Traces of the Second Demographic Transition in four selected countries in Central and Eastern Europe: Union formation as a demographic manifestation

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Traces of the Second Demographic Transition in four selected countries in Central and Eastern Europe: Union formation as a demographic manifestation

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Abstract. Using data from the first round of the national Gender and Generations Surveys of Russia, Romania, and Bulgaria, and from a similar survey for Hungary, which were all collected in recent years, we study rates of entry into marital and non-marital unions. We have used elements from the narrative of the Second Demographic Transition (SDT) as a vehicle to give our analysis of the data for four countries some coherence, and find what can be traces of the SDT in these countries. The details vary by country; in particular, latter-day developments in union formation patterns did not start at the same time in all countries, but in our assessment it began everywhere before communism fell and thus before the societal transition to a market economy got underway in 1990.

1. Introduction

In this paper we study trends in family-formation behavior since 1960 in the countries that used to be called the Eastern Bloc. In this connection the account of the Second Demographic Transition (SDT) is very attractive, both as a generalized summarizing description and because of its underlying theory of value change in the direction of increasing tolerance in family matters and of women's increasing autonomy (Lesthaeghe and van de Kaa, 1986; Lesthaeghe and Surkyn, 2002; for a recent, independent assessment, see Sobotka, 2008). The SDT account consists of a narrative of changes in demographic behavior and of an explanation for those changes. The changes on which the narrative focuses are a decline in marriage formation, an increase in non-marital cohabitation, a general decrease in fertility (particularly at higher birth orders) but an increase in non-marital childbearing, an increase in union disruption, and a postponement of marriage and childbearing. Briefly stated, the explanation given is that these developments are caused by ideational changes regarding family life and childbearing, i.e., changes in norms, values, beliefs, and attitudes, sometimes operating in tandem with political, economic, and social changes.

There is ample evidence of most of the demographic developments in the SDT narrative all over Europe, particularly concerning fertility trends; see for instance Frejka et al. (2008). There is also already quite some literature on recent changes in union-formation behavior in Central and Eastern Europe and on their interpretation (Carlson and Klinger 1987, Lesthaeghe and Surkyn 2002, Aassve et al. 2004, Spéder 2004 and 2005, Zakharov 2005, Gerber and Berman 2005, Koytcheva 2006, Thornton and Philipov 2007, Kostova 2007, Mureşan 2007, and Bradatan and Kulcsar 2008).[1] Too little attention has been given so far to the finer structure of the trends in union-formation risks in the region, however, which is surprising, given that a growth in non-marital cohabitation can only come about through ideational change and therefore is a prime indicator of the very explanation given for the SDT. In the present paper we focus therefore on the growth of non-marital cohabitation as a competitor to conventional marriage. Our account is for Bulgaria, Romania, Russia, and Hungary, for we are fortunate in having early access to the data from the first round of the Gender and Generations Surveys (GGS) in the first three of these countries and to their close counterpart for Hungary.[2] All surveys have used a random sample of women and men of all relevant ages. In the present study we use the data for women only; for sample sizes (in terms of years of exposure) see Table 1.

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[1] A Russian-reading colleague has also made us aware of Maleva and Sinyavskaya (2007).

[2] For a description of the GGS program, see Vikat et al. (2007); for a description of the Hungarian survey, see Spéder (2001).

(Table 1 about here)

We started this investigation in a descriptive mood and without any strong preconceived ideas or hypotheses about union-entry trends, but with a series of open questions. We were curious to see to what extent the fall of communism around 1990 might have given a particular impetus to developments in union formation across the four countries, and what commonalities we could find in the patterns of such developments. The single-country background and previous literature has been described succinctly for Russia and Bulgaria by Philipov and Jasilioniene (2007),[3] for Romania by Mureşan (2007),[4] and for Hungary by Spéder (2005). The three former authors have provided extensive survival tables for Russia, Bulgaria, and Romania in the spirit of Andersson and Philipov (2002), who gave such tables for Hungary and fifteen other European countries for an earlier period. Following Carlson and Klinger (1987), Spéder (2004, 2005) maintained that post-divorce non-marital cohabitation has old roots in Hungary and that consensual first unions gained considerable ground in that country well before the regime change. We focus on first unions and find similar patterns also for Russia, Romania, and Bulgaria.

Table 2 contains some highlights for the three GGS countries for which period survival tables are available,[5] and we see that there was considerable cohabitation around in the late 1980s already and that in Bulgaria and Russia it had outflanked direct marriage at least by the early 21st century. According to this table, Romania seems to be in a different category, where marriage had held up much better that in Bulgaria and Russia. Statistics like those of Table 2 have been derived from straightforward occurrence/exposure rates, with no standardization or other attempt at hedging against compositional effects, so we started out wondering to what extend the considerable differences between the countries would hold up to closer scrutiny.

(Table 2 about here)

2. Method and covariates

In demography, one of the ways to handle compositional effects is by using standardization, and we have applied this method in the form of an unusual variant of intensity

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[3] See also Zakharov (2008) and Koytcheva and Philipov (2008).

[4] See also Mureşan et al. (2008).

[5] We have not found comparable information for Hungary, also because Spéder (2005 etc.) has worked consistently with birth cohorts rather than with calendar periods.

regression where entry into marriage and into a non-marital union are studied jointly as competing risks in a manner that permits direct comparison between the two types of union formation in each of the four countries.[6] This procedure has been described most fully by Hoem and Kostova (2008), to whom we refer for mathematical aspects of the approach we use. (We give some further discussion of such items in an appendix to the present paper.) They also gave a first application to the Bulgarian GGS data already. This paper can be regarded as a further elaboration of the Bulgarian data and an extension to the three other countries we also have data for.

Based on data in a monthly format for the years since 1960 we have used proportional-hazards event-history analysis with a piecewise constant baseline intensity to reflect the impact of a woman's age, formally using the decrement as a fixed covariate in addition to the other fixed and time-varying covariates available to us (the determinants). Among the determinants we have included a time-varying covariate that we call pregnancy-and-parity status. It provides a differentiation between (i) non-pregnant childless women, (ii) pregnant childless women,[7] and (iii) mothers, i.e., women at parities 1 and above. The first of these three groups overwhelmingly dominates the exposures to the risk of first-union formation (Table 1) and we report most of our results for this group alone. Since our focus is on the changing trends in union formation, we display the interaction between calendar time and union type in our descriptions below, and let the other available covariates appear as control variables. These are most importantly (self-reported) ethnicity, but also a number of covariates that are intended to reflect other aspects of the respondent's background, namely whether she grew up in an urban or rural region, whether she lived with both parents at age 15, her number of siblings, her own educational attainment, and the educational attainments of her mother and father.[8] These are standard covariates readily available in our data, except the respondent's own educational attainment. We would have used it more extensively if we had had enough information to make it a genuine time-varying covariate, but the data only contain the attainment made by the time of data collection plus the time at which the respondent had reached this level of education (according to her own report), so we have had to impute a non-fixed covariate using a method developed by Hoem and Kreyenfeld (2006). Since this is not the real thing, we do not report the outcome here, nor do we

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[6] We also call the type of union formed (marital or non-marital) a <u>decrement</u>, in line with actuarial practice.

[7] In line with much demographic practice, we have counted a woman as pregnant during the seven calendar months just before the month in which she gave birth. We have used seven instead of nine months to allow for the fact that few women can be sure they are pregnant in the first one or two months of a pregnancy.

[8] Some of these covariates have not been available for Hungary. For Romania we did not include the parents' educational attainments because of data-quality problems.

report the risk patterns for our other control variables, mainly in order not to detract attention from our main focus on union-entry trends, but also because they do not contain any really notable surprises, particularly since Bradatan and Kulcsar (2008) went this way before us. Among the findings that we do report is a strong drop in the marriage-formation risk in all four countries and a counterpart increase in the risk of entry into non-marital unions, though surprisingly in Bulgaria (and possibly Hungary) this increase turned into a drop at the beginning of the 2000s. (For the same feature see also Hoem and Kostova 2008, Figure 3.)

To give a feeling for the size order of the relative union-formation risks in our four data sets in the twilight years of state communism, we attach Table 3, where for each country we display the (two-way) empirical interactions between the type of union formation (marital and non-marital) on the one hand and pregnancy-and-parity status on the other. The estimates have been produced by an intensity regression where age and calendar time appear formally as (time-varying) control variables not involved in any interactions, so the figures represent a kind of average over active childbearing ages and over the forty-odd years since 1960.

(Table 3 about here)

The general pattern is that as long as a woman was childless and not pregnant, the risk of entry into a non-marital union most often was low by comparison to the risk of marriage formation. Bulgaria constitutes an exception, in that entry into cohabitation was the higher. (We return to this deviation below. Note that our method allows for a direct comparison of the union-formation risks across the two types of unions in each country.) Not surprisingly, the union-formation intensities increased strongly if the woman became pregnant, and the increase was particularly strong for marriage formation. If she did not form a union before she had her (first) child, then the entry intensities largely went back to the size order they had before she became pregnant, or even to something smaller. It is as if the arrival of the first child is some kind of watershed, after which the woman was less attractive as a partner, or alternatively that the remaining women were less attracted by partnership. Only in Hungary, mothers still ran a (somewhat) higher risk of entry into a union, especially a marital union, than before they became pregnant.

3. Trends over the years since 1960

To get closer to the changing dynamics of union formation we report the trends in (standardized) entry rates since 1960 in Figure 1, computed separately for each of the four countries. These are relative risks of entry into cohabitation and into marriage for childless non-pregnant women[9] in a two-way interaction between calendar period and decrement type, standardized for the control variables listed above. The basis of comparison is the country-specific risk of entry into a marital union for childless non-pregnant women in 1960-64.[10,11]

[Figure 1 (concerning standardized entry rates for the two competing risks) about here]

The following patterns strike the eye:

In *Bulgaria* and *Hungary*, marriage risks have decreased over time ever since the early 1980s (roughly); in *Russia* they have decreased strongly since half a decade later, and in *Romania* since another half a decade later. In all countries the risks of entry into non-marital unions have increased ever since the 1960s, much as one would expect from descriptions of the Second Demographic Transition. Taken together, these manifestations started well before the fall of communism, particularly for entry into consensual unions. Developments of this nature have been noted before by Gerber and Berman (2005) and by Spéder (2004, 2005).

Bulgaria seems to be a case of its own. As we just said, the marriage risk fell since the early 1980s, but the entry risk for cohabitation stabilized in the 1980s and 1990s. If anything, it *dropped* after the turn of the century. This looks like a deviation from (standard) patterns in the Second Demographic Transition, though one should note that the cohabitational entry risks continued to increase *relative to* the marriage risks through our whole period of observation[12].[13]

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[9] Because of the exposure dominance of the non-pregnant childless women, the interaction would not have been much different if we had disregarded pregnancy-and-parity status.

[10] The diagram for Bulgaria deviates somewhat from the corresponding diagram in the paper by Hoem and Kostova (2008) because their computations were for cohorts born in 1955 and later while the present diagram is for *all* cohorts in our data, as it is for all our countries.

[11] We have further experimented with an intensity model that also contains (i) an interaction between the type of union formation and age attained as well as (ii) an interaction between union type and the control covariates. We have relegated an account of the mathematics involved to our appendix, which also contains a discussion of the items plotted in Figure 1 and the subsequent Figure 2 in terms of relative risks.

[12] The entry risk for cohabitation relative to the corresponding risk of marriage formation in Bulgaria rose steadily as follows:

1960-64	1965-69	1970-74	1975-79	1980-84	1985-89	1990-94	1995-99	2000-04
0.5	0.72	0.78	1.07	1.17	1.31	1.77	2.81	4.44

Romania is another exception from the general trend in the risks of entry into cohabitation, relative to that of marriage formation. Even if the process of first union formation largely follows the trends observed in the other three countries, marriage was the dominating type of first union throughout the entire period of observation.

If we add an interaction between age attained and decrement (union type) in the intensity regression that produces the standardized risk trends mentioned above, we get age profiles for the two entry risks as an extra bonus (Figure 2). (For the mathematics, see our appendix once more.) We had expected entry into cohabitation to be shifted toward younger ages than the age profile for marriage formation, much as in the diagram for Bulgaria, but the diagrams for Russia, Hungary and Romania shows how incorrect such a preconception can be.

[Insert Figure 2 (national age profiles of entry risks) about here.]

4. Shifting age profiles

The findings presented in Section 3 provide a neat and compact description of entry trends in our four countries, based on a standardization technique of a type that is ubiquitous in demography.[14] Standardization is known to summarize risk trends and differentials well under wide conditions, and to be robust against mild deviations from those conditions.

One of the conditions that we have not addressed above is the assumption of a stable age profile in the risks, i.e., we have behaved as if each of the two piecewise constant baseline hazards (one for each decrement) were the same for all calendar periods in the analysis. This may have simplified matters unduly; after all many authors document to their satisfaction that there has been a delay in union formation, so marriage and perhaps entry into cohabitation occur progressively later in life as calendar time increases. One question is, therefore, how robust the results above are against what may be a misspecification.

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[13] There is a hint of a drop in the entry risk for cohabitation between 1995-99 and 2000-2001 in Hungary as well, but we do not pay much attention to this since the latter period is only two years long in Hungary as against five years for other periods and countries. Random variation may therefore play a greater role than otherwise at the tail end of the curve for Hungary.

[14] Using hazard regression in our situation is just a practical manner of applying indirect standardization.

To check on this question we have estimated the hazard parameters once more, but now with a *three*-way interaction between age, period, and decrement.[15] The outcome is given in Figure 3, where to avoid needless complication we have temporarily used five-year age groups and have concentrated on the years between 1980 and the survey date.[16] For each country we have plotted the age profiles of the rates of union formation for each period *k,* and we get the following graphical patterns, which can serve as a simple optical goodness-of-fit test of our basic specification.

[Figure 3 (age-period-decrement profile) about here]

For *Hungary* the entry risks have indeed shifted steadily toward higher ages. For *Romania* we see a bit of a shift towards later ages in the risk of entry into marriage,[17] while in *Bulgaria* we can see a similar shift in the risk of entry into non-marital cohabitation. With some good will one can even discern some tendency for the profile to shift a little toward *younger* ages in the risk diagrams for Russia. All in all, perhaps there is only a mild deviation from the requirement of a stable age profile in *Bulgaria, Russia*, and *Romania*.

By way of conclusion, to get a realistic representation it looks as if we may be able to make do with our original intensity specification for Russia, Romania, and Bulgaria, but not necessarily for Hungary. For the latter country we have therefore tried the specification with a three-way interaction between age, period, and decrement once more, but now with our finer age specification and with periods back to 1960. The result is that for each age group we can essentially draw a diagram like that of the corresponding panel in Figure 1 (details available from the first author). In our view, therefore, the whole story of the entry trends in Hungary since the 1960s is adequately represented in Figure 1 in any case. Except for details we draw the same conclusion concerning the intensity age profiles in Figure 2.

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[15] In the mathematical terms of our appendix we specify $\mu = ACD + B$ and see whether we can actually decompose ACD as $AD + CD$ in our data. In a different formulation this means that instead of a simple model $\mu_{ijkh} = a_{ih}b_j c_{kh}$ we fit $\mu = a_{ikh}b_j$ and see how realistic it is to suppose that a_{ikh} can be split as $a_{ikh} = a_{ih}c_{kh}$ (if we allow ourselves some lenience in mathematical representation).

[16] The use of five-year instead of the shorter age groups is intended to avoid an overly strong impact of random variation. An extension back to 1960 and the use of shorter age groups essentially gives the same picture (not documented here). The results remain standardized with respect to the control variables in our analysis.

[17] For Romania, we see a shift in the risk of entry into marriage towards later ages also in 1990-1994. This corresponds to what one sees in official statistics: the mean age at first marriage increased from 22.1 in 1989 to 22.8 in 1990 and then went back to 22 years in 1991 (Generations and Gender Contextual Database). Thus, it was a phenomenon that lasted only a year, nevertheless it is caught by the GGS data.

5. Conversion of non-marital into marital unions

As we mentioned toward the end of Section 3, we have found that lately the risk of entry into cohabitation has dropped some in Bulgaria. To see whether this means that Bulgarian women have given up on the Second Demographic Transition, at least as far as union formation is concerned, it pays to introduce an additional dimension, namely the conversion of non-marital unions into marriages. One take on this is our Figure 4a, which is similar to a corresponding figure presented by Hoem and Kostova (2008, Figure 4), but which is now constructed in a way that covers the whole period and the entire population of the present study. In Figure 5 the same data are seen from a different angle, but it tells the same story, namely that the SDT remains in progress in Bulgaria. Here is some further background information.

[Figures 4a and 4b about here.]

Consensual unions seem to have been entrenched in Bulgaria for a long time. (Note how high the Bulgarian curve for entry into cohabitation is in Figure 1.) Anecdotal evidence suggests that there may have been a long-standing pattern where couples who are engaged to be married, move in with one set of their parents and then marry only subsequently, when this fits the family economy and other practical circumstances (observation by Kostova 2007). [This fits well to the quick conversions of consensual to marital unions noted by Koytcheva (2006, Section 7.1.1) based on Bulgarian data sets different from the GGS.] In our data, this would be recorded as an entry into a consensual union and a later conversion of the union into a marriage. Figures 4a and 5 show that after the fall of communism, the conversion activity was scaled down considerably. A consensual union became a much more durable arrangement, fully in agreement with what a description of the Second Demographic Transition would predict. Figure 4b extends this painlessly to Hungary, for which as we remember we have found a similar drop in the two years right after the turn of the century (Figure 1). Extensions to the data for Romania and Russia largely show the same pattern for conversion risks (not documented here).

[Figure 5 about here.]

6. Summary and reflections

The union-formation trends that we have revealed in this descriptive study of four countries in Central and Eastern Europe turn out to have several features in common. Marriage formation has dropped in all countries since the fall of communism, and sometimes before.

Consensual unions have gained ground all the time until the end of the twentieth century, and only in Bulgaria and Hungary does popular interest in consensual-union formation seem to have been reduced somewhat thereafter. In all four countries, the wind has gone out of the sails of conversions of consensual unions into marriages; so non-marital unions have progressively stayed consensual longer.

Despite all commonalities, it is evident that the Second Demographic Transition, of which we have found some traces, is *not* a unitary movement that reached all countries in Central and Eastern Europe roughly at the same time and had the same features throughout, no more than it was in Western Europe. If anything, such a transition did not start simultaneously in all of our countries, and above all it began well before the fall of communism and before the societal transition to a market economy got underway around 1990. If we take the distinct drop in marriage formation as a main marker of the start of the Second Demographic Transition as we study it in this paper, then a rough estimate would be that it started in Hungary and Bulgaria after the early 1980s and in Russia and Romania half-a-decade and a full decade later, respectively. Such differences should fit with the economic and social developments in the countries, but establishing such a correspondence is a matter of future research.[18] In particular the special trends in Bulgaria (and possibly Hungary) need further investigation, most likely by bringing in further dimensions in the analysis. We doubt that it will be enough to continue to study standardized trends in decrement-specific union formation. In any case our empirical findings have put similar observations made by Lesthaeghe and Surkyn (2002), Gerber and Berman (2005), Zakharov (2005), and Spéder (2004, 2005) on a firmer empirical ground than before.

As a final reflection on our findings we want to underline that interpretations should be made with some prudence, for it is possible that the perception of what constitutes a consensual union has varied across countries and has changed over time, and also that reporting inaccuracy may have exaggerated the early growth of entry risks for consensual unions. Briefly, the reporting accuracy depends on the respondents' ability to recall and willingness to reveal cohabitational episodes. It is possible that cohabitational episodes that occurred long ago may have been forgotten or suppressed more often than more recent episodes,[19] and if this is the case, cohabitational behavior at the beginning of our period of observation may have been more

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[18] We have been surprised by the relatively low rates of entry into consensual unions in Hungary. From general impressions we would have expected this country to be on the liberal side among the countries that we analyze, but it does not seem to be as far as the formation of non-marital unions is concerned.

[19] Compare the reflections and findings of Hayford and Morgan, 2008, based on similar data for the United States.

extensive than what we can report. If so, then the value *change* central to the SDT explanation may have been smaller than what meets the eye.

Acknowledgements

References

Andersson, G. & Philipov, D. (2002). Life-table representations of family dynamics in Sweden, Hungary, and fourteen 14 other FFS countries. A project of descriptions of demographic behavior. *Demographic Research* 7 (4), 67-144.

Aassve, A., Billari, F. C. & Spéder, Z. (2004) . Family Formation during the Hungarian Societal Transition: Trends in Postponement and the Impact of Policy Changes. Presented at the annual meeting of the Population Association of America (PAA).

Bradatan, C. & L.J. Kulcsar (2008). Choosing between marriage and cohabitation: Women's first union patterns in Hungary. To appear in *J. Comp. Family Stud.* 39 (4).

Carlson, E. & Klinger, A. (1987). Patterns in life: unmarried couples in Hungary. *European Journal of Population* 3, 85-99.

Frejka, T., Sobotka, T., Hoem, J. M. & Toulemon, L. (Eds., 2008). Childbearing trends and policies in Europe. *Demographic Research* 19, Special Collection 7, Parts 1 to 3.

Gerber, T. P. & Berman, D. (2005). Economic crisis or Second Demographic Transition? Trends and correlates of union formation in Russia, 1985-2001. Presented at the annual PAA meeting.

Hayford, S. R. & S. P. Morgan (2008). The quality of retrospective data on cohabitation. *Demography* 45 (1), 129-141.

Hoem, J. M. (1976). The statistical theory of demographic rates: A review of current developments (with discussion). *Scand. J. Statist.* 3 (4), 169-185.

Hoem, J. M. & Kostova, D. (2008). Early traces of the Second Demographic Transition in Bulgaria: A joint analysis of marital and non-marital union formation. *Population Studies,* 62 (3), 1-13. [forthcoming]

Hoem, J. M. & Kreyenfeld, M. (2006). Anticipatory analysis and its alternatives in life-course research. Part 2: Marriage and first birth. Reflexions. *Demographic Research* 15 (17), 485-498.

Kostova, D. (2007). The rise of cohabitation in Bulgaria: Who are the forerunners of the new family model? Paper presented to the annual PAA meeting.

Koytcheva, E. (2006). *Social-demographic differences of fertility and union formation in Bulgaria before and after the start of the societal transition.* Doctoral dissertation from the University of Rostock. http://www.demogr.mpg.de/publications/files/ 2318_1153389353_1_Full%20Text.pdf

Koytcheva, E. & Philipov, D. (2008). Bulgaria: Ethnic differentials in rapidly declining fertility. In Frejka et al. (Eds.) Childbearing trends and policies in Europe. *Demographic Research*, Volume 19, Special Collection 7 (pp. 361-402).

Lesthaeghe, R. & van de Kaa, D. J. (1986). Twe demografische transities? In Bevolking: groei en krimp, R. Lesthegehe & D. J. van de Kaa (Eds.), *Mens en Maatschappij*, 9-24. Deventer: Van Loghum-Slaterus.

Lesthaeghe, R. & Surkyn, J. (2002). New forms of household formation in Central and Eastern Europe: Are they related to newly emerging value orientations? *Economic Survey of Europe*, pp. 197-216.

Maleva, T.M. & Sinyavskaya, O.V. (eds, 2007). *Parents and Children, Men and Women in Family and Society* (in Russian). Moscow: Independent Institute for Social Policy.

Mureşan, C. (2007), How advanced is Romania in the Second Demographic Transition? *Romanian Journal of Population Studies* 1 (1-2), 46-60.

Mureşan, C. (2007), Family dynamics in pre- and post-transition Romania: a life-table description. MPIDR Working Paper 2007-018.

Mureşan, C., Hărăguş, P.T., Hărăguş, M., & Schroeder, C. (2008), Romania: Childbearing metamorphosis within a changing context. In Frejka et al. (Eds.) Childbearing trends and policies in Europe. *Demographic Research*, Volume 19, Special Collection 7 (pp. 855-906).

Philipov, D. & Jasilioniene, A. (2007). Union formation and fertility in Bulgaria and Russia: a life table description of recent trends. MPIDR Working Paper 2007-005.

Sobotka, T. (2008). The diverse faces of the Second Demographic Transition in Europe. Overview Chapter 6 in Frejka et al. (2008), Part 1, 171-224.

Spéder, Z. (2001). Turning points of the life course. Concept and design of the Hungarian social and demographic panel survey (in Hungarian). *Demográfia* 44 (2-3), 305-320. www.dpa.demografia.hu.

Spéder, Zsolt (2004). Cohabitation and marriage — facts, opinions, trends and transitions. Presentation to the annual PAA meeting.

Spéder, Z. (2005). The rise of cohabitation as first union and some neglected factors of recent demographic developments in Hungary. *Demográfia* 48, 77-103.

Spéder, Z. (2006). Rudiments of recent fertility decline in Hungary. *Demographic Research* 15 (8), 253-288.

Thornton, A. & Philipov, D. (2007), Developmental idealism and family and demographic change in Central and Eastern Europe. European Demographic Research Papers 2007/3, Vienna Institute of Demography. http://www.oeaw.ac.at/vid/download/edrp_3_07.pdf

Vikat, A., Spéder, Z., Beets, G., Billari, F.C., Bühler, C., Désesquelles, A., Fokkema, T., Hoem, J. M., MacDonald, A. L., Neyer, G. R., Pailhé, A., Pinnelli, A., & Solaz, A. (2007). Generations and Gender Survey (GGS). Towards a better understanding of relationships and processes in the life course. *Demographic Research* 17 (14), 389-440.

Zakharov, S. (2005). Recent trends in first marriage in Russia: Retarded second demographic transition. Paper presented to the 25th International Population Conference, Tours, France.

Zakharov, S. (2008). Russian Federation: From the first to second demographic transition. In Frejka et al. (Eds.) Childbearing trends and policies in Europe. *Demographic Research*, 19, Special Collection 7, 907-972.

Appendix: The specification and interpretation of our period coefficients

A1. Stable and uniform age profiles

For a discussion of the mathematical structure of our transition intensities we note that the quantities plotted in Figures 1 and 2 are maximum-likelihood estimates of parameters c_{kh}

and a_{ih} of multiplicative intensity functions that in its most general four-factor representation has the form

$$\mu_{ijkh} = a_{ih} b_{jh} c_{kh} \qquad (1)$$

for age group i, background group j, calendar period k, and decrement h.[20] As we have noted, in the present application the latter stands for the type of union formation, i.e., for entry into a marital or non-marital union, represented by $h=1$ and $h=2$, respectively, say. If we let A stand for the age factor, B for the combination of all the background factors that we mentioned as covariates in Section 2, C for the calendar period, and D for the decrement,[21] then the above specification of the union-formation intensity can be written symbolically as

$$\mu = AD + BD + CD,$$

where a double letter like AD or CD indicates that we include an interaction between the two factors involved (A and D, say, represented in (1) by a double subscript on the a parameter) and a plus sign indicates that an interaction has not been included. We then note that a condition that makes the items in Figures 1 and 2 work as a fair representation of our data is that for each type h of union entry, *the age effect $\{a_{ih}\}$ is the same in all periods k and for all levels j on the background factor*, as is indicated here by the lack of subscripts k and j on the a parameter in (1). Briefly, there is a requirement of (i) *stability* and (ii) *uniformity* in the age effects for our standardization to work without problem. (Similar requirements must be satisfied for factor B.) As we indicated in Section 4, we are of the opinion that the requirements on Factor A cause no essential problem for our empirical analysis.

A2. The interpretation of our period coefficients as relative risks

We now turn to the issue of the interpretation of the intensity parameters as relative risks. If we had been willing to analyze each decrement <u>separately</u>, then we would be dealing with two individual intensities μ_{ijk1} and μ_{ijk2}, and the very multiplicative specification of each of them, as in (1), would make sure that the parameters b_{jh} and c_{kh} could be interpreted as relative risks, in the usual manner for single-decrement intensities (Hoem 1976). (For the baseline factor A, we operate with absolute risks, per 1000 person-months for instance, and the

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[20] Relation (1) is a suitable starting point for a general discussion of the issues that we raise. For reasons that we will make clear as we go along, Figures 1 and 2 have actually been based on simpler specifications, namely $\mu_{ijkh} = a_i b_j c_{kh}$ for Figure 1 and $\mu_{ijkh} = a_{ih} b_j c_{kh}$ for Figure 2. Thus for the diagrams we use fewer factor interactions (double-subscript parameters) than the general theory allows for.

[21] Note how we use a mnemotechnical device in the naming of the various factors involved.

issue of relative risks does not concern the *a* parameters.) To secure parameter identification we would impose side conditions of the type $c_{k_0 1} = 1$ <u>and</u> $c_{k_0 2} = 1$ for a suitable period k_0, and we would use similar side conditions for the *b* parameters.

Since we now want to analyze μ_{ijk1} and μ_{ijk2} <u>jointly</u> for the purpose of seeing how one of them develops over time (i.e., as a function of *k*) <u>relative to the other</u>, things turn out to be a bit more complicated. First we drop one of the side conditions on each parameter set, and only require that $c_{k_0 1} = 1$ for the *c* parameters, say. For the period factor *C* we are faced with two types of relative risks, as follows:

(i) For any given type *h* of union formation, the intensity of union entry for a factor combination (*i, j, k*), relative to the combination (*i, j, k_0*), is

$$\mu_{ijkh} / \mu_{ijk_0 h} = \frac{a_{ih} b_{jh} c_{kh}}{a_{ih} b_{jh} c_{k_0 h}} = \frac{c_{kh}}{c_{k_0 h}}. \qquad (2.h)$$

Thus in particular for *h*=1,

$$\mu_{ijk1} / \mu_{ijk_0 1} = c_{k1}, \qquad (2.1)$$

(because of the side condition $c_{k_0 1} = 1$) and we see that c_{k1} is a relative risk in its own right. Furthermore,

$$\mu_{ijk2} / \mu_{ijk_0 2} = c_{k2} / c_{k_0 2}, \qquad (2.2)$$

which shows that up to a divisor $c_{k_0 2}$, the c_{k2} are relative risks also. All-in-all we have established that the items c_{kh} can essentially be interpreted as relative risks <u>along each curve</u> for every country in Figure 1. The curves faithfully represent the trend in each competing risk separately. Thus (2.1) and (2.2) show that the <u>form</u> of the trends curves remains independent of the specification of the age and background parameters (*a* and *b*).

(ii) It remains to compare corresponding points on the two curves for each country, i.e., to compare the curve point for the coordinate (*k*, 1) with the curve point for (*k*, 2) for each period *k*. Note that as has essentially been shown before by Hoem and Kostova (2008, end of Appendix),

$$\mu_{ijk2} / \mu_{ijk1} = \frac{a_{i2} b_{j2} c_{k2}}{a_{i1} b_{j1} c_{k1}} = s_{ij} \frac{c_{k2}}{c_{k1}} \text{ with } s_{ij} = \frac{a_{i2} b_{j2}}{a_{i1} b_{j1}}. \qquad (3)$$

With the model specification $\mu = A+B+CD$, $a_{i2} = a_{i1}$ for all i and $b_{j2} = b_{j1}$ for all j, and $s_{ij} \equiv 1$. We have used this model specification to produce all the curves in Figure 1, and see that we can therefore compare directly the trend <u>and level</u> of the risk of entry into cohabitation with the corresponding risk of marriage formation for each country. Our comments in the text proper have been based on this insight.

      Our computer program will provide estimates for the *a*, *b*, and *c* parameters even if we do not have $s_{ij} \equiv 1$, and the estimates may have some interest in their own right, but we can no longer automatically interpret the *c* estimates as relative risks and they may easily deviate considerably from those produced by the specification $\mu = A+B+CD$, except in special cases. For instance, the specification $\mu = AD+B+CD$ produces separate age profiles for the two decrements, for it means that we have let $\mu_{ijkh} = a_{ih}b_{j}c_{kh}$, with two separate age profiles $\{a_{i1}\}$ and $\{a_{i2}\}$, which we have plotted for each country in Figure 2. We see that for each country the age profiles of the two entry risks largely coincide except in details (i.e., $a_{i1} \approx a_{i2}$ for all *i*), i.e., the near-equality of the age profiles need not be such a terrible approximation, though for the details it manifestly is a tall order. So long as we are willing to accept approximations liberally, as one generally does when one practices standardization, the *c* parameters can therefore still be interpreted as relative risks, because according to (3), c_{k2}/c_{k1} largely represents the relative risk μ_{ijk2} / μ_{ijk1} as desired. (Remember that $b_{j1} = b_{j2}$ for all j with the given intensity specification, so b_{j1} and b_{j2} cancel in s_{ij}.) This also shows up in risk-trend diagrams that are much like those in Figure 1 except for minor details (not displayed here).

      We run into trouble if we try to extend these ideas to the specification $\mu = AD + BD + CD$. (When *B* represents several background factors, each of them is interacted separately with the decrement factor.) For Bulgaria and Romania the corresponding *c* plots are much like those in Figure 1, but for Russia and Hungary the inclusion of the interactions between the decrement and all the background factors produces *c* plots that really fail to represent properly the trends in union-formation risks. Allowing the background factors to influence the two competing risks differentially (as is the purpose of letting $\mu_{ijkh} = a_{ih}b_{jh}c_{kh}$ with genuinely *h*-dependent b_{jh}) results in a loss of control over the interpretation of the intensity parameters. We do not know how to interpret our parameters if they cannot be taken as relative risks. If a differential impact of the background factors is important, one should probably abstain from a joint analysis of the competing risks and instead study them separately until one understands better the implications of our formula (3).

Table 1: Person-years of exposure since 1960

	Total person-years of exposure			Year of data collection
	as childless, not pregnant	as childless, pregnant	at parity ≥1	
Russia	35161	373	3865	2004
Bulgaria	40057	360	1989	2004
Romania	33931	290	1416	2005
Hungary	49747	455	951	2001[1]

Note: [1]The first wave of the Hungarian GGS (originally called "Turning points of the life-course") was conducted in November 2001 through January 2002, but we do not use data collected in 2002 in our study.
Source*: Own calculations based on GGS data.

Table 2: Entry into marital and non-marital unions as competing events in Bulgaria, Romania, and Russia. Period survival-table estimates. Percent ever entered by age 35. Women

	Bulgaria[a]		Russia[a]		Romania[b]	
	Ever entered into		Ever entered into		Ever entered into	
Period	cohabitation	marriage	cohabitation	marriage	cohabitation	marriage
1985-1989	54	37	34	62	20[c]	76[c]
1990-1994	60	32	46	50		
1999-2003	63	14	62	33	35[d]	56[d]

Notes:
a: Source: Philipov and Jasilioniene (2007), Table A8.
b: Source: Muresan (2007), Tables 5.5 and 5.6.
c: 1980-89.
d: 1996-2005.

Table 3: Relative risk of first-union formation by parity-and-pregnancy status, for each type of union. Our selected countries, 1960-ca. 2004

	Childless, not pregnant	Childless, pregnant	Parity ≥1 (mother)
Entry into:			
RUSSIA 1960-2004			
cohabitation	0.50	2.34	0.48
marriage (direct)	1	8.40	0.42
BULGARIA 1960-2004			
cohabitation	1.31	11.69	0.64
marriage (direct)	1	17.07	0.47
ROMANIA 1960-2005			
cohabitation	0.24	1.70	0.16
marriage (direct)	1	8.47	0.73
HUNGARY 1960-2001			
cohabitation	0.32	1.36	0.50
marriage (direct)	1	17.69	1.25

Note: Standardized with respect to age, ethnicity, calendar period, character of region where respondent grew up (urban/rural), whether respondent lived with both parents at age 15, number of siblings, own educational attainment, and mother's and father's educational attainments. Some of these covariates have not been available for Hungary. For Romania the parents' educational attainments were not included because of data quality problems.
Source: Own calculations based on GGS data.

Figure 1: Trends in the rates of union formation, by type of union. Non-pregnant childless women in Russia, Romania, Bulgaria, and Hungary, since 1960. Rates relative to that of direct marriage in 1960-64, separately for each country
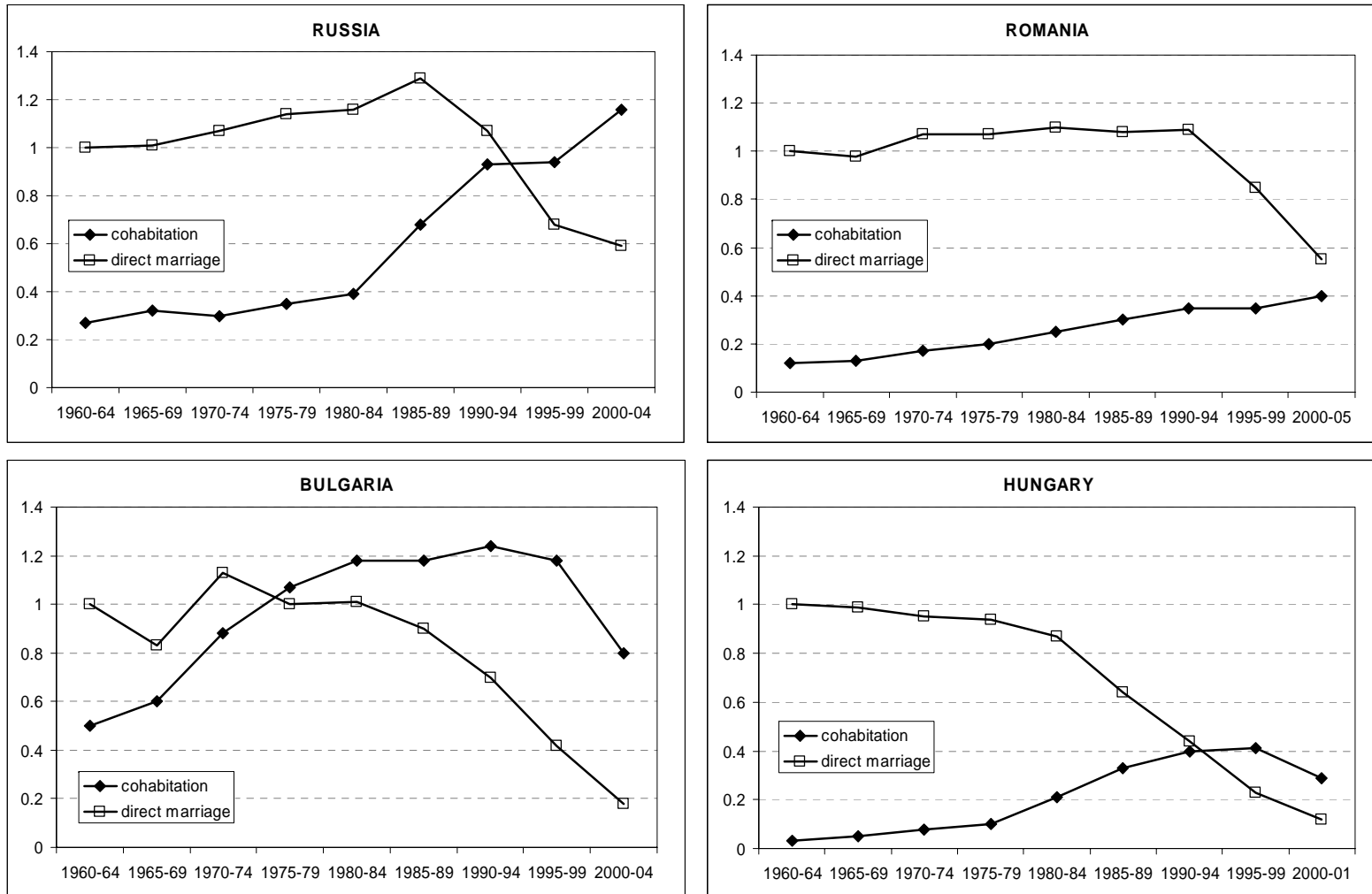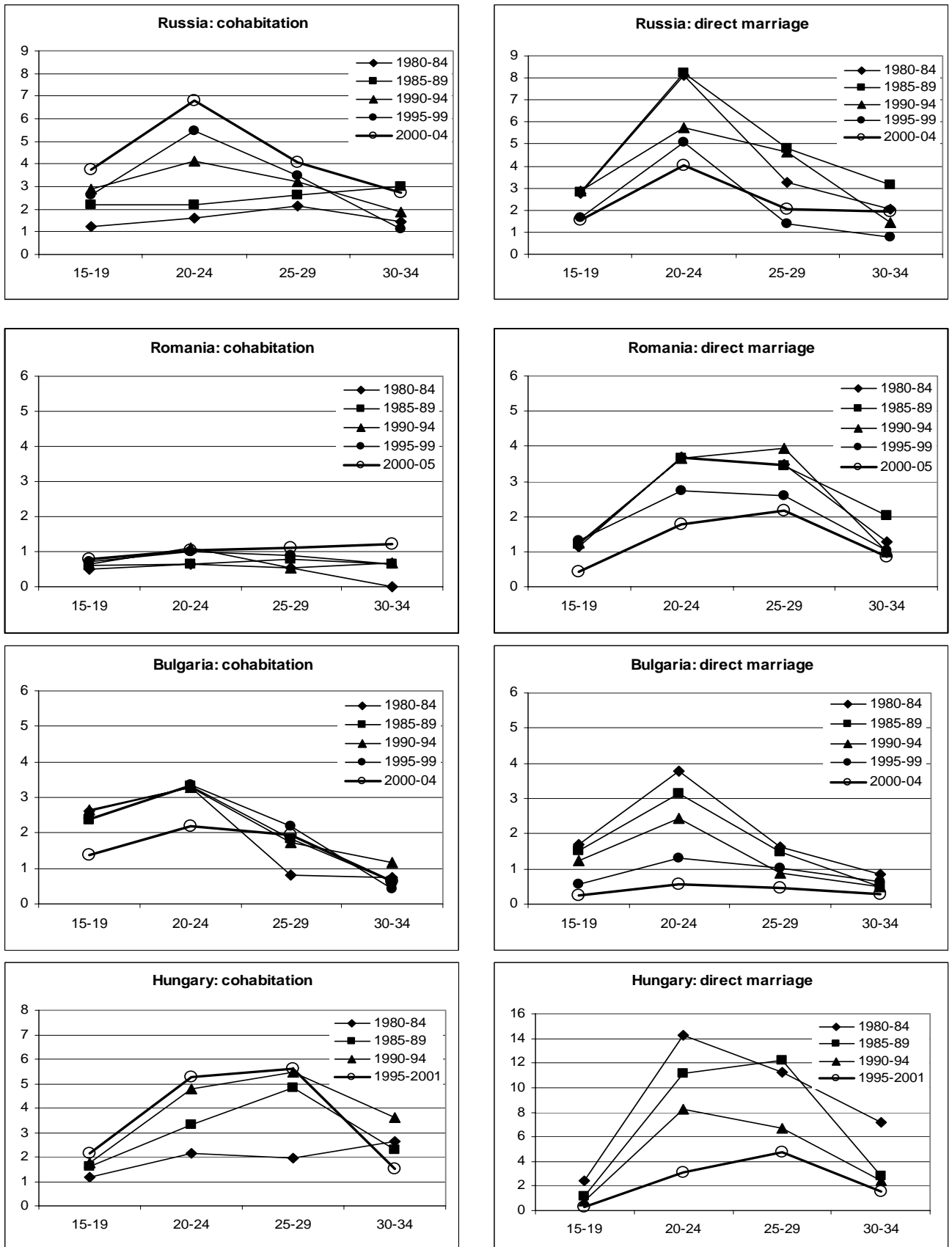

Source: Own calculations based on GGS data.

Figure 2: Age profiles of entry risks of union formation, by type of union. Non-pregnant childless women in Russia, Romania, Bulgaria, and Hungary, 1960-ca. 2004. Absolute risks per 1000 person-months
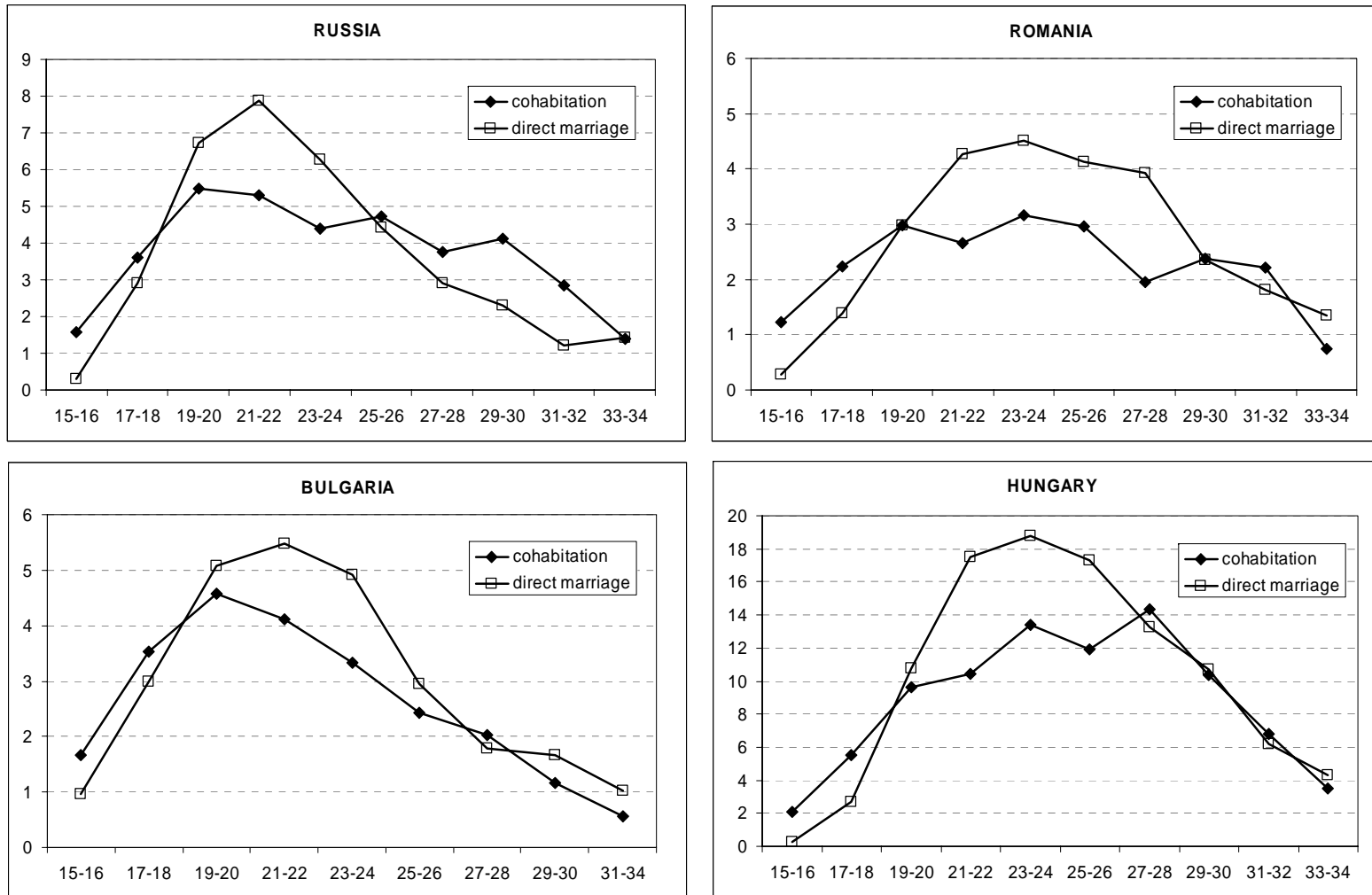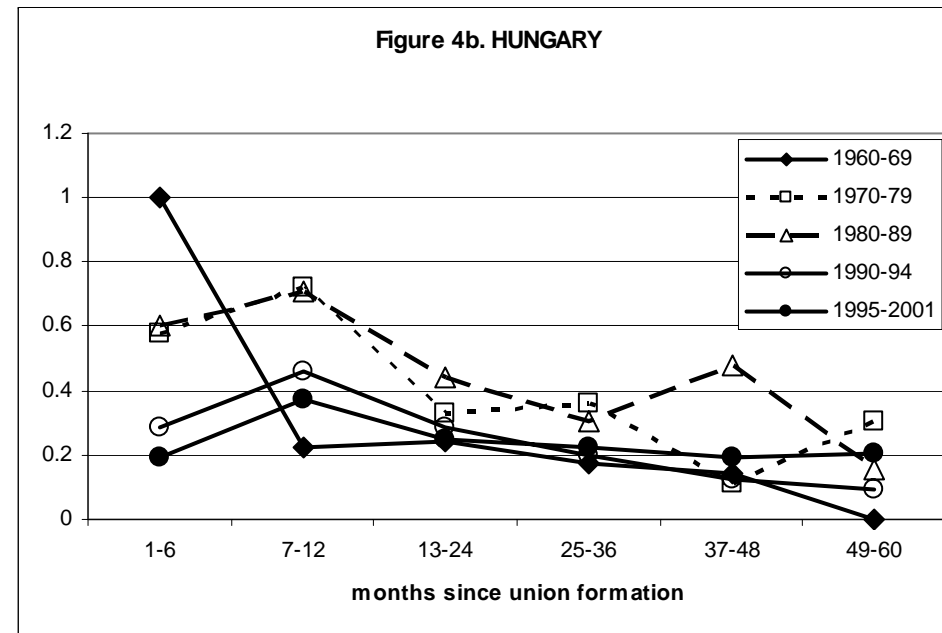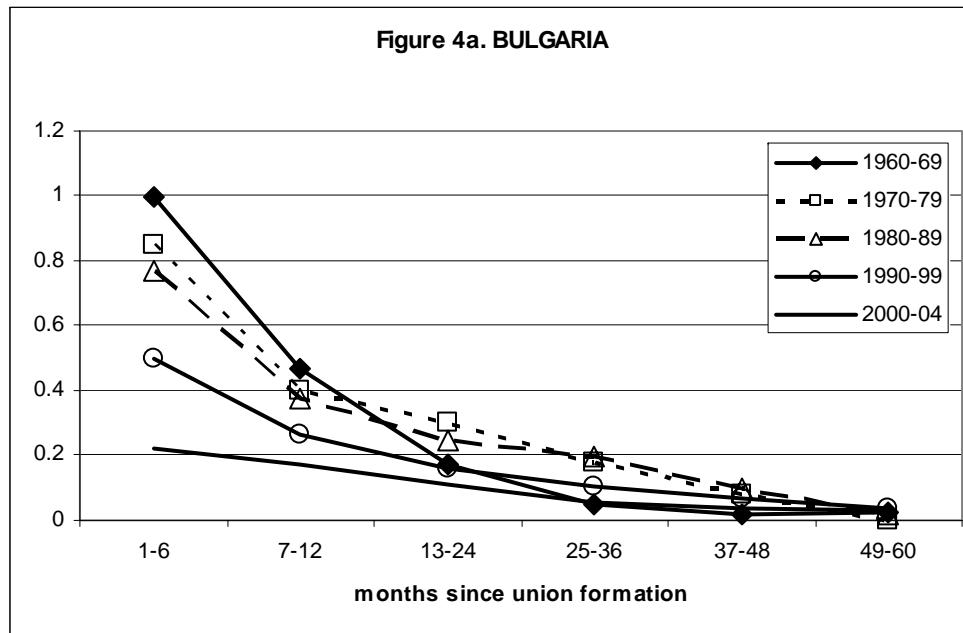

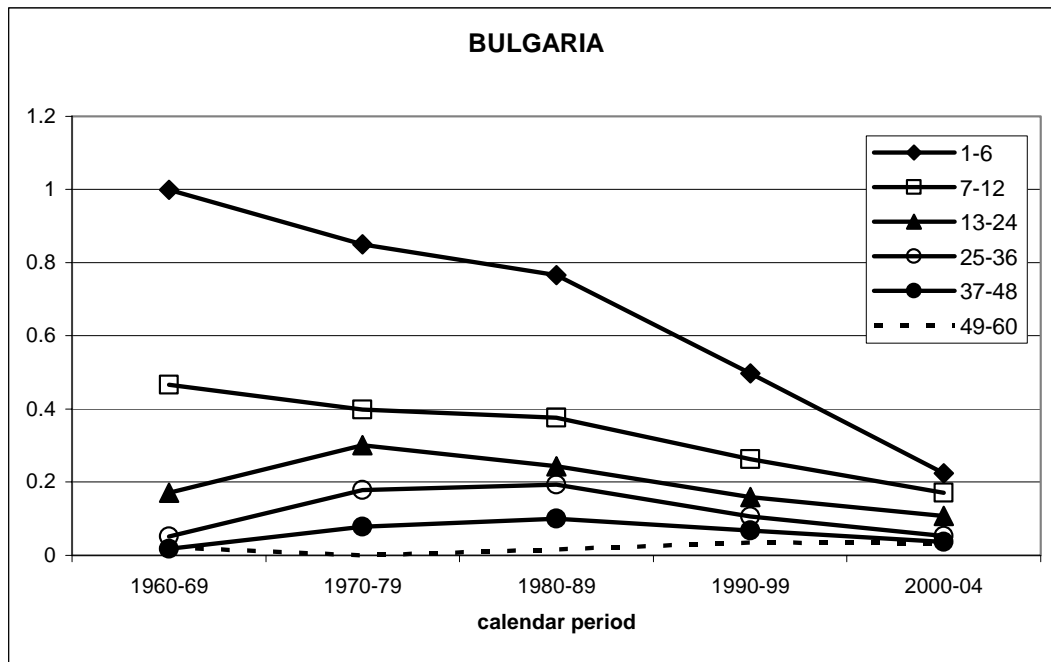**Source:** Own calculations based on GGS data

Figure 3: Age profiles of entry risks of union formation, by type of union and period. Non-pregnant childless women in Russia, Romania, Bulgaria, and Hungary, 1980-ca. 2004. Absolute risks per 1000 person-months

Figure 4: Relative rates of conversion of cohabitation into marriage, by time since entry into cohabitation for each calendar period, women in Bulgaria and Hungary, 1960-ca. 2004. Rates relative to a conversion during the first 6 months in the period 1960-1969.


Figure 4a. BULGARIA

Figure 4b. HUNGARY

Source: Own calculations based on GGS data.

Figure 5: Relative rates of conversion of cohabitation into marriage, by calendar period for each duration since entry into cohabitation, Bulgarian women, 1960-2004. Rates relative to a conversion during the first 6 months in the period 1960-1969.

Source: Own calculations based on GGS data.
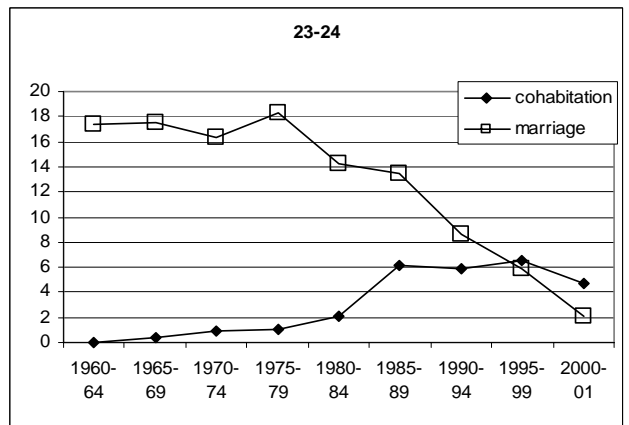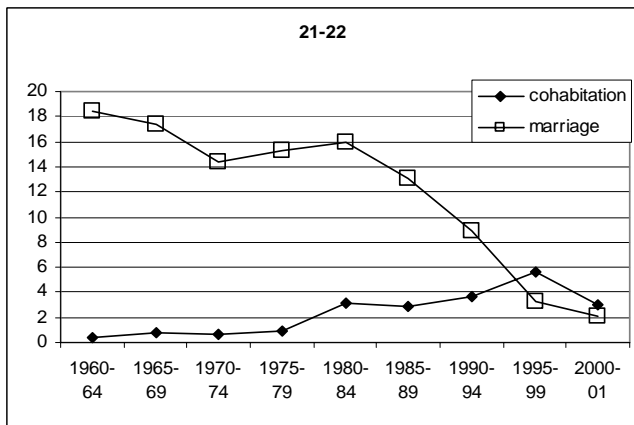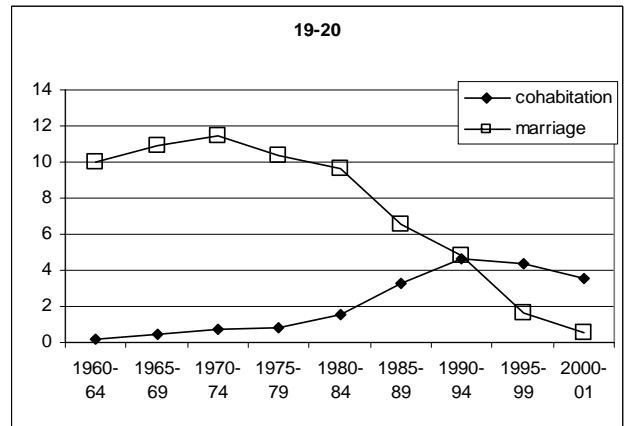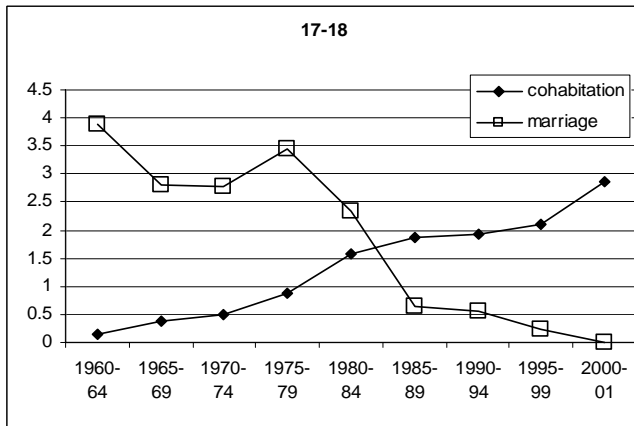
Hungary: childless, non-pregnant women, interaction from the model ACD+B.
Results presented for each age-group separately

Hungary: childless, non-pregnant women, interaction from the model ACD+B. Results presented for each calendar period separately