New Partners, New Children in Sweden

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ABSTRACT

The increasing incidence of childbearing with more than one partner is the feature of the Second Demographic Transition that has perhaps the greatest implications for family life. In this paper, we use data from the 2000 Swedish Level of Living Survey (LNU) to update research on stepfamily childbearing in Sweden. We model parity-specific birth risks in unions to test the hypothesis that birth risks are elevated if the child will be the first in a new union or if she will be the first full sibling to the couple's shared child. We also show how pace of childbearing immediately after stepfamily formation and the age of the youngest stepchild alter differentials in birth risks between couples with and without stepchildren. We identify shifts in the stepfamily differential from the 1950s to the 1990s in Sweden.

INTRODUCTION

The past three decades have been a time of rapid family change in Europe, often referred to as the Second Demographic Transition (Lesthaeghe and van de Kaa 1986). European fertility rates have dropped to below-replacement levels (2.1 children per woman), individuals are marrying later and less often, cohabitation has become a precursor to and a substitute for marriage, divorce and union dissolution rates have risen, and there is an increasing incidence of childbearing with more than one partner. This last phenomenon has perhaps the greatest implications for family life.

In this paper, we use data from the 2000 Swedish Level of Living Survey (LNU) to update and extend previous research on stepfamily childbearing in Sweden (Vikat, Thomson & Hoem 1999). The LNU covers not only more recent but also older cohorts and thus enables us to observe changes over time in the risk of stepfamily childbearing. We also examine in greater detail than previous studies of stepfamily fertility the timing of births in relation to union formation and the age of youngest stepchild.

PARTNERSHIP AND PARENTHOOD IN THE SECOND DEMOGRAPHIC TRANSITION

The central feature of the Second Demographic Transition is the changing relationship between partnership and parenthood. Childbearing outside of marriage, in particular within cohabitation, has become increasingly common, in some countries more common than childbearing in marriage. Parental separation is also more common, for the most part because cohabiting parents are more likely to separate than married parents, but also because divorce rates have risen in many countries. These changes increase the pool of single parents, thus increasing the possibilities for formation of new partnerships. And because much of the ‘churning’ of partnerships occurs during the childbearing years, the new partnerships present opportunities for further childbearing. Thus, we also observe an increase in stepfamilies and in childbearing with more than one partner.

If individual desires for particular numbers of children were independent of the partnerships in which children are born, we would expect childbearing in stepfamilies to depend only on the number of previous children. Those who wanted two children altogether but had only one would have the second in a stepfamily; those who already had two would not. Griffith, Koo & Suchindran (1985) argued, however, that new partnerships produce new motives for further childbearing. As in first unions, having a child together establishes the legitimacy of the parents’ union (termed *commitment* value), while having a second child together provides a full sibling for the first (*sibling* value). Because couples in stepfamilies may already have at least one child and may have additional children within the new union, such motives can produce ‘extra’ births when stepfamilies are formed.

By specifying parity progressions in terms of shared and separate children, Thomson and colleagues (Vikat, Thomson and Hoem 1999; Thomson et al. 2002; Vikat, Thomson & Prskawetz 2004; Henz and Thomson 2005), found very strong support for the commitment value of a common child and, in some instances, for the value of a full sibling for the first shared child in a stepfamily. Thomson (2004; Thomson and Li 2002) demonstrated that patterns were similar for birth intentions as for observed births, supporting the underlying motivation-based theory.

All of these studies faced a common problem – how to estimate a stepfamily fertility differential while taking into account the fact that stepfamilies start out with at least one child while couples without stepchildren start out with none. Children born in stepfamilies will always be of higher birth order to at least one of their parents than children born to couples who have only shared children. To take into account the effect of previous parity on childbearing, one must compare stepfamilies to couples with the same number of children, all of whom are shared (Vikat et al. 1999; Thomson et al. 2002). That is, first births to couples without stepchildren do not contribute to the comparisons. Instead, the risk of a first birth to stepfamily couples is compared to the risk of a higher-order birth to couples who have as many children together as the stepfamily couple has before their union.[3]

This type of comparison means that the birth risk for stepfamily couples and for couples without children may have different duration dependencies. Lillard and Panis (2003) noted that the risk of any particular event may depend on more than one duration parameter, or 'clock'. In order to fully understand duration dependence of a particular event, one must identify the 'baseline clock' and disaggregate any associated 'clocks' that could confound estimated effects of fixed or time-varying characteristics (e.g., stepfamily status and parity) on the event's risk.

Birth risks may be influenced by at least three types of clocks – the prospective parent's age, the duration of a partnership, and duration since a previous birth, i.e., the age of the youngest child. The parental age clock is likely to keep the same time for couples in stepfamilies and those with only shared children -- the older the parent, the less likely she/he is fecund and/or to want to begin a new long-term commitment to

[3] Some studies investigate parity progressions at the couple level, i.e., the effect of prior children on a couple's first or second birth. These estimates always show that birth risks are lower or at least no greater in stepfamilies. Lower risks could be due entirely to the fact that the birth is of higher order when it occurs to a stepfamily couple than to a couple without stepchildren; equal risks may reflect the added value of a shared birth overcoming the expected negative effect of prior parity.

childrearing. The other two clocks may keep very different time, however, for new stepfamilies and for couples with shared children.

For first births in stepfamilies, the period of observation is *union duration*. After the birth of a shared child, the period of observation is the *youngest child's age*. Although the age of the youngest step-child may also influence the birth risk in newly formed stepfamilies, the risk is observed only at older ages of youngest children, not from the month of the child's birth. Thus, the duration dependence of stepfamily childbearing on the youngest step-child's age is likely to differ from that for couples having a child together. In addition, because stepfamily couples may have 'delayed' further childbearing until the union was formed, we may observe a speeding up of childbearing shortly after the beginning of the union, particularly in relation to couples who have just had a birth and wish to space the next child.

Previous studies of the potential 'extra' births produced by stepfamilies specified duration dependence in different ways. Vikat and colleagues (1999) used a discrete-time model in which both union duration and age of youngest child (shared or step) were included. This specification does not enable us to identify differences in either of these two duration dependencies for those with and without shared children. Thomson and colleagues (2000; Henz & Thomson 2005) specified the baseline clock as age of youngest child after a shared birth but as union duration in a new stepfamily, i.e. assuming that the pace of childbearing was the same after stepfamily formation as after the birth of a shared child. They did not consider effects of the age of youngest child in stepfamilies or effects of union duration for a second- or higher-order shared birth. Henz (2002) tested the interaction between duration and whether a spell began with a birth or the formation of a stepfamily union. She found in West Germany that couples with one child were more likely to have a second child if the first was not shared, but only during the first three years of the union. The different patterns suggest that for some couples partnership formation may be endogenous to the fertility process. Using a large data set from the United States, Li (2006) demonstrated that the pace of childbearing in new marital stepfamilies was faster than the pace of childbearing after a marital birth. Once stepfamily couples had a common child, however, they exhibited the same duration dependence in having subsequent children as couples without stepchildren.[4] Although most of the previous research on stepfamily childbearing has controlled for period effects on fertility, none have examined possible changes in the level or pace of childbearing in stepfamilies across time. Delays in family formation produce later separations and re-partnering and may therefore have made it harder for stepfamily couples to have one or two children together, even if they would like to do so. In addition, in 1980 Sweden formalized a parental leave policy that encouraged closer spacing of second- and higher-order births and thereby increased overall the risk of second births (Hoem 1993; Andersson 2004). Because most stepchildren are more than two years old at the time of stepfamily formation, very few stepfamily couples could benefit from the spacing incentive. This means that in the later decades any 'extra' incentive to have a second child who would be the first in a union would be counterbalanced by strong incentives to have a second shared child at a rapid rate after the first.

[4] The 'clock' problem also applies when comparing the risk of a first shared birth to couples with and without stepchildren. For both groups, the clock is union duration, but stepfamily couples already have an older child and may have an incentive to 'speed up' their childbearing. Buber and Prskawetz (2002) tested the interaction between union duration and having prior children and found an early peak in birth risks for stepfamily couples in which only the man had previous children.

DATA AND METHODS

Data

We use data from the 2000 Swedish Level of Living Survey (LNU). The LNU is a 1/1000 random sample of the Swedish population between 15 and 75 years of age, first conducted in 1968. The survey was replicated in 1974, 1981, 1991 and 2000. In the last two waves the lower age limit was increased to 18 and complete birth and union histories were collected.

We restrict our sample to Swedish-born respondents who had only singleton births, did not report an adopted or a deceased child and did not report births less than six months apart. A handful of respondents who did not report the year of each birth are also excluded.[5] The final analytic sample consists of 3,023 respondents and 3,247 unions at risk for a second birth and 2,275 respondents and 2,478 unions at risk for a third birth. The LNU did not provide sufficient observations to obtain robust estimates for stepfamily effects on the risk of a fourth or higher-order birth.

We combined union and birth histories in order to assign each birth to the respondent's union(s) or periods of living alone. In some cases union start and end months are reported by quarter rather than month or are missing altogether. We impute union start and end months where missing or reported as a quarter; imputations were made for 267 unions and 224 respondents. Children are 'assigned' to a union if they were born within nine months of the union's end. Children not assigned to previous unions but born within 12 months prior to a new union are assigned to that union. Only

[5] In two instances the birth month was missing; we assigned the birth to June.

children born more than nine months after a union and more than 12 months before a subsequent union are considered to be born ‘out’ of any union.[6]

Methods

We estimate second- and third-birth risks for respondents who reported, respectively, a first or second birth. For the second birth risk we observe two different types of spells: (1) intervals after the first birth in a union without stepchildren; and (2) intervals after the respondent enters a new partnership with one child. For the third birth risk we observe three types of spells: (1) intervals after the second birth in a union without stepchildren; (2) intervals after the first birth in a union with one stepchild; and (3) intervals after the respondent enters a new partnership with two children. Spell types (1) for first births and (1) and (2) for second births begin with a birth. Spell types (2) for first births and (3) for second births begin with a union. All spells are censored at interview and at the end of the childbearing years: age 45 for women and age 50 for men.[7] In total, the 3,247 respondents with a first birth contribute 180,461 person months of observation and 2,281 second births, while the 2,478 respondents with a second birth contribute 255,807 person months of observation and 820 third births.

We model the risk of a birth in continuous time using Cox proportional hazards models (Cox 1972; Blossfeld, Golsch and Rohwer 2006). The duration variable is not parameterized, and thus there is no assumption about the underlying relationship between the hazard function and time. Following Buber and Prskawetz (2000) and Li (2006), we

[6] This assignment has also been used by most of the stepfamily analyses referenced herein.
[7] As a result of censoring at age 45 for women and 50 for men, we fail to observe 11 births.

test for differences in the pace of childbearing after union formation versus shared birth by specifying interactions between stepfamily status and the duration parameter.[8] .

Our central interest is in the relationship between the union status of prior births and the risk of a subsequent birth. For the second birth risk, we specify whether the union is or is not the same as that producing the first child. For the third birth risk, we specify whether the union is the same as that producing the second child, the second and first children, or neither.

Because we do not have information on the number of children born to prior partners, we cannot estimate a 'full' parity model as in several previous analyses (e.g., Thomson et al. 2002; Thomson 2004; Vikat, Thomson & Prskawetz 2004; Henz 2002; Henz & Thomson 2005). This is the same limitation faced by Vikat, Thomson & Hoem (1999) in their previous analysis for Sweden.

Control variables in our models include the respondent's sex and age, the period of observation (1950s and 1960s, 1970s, 1980s, 1990s) and, for the first birth in stepfamilies, the age of the youngest stepchild.[9] Respondent's age and period of observation are specified as time-varying linear splines.[10]

RESULTS

Table 1 presents descriptive statistics for spells at risk for a second birth. Spells where the first birth was in the current union are more likely to end in a birth than

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[8] We do not have sufficient observations to apply the more fine-tuned models presented in Li (2006).
[9] Note that this specification assumes that the age of the child *at stepfamily formation* has a constant proportional effect on the birth risk, even as the child ages further over the period observed.
[10] The spline specification and the particular nodes chosen were based on the estimated AIC and BIC selection criterion from several specifications (continuous, dummy variable, splines with alternative nodes), not shown here.

stepfamily spells, with 73% and 51% of spells ending in a birth, respectively. Duration also varies by spell type. Approximately one-quarter of spells where the first birth took place in the same union last 24 months or less, 25 to 36 months, 37 to 60 months or 61 months or more. On the other hand, spells beginning with stepfamily formation are more likely to be shorter, with 38% of spells lasting 24 months or less. Additionally a higher proportion of these spells (30%) last more than 60 months. The sex ratio of respondent is balanced across the two types of spells. In stepfamily spells, respondents are older on average than in intact unions: in 70% of stepfamily spells, respondents are over the age of thirty while only about 50% of respondents are over thirty across second birth intervals in intact unions. Stepfamily spells are differentiated by the age of youngest child at the start of the spell: one-quarter of stepchildren are under age 3, one-fifth are between 3 and 5 years old, 30% are between 5 and 10 and just under one-fifth are over 10 years old. Finally, intact union spells are more likely to be observed before 1980, while stepfamily spells are observed more frequently in later periods.

Table 2 presents descriptive statistics for spells at risk for a third birth. For third birth risk, we characterize union type in three ways: spells in which both the first and second children were born in the current union (84% of spells), spells where the second birth occurred in the current union but the first birth occurred before the current union (7% of spells) and spells where all children were born previous to the current union (9% of spells). Both the second and third types of spell are considered stepfamily spells, but only in the third type is union duration the underlying clock. Spells where the previous birth was to the current union but the first birth before the union are most likely to end in a birth (44% of spells), while approximately 30% of spells with both of the children or none of the children born in the current union end in births. Spells following the second birth in a union are the longest, with 80% lasting more than 36 months. Among spells following the first shared birth (with an older stepchild) , 63% last more than 36 months. Among spells following stepfamily formation (both children born before the union), duration is more equally distributed, with half of the spells ending before or after 36 months. A marginally higher proportion of stepfamily spells, with either one or no children born in the current union, are contributed by women. Respondents are youngest, on average, in spells with one birth in the current union, with 70% over the age of thirty, followed by spells with two births in the current union (80% over thirty) and spells with all children born previous to the union (92% over thirty). When spells begin with a union, the younger stepchild tends to be older on average than in the case of second birth risks, with 15% of children under the age of 3, 16% between 3 and 5 years old, 32% between 5 and 10 and 37% over the age of 10. Finally, with regard to period of observation, about 45% of spells following a birth – whether the first (with a stepchild) or second in a union – were observed before 1980. On the other hand, the large majority of spells beginning with union formation – after the birth of two children to the respondent – are observed after 1980 (77%).

Estimates for the risk of a second lifetime birth are presented in Table 2. Model 1 estimates the differential in the birth risk between those entering a stepfamily with one child and those having had their first child in the same union. The estimated birth risk is *lower* in stepfamilies than for couples with one shared child. When we control for characteristics associated with stepfamilies and for the age of the youngest stepchild (Model 2), the estimated birth risk for stepfamilies is no different from that for a second shared birth to couples without stepchildren. As expected, birth risks are lower for stepfamilies with older children, as well as for both types of families with older parents. Keep in mind that the estimated difference in birth risks between stepfamily couples with one stepchild and no shared children and couples with one shared child depends on the assumption that the pace of childbearing is the same for couples having just formed a stepfamily union or having just had a first shared birth.

In Model 3, we relax the assumption of common duration dependence by distinguishing the step-family difference at different durations. During the first 24 months of a stepfamily union, the risk of a birth is twice as high as during the first 24 months of a shared child's life. Thereafter, the birth risk for stepfamilies is no more than half the risk for couples whose child is age two and above. When we add control variables, the stepfamily birth risk is more than three times as high during the first 24 months of observation, but after 24 months, the differential reverses with lower second birth risks in stepfamilies.

Note that in both Models 2 and 4 the risk of second births was greater in the 1980s and 1990s than in previous decades. The increase is consistent with the introduction of a 'speed premium' in parental leave formalized in 1980 and the resulting decrease in second birth intervals and increase in second births overall (Hoem 1993; Andersson 2004). As noted above, the speed premium would not likely influence birth risks in stepfamilies because by the time a stepfamily is formed, the first child would have been too old for the mother or father to claim the premium. Thus, we might expect differences in the birth risk to be reduced in later periods when strong incentives existed for short birth intervals among couples with shared children. In Model 5 we present results in which period (before 1980, 1980 and later) interacts with the duration-specific stepfamily parameters.[11] In the 1950s, 1960s and 1970s, the birth risk for stepfamilies during the first 24 months of their union was more than 4 times that of the risk for couples who had a child under 24 months; the differential was reduced to about 1.6 (4 x .4) for the 1980s and 1990s. Between 24 and 36 months, stepfamilies were equally likely to have a child as were couples with a child 24-36 months old in the earlier period, but half as likely during the later period. Similar patterns are found for the relative risks at later durations.

Table 4 presents models for the risk of a third life-time birth. In Model 1, we find higher birth risks for stepfamily couples, whether they have no child or one child together. When we control for parental age, sex, age of youngest child and period (Model 2), the higher risk of a third birth remains for those who have one shared child and one step-child. The differential for couples with two stepchildren is now reflected in the set of coefficients that include age of youngest stepchild at the time of union formation – defined only for the spells beginning with a union rather than a birth. Couples with two stepchildren, the youngest of which is between one and two years old[12] are three times as likely to have a birth as couples with two shared children and twice as likely as couples with one shared child and one stepchild. The risk is even greater when the younger stepchild is between 3 and 6 years old. All of these estimates, however, depend on the assumption that the baseline hazard is the same for new stepfamilies as for couples after a shared birth. Model 3 relaxes that assumption by specifying an

[11] We tested the full interaction using five 10-year periods, but found that the shift in effects occurred between the 1970s and 1980s. The more detailed interaction is less robust because of a limited number of observations.

[12] Because we counted children born up to 12 months before the union as shared children, no stepchild can be less than 13 months old at the time of stepfamily formation.

interaction between stepfamily status (new stepfamily versus other couples) and the period of observation. We do not specify an interaction with duration for stepfamily couples with one shared child because the baseline hazard after that child's birth is virtually identical to the hazard for couples who have had a second child together. (See also Henz 2002; Li 2006.) As for second births, the risk of a first birth is particularly high for couples with no shared children during the first 24 months of their union. It drops to the level of couples with one shared child during the succeeding year and is essentially the same as that for couples with two shared children after three years of observation.

When we add additional covariates, we find that the relative risks of childbearing for new stepfamilies increases, especially during the first 24 months, when the younger stepchild is between one and two years old or between three and six years old. We also tested the interaction between the two types of stepfamilies and period but found no significant differences. That is, the 'extra' third births produced by stepfamilies early in their union, especially when the stepchildren are quite young, were as common in the later as in the earlier period.

CONCLUSIONS

This paper finds results consistent with previous research on stepfamily fertility in Sweden. We find evidence of an increased risk of a birth when it is the first shared child in a couple. However, this finding of an increased birth risk is conditioned upon differential effects of pace of childbearing, various underlying "clocks," such as age of parent and age of youngest child, and period.

Consistent with previous work by Li (2006) and Buber and Prskawetz (2000), we find that the pace of childbearing may vary across spell types. When considering the risk of a second birth under an assumption of common duration dependence we find little evidence of an increased risk of a second birth for stepfamilies. When we allow the duration dependence after a shared first birth to differ from that after a stepfamily is formed, however, we find an increased risk of a second birth for stepfamilies only in the first 24 months of observation. After two years, we find no difference in the risk of a birth for one-child stepfamilies relative to unions with a shared first birth. We find a similar increased pace of fertility for stepfamilies with no shared children at risk for a third lifetime birth. Once we disaggregate these spells by duration and include covariates, we find that the risk of a third birth is particularly high early on in the union, falling in subsequent years and, after five years of union duration, the risk is no different from that for couples with two shared children. These findings provide evidence of an increased pace of fertility in stepfamilies without shared children; that is, the 'commitment value' of children appears to operate most strongly within the first 24 months after the formation of stepfamilies. Henz (2002) suggested that childbearing shortly after the formation of a stepfamily union indicates that union formation is endogenous to the couple's decision to have children together. This does not negate, but is in fact consistent with the 'commitment value' of a first shared child. It does, however, suggest that the 'clock' putting couples at risk of a shared birth precedes their coresidence.

The results presented here also highlight the importance of disaggregating the multiple clocks underlying the risk of a birth in a step union. As expected, we find that the risk of a birth is inversely related to the age of the parents. Age of youngest child operates independently of union duration for families with no shared births with regard to the risk of both a second and third birth. We find that having an older child reduces the risk of a second birth. For third births, we also find that having an older child (over the age of 10) reduces the risk of a birth. However, the increased birth risk for stepfamilies with youngest children aged 6 to 10 is more puzzling.

Finally, the LNU allows us to consider the risk of a stepfamily birth across a wider range of cohorts. We find the risk of any second birth was greater in the 1980s and 1990s than in previous decades. Additionally, we identify a change in the relationship between union type and birth risks for the second births over time. In the period after 1980, we observe an increased risk for a stepfamily birth in the first 24 months of the union relative to intact families. However, this risk is reduced as compared to the risk of such a birth in the pre-1980s period. For subsequent relationship durations, the risk of a birth in a stepfamily is half that of an intact family after 1980, while a birth was equally likely in intact and stepfamilies in the earlier period. Altogether, these findings provide some evidence for a changing relationship between union status and birth risk and evidence for a relationship between policy and individual family formation decisions. It is likely that the observed increased pace of fertility among intact couples is a result of policy shifts in the 1980s introducing a “speed premium” in Sweden. This increase in earlier birth risk for couples with a shared birth, in turn, reduces the relative risk of a birth in stepfamilies.

We expect that our estimates of the commitment effect increasing the risk of births in stepfamilies are conservative. Because we rely on retrospective fertility and relationship histories of the respondent, we are unable to take into account on partner's non-residential children from prior unions. Consequently, there may be some stepfamilies wrongly categorized as families where all children are shared, thus somewhat diminishing the difference in birth risk between intact and stepfamilies. As a result, we may underestimate any increased risk of a birth identified among stepfamilies.

REFERENCES

Andersson, Gunnar. 2004. “Childbearing developments in Denmark, Norway, and Sweden from the 1970s to the 1990s: A comparison”. *Demographic Research* Special Collection 3(7):155-176.

Blossfeld, Hans-Peter, Katrin Golsch and Götz Rohwer (2006). Event History Analysis with Stata. Mahwah, NJ: Erlbaum.

Buber, I. and A. Prskawetz. 2000. “Fertility in Second Unions in Austria: Findings from the Austrian FFS.” *Demographic Research* 3(2).

Cox, D. R. 1972. “ Regression Models and Life Tables” Journal of the Royal Statistical *Society Series* B34: 187-220.

Griffith, Janet D., Helen P. Koo, and C. M. Suchindran. 1985. “Childbearing and Family in Remarriage.” *Demography* 22:73-88.

Henz, U. 2002. “Childbirth in East and West German Stepfamilies: Estimated Probabilities from Hazard Rate Models.” *Demographic Research* 7(6).

Henz, Ursula, and Elizabeth Thomson. 2005. “Union Stability and Stepfamily Fertility in Austria, Finland, France & West Germany.” *European Journal of Population* 21: 2-29.

Hoem, J.M., 1993. “Public policy as the fuel of fertility: effects of a policy reform on the pace of childbearing in Sweden in the 1980s.” Acta Sociologica 36: 19-31.

Lesthaeghe, R. and van de Kaa, D. J., 1986. 'Twee Demografische Transities?', in D. J. van de Kaa and R. Lesthaeghe (eds), Bevolking: Groei en Krimp. Van Loghum Slaterus, Deventer, 9–25.

Li, Jui-Chung Allen. 2006. “The Institutionalization and Pace of Fertility in American Stepfamilies.” *Demographic Research* 14(12): 237-66.

Lillard, L. and C.W.A. Panis. 2003., aML Multilevel Multiprocess Statistical Software, Version 2.0. Los Angeles: EconWare.

Oláh, L. Sz. 2001. Gendering family dynamics. The case of Sweden and Hungary. Stockholm: Demography Unit, Stockholm University. Dissertation Series No. 3.

Thomson, Elizabeth. 2004. “Stepfamilies and Childbearing Intentions in Europe.” *Demographic Research* S3-4 [www.demographic-research.org]

Thomson, Elizabeth, Jan M. Hoem, Andres Vikat, Isabella Buber, Alexia Fuernkranz-Prskawetz, Laurent Toulemon, Ursula Henz, Amy L. Godecker and Vladimira Kantorova. “Childbearing in Stepfamilies: Whose Parity Counts?” 2002. In E. Klijzing and M. Corijn (eds.), Fertility and Partnership in Europe: Findings and Lessons from Comparative Research. Volume II. Geneva/New York: United Nations.

Thomson, Elizabeth, Maria Winkler-Dworak, Martin Spielauer and Alexia Prskawetz. 2007. “Union Instability as an Engine of Fertility? A Micro-simulation model for France.” Presented at the Population Association of America Annual Conference, New York, NY.

Vikat, Andres, Elizabeth Thomson and Jan M. Hoem. 1999. “Stepfamily Fertility in Contemporary Sweden: The impact of childbearing before the current union.” *Population Studies* 53(2): 211-25.

Vikat, Andres, Elizabeth Thomson and Alexia Prskawetz. 2004. Childrearing Responsibility and Stepfamily Fertility in Finland and Austria. *European Journal of Population* 20:1-21.

Table 1: Characteristics of spells for the second birth risk

	Number of Births in Current Union		Total
	One	None	
Proportion of Spells (%)	89.44	10.56	100.00
Proportion of Spells that End in a Birth (%)	72.52	51.02	70.25
Duration of Spell (%)			
24 months or less	22.76	38.48	24.42
25 - 36 months	27.07	15.45	25.84
37 - 60 months	27.79	15.74	26.52
61 or more months	22.38	30.32	23.22
Female (%)	52.44	52.19	52.42
Age (%)			
<25	18.30	10.58	17.50
25 to <30	30.49	18.85	29.28
30 to <35	22.84	19.56	22.50
35+	28.38	51.00	30.72
Age of Youngest Child at Spell Start (%)			
36 months or less	-	26.87	-
37 - 60 months	-	18.50	-
61 - 120 months	-	31.76	-
121 months or more	-	22.87	-
Period (%)			
1950s, 1960s	30.66	15.16	29.05
1970s	27.61	19.84	26.81
1980s	22.94	34.70	24.15
1990s	18.80	30.30	19.99
Total Spells	2,904	343	3,247
Total Person Months	161,795	18,666	180,461

Source: LNU 2000. Authors' Calculations

Note: Spells for those with no children born in union start at the beginning of the union rather than at the birth of the previous shared child.

Table 2: Characteristics of spells for the third birth risk

	Number of Births in Current Union			Total
	Two	One	None	
Proportion of Spells (%)	84.02	6.90	9.08	100.00
Proportion of Spells that End in a Birth (%)	32.33	44.44	31.55	33.1
Duration of Spell (%)				
24 months or less	9.80	16.37	34.67	12.51
25 - 36 months	10.85	21.05	18.22	12.23
37 - 60 months	17.20	18.13	18.22	17.35
61 or more months	62.15	44.44	28.89	57.91
Female (%)	52.02	57.89	56.44	52.82
Age (%)				
<25	3.28	6.94	1.60	3.41
25 to <30	16.55	22.08	5.99	16.41
30 to <35	26.49	26.61	19.63	26.19
35+	53.69	44.36	72.78	53.99
Age of Youngest Child at Spell Start (%)				
36 months or less	-	-	15.01	-
37 - 60 months	-	-	16.07	-
61 - 120 months	-	-	32.13	-
121 months or more	-	-	36.78	-
Period (%)				
1950s, 1960s	17.05	16.71	5.70	16.53
1970s	27.15	30.18	16.93	26.87
1980s	30.84	30.98	40.19	31.26
1990s	24.96	22.12	37.18	25.34
Total Spells	2,082	171	225	2,478
Total Person Months	229,986	14,604	11,217	255,807

Source: LNU 2000. Authors' Calculations

Note: Spells for those with no children born in union start at the beginning of the union rather than at the birth of the previous shared child.

Table 3: Relative Risk of a Second Birth

	Model 1	Model 2	Model 3	Model 4	Model 5
Relationship Characteristics (time varying, 0/1)					
Same union 1st Birth	1.00	1.00	1.00	1.00	1.00
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
Different/Non Union than 1st Birth	0.73 ***	1.09	-	-	-
	(0.06)	*(0.13)*			
Duration 0 - 24 months	-	-	2.26 ***	3.22 ***	4.13 ***
			(0.26)	*(0.47)*	*(0.68)*
Duration 25 - 36 months	-	-	0.51 ***	0.75	0.97
			(0.09)	*(0.15)*	*(0.23)*
Duration 37 - 60 months	-	-	0.28 ***	0.40 ***	0.68
			(0.06)	*(0.09)*	*(0.19)*
Duration 61 or more months	-	-	0.48 **	0.69	0.98
			(0.11)	*(0.17)*	*(0.34)*
Different/Non Union than 1st Birth * Period	-	-	-	-	-
Duration 0 - 24 months, 1980's-1990's	-	-	-	-	0.45 ***
					(0.10)
Duration 25 - 36 months, 1980's-1990's	-	-	-	-	0.45 *
					(0.16)
Duration 37 - 60 months, 1980's-1990's	-	-	-	-	0.29 **
					(0.12)
Duration 61 or more months, 1980's-1990's	-	-	-	-	0.45 +
					(0.20)
Female (0/1)	-	0.91 *	-	0.91 *	0.91 *
		(0.04)		*(0.04)*	*(0.04)*

Table 3: *Continued*

	Model 1	Model 2	Model 3	Model 4	Model 5
Age (splines)					
<25	-	1.01	-	1.01	1.01
		(0.02)		*(0.02)*	*(0.02)*
25 to <30	-	0.98 $^{+}$	-	0.98	0.97 $^{+}$
		(0.01)		*(0.01)*	*(0.01)*
30 to <35	-	0.96 *	-	0.96 *	0.96 *
		(0.02)		*(0.02)*	*(0.02)*
35+	-	0.81 ***	-	0.81 ***	0.81 ***
		(0.02)		*(0.02)*	*(0.02)*
Age of Youngest Child (0/1)					
36 months or less	-	1.00	-	1.00	1.00
		(0.00)		*(0.00)*	*(0.00)*
37 - 60 months	-	0.90	-	0.96	1.16
		(0.17)		*(0.18)*	*(0.22)*
61 - 120 months	-	0.70 $^{+}$	-	0.73	0.96
		(0.13)		*(0.14)*	*(0.19)*
121 months or more	-	0.41 *	-	0.42 *	0.53 $^{+}$
		(0.14)		*(0.15)*	*(0.18)*
Period, time varying (0/1)					
1950s, 1960s	-	1.00	-	1.00	1.00
		(0.00)		*(0.00)*	*(0.00)*
1970s	-	1.04	-	1.03	1.02
		(0.06)		*(0.06)*	*(0.06)*
1980s	-	1.40 ***	-	1.43 ***	1.52 ***
		(0.08)		*(0.09)*	*(0.09)*
1990s	-	1.54 ***	-	1.55 ***	1.65 ***
		(0.10)		*(0.10)*	*(0.10)*
Log Likelihood	-16735.31	-16572.90	-16676.73	-16515.94	-16501.81
AIC	33472.61	33169.79	33361.46	33061.88	33041.61
BIC	33482.71	33291.03	33401.87	33213.43	33233.57
df	1	12	4	15	19
N (Century Months)	180,461	180,461	180,461	180,461	180,461
N (Individuals)	3,247	3,247	3,247	3,247	3,247

Source: LNU 2000. Author's Calculations

Note: Hazard ratios estimated using Cox proportional hazards regression; regression coeeficient standard errors in parentheses

*** $p<0.001$ ** $p<0.01$ * $p<0.05$ $^{+}$ $p<0.1$

Table 4: Relative Risk of a Third Birth

	Model 1	Model 2	Model 3	Model 4
Relationship Characteristics (time varying, 0/1)				
Same union 1st, 2nd Birth	1.00	1.00	1.00	1.00
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*
Same union 2nd birth; different union 1st birth	1.76 ***	1.74 ***	1.77 ***	1.74 ***
	(0.21)	*(0.21)*	*(0.21)*	*(0.21)*
Different/Non Union than 1st, 2nd birth	1.78 ***	3.23 ***	-	-
	(0.22)	*(0.78)*		
Duration 0 - 24 months	-	-	5.32 ***	9.02 ***
			(1.08)	*(2.64)*
Duration 25 - 36 months	-	-	1.61 +	2.97 ***
			(0.40)	*(0.95)*
Duration 37 - 60 months	-	-	1.01	2.15 *
			(0.28)	*(0.74)*
Duration 61 or more months	-	-	0.64	1.19
			(0.29)	*(0.58)*
Female (0/1)	-	0.72 ***	-	0.72 ***
		(0.05)		*(0.05)*
Age (spines)				
<25	-	0.90	-	0.90
		(0.07)		*(0.07)*
25 to <30	-	1.00	-	0.99
		(0.03)		*(0.03)*
30 to <35	-	0.91 ***	-	0.91 ***
		(0.02)		*(0.02)*
35+	-	0.81 ***	-	0.82 ***
		(0.02)		*(0.02)*

Table 4: *Continued*				
	Model 1	Model 2	Model 3	Model 4
Age of Youngest Child (0/1)				
36 months or less	-	1.00	-	1.00
		(0.00)		*(0.00)*
37 - 60 months	-	0.75	-	0.73
		(0.28)		*(0.28)*
61 - 120 months	-	1.91 *	-	1.73 +
		(0.57)		*(0.52)*
121 months or more	-	0.53 *	-	0.51 *
		(0.26)		*(0.25)*
Period, time varying (0/1)				
1950s, 1960s	-	1.00	-	1.00
		(0.00)		*(0.00)*
1970s	-	0.57 ***	-	0.58 ***
		(0.06)		*(0.06)*
1980s	-	0.97	-	0.97
		(0.09)		*(0.09)*
1990s	-	0.85	-	0.85
		(0.09)		*(0.09)*
Log Likelihood	-6038.92	-5919.94	-6020.56	-5904.73
AIC	12081.84	11865.89	12051.12	11841.46
BIC	12102.74	12001.77	12103.38	12008.69
df	2	13	5	16
N (Century Months)	255,807	255,807	255,807	255,807
N (Individuals)	2,478	2,478	2,478	2,478

Source: LNU 2000. Author's Calculations

Note: Hazard ratios estimated using Cox proportional hazards regression; regression coeeficient standard errors in parentheses

*** $p<0.001$ ** $p<0.01$ * $p<0.05$ $^{+}$ $p<0.1$