

## The total marital fertility rate and its extensions

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**The Total Marital Fertility Rate and its Extensions**

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## The Total Marital Fertility Rate and its Extensions<sup>1</sup>

### Abstract

What we will call the *age-based TMFR* is computed conventionally by adding up age-specific marital-fertility rates in the hope of estimating the number of children ever born to a woman who is married throughout her childbearing years. Demographers have long been strongly skeptical about this quantity because it normally indicates implausibly many children. Our analysis of data from the Romanian GGS confirms this finding, and we propose an alternative *duration-based TMFR* computed in the spirit of parity progression ratios. At the same time we extend the method to cover any type of living arrangement (cohabitation, marriage, non-partnered arrangement, and so on). Because each resulting Total Union-type Fertility Rate (*TUFR*) explicitly accounts for the living arrangement, it improves on the conventional Total Fertility Rate, which does not. We embed the investigation in an event-history analysis with fixed and time-varying control covariates and find patterns of relative risks for such variables that reveal interesting features of childbearing behavior in the Romanian data, which we use to illustrate the method. In most cases these patterns are quite robust against model re-specification, including the shift from the age-based to the duration-based approach. Since the number of female respondents is 'only' about 6,000 (minus records that cannot be used for the current purpose) in a normal single-round GGS, there is considerable inherent random variation in the data set, but we show that simple few-term moving-average graduation suffices to overcome this problem.

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<sup>1</sup> The authors have contributed equally to this paper.

## **Le taux de fécondité totale dans le mariage et ses extensions<sup>2</sup>**

Le taux de fécondité totale en mariage (TFTM) selon l'âge est calculé par convention en sommant les taux de fécondité par âge dans le mariage en vue d'obtenir une estimation du nombre total d'enfants nés d'une femme qui aurait été mariée tout au long de sa vie reproductive. Depuis longtemps les démographes considèrent cette mesure avec scepticisme car elle aboutit souvent à un nombre total d'enfants beaucoup trop élevé. Notre analyse des données du EGG roumain confirme cette constatation et nous proposons dès lors, comme alternative, un TFTM selon la durée, dans l'esprit des probabilités d'agrandissement des familles. Par ailleurs, nous étendons la méthode à tous types de situation de couple (cohabitation, mariage, sans partenaire, etc.). Comme le taux de fécondité totale selon le type d'union (TFTU) tient compte explicitement de la situation de couple, il doit être préféré au TFTM qui ne tient pas compte de ce critère. Notre étude est conduite dans le cadre d'une analyse biographique tenant compte de covariables fixes ou dépendantes du temps. Les résultats de l'analyse nous permettent de découvrir des caractéristiques intéressantes de la fécondité roumaine, que nous utilisons pour illustrer la méthode. Dans la plupart des cas, ces caractéristiques sont robustes face à une re-spécification du modèle, notamment le passage de l'approche basée sur l'âge à l'approche tenant compte de la durée. Comme le nombre de répondants à l'enquête EGG à un passage n'est 'que' de l'ordre de 6000 (moins les cas qui n'ont pas pu être utilisés pour le présent travail), il existe une fluctuation aléatoire importante dans les données. Nous montrons toutefois qu'un lissage par moyenne mobile à quelques termes seulement nous permet de surmonter cette difficulté.

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<sup>2</sup> Les auteurs ont contribué de manière égale à ce texte.

## 1. Introduction

In their recent discussion of the Total Fertility Rate (TFR), Hoem and Mureşan (forthcoming) showed that this concept can be extended to also account for fixed and time-varying factors other than age attained. This produces an indirectly standardized TFR that covers impacts from factors other than the population age composition and that is therefore less subject to the criticism that the fertility measure is overly rigid. The two authors deliberately left aside the composition by civil status to avoid needlessly complicating their presentation. In the present paper we pick up that thread where they left it. Our purpose is to study how the TFR can be extended to account explicitly for the various types of living arrangements that are available to individuals in a population. Such an extension should greatly improve on the TFR, because (as is well known) the childbearing “risk” is strongly influenced by the type of union.

The Total Marital Fertility Rate must originally have been invented to provide a parameter to measure the level of marital fertility on the basis of information that was available in good national statistical systems at the time of its inception. Since it tended to indicate implausible levels of fertility in marriages, it got a bad press and fell into disrepute, however. In the present paper we suggest three developments, namely (1) that the *age-based* total rate be replaced by a similar *duration-based* rate in the spirit of the analysis of parity progression ratios, (2) that the study of the fertility effect of the living arrangement be extended to also include types of unions other than marriage (in fact we operate with two types of marriage plus non-marital unions), and (3) that the whole investigation be embedded in an event-history analysis which includes fixed and time-varying covariates beyond union type, age attained, and union duration. Step 1 removes the most flagrant weakness of age-based analysis, namely that its outcome is dominated by the high fertility at young ages and short union durations which it fails to take into account. Step 2 broadens the focus and includes union forms that have grown tremendously in importance in recent years and that may have received less attention than they deserved even before modern developments. This step produces a union-specific TFR for each type of union considered. Step 3 allows the analyst to focus more on features of behavior that may have been overshadowed by the interest in fertility *levels* at a time when there was less focus on fertility *patterns*. It turns out that this latter step has the added advantage that it reveals how robust fertility patterns seem to be against model re-specification (if our empirical findings can be generalized to other data sets), in that we largely find the same patterns of effects of our control factors in our duration-based as in our age-based analysis. This suggests to us that such robust features should be included in a future re-specification of the content of official statistics.

The use of the event-history approach supersedes the various applications and extensions of ordinary life-table analysis that have appeared in the literature on parity progressions over the last thirty-odd years.

To avoid overburdening our presentation we restrict ourselves mostly to a partition of the respondents into four groups, namely into those who at any given time (i) have never lived in a marital or consensual union, (ii) currently live in a non-marital union (the cohabitants), (iii) have converted a cohabitational union into a marriage, and (iv) have married directly (i.e., without a consensual union that has been converted into a marriage). The subdivision can be extended trivially to also cover individuals who live in other union forms, such as widows, divorcees, other non-partnered individuals who have previously lived in a consensual union, and so on. As

is common in investigations of total fertility we limit ourselves to women, though the same methodology can be applied to data for men. If the data permit, further living arrangements, like couples who “live apart together” or live in same-sex unions can additionally be covered. For suitable generality we shall use the term ‘union types’ for the living arrangements that the analysis covers. We avoid the term ‘civil status’ because we are not certain that this term can be used for a consensual union or some other arrangement that does not involve official registration. We shall use the abbreviation TUF<sub>R</sub> (as in Total Union-type Fertility Rate) to cover a TFR-type quantity produced for any kind of union type, in analogy with the more conventional TMFR (the Total Marital Fertility Rate).<sup>3</sup>

The time-varying factors involved in the extension of the Total Fertility Rate studied by Hoem and Mureşan (forthcoming) typically depend meaningfully on age  $x$  attained and are features like parity (at age  $x$ ), educational attainment (at the given age), or the calendar period in which an individual attains age  $x$ . It is not quite as straightforward to incorporate time-varying control variables in a study of the effect of union status. Individuals who live in a marital or non-marital union are likely to have a fertility that depends more strongly on how long that union has lasted than on how old they are,<sup>4</sup> though their age is sure to have an impact also. To cover all bases we could operate with fertility rates that depend *both* on age *and* union duration, but to avoid overkill, we will work with age-dependent and duration-based fertility rates separately. Section 2 contains what we have to say about age-based fertility and Section 3 covers duration-based fertility. Section 4 compares outcomes of the two approaches.

We use data from the Romanian Gender and Generations Survey of 2005 to illuminate our ideas. Its sample consists of 11,986 respondents (5,977 men and 6,009 women) aged from 18 to 79 years at the time of interview, but our interest is focused on the 5,847 women belonging to the Romanian or Hungarian ethnic groups. We excluded 90 Roma and 72 individuals of other ethnicities because of their very special union formation and childbearing behavior; they are too few for a reliable separate analysis. Our initial sample had 5,824 female respondents, a further 23 having been excluded for the following reasons: 16 respondents had **incomplete** marital histories, 5 had **incomplete** educational histories, 1 reported she had become a mother before age 12, and 1 spent her childhood in another country. Our final sample was further reduced to 4,055 women, since we only include women who spent some time (at least one month; the maximum period is 21 years) at ages 15-39 in our restricted period of investigation (1985- 2005). An analysis that involves age- and duration-specific fertility dependence simultaneously would require a much larger data size; essentially register data from a national population. Even when we keep the two features separate, we have needed to use curve smoothing to overcome some of the effects of random variation in our sample data. With this simple device we have been able mostly to use single-year age groups and single-year calendar periods and duration intervals.

In Sections 2 to 4, childbearing parity appears as one among several control variables in the simplest manner possible, i.e., it figures as a time-varying covariate where any effect of

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<sup>3</sup> Though the concept of the TMFR has been around for a long time, the literature on its interpretation and properties seems to be very sparse. The most insightful contributions known to us are by Rodriguez and Cleland (1988) and particularly by Hinde (1998, Section 10.3).

<sup>4</sup> This observation is in fact the basis for the study of Parity-Progression Ratios, in a tradition that goes back to Henry (1951, 1952) and Ryder (1951). An overview has been given by Ni Bhrolchain (1987); for some later contributions see Feeney (1991) and Hosseini-Chavoshi et al. (2006).

duration since last previous parity change is ignored. In Section 5 we extend that analysis to include duration since last previous birth after entry into motherhood.

Hoem and Mureşan (forthcoming) put much emphasis on showing how factors other than age (and other than union status) can be drawn into the picture, and therefore worked with control factors and with an *indirectly standardized TFR anchored* in a selected year. (The anchoring had the purpose of removing the effect of the choice of baseline levels for the control factors, an effect that would be regarded as disturbing for a neat theory.) In the present paper we are more interested in the various ways in which union status can be taken into account. It then simplifies matters if we leave out anchoring in our arguments. The ideas of the earlier paper can be used to cover such features without difficulty, as we demonstrate in Section 2 by including selected control factors in additional specifications.

All computations in this paper are for a woman's biological children. In order not to complicate matters unduly, we have ignored any stepchildren and adopted children, as is universally done in the literature we are aware of concerning total fertility rates.<sup>5</sup> (Her biological children may of course become her new husband's stepchildren if she enters a later union.)

## 2. An age-based TUF

Relying on the event-history approach with a piecewise-constant childbearing intensity, we let the intensity at age  $x$  for union type  $u$  be  $\varphi_{xu}}$  in year  $t$ . This parameter is estimated by the corresponding childbearing rate. Based on this specification we define the Total Union-type Fertility Rate for union-type  $u$  in year  $t$  as  $TUFR_{tu} = \sum_x \varphi_{xu}}$ , which we have plotted against  $t$  for four union types  $u$  in Figure 1.<sup>6</sup> To reduce the strong effect of random variation and reveal the basic structure of the curves, we have smoothed each of them by means of an unweighted seven-term moving average.<sup>7</sup>

We are struck by the following features in this diagram:

- (a) At each age, the fertility of cohabiting women does not seem to be all that much lower than for married women.
- (b) For married women there does not seem to be much difference between fertility in post-cohabitational marriages and direct marriages. This would fit with a speculation that women who marry after cohabiting are mostly those who plan to have children *in* the marriage and not before it.

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<sup>5</sup> For a recent contribution concerning stepchildren, see Holland and Thomson (2011). Thomson and Li (forthcoming) give an interesting overview of the empirical literature.

<sup>6</sup> The  $TUFR_{tu}$  is interpreted in analogy with the usual understanding of the common TFR. We pay some further attention to its interpretation in Section 3 below.

<sup>7</sup> This means that a smoothed curve value for any year is computed as the straightforward mean of seven annual values, centered on the year in question. The first and last smoothed values are means of five annual values. Two smoothed annual values are then lost in each tail of the curve, and in the diagrams we have used unsmoothed values in those positions. To avoid smoothing away interesting irregularities caused by the strong variations in Romanian public policies over previous years (Mureşan and Hoem, 2010) rather than by random variation, we have only included the years between 1985 and 2005 here.

(c) Except for minor details, the estimated fertility level is about 3 for both married and cohabiting women.<sup>8</sup> (We will show later that this is a strong overestimate.) For the never married, the level is about 0.2.

To reveal more of the information in the data, we have also carried out alternative analyses based on an event-history model with the following additional specifications:

1. We have added the following control factors: ‘parity’ (0, 1, 2, 3, and 4+), ‘character of place of origin’ (urban vs. rural), ‘union order’ (never in a union, first union vs. higher-order unions), and ‘educational level attained’ (low, middle, high, in education). The educational levels are defined as follows: low means “no formal qualifications”, middle means “high-school or vocational certificate”, and high means “degree in higher education”.<sup>9</sup> We have kept the factor ‘union type’ (never in a union, currently in a consensual union, in a direct marital union, in a marital union after a consensual union) as a time-varying covariate as in the previous analysis. We have deleted periods with other living arrangements, as before.
2. We have included an interaction between age attained and educational level.
3. For women in a union, we have added an interaction between union order and union type.

To get a simple structure, we have avoided involving period and union type in any other interaction (even with each other, as in our definition of  $TUFR_{it}$  above), and we have kept the rest of the control variables outside any interactions. After smoothing, the age-vs-education interaction has the structure given in Figure 2, which shows that in Romania, fertility decreased with increasing educational attainment in our period of observation (1985-2005), a feature that Mureşan and Hoem (2010) have previously displayed as an average for the whole period 1950-2005.<sup>10,11</sup> We now see that this feature is particularly prominent at young ages, which is puzzling since educational differences are supposed to be weak in Eastern Europe due the history of state socialism in these countries.<sup>12</sup>

For ages 28 through 34, highly educated Romanian women actually have the *highest* estimated fertility among the educational groups. In an environment of nearly universal ultimate motherhood, many of the highly educated in this age interval who have not had a child yet, will surely try to attain motherhood, which is reflected here as a rather high fertility in this group. Muresan (2009) has also shown that the negative educational gradient in the transition to motherhood reverses to a positive gradient after age 30, but only in the period 1990-2005 and therefore not before the political turnaround. Women with a low education most likely have

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<sup>8</sup> In the mean, it varies between a low of 2.93 for cohabiting women and a high of 3.17 for women who have married without earlier cohabitation.

<sup>9</sup> See Mureşan and Hoem (2010) for more exact definitions.

<sup>10</sup> Mureşan (2007 and 2009) has shown that this gradient actually appeared only after the fall of state socialism for the first birth, and only 5 years before the fall for the second birth (i.e. after 1985, which is also the starting year of the current investigation).

<sup>11</sup> We doubt that the negative educational gradient really is specific to Romania. It may extend to all countries where highly educated women find it difficult to establish themselves in the labor market. For the Czech Republic, see Kantorova (2004b).

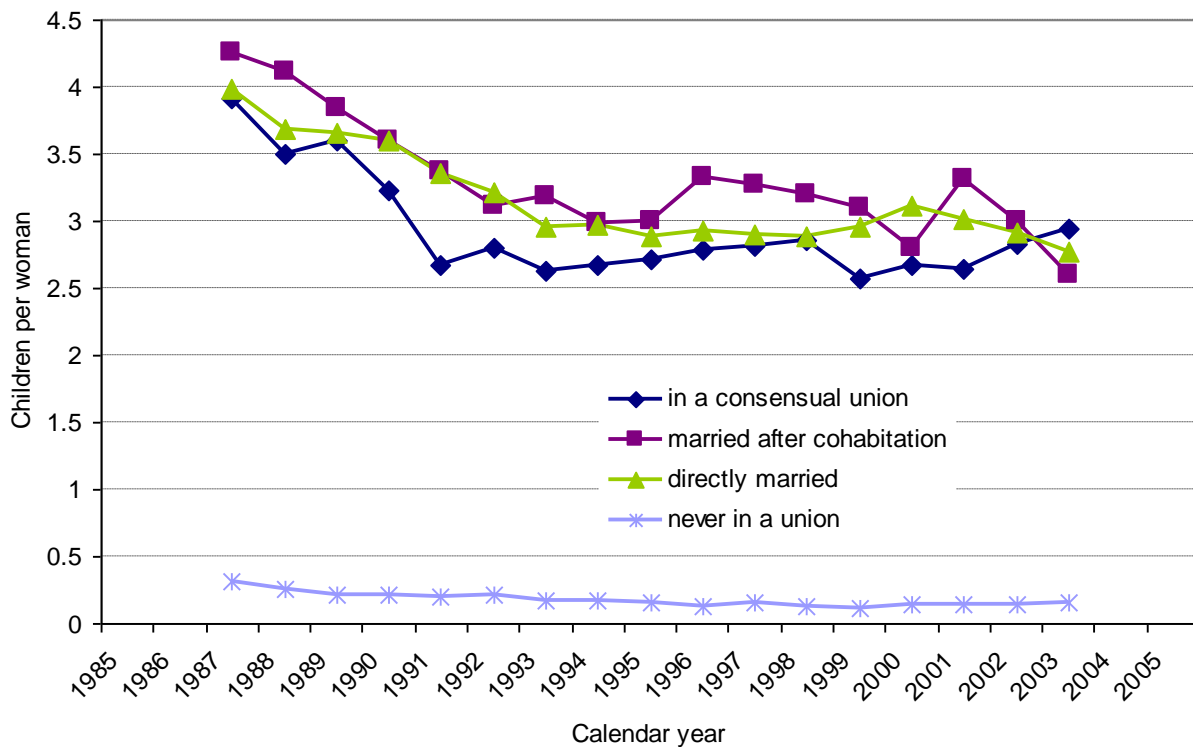
<sup>12</sup> For a nice counterexample and some theoretical argumentation see Kantorova (2004b).



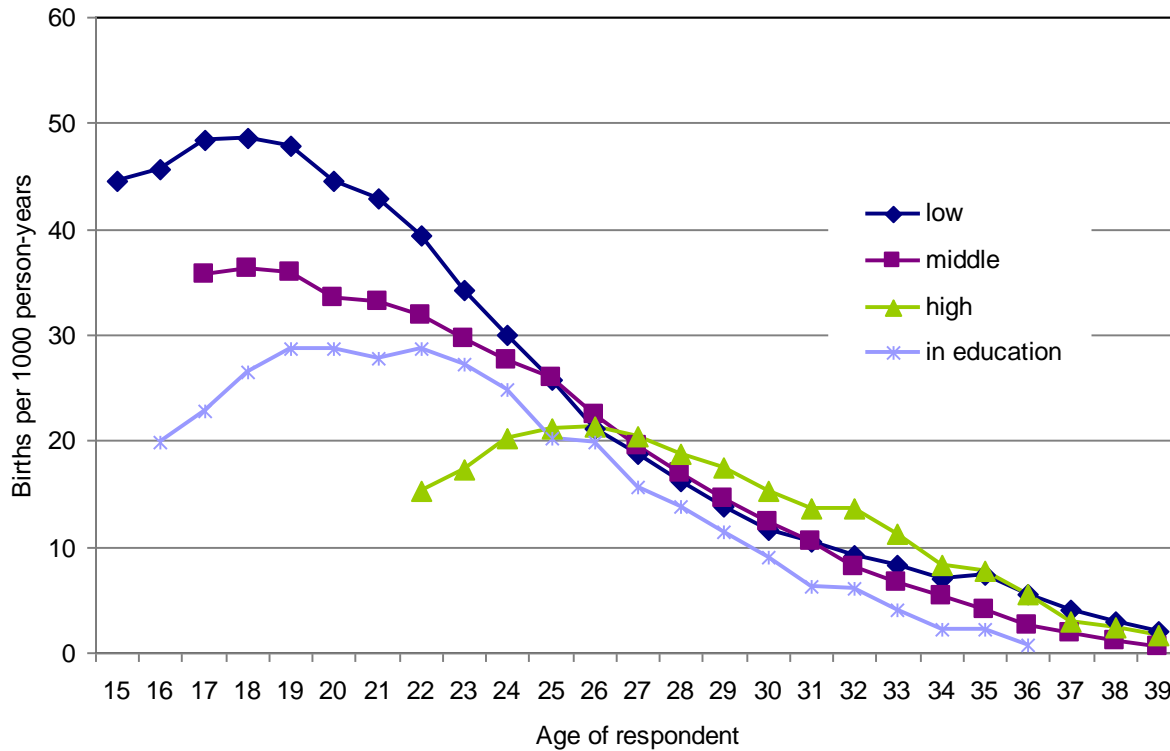
already finished their childbearing before age 28, because most women have their children relatively soon after finishing education and the less educated have completed their education at a much earlier age.

Figure 2 also shows that Romanian women appear to have quite high fertility even when they are enrolled in education. This feature may actually be misleading and may have been produced by our data construction, at least in part. Due to data limitations, we have been forced to impute educational careers in a manner that disregards any breaks in educational activity. Since the late 1990s, alternatives to daytime higher education have emerged with some regularity, so after about 2000 anticipatory bias may have interfered with our findings (Mureşan and Hoem, 2010). The best way to settle this issue will be to include complete educational histories in future data collection.

**Figure 1. Trends in age-based Romanian TUF<sub>R</sub> for four union types, 1985-2005**



NOTES: Curves smoothed by an unweighted seven-term moving average. No standardization. Computed by the authors from the Romanian GGS of 2005.

**Figure 2. Fertility rates by educational level, plotted against age attained**

NOTES: Curves smoothed by a five-term unweighted moving average. Indirect standardization with respect to calendar year, place of origin, union order, union type, and parity.<sup>15</sup> Computed by the authors from the Romanian GGS of 2005.

Unfortunately, demographers have long rightly been of the opinion that the (age-based) Total Marital Fertility Rate tends to overestimate marital fertility,<sup>14</sup> perhaps strongly.<sup>15</sup> We will show below that there is a corresponding overestimation of the TUFMR for the other union types, and that the bias even differs from one union type to another. Consequently, there seems to be something structurally wrong with the pattern in Figure 1. We hope to improve on this weakness in what follows.

<sup>13</sup> Note that the indirect standardization consists in fitting a hazard regression model with the given control variables as fixed and time-varying covariates and (in the present case) age attained as process time. Using indirect standardization avoids the need to specify a 'standard' as is used in direct standardization, and the result is superior to the latter in terms of variance minimization (Hoem, 1987, 1991). This technique has been used frequently, and recently by Gabrielli and Hoem (2010) and by Hoem et al. (2009).

<sup>14</sup> For a particularly sharp-worded expression of this position, see Pressat (1972, page 190).

<sup>15</sup> As one among many examples, we mention that the age-based TMFR for Norway was about 7 in 1959-62 and that from there to 1969-70, 1971-75, and 1976-80 it dropped by about 0.8 from one period to the next, to reach 4.6 by 1980. (Computed by the present authors from Table 3.6 in Statistisk Sentralbyrå, 1984.) Very few marriages had such high numbers of offspring at the time. (The TFR for Norway was 1.7 in 1980.)

### 3. A duration-based TUF

We have already indicated that the fertility of partnered women is likely to depend strongly on the duration of the partnership, and we display the general structure of the duration dependence in Figure 3, in which there is one curve for each age group at union formation, and in Figure 4, which has one curve for each type of union. (For a marriage converted from a pre-marital consensual union, we have used the time of marriage formation as our starting point in these computations.)<sup>16</sup> We note that fertility generally decreases as the starting age increases, but that the rates are high during the first two years of the union/marriage for all starting ages and all types of unions.<sup>17</sup> The most striking feature in Figures 3 and 4 is how systematically fertility declines with increasing duration after the first year of the union/marriage.<sup>18</sup>

For any population group of partnered women we can add up all the duration-specific fertility rates to get a TFR-like parameter, which we call the *duration-based* TUF and which will have the same type of interpretation as a regular Total Fertility Rate, namely the number of children theoretically ever born to women who survive and stay in the group, if their childbearing proceeds according to the duration-specific fertility used to compute the TUF parameter.<sup>19</sup> For instance, adding up the rates for each of the four groups in Figure 3 will give the TUF by starting age for each union type, restricted to periods in which the women are partnered (and restricted to periods where they are married for post-cohabitational marriages). For our Romanian GGS-data we have done so separately by union type for each calendar year from 1985 through 2005 (irrespective of age at union/marriage formation) and have plotted the outcome by calendar year, as in Figure 5,<sup>20</sup> which contains separate time series of TUF values for women in consensual unions and for women in our two kinds of marriages. Unlike Figure 1, this diagram shows that directly married women clearly had *more* children in the marriage than women in post-consensual marriages throughout the period 1985-2005; we speculate that women in post-consensual marriages are less oriented toward large families than women who marry

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<sup>16</sup> This means (i) that any children born before marriage are not included in the count of children *in* the marriage, and (ii) that the duration in Figure 3 is counted from marriage formation and not from union formation. Alternatively we could have used the formation of the consensual union as a starting time, but we wanted to avoid some minor corresponding complications at the present stage. Our simplification is unlikely to be important in a population where a consensual union is quickly converted to a marriage when a pregnancy occurs.

<sup>17</sup> The increase from the first to the second year of union for ages below 30 most likely reflects the arrival of children conceived shortly after union formation. We speculate that women who start their union/marriage at ages 30-34 are less fecund or less in a hurry to start a pregnancy since their curve does not have a similar peak.

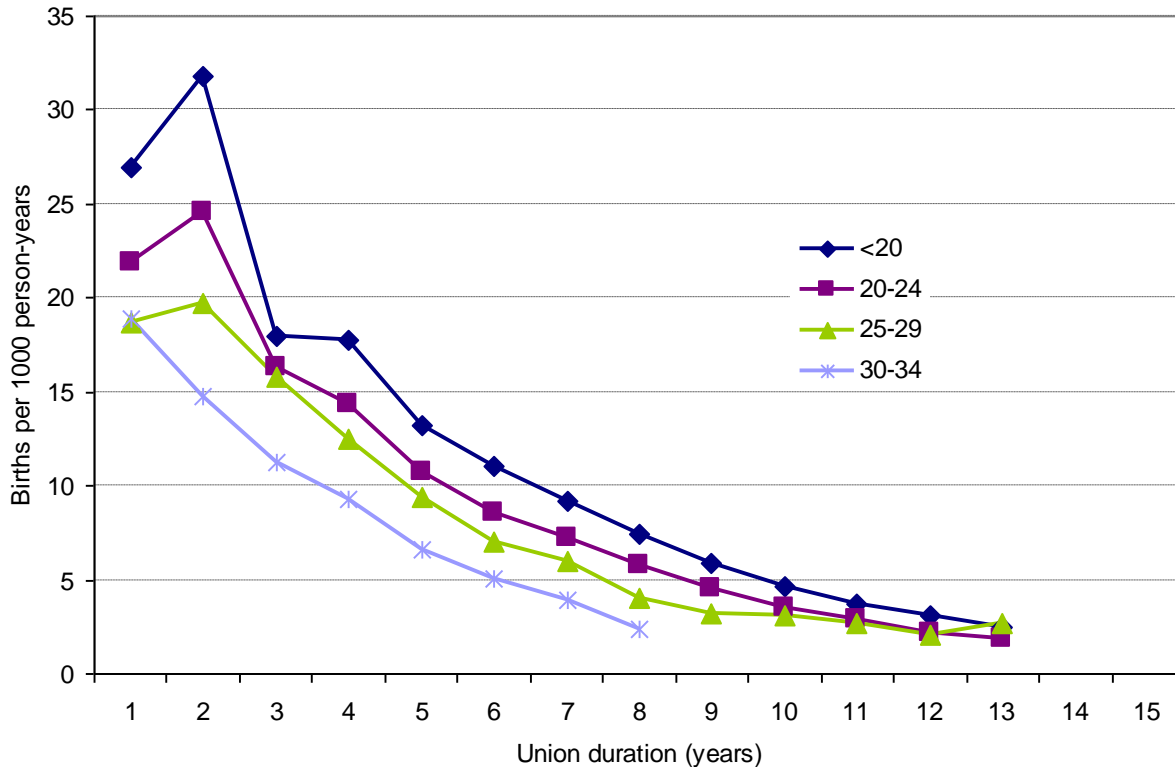
<sup>18</sup> The pattern of these types of diagrams has been quite stable over the sub-periods 1985-89, 1990-99, and 2000-2005 (documentation not shown here).

<sup>19</sup> This conceptual construct may look strange at first sight (how many women who enter a cohabitational union stay in the union without marrying for the fifteen years during which the group remains under observation as a collective?), but it ought to be no stranger than many other thought-constructs used by demographers. Just think of cause-deleted mortality concepts, and of the Gross Reproduction Rate, which is simply a mortality-deleted Net Reproduction Rate. Our purpose in suggesting the TUF is to provide an interpretable statistical parameter that can serve as a measure of the fertility level for women in a given type of marriage/union, while not suggesting that surviving for all those years in a union of the given type is a common or even frequent behavior.

<sup>20</sup> In most of our diagrams, we have used a *five*-term moving average to filter away random variation and better bring out the patterns in the curves. In Figures 1 and 5 we have used *seven*-term smoothing to get sufficiently smooth curves. We have also started the smoothing at the point for the fourth year of union duration to avoid smoothing away the pattern at brief durations.

directly.<sup>21</sup> More surprisingly, the total fertility of cohabiting women had a bulge starting in the mid-1990s, to such an extent that they had a higher fertility than married women during half a decade around the year 2000. The latter feature did not appear in Figure 1.

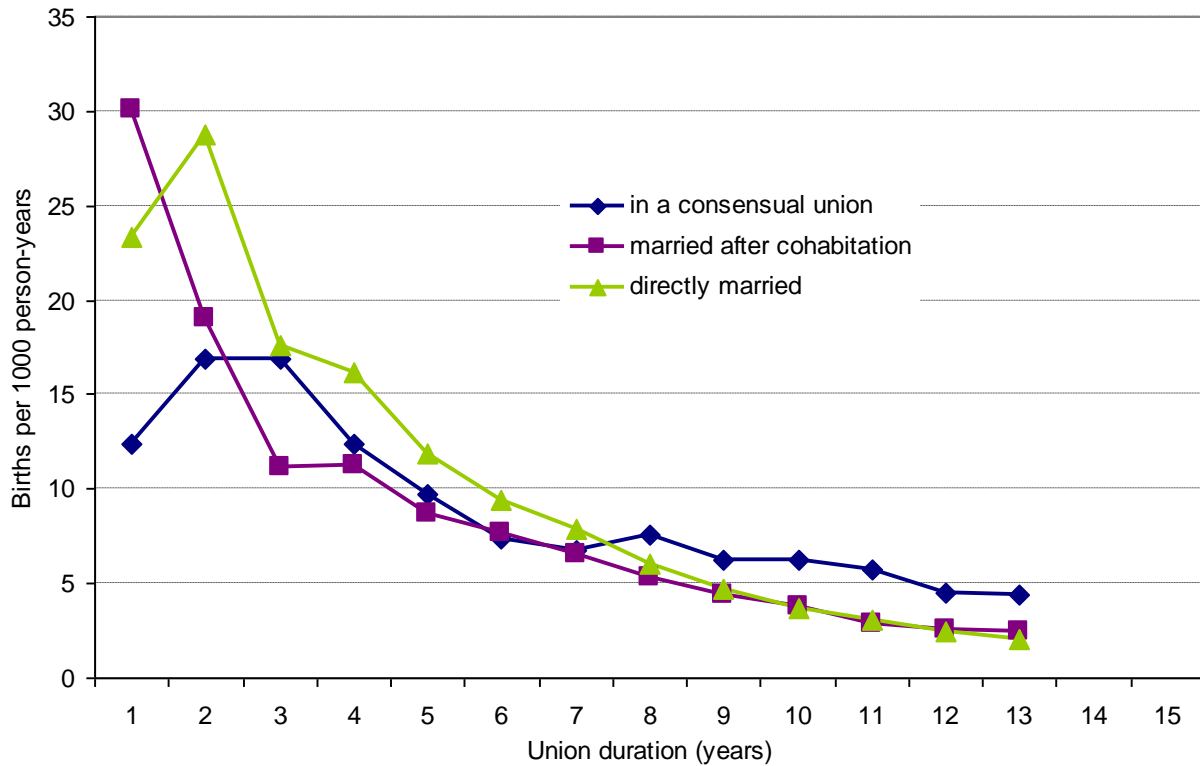
**Figure 3. Fertility rates for partnered Romanian women, 1985-2005, by union duration and age at union formation**



NOTES: Rates not standardized. Smoothed by a five-term moving average starting in the fourth year of union duration. Computed by the authors from the Romanian GGS of 2005.

<sup>21</sup> Some of the women in post-cohabitational marriages may have had children in the consensual union before marriage, i.e. not in the marriage itself. Such children are not counted in their TUFRR for the marriage, and this might reduce the child-count *in* the marriage. However we believe this to be a minor element in the difference observed. If a cohabitational union is transformed into a marriage, this is most likely to happen *before* the first birth in the union, while women who have a birth during cohabitation are less likely to convert their union into a marriage later.

**Figure 4. Fertility rates for partnered Romanian women, 1985-2005, by union duration and type of union**



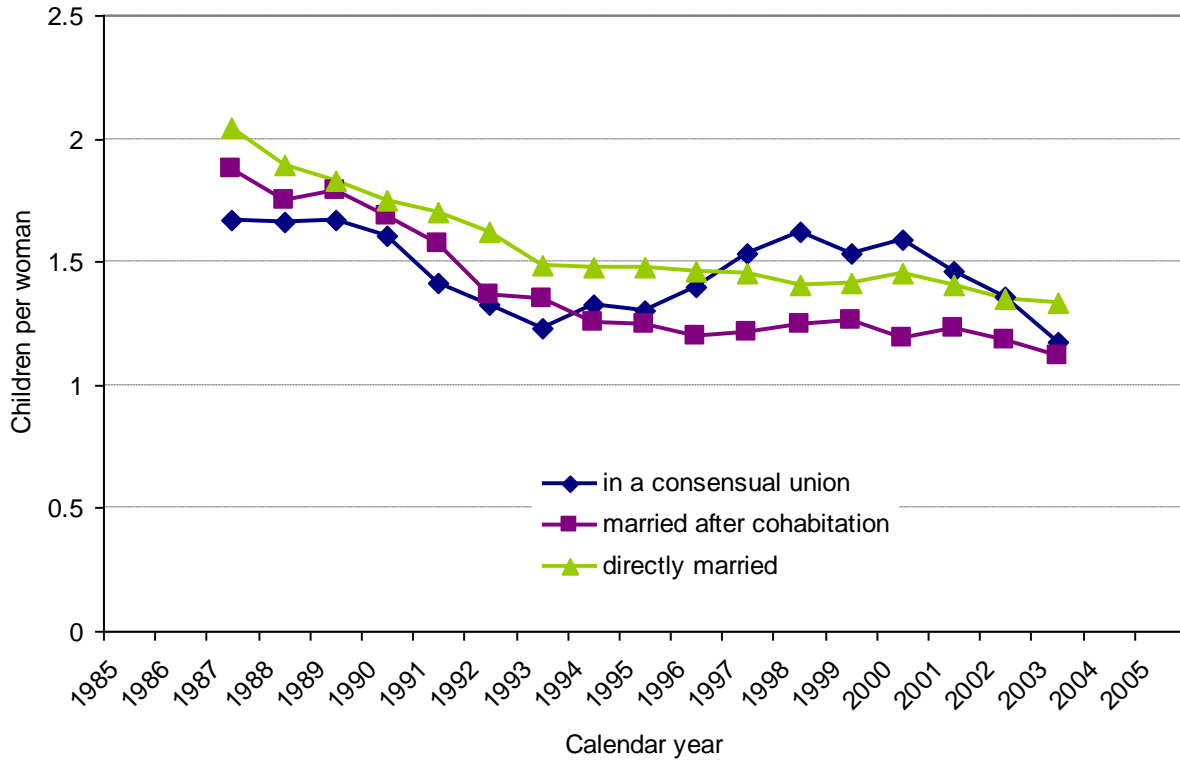
NOTES: Rates not standardized. Smoothed by a five-term moving average starting in the fourth year of union duration. Computed by the authors from the Romanian GGS of 2005. For married women, the *x*-axis gives the duration of the *marriage*. For women married after cohabitation it indicates the duration of the marriage and does not count in the preceding period of premarital cohabitation.

In Figure 5, the level of the duration-based TUFMRs for partnered women is around 1.5,<sup>22</sup> which is also roughly the level of the TFR for the same period (as given by Hoem and Mureşan, forthcoming, their Figure 5), against the level 3 in our earlier age-based computations. This demonstrates that (as suspected) the age-based levels tend to overestimate fertility in marital unions, but also in cohabitational unions, probably because age-based analysis is dominated by the very high fertility at young ages and short durations at these ages.<sup>23</sup>

<sup>22</sup> The estimated mean levels are 1.57 for direct marriages, 1.44 for consensual unions, and 1.39 for post-consensual marriages.

<sup>23</sup> See Hinde (1998, p. 124) for a particularly clear discussion of issues concerning the conventional TMFR. Like several others he suggests that it may make sense to aggregate age-specific marital fertility rates from age 20 only (or from a later age) in order to reduce the over-estimation inherent in the age-based TMFR.

**Figure 5. Trends in duration-based TUFs for partnered Romanian women, for three types of union, 1985-2005**



NOTES: Curves smoothed by an unweighted seven-term moving average. No standardization. Computed by the authors from the Romanian GGS of 2005.

#### 4. Relative risks

Except for Figure 2, the analysis reported in Sections 2 and 3 does not make use of control variables like parity, educational attainment, and union order. When we do draw control variables into the analysis, their patterns of relative risks appear as in Tables 1 and 2. These patterns closely resemble features that have mostly been reported in previous work on recent Romanian fertility behavior, such as Mureşan and Hoem (2010) and Hoem and Mureşan (forthcoming). Our tables show that the *patterns* are mostly robust against a change-over from age-based to duration-based analysis even though the fertility *levels* are badly over-estimated in the age-based approach.<sup>24</sup> The over-estimation differs by union type as a comparison of Figures 1 and 5 shows. Note that the negative educational gradient of fertility appears more clearly in the duration-based than in the age-based analysis.

<sup>24</sup> The patterns of relative risks are also very robust against changes in the selection of covariates with respect to which we choose to standardize (not documented here).

The effect of union order combined with union type appears in Table 2, which shows that in first unions, duration-based fertility is some 40% lower in consensual unions than in marriages and that in the age-based computation it is some 30% lower. Together, these findings suggest that consensual unions have become a rather common framework for childbearing in Romania. The low fertility of women who have never lived in a union suggests that single motherhood is very rare in Romania.<sup>25</sup>

Otherwise Table 2 displays the one case that we have discovered where the similarity between the outcomes of the two approaches really is questionable. In the age-based analysis, women in second or later unions have higher fertility than women in first unions, a feature that has appeared often in the literature.<sup>26</sup> This feature does not carry over to the duration-based analysis. A comparison of the two patterns for second and higher-order unions suggests that occasionally the age-based computations can produce pretty biased results.

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<sup>25</sup> See also the reflections by Rotariu (2009). Note that the age-based and the duration-based approaches coincide for never-partnered women.

<sup>26</sup> For an overview, see Thomson and Li (forthcoming).

**Table 1. Age-based vs. duration-based effects of selected control variables on childbearing intensity**

**1a. Effects of parity**

Parity	Age-based	Duration-based
0	1	1
1	0.36	0.38
2	0.14	0.19
3	0.30	0.38
4+	0.42	0.54

**1b. Effects of place of birth**

Character of place of birth	Age-based	Duration-based
urban	1	1
rural	1.14	1.17

**1c. Effects of educational attainment**

Ed. level	Age-based	Duration-based
low	1	1
middle	0.82	0.79
high	0.78	0.60
in education	0.63	0.58



**Table 2. Age-based and duration-based Romanian fertility 1985-2005.  
Interaction between the factors ‘Union order’ and ‘Union type’**

**2a. Age-based fertility**

	Never in a union	In a consensual union	Married after cohabitation	Directly married
Not in union	0.018			
First union		0.705	1.076	1
Second or later union		1.745	1.521	1.786

**2b. Duration-based fertility**

	Never in a union	In a consensual union	Married after cohabitation	Directly married
Not in union	0.017			
First union		0.614	1.020	1
Second or later union		1.035	0.824	1.054

NOTES FOR TABLES 1 AND 2: The age-based coefficients are for the specification in Figure 2. The duration-based coefficients are for a specification with the factors calendar year, age at union formation, educational attainment, and character of place of origin, as well as for union type in interaction with union order, in addition to parity. Computed by the authors from the Romanian GGS of 2005. The table entries are “risks” relative to the baseline indicated by a table entry of 1 (without decimals).

**5. Separate analysis by birth order**

In our account so far, parity attained has just been a control variable which as usual has appeared in a multiplicative manner in our underlying model specifications, as in the statistical analysis on which Figures 2 and 3 are based. Its effects then appear as relative risks that depend to some extent on how the underlying event-history model is specified, but like other risk patterns they are quite robust against model re-specification. The typical pattern is apparent in Table 1a. It was discussed previously by Hoem and Mureşan (forthcoming).

To use parity as a straightforward control variable is sufficient for a discussion of Total Union-type Fertility Rates, as in Section 2 and 3 above. Nevertheless, it would be more satisfactory to pay closer attention to the parity factor, particularly as the standard demographic approach would be to analyze each birth order separately, and we now turn to this approach. This will open the possibility of systematically taking into consideration new time factors such as *age attained* in the study of the first-birth intensity and *time since last previous birth* in the study of births of higher orders, typical features of the analysis of parity progressions. At the same time we want to retain *union duration* (as in Figures 3 and 4) and *calendar time* (as in Figures 1

and 5), and we want to account for all of these factors simultaneously.<sup>27</sup> This would amount to operating with several simultaneous clocks, the effect of each being represented by a linear spline function (say), a possibility that has been made commonplace by Lillard and Panis (2003).

For this purpose we have estimated separate models for first, second, and third births, and since we are mainly concerned with the effect of union type at this point, we display the corresponding duration splines in Figure 6.1 through 6.3.<sup>28</sup> (As before, “union duration” remains the duration since marriage formation for post-cohabitational marriages.) In this connection we have also included the (low) fertility of those who are currently no longer partnered after having left a marital or non-marital union (called “in between unions” in our diagrams),<sup>29</sup> as before we have also included the effect of never having lived in a marriage or a consensual union. The disk representing “never in a union” in Figure 6.1 is centered one unit above the origin on the ordinate axis. All other curves in Figure 1 have starting points for risks that are computed relative to this one. For instance, the starting point of the curve marked “married after cohabitation” is some 25 units higher on the ordinate axis, indicating that the corresponding risk is about 25 times as high as for women who never lived in a union.<sup>30</sup> For the directly married, for those who have left a union without entering another (the disrupted, separated, divorced, or widowed), and for women who have just started a consensual union, the initial relative risk is about 4 in each case.

In Figures 6.1 through 6.3, the subsequent effect of union duration is represented by a linear spline. Figure 6.1 is standardized for age attained and calendar year (both in a linear spline representation). Figures 6.2 and 6.3 are standardized for age at first birth, duration since last previous birth, and calendar year (again all in a linear spline representation). All diagrams are also standardized for educational attainment and for character of place of birth.

For partnered women, the first-birth intensity has a distinct peak a year after union formation. Once they have entered a union (and in particular a marriage), these women will also have a child. Union formation and childbearing are not strongly “separated” as in most Western countries. For post-consensual marriages the pattern is the same, but it starts even earlier with marriages of already pregnant women (Mureşan 2008); many first births appear shortly after the conversion of a consensual union into a marriage.

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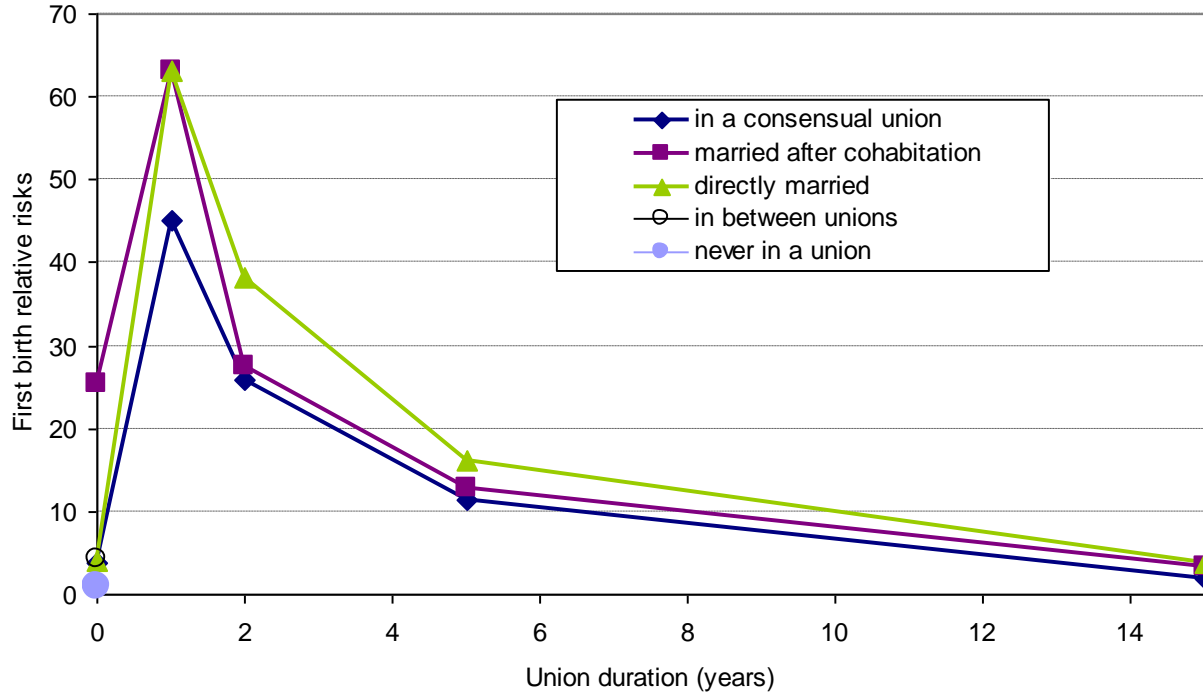
<sup>27</sup> We could also include time since completion of current educational level, as has been done by Kantorova (2004a, 2004b), Skirbekk et al. (2004), and possibly others, but we will leave this possibility aside in the current context.

<sup>28</sup> As in the rest of this paper, we have computed these estimates from the data of the Romanian GGS of 2005. More details and the rest of our empirical results are available from the authors on request.

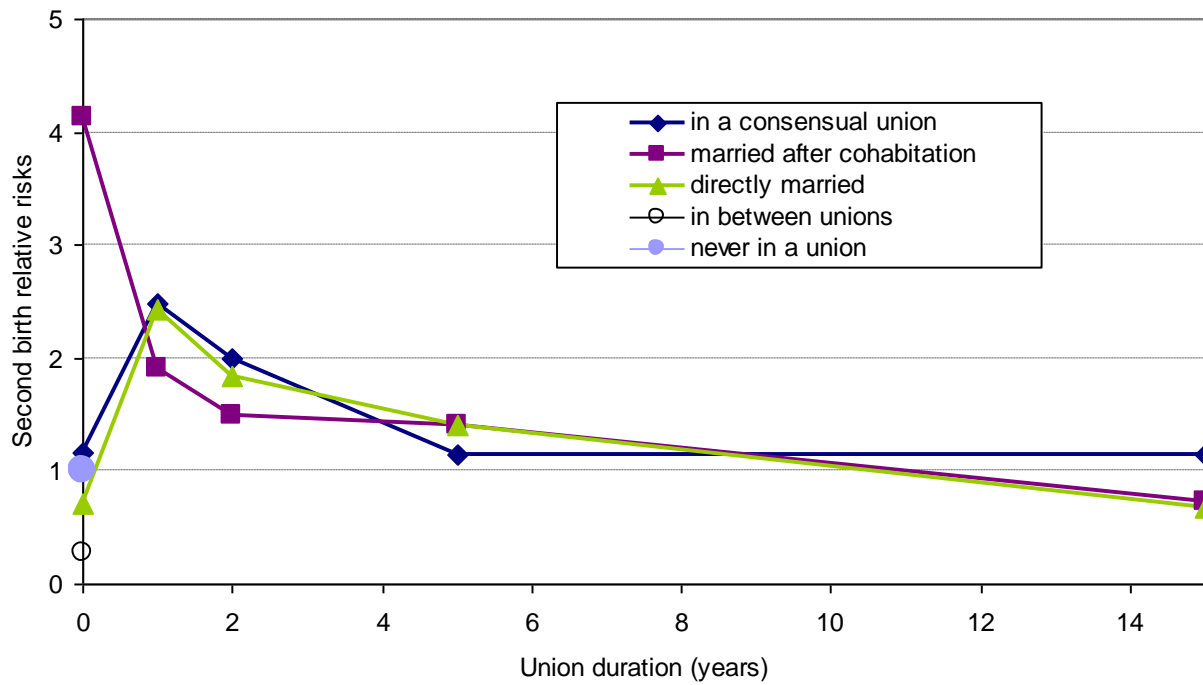
<sup>29</sup> A period “in between unions” starts at the end of the last previous union and ends at the formation of a new union or at censoring. (To call it “in between unions” may actually be a misnomer since the period need not end by the formation of a new union.) There are so few births “in between unions” that we have not estimated a childbearing risk that may change over time but have assumed that it is constant until the start of any new union.

<sup>30</sup> Note that (as we have discussed above) this particular rate is computed (and plotted) by the duration of the *marriage*, not including the duration of the pre-marital consensual union. The same comment pertains to Figures 6.2 and 6.3.

**Figure 6.1. First-birth duration spline for time since age 15, by union type**



**Figure 6.2. Second-birth duration spline for time since union formation, by union type**



**Figure 6.3. Third-birth duration spline for time since union formation, by union type**

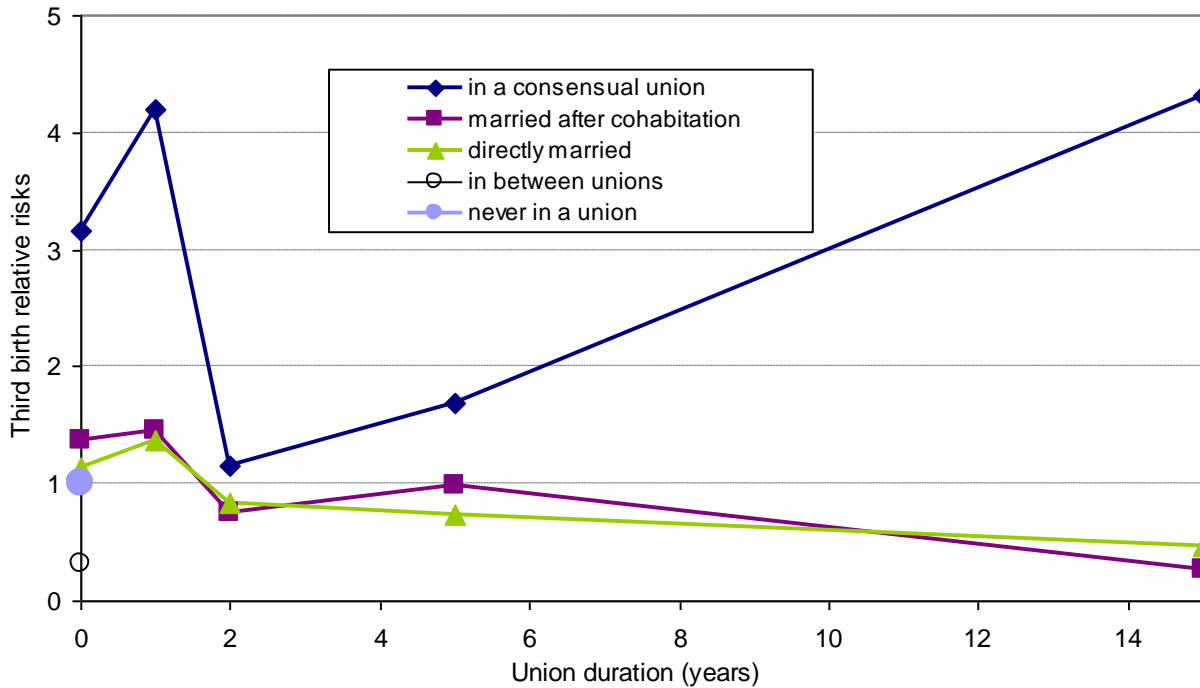


Figure 6.2 can be read similarly. (Note the very different y-axes in the two diagrams.) The disk for those who have never lived in a marital or non-marital union is at level 1 on the y-axis, as it is in Figure 6.1. The circle for women “in between unions” is at a level about one quarter above the y-origin. The post-consensually married start at a relative risk of a little over 4, indicating that they have a markedly higher immediate risk of second childbearing than the other groups right after marriage formation (but not nearly as high as for first births). Again, marriage and childbearing are very closely related.

For third births (Figure 6.3), the story is similar. Women “in between” unions have a relative third-childbearing risk of around one-third. The third-birth risk for cohabiting women is much higher than the risk for the other groups. Presumably, in most cases the first two children were born in previous unions (or outside of any union). We have two explanations for the rather steeply rising right tail of the curve for cohabitants with two children. *First*, the “old fashioned” consensual union has mostly been seen as a partnership for people who were very poor (so could not afford the cost of a wedding) and/or who have had very little education (so without the means, ability, or possibility to control their reproduction, and possibly with a fatalistic attitude to childbearing). In any case, one could expect their fertility to be high relative to other women, and progressively more so as their union matures. *Second*, there are not many individuals in this part of the GGS sample, so our estimates must be strongly influenced by random variation. In fact the slopes for their duration spline are not significantly different from 0 when significance is computed in the conventional manner.

## 6. Discussion

In the present paper we have noted existing methods for estimating the impact of living arrangements on levels of childbearing and have developed the methods further by embedding them in an event-history analysis and by including additional types of living arrangements. We have extended the theory of the TFR by now taking union type into account, and in some sense we hope to have restored part of the honor of the ill-reputed TMFR by showing that in the Romanian data most of the patterns of relative risks for fixed and time-varying control variables are robust against the change from an age-based to a duration-based approach. If such robustness can be shown to be present in data from populations other than the Romanian, then fertility patterns may be explored more easily for correspondingly more populations. Note that we do not claim that good estimates of fertility *levels* would also be more easily available by the age-based approach, only that they will result from the duration-based version. The criticism against age-based estimates of fertility levels remains valid because the age-based approach does not account for the effects of the duration patterns of union-type fertility. In particular, it therefore does not capture direct effects of the high fertility at short union durations nor the reduction of fertility at long durations caused perhaps by fecundity problems in many cases.

At least one feature of Romanian fertility is probably particular to this and closely similar populations, namely the very low childbearing outside of unions. For such sub-groups “duration” is re-defined as “duration of the current living arrangement”, which might mean “duration since age 15 (say)” for women who never were in a union, and “duration since the end of the last previous union” for women “between” unions.<sup>31</sup> We believe that there is much more childbearing outside of stable unions in the United States and in the West Indies, say, and that non-marital childbearing in Germany has a much stronger non-union element than in Romania. Among items that remain open, therefore, is how many of our empirical findings (on pattern robustness, say) can be generalized to other populations, or to populations in general.

Finally, one “technical” item: Using a smoothing technique no more difficult than simple few-term unweighted moving averages has enabled us to get pretty sharp age and period profiles from computations using single-year age and duration groups and single-year calendar periods. We have seen no need to combine ages, durations, and periods into longer groups.

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<sup>31</sup> We have actually made use of the first of these definitions in part of our work for the present paper, viz., in the computation of the TUFMR values for those never in a union, given as the lowermost curve in Figure 1. We have then operated with much longer durations for women in this group (in order to cover all fertile ages) than the duration span for partnered women (as in Figures 3 and 4). For the never-partnered women, there is no difference between proper age-based and the duration-based computations.

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