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ON POPULATION, FAMILY AND WELFARE

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ABSTRACT

Post-communist governments in Hungary made serious efforts to stop the massive fertility decline that started with the fall of the communist reproductive system, or at least to reduce it. Two of several interventions by the Hungarian government – the lengthening of child-raising allowance (*'full time motherhood'*) and of a new tax-relief system –, were aimed specifically at supporting those having three and more children. However, the relevant statistics have not shown a growing ratio of third or further children. Here the authors employ an event-history method and use the data of the *Hungarian Generation and Gender Survey* to examine whether the government interventions had an effect of encouraging partners with two children to have a third birth, and whether specific social groups displayed different behaviour in this respect. The policy interventions have been measured by period indicators. After controlling for standard factors affecting third birth, the authors concluded that indeed, the government interventions have a significant, and differentiated effect on third- birth risks. While the massive lengthening of the child-raising allowance increased the third-birth risk of those with low educational attainment, the introduction of a generous tax relief seems to have done likewise for those with a tertiary education.

Keywords: population policy, low fertility, third births

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INTRODUCTION

Two government measures intended to encourage the birth and upbringing of a *third* or further children were introduced in Hungary in the quarter-century following the change of system in 1989. *In 1993*, during the most intensive period of socio-political change, *monthly cash allowance* for parents of a third or subsequent children was extended from the third up to the eighth birthday of the youngest child. The other government measure, *in 1999*, introduced a *comprehensive tax-relief system,* where the relief was generously increased. The main beneficiaries were parents raising three or more children who had taxable income. So the first measure aimed mainly to increase the welfare of large, disadvantaged families, while the second sought directly to raise fertility. The prime purpose of this analysis is to pinpoint the effect these measures had on fertility, namely on third birth risk.

A period beset with social transformation was in no way favourable to measures of socio-political intervention. The first such measures came in the most intensive phase of economic transformation, in the very beginning of the 90s. Privatization and initiation of a free market were proceeding at full blast: GDP sank by 11.9 per cent in 1991 and 3.1 per cent in 1992; joblessness became a personal experience of many people in 1992, as the unemployment rate climbed from virtually zero to 12.7 per cent. The 1999 tax concession likewise appeared in a social troubled period. A couple of years earlier, in 1995, the overall principle of universal assistance under the family allowance system had been replaced by means test, but this was reversed by the new government of 1998: the need principle gave way to a still more comprehensive system of universalism. There was also little cause for optimism in economic conditions at the turn of the millennium. Although performance was growing and unemployment falling, it was not until 2000 that GDP returned to its 1989 level.

So questions and doubts are in order (Gerber and Perelli-Harris, 2012; Neyer and Andersson, 2008); can specific family-policy measures be effective in a period when almost *all aspects* of social structure and social coexistence are undergoing deep changes? Can one expect family policy to be effective when *child-bearing behaviour is altering*, when society is at the beginning or in the midst of a process switching from an early child-bearing pattern to a later one? Of course the questions can be put from other angles– is administrative intervention not intended to offer some security against unpredictability, to provide a safety net? In other words, is it not the right time for expecting the welfare provisions of the state to meet their goals?

Nor does literature on the effectiveness or efficiency of population policy offer convincing arguments (Demény, 1987; Gauthier, 2007). Although the efficacy of some family-policy measures has been shown in Sweden (Hoem, 1993), Austria (Hoem et al., 2001; Lalive and Zweimüller 2009), Canada (Milligan, 2005) and Israel (Cohen, Dehejia and Romanov, 2013), the results of international comparative studies are mixed or inconsistent. As a comprehensive study by Gauthier (2007) puts it, "The analysis […] suggests that policies may indeed have an effect on families, but that effect tends to be [of] a small magnitude and […] possibly have an effect on the timing of fertility, rather than completed family size" (p. 339).

Results of analyses on population policy interventions in individual *post-communist countries* are mixed as well. Results suggesting positive effects tend to refer overwhelmingly to periods before the change of system (Andorka, 1996; Bütner and Lutz, 1990; Kapitány, 2008; Gábos et al., 2009). Analyses focused on the period after 1989 show ineffectiveness of administrative measures (Stropnik and Sircelj, 2008; Frejka and Zaharov, 2013).

The analysis here, focusing on Hungary, shows that despite unfavourable conditions, the two sets of measures mentioned, to differing extents, could

indeed influence the third birth risk. So they must, with qualifications, be classed as successful.

The analysis is built up as follows. First comes a non-comprehensive survey of the literature on the efficacy of the monetary measures of family policy. Second is a short introduction to the features of Hungary's family-allowance system, including details on the key institutional changes affecting the decision to have a third child. Here we present our hypotheses on these institutional changes. Then come the data, the strategies of analysis, and the dependent and independent variables. Discussion of the results concentrates on the individual social-policy periods, but not omitting the relationships that the control variables help discover. Finally, the results are presented.

THE EFFECTS OF FAMILY-POLICY INTERVENTIONS – A SELECTIVE REVIEW

THE COMPARATIVE VIEW

This review of the literature is not aimed at completeness. That has been covered by others (Gauthier, 2007; Thévenon and Gauthier, 2011), whose work is used here as a basis. Both cited summaries show clearly that family-policy interventions have an effect on fertility, but its scale is small and unsystematic. The latter deficiency can be put down to the heterogeneity of the policies and effects, and the differing results according to social status and number of children. However, the recent analysis by Luci-Greulich and Thévenon (2013) differentiates five characteristic types of family support that all affect the course of fertility in developed countries. They underline the strong significance of the "mix" of single measures. No such comprehensive, all-embracing analysis is available for the post-communist countries. Frejka and Zakharov (2013) identify a timing effect of the "baby bonus" introduced in Russia, but express scepticism about the cohort effects of the measures. Frejka and Gietel-Basten (2016) employs an "intuitive" way to link the trend in fertility to the various approaches of state intervention in Central Eastern European countries and ultimately underlines their general inefficacy (p. 35). The following review focuses on analyses that examine the effect of *lengthening the period of childcare leave (cash allowance)* and the impact of the direct *monetary or tax-relief increases*. Special heed is paid to the findings that examine the chance of a third child being born.

PARENTAL LEAVE

The basic function of paid parental leave is to enable employment and reconcile it with child-raising. Three vital features of the system are 1) to allow the parent to be with the child in the leave period, 2) largely to compensate the parent for lost earnings, and 3) to ensure a return to the same job after the leave period. Without such a scheme the high level of female employment and the double-earner family model would be jeopardized. While job guarantees are widespread and play a clearly positive role, countries differ markedly in the length of the leave period and the extent of compensation for the wages lost. Nor is there agreement on their effect on fertility. It is assumed that a *moderate period of leave* with a *high offset against lost earnings* promotes fulfilment of parenthood plans and a return to the former place of work (D'Addio and D'Ercole, 2005; Matysiak and Szalma, 2014), so keeping the indirect opportunity cost of child-raising low. Yet low or partial compensation – whether the leave is long or short – hardly reduces such child-raising costs. The drawbacks of excessive parental leave usually mentioned are mainly

long-term ones (less chance of return to a former place of work and reduced career earnings).

Empirical analyses of the effects of extending parental leave on childbearing in specific countries point the same way, but with some inconsistencies. Interesting for Hungary is Austria's case, where a relatively high flat rate of parental payment (some 40 per cent of female net income) was extended in 1990 from one year to two. Both Lalive and Zweimüller (2009) and Hoem, Prskawetz and Neyer (2001) concluded that this extension had increased the chance of a third child being born: the former study looked at the birth of the "next child", while Hoem, Prskawetz and Neyer (2011) looked at the possible birth of a third child. However, an analysis of this measure by Stasna and Sobotka (2009) stated it as a matter of timing, but not a quantum effect. Lalive and Zweimüller (2009), dissecting the effect mechanisms, concluded that higher paid leave and a longer job-retention guarantee each had an incentive effect separately. Introduction of home-cash-for-care benefit in Finland increased the number of third children born (Vikat, 2004). Aassve and Lappegård (2011) reached similar findings in Norway, again for home-cash-for-care benefit, introduced to compensate those staying at home after their child was one.

Andorka, looking at former socialist countries, found that introducing the paid parental leave in communist Hungary raised fertility among those with higher educational attainment (Andorka, 1996). Based on modelling, Kapitány concluded that the wage-related paid parental leave, introduced in 1985, had raised the number of children born between 1986 and 1996 by an estimated 10.2 per cent, and although timing of births was a factor in some cases, so was an increase in fertility among older cohorts (Kapitány, 2008). In a comparative study, Matystiak and Szalma (2014) showed there were greater chances of a second child being born in Hungary, where paid parental leave was a much longer 36 months, than in Poland, where it was only 6 months. Studies by Gerber and Perelli-Harris (2008) suggest that parental leave in Russia had a decisive effect on second births in the "turbulent" period of social changes, by easing considerably the costs of the social transition for those who undertook to bear children.

CASH BENEFITS

Cash benefits aim to reduce the direct costs of child-raising. Though many forms exist, it is worth grouping them by frequency. The regularly received, by and large monthly benefits normally contribute to child-raising costs up to adulthood (the end of the 18th year), whether as direct payments or as tax concessions. One-time benefits are tied to the birth event and have sought traditionally to bridge difficulties around that time. Among the favoured forms of birth stimulants these days are large sums paid as a "baby bonus". International comparisons for the form lead to the plain conclusion that fertility is stimulated by the scale (size) of such payments (D'Addio and D'Ercole, 2005; Luci and Thévenon, 2011; Thévenon and Gauthier, 2011).

For the subject of this paper special importance attaches to analyses or "natural experiments" focused on institutional change in a given country– whether it takes the form of a new scheme or increases benefits under an existing one. Examining the French institutional system is particularly popular because family policy there has a long history, the benefits are generous, and the changes made have been appreciable. It is generally accepted that French family policy contributes strongly to French fertility, which is high by European standards (Toulemon et al., 2008; Thévenon, 2011). Especially relevant here is the work of Breton et al. (2005), presenting the fertility effects of measures taken in 1978–82, which were notably favourable to bearing a third child, as both monetary benefits and tax concessions were raised. Breton et al. showed how the parity progression ratio of the third birth clearly increased. Laroque

and Salanié (2008) were able to identify positive effects from the tax concession on the birth of a first and a third child.

On the other hand, a mixed picture emerges of the fertility effects of the Working Family Tax Credit introduced in the United Kingdom in 1999. This increased the child-related tax concession for lower income brackets. Brewer et al. (2009) concluded that fertility in lower-income, double-earner families had increased as a result, but Ohinata (2009) found the fertility effect on double-earner families imperceptible and saw at most a timing effect. A very limited rise in fertility was also found in one recent analysis of the effect of a tax concession in Spain (Azmat and González, 2009). Gábos et al. (2009), examining the effect on total fertility rate of the direct monetary benefits and tax concessions in Hungary, found that a change in the real value of these produced a change in fertility in the following year.

Several analyses are available of the fertility effects of single, lump-sum payments associated with birth event. The Allowance for Newborn Babies (ANB) introduced to stimulate fertility in the Quebec province of Canada was examined by Milligan (2005) using the difference-in-differences method. This scheme paid once for first and second children and a larger sum several times to those parenting a third child. It is relevant here to emphasize that ANB encouraged the birth of third children. The "baby bonus" introduced in Australia and subsequently popular elsewhere was likewise found to promote fertility (Drago et al., 2013; Sinclair et al., 2012). Not so the "baby bonus"-type "maternal capital" introduced in Russia expressly to raise fertility and offering sizeable monetary support, which according to Frejka and Zakharov (2013) simply affected timing (brought births forward). All in all, innumerable family-policy changes were made in the former socialist countries after the change of system (Frejka and Gietel-Basten, 2016), but no empirical analyses of their effects are known.

SOCIAL STATUS AND ITS EFFECTS ON FERTILITY

There is good reason to assume that incentives of childbearing, in this case the birth of a third child, will affect the various groups in society to differing extents. Flat-rate, lump-sum support tends to encourage low-income groups, and income-related support those of middle or high income. Although the studies mentioned so far are inconsistent in some respects, they offer indications on this matter as well. The effect of lengthening the period of parental leave in Austria was a greater rise in fertility among those of low income (Lalive and Zweimüller, 2011). In Norway similarly, those of low income were more prone to extend their period of paid leave and raised their chances of having a second or third child (Aassve and Lappegård, 2009).

In Hungary, the 1996 change from wage-related leave to an income-tested one-time payment produced a drastic fall in first childbirths among those of higher income status. Their childbearing behaviour returned to its earlier level after the measure was reversed in 1998 (Aassve et al., 2006). The French analysis did not deal explicitly with the structural effects of policy, but it was assumed that all social groups would be affected equally. It emerges from the models that two-child families in France with high educational attainment were more likely to have a third child; this pattern remained after the measures were taken. The Australian studies found that the "baby bonus" gave lower income groups greater encouragement to have further children (Drago et al., 2013). In Quebec, on the other hand, it was found that the similar ANB tended to increase fertility more among higher income groups (Milligan, 2005). So at first sight the findings seem inconsistent with each other. Perhaps they would be less so if it were possible to include as a factor the essential differences of the available welfare mix in the mentioned countries.

There is good reason, therefore, to pay attention to the differing fertility consequences of specific welfare schemes.

THE HUNGARIAN SYSTEM OF FAMILY ALLOWANCES

GENERAL FEATURES AT THE TURN OF THE MILLENNIUM

Hungarian family policy can be seen in general as generous (Szelewa and Polakowski, 2008; Thévenon, 2011:71), but difficult to predict (Spéder and Kamarás, 2008).[1] The longest-standing and best-known benefit is *family allowance*, paid to all monthly up to the 18th birthday of the child. Mothers receive a small *single payment* of maternity benefit on giving birth. Mothers (in some cases fathers) may claim *parental leave* to look after their child up to the end of its *third year* and receive *childcare support* during that period. Those hitherto employed receive 75 per cent of previous earnings in a first period of six months, 65 per cent *with a ceiling* in a second period (seven to 24 months), and a flat rate irrespective of previous earnings in a third period (25 to 36 months). Parents decide how long they will stay at home with the child and when they will return to their job. Those without previous employment receive flat-rate pay during the first 36 months. The Hungarian family-benefit system also supports parenting through *the tax system*, albeit to a varying extent (see later). State assistance with setting up a family is also given through various *housing-support schemes*. Finally, parents or children receive several other types of assistance in kind (largely free day care, pre-school and school education). On the whole, Hungary's cash benefits can be seen as generous even among the OECD countries, as it devotes 3.57 per cent of GDP (2013) to such family assistance, giving it a permanently prominent rating.

Hungarian family policy has undergone several radical changes in the last quarter-century (Spéder, 2016), so that it qualifies as *unpredictable*. The subject is always prominent in electioneering, and regular changes are made with each new government. The change of system saw a broadening of *universal* benefits, but this principle gave way to one of *need* in 1995. In 1999, a still broader principle of *universalism returned*, complemented by tax concessions. The spread of universalism continued with the 2002 and 2008 governments. (Looking at the two decades, they can be seen as a period of stop-and-go in family policy.) To sum up, the generosity of Hungarian family policy may have been overlooked by its potential beneficiaries because the continual changes and sharp debates led to an image of unreliability.

"FULL-TIME MOTHERHOOD" – BASIC INCOME-SUPPORT FOR LARGE FAMILIES (1993)

The conservative political commitments of the first freely elected government (1990–94), with squeezing (radical constriction) of the labour market and widening of income inequalities, bore heavily on families with several children and prepared the way for a new scheme of family policy. *Child-raising support (GYET)*, introduced in 1993, assured families a *fixed monthly sum* for a third child and subsequent children *up to the age of eight*. The payment matched the childcare allowance *(GYES)* available until the age of 36 months for the newborn. On introduction it amounted to two-fifths (39 per cent)

1 Makay (2015) presents details of the current Hungarian system of family benefits; Szelewa and Polakowski (2008) compares the post-communist countries in this respect. Spéder (2015) looks over the long-term development in the links between demography, fertility and state intervention.

of average female pay and from 2000 to one-third (33 per cent) of it.[2] So it was not negligible– could be called significant, in fact. Initially the scheme paid those who had had a job (were socially insured) and where the family's per capita income did not exceed three times the minimum old-age pension (an income test). However, the second qualifying criteria hardly reduced the number of claimants in practice.[3] The requirement of a previous job (social insurance) was dropped in 1999, but the income ceiling remained. Nonetheless, the number of potential claimants increased somewhat. The scheme works as an alternative to employment and also to unemployment. So understandably, the expression *"full-time motherhood"* has become widespread.

The new scheme gave women a *choice*: either send a third child reaching the age of three to kindergarten and herself return to the labour market, or raise the children at home. These were real alternatives in institutional terms. Firstly, there has been almost full availability of places in kindergarten for the over-threes, although there may be local anomalies. Secondly, previous employers are required to restore the mothers of small children to their jobs after their 36-month parental leave (job security). Of course labour-market fluctuations may prevent this happening in each and every case, but in general the acceptance of child-raising support as some kind of "extension" of unpaid leave is not compulsory, but the result of a free choice. It may also reflect that parents of large families prefer *personal provision and upbringing* to institutional ones. Furthermore, "full-time motherhood" *may avert* the possibility of *unemployment* in periods of recession. So willingness to have and raise a third or further children may mark a response to labour-market problems: parenthood becomes an alternative to unemployment and inactivity, and one that allows a legitimate female role to be performed (compare Vikat, 2004).

Finally, "full-time motherhood" does not rule out ultimately the role of employment. On the one hand, restricted amounts of work can be done alongside it. When it was introduced, claimants could work *part time* (up to 20 hours a month). This restriction has been eased since. At present they may even do a full-time job if it is done from home, i. e. the place where the childcare occurs.[4] On the other hand, agreement between employer and employee can be reached to keep the job open by classing the whole motherhood period as unpaid leave.

Though Hungary's family policy between 1990 and 2010 was marked by *unpredictability*, punctuated by 180-degree turns, the institution of providing a basic income to large families *(GYET)* until the youngest child reaches the age of eight was never questioned or disputed or subjected to change. Throughout the period, the eight-year-long "full time motherhood" *(GYET)* programme remained as a *reliable element of Hungarian family policy*, even though there were continual and profound changes within it. In times of stop-and-go policy, to postpone childbirth is not only advantageous, but rational (Rodin, 2011), but the unquestioned existence of the paid support outlined above may well have *increased the inclination to have a third child (1st hypothesis)*.

The universalism of "full time motherhood" has been emphasized here, but presumably the rate of claim *differs also by social status* due to the substantial opportunity cost it may involve. For some, a long break in their working career may lose them workplace promotion or lessen their life-time earnings, or for weak employees, cause exclusion from the labour market. Nor can decades of absence from the labour market be afforded by those who are career-oriented, or those for whom doubling their - earnings is essential to maintaining their standard of living. Meanwhile those without a job or in a job that is peripheral to the labour market may find child-raising support a lucrative solution. So

2 This was a time of strong inflation in the Hungarian economy (CPI 1993: 22.5%; 2000: 9.8%), to which the benefits were adjusted occasionally, not continually.

3 At the time the decisive majority of such women had years of service to qualify as claimants, and the generally low pay rates meant few large families reached the per capita income ceiling.

4 In 2005, 8.6% of beneficiaries did paid work regularly and 8.5% occasionally (Váradi, 2006).

we assume that the provision motivates having a third child to different degrees among different social groups. In other words, we assume that *it is among those with lower income – in our case lower educational attainment– that this policy measure increased the chances of a third birth (2nd hypothesis)*.

CHILD-RELATED TAX CONCESSIONS (1999)

Personal income taxation was introduced in Hungary in 1986. Various concessions were made on various grounds, but initially, child-raising costs were considered only to a symbolic extent. The personal income tax paid was largely unaffected by the number of the taxpayer's children.[5]

According to 1999 rules, valid from 1 January 2000, child-raisers could claim a tax rebate depending on the *number of children*. The concession per child rose with the number of children. This progressiveness was enhanced in 2001: the monthly relief for a single-child family became 3000 HUF, for a two-child family 4000 HUF per child, and for three or more children 10,000 HUF per child.[6] This clearly favoured large families. The full concession for a three-child family in 2000 came to almost an eighth of average female pay (18 per cent) and in 2001 to *half of it (51 per cent*).

The relief for a three-child family was a *universal*, *flat-rate type* of benefit, *yet indirectly it was strongly differentiated*, as family income had to be quite high before the whole amount could be claimed. This still applies. Those in lower income brackets pay either no tax or a low amount,[7] and the figures mentioned cannot be claimed back. There are no data available on what proportion of child-raising families failed to earn enough for the full concession, but certainly those in the lowest third of income dispersion have been unable to claim the full amount, for which medium or high taxable income was required. Flat-rate benefits (including the tax concession) generally encourage those of lower income and social status to have more children (Aassve and Lappegård, 2011; Thevénon and Gauthier, 2011), but the income distribution and taxation system of Hungary enables those of *medium or higher income* (in this case medium or higher educational attainment) *to receive benefit and may encourage them to have the third child (3rd hypothesis)*.

ANALYTICAL STRATEGY AND DATA

DATA AND METHODS

For the analysis we use an event-history database created from the first four waves of the Hungarian Generations and Gender Survey (Vikat et al., 2007; Murinkó and Spéder, 2016). The first wave of the survey took place in 2001 and the fourth one in 2011–12 (their main features appear in *Table A2* of the Appendix.) Our sample is confined to women and men who had two children in 1980 or onwards and who took part in the first and second wave of data collection. The dependent variable for the event-history analysis is *the month of the conception of the third child* (date of the third birth minus nine months). The risk period, measured in months, begins with the birth of the second child and ends with the conception of the third one. Observations are censored in three cases: if the event failed to occur within eight years of the birth of the second child, if the marriage or partnership broke up, or at the time of the last interview, whichever happened earlier.

5 The study cannot cover the tax system in detail. For information on this, see Kupa 1988, EC 2016:106-111.
6 The scale and development of the tax relief appear in *Table A1* of the Appendix.
7 It is not possible to cover here how many people receive the minimum wage in Hungary or why.

Eventually 2329 cases were included in the analysis, and a third child was born in 745 cases. The number of person-months is 148,825. The number of excluded cases and the reasons for doing so can be found in *Table A3* in the Appendix. In the analysis we employed the method of piecewise constant event history modelling, which assumes that the risk is constant in each category of the independent variables.

THE CENTRAL VARIABLE IN THE ANALYSIS: POLICY PERIOD

One important aim of our analysis is to divide the examined period (1980–2012) into shorter periods relevant to population policy and third birth. We assume that if the risk of a third birth increased or decreased significantly (considering the control factors), then it was affected by *measures introduced at the beginning of the period* or closely related to them. We are aware that this assumption is not free of problems, since numerous other events in the period (confounding factors) may have influenced birth risks (cf. Milligan, 2005). One serious argument in favour of the assumption is that the benefits introduced (or withdrawn) were strong enough for having such effects. No less important is that others have successfully used this analytical strategy in examining population-policy effects (Hoem, 1993; Hoem, Praskavetz and Neyer, 2001; Aassve, Billari and Spéder, 2006; Lappegård, 2010).

Table 1

The division of the analysed period into policy periods

Policy periods	Main features of family policy changes
January 1980 – December 1988	Period of state socialism, a complex system of reproduction
February 1989 – January 1993	Intensive period of social transformation but no change in the institutional setting of family policy
February 1993 – December 1998	Introduction of "full-time motherhood": five-year extension of flat-rate payments for third or subsequent children up to the 8th birthday of the youngest child
January 1999 – December 2005	Generous increase in the tax relief for families with three or more children
January 2006 – March 2009	High universal family allowance (unchanged "full-time motherhood", decreased tax concessions, increased family allowance)
April 2009 onwards	The length of parental leave was reduced from 36 to 24 months

CONTROL VARIABLES

As far as the data makes it possible, the analysis includes control variables thought advisable in the literature on third birth (Berghamer, 2009; Breton et al., 2005; Corman, 2002; Hoem et al., 2001). These findings impinge only slightly on the analysis in this paper, and so the reasons for choosing the control variables are not given in detail. The statistics describing them appear in *Table A4* of the Appendix.

Of the strictly *demographic variables*, heed is taken of the sex of the respondent, age at second birth (a categorical, dynamic variable), the sex of the first two children (whether the same or different), and the number of the respondent's siblings (cf. Breton et al., 2005). It should be stressed that the age when the second child was born was viewed in relation to educational attainment (see *Table 2* below). The chance of bearing a child is strongly linked to and more appreciable after schooling is completed, and so taking the relevant literature (see Corman, 2002; Breton et al., 2005; Hoem et al., 2001), we employed relative age categories (*early, average, late*) instead of the absolute

age of the respondent (see *Table 2*). So the variable is set as a function of educational attainment, with groups having a second child at an early, average or late stage in their life. The literature shows that parents whose second child was born when they were young or in an early stage of their life-course (see above) are more likely to have a third child.

Table 2

Age categories for the variable "relative age at second birth" by sex and education

	Women			Men		
	Early	Average	Late	Early	Average	Late
Primary or less	16–19	20–23	24+	16–22	23–26	27+
Vocational training school or secondary	16–22	23–27	28+	16–25	26–30	31+
Tertiary	16–24	25–30	31+	16–27	28–33	34+

A problem that has become a theme in demographic literature more recently is whether it is more common for the divorced and the re-partnered to have a second child than in intact relationships (Billari, 2005). The analysis here does not include two-child families in which the parents separated and established a new relationship after the birth of the second child (these cases are censored). However, information is available on those whose relationship dissolved between the first two births, so that the second child was born in a new union. We assume that the chance of a third birth is higher among such people because it would mean having the second common child in the new relationship.

There is strong interest in the role of *educational attainment* in child-bearing decisions, including a third child. Numerous analyses have shown fertility as inversely proportionate to female educational attainment. This was the trend with the birth of a third child as well. However, several analyses show that when the chance of a third birth is examined among two-child individuals, the rate is higher for those with high educational attainment than for those with medium one (Kravdal, 2001; Breton et al., 2005). Hoem et al. (2001), in examining changes in family policy in Austria, looked comprehensively and meticulously at the role of educational attainment in the chance of a third child. They found that the educational level of the women plays no significant role. We assume, based on earlier analyses (Bartus et al., 2013; Spéder and Kapitány 2007.), that those with a medium level of education will have the lowest risk of having a third birth in Hungary.

The analyses assigns a central role to the educational attainment variable, *notably the combined effect of educational attainment and social-policy periods*. We hypothesize that the extension of flat-rate paid parental leave raises the chance of a third birth among those of lower status the most *(2nd hypothesis)*, as does tax relief among those with medium and/or high status and income *(3rd hypothesis)*. Since no data on income or social stratification are to hand, the aim is to proxy income status with a three-level educational attainment variable: primary, secondary and tertiary. We expect the attainment to exert different effects on the birth of a third child in different policy periods.

Finally, ethnicity and religiosity are also included in the control variables. It is expected that those of *Roma ethnicity* (Koycheva and Philipov, 2008) and those with *religious* beliefs will have a notably higher chance of a third child (Berghammer, 2009).

ANALYSIS AND RESULTS

The models in the study build on each other. Our main goal is to examine the period effects. Initially only the variable of policy periods appear in the model, joined step by step by the control variables, and finally by the interaction effects with the periods.

The first model, including only the family-policy periods, suggests that, compared with the socialist period, fewer two-child families eventually had a third birth under the reproduction system that developed out of the transition. Based on the raw data, the two examined policy interventions did not lead to higher risk of third birth but to a lower one.

The second model shows the effect of the *demographic control variables*, which largely meet our expectations about the extent and manner of influences and hardly change in the further steps of modelling. The risk of a third birth is higher for those who had the first two children relatively early and whose first two children are of the same sex. The highest risk of a third child appears within three years of the birth of the second one– conception within 15 and 27 months. (This is discussed again later in the study.) There are higher risks of a third child if parents come from large families (4+ children), and the risk is lower if they were only children. Finally, the likelihood rises if the first two children were born indifferent partnerships. In the third model, where the period and the individual variables occur together, the direction, scale and significance of the individual variables hardly change, but coefficients of individual periods do so and lose their statistical significance.

In the next stage, *ethnicity and religious beliefs* are added to the model as time-constant variables. The effect of each meets our expectations. The chance of a third child is higher among the Roma and among religious believers. The period variable continues to lack significance. Although in two periods (4 and 5) the coefficients are higher compared with the reference period, the differences are still not significant. At the same time, the p-value in the period of tax relief (January 1999 – December 2005) approaches the level of 0.1 (0.113).

The final step is to include *educational attainment* in the model. As a result, the risk of having a third child is elevated in the two periods mentioned earlier. In the period of generously increased tax relief (January 1999 – December 2005) and in that of universal family benefit (January 2006 – March 2009), the risk of a third birth was one-and-a-half times higher than in the period of state socialism. As a first approach, therefore, it has to be said that introducing tax relief increased the probability of a third child, but lengthening parental leave had no fertility effect.

Figure 1 demonstrates the effect of the policy periods by comparing coefficients of the first model (without controls) with the last model (containing all the control variables). The declining proportions of two groups– those with elementary education and those having a second child relatively early– cause the effect to be reversed if the control variables are omitted. These two groups steadily decrease in size over time, while the chance of a third childbirth among them remains relatively high.

The behaviour of two-child families differs clearly according to educational attainment. Those with medium attainment (vocational training school or secondary school) are least likely to take on a third child. Those with low attainment are more than twice as likely to do so (2.122), and those with high attainment are somewhat more likely (1.317) than those with medium attainment. Thus the chance of a third child in relation to educational attainment forms an *inverted "J" shape*.

It was detailed in the Introduction how the two measures introduced – longer and flat-rate paid leave, and tax relief– had differing effects at various levels of the social structure (*hypotheses 2 and 3*): the former tended to motivate those of lower status and the latter one those of medium or higher status. To test these hypotheses, the interaction

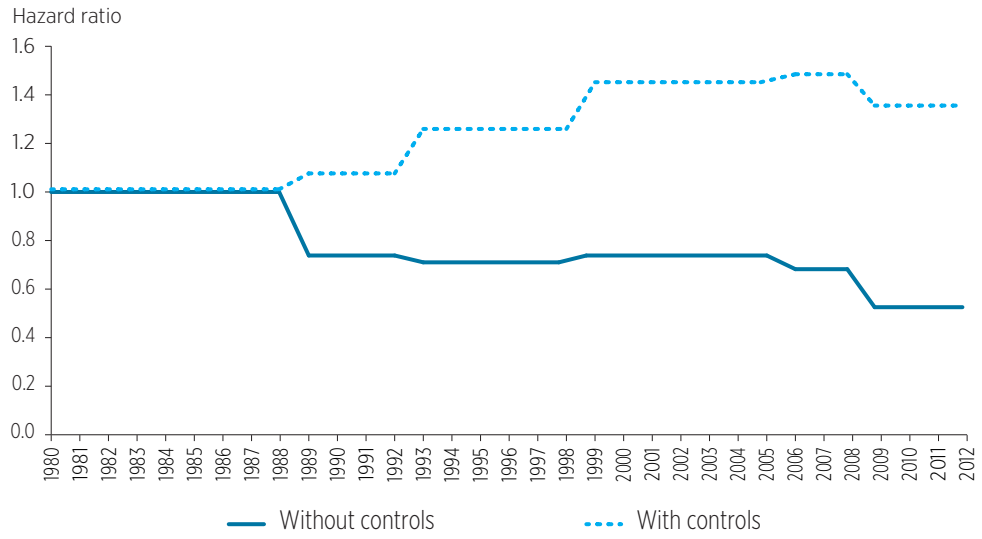
Table 3

Determinants of third births: results of event-history models

	Model 1	Model 2	Model 3	Model 4	Model 5
INDIVIDUAL COVARIATES					
Age of the second child in months					
<15		0.648 ***	0.651 ***	0.631 ***	0.627 ***
15–27		(ref.)			
28–51		0.733 **	0.726 **	0.744 **	0.745 **
52+		0.455 ***	0.440 ***	0.455 ***	0.458 ***
Sex					
Woman		(ref.)			
Man		1.122	1.090	1.115	1.145
Relative age at second birth					
Early		1.520 ***	1.594 ***	1.512 ***	1.590 ***
Average		(ref.)			
Late		0.836 †	0.805 *	0.811 †	0.741 **
Sex of the first and the second child					
Different		(ref.)			
The same		1.333 ***	1.327 **	1.327 **	1.351 ***
Number of siblings the respondent grew up with					
0		0.626 **	0.634 *	0.623 **	0.636 *
1		(ref.)			
2		1.124	1.119	1.035	1.015
3 or more		2.132 ***	2.185 ***	1.564 ***	1.370 **
Union dissolution between the first and the second birth					
No		(ref.)			
Yes		1.798 ***	1.793 ***	1.671 ***	1.679 ***
Ethnicity					
Non-Roma		–	–		
Roma				2.804 ***	1.929 ***
Religiosity (follows the teaching of the church)					
No		–	–	(ref.)	
Yes				1.499 ***	1.471 ***
Highest educational attainment at second birth					
Primary or less		–	–	–	2.122 ***
Vocational training school or secondary					(ref.)
Tertiary					1.317 *
POLICY VARIABLES					
Policy period					
1980–1988	(ref.)	–	(ref.)		
January 1989–January 1993	0.736 †		0.930	0.970	1.078
February 1993–December 1998	0.704 *		1.066	1.108	1.259 (0.138)
January 1999–December 2005	0.728 *		1.254 (0.159)	1.290 (0.113)	1.456 *
January 2006–March 2009	0.677 *		1.250 (0.254)	1.298 (0.183)	1.481 *
April 2009–	0.521 **		1.106	1.148	1.351 (0.202)
Log likelihood		–1890.0	–1887.2	–1843.4	–1823.4
N (person months)		145,825	145,825	145,825	145,825

Notes: hazard ratios; † p<0.1; * p<0.05; ** p<0.01; *** p<0.001; () p>0.1

between period and education were added to the last model in order to clarify main and interaction effects. The interaction effects can be shown several ways, of which we chose to devise one single interaction variable, consisting of 6 x 3 = 18 categories, replacing the separate education and period variables in the last model. This way the effect of educational attainment could be judged for each individual policy period. *Table 4* (and *Figure 2*) presents the coefficients for the values of this variable (coefficients for the other variables in the model are not presented here).



Source: Own calculation, Hungarian GGS.
Note: The model without control variables agrees with Model 1 of *Table 3,* and that with them with Model 5.

Figure 1
Risk of third birth by policy periods, 1980–2012 (regression estimates)

Those with a *medium (vocational and secondary) level of education in the communist period* are taken as the *reference category*. The results show that those with basic schooling in the communist period were clearly more inclined to have a third child. In the early period after the change of system (January 1991–January 1993), those with various levels of educational attainment did not differ from the reference group or from each other. However, in the following periods, the pattern differed among those with primary and with medium levels of education: the risk of a third birth among the lowest educated was noticeably high in three consecutive periods, arriving at three times the risk for the reference group. These coefficients are presented visually in *Figure 2*.

As stated earlier, long-term, flat-rate paid parental leave was introduced in January 1993, so opening a period in which the risk of a third birth at first rose to three times that of the reference category. It is clear to the authors that lengthening the period of benefit from three to eight years (an additional 60 months) after the birth of the third child *increased the fertility of those of low status*. Somewhat discouraging is the fact that the value of the coefficient fell in the last examined period, but it should not be forgotten that the 36 months of paid leave awarded for several decades were reduced to 24 in April 2009. This change marked a curb in family-policy provisions in general and for the group in question as well.

A significant rise in third birth risk came to those with high educational attainment when taxable earnings gained substantial tax relief in 1999. The group's coefficient remained significantly higher than the reference group's throughout (coefficient values in order: 2.106; 2.104; 1.881). From the point of view of our third hypothesis it is important to note how the risk of a third birth among those with medium educational attainment

Table 4

Interaction between policy period and educational attainment

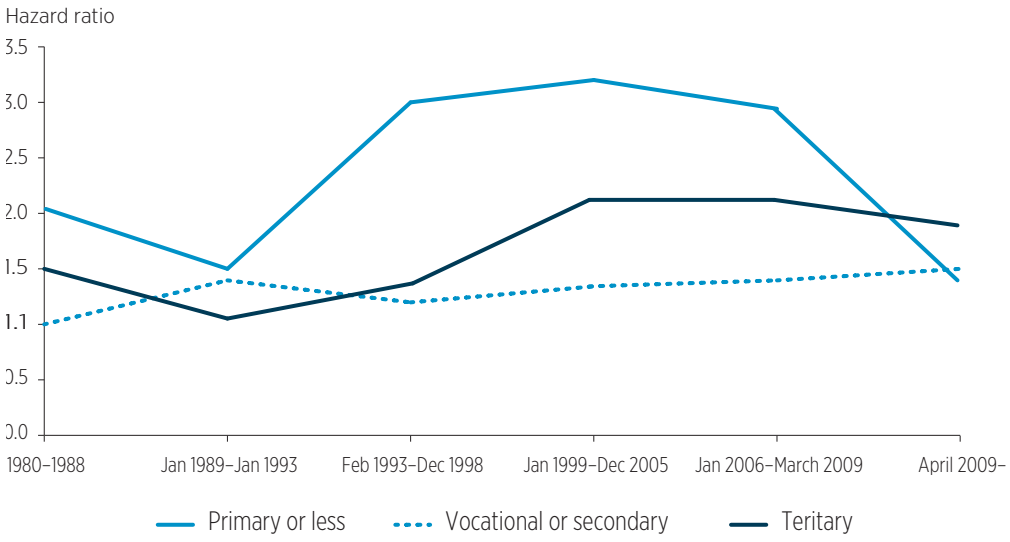
	Hazard ratio	p
Policy period & education interaction		
Before 1989 # primary or less	2.022	0.014
Before 1989 # vocational or secondary	(ref.)	
Before 1989 # tertiary	1.500	0.416
Jan 1989–Jan 1993 # primary or less	1.468	0.283
Jan 1989–Jan 1993 # vocational or secondary	1.395	0.222
Jan 1989–Jan 1993 # tertiary	1.013	0.978
Feb 1993–Dec 1998 # primary or less	3.022	0.000
Feb 1993–Dec 1998 # vocational or secondary	1.191	0.474
Feb 1993–Dec 1998 # tertiary	1.355	0.292
Jan 1999–Dec 2005 # primary or less	3.183	0.000
Jan 1999–Dec 2005 # vocational or secondary	1.310	0.282
Jan 1999–Dec 2005 # tertiary	2.106	0.007
Jan 2006–March 2009 # primary or less	2.937	0.003
Jan 2006–March 2009 # vocational or secondary	1.390	0.267
Jan 2006–March 2009 # tertiary	2.104	0.024
Apr 2009–interview # primary or less	1.396	0.599
Apr 2009–interview # vocational or secondary	1.468	0.247
Apr 2009–interview # tertiary	1.881	0.082

Note: The control variables in the model are sex, relative age, age of second child, sex of 1st and 2nd child, number of siblings, union dissolution, ethnicity, and religiousity.

hardly changed, or not to a significant extent. These findings show that *fertility was obviously affected by the introduction of the tax relief* and that its effect was strongest on those *with high educational attainment*.

The statistical analysis here obviously fails to explain what exact measures and other phenomena lay behind the higher number of third children born to those with low and high educational attainment between January 1999 and December 2005. However, it can be stated clearly that it was higher than for the reference group. Nonetheless, looking at the whole period (*Figure 2*), it would be hard to conclude otherwise on theoretical grounds than that the two above-discussed policy interventions successfully increased third birth risks.

As for the part played by the Hungarian family-benefit system, especially the paid parental-leave system, it is worth recalling one constituent of the multivariate analysis used here: the effect of the time since the birth of the second child. It was found that the chance of a third child being conceived was greatest when the second child was about two years old (15–27 months), i. e. would be born before the second child's third birthday (if all other factors are controlled). Based on the literature, the expectation will be that the chance of a third birth diminishes as the period since the birth of the second increases (see Breton et al 2005). However, it should be remembered that parental leave in Hungary for the second child can be claimed for 36 months: up to then the benefit is paid and the job is held open. According to the results of this analysis, the third child is *timed rationally*: the benefits from the second child are used fully. This accords with the findings of Makay (2017), which shows that the chance of a woman on parental leave to return to work is highest in the 37th month after giving birth, and also with those of Matysiak and Szalma (2014), who examined the timing of the second child in relation



Source: Own calculation, Hungarian GGS.

Figure 2
Interaction effects of policy period with educational attainment (results of the event-history modelling)

to different lengths of childcare leave through a Hungarian–Polish comparison. Finally, Vikat (2004) shows that the introduction of the home-cash-for-care in Finland logically raised the chance of a third child being born around the 36th month. All these findings point to a clear role for the character of the family-benefit system and the length of the paid leave for parental care in the timing of a third child.

SUMMARY CONCLUSIONS

There has been great interest in the measures of population policy in the period of very low fertility, particularly in the post-communist countries of East Central Europe, irrespective of what values or motives lie behind specific political interventions. The interest is great, but the likely results are hard to calculate; there are few studies available for analysing empirically the success of such interventions. Even in a period of relatively calculable economic growth within the welfare-state context of Western Europe, moderation pertains in gauging the effectiveness of family and/or population-policy interventions, although there exist numerous analyses whose findings point to the success of institutional changes and a positive effect on fertility (Hoem, 1993; Hoem, Praskawetz and Neyer, 2001; Lalieve and Zweimüller, 2009).

This analysis on the effectiveness of family-policy intervention was done amidst "unusual" conditions: during the transition from state socialism to a market economy, when the Hungarian population was moving from a model of early to later child-bearing, with a strong process of postponement at work (Sobotka, 2004). Successive changes of government meant that family policy underwent frequent, stop-and-go changes, although the proportion of family-policy benefits to GDP remained high by OECD standards (OECD, 2011).

Earlier it was shown how the population reacted to key family-policy change in the period: the abolition of earnings-related paid parental leave induced a radical fall or postponement of parenthood among the more educated women (Aassve, Billari and Spéder, 2006). However, we also assumed elsewhere that demographic policy measures promise little potential reward amidst turbulent economic transformation and an unstable institutional environment (Spéder, 2016).

Finally, the present empirical analysis leads us to the conclusion that the *population-policy interventions had a clearly positive effect.* This is probably due to three features of the interventions: their sizeable scale, predictability and purposiveness. First, their *scale was indeed great*, with the period of support increasing from three to eight years, at a time when it amounted to a third to half of average female pay, and in a form that did not exclude claimants from the labour market. No less significant was the scale of tax relief, which was around 50 per cent of average female earnings in 2001. Secondly, despite the general stop-and-go feature of family policy in the period, the withdrawals and reductions did not affect the provisions to three-child families. The family-policy conditions for large families remained *relatively predictable* even in the period of general unpredictability. Thirdly, the measures examined were *aimed at* families with three or more children. This group too was affected by the unstable economic conditions, yet postponement as a response was inadequate for them. Childless couples could indeed respond to uncertainties by postponing parenthood, but this was not an attractive alternative for two-child families. The decision to have a third child was hardly about timing but mostly about quantum.

The effect is *differentiated* and fits our initial expectations. Those of lower socio-economic status were probably motivated to have more children by a flat-rate benefit coupled with longer paternal leave. There is reason to assume this may have allowed them to avoid some negative consequences of the economic transition (unemployment). Nor was the behaviour confined to the post-communist transformation. Vikat (2004) draws attention to this possible explanation in relation to Finland. It also coincides with the theoretical assumption that low-income families respond keenly to a set, lump-sum type of benefit. On the other hand the prospect of appreciable tax relief effectively encourages those of high educational attainment to have a third child.

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APPENDICES

Table A1

The changing amount of tax relief (or the decrease of taxable income), according to the number of children in families, Hungary, 1988–2015

Year	Number of children			Remarks
	1 child	2 children	3+ children	
1988–	NO	NO, but 1000 HUF/child for lone parents	1000 HUF/child	decrease taxable income
1991	1300 HUF/child	1300 HUF/child	1300 HUF/child	decrease taxable income
1993	300 HUF	300 HUF/child	400 HUF/child	tax relief
1995–1998	NO	NO	NO	
1999	1700 HUF	1700 HUF/child	2300 HUF/child	tax relief
2000	2200 HUF	2200 HUF/child	3000 HUF/child	tax relief
2001–2004	3000 HUF	4000 HUF/child	10,000 HUF/child	tax relief
2005	3000 HUF	4000 HUF/child	10,000 HUF/child	tax relief, income ceiling
2006–2010	NO	NO	4000 HUF/child	tax relief
2011–	10000 HUF	10,000 HUF/child	330,000 HUF/child	tax relief

Table A2

Main characteristics of the "Turning Points of the Life Course" (Hungarian Generations and Gender Survey) panel survey

Waves	Dates of fieldwork	N	Response rate (%)*	Response rate excluding deaths (%)	Age range
Wave 1	November 2001 – February 2002	16,363	67.9	–	18–75
Wave 2	November 2004 – May 2005	13,540	82.7	85.8	21–78
Wave 3	November 2008 – February 2009	10,641	78.6	83.3	25–82
Wave 4	November 2012 – March 2013	8,103	76.1	82.2	29–86

Note: * compared to the reference population (in Wave 1) or to the previous wave.

Table A3

Reasons for exclusion from the working sample and the number of cases

Reasons for exclusion	n
Respondent was born before January 1960	7,418
Invalid partnership history (negative union length, marriage before the start or after the end of the union, no end date but there is a next relationship, union without start date but with end date, start of a union precedes the end of the previous one)	584
Respondents who have had only one or no biological children	2,945
Women older than 35 and men older than 38 at second birth	64
Not living in a co-residential union at second birth or no later than 12 months after the second birth	82
Having the second child before January 1980	5
Having twins at first or second birth	83
Invalid birth history (9 months or less difference between births, respondent was younger than 15 at any of the births)	24
The first or the second child was adopted	6
Total number of excluded cases	11,211
Final n of the working sample	2,329

Note: the initial working sample included those respondents who participated in Wave 1 and 2 of the survey.

Table A4

Descriptive statistics of the independent variables (%)

	Total	Women	Men
INDIVIDUAL COVARIATES			
Age of the second child in months (duration; dynamic)			
<15	21.6	21.2	22.1
15–27	17.9	17.9	17.8
28–51	27.6	27.6	27.6
52+	32.9	33.2	32.5
Sex of the respondent (time-constant)			
Woman	60.4	–	–
Man	39.6	–	–
Relative age at second birth (time-constant)			
Early	20.9	17.7	25.7
Average	50.3	53.1	46.0
Late	28.8	29.2	28.3
Sex of the first and the second child (time-constant)			
Different	49.3	48.8	50.1
The same	50.7	51.2	49.9
Number of siblings the respondent grew up with (time-constant)			
0	10.7	11.2	10.0
1	48.9	48.7	49.3
2	21.5	21.6	21.3
3 or more	18.9	18.6	19.4
Union dissolution between the first and the second birth (time-constant)			
No	92.4	91.9	93.1
Yes	7.6	8.1	6.9
Ethnicity (time-constant)			
Non-Roma	91.5	92.3	90.5
Roma	8.5	7.7	9.5
Religiosity (follows the teaching of the church; time-constant)			
No	86.7	85.3	88.7
Yes	13.3	14.7	11.3
Highest educational attainment at second birth (time-constant)			
Primary or less	16.5	17.1	15.6
Vocational training school or secondary	62.5	59.6	66.9
Tertiary	21.0	23.3	17.5
POLICY VARIABLES			
Policy period (dynamic)			
1980–1988	7.2	9.5	3.5
January 1989–January 1993	9.3	10.6	7.3
February 1993–December 1998	35.0	35.6	34.2
January 1999–December 2005	30.2	27.4	34.6
January 2006–March 2009	10.5	9.5	12.0
April 2009–	7.8	7.4	8.4

Notes: % for time-constant covariates, % distribution of exposure time for dynamic covariates

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