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Tempo Effects in the Fertility Decline in Eastern Europe: Evidence from Bulgaria, the Czech Republic, Hungary, Poland, and Russia

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TEMPO EFFECTS IN THE FERTILITY DECLINE IN EASTERN EUROPE:

EVIDENCE FROM BULGARIA, THE CZECH REPUBLIC, HUNGARY, POLAND, AND RUSSIA

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1. Introduction

The enormous social, economic, and political changes in Central and Eastern Europe after 1989 were associated with an abrupt fall in fertility. In 1997 the total fertility rate (TFR) reached its lowest level so far in many countries, e.g., 1.09 in Bulgaria, 1.17 in the Czech Republic, 1.24 in Estonia, 1.38 in Hungary, 1.11 in Latvia, 1.32 in Romania, 1.28 (in 1996) in Russia (Council of Europe, 1998). This fall in fertility has meant a rapid population decrease in recent years. For instance, in 1996 the annual population growth rate was -0.53% in Bulgaria, -0.37% in Hungary, and -0.32% in Russia (the respective rates of natural increase were -0.54%, -0.37%, -0.55%; Council of Europe, 1998). This observed decline in population frequently prompted a pessimistic debate in the popular media, and sometimes also in the scientific literature. The observed demographic trends were associated with phrases such as "demographic crisis", "breakdown", "catastrophe".[1]

Two arguments serve as the background for such extreme views. One is connected to changes in reproductive behaviour, the other to the measurement of fertility. According to the first argument, the abrupt decline in the number of marriages and births is a direct consequence of the social and economic difficulties experienced by the majority of the population during the transition towards a more democratic society. Adherents of this view would state that fertility would not have fallen to its present low levels if there had been no difficulties (i.e. transition) in recent years. The second argument rests either on the overly simplistic observation of the falling absolute number of births, or on the widely-used interpretation of the total fertility rate (TFR) as the average number of children per woman. Since the observed TFR is considerably below two, the view emerges that the average number of children in a family is considerably below two and hence the population will substantially decline in the foreseeable future.

Demographers usually do not share such extreme views. Terms such as "crisis" are either questioned (e.g., Zakharov, 1997) or placed in quotation marks (DaVanzo, ed., 1996). The two arguments presented above have been discussed in the demographic literature as follows.

Consider first the argument about reproductive behaviour. There has been considerable debate about the specific interrelation of fertility and the socioeconomic changes in Central and Eastern Europe. For instance, are the recent demographic changes in this part of Europe the result of a diffusion of Western European behaviour, or do they represent a reaction on the part of individuals to the difficulties associated with the transition towards a market economy and a democratic society? (e.g., Kohler and Kohler 1999; Zakharov 1997; Zakharov and Ivanova, 1996). There seems to be a preference for assessing these recent demographic changes as a change towards patterns observed earlier in Western, Northern, and more recently, Southern Europe. It is argued that postponement of childbearing and a lower level of fertility, a decrease in marriage rates, the increased incidence of divorces, the spread of voluntary childlessness, and an increase in non-marital unions and extra-marital births are typical behavioural changes associated with the "second demographic transition" (Lesthaeghe and Van de Kaa, 1987; Van de Kaa, 1987). Low fertility levels may then be the result of such a transformation of demographic behaviour towards more "modern", or more "western", patterns. Instead of representing a crisis, current fertility patterns in Central and Eastern Europe are therefore a more-or-less expected change associated with the transformation in Eastern Europe. The latter has merely speeded up the expected demographic change.

The “measurement argument”, i.e. the common interpretation of the public of the TFR as the total number of children per woman, can be called into question by considering the well-known demographic fact that a postponement of births not only leads to a rising mean age at birth, but it also affects the observed fertility rates and thus the TFR. In particular, a postponement of fertility leads to a decrease in the observed age-specific and total fertility rates because events are spread out over a longer period of time. Since the decline in the TFR in Central and Eastern Europe has often been associated with a rapid increase in the mean age at birth (by birth order), the decline in fertility could merely be a reflection of the postponement of births, which does not affect the quantum of fertility. This reasoning then suggests that the total fertility rate will increase when the rate of postponement of births slows down. Therefore, both the observed number of births and the observed TFR represent distorted measures of the level of fertility, as far as long-term population growth is concerned. (See, for example, Zakharov 1997; Zakharov and Ivanova, 1996 for more details on both arguments.)

This paper makes use of a recently proposed *adjusted total fertility rate* (Bongaarts and Feeney, 1998) that accounts for the impact of changes in the timing of births on the TFR. Although the general relation between timing effects and observed rates has been known for at least half a century and adjusted fertility measures were available (Ryder, 1956), a quantitative evaluation of this effect has usually been missing in discussions of fertility trends. One of the reasons for this is that the application of completed cohort fertility can only be used for cohorts with (almost) completed fertility . Thus, Ryder’s approach is not applicable to a study of the timing of fertility in the recent decade.

The Bongaarts-Feeney formula provides a simple and relatively intuitive measure for overcoming this limitation. In particular, the adjusted total fertility rate equals the total fertility rate that would have been observed if there had been no postponement of births within a given year. The adjusted TFR therefore makes it easier for demographers to evaluate the effect of changes in the timing of births on the period fertility level. Moreover, it augments analyses based on the standard total fertility rate with a new measure, and hence allows for a new study and interpretation of aggregate fertility trends.

In this paper we apply the Bongaarts-Feeney formula to five countries from Central and Eastern Europe (Bulgaria, the Czech Republic, Hungary, Poland, and Russia). The results reveal that the adjusted total fertility rate declines substantially less than the observed TFR in all of these countries with the exception of Russia. In some instances, the adjusted TFR even increases after 1990. This finding re-emphasises the relevance of tempo effects for understanding the recent fertility declines in Central and Eastern Europe.

The implications of these tempo effects would have been much more difficult to detect and evaluate without the adjusted total fertility rate suggested by Bongaarts and Feeney (1998). Nevertheless, considerable caution is necessary when interpreting the results of this adjustment. In particular, the Bongaarts-Feeney formula is derived under assumptions about a specific postponement of fertility that changes the mean age at birth (by order) without affecting other aspects of the fertility schedule. Thus, the TFR adjustment assumes the absence of cohort effects. As pointed out by Bongaarts and Feeney, this assumption may be violated in periods of rapid fertility change or in periods of social turmoil. Kohler and Philipov (1999) have shown formally that increases in the variance of the fertility schedule, which constitute one particular violation of the assumptions underlying the TFR adjustment, lead to a downward bias in both the estimated tempo effect and in the adjusted TFR. In this paper we therefore test the assumption of “no cohort effects” by comparing the shape of the fertility schedule in consecutive years by means of survival analyses.

2. Tempo and Quantum in Fertility

The total fertility rate (TFR) is the most widely used indicator of the level of fertility today, and its interpretation as the expected number of children per woman that survives through the reproductive years and experiences the fertility rates observed in the current period is well-known.

By considering births separately by birth order, the total fertility rate can be expressed as

$$TFR_t = TFR_t(1) + TFR_t(2) + TFR_t(3) + \ldots \tag{1}$$

where $TFR_t(i)$ is the i-th order TFR in period *t*. In the following discussion we will omit the subscript for the time period *t* since all rates in equation (1) refer to one and the same period.

TFR(i) has an interpretation similar to that of the TFR, but referring specifically to the i-th birth. Although from a cohort perspective every woman can have only one first birth, the TFR(1) can nevertheless be greater than one. In particular, this can be observed in periods when the mean age at first birth is decreasing. Thus a tendency towards a younger age at birth brings about an increase in the TFR. Inversely, a deferral of births leads to a decrease in the observed TFR. The change in the period fertility level that is due to changes in the timing of births is known as the *tempo effect*.

From the cohort perspective, changes merely in the timing of births have no effect on the level of fertility. Completed fertility remains unchanged, because it depends only on the number of live births per woman, not on the age at which these births occur. The level of completed fertility changes only when the number of children born to a woman over her entire reproductive period changes – an effect that is known as the *quantum* of fertility.[2]

The separation of tempo and quantum effects is especially relevant for understanding the fertility patterns in Central and Eastern Europe, as the mean age at birth has been rapidly increasing and fertility has approached unprecedented low levels. If tempo effects are the primary reason for the recent decline, rather than quantum effects, then the fertility behaviour is characterised by a postponement of births, rather than a permanent reduction of births. If this is indeed the case, a modest recovery to somewhat higher fertility levels is more likely than the long-term persistence of these historically low fertility levels.

3. The Bongaarts-Feeney formula

The problem with the distinction between tempo and quantum effects is how to control for the tempo effect when using period data. “It is possible to make progress by assuming that fertility changes in a structured way. Specifically, we will assume that fertility may be influenced by period, age, parity, and duration since last birth, but not by cohort.” (Bongaarts and Feeney, 1998, p.275). Hence the basic assumption is that during the period under study there were no cohort- or age-differentials in fertility change. This implies that the fertility schedule can move to the right or left (due to tempo effects) or up or down (due to quantum effects). The shape of the fertility schedule, however, must remain unaffected by the changes in both the tempo and quantum of fertility. The assumption need only be valid for the period in which the fertility is adjusted, which is one year. However, since we usually observe only annual data, the assumption for the practical implementation of the formula requires that the shape of the age- and order-specific fertility schedule observed in year *t* be identical with that observed during the next year *t+1*.

On the basis of these assumptions, Bongaarts and Feeney derive an adjusted total fertility rate that equals the TFR that would have been observed during a period if there had been no change in the timing of births (for an alternative derivation and extension of the formula, see Kohler and Philipov 1999). The adjusted total fertility rate is calculated as

$$\text{(2)} \qquad \mathrm{TFR(adjusted)} = \frac{\mathrm{TFR(1)}}{1-r_1} + \frac{\mathrm{TFR(2)}}{1-r_2} + \frac{\mathrm{TFR(3)}}{1-r_3} + \ldots$$

where TFR(adjusted) is the TFR without tempo effect, and r_i is the change in the mean age at birth at the i-th order (all TFRs and r_i's refer to the same period of time t).

The terms on the right-hand side of (2) are the adjusted order-specific TFRs. For example, $TFR_{adj}(1) = TFR_{obs}(1)/(1-r_i)$. The change r_i in the mean age at birth-order i is estimated by

$$\text{(3)} \qquad r_i(t) = (a_i(t+1) - a_i(t-1))/2$$

where a_i is the mean age of the births of order i. Note that the adjusted TFR for the first and the last year in a series cannot be computed due to the way $r_i(t)$ is calculated in (3).

4. Application of the Formula

In this section we apply the Bongaarts-Feeney formula to Bulgaria, the Czech Republic, Hungary, Poland, and Russia. All these countries have experienced a substantial decline in fertility since 1989. Although all these countries were in the process of transformation towards a market economy and a democracy, their specific socioeconomic and demographic contexts differed. In the following analyses we therefore present for each country the observed and adjusted total fertility rate, separately by birth order as well as combined. We also comment on recent fertility patterns in these countries, using the new insights that are obtained from the adjustment of the total fertility rates.

The observed TFRs and mean ages are illustrated in figures 1 through 5. In each figure the top-left graph gives the trends for all births combined and the remaining three graphs report the trends for birth orders 1 to 3. In each of the graphs the thin solid line represents the observed TFR, the bold full line the adjusted TFR (for both TFRs the scale is on the left-hand side), and the dotted line gives the mean age at birth (scale is on the right). The scaling interval is the same for all indicators, which allows for a direct visual comparison of changes among the five countries. The numeric results that correspond to the figures below are given in Tables A1-A5 in the Appendix.[3]

(a) Bulgaria

As already mentioned in the introduction, Bulgaria has recently become one of the countries with the lowest total fertility rates, not only in Central and Eastern Europe, but also worldwide. Since these most recent fertility changes are particularly striking, we first concentrate on the period 1988–1997. The next section gives a discussion on the population policy effects in a longer period.

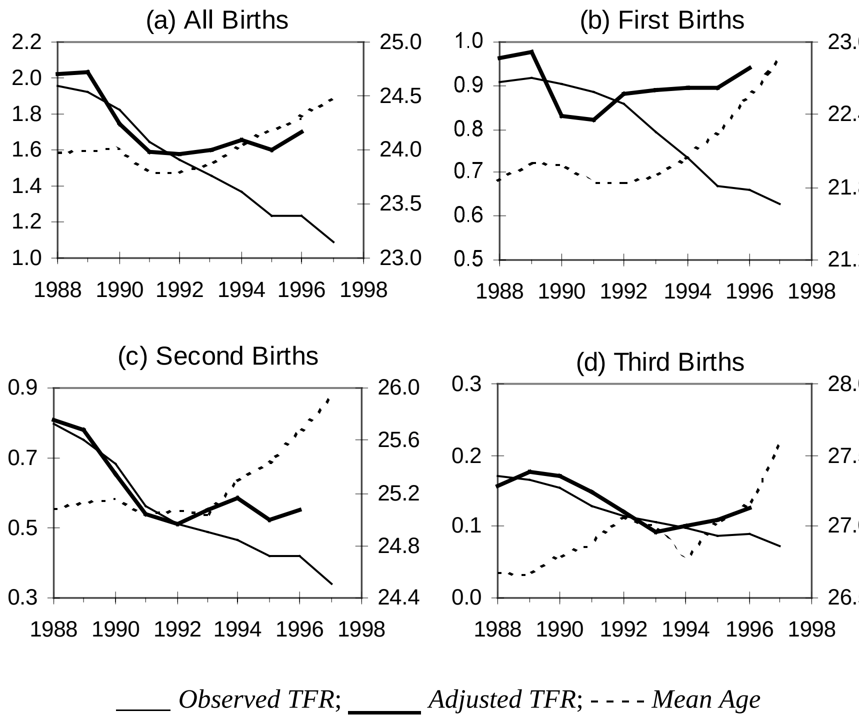


Figure 1 : ***Observed and adjusted TFRs and mean ages of childbearing for (a) all births and (b-d) order-specific births, Bulgaria 1988-1997.*** *(TFRs are plotted on the left axis; mean ages are plotted on the right axis)*

Figure 1 clearly reflects the precipitous fall in fertility after 1990, and it also reveals that the decline affected all birth orders to a similar extent. Somewhat surprising is the fact that the order-specific mean ages for first and second births started to increase mainly after 1993, although the decline in fertility occurred already in 1990. Until about 1993 the changes in the mean ages are relatively small, despite a substantial drop in fertility.

This finding suggests that the fertility decline was dominated by quantum effects in the years prior to 1993, and that tempo effects emerged as an important factor only afterwards. The adjusted total fertility rate confirms and quantifies this interpretation. The adjusted order-specific TFRs given in figure 1 reveal a fertility trend that differs – sometimes substantially – from the respective observed total fertility rate. While after 1992–1993 the observed TFR continues to decline, the adjusted TFR stagnates or even increases modestly. The largest difference between observed and adjusted TFR occurs for the first order in 1996. In that year the adjusted TFR(1) exceeds the level of 0.9, indicating that some 90% of all women would have had a first child if the tempo effect had been absent. This high incidence of first births was typical for Bulgaria prior to 1989, and the adjusted TFR(1) returned to such high levels despite the fact that the observed TFR(1) declined to a level just over 0.6.

The finding that the adjusted TFR exceeds the observed total fertility rate also holds when births of all orders are considered. While the observed TFR for all births in the top-left graph exhibits a continuous decline to a level below 1.2 during the period 1989–1997, its adjusted counterpart stops decreasing at around 1991 and remains more or less constant at approximately 1.6 to 1.7. Thus, according to the Bongaarts-Feeney formula, the decline in fertility after 1992 seems to be due almost exclusively to tempo effects.

Our application of the TFR adjustment to Bulgaria points to a change in the pattern of fertility decline during the 1990s: at the beginning of this decade Bulgarian fertility declined predominantly due to quantum effects, while in the middle of the 1990s it decreased mainly because of changes in the timing of fertility.

(b) Czech Republic

The observed and adjusted total fertility rates for the Czech Republic are depicted in figure 2. As in the case of Bulgaria, the transformation after 1990 led to a substantial decline in fertility. Nevertheless, some significant differences exist between the two countries. For instance, unlike in Bulgaria, the mean ages did not decrease at all, and they started rising somewhat earlier.

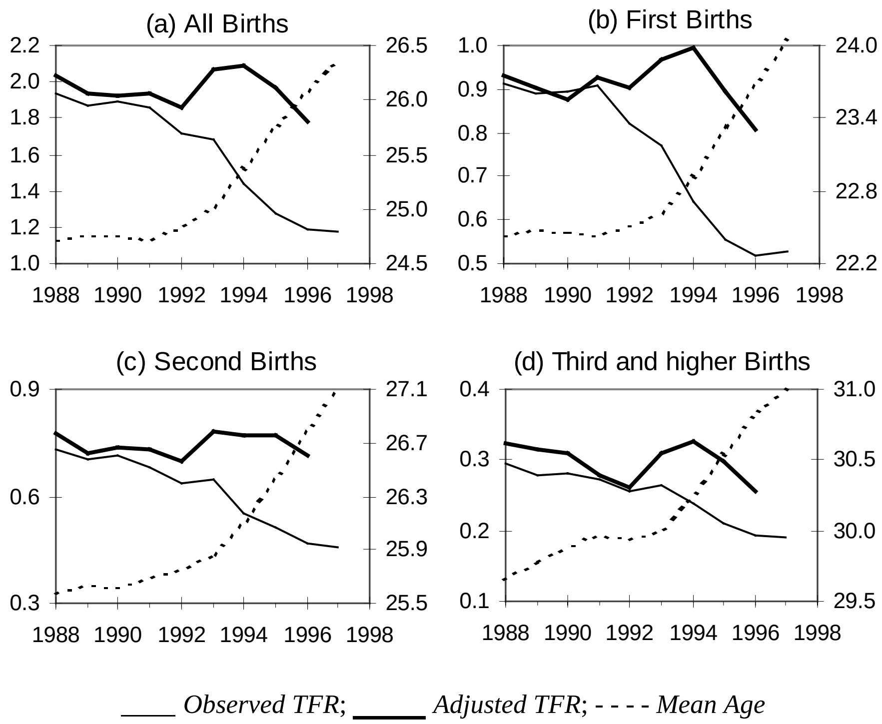


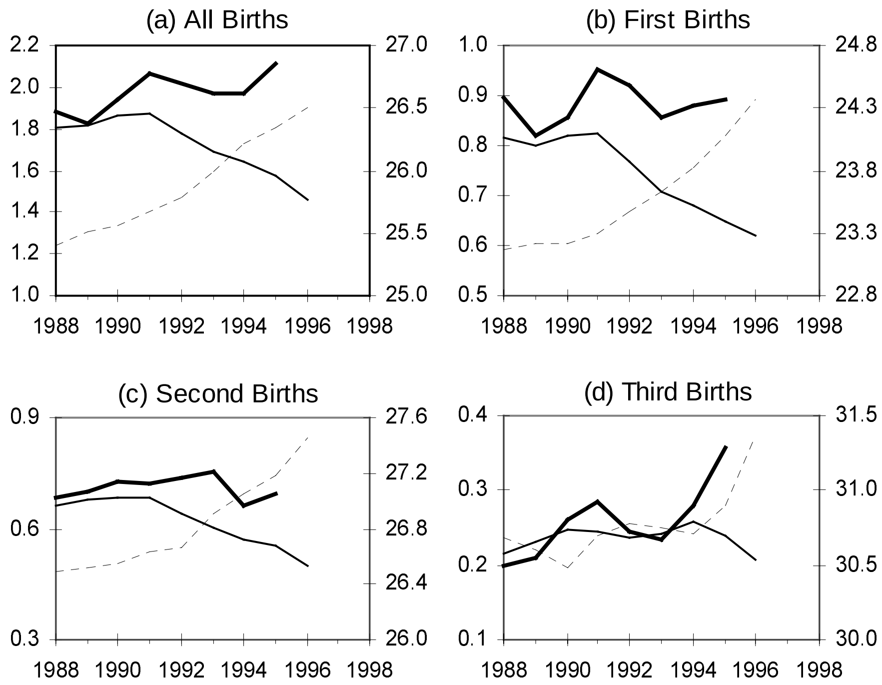
Figure 2: ***Observed and adjusted TFRs and mean ages for (a) all births and (b-d) order-specific births, Czech Republic 1988-1997.*** *(TFRs are plotted on the left axis; mean ages are plotted on the right axis)*

Panels (b)–(d) in figure 2 reveal that the postponement of births was initiated almost concurrently with the decline in the TFR: the trends in the order-specific TFRs are virtually a mirror image of the trends in the mean age at birth of the corresponding order. This inverse relation between the mean age at birth and the fertility level also persists when all births are considered (panel (a)): in periods when the mean age was constant, so was the TFR; in periods when the mean age increased, the TFR decreased.

A comparison between the adjusted and the observed total fertility rate reveals that the fertility fall during the early 1990s was primarily due to a tempo effect, while the quantum effect was rather small or absent. Unlike in Bulgaria, between 1993 and 1995 the adjusted TFR in the Czech Republic recovered to its level of before 1989. In 1993–1994 it even exceeded a level of 2.00, as is shown by panel (a) in figure 2. After 1994 strong quantum effects – primarily for first and, to a lesser extent, for higher-order births – reduce the adjusted total fertility rates, which drop below the 1989 level; nevertheless, the adjusted TFR remains substantially above the observed TFR.

(c) Hungary

Figure 3 presents the observed and adjusted total fertility rates for Hungary.



____ *Observed TFR*; ______ *Adjusted TFR*; - - - - *Mean Age*

Figure 3: *Observed and adjusted TFRs and mean ages for (a) all births and (b-d) order-specific births, Hungary 1988-1996.* *(TFRs are plotted on the left axis; mean ages are plotted on the right axis)*

Observed fertility in Hungary reflects a pattern that is similar to those of Bulgaria and the Czech Republic: the 1990's are characterised by a substantial decline in fertility, although this decline is somewhat more modest in Hungary than in the other two countries. When the tempo effects are taken into account, the picture becomes more diverse. In particular, the adjusted TFR for all births reflects a slight increase throughout the period, indicating that the decline in fertility is accounted for almost entirely by tempo effects. This trend corresponds to a rise in the mean age at birth throughout the period 1988–1996.

The order-specific adjusted TFRs in figure 3 show a less regular pattern and exhibit substantial fluctuations. For instance, the post-1989 increase in the adjusted TFR for first and second births is temporary. The trend reversed itself in the mid-1990s, leading to a decline and later stagnation of the order-specific adjusted TFRs (panels a and b). In contrast to the pattern at lower orders, the adjusted TFR for third births rose substantially during 1994–1995 (panel d). Moreover, the mean age for third births also exhibits the highest annual change that we encountered in our analyses for this paper: in 1996 the mean age at third birth increased by 0.5 years (table A3), which is striking compared to the more commonly encountered annual increases in mean age of 0.1 to 0.2 years.

(d) Poland

Poland is no exception from the above-mentioned three countries, in that the adjusted TFRs are substantially higher than the observed ones (figure 4).

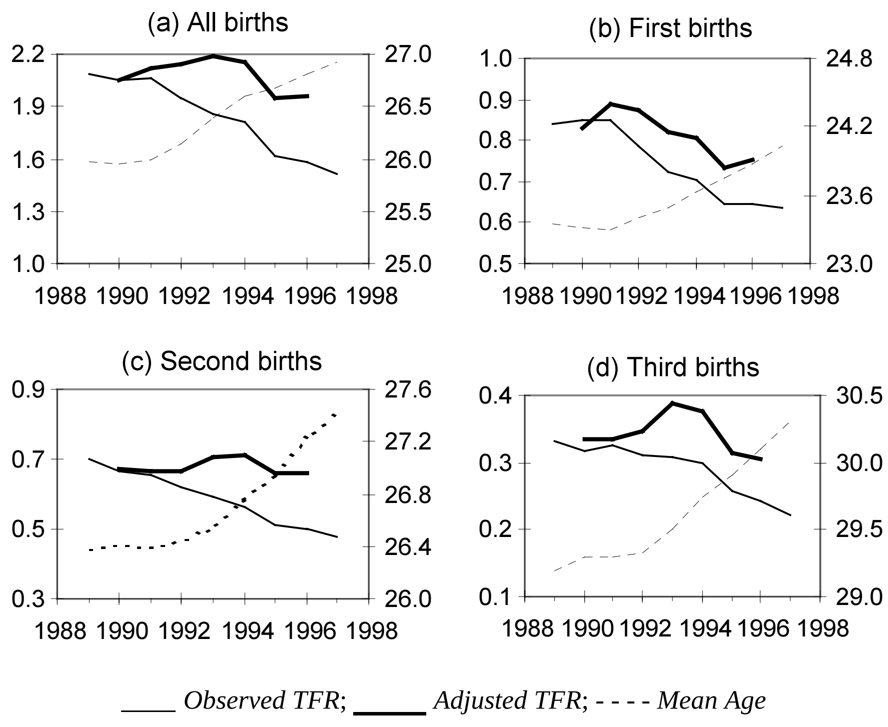


Figure 4: ***Observed and adjusted TFRs and mean ages for (a) all births and (b-d) order-specific births, Poland 1988-1997****. (TFRs are plotted on the left axis; mean ages are plotted on the right axis)*

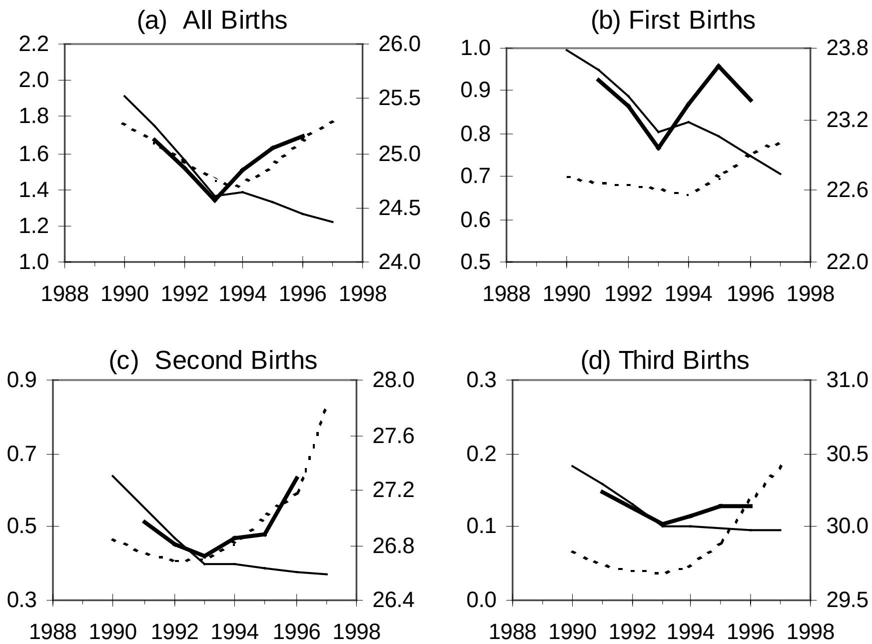
This country has traditionally had somewhat higher fertility than other countries in Eastern Europe. Hence it is no surprise that the overall fertility level adjusted for the tempo effect would be as high as 2.2 at the beginning of the 1990’s, as can be seen in panel (a) in figure 4. However, the quantum effect became stronger during the last couple of years of the period of study, and the adjusted TFR dropped to 1.9 in just one to two years.

The trends by order of birth differ from those observed in the previous three countries studied here. First births decline even after adjustment for tempo (panel (b) in figure 4). Traditionally, the proportion of women who ever have a first birth is slightly lower in Poland than in other countries in the region, being between 0.8 and 0.9. Thus the decrease in the adjusted TFR(1) may mark an important change in this aspect of fertility behaviour.

The trend in second births reveals hardly any change when the adjusted TFR is considered, except for a very slight rise in 1993 and 1994. This rise is far more significant for third births.

(e) Russia

Since Bulgaria, Hungary, the Czech Republic, Poland, and Russia were all experiencing a transition towards a market economy and a democracy, one might expect that the adjusted total fertility rates would be similar. However, figure 5 does not confirm this.



____ *Observed TFR*; ______ *Adjusted TFR*; - - - - *Mean Age*

Figure 5: *Observed and adjusted TFRs and mean ages for (a) all births and (b-d) for order-specific births, Russia, 1990-1997.* *(TFRs are plotted on the left axis; mean ages are plotted on the right axis)*

The pattern of fertility decline during the 1990s diverges surprisingly after tempo effects are included in the analyses. Contrary to the pattern we observed in the Czech Republic, Hungary, and Poland, there is little indication of a postponement of births between 1990 and 1993 in Russia. Up until 1993–1994, the mean age at first or second birth even decreased, and only afterwards did a trend towards a delay in childbearing emerge. The following years, however, exhibit an effective increase in the mean ages, and the impact of the tempo effect on the fertility level seems to have increased in importance. Thus, fertility in Russia changed primarily due to quantum effects up until 1993 and primarily due to tempo effects after that. The same pattern of change was observed to a lesser extent in Bulgaria.

In general, the differences between the observed and the time-adjusted TFR by birth orders are small in Russia up until 1993, but since then they have diverged due to the sudden start of postponement of births. Especially notable is the rise in the second-order adjusted TFR. Hence it is above all second births that are postponed.

5. Usage of the adjusted TFR for the assessment of population policy effects

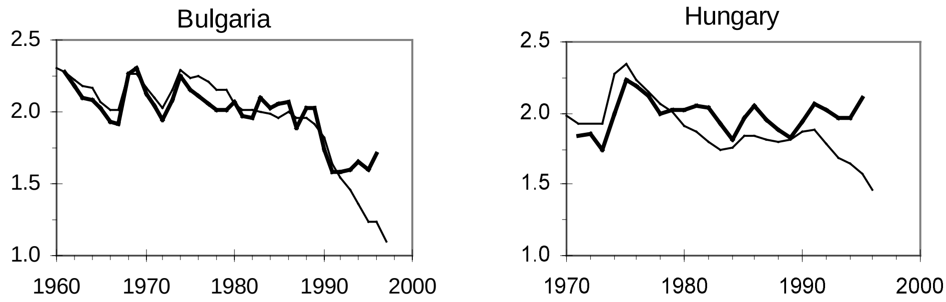
The longer data series for Bulgaria (1960–1997) and Hungary (1970–1997) allow us to study the effect of pro-natal policies during the socialist period on the timing of fertility.

Pro-natal population policies were implemented in most Eastern European countries starting at the end of the 1960s and mainly during the 1970s. Klinger (1991) gives a concise description of the most popular policy instruments. In Bulgaria this policy was introduced first in 1967, predominantly through repressive measures, the most effective being the strict restriction on abortions. In 1973 the general policy was enhanced by positive measures, such as generous maternity leaves and job security for mothers, child allowances, etc. In Hungary, a pro-natal population policy was first adopted in 1973 with instruments and measures similar to those included in the Bulgarian policy of the same year. The restriction on abortions was also introduced but it was not as rigorous. Generally these population policies are order-specific, with stronger incentives for second- and particularly for third-order births.

The effect of the pro-natal population policies has been a topic of numerous discussions. Many authors assess them as successful, as far as inferences could be made from the rise in the TFR. It is well-known though that this rise was often temporary, followed by a decrease in the TFR, which is known as the compensation effect. Thus cohort fertility changed only slightly. Moreover, the mean age of childbearing usually dropped during the first few years after the adoption of the policies, either due to the restrictive measures or to the wish of couples to benefit from the advantages offered.

One may therefore conclude that the effect of the pro-natal population policies could be just in the timing of births. In order to carry out more rigorous analysis some authors applied age-period-cohort (APC) methods, following up on the work of Willekens and Baydar (1984). Büttner and Lutz (1990), for example, studied the effect of the pro-natal policy in the GDR. Philipov (1993) looked at Bulgaria and Stloukal (1998) at the Czech Republic. In each case the policy was found to be effective, contributing to a steady rise in fertility of approximately 10 to 15 per cent.

The APC model used by the above authors does not explicitly distinguish between tempo and quantum effects (see Bongaarts and Feeney, 1998 for a general discussion). The adjusted TFR successfully makes this distinction and hence can contribute to the analysis of the effects of population policies. The adjusted and observed TFRs are illustrated in figure 6 below for Bulgaria and Hungary.



____ *Observed TFR*; ______ *Adjusted TFR*

Figure 6: *Observed and adjusted fertility for Bulgaria and Hungary*

The 1967 pro-natal policy in Bulgaria, which consisted mainly of anti-abortion measures, did not cause substantial timing effects. The values of both adjusted and observed TFRs show a sudden increase for just a couple of years. Data not shown here reveal that the same heap was exhibited clearly for births of order two and higher, and not as much for first births. Hence the restrictions on abortions brought about a significant increase in unwanted births that would perhaps otherwise not have occurred at all. This explains the lack of timing effects. People became accustomed to the restrictions and found ways to control their fertility within few years.

The incentive-based 1973 policy package induced earlier fertility. People who wished to have children decided to have them at an earlier age in order to benefit from the provisions of the policy. This timing effect made the observed TFR stay for several years at a higher level that was taken as a proof of the effectiveness of the policy. As a result, the adjusted TFR shows a very short-term quantum effect and then quickly very nearly returns to the pre-policy level. (The order-specific data, not shown here, indicate that the temporary quantum growth was due to the increase in second-order births, while the third-order births marked no increase. Hence the policy that aimed specifically at increasing the number of third-order births did not achieve its goal. Although this fact was already known through the observed third-order TFR, it becomes more evident after the timing effect is removed.)

In Hungary the policy adopted in 1973 caused a significant increase in observed fertility. Its adjusted counterpart reveals that this rise was due primarily to quantum effects. Tempo effects were present but they were minor. (Data not shown here reveal that the quantum change was due to a rise in quantum of second- and third-order births, while the rise in quantum of first births appeared some years later.)

These policies were subsequently amended in both countries several times to introduce new measures. Increases and decreases in observed fertility were moderate, and an overall near-replacement level was achieved. It is difficult to judge to what extent this was the result of the population policies, and/or of realised fertility behaviour uninfluenced by this policy. One could perhaps conclude from figure 6 that the policy was actually effective to a certain extent in Hungary in the long run, because the adjusted fertility was higher than observed fertility during the 1980s.

6. Testing the assumptions of the formula

As we discussed earlier, the derivation of the Bongaarts-Feeney formula relies on the validity of certain assumptions, most importantly on the assumption that there are no cohort-effects (or age-period interactions). Only if these assumptions are empirically justified can we interpret the adjusted TFR as the total fertility rate that would have been observed in the absence of tempo effects.

In order to verify the assumptions of the adjustment formula (2) we need to check whether the shape of the observed order-specific fertility schedules remain unchanged during adjacent years. Since variations in the mean age and in the level of fertility may influence the TFR adjustment, a test of the “equal-shape” assumption needs to check the significance of these influences.

Bongaarts and Feeney (1998) did not test this assumption directly. Instead, they verified the adjusted TFR by comparing the completed fertility of cohorts to the weighted average of the adjusted TFRs during the childbearing years of this cohort. Bongaarts-Feeney find a relatively close correspondence between these two measures and view this as a validation of the TFR adjustment.

In this paper we pursue a different test, which focuses more directly on the “no cohort-effects” assumption, or equivalently, on the assumption that the shape of the fertility schedule does not change over time. In particular, we test this assumption by comparing two pairs of schedules: those for the years t-1 and t, and those for the years t and t+1. Because r_i is estimated in equation (3) on the basis of the mean age in the years t-1 and t+1, our approach therefore tests whether the fertility schedule in the year of interest t has the same shape as the fertility schedule in the immediately preceding and following years. For example, consider first-order fertility in the period 1994–1996 for the Czech Republic. This period is chosen because it exhibits the largest difference between the observed and the adjusted TFRs. The observed schedules are depicted in the left-hand panel in figure 7. These schedules clearly reflect the trend towards a higher mean age at first birth and the decline in fertility from 1994 to 1996.

In order to remove the shifts in the mean age and in the change in the fertility level, these schedules were standardised so that they exhibit the same mean age (equal to the 1995 level) and an equal fertility level (which was chosen as 1). These standardised fertility schedules are shown in the right-hand panel. This standardisation introduces some changes in the shape of the schedule, as can be seen if one compares the peaks of the corresponding schedules in the two panels in figure 7. This is due to the fact that we have discrete data and that we assume linearity between adjacent age groups.

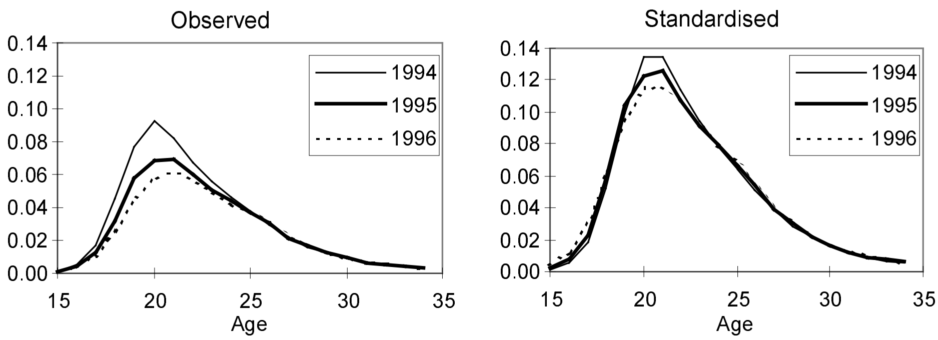


Figure 7: ***Observed and normalised (equal means and areas) first-order fertility schedules, Czech Republic, 1994–1996***

The assumptions underlying the Bongaarts-Feeney formula would hold exactly if the fertility schedules in the right-hand panel were equal. It is of course unrealistic to expect absolute equality, and minor variations in the fertility schedule can occur solely because of the randomness of the birth process. To evaluate the differences in the shape of the fertility schedules we therefore use two statistical procedures: a descriptive one and survival analysis.[4]

The "no cohort-effect" assumption can be relaxed by the incorporation of more general tempo effects that allow for age-period interactions as part of the fertility change. Kohler and Philipov (1999), for example, allow not only for a changing mean age of the fertility schedule, but also for a changing variance of the schedule. Hence, it is not just the case that births can be postponed. In addition, fertility behaviour can become more diverse in the sense that fertility becomes distributed across a wider age range. The application of these "variance effects" is beyond the scope of the present paper, and below we restrict ourselves to a test of the "no cohort-effect" assumption.

The sum of the absolute values of the differences between the standardised age- and order-specific fertility schedules for 1995 and 1994 in figure 7 is 0.060. The same sum of differences is also obtained for the 1995 and 1996 fertility schedules. When compared to the total area underneath the fertility schedule, the differences are therefore only 6%. Moreover, these 6% differences are dispersed across different years, which creates the visual impression that the later schedule has become wider, i.e. its variance has increased. (The standard deviations are 3.96 in 1994, 4.04 in 1995, and 4.11 in 1996.) The largest single-age difference (0.01) was observed at age 20, with the corresponding rates for the three years being 0.134, 0.123, and 0.115. From a demographic perspective, the difference between any of the two pairs of fertility schedules is small and can probably be neglected.

Survival analysis provides a more rigid statistical procedure for the justification of the equality of schedules. The events underlying each fertility schedule are births of a specific order occurring at a specific age. The application of survival methods requires that data on events be counts rather than rates. Consider an initial cohort at the age when childbearing begins. It decreases through age with the corresponding number of births until all members of the cohort disappear (or "die", in the language of survival analysis). The pattern of decrease should correspond to the normalised curves. That is, the initial cohort should be multiplied by the age-specific normalised rates in order to arrive at the number of births distributed by age.

The size of the cohort depends on the level of accuracy we wish to achieve. For example, if we wish to compare the curves up to the 4th decimal point of their rates, we should choose an initial cohort of 10,000. If we wish to compare the curves up to the third decimal point only, then an initial cohort of 1,000 is sufficient. The size of the cohort influences the results of the statistical test in that the larger the size the better the asymptotic properties of the statistical model. Perhaps the most appropriate initial cohort should be equal to the number of births of the particular order in the given year.

We used a log-rank test for the comparison of each pair of adjacent curves, for each order, and over the entire period for which data were available, for the Czech Republic, Hungary, and Russia. The test was not applied to Bulgaria, where the data were available by 5-year age groups only, and to Poland, where the changes are less abrupt.[5]

First, we consider Russia, where the number of births is largest. Each initial cohort was made equal to the average of the observed number of births (table 1). Two different age spans were considered. The first is 16–50 years of age, which practically covers all observed births. The second is 16–40, which means that we neglect the very small probabilities for a birth in the age interval 41–50. The latter are often neglected in demographic studies of fertility, particularly for the countries of Central and Eastern Europe, where mean ages of childbearing are low in comparison to other European countries.

Table 1 illustrates some of the smallest p-values of the log-rank test for the equality of the survival curves in two adjacent years. In the case of first births the smallest p-values were found for the pair of curves from 1994 and 1995. They are large enough that one cannot reject the null hypothesis of equality between the survival curves (and hence, the equality of the shape of the standardised fertility schedule).[6] Hence, the analysis does not provide evidence that the shape of the fertility patterns pertaining to first births is significantly different between any two adjacent years.

Let us now consider the second births. Where the initial cohort equals the observed number of births and the age span is 16–50, two cases were found in which the null hypothesis should be rejected, namely, for the pairs of curves 1990–1991 and 1991–1992. Shortening the age span increased the p-value for the second pair such that it is larger than any reasonable level of significance. The smaller initial cohort of 10,000 raised the p-value substantially, such that the null hypothesis cannot be rejected.

Table 1: *Low P-values of the log-rank test for equality of two survival curves, by order of birth, Russia, 1990-1996*

Years and order of lowest p-values:	Initial cohort equal to: 10,000. Age span:		observed births. Age span:		observed births, in thousands:
	16-50	16-40	16-50	16-40	
First births					
1994-1995	0.860	0.801	0.439	0.222	840 and 814
Second births					
1990-1991	0.498	0.705	*0.000*	*0.020*	680 and 573
1991-1992	0.657	0.856	*0.003*	0.*519*	573 and 475
Third births					
1990-1991	0.440	0.970	*0.001*	0.901	208 and 176
1991-1992	0.377	0.850	*0.001*	0.783	176 and 141
1992-1993	0.511	0.924	*0.014*	0.719	141 and 109

Figure 8 plots the standardised schedules for 1990–1991, thus helping us to understand what happened in these two years. It shows that women aged 25 in 1990 had relatively fewer second births than women in the adjacent age groups. They also had relatively fewer second births in 1991, at age 26. This is a cohort change. The assumption of no cohort change for these particular years does not hold. The value of the adjusted TFR(2) for 1991 is biased, as is the adjusted TFR. The magnitude of the bias seems to be small, though, because the cohort change is very minor. Moreover, if a smaller initial cohort size of 10,000 is chosen, the null-hypothesis of equal shape can no longer be rejected.

In the case of third births the p-values are low for three pairs of years, but a shorter age span shows that the schedules differed significantly in the age interval 41–50, where the number of births is very small.

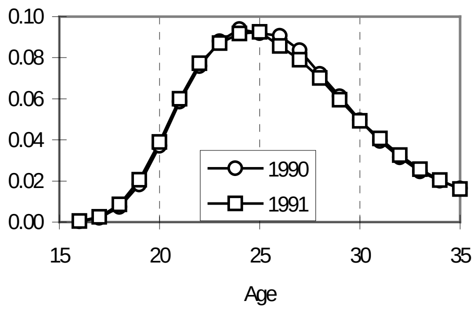


Figure 8: ***Second-order standardised fertility schedules, Russia, 1990–1991***

As far as the data for the Czech Republic and Hungary were concerned, we found no case where the p-value is small enough to reject the null hypothesis.

The results of the survival analysis can be summarised as follows. If one is interested in obtaining strict statistical equality of the age- and order-specific fertility schedules, the initial cohort should be that of observed births. Then one must reject the null hypothesis for the equality of the schedules in very few cases for the Russian data. If one is satisfied with less accuracy, while still preserving the demographic meaning of the schedules, then the null hypothesis cannot be rejected and the assumptions of the formula should be accepted for all the years studied here. In addition, the results change by reducing the span of the childbearing years. Such a reduction is appropriate in cases in which the researcher prefers to avoid the impact of variation due to very low rates.

In this section we have shown that the abrupt decline in fertility observed recently in Eastern Europe is not an obstacle to the use of the adjusted TFR. Since the basic assumption of no age-period interactions does not hold exactly, the estimated adjusted TFRs are approximations of their values that would have been achieved when all age-period interactions were controlled. We believe though that these approximated values and their unknown real counterparts will depict the same trends.

7. Conclusions

The adjusted total fertility rate suggested by Bongaarts and Feeney provides an effective fertility measure to augment the commonly used total fertility rate. In particular, the comparison between the observed and the adjusted TFR reveals the influence of tempo effects on fertility patterns, and it allows us to determine to what extent the recent fertility declines in Central and Eastern Europe are due to a postponement of births.

The application of the Bongaarts-Feeney adjustment formula reveals some important differences in the patterns of fertility decline in Eastern Europe. One can be observed in Bulgaria and in Russia up until about 1993. It is characterised by strong quantum effects: people preferred to have fewer children, while any postponement of births was weak or absent. In fact, in Russia (and to a lesser extent in Bulgaria), we can even observe a decline in the mean age at birth after 1990. In the Czech Republic this pattern is evident to a much lesser extent, in that a slight tempo effect was present and fertility did not decrease as drastically. In Hungary and Poland tempo-adjusted fertility began to increase right from the start of the transition period. A comparison of tempo-adjusted fertility with the observed TFR reveals a significant increase in tempo effects.

During the next few years (1993–1996) postponement of births comes to dominate in all five countries. It constitutes a primary change in fertility behaviour. The adjustment of the effect of postponement would increase the observed TFR by 0.5 to 0.6 during the last year of each of the five data series. Compared to the previous period, the adjusted TFR indicates a substantial increase in fertility in Russia and a sustained level in Bulgaria and Hungary, while a moderate decrease is to be observed in the Czech Republic and Poland.

These general findings indicate that the patterns of fertility change correspond to some extent to the overall social, economic, and political developments. The Czech Republic, Hungary and Poland are all doing better than Bulgaria and Russia, and their fertility is considerably higher. The transition period has been considerably more difficult for the people of Bulgaria and Russia right from the start. This may explain the decrease in quantum fertility in these two countries and the prevalence of lower levels of fertility even after the adjustment for timing. On the other hand, once tempo effects are removed, the Czech Republic, Hungary and Poland exhibit a level of the adjusted TFR that hardly anybody would associate with a demographic crisis.

This summary is of direct relevance to the two arguments mentioned in the introduction. In fact, the first argument can be rephrased as the appearance of an abrupt *quantum* fall in fertility. The findings here indicate that this was the case for just a few years and only in Bulgaria and Russia, i.e. the countries that experienced more difficulties during the transition than the other 3 countries. However, quantum fertility has been increasing in these two countries during the last few years. Thus, the demographic “crisis” was a temporary phenomenon, characteristic only of the start of the transition. The postponement of births is a typical characteristic of the “second demographic transition”, and its impact on the fertility level serves as additional evidence for the validity of the latter as an explanation of recent fertility changes.

It would make sense to make known to the general public the fact that the adjusted TFR is a more precise indicator of the reproductive potential of a population than the period TFR. The fact that it is considerably higher than the observed TFR will help correct false notions concerning demographic matters and make more precise important aspects of population policies. Nevertheless, the adjusted TFR remains purely a period measure, and any assumptions that later cohort fertility will converge to the levels of the adjusted TFR need to be expressed with considerable caution.

Acknowledgement

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NOTES

[1] Such opinions and phrases are often used in the mass media by persons who have an influence on the public, such as politicians, journalists, scientists. Evidence for such public debates can be found in English for instance in the *Current digest of Post-Soviet Press*. We list below a few examples. Declining birth rates are said to threaten the country's security (*Current Digest of Post-Soviet Press* 1997, 2). The *Current Digest of Post-Soviet Press* (1997, 1) gives a description of a report prepared by the President's Commission on Problems of Women, the Family and Demography. See also *Current Digest of Post-Soviet Press* (1996). Some members of the Russian Duma sought the President's impeachment on the basis of five accusations, one of which was the low birth rates in Russia. UNDP (1995) describes concisely the demographic crisis in Bulgaria.

[2] A more detailed review of the study of tempo and quantum effects can be found in Ní-Bhrolchain (1992) and Bongaarts and Feeney (1998).

[3] The data for the 5 countries were obtained as follows. Bulgaria: the order-specific births and the female population by five-year age groups are available in the Demographic Yearbooks, issued by the National Statistical Institute in Bulgaria, 1960 through 1997; Czech Republic: single-age and order-specific fertility rates for the period 1986-1997 were received from the Department of Demography and Geodemography, Charles University, Prague (courtesy of T. Kucera); Hungary: order-specific live births and number of females by single-year age groups can be found in the Demographic Yearbooks issued by the Hungarian Central Statistical Office, 1970-1997; Poland: as in Hungary, period 1988-1997 (courtesy of E. Fratczak); Russia: we received single-age and order-specific fertility rates over the period 1990-1996 from the Database of the Center for Demography and Human Ecology, Institute for Economic Forecasting, Russian Academy of Sciences, Moscow (courtesy of S. Zakharov).

[4] A third way to go about testing this is to employ analytical expressions developed in demography. Models like those of Coale and Trussell (1974), Rogers (1986), the Gompertz relational model due to Brass (1984), and others could be of use, after they have been fitted to order-specific fertility rates. Relational models using standard schedules might well prove to be of particular use here.

[5] The log-rank test makes use of the χ^2 statistic. This is why the size of the initial cohort matters: it defines the degrees of freedom in χ^2 through the consecutive age groups.

[6] Assume a high level of significance, such as α=0.1 or 0.05.

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Appendix: Tables with observed and adjusted TFR for all births and by birth order, and mean ages - Bulgaria, Czech Republic, Hungary, Poland, Russia

(Note: applying formulas (2) and (3) to the data in the tables may yield results that differ slightly from the reported adjusted TFRs due to rounding errors.)

Table A1: Bulgaria, 1988-1997

Year	**1988**	**1989**	**1990**	**1991**	**1992**	**1993**	**1994**	**1995**	**1996**	**1997**
All births										
TFR obs	1.96	1.92	1.82	1.64	1.55	1.46	1.37	1.24	1.23	1.09
TFR adj	2.02	2.03	1.74	1.58	1.58	1.60	1.66	1.60	1.70	-
M. Age	24.0	24.0	24.0	23.8	23.8	23.9	24.1	24.2	24.3	24.5
First births										
TFR obs	0.91	0.92	0.90	0.88	0.86	0.79	0.73	0.67	0.66	0.63
TFR adj	0.96	0.98	0.83	0.82	0.88	0.89	0.90	0.89	0.94	-

M. Age	21.9	22.0	22.0	21.8	21.8	21.9	22.0	22.2	22.6	22.8
Second births										
TFR obs	0.80	0.75	0.68	0.56	0.51	0.49	0.47	0.42	0.42	0.34
TFR adj	0.81	0.78	0.65	0.54	0.51	0.55	0.59	0.52	0.55	-
M. Age	25.1	25.1	25.2	25.0	25.1	25.0	25.3	25.4	25.7	25.9
Third births										
TFR obs	0.17	0.17	0.15	0.13	0.11	0.11	0.10	0.09	0.09	0.07
TFR adj	0.16	0.18	0.17	0.15	0.12	0.09	0.10	0.11	0.13	-
M. Age	26.7	26.7	26.8	26.9	27.1	27.0	26.8	27.0	27.2	27.6
Fourth and higher-order births										
TFR obs	0.09	0.09	0.09	0.08	0.07	0.07	0.07	0.06	0.06	0.05
TFR adj	0.09	0.10	0.09	0.07	0.07	0.07	0.08	0.07	0.08	-
M. Age	29.4	29.6	29.6	29.6	29.6	29.5	29.5	29.7	29.9	30.3

Table A2: Czech Republic, 1986-1997

Year	**1986**	**1987**	**1988**	**1989**	**1990**	**1991**	**1992**	**1993**	**1994**	**1995**	**1996**	**1997**
All births												
TFR obs	1.94	1.91	1.94	1.87	1.89	1.86	1.71	1.68	1.44	1.28	1.18	1.17
TFR adj	-	1.99	2.03	1.94	1.92	1.93	1.86	2.06	2.09	1.96	1.78	-
M. age	24.6	24.7	24.7	24.7	24.8	24.7	24.8	25.0	25.4	25.8	26.1	26.4
First births												
TFR obs	0.91	0.90	0.91	0.89	0.90	0.91	0.82	0.77	0.64	0.56	0.52	0.53
TFR adj	-	0.92	0.93	0.90	0.87	0.93	0.90	0.97	1.00	0.90	0.81	-
M. age	22.4	22.4	22.4	22.5	22.5	22.4	22.5	22.6	22.9	23.3	23.7	24.0
Second births												
TFR obs	0.73	0.72	0.73	0.71	0.71	0.68	0.64	0.65	0.55	0.51	0.47	0.46
TFR adj	-	0.76	0.78	0.72	0.74	0.73	0.70	0.78	0.77	0.77	0.71	-
M. age	25.5	25.5	25.6	25.6	25.6	25.7	25.8	25.9	26.1	26.4	26.8	27.1
Third and higher-order births												
TFR obs	0.29	0.28	0.29	0.28	0.28	0.27	0.26	0.26	0.24	0.21	0.19	0.19
TFR adj	-	0.31	0.32	0.31	0.31	0.28	0.26	0.31	0.33	0.30	0.25	-
M. age	29.5	29.6	29.7	29.8	29.9	30.0	29.9	30.0	30.2	30.5	30.8	31.0

Table A3: Hungary, 1988-1996

Year	**1988**	**1989**	**1990**	**1991**	**1992**	**1993**	**1994**	**1995**	**1996**
All births									
TFR obs	1.80	1.82	1.87	1.88	1.78	1.69	1.64	1.57	1.46
TFR adj	1.88	1.83	1.94	2.07	2.02	1.97	1.97	2.11	-
M. age	25.4	25.5	25.6	25.7	25.8	26.0	26.2	26.3	26.5
First births									
TFR obs	0.82	0.80	0.82	0.82	0.77	0.71	0.68	0.65	0.62

TFR adj	0.90	0.82	0.85	0.95	0.92	0.86	0.88	0.89	-
M. age	23.2	23.2	23.2	23.3	23.5	23.6	23.8	24.1	24.4
Second births									
TFR obs	0.66	0.68	0.68	0.68	0.64	0.60	0.57	0.55	0.50
TFR adj	0.68	0.70	0.73	0.72	0.74	0.75	0.66	0.69	-
M. age	26.5	26.5	26.6	26.6	26.7	26.9	27.1	27.2	27.5
Third births									
TFR obs	0.21	0.23	0.25	0.24	0.24	0.24	0.26	0.24	0.21
TFR adj	0.20	0.21	0.26	0.29	0.24	0.24	0.28	0.36	-
M. age	30.7	30.6	30.5	30.7	30.8	30.7	30.7	30.9	31.4
Fourth and higher births									
TFR obs	0.11	0.11	0.12	0.13	0.13	0.13	0.14	0.13	0.13
TFR adj	0.10	0.10	0.10	0.11	0.12	0.12	0.15	0.17	-
M. age	35.2	35.1	34.9	34.7	34.6	34.5	34.4	34.7	34.8

Table A4: Poland, 1989-1997

	1989	**1990**	**1991**	**1992**	**1993**	**1994**	**1995**	**1996**	**1997**
All births									
TFR obs.	2.09	2.05	2.07	1.95	1.86	1.81	1.62	1.59	1.51
TFR adj.	-	2.05	2.12	2.14	2.19	2.16	1.94	1.96	-
M. age	26.0	26.0	26.0	26.1	26.4	26.6	26.7	26.8	26.9
First births									
TFR obs.	0.84	0.85	0.85	0.78	0.73	0.70	0.65	0.65	0.64
TFR adj.	-	0.83	0.89	0.87	0.82	0.81	0.73	0.75	-
M. age	23.3	23.3	23.3	23.4	23.5	23.6	23.7	23.9	24.0
Second births									
TFR obs.	0.70	0.66	0.65	0.62	0.59	0.56	0.51	0.50	0.48
TFR adj.	-	0.67	0.66	0.67	0.7	0.71	0.66	0.66	-
M. age	26.4	26.4	26.4	26.4	26.5	26.8	26.9	27.2	27.4
Third births									
TFR obs.	0.33	0.32	0.33	0.31	0.31	0.30	0.26	0.24	0.22
TFR adj.	-	0.33	0.33	0.35	0.39	0.38	0.32	0.31	-
M. age	29.2	29.3	29.3	29.3	29.5	29.7	29.9	30.1	30.3
Fourth and higher-order births									
TFR obs	0.22	0.22	0.23	0.23	0.24	0.24	0.20	0.19	0.18
TFR adj	-	2.05	2.12	2.14	2.19	2.16	1.94	1.96	-
M.age	32.4	32.4	32.4	32.4	32.6	32.7	32.8	32.9	33.1

Table A5: Russia, 1990-1996

	1990	1991	1992	1993	1994	1995	1996	1997
All births								
TFR obs	1.91	1.75	1.56	1.36	1.38	1.33	1.27	1.22
TFR adj	-	1.67	1.52	1.34	1.50	1.62	1.69	-
M.age	25.3	25.1	24.9	24.7	24.7	24.9	25.1	25.3
First births								
TFR obs	0.99	0.95	0.89	0.81	0.83	0.79	0.75	0.71
TFR adj	-	0.93	0.87	0.77	0.87	0.96	0.88	-
M.age	22.7	22.7	22.7	22.6	22.6	22.7	22.9	23.0
Second births								
TFR obs	0.64	0.55	0.47	0.40	0.40	0.39	0.37	0.37
TFR adj	-	0.51	0.45	0.42	0.47	0.48	0.63	-
M.age	26.9	26.8	26.7	26.7	26.8	27.0	27.2	27.8
Third births								
TFR obs	0.18	0.16	0.13	0.10	0.10	0.10	0.10	0.09
TFR adj	-	0.15	0.13	0.10	0.11	0.13	0.13	-
M.age	29.8	29.7	29.7	29.7	29.7	29.9	30.2	30.4
Fourth and higher-order births								
TFR obs	0.10	0.09	0.07	0.05	0.05	0.05	0.05	0.05
TFR adj	-	0.09	0.07	0.05	0.05	0.06	0.06	-
M.age	32.8	32.7	32.7	32.7	32.6	32.7	32.9	32.9