Fertility trends and its determinants in Spain and Europe

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CHAPTER 5

Does Union Formation Change Attitudes towards Childbearing in Bulgaria? A Propensity Score Analysis.

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Abstract

This paper explores how attitudes towards childbearing change after cohabitation and marriage. Entering in a union may affect attitudes towards childbearing, but at the same time individuals who are more oriented towards childbearing may be more determined to form unions. In order to disentangle the interplay between union formation and subsequent changes in attitudes towards childbearing we implement propensity score matching, which is applied to a panel data coming from a survey in Bulgaria to identify the effect of union formation of attitudes related to childbearing. This method controls for observable confounders which may affect both the probability of entering into a union and childbearing preferences, reducing selection bias. We find a positive significant causal relationship of entering into union on attitudes towards parenthood among men, whereas the effect is weaker and often uncertain for women.

5.1. Introduction

Union transition is a pivotal event in an individual's life and represents an important first step towards family formation. The literature points to transition to union formation as one of the strongest predictors for parenthood (Rindfuss & Van den Heuvel, 1990; Thomson, 1997; Thornton, Axinn, & Teachman, 1995).

Since union formation is often followed by the birth of a child, an interesting and under investigated point in the literature consists in addressing whether union formation activates a change in individuals attitudes that is supposed to lead to the decision of planning to have a child. The formation of positive attitudes towards childbearing is considered in the literature as a premise for the formation of positive intentions of giving birth to a child. Hence, mapping changes attitudes towards parenthood is a way to establish the main determinants of fertility.[7]

The aim of this paper is, thus, to verify whether union formation affects attitudes towards childbearing, intended as a predictor of parenthood. This is done using panel data that collects information on attitude changes and life course paths between 2002 and 2005 of a sample of young Bulgarian men and women aged between 18–34 year old. In this framework, the main problem in estimation is that the identification of the effect of union formation on attitudes towards childbearing may be influenced by selection bias and reverse causality. The first refers to the fact that there could be other factors, such as in this case, risk of poverty, job position and age, which affect both the choice of a marital or non-marital cohabitation and having a child (Aassve & Lappegård, 2008). In this case, the individuals in the sample who decided to marry or cohabit with their partners were also more likely to have previous positive attitudes towards childbearing, leading to a potential underestimation or overestimation of the causal effect among them. The second issue regards the fact that it is necessary to disentangle causal effects, by understanding whether individuals enter into a union because their beliefs are already oriented towards parenthood, or whether union formation is a predictor in a process that activates a change in individuals' attitudes towards parenthood. The literature, for example, gives evidence that positive attitudes towards childbearing also predict the choice of entering into a union (Barber, Axinn, & Thornton, 2002).

7 This scheme has been applied to the choice of giving birth to a child as described by Ajzen & Klobas (2013) and as applied by Billari, Philipov & Testa (2009) with respect to the Bulgarian case.

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In order to disentangle the causal interplay between union formation and attitudes towards parenthood and to quantify the effect of union entry on attitudes towards childbearing, we combine propensity score (Rosenbaum & Rubin, 1983) and difference-in-difference estimation. Propensity score matching is widely used in causal inference and in program evaluation practice in order to estimate the effect of a binary treatment variable (union formation) on a continuous response outcome (attitudes towards parenthood), giving the opportunity to reduce bias due to observable confounding factors which may affect both the probability of entering a union and the level of attitudes. This purpose is realised through the comparison of individuals with a similar probability of union entry based on a matrix of observed determinants. Nevertheless, propensity score matching fails in the presence of unobserved confounders associated with both union entrance and attitudes towards childbearing: we refer to a set of variables such as values, personality traits, social norms and other beliefs which are not fully captured in the set of observed determinants. The presence of unobserved characteristics that simultaneously explain the union formation and the attitudes towards childbearing produces biased estimates, either underestimating or overestimating the causal effects of the union entry on the expected benefits and costs of giving birth to a child. In this case a difference-in-difference estimator combined with propensity score allows for neutralising the effect of unobserved factors if they are timeinvariant between 2002 and 2005. A sensitivity analysis has the role of testing the robustness of union entrance effects if unobserved confounders are time-variant instead. Two types of union formation are taken into account as a "treatment": marriage and non-marital cohabitation. The distinction between the two is justified by empirical findings denoting a rise of childbearing outside of marriage in

Bulgaria (Koytcheva, 2009), as well as theoretical reasons: cohabiting relationships, in terms of childbearing rewards, cannot be compared to those of married couples, because while the former provide a net gain in terms of lower regrets for the loss of freedom (Barber, 2001), marriage has the advantage of providing a supportive emotional and economic network (Clarkberg, Stolzenberg, & Waite, 1995; Waite, 1995). Finally, given that marriage and cohabitation seem to have a different impact on men with respect to women in terms of economic and health outcomes, they may experience a different push towards childbearing after entering a union. In more detail, men's wage outcomes seem to take more advantage than women's from entering a union. With respect to health outcomes, union formation also provides a higher level of social support and protection from which men would seem to benefit more. In order to control for these hypothetical differences, we will carry out the analysis by separating the effect of entering a union for men and women.

The remainder of the paper is organised as follows: the second section summarises the theoretical background of the study, giving additional details about the existing literature;

Section three presents the data; and Section four introduces attitudinal indices as outcome variables. Section five presents propensity score matching, while the estimates are reported in sections six and seven. The final section discusses the results and concludes the study.

5.2. Theoretical background

The aim of the paper is to address the causal effect of entering into a co-residential union (distinguishing then between marriage and non-marital cohabitation) on attitudes towards childbearing, intended as an individual's positive or negative evaluation of the birth of a child over three years. The interest in attitudes towards childbearing refers to the fact that attitudes are the main antecedents of parenthood (Billari, Philipov, & Testa, 2009; Fishbein & Ajzen, 1975): positive attitudes towards childbearing, along with subjective norms, predict intentions of having a child, and intentions predict behaviour (Barber et al., 2002). Given this linkage between attitudes and behaviour, the paper focuses on the process which contributes to forming positive or negative attitudes towards childbearing. The birth of a child is an important milestone in the process of transition to adulthood, which is often preceded by a union formation. Hence, we are interested in testing whether a union formation process has an effect on modifying attitudes towards childbearing. Axinn and Barber (1997) provide a theoretical framework for the explanation of attitude adaptation considered to be the end of a complex mechanism involving several steps: attitude changes through an individual learning process that is a consequence of the social interaction and new life course events experienced by the respondents. In addition, the existence of a process of cognitive consistency may convince individuals that past life course paths they experienced were legitimate.

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In the case of attitudes towards childbearing, the interaction with a new partner in a different living arrangement, which exposes individuals to new ideas and values, may change attitudes towards childbearing and family formation. At the same time, experiences of new living arrangements different from those of singlehood, also including a life without children, may have some advantages that were unknown in early adulthood. In this regard, Axinn and Barber (1997) stated that learning processes, perspectives and cognitive consistency produce more negative attitudes towards childbearing and fertility preferences in people experiencing non-family living arrangements in early adulthood than those experiencing cohabitations.

Literature dealing with attitude adaptation both in general and specifically due to union formation is scarce, but we found significant evidence for the recursive relation that connects union formation and subsequent value or attitude adjustment in Cunningham and Thornton (2005) and in Clarkberg (2002). The latter finds that the union experience changes orientations towards "the important things in life", including family formation. Results emphasise differences between the effect of cohabitation and marriage on a shift in attitudes. On one hand, marriage increases orientation towards marriage and childbearing; on the other hand, non-marital cohabitation leads individuals to give more importance to less conventional living arrangements. Similarly, Mulder (2003) found a remarkable difference in family orientations among young adults in the Netherlands between couples starting their first union being married and those who are cohabiting. Marriage remains the first choice for family-oriented individuals. Fertility intentions seem to be influenced not only by the union status per se, but also by the meaning attached to cohabitation. Hiekel and Castro-Martin (2014) found a strong association between fertility intentions and marriage intentions among cohabitors in Eastern Europe. Indeed, the proportion of cohabiters who viewed their union as an intermediate stage in the marriage process is higher in Eastern than in Western Europe. This is confirmed also for Bulgaria: Di Giulio, Impicciatore, and Sironi (2019) analyse cross national (including Bulgaria) survey data about women who started their first cohabitation between 1970–2005 and find some interesting differences between Bulgaria and other Countries. According to their results, 75.9% of the interviewees start a cohabitation as a "prelude to marriage" (whether with children or not), against the other Countries ranging from 28.5% to 40%. Besides, only 3.1% consider cohabitation as a temporary union, i.e. a lower quality relation, against the other Countries ranging from 14.9% to 27.6%.

Other studies emphasise this distinction between marriage and cohabitation with respect to the effect on fertility in Eastern European countries, where out-ofwedlock births have risen substantially, partly as consequence of the fall of the normative restriction of births within marriage. Among others, Philipov, Spéder, and Billari (2006) reported the effect of anomie on fertility intentions in Bulgaria, which might have led to a substantial increase in extramarital births. Spéder and Kapitány (2009) underlined how out-of-wedlock childbearing is becoming increasingly common in Hungary, even though married couples have a better probability of realising their fertility intentions.

From a theoretical perspective, marriage and cohabitation differ not only in the individual commitment, but also in the legal framework and social acceptance. On one hand, marriage establishes "enforceable trust" (Cherlin, 2004), that is, a strongly cohesive family unit, which is intended to jointly foster the long-term investments in children (Berger, Carlson, Bzostek, & Osborne, 2008; England & Farkas, 1986), while cohabitation is usually characterised by a temporary nature and uncertainty. Indeed, cohabitation is associated with lower childbearing preferences (Barber, 2001; Barber et al., 2002; Cunningham, 2005), as there is a negative effect of considerations about childbearing on intentions to cohabit

(Liefbroer & de Jong Gierveld, 1993). On the other hand, the legal, economic and social context plays a relevant role in shaping couples' decisions (Lesthaeghe, 1980; Liefbroer & Billari, 2010). As found by Lappegård, Klüsener, and Vignoli (2014), in Europe social norms and economic conditions are relevant for the decision to have children within cohabitation or within marriage. The status of the parents' union is relevant for paternity's establishment, joint custody and family name, so that cohabiting couples face some bureaucratic obstacles that may influence their attitudes towards childbearing (Perelli-Harris & Sánchez Gassen, 2012; Perelli-Harris et al., 2010). Furthermore, unlike married couples, cohabiting parents may be subject to social pressure and disapproval for their non-marital status that can negatively affect their attitudes towards having children. For example, in the Bulgarian case parental responsibilities towards children are established by the Bulgarian Family Code, which makes no distinctions with respect to marital status of parents. However, fatherhood is not automatic if the couple is unmarried, but can be established both through official recognition and by means of judicial decision. In this case, children are entitled to receive a monthly maintenance allowance from the parent in case of union breakdown. No specific provisions exist in the Bulgarian Family Code on the regulation of the relations between unmarried couples. For instance, unmarried mothers are not entitled to any alimony in case of union disruption, i.e. only the child is protected.

These findings on the effects of different living arrangements induce us to examine the potential outcome in terms of attitudes towards childbearing distinguishing between different union status. If, in fact, union formation has until recently been positively identified with marriage, as opposed to singlehood (Leplae, 1964), in recent decades, the diffusion and acceptance of non-marital cohabitation has rendered union status conceptualisation more challenging, as the contraposition between marriage and singlehood is now intertwined with alternatives (De Jong Gierveld, 2004). In this study, we use a three-fold classification of union status: single, cohabiting and married. We will consider non-marital cohabitation and marriage as two independent alternatives to singlehood, as described in Thornton, Axinn, and Xie (2007). In particular, we define single as anyone not living with a partner, following Thornton et al. (2007); thus, being single means either being unpartnered or having a partner, but not living with him/her (living apart together).

Marriage is an easily defined transition, since married status is a legally recognised category. On the other hand, cohabitation is more open-ended, and its boundaries are sometimes difficult to define and can be more blurred (Brown & Booth, 1996; Thornton et al., 2007). Indeed, cohabitation could be intended as an alternative to being single, thus having no clear plans for a future together, while it could also be intended as a more stable plan, an alternative to marriage or (more frequently) a marriage trial (Manning &

Smock, 2002; Manting, 1996; Thornton et al., 2007). Even though many approaches exist for defining various types of cohabitation and ways in which it is entered (Huang, Smock, Manning, & Bergstrom-Lynch, 2011; Sassler, 2004), due the features of the sample analysed in this study, we do not differentiate between different trajectories of cohabitation but rather regard it as one status, though we acknowledge its many facets (Brown & Booth, 1996; Clarkberg et al., 1995; Sassler, 2004; Thornton et al., 2007; Trost, 1978).

The effect of union formation on attitudes towards childbearing may be different not only with respect to the type of union, but also among men and women. Focusing on marriage, Goldscheider and Waite (1991) found opposite effects with respect to gender attitudes towards childbearing: positive for men and negative for women. A possible explanation for this difference could be the fact that men seem to benefit from marriage more than women. Other studies found that wages of men increase after marriage (Hewitt, Western, & Baxter, 2002), because men are more motivated to work harder (Becker, 1981), can specialise in market work (Korenman & Neumark, 1991), or are favoured by employers (Hill, 1979). In contrast, empirical findings on women's wages after marriage are mixed, that is data show both a decrease and a not significant impact of union formation on women's earnings (Loughran & Zissimopoulos, 2009; Neumark & Korenman, 1994). Furthermore, after marriage, risky and unhealthy behaviours decrease, and consequently health improves, in particular for men who can rely on their partner for support (Waite, 1995). With respect to those topics, cohabitants also benefit from the union in terms of reduction of risks due to the presence of partner's support. Nevertheless, some differences refer to cohabitation and marriage in a gender perspective: cohabiting couples seem to be more likely to show a more egalitarian division of domestic labour with respect to married ones. Using data from US, South and Spitze (1994) found evidence that married couples stick to a more traditional division of household labour if compared to other types of unions; Davis, Greenstein, and Gerteisen Marks (2007) in a cross-country study, which includes also Bulgaria, show that cohabiting men perform more housework than cohabiting women. The expectation of a greater support of males in the household work is expected to promote females' attitudes toward childbearing.

5.3. Sample description

The data used in this study come from a survey carried out in Bulgaria entitled The Impact of Social Capital and Coping Strategies on Reproductive and Marital Behavior, sponsored by the Max Planck Institute for Demographic Research in Rostock, Germany. The survey explored the impact of significant economic, cultural, social and institutional changes that affected Bulgarian society during the transition to an open market economy. In particular, it explored topics connected to the transition to adulthood and family formation as well as collected information on the social and economic background of the respondents, thus complementing the micro-level data with macro-level information on policies. The survey also explored the ideational and decisional side of demographic issues, namely attitudes, value orientations, religiosity and social capital of individuals.

Data were collected on a sample of respondents aged 18–34 in two rounds, first in 2002, then during the winter of 2005/06. The sample used in the analysis contains 2,563 individuals who were not in a union at the time of the first wave. Individuals who committed to non-marital cohabitation or marital union, that is to say a "coresidential union", between the two waves of the survey are labelled Treated Units; otherwise, they are labelled Untreated Units. The respondents are thus grouped as follows (Table 5.1):

Table 5.1. Sample used in the analysis by status during the first and second waves

	In union in 2002–2005	**Not in union in 2002–2005**	**Total**
Partnership	615 (24.15%)	*1932* (75.85%)	*2547* (100.00%)
Married	280 (10.99%)		
Cohabiting	272 (10.68%)		
Cohabitation turning into marriage	63 (2.47%)		

Respondents who had never been in a co-residential union before 2002, represent a substantial part of the entire sample. One-quarter of them experienced a new union formation between the two survey waves: 10.99 % of all respondents experienced marriage between the waves, whereas 10.68 % entered into a new cohabitation. A specific subcategory is represented by those who started a cohabitation that turned into marriage before the time of the second interview. Nevertheless, this category is numerically scarce; therefore, these individuals are included in the married subsample for subsequent analyses. Finally, widowed, separated and divorced respondents, who originally represented less than 5% of the entire sample, are excluded from the propensity score analysis due to their extremely different life experiences.

5.4. Attitude indices

The surveys carried out in 2002 and 2005/06 among Bulgarian respondents addressed several issues, including childbearing desires. These items evaluate the consequences, either positive or negative, of having a child in the following two years irrespective of the respondents' wish to have a child or not. Respondents were asked to give an assessment for every item based on a Likert scale. This set of items forms attitude indices in an agree/disagree format. Saris, Revilla, Krosnick, and Shaeffer (2010) have criticised questions with agree/disagree options because they are more likely to produce acquiescent response bias. However, as stated by the same authors, this effect occurs when people lack an opinion on an issue and have an inclination to acquiesce. Given the age of the respondents and the issue of the items, it could be assumed that this bias does not occur in this case. Anyway, we have implemented an equivalent factor analysis that is the solution to the potential bias proposed by the authors.

Factor extraction is obtained via a principal components analysis and is performed through a joint sample that includes both men and women (if we separate genders, results do not change). The answers that we considered for principal components analysis belong to the first wave of the survey.

Table 5.2. Results for factor analysis with respect to attitudes towards childbearing

Items	Factor 1	Factor 2	Uniqueness
	Positive attitudes	Negative attitudes	
1) Having children increases your economic difficulties.	-.128	.618	.602
2) Having children decreases your chances in your working career and/or education.	-.202	.534	.674
3) Having children increases your security that at old ages there is someone to care about you.	.597	-.101	.634
4) Having children increases joy and satisfaction in your life.	.659	.093	.557
5) Having children increases worries and preoccupations in the course of daily life.	.037	.792	.371
6) Having children decreases the time you have to pursue personal interests or be with friends.	.046	.777	.394
7) Having children increases certainty in your life.	.726	-.127	.457
8) Having children increases the closeness between you and your partner.	.815	-.034	.335
9) Having children increases the closeness between you and your parents and relatives.	.773	-.023	.401
10) Having children means that a part of you continues into the future.	.695	.102	.507

Results suggest the retention of two principal components[8], whose loadings are displayed in Table 5.2 after a varimax rotation. As clarified in the table, two latent dimensions (labelled as positive and negative attitudes) are set as proxies for individuals' attitudes related to fertility.

As in Billari et al. (2009), items 3, 4, 7, 8, 9 and 10 have high factor loadings on Factor 1, whereas the opposite pattern can be observed for the remaining four items. Indeed, people scoring high for this factor are strongly oriented to associate childbearing with satisfaction in life and to give more emphasis to the benefits of having a child. Therefore, Factor 1 can be interpreted as a proxy for positive attitudes towards childbearing according to Billari et al. (2009). In contrast, people with a high score for Factor 2 strongly agree with items describing the costs of having a child: they remark on the loss of career achievements and social life, and they especially focus on an increase in worries and distress, stressing the perception of childbearing (Barber et al., 2002; Crimmins, Easterlin, & Saito, 1991). More generally, people scoring high for this factor do not reject childbearing, but they are more likely to hold negative attitudes towards childbearing. Due to the specific formulation of the question, these orientations express preferences for an interval of approximately two years, starting from the date of the first interview. Because of the statistical construction of indicators, which are orthogonal by definition, the case of individuals scoring high values for both factors is admissible, that is, two opposite and competing tendencies may persist in each individual as also shown by Barber (2001).

The next step is to evaluate how attitude indicators changed by the time of the second interview. Factor represents a linear combination of J items at the time of the first interview t and is defined as:

$$(1) \quad Y_i^t = \sum_{j=1}^{J} a_{ij}^t Z_{ij}^t$$

where Y_{it} are standardised factor scores for the attitude index, a_{ij}^t is the coefficient of the linear combination estimated via principal component analysis for the item j and individual i at the time of the first wave, and Z_{ij}^t is the matrix of attitudinal items explored in Table 5.2. The indicators used for the second wave maintain the value of a_{ij}^t computed in t and replace the observation vector in time t with that observed in $t + \Delta t$:[9]

8 The number of factors extracted corresponds to the number of eigenvalues bigger than one.

9 A robustness check was implemented using $a_{ij}^{t+\Delta t}$ instead of a_{ij}^t in order to compute Y_i. The results obtained are similar.

$$(2) \quad Y_i^{t+\Delta t} = \sum_{j=1}^{J} a_{ij}^t Z_{ij}^{t+\Delta t}$$

Eq. (2) expresses the value of attitude indices for each individual in the second wave, which represent the dependent variables of our survey; the panel design allows for searching for a causal effect of a new $t+\Delta t$ union formation between the two interviews on the variable $Y_i^{t+\Delta t}$.

5.5. Statistical methodology

The aim of the empirical analysis is to identify the causal impact of co-residential union formation on individuals' attitudes towards childbearing after having experienced this life course event. To measure the effect of marriage and cohabitation, we use attitude indices. For this reason, we assume that each individual has two potential attitudinal outcomes: $Y_{1i}^{t+\Delta t}$ if the individual entered a co-residential union and $Y_{0i}^{t+\Delta t}$, if he or she did not. The causal impact of a union formation for each respondent included in the survey is $Y_{1i}^{t+\Delta t} - Y_{0i}^{t+\Delta t}$. Since this is an individual-specific variable, Rosenbaum and Rubin (1983) suggested focusing on the $E[Y_{1i}^{t+\Delta t} - Y_{0i}^{t+\Delta t} | T_i = 1]$, which in econometric literature is defined as the Average Treatment Effect on treated (ATT); in our case, this is the expected effect of a co-residential union formation on all individuals.

Let T_i be a dichotomous variable, denoted treatment, which takes value 1 if an individual got married or started cohabiting between the two waves and 0 otherwise (see Table 5.1). Heckman (1997) proposed restricting the analysis only to those who are actually eligible for the treatment. Hence, the main quantity of interest is:

$$(3) \quad ATT = E[Y_{1i}^{t+\Delta t} - Y_{0i}^{t+\Delta t} | T_i = 1] = E[Y_{1i}^{t+\Delta t} | T_i = 1] - E[Y_{0i}^{t+\Delta t} | T_i = 1]$$

However, $E\,[Y^{t+\Delta t} | T = 1]$ is unobservable because only one of the potential outcomes can be observed for each individual. A possible solution to overcoming this problem and finding the real value of the ATT is to replace $E\,[Y^{t+\Delta t} | T = 1]$ with $E\,[Y^{t+\Delta t} | T = 0]$. This hypothesis assumes that there is no selection bias, which means that those who experienced union entry are randomly selected from the population so that the two groups may be considered comparable in all other relevant characteristics. This assumption is not realistic because the two groups are different in terms of observable characteristics included in a set of variables X. Hence, according to Rosenbaum and Rubin (1983), identification of the ATT in (3) is feasible if we compare individuals with identical characteristics (Smith & Todd, 2005), that is with the same probability of receiving the treatment (in our case union entry), conditionally on the set of variables X. This probability is called propensity *t*4 score and is analytically defined as: $P\,[T_i = 1 \mid X_i{}^t]$. Propensity score can be easily estimated for each individual i using a logistic regression, where the dependent variable is T_i and the set of t covariates is given by X_i .[10] In more detail, Rosenbaum and Rubin (1983) state that comparing individuals with the same propensity score is equivalent to comparing them on vector X proposed matching based. Therefore, the ATT can be formalised as:

$$(4)\quad ATT = E_{p(Xi)}\{E[Y_{1i}^{t+\Delta t} \mid p(X_i^t), T_i = 1] - E[Y_{0i}^{t+\Delta t} \mid p(X_i^t), T_i = 0]\}$$

The method illustrated above is developed for all respondents. As we have to obtain sample estimates of ATT, we cannot match individuals with exactly equal values of propensity score. Therefore, we have to implement matching algorithms (Sianesi & Leuven, 2003), in order to overcome the problem and compare similar respondents. In this framework, we implement nearest neighbour matching, which consists of comparing each treated unit i, namely individuals who entered a coresidential union between the two interviews, with the closest control unit in terms of propensity score, that is, individuals who remained single between the two waves. In this case, we will have the following:

$$(5)\quad A\hat{T}T = \frac{1}{N_1} \sum_{i \in \{T_i = 1\}} \left[y_{1i}^{t+\triangle t} - \sum_{j \in \{T_i = 1\}} w_{ij}\, y_{0i}^{t+\triangle t} \right]$$

where N_i indicates the number of units that experienced a co-residential union formation in the considered time interval, and w_{ij} represents a sample weight for control units used in the matching procedure, which is usually equal to one.[11] Propensity score matching presents a main disadvantage: its estimates give robust results when matching on observables, but fails in the presence of unobservable confounding factors. Reliability of estimates holds if the selection bias is due to observable characteristics. In the presence of unobservables affecting assignment into both the treatment and outcome variable simultaneously, a hidden bias might arise. In this paper, we can mediate through two complementary methods in order to preserve estimates from irregular assignment to treatment due to hidden bias:

1) The richer the set of pre-treatment variables included in the matching, the lower the probability of omitting relevant unobserved variables. Thus, some latent traits could be captured by observable variables that are in some way related to the unobservables.

10 Propensity score can be easily estimated for each individual using logistic regression where the dependent variable is Xi and the set of covariates is given by X^t_i

11 FollowingCaliendo and Kopeinig (2008), we estimate PSM requiring that the “common support condition” has been satisfied. It ensures that persons with the same X values have a positive probability of being both participants and non-participants.

Hence, the first suggestion we have implemented is to use a rich vector of covariates to estimate the propensity score, including proxies for potentially unobservable dimensions, such as measures of attitude indicators after the first wave and other variables capturing the psychological well-being of respondents as suggested in the literature (Philipov et al., 2006);

2) The approach described above is used in cross-sectional surveys. Conversely, this study is based on longitudinal data and measures each available variable before and after the treatment. This allows us to compare the mean change of attitudes from the first wave t to the time after the second wave t+Δt. The obtained estimator is defined as the difference in difference estimator and is used to estimate the differences of the mean of outcomes before and after Δt:

(6) $$DD = E(Y_{1i}^{t+\Delta t} - Y_{1i}^{t}) - E(Y_{0i}^{t+\Delta t} - Y_{0i}^{t}) = E(\Delta_{1i}) - E(\Delta_{0i})$$

An important advantage of the DD estimator is that it allows for controlling the selection due to time-invariant sources of bias as in Smith and Todd (2005). Of course, even this assumption might be violated if some time-varying confounders exist. Thus, it is possible to combine the DD with the propensity score matching estimator, joining the advantages of the two approaches. The DD-PSM combined estimator for the ATT can be implemented to estimate the following quantity:

(7) $$ATT_{DD} = E_{p(Xi)}\{E[\Delta_{1i} \,|\, p(X_i^t), T_i = 1] - E[\Delta_{0i} \,|\, p(X_i^t), T_i = 0]\}$$

In Eq. (7) individuals are always matched on the basis of estimated propensity score, but the analysis is done by comparing the average change that occurred between time and time-tt in the outcomes variable for the treatment group, compared to the average change overtime for the control group (instead of comparing the values of the outcome only in-tt). In presence of unobservable factors that bias the results of the propensity score matching and that are invariant overtime, the DD method allows us to overcome this problem, because constant unobservable characteristics in the two waves are dropped out when average changes are computed. Results presented in Section 5.7 provide estimates for the quantity expressed in Eq. (7), that is, the average increase of attitude indices from the first to the second interview due to a union formation occurring between the two waves.

5.6. Propensity Score Specification

Table 5.3 (column 1 and 2) presents the results of the logit models used for estimating the propensity score whose results will be displayed in Section 5.7. In practice, this set of models addresses the determinant of a new union formation happening between the two waves of the interview for each sampled individual. It also presents the specification of propensity score used in the matching stage. Given that a subsequent analysis devoted to separate different types of union has been implemented, the model has been run also for separately estimating the determinants of marriage and non-marital cohabitation.[12] The propensity score theory suggests that the set of covariates to be included in the model should include all variables that supposedly affect both attitudes towards childbearing and the likelihood of entering a co-residential union. Indeed, once we have established that attitudes may change after union formation, we are in the presence of possible confounders that may simultaneously affect both the likelihood of union formation and attitudes towards parenthood, which may under-mine the estimates of ATT. The demographic literature provides sup-port in the choice of a correct set of covariates in order to implement the matching. According to the literature, factors such as education, economic conditions (e.g. household income, measured in quartiles, and work-force participation) and values play a key role both in building attitudes towards childbearing and in affecting the choice of entering into union (2000, Oppenheimer, 1988, 2003; Oppenheimer, Kalmijn, & Lim, 1997; PerelliHarris et al., 2010). Women's education and access to the labour market are, in fact, usually indicated as the reason for the postponement of marriage (Becker, 1973). Age represents, obviously, a relevant factor for the decision of entering into union, where we expect younger and older cohorts to have less probability of marrying. Studies on the intergenerational transmission of behaviour show that fertility behaviours of parents and children are positively correlated (Seltzer,1994), thus a variable on the number of siblings is included.

12 Individuals who started a cohabitation that turned into marriage in the same time window (2002-2005) were considered as married.

Table 5.3. Odds ratios for determinants of entering a union: results of logit model.

Explanatory variables (all measured at time t)	Entering in a union (marriage + cohabitation) vs. never in union between t and t+Δt		Getting married vs. c never in union between t and t+Δt		Starting a non-marital ohabitation vs. never in union between t and t+Δt	
Attitudes	**Males**	**Females**	**Males**	**Females**	**Males**	**Females**
Positive attitudes towards childbearing	1.153	0.988	1.342***	0.988	1.037	1.169
Negative attitudes towards childbearing	0.929	0.984	0.744**	0.984	1.174	1.261
Attitudes towards partnership	0.974	1.050	1.060	1.050	0.980	1.027
Past partnership						
No (ref.)	1	1	1	1	1	1
Yes	2.824**	2.044**	1.537	2.044**	3.867**	2.351*
Household income						
1st quartile (ref.)	1	1	1	1	1	1
2nd quartile	1.047	1.179	0.851	1.179	1.106	1.301
3rd quartile	0.634**	1.110	0.822	1.110	0.393***	0.866
4th quartile	0.701	1.407	0.885	1.407	0.509*	1.937*
Working condition						
Does not work or study (ref.)	1	1	1	1	1	1
Studies, does not work	0.434**	0.66	0.595*	0.66	0.253**	0.546
Works in a private firm	1.215	0.997	1.694*	0.997	0.893	1.008
Works in a state firm	1.040	0.642	1.469	0.642	0.739	0.656
Education						
Primary or less (ref.)	1	1	1	1	1	1
Basic	0.656	0.205**	2.107	0.205**	0.551	0.242*
Secondary	0.484	0.336*	1.773	0.336*	0.322*	0.230*
Higher	0.752	0.337*	2.411	0.337*	0.480	0.175**
Number of siblings	1.017	0.905	0.960	0.905	1.141	1.015
Father's Education						
Primary or less (ref.)	1	1	1	1	1	1
Basic	0.981	1.435	0.804	1.435	1.266	0.938
Secondary	1.110	0.865	0.980	0.865	1.477	0.800
Higher	0.896	0.792	0.688	0.792	1.598	0.776
Religiosity						
Not religous (ref.)	1	1	1	1	1	1
Religious	1.043	1.500**	1.277	1.500**	0.868	1.867**

Age						
18-21 (ref.)	1	1	1	1	1	1
21-27	1.518**	1.01	1.171	1.01	2.051**	1.141
27-30	1.981***	1.364	1.897**	1.364	1.992**	1.607
30-34	0.795	0.611	0.823	0.611	0.846	1.100
Anomie indices						
Psychological well-being	1.158	1.214**	1.168	1.214**	1.141	1.065
Disorientation	0.996	1.193	1.154	1.193	0.793	1.054
Social capital index						
Exchange of help	0.927	1.228	0.850	1.228	1.023	1.152
Economic constraints	0.959	1.094	0.851	1.094	1.146	1.011

***** $p < 0.01$, ** $p< 0.05$, * $p<0.1$** [a] *The presence of children in the first wave predicts the value zero for the dependent variable for cohabiting models. Hence, this covariate, used for predicting union entrance, has been omitted.*

Cultural and religious factors are added taking into account cultural influences, as well as a dimension on past-experiences, such as previous relationships. Finally, psychological and behavioural characteristics of the respondents may influence individuals' inclination and suitability for marriage and parenthood, which has been indicated in past studies using the same set of data (Philipov et al., 2006).

First of all, childbearing attitudes mea-sured at time are also included among the explanatory variables. Indeed, attitudes may be formed early in the life course (before t) and they may drive people's life choices, in particular regarding union formation. This is even more true for attitudes towards partnerships that are directly related to union entry. In addition, that set of variables includes the following factors extracted with the principal components methodology: 1) an indicator for "Psychological well-being", which appears in the analysis, combining the responses to the items "During The past month have you ever felt very lonely or remote from other people?" and "During the past month have you ever felt depressed or very unhappy?"; 2) a "Disorientation indicator", that is obtained combining the answers to the following statements: "I have no influence over my everyday affairs"; "Life is so complicated nowadays that most of the time I don't know what to do" and "No One Cares What Happens To Other people"; and, finally; 3) it has also been included a measure of "Social Capital"defined as a factor combining two sub-indicators, which are "help received" and "help given". Finally, 4) a variable defined as "economic constraints" (Billari et al., 2009). That variable has been computed from the following question: "How much would your decision on whether or not to live with a/your friend during the next two years depend on: a) your income, b) your working or educational situation, c) your housing conditions, d) your health status". Table 5.3 offers a separate analysis for men and women. The results show a relevant heterogeneity between the determinants of family formation for males and females. Estimates substantially confirm the general picture: enrolment in the educational system penalises the opportunities of those involved in co-residential union formation, in case of for women (Thornton et al., 1995), and levels of education above primary are linked to a lower likelihood of forming a union; the likelihood reaches its minimum at the secondary education level in the model for determining union formation, confirming the estimates from Philipov et al. (2006). The permanence in the education system has important direct effects on marriage, non-marital cohabitation and the timing of first childbirth. With respect to the role played by ideational factors, attitudes towards childbearing do not change the propensity of entering a new union, but become significant in predicting marriage between the two waves. Other socio-demographic characteristics seem to play a significant role in the process selection on observables: gender plays a decisive role in determining union entry. Religiosity increases the likelihood of forming a union and, particularly, of getting married in the case of women and agrees with the general framework, which implies that women who describe themselves as religious are more likely to marry, have children and support gender division of housework (Hayford & Morgan, 2008). Social capital and economic constraints seem to play a minor role in this context but have been included in the model according to previous studies (Philipov et al., 2006). In addition, estimates confirm that women are more likely to enter into a union if their level of psychological well-being is higher. Philipov et al. (2006) also stresses the dependence of Bulgarian women's behaviour on psychological well-being in determining the intentions of having children. Finally, a relevant role has been played by past partnerships in the logit estimates. This variable is a dummy taking value 1 if an individual has had a previous non-marital cohabitation and 0 otherwise. Results show that individuals with past experiences in a union are more likely to re-enter into union between the two waves. In conclusion, estimates deriving from Table 5.3 confirm that several confounders have an impact on the likelihood of entering into a co-residential union, supporting the idea that selection bias may arise, even if the pattern of significance changes across the models. For example, attitudes seem to be significant only in predicting marriage, while psychological wellbeing and religiosity play a key role only for females. Anyway, a coherent analysis requires to adopt the same specification in order to model the different union types. Results of matching estimates (matched sample) and descriptive differences in means (full sample) are given in the next section, in order to be compared.

5.7. Matching Estimates

Considering individuals not in a co-residential union in the first wave, Table 5.4 portrays the effect of experiencing union entry between the two waves on factor score changes for men and women included in the sample, comparing individuals with the closest values of estimated propensity score. As shown in Table 5.4, an unmatched *ATT* does not evidence differences among individuals who entered into a co-residential union between the two waves of the survey. Propensity score estimates, which are robust to observable confounders, confirm previous results with the exception of the effect of co-residential union entry on the *"benefits" factor*. Indeed, the matched men sample shows increased positive attitudes towards childbearing. This effect emerges when selection bias due to observables is reduced through the propensity score.

Table 5.4. Effect of union formation on childbearing attitudes, whole sample

		Males		Females	
Union formation on attitudes		**Full sample**	**Matched sample**	**Full sample**	**Matched sample**
Union formation on positive attitudes	ATT	.102	.294**	.118	.146
	(t stat)	(1.19)	(2.39)	(1.27)	(1.12)
	Γ critical		1.25		1
	Obs (T=1)	217	214	222	218
	Obs (T=0)	977	214	559	218
Union formation on negative attitudes	ATT	-.114	-.114	.016	-.070
	(t stat)	(-1.28)	(-.96)	(.16)	(-.48)
	Γ critical		1		1
	Obs (T=1)	217	214	222	218
	Obs (T=0)	977	214	559	218

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

Table 5.5. Decomposition of causal effect of union formation (cohabitation and marriage) on childbearing attitudes, whole sample

		Males		Females	
Cohabitation on **positive attitudes**		**Full sample**	**Matched sample**	**Full sample**	**Matched sample**
Cohabitation on positive attitudes	ATT	.127	-.169	.133	.120
	(t stat)	(1.07)	(-.90)	(.99)	(.66)
	Γcritical		1		1
	Obs (T=1)	101	100	87	87
	Obs (T=0)	977	100	559	87
Cohabitation on negative attitudes	ATT	-.117	-.058	-.082	.312
	(t stat)	(-.94)	(-.33)	(-.59)	(1.54)
	Γcritical		1		1
	Obs (T=1)	101	100	87	87
	Obs (T=0)	977	100	559	87
Marriage on positive attitudes	ATT	.042	.317**	.129	.071
	(t stat)	(.37)	(2.00)	(1.13)	(.41)
	Γcritical		1.25		1
	Obs (T=1)	109	109	129	129
	Obs (T=0)	977	109	540	129
Marriage on negative attitudes	ATT	-.115	-.308*	.070	.004
	(t stat)	(-.95)	(-1.78)	(.56)	(.02)
	Γcritical		1.1		1
	Obs (T=1)	109	109	129	129
	Obs (T=0)	977	109	540	129

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

Nevertheless, the estimates of average treatment effects are reliable if the outcome variable is largely free of bias arising from an association between the choice to have children and unobservable factors (Di Prete & Gangl, 2005; Rosenbaum, 2002). Since the approach to value orientations and attitudes may be consistently affected by the presence of unobservable covariates, we need a robustness analysis of matching results, implementing a difference-in-difference estimator and a sensitive check with respect to unobservable confounders.[13] In addition, the use of union formation as a unique treatment for the model means hiding the difference that exists between marriage and cohabitation with regards to the childbearing desires of Bulgarians. Hence, in Table 5.5 we split individuals experiencing union entry in two subsets: those who are cohabiting and those who are married by the time of the second interview, in order to decompose the causal effect of union formation.[14]Even though marriage is still the preferred environment, to some extent, to raise children (Thornton et al., 2007), the importance of cohabitation cannot be ignored. Indeed, cohabiters are less likely than non-cohabiters to hold traditional family values (Bumpass, Sweet & Cherlin, 1991), less likely to follow traditional gender roles than married couples (Lesthaeghe & Surkyn, 1988), and less committed to the idea of durability and permanence of the union (Axinn & Thornton, 1992; Bumpass et al., 1991).

The results underline differing among cohabiting and married couples as the mechanism through which gender operates. The results also emphasise the role of marriage in increasing positive attitudes towards childbearing for the men's subsample: the *ATT* increases in magnitude and shows a strong reliability to hidden bias. Conversely, cohabiting couples do not show any significant modification in attitudes towards parenthood for either gender. With respect to this point, Sassler and Cunningham (2008) point out that marriage is still the preferred choice for childbearing among both singles and cohabiters. Moreover, the lack of childbearing preferences among cohabiters might also result from ignorance by the couple in terms of their plans if they were not discussed before entering the cohabitation. Besides the problems of hidden bias and of splitting marriage and cohabitation,[15] another issue concerns the estimates of causal effects in Tables 4 and 5: union entry is often linked to concomitant childbearing. Indeed, the presence of children changes patterns of union

13 Table 5.4 shows the results of Wilcoxon sign-rank tests. As displayed in the table, the estimates of *ATT* for the positive attitudes do not seem to be relatively robust to hidden bias: the critical level of , at which the significance of *ATT* is questioned, is less than 1.25.

14 According to the summary statistics, married individuals include also those who were not in union in 2002 and who started a cohabitation that turned into a marriage within the time of the second interview.

15 The choice of splitting the effect of marriage and cohabitation on changes in attitudes requires rerunning the propensity score procedure with the same set of explanatory variables reported in Table 3. The only change concerns the dependent variable of propensity score that compares married vs single individuals in the model measuring the ATT for marriage as well as cohabitants vs singles in the model measuring the ATT for cohabitation.

entry (Billari et al., 2009), speeding up the process of getting married or cohabiting (Manning, 1993). Therefore, the presence of concomitant childbearing between 2002 and 2005 may bias the ATT of union formation, especially whenever we do not know with certainty whether the birth of a child occurred before or after entry into union. This problem of reverse causality may underestimate or overestimate the effect of the union event. Indeed, the theory suggests that a new union is supposed to increase benefits and decrease the costs associated with having a child if union formation unequivocally starts before the woman's pregnancy. Conversely, the effect on attitudes towards childbearing may change if childbearing advances the union formation, especially in the case where pregnancy is unexpected. In addition, a further reason for excluding individuals with children between the two waves is in the formulation of attitude items. The focus of the research is on the variation in *attitudes towards parenthood*, in which the attitude is referred to as the birth of a child since the first interview. The birth of a first child between interviews changes the meaning of the attitude in $t{+}\Delta\ t$, which would be referred to as the birth of a second child, giving a different meaning to the item measured in t than to that measured in $t{+}\Delta\ t$ and referring to a subsequent child. Estimated results in Table 5.6 confirm this idea, showing higher *ATT* scores. In particular, considering only childless individuals and studying the impact of union formation among this subsample improves the significance of the estimates of the model.

5

Table 5.6. Decomposition of the causal effect of union formation (cohabitation and marriage) on childbearing attitudes, childless sample

		Males		Females	
Marriage/cohabitation on attitudes		**Full sample**	**Matched sample**	**Full sample**	**Matched sample**
Cohabitation on positive attitudes	ATT	.183	.401*	.229	.443*
	(t stat) Γcritical	(1.30)	(1.74)	(1.43)	(1.78)
	Obs (T=1)	69	1.3	57	1.05
	Obs (T=0)	977	67	547	55
			67		55
Cohabitation on negative attitudes	ATT	-.423***	.039	-.221	.097
	(t stat) Γcritical	(-2.86)	(.17)	(-1.29)	(.36)
	Obs (T=1)	69	1	57	1
	Obs (T=0)	977	67	547	55
			67		55
Marriage on positive attitudes	ATT	.427***	.723***	.226	.311
	(t stat) Γcritical	(2.58)	(3.17)	(1.23)	(1.14)
	Obs (T=1)	49	1.75	41	1
	Obs (T=0)	977	49	530	40
			977		530
Marriage on negative attitudes	ATT	-.167	-.651***	.098	-.369
	(t stat) Γcritical	(-.96)	(-2.73)	(.50)	(-1.12)
	Obs (T=1)	49	1.65	41	1
	Obs (T=0)	980	49	530	40
			49		40

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

Results presented in Table 5.6 reinforces previous results and are coherent with them, even if a certain caution has to be mentioned before commenting the empirical results: first of all, dropping out individuals who experienced a childbirth between the two waves introduce a selection bias, which counterbalance some incoherency in the wording of attitudes in t and t+Δt. If a new child was born between the two waves, the items used for building the attitudes refer to two different situations: the birth of a first child in t and the birth of a second child in t+Δt. The second caution refers to the sample size that has been reduced when the analysis is performed both by gender and by type of union and excluding from the sample those that gave birth to a child between the waves, as underlined in the Table 5.6. Important reductions of sample size are quite usual in propensity score settings and some literature guarantees the robustness of propensity score estimates in presence of sample sizes equal to forty units, as underlined by Pirracchio et al. (2012). Anyway, results are always coherent to those that are obtained in Table 45. and 5.5.

Results suggest a substantial impact of marriage on both positive and negative attitudes towards childbearing among men. Marriage is still the living arrangement most connected to childbearing. The effect seems to be significant only in the men's subsample with respect to positive attitudes. The effect is milder among women, as evidenced in the previous results (Tables 5.4 and 5.5).

Results do not exclude the scenario in which only entering a union with a marriage may activate attitudes towards parenthood. It is likely that marriage has been considered to be a more binding institution that ensures more stability of the relationship, furnishing better guarantees regarding the collaboration of both parents in the future expenditures required for childbearing (Cherlin, 2004; England & Farkas, 1986). So after the celebration of the marriage, partners may decide to take more into account the idea of giving birth to a child, increasing the benefits of a future child and starting to feel a decreased risk of becoming a parent. These findings support the literature proposing marriage as *the institution* for childbearing, in which most couples desire to have children (Barber, 2001; Barber et al., 2002; Cunningham, 2005). The entry into non-marital cohabitation changes only positive attitudes towards childbearing, in comparison to single individuals for both males and females.

Summarising the decomposition of the causal effect of union formation on the impact of marriage and cohabitation on attitude changes, we found that, in general, marriage increases attitudes towards childbearing. The effect of cohabitation on increasing attitudes is often not significant and uncertain, even if generally stronger once again in men than in women. Although few papers provide evidence of the effect on attitudes considering gender specificity in the case of marriage, there are studies in the literature that support and explain our findings. Seccombe (1991) analysed the costs and benefits of having children after entering marriage for childless individuals (as in our paper in Table 5.6), comparing married and cohabiting individuals and differentiating by gender. The results reject the pronatalist ideology that stresses the strong relation between women and motherhood: husbands attach greater importance to having children with respect to wives after the entrance into a union. As improvements in women's employment status in recent years increased the financial power and the economic role of a woman in a union, childbearing may disrupt the balance of power. On the contrary, men have less to lose and more to gain economically and socially by the birth of a child.

5.8. Discussion and Conclusions

The analysis we have presented offers several contributions to the understanding of how union formation influences attitudes towards childbearing. In order to identify the process of attitude adaptation, we explored several attitudinal items directly related to fertility, summarised in two distinct dimensions. The first dimension was linked to the benefits of having a child at the second interview, while the second emphasised the related costs. The analysis was conducted by means of propensity score matching algorithms to identify the causal effect of union formation on attitude shifts, controlling for individual constraints and demographic characteristics.

Our results show that, in general, the effect of union formation is positive and significant on positive attitudes toward childbearing, but only for men (Table 5.4). However, the decomposition of the causal effect by type of union formation and gender shows that the impact differs not only between men and women, but also between cohabiting and married couples. Also in this case, there is no effect among women. As regards men, entering in a union formation not only increases positive attitudes toward childbearing, but also decreases the negative ones. Although, this effect is statistically significant only if the union is a marriage.

Gender differences in the estimates are not explained by an ex ante difference between men and women on family formation desires. Indeed, the attitude levels at the beginning of the time window taken into account are not different if we compare men and women. If women would be more family oriented, than men independently of the relevant life course events, then our data would show that men and women start with unequal levels of family orientation, but this is not the case. Several studies show that women continue to carry the primary burden of childcare (Bielby & Bielby, 1989; Biernat & Wortman, 1991; Hochschild, 1989).

Loughran and Zissimopoulos (2009), for example, show that a first birth reduces female wages by 2–3 percent while male wages are unaffected by childbearing. Furthermore, others demonstrate how the work/family conflict is greater in working women than in men

(Cleary & Mechanic, 1983; Duxbury & Higgins, 1994; Wortman, Biernat & Lang, 1991); this aspect is of particular interest in a post-communist country such as Bulgaria, where the gap between female and male workers is lower than in Western countries (European Commission, 2013). As mentioned in the theoretical section, marriage seems to exacerbate, in some cases, this reduction in female wages even before the birth of the first child. On the contrary, men seem to experience some benefits on wealth and health that also have an impact on attitudes toward childbearing. The results could be, then, explained by the fact that after marriage men may benefit more from the role division within the couple and are more confident to concentrate themselves in pursuing success in their career development, thus they feel more inclined toward having children and less worried about the costs of rearing them. These benefits do not involve women, when they start a new union, especially if they foresee their commitment in the early stages of the life of a baby, that is perceived as higher than what happens to men: this may justify why they do not change their attitudes toward children after marrying. Since cohabitation does not require the same level of commitment, it could be that men feel less confident to invest in the choice of giving birth to a child. In addition, cohabitation in some cases represents a trial period that anticipates marriage and that is focused on testing whether the partner relationship works well more than planning to have children in the short run; this set of reasons may justify why we have not not found the same effect among cohabitants.

This result is also coherent with previous literature that describes how married people exhibit stronger attitudes towards childbearing. The given explanations are that, on one hand, marriage furnishes greater guarantees regarding the collaboration of both parents in the future expenditures necessary for childbearing (Cherlin, 2004; England & Farkas, 1986). On the other hand, the literature has shown that married couples have better life conditions than cohabiting individuals, both in terms of health and longer life experience (Lillard & Waite, 1995) and higher average wealth (Lupton & Smith, 2003; Manning & Brown, 2006).

Therefore, even if marriage is progressively weakening its central role as the first step in union formation in the Bulgarian society, as confirmed by the high prevalence of children born from parents living in non-marital cohabitation, but it still holds a positive and fully significant effect in augmenting positive attitudes towards parenthood.

The first part of the analysis involved the effect of union formation on changes in attitudes towards childbearing, without accounting for the concomitant birth of a child. However, including individuals who experienced childbearing between the waves may spuriously and negatively influence the effect of union formation on attitudes towards parenthood: indeed, individuals who had a child exhibited a decline in positive attitudes towards parenthood immediately after childbirth. This could be explained by the fact that the choice of having a second child differs substantially from having the first one: Bühler (2008) stated that a first child plays a key role in strengthening the feeling of closeness between the partners in Bulgaria. Once members of the couple have satisfied their needs regarding childbirth, we observe less incentives in the decision of giving birth to a second child, especially in the short run. In addition, higher educated individuals, especially women, perceived increased opportunity costs of childbearing in terms of loss of working opportunities and the risk of worsening economic family conditions, which are a result of the low public support and the decline of the pronatalist policy in the country after the fall of the Iron Curtain. This is reflected in the strong preference for a single child in Bulgaria (Billari et al., 2009; Koytcheva, 2006; Koytcheva & Philipov, 2008; Philipov & Kohler, 2001; Philipov et al., 2006), where couples are oriented to forming a family with only one child. Hence, in the second part of the analysis we have considered only respondents without children. The results are in line with our expectations: marriage substantially increases positive attitudes and decreases the negative ones for men, while it has no impact on women. With respect to the previous result, cohabiting affects the positive attitudes toward childbearing of childless couples, for both men and women.

One possible explanation of the fact that, when removing couples with children, also cohabitation becomes significant can be found in the already discussed difference between cohabitation and marriage. Cohabitants are less likely than married couples to pool financial resources, to support the partner financially, to spend free time together and to agree with the future of the relationship (Waite, 1995). Hence, it is not surprising that this uncertain situation increases the belief that a child would give them certainty, security and closeness with the partner, and thus positive attitudes toward childbearing increase, for both men and women.

Many analyses do not distinguish the effect of attitudes on behaviour from the effect of behaviour on attitude orientations due to a general presumption that causality runs mainly from attitudes to behaviour. Nevertheless, the theoretical literature notes that the causality process also runs in the reverse direction, proceeding from behaviour to attitudes (see a brief literature review in Axinn, Emens & Mitchell, 2008).

The relevance of this study is given by the in-depth analysis of the complex and under-investigated subject regarding the impact of union formation on changes in attitudes towards childbearing. This relationship is potentially affected by endogeneity problems due to reverse causality and selection bias distorting the results. Nevertheless, given the presence of a unique longitudinal dataset and the combination of different statistical techniques, we overcome these issues and compute the unbiased magnitude of the causal effect of union formation on attitudes towards childbearing.

5.9. References

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