Max-Planck-Institut für demografische Forschung
Max Planck Institute for Demographic Research
Doberaner Strasse 114 · D-18057 Rostock · GERMANY
Tel +49 (0) 3 81 20 81 - 0; Fax +49 (0) 3 81 20 81 - 202;
http://www.demogr.mpg.de

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Twins or two single children: the influence of the multiplicity of the first birth on the divorce risk of Swedish women

Rainer Walke (walke@demogr.mpg.de)

This working paper has been approved for release by: Jan M. Hoem (hoem@demogr.mpg.de)
Head of the Laboratory of Contemporary European Fertility and Family Dynamics.

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Abstract

Based on Swedish register data, we compared the influence of a twin birth on the divorce risk with the influence of the sequential birth of two single children. The divorce risk for a woman with a very young child was lower than the risk for women without children or women with children older than 3.5 years. This behaviour was essentially independent of the number of children and whether or not the woman gave birth to twins. The effect of parity was much smaller than the effect of child age. The divorce risks for mothers of twins appeared to be between that of a mother with one child and a mother of two children.

Correspondence should be directed to Rainer Walke, Max Planck Institute for Demographic Research, Doberaner Str.114, D-18057 Rostock, Germany. E-mail: walke@demogr.mpg.de.

1 Introduction

The main purpose of this paper is investigating the difference between the intensity of divorce for Swedish women who gave birth to twins, and women who gave birth to two children in separate births. Often, mothers of twins are excluded from marriage disruption studies (Andersson, 1997; Andersson and Woldemicael, 2000) due to a presumption that they behave differently, and yet one does not really know where they fall within the established categories of parity in such studies.

To our knowledge, there is no study that compares the varying influence of twins and singletons on the divorce risk for mothers. Other types of studies do use twins as a natural experiment. Bronars and Grogger (1994) investigated the influence of unplanned children on the life choices of unmarried women. Using a technique derived by Rosenzweig and Wolpin (1980) they compared these mothers with a randomly selected control sample of unmarried women with single first births. In particular, they examined the potential effect upon labour-force participation, education, the propensity for a later marriage, poverty, and welfare contributions. They considered the birth of twins to be an 'exogenous' fertility event. Further more, they assumed that parents distribute their attention and energy equally among their children. This method proved useful in investigating the economic impact of unplanned births. Nevertheless, its assumptions hold only for economic events. Though divorces are certainly influenced by economic factors, these are not the only factors.

Many studies suffer from a small data basis. Fortunately, the Swedish register data set is large and comprehensive. Therefore, it is possible to investigate statistically rare events occurring in small sub-populations, as we do.

2 Data and methods

We have used a data set derived from Swedish population registers. This set covers all women born between January 1945 and December 1985 who have ever been registered as living in Sweden. For every woman we know her date of birth, her country of birth, and all of her children's dates of birth. Additionally, we know whether and when she either immigrated to Sweden or emigrated away from Sweden, her dates of marriage and divorce, and whether she herself or any current or former husband is deceased. For all these events, we know the exact year and month.

We then created a data subset containing all women with twins at first parturition and a random sample drawn from all other women. This subset contains a total of 104,970 women and it includes 14,536 mothers with twins at first birth. The size of the comparison group is six times the 'twins-at-first-birth' group. Therefore, it includes a sufficient number of parity 2 women with two single children. About twenty-eight percent of all women have parity 2 at the end of the study. A slight difference between the group sizes does not result in a substantial loss of analytical power (Piantadosi 1997).

For the hazard or intensity regression analysis of the risk of divorce in a woman's first marriage, we restrict the set to those women who married at least once. Only married women can divorce. The marriage ends either with a divorce, through the death of either spouse, with a third pregnancy, or emigration, or is censored nine months before December 1999, which is the end of the period of data collection. Within these parameters, there are 40,226 valid intervals of marriage. A total of 7,850 mothers gave birth to twins (at first birth), but as many as 6,686 of our mothers-of-twins were unmarried during our period of observation (1968–1999).

For our analysis we use a proportional hazard model. In this model, the risk probability of a woman to get divorced depends on the following factors: duration of the marriage, current age, calendar year, whether she was born in Sweden, and whether or not she has had children before her first marriage. Additionally, if she has children, the divorce risk depends on the age of the first single child, the age of the second single child, and in case of twins at first birth, on their age as well. For most of these properties, we use piecewise-linear duration splines. Assuming you have enough nodes, they adjust to any pattern in the data set.

We use the software package aML Release 1.0, created by Lillard and Panis (2000). Among other capabilities, it supports hazard models with piecewise-linear log-hazard duration dependencies.
The mathematical representation of the exponential-regression formula (cf. Hoem 1987) for the divorce intensity of a woman *i* at exact marriage duration *t* (years) is:

$$\mu_i(t) = \mu_0(t)\exp\left\{\sum_{j=1}^{m} \beta_j x_{ji}(t)\right\}.$$

We used a piecewise linear approximation $\gamma(t)$ to the baseline log-intensity $\ln \mu_0(t)$. In the resulting model, we see two binary variables and five duration splines in the regressor set. The binary variable for unmarried births and the calendar year spline are connected in an interaction term. Therefore the intensity formula reads:

$$\ln \mu_i(t) = \gamma(t) + \beta_1 x_{1i} + s_3(t - t_{3i}) + s_4(t - t_{4i}) + s_5(t - t_{5i}) + s_6(t - t_{6i}) + s_7(t - t_{7i}, x_{2i}) .$$

The binary variable x_{1i} is 1 if woman *i* was not born in Sweden. The following spline s_3 covers the dependence on the current age of woman *i*. It refers to the risk of a woman 30 years of age i.e., t_{3i} is the duration of the marriage of woman *i* at her 30th birthday.

The following three splines are conditional splines. Until a certain event, they are zero. When that event occurs, they spike and continue as a piecewise linear spline. The spline s_4 starts at the birth of the first singleton child if any, and t_{4i} is the duration of marriage of woman *i* at the time of birth. Similarly, the spline s_5 starts at any second singleton birth. If a woman has twins in the first birth, the spline s_6 starts.

The last spline s_7 forms an interaction term. A binary variable x_{2i} is 1 if woman *i* gave birth to a child before her first marriage. This variable interacts with a spline that denotes the influence of the specific calendar year. Therefore we have two different splines depending on x_{2i}, s_{70} in case of a woman without a premarital birth and s_{71} in case of a woman with a premarital birth. This interacting term is normalised to women without a child before marriage and to the basis year 1980 i.e., t_{7i} is the duration of the marriage of the woman *i* as of 1 January 1980.

Each spline, $s_3 \ldots s_7$, has the same shape for all women. They shift over time by $t_{3i} \ldots t_{7i}$ depending on the history of woman *i*. The nodes of all splines were pre-specified by us. We adjusted them by hand to achieve a sufficient representation of the spline shape. A spline with *n* nodes needs *n+1* slope parameters; but the s_7 interaction term contains *n+1* slope parameters for each spline s_{70} and s_{71}. Additionally, the splines γ, s_4, s_5, s_6 and s_{71} have an intercept parameter. The aML program varies all of these slope parameters as well as β_1 to get the maximum likelihood estimates for these parameters.

3 Impact of twins on the divorce risk

Figure 1 shows the logarithmic divorce intensity dependent on marriage duration expressed in years. This is the baseline function. For a thirty-year-old woman born in Sweden, without children, it shows the real divorce risk per year valid for 1980. For example, in the fifth year of marriage 3.3 percent of all married such women divorce. Understandably, the divorce probability is very low immediately after marriage. It increases rapidly and reaches a maximum after about three years, after which it stays roughly constant and decreases slowly only after fifteen years of marriage. Detailed estimate values can be found in our appendix.

Figure 1: *Divorce risk per year for Swedish women, by duration of first marriage.*

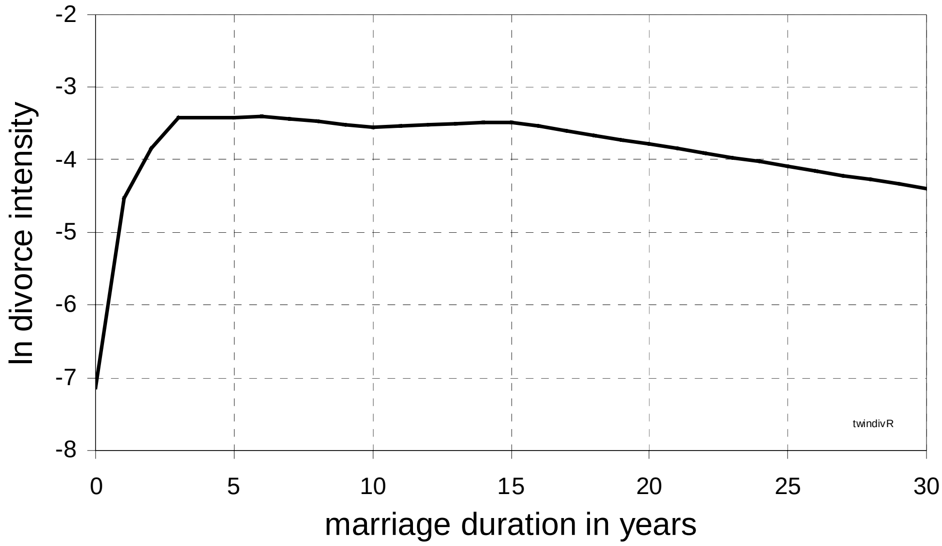


Figure 2 shows the logarithmic effect of the calendar year on the divorce intensity. It uses a baseline year of 1980 and the baseline intensity mentioned above. The black curve denotes women who had not given birth before marriage, while the grey curve shows the calendar effect for women who had at least one child before marriage. The divorce risk (black curve) increases rapidly between 1970 and 1973. There is a jump at the end of 1973. This is very likely due to the liberalisation of the divorce law in Sweden at this time. After 1973, there is a slower increase in the risk until 1990. The grey curve does not have the same jump. The divorce risk gradually rises with the calendar year for women with a child before marriage. This difference between the two curves is an effect of parity. Andersson (1997) showed that the strongest increase in divorce risk in 1974 was for parity 0.

Figure 2: *Divorce risk dependence on calendar year.*

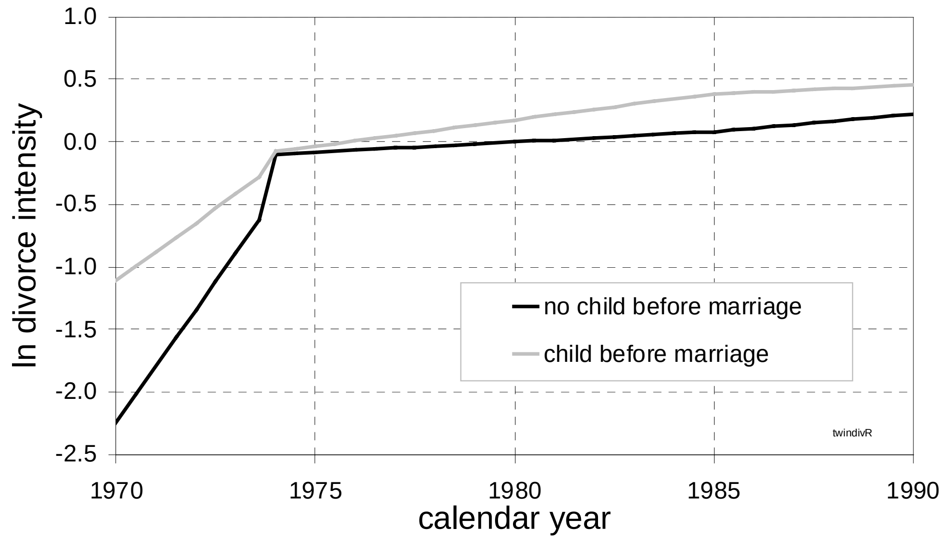
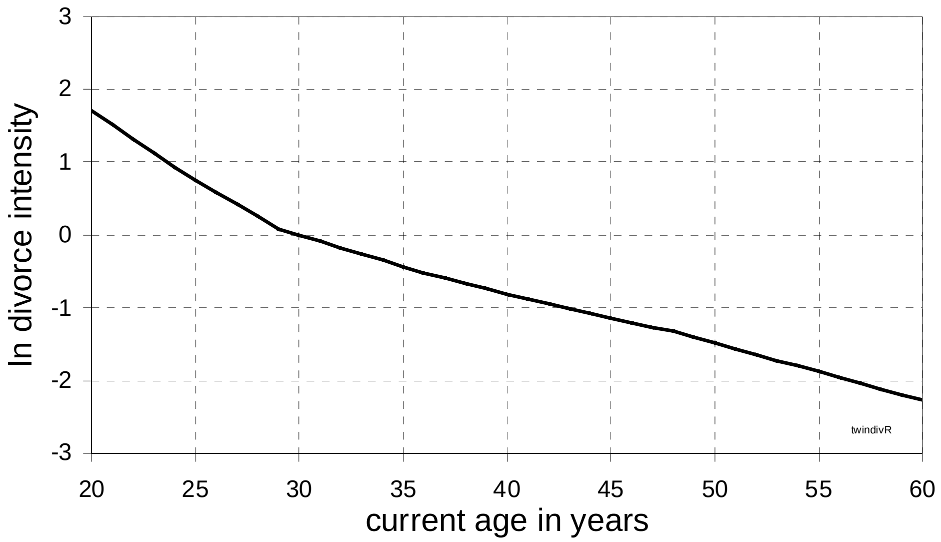


Figure 3 shows the effect of the age of a woman on the divorce risk. The baseline age is 30. There is a significant additional divorce risk for younger women i.e., the risk for a twenty-year-old is approximately five times higher than that of a thirty-year-old (exp(1.70) = 5.5). This risk decreases rapidly between the ages of twenty and twenty-nine. The slope then flattens, but it continues to decrease with increasing age.

Figure 3: *Divorce risk dependence on the age of the woman.*



Our last figure, *Figure 4*, shows the effect of the age of the youngest child by parity and singleton/twin mother status. This is the main variable of interest to us. The baseline is set at women without children. The model is extended by an additional duration spline at each birth. This spline differs depending on which type of birth (multiple or singleton) occurs. After parturition the divorce risk is lower than for the childless baseline group. The risk then begins to increase again and plateaus after the child reaches 3.5 years. Three different curves are given in order to depict the varying effects of child age at different parities and singleton/twin status.

All three curves show a similar behaviour. Mothers with a singleton first birth have the lowest divorce risk in the months immediately following birth. The risk is higher for women who gave birth to twins. It is further higher for mothers who had a second singleton. The effect of the type of birth is reversed after 3.5 years. For mothers of only one child, it reaches a level slightly above the baseline level. In case of twins, it does not reach this baseline intensity and for two singletons, it is clearly lower. Roughly spoken, the divorce risk of mothers of twins lies between mothers of a single child and mothers with two singletons.

Figure 4: *Divorce risk dependence on the age of the children.*

It remains for us to report the result for the binary variable that represents where the woman was born. The log-divorce-intensity of women not born in Sweden is 0.216 (0.037) higher then those born in Sweden. This translates into an additional risk of twenty-four percent, because exp(0.216) = 1.24.

4 Conclusions

The divorce intensity of Swedish women is strongly reduced when a child below age 3.5 years is present. Andersson (1997) showed similar results. This does not depend on whether the woman has twins or singletons. Quantitatively, there is a smaller effect of parity. The divorce risk of parity 2 mothers of twins falls somewhere between that of mothers of parity 1 and parity 2 with singletons.

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Appendix

The following tables contain information produced by the aML program for our diagrams above. It includes the node positions, the estimates of intercepts and slope parameters, and a standard error derived by numerical computation of the Hessian matrix (Lillard and Panis 2000).

Marriage duration (Figure 1):

	log-intensity	standard error based on numerical Hessian
intercept	-7.1395	0.1799
slopes		
0–1 years	2.6114	0.2062
1–2 years	0.6779	0.0929
2–3 years	0.4189	0.0646
3–6 years	0.0062	0.0179
6–10 years	-0.0340	0.0141
10–15 years	0.0132	0.0132
above 15 years	-0.0612	0.0109

Calendar year and child before marriage (Figure 2):

	child before marriage		no child before marriage	
	log-intensity	standard error based on numerical Hessian	log-intensity	standard error based on numerical Hessian
intercept	0.1759	0.0465	–	–
slopes				
up to 1973.5	0.2315	0.1630	0.4505	0.1121
1973.5–1974.0	0.5085	0.5367	1.3064	0.3269
1974–1985	0.0417	0.0096	0.0163	0.0059
over 1985	0.0142	0.0050	0.0286	0.0039

Current age (Figure 3):

slopes	log-intensity	standard error based on numerical Hessian
up to 24 years	-0.1947	0.0212
24–29 years	-0.1673	0.0105
29–36 years	-0.0870	0.0076
36–42 years	-0.0720	0.0108
42–48 years	-0.0624	0.0171
over 48 years	-0.0785	0.0453

Age of child (Figure 4):

	mothers of one single child		mothers of two singleton children		mothers of twins at first birth	
	log-intensity	standard error based on numerical Hessian	log-intensity	standard error based on numerical Hessian	log-intensity	standard error based on numerical Hessian
intercept	-3.8652	0.4964	-2.4769	0.2533	-3.0955	0.6954
slopes						
0–1 years	1.4508	0.5720	0.3757	0.3248	0.5203	0.8719
1–2 years	1.5226	0.2287	1.0359	0.1943	1.5672	0.4601
2–2.5 years	0.5895	0.2861	0.3253	0.2839	0.8457	0.5571
2.5–3.5 years	0.6875	0.0942	0.3272	0.0947	0.2132	0.1676
over 3.5 years	0.0236	0.0054	0.0193	0.0062	0.0361	0.0082

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