by educational level to micro- and macro-level economic circumstances. In addition, some hypotheses addressed differences in the impact of economic insecurities on male and female transition hazards and the varying role of economic conditions in different kind of transitions.

In first instance we find some support for our hypothesis that the timing of first cohabitation is more strongly related to individual-level employment status than marriage with pre-marital cohabitation (H1). At least, this finding prevails among men. Educational enrolment and an inactive position on the labour market puts off the entry into a first cohabitation. Employment thus constitutes an important prerequisite for this transition. Although similar results apply for the transition from cohabitation to marriage, the magnitude of these effects remains smaller. In accordance, we find that deteriorating macro-level economic conditions affect the male first union hazards negatively throughout the period considered (H2a). An adverse economic context even impedes the first cohabitation transition regardless of the employment situation of men. In other words, our results show that the impact of aggregate-level unemployment is not channelled through employment status. Macro-level economic conditions exert, however, relatively weak impacts on the conversion of cohabitation into marriage (H2b). Our analyses provide also evidence for the fourth hypothesis (H4) which accounts for a necessary elaboration of the first and second hypotheses. Particularly amidst men with a high educational degree economic uncertainty entails a postponement strategy vis-à-vis union formation (Bucholz et al. 2009). Career maturity and individual economic outlook appear to be of particular significance for men with a more extended education. Pertaining to men with a low education the results suggest that co-residing as a first union might be increasingly common within inactive labour market positions (H5). For this group forming a union characterised by a flexible nature is part of a range of rational responses to economic insecurity (Mills and Blossfeld 2005). With regard to the interaction between education and the aggregate-level economic effects our outcomes confirm these conclusions only partly. The economic context-induced postponement of first union is the strongest for the more educated although differences lack statistical power. Counter-intuitively, the effect employment renders on marriage following cohabitation diverges similarly by education compared to the results for the transition from singlehood to first cohabitation. Hence, economic security predominantly facilitates the marriage transition for men with a high education, whereas economic insecurity yields similar transition hazards among their low educated counterparts. Also it is demonstrated that economic adversity at the macro-level does not differ significantly by educational credential. As mentioned earlier the estimates of both the micro- and cross-level interaction terms might be less stable and reliable since a low amount of cases weakens statistical power.

For women our findings advocate that employment plays a different role in the transition to a first cohabitation. Notwithstanding the developments towards more gender symmetry with respect to financial contributions to - and the role divisions within the household in France (Winkler-Dworak and Toulemon 2007), the impact of employment on female first cohabitation is comparable to inactivity. Correspondingly, macro-level economic insecurity exerts negligible postponement effects on the entry into a first union. As a result, the hypothesis predicting a gender similarity in the outcomes of economic (in)security on first cohabitation is discarded (H3a). With regard to the transition from cohabitation to marriage mainly equivalent conclusions can be drawn. In terms of transition risks homemakers and inactive women do no differ substantially from employed women. Yet, having no job (except for studying women) somewhat increases the likelihood of entering a marriage. The same applies for macro-level economic adversity. Under a less prosperous economic context a limited surge in marriage hazards has been observed. In summary, the first two hypotheses (H1, H2a and H2b) are not supported for women. The same goes for the second part of the third hypothesis (H3b). Despite the weaker impact of employment on marriage among men, associations between marriage and economic uncertainty tend to run in opposite directions for men and women. Further, neither the outcomes for entry into a first cohabitation nor the results for marriage with pre-marital cohabitation show compelling differences of employment status by education (H4 and H5). Being employed is for the better educated fairly more important with regard to the first cohabitation transition whereas being inactive enhances the marriage transition in this group. Concerning the cross-level interactions the positive relationship between a deteriorating economic context and marriage is principally attributed to intermediate and high educations.

5.2 Limitations

To conclude the discussion of our results we address some important limitations of this paper. First, gender differences in the effect of economic uncertainty on union formation hazards are likely to remain somewhat ambiguous as long as we focus on individual men and women. Addressing individuals without taking both the man's and the woman's economic position into account casts doubt on whether economic security is important on either one or both sides. From a couple perspective both male and female's labour market positions could be investigated simultaneously (Kalmijn 2011). It could be for instance hypothesised that it is the accumulation of two insecure employment states within a couple that hampers union and family formation (De Lange et al. 2014).

A substantial amount of previous studies have looked in the role of more qualitative job indicators on partnership transition and parenthood (Francesconi and Golsch 2005; De Lange et al. 2014; Kieffer et al. 2005; Mills et al. 2005). Information on part-time work, duration of contracts, wages, the sector of jobs, employment history, etc. should likewise be taken into account considering the impact of economic insecurity on union formation. One possibility might be that a majority of jobs among the men with a low education are only temporary. This could (at least partly) explain the similar partnership transition hazards for low educated employed and inactive men. Hence, extending the information our variable of interest (i.e. employment status) provides is necessary in future research. To estimate similar interaction effects, however, the analyses require enlarged sample sizes as well.

The presence of children possibly modifies the net effect on transitions between living arrangements. Particularly for women the interaction between parenthood (e.g. during pregnancy) and employment status might be pivotal regarding the transition from cohabitation to marriage. Given that partnership transitions and childbearing are part of the same family formation process (Baizán et al. 2003), future research should examine the interplay of economic effects on union and family formation simultaneously.

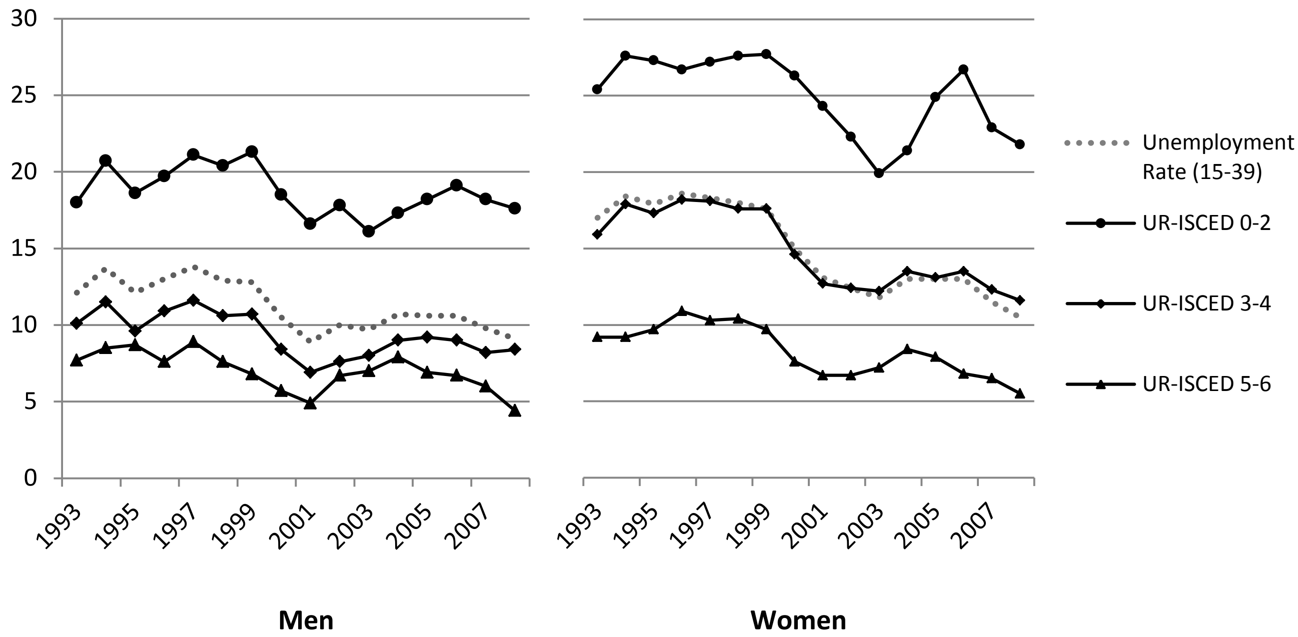
Finally, institutional factors have not been considered in the analysis, although they may equally be relevant to account for a mediating effect of social policy on economic insecurity and union formation.

6. Conclusion

The main finding in this contribution is that economic conditions particularly determine transition hazards of men with a high education. We provide support that unfavourable labour market positions are considered as an inappropriate setting for union formation in this group. Adverse prospects may strongly contradict the aspirations of better-educated men. Therefore, uncertainty-induced postponement of adulthood markers holds especially for this category (Mills et al. 2005). Additionally, it is an important outcome that the negative effect of macro-economic uncertainty establishes independently from individual employment. Put differently, the macro-economic association with first union entrance holds when controlling for time-varying employment status (De Lange et al. 2014). For the least educated we find that employment is less decisive for partnership transitions. Not only might the flexibility of a non-marital union attract this group to cope with economic uncertainty, but also are marriage risks alike for inactive and employed cohabiting men with a low education. Latter finding needs further investigation by the means of qualitative job indicators and an analytical sample size of unmarried cohabitants. Contrary to the expectations our results suggest particular gender-specific effects. The non-participation of women in the French labour market disappears and the traditional family model in France is clearly fading. Nevertheless, the idea that female employment is a necessity to enter a first union is opposed. Also, female union transitions are not postponed if the economic circumstances are not prosperous. Our results are however in line with previous French research reporting an exaggerated negative impact of employment status on partnership formation among men and an increased likelihood of entering a union (and parenthood) for very low paid female jobs and female inactivity (Kieffer et al. 2005). France's public family policies including pro-natalistic policies and measures to promote female employment might change these reversed relationships gradually (Blossfeld 1995).

7. Appendix

Figure A1. Yearly unemployment rates in France (15-39) by highest education and sex (1993-2008)



Source: Eurostat (2014b)

Table A1. Odds ratios for entering a co-residential union among never-partnered singles (Models 1 and 3) and for entering a marriage among cohabitants (Model 5 and 7), men and women aged 16-39 in France (1993-2008)

	Model 1				**Model 3**				**Model 5**			
	Men		Women		Men		Women		Men		Women	
	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig	e(b)	Sig
Duration since graduation												
Linear	1.24	**	1.17	*	1.23	**	1.18	*	1.39	**	1.25	*
Quadratic	0.98	**	0.98	-	0.98	**	0.98	*	0.95	***	0.96	***
Cubic	1.00	*	1.00	-	1.00	*	1.00	-	1.00	***	1.00	***
Age (centered around 16)												
Linear	2.12	***	2.43	***	2.11	***	2.41	***				
Quadratic	0.94	***	0.91	***	0.94	**	0.92	***				
Cubic	1.00	***	1.00	***	1.00	***	1.00	***				
Duration since cohabitation												
Linear									1.41	***	1.26	***
Quadratic									0.98	***	0.98	***
Age at cohabitation												
Linear									1.26	-	0.95	-
Quadratic									0.99	-	1.00	-
Educational attainment												
Low	0.92	-	0.90	-	3.85	**	0.91	-	0.81	-	0.72	-
Middle	1.11	-	1.12	-	1.47	*	1.12	-	0.98	-	0.85	-
High	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*
Living at parental home												
Yes	1.27	*	0.92	-	1.27	*	0.93	-				
No	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*				
Calendar time												
Year	1.13	*	1.10	***	1.08	-	1.10	**	0.73	**	0.78	***
Calendar time* age												
Linear	0.99	-	0.99	-	0.99	-	0.99	-	1.04	**	1.05	***
Quadratic	1.00	-	1.00	-	1.00	-	1.00	-	0.99	*	0.99	***
Employment status												
Student	0.66	*	0.60	***	0.65	*	0.61	***			0.85	-
Employed	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*	1.00	*Ref.*
Inactive	0.44	***	1.20	-	0.43	***	1.17	-	0.54	-	1.27	-
Homemaker											1.34	-
Wave												
1					1.00	*Ref.*	1.00	*Ref.*				
2					2.06	***	1.90	***				
Unemployment rates												
UR (15-39)					0.88	**	0.99	-				
N person-years	5,978		7,493		5,978		7,493		4,014		5,913	
N respondents	837		1,138		837		1,138		782		1,111	
N events	616		861		616		861		294		425	
Sigma u	0.0264		0.0773		0.0000		0.0369		0.0000		0.0603	

Source: French Harmonized Histories and French GGS Wave 2, calculations by author

Significance levels: not significant (-), p < 0.050 (*), p < 0.010 (**), p < 0.001 (***)

Table A2. Continued

	Model 7			
	Men		Women	
	e(b)	Sig	e(b)	Sig
Duration since graduation				
Linear	1.38	**	1.25	*
Quardratic	0.95	***	0.96	***
Cubic	1.00	***	1.00	***
Age (centered around 16)				
Linear				
Quadratic				
Cubic				
Duration since cohabitation				
Linear	1.41	***	1.26	***
Quadratic	0.98	***	0.98	***
Age at cohabitation				
Linear	1.25	-	0.96	-
Quadratic	0.99	-	1.00	-
Educational attainment				
Low	1.32	-	0.22	*
Middle	1.08	-	0.61	-
High	1.00	*Ref.*	1.00	*Ref.*
Living at parental home				
Yes				
No				
Calendar time				
Year	0.73	**	0.79	***
Calendar time*age				
Linear	1.05	**	1.04	***
Quadratic	0.99	*	0.99	***
Employment status				
Student			0.86	-
Employed	1.00	*Ref.*	1.00	*Ref.*
Inactive	0.53	-	1.26	-
Homemaker			1.33	-
Wave				
1	1.00	*Ref.*	1.00	*Ref.*
2	0.85	-	1.19	-
Unemployment rates				
UR (15-39)	0.96	-	1.07	-
N person-years	4,014		5,913	
N respondents	782		1,111	
N events	294		425	
Sigma u	0.0006		0.0572	

Source: French Harmonized Histories and French GGS Wave 2, calculations by author

Significance levels: not significant (-), p < 0.050 (*), p < 0.010 (**), p < 0.001 (***)

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