On the Tempo and Quantum of First Marriages in Austria, Germany, and Switzerland: Changes in Mean Age and Variance

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Abstract

Period marriage rates have been falling dramatically in most industrial societies since the beginning of the 1970s. As has been shown in the literature, part of this decline is due to the postponement of marriage to later ages. However, the change in variance has been ignored so far. In the case of Austria, Germany, and Switzerland, this paper explores how much of the change in female first marriage rates can be attributed to tempo effects caused by changes in the mean age and variance, and how much of it is due to quantum effects, i.e., the proportion of women who ever marry. In all three countries we find a significant share of the decline in first marriage rates due to tempo distortions, though on different levels.

European Demographic Research Papers are working papers that deal with all-European issues or with issues that are important to a large number of countries. All contributions have received only limited review.

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INTRODUCTION

The apparent decline in first marriage rates and the increase in the age at first marriage in Germany, Austria and Switzerland as well as in most other industrial societies has been described as one of the great social changes of our time, with some calling it a feature of the "second demographic transition" (Lesthaeghe 1995). Numerous studies show that marriage levels influence fertility because married fertility is still higher than unmarried fertility (Goldstein 2002), although with increasing non-marital childbearing in many Western European countries (Kiernan 2001). Hence, the mean age of marriage affects the average number of children, the timing and spacing of births (Heckman et al. 1985), and thus the mean interval between successive generations (Lutz et al. 2003).

Other studies show a negative relation between the individual marriage age and the risk of divorce, negatively contributing to the increase in divorce rates (Engelhardt 2002). Moreover, in many industrialised countries the first marriage rate is a social indicator affecting the welfare of adults and children since married persons and their children are on average wealthier than unmarried individuals and children with single parents (Waite 1995, McLanahan and Sandefur 1984). In addition, from a sociological perspective, first marriages are of interest as an indicator of the degree of individualisation of a society, in which unmarried individuals as well as cohabiting couples live with less connexion to the traditional norms of their society (Beck-Gernsheim 1998).

Several theories have been developed about why people marry and what factors influence the timing of marriage. These theories range from purely economic explanations to institutional approaches. In family economics, marriage is seen as a rational choice made by individuals for whom the benefits of getting married outweigh the benefits of staying single (Becker 1973, 1974, 1991). Given complementarity of men and women in the household production, these individuals would be more productive in a joint household than they would be if they remained single. If each sex specialises in its comparative advantage, then the sexual division of labour within households creates gains within marriage. The theory thus ascribes recent declines in marriage (and increases in divorce) to reduced gains from marriage by a rise in women's labour force participation and earnings (economic independence) and by a fall in fertility because a sexual division of labour becomes less advantageous. However, the specialisation model has been heavily criticised because of two reasons. First, "contrary to the theory, women with greater labour market potential are now more likely to marry" (Cherlin 2000, p. 127). Oppenheimer (2000, p. 285) explains the latter by the facts that "marriage, more than cohabitation, encourages the accumulation of pooled savings and capital investment that provide greater long-term economic security. [S]ex role specialisation is essentially a high risk and inflexible family strategy in an independent nuclear family system". Thus, highly educated women with greater labour market potential are more attractive to their future spouses than less educated women with poor employment prospects.

Secondly, the specialisation model of Becker assumes common preferences and interests of men and women and strictly divided roles inside and outside the house by gender. An alternative framework, which takes the latter criticisms into account, is represented by the so-called bargaining models, where spouses differ by their preferences and rather establish their roles through a process of bargaining (Cherlin 2000, Lundberg and Pollak 1996). Within this perspective, Cherlin (2000) claims that delayed marriage can be explained by the fact that women's bargaining position has improved. Moreover, women are incorporating premarital cohabitation into the search and bargaining processes because cohabitation provides a better opportunity to observe men's earnings potential and willingness to share household and childraising tasks (Cherlin 2000).

The earnings potential of young men is particularly stressed by Oppenheimer (1988, 2000). In particular, Oppenheimer argues that the pace of marriage formation is affected by the pace and difficulty of the transition to a stable work career. However, Oppenheimer and Lewin (1999, p. 193) find that it is still unlikely that "women's familial roles are normatively defined in terms of their ability to make a major and long-term stable income contribution to the family to the same degree as men's." Hence, men's work careers and career maturity are playing a more important role than women's for the timing of marriage of both men and women.

Recalling that Oppenheimer's arguments are based on the normative family roles of men and women leads to the second approach as to why and when people marry, i.e., to the institutional perspective. According to Goode (1982, p. 11, in Goldstein und Kenney, 2001), marriage is supported by "a structure of norms, values, laws, and a wide range of social pressure". Thus, the institutional perspective locates the decline of marriage rates in changing social norms, values, attitudes and preferences as well as in changing laws, which enable unmarried couples living legally together in consensual unions. In Germany, for instance, a law (the so-called "Kuppeleiparagraph") prohibited renting out any accommodation to unmarried couples until its abolition in 1974.

In fact, there is a strong empirical association between increasing cohabitation rates and decreasing marriage rates in many Western societies (e.g., Bumpass and Lu 2000, Kiernan 1999). Besides reducing the number of actual marriages since couples move together without marrying, cohabitation might also be viewed as a temporary phase before marriage and might be interpreted as a cause of the postponement of marriage (cf. Manting 1996). In fact, in 60.7% of the Austrian marriages in the year 2000, the groom and bride had been living together before (Statistik Austria 2001). Since most cohabitations end in marriages (Kiernan 1999, Murphy 2000), the shift to cohabitation as the dominant mode for first partnership plays an important role in the delay of first marriage (Bumpass and Lu 2000, Ermisch and Francesconi 2000, Toulemon 1997). Hence, cohabitation has a quantum ('couples do not marry') and a tempo effect ('couples postpone marriage') on the number of marriages.

The extent of quantum and tempo changes is also interpreted differently by the theories discussed above. On the one extreme, researchers following the school of Becker argue that the increasing economic independence of women due to longer educational investment, higher educational attainment, and greater labour-force participation will lead not only to delayed marriage but also to a decline in the proportion of women who ever marry (e.g., Bloom and Bennett 1990). On the other hand, Oppenheimer (1988, p. 587) claims that “[t]he consequence [of women’s greater economic independence] is an increase in delayed marriage with some accompanying greater risk of non-marriage [...]. But all this is consistent with continued high gains to marriage as well as with a continued desire to marry.” Motivated by this theoretical debate, we aim to quantify both tempo and quantum effects in the decline in female first marriage rates for Austria, Germany, and Switzerland.

The most well-known methodology to correct for tempo distortions was derived by Bongaarts and Feeney (1998), originally developed in order to adjust total fertility rates. However, Bongaarts and Feeney (1998) took into account only changes in the mean age of the fertility rates and assumed no age-period interactions. Kohler and Philipov (2001) overcame these restrictions by assuming tempo effects as well which are caused by changing variances of the fertility schedule. Kohler and Ortega (2002) refined the adjustment for tempo distortions caused by mean and variance changes and applied them to parity-specific fertility rates.

Goldstein (2002) already applied tempo adjustments to first marriage rates in France. However, he restricted the analysis to changes in the mean age by applying only the methodology derived by Bongaarts and Feeney (1998). To our knowledge, both mean and variance effects have never been considered for first marriages yet. Therefore we try to quantify possible tempo effects by accounting for changes in the mean age and variance of first marriages in Austria, Germany and Switzerland.

In section 2, we first review the decline in observed period marriage rates that has occurred in Austria, Germany and Switzerland since the early 1970s. In section 3, we outline the theory of tempo and variance adjustment showing the effects of changes in tempo and variance on observed period rates. In particular, section 3.1 discusses the different methods of adjustment developed by Bongaarts and Feeney (1998), Kohler and Philipov (2001), and Kohler and Ortega (2002). Section 3.2 presents the results of applying the tempo and variance adjustment to female first marriage data from Austria, Germany and

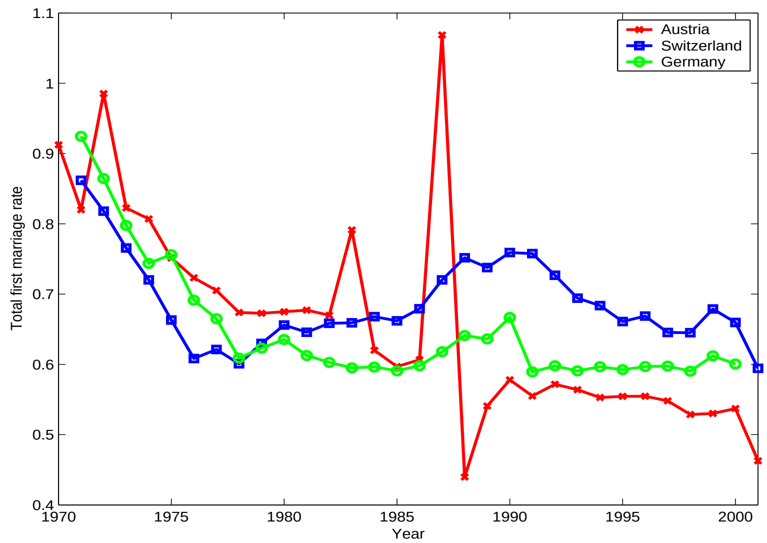


Figure 1 Female total first marriage rate for Austria, Germany and Switzerland.

Switzerland. We also compare, in section 3.3, the adjusted period marriage rates with those observed and estimated for the cohorts. Finally, we discuss the implications of the mean and variance adjustment for the quantum of first marriage decline in section 4.

EMPIRICAL EVIDENCE ON FEMALE FIRST MARRIAGES RATES

In our empirical analysis we focus on the countries Germany, Austria and Switzerland, for all three of which demographic statistics in first marriages were made available to us by their statistical offices.[1] Austria represents a

[1]We would like to thank Statistik Austria, the Swiss Federal Statistical Office, and the German Statistical Office for supplying the number of female first marriages by age and the number of the female population by age and marital status on an annual basis. For Switzerland, we were fortunate to also get cohort data from Calot (1998).

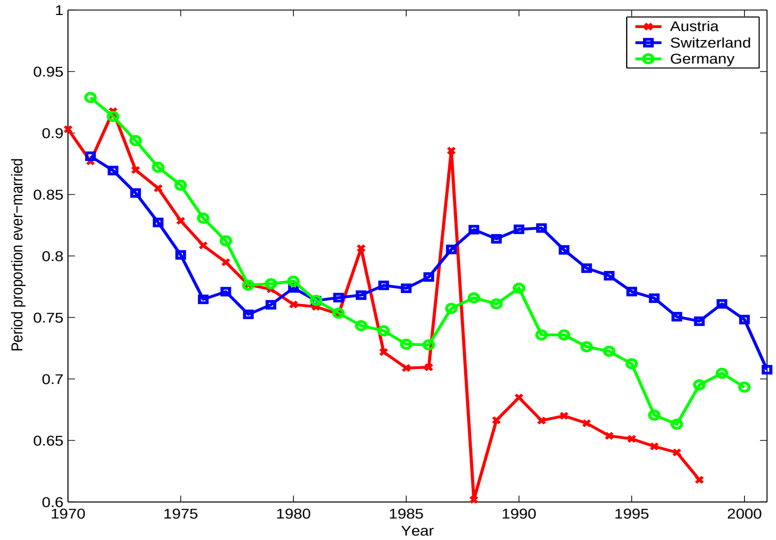


Figure 2 Female period proportion ever married for Austria, Germany and Switzerland.

particular case where the booms and busts in first marriage rates have often been attributed to changes in public policy (for details see below). Germany and Switzerland in comparison are characterised by an almost steady decline with a slight marriage boom in the 1980s in first marriage rates since World War II without showing extreme peaks like Austria. We restrict our analysis to female first marriages. Males are on average two to three years older than females at first marriage. This result seems to be valid over space and time in Western Europe (Diekmann 1987, United Nations 2000).

The observed decline in period marriage rates for Austria, Germany and Switzerland is illustrated in Figure 1 and Figure 2. Two measures are shown. The first is the female total first marriage rate, TFMR (Figure 1). Analogous to the total fertility rate, the index is estimated as the sum of the female age-specific rates of the second type[2] of first marriage, F_x, given by the ratio of the number of marriages in the age group x to $x+1$, D_x, and the number of women aged x to $x+1$, K_x.[3] Hence,

$$\mathrm{TFMR} = \sum_{x=15}^{49} \frac{D_x}{K_x} = \sum_{x=15}^{49} F_x. \tag{1}$$

The second measure is the period proportion ever married, PPEM, from the period nuptiality table:

$$\mathrm{PPEM} = 1 - l_{50}, \tag{2}$$

where l_{50} is the probability of still being single at age 50 as given by

$$l_{50} = \prod_{x=15}^{49} (1 - q_x). \tag{3}$$

Assuming that the rate of marriage is constant within each single year, we obtain the usual exponential model for the life table probability of marrying aged x to $x+1$, q_x,

$$q_x = 1 - \exp\left(-\frac{D_x}{N_x}\right) = 1 - \exp\left(-M_x\right), \tag{4}$$

where M_x are the age-specific rates of the first type[4], which are given by the ratio of the number of marriages in the age group x to $x+1$, D_x, and the number of single women aged x to $x+1$, N_x. Taking equations (3) and (4) into account, equation (2) modifies to

$$\mathrm{PPEM} = 1 - \exp\left(-\sum_{x=15}^{49} M_x\right). \tag{5}$$

Both indicators are synthetic period measures, applying to a fictional cohort which over its lifetime experienced the age-specific rates observed in

[2] Analogous expressions are 'frequencies' and 'incidence rates'.

[3] As common in the demographic literature, we restrict our analysis to women between ages 15 to 49 and neglect first marriages of women aged 50 and older.

[4] Other expressions commonly used for these rates are 'occurrence-exposure rates' and 'intensities'.

a particular period. When rates are unchanging over time, the two measures will both be equal to the experience of actual cohorts. However, when marriage rates are changing over time, the total first marriage rate and the period proportion ever married will not necessarily follow each other. Péron (1991) contrasts the information given by age-specific first marriage rates of the second type and current nuptiality tables, where the sum of the former defines the period TFMR and the period PPEM is derived by the latter (cf. equation (2)). Rewriting the age-specific first marriage frequencies

$$F_x = \frac{D_x}{K_x} = \frac{D_x}{N_x}\frac{N_x}{K_x} = M_x\frac{N_x}{K_x}, \tag{6}$$

yields that they can be decomposed into the age-specific first marriage intensities times the share of single women aged x to $x+1$, N_x/K_x. While the first expression represents current nuptiality behaviour, the second results from previous marriage behaviour. Péron (1991, pg. 1440) concludes that the age-specific marriage frequencies can be used to "trace movements in the number of marriages resulting from the link which exists between present and past nuptiality, current nuptiality tables provide information about the nuptiality behaviour of individuals during a short period of observation".

Since the period TFMR sums over age groups which are born in different years, the TFMR can exceed one. Summing over the same rates in a cohort always implies a total first marriage rate less than one. In contrast, the PPEM is restricted by one by definition from the nuptiality table. However, the TFMR is the most widely cited measure on first marriage and is also used by Eurostat for comparing period levels within Europe (Goldstein 2002).

As we can see in the figures for Austria, according to either measure the level of first marriage indicated by period rates has fallen dramatically since the early 1970s with three peaks in 1972, 1983 and 1987. Both measures begin with marriage levels above 90 per cent. Marriage rates then fall either to a level of only 50 per cent, according to the TFMR, or to a level of about 65 per cent according to the PPEM. In either case, the decline was dramatic, with the predicted proportion ever married reaching historic lows. The peaks are related to the introduction and abolition of a marriage grant of about

EUR 545 given to all first married persons (Gisser et al. 1990, Prioux 1992). The introduction in 1972 led to a postponement of marriages from the previous year. The second peak is most likely caused by marriages brought forward due to political discussions about abolishing the marriage grant in mid-1983. The marriage boom in the second half of 1987 and the following bust in 1988 was caused by the announcement in August 1987 to cut the marriage grant by January 1988.

In Germany and Switzerland, the first marriage rates do not show extreme peaks like in Austria. In Germany, we observe an almost steady decline in female first marriages during the 1970s which levelled off during the 1980s with a slight increase at the end of the 1980s. Unfortunately, we were not able to get separate data for former West and East Germany. Therefore the decline in marriage age in the beginning of the 1990s might be due to the unification of West and East Germany. The inclusion of East Germany in the official statistics in 1990 lowered the total first marriage rate considerably because of the dramatic drop in marriages in East Germany after unification.

In Switzerland the first marriage rate for women had declined to a very low level in the mid-1970s, then rose until 1988, only to fall again substantially back to the mid-1970s levels by the end of the 1990s. Just like in Austria, the shapes for TFMR and PPEM look quiet similar in Switzerland and Germany.

As stated before, the fall in the first marriage rates was accompanied by a postponement of first marriages. The mean age of first marriage is the most common measure for first marriage timing. If the mean age is rising from year to year, this is taken as evidence for events being delayed. We calculated the mean ages from the age-specific first marriage rates in order to be consistent with the derivation of the adjustment formula by Bongaarts and Feeney (1998) and Kohler and Philipov (2001). The resulting figures may be different from those published by national statistical offices, since they sometimes give the median instead of the mean. Moreover, as Goldstein (2002) finds, sometimes the "mean age is calculated as the mean of the people who actually married", where the latter is sensible to changes in the age structure of the population.

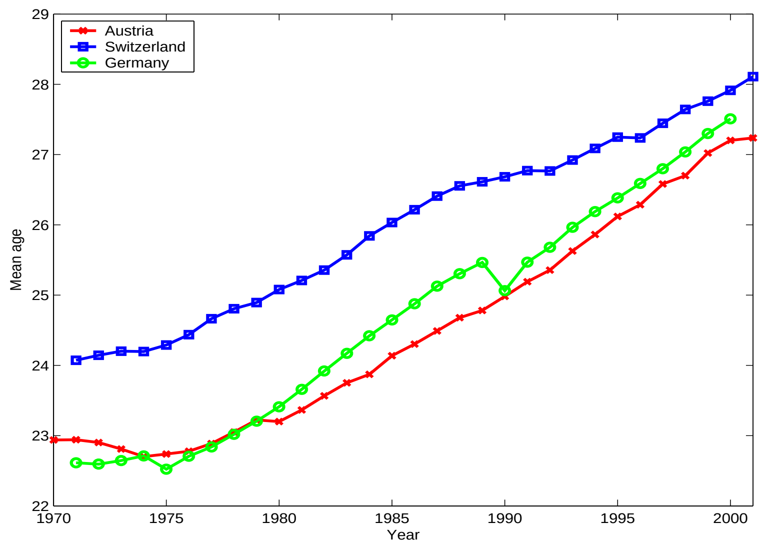


Figure 3 Mean age of first marriage (based on age-specific first marriage rates of the second type for Austria, Germany and Switzerland).

As we can see in Figure 3, changes in marriage timing have indeed paralleled the decline in marriage levels, although the two processes neither started nor ended at the same time. Age at first marriage for Germany, Austrian and Swiss women rose rapidly, beginning in the mid-1970s, and this rise has continued to the present. In Austria, for instance, the 25-year period between 1975 and 2000 brought a 4.3 year increase in the average age of first marriage, or a delay of roughly 0.17 years per year. While the mean age at first marriage is quite similar in Germany and Austria, it is at a considerably higher level in Switzerland. The latter result is also consistent with Calot's (1998) findings.

From a historical perspective, age at first marriage was exceptionally low in the early 1970s. Calot (1998, pg. 41) finds that "[f]rom the end of the 1930s, age at first marriage began to fall sharply and continued so until the 1970s[, when] the marriage rate particularly rose among the younger age groups." Similarly, Heilig (1985) reveals that during the 1950s and 1960s the number of early marriages, particularly before the age of 22 years, increased in Germany.[5] These early marriages were also popular during the 1960s in Austria (Gisser et al. 1990). However, in the 1970s the age at marriage started to rise, where Calot (1998, pg. 41) finds, for instance in Switzerland, that "the [downward] trend [age at first marriage] reversed more quickly than it had developed".

Standard deviations of the age-specific first marriage rates for Germany, Austria and Switzerland are shown in Figure 4. While the mean age at first marriage increased over time, the variance decreased till the first half of the 1980s and increased thereafter in all three countries. Interestingly, the time series of the standard deviation of age-specific first marriage rates in Austria, Germany and Switzerland are very close to each other over the three decades. The pace of change of the mean age and the variance can be used to estimate

[5] One explanation for the very early age at first marriage in 1970, according to Scheller (1985), is the result of a cultural lag. In particular, a change in attitudes towards sexual behaviour took place in the 1960s, implying an increase in pre-marital sexual relationships. However, norms about births out of wedlock did not change in the same extent and single mothers as well their illegitimate children were discriminated.

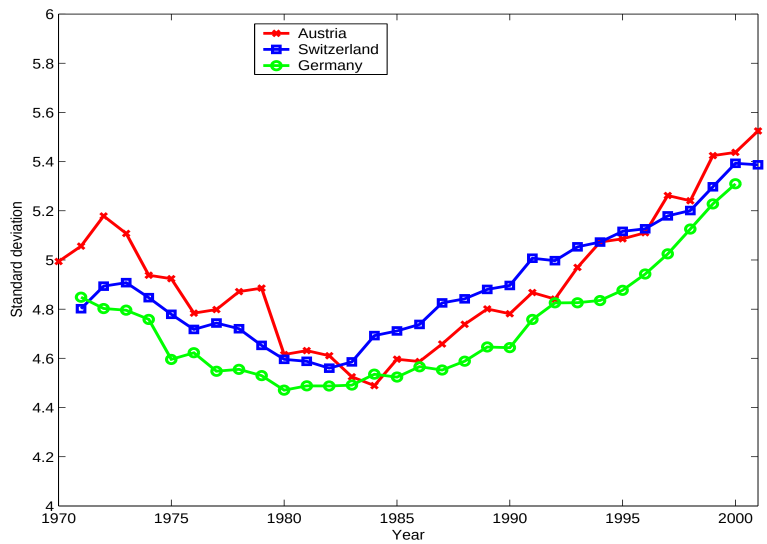


Figure 4 Standard deviation of age-specific first marriage rates for Austria, Germany and Switzerland.

rates adjusted for tempo and variance as we will see in the following section.

ADJUSTING FOR TEMPO AND VARIANCE

Methods

In order to correct for tempo effects in the time series of the TFMR and the PPEM, we apply the method of Bongaarts and Feeney (1998), which was extended by Kohler and Philipov (2001) (and further refined by Kohler and Ortega 2002) for variance effects and originally developed for fertility rates.

Bongaarts and Feeney (1998) present a simple formula to estimate distortions in the total fertility rate caused by tempo effects of childbearing. Translating fertility rates into first marriage rates, we may employ this formula for analysing tempo distortions in first marriage. In particular, if marrying is postponed and, subsequently, the mean age at first marriage increases, the observed total first marriage rate is lower than in the absence of such timing changes. In the opposite case, if first marriage occurs at an earlier age, the mean age at first marriage decreases, and hence the observed total first marriage rate is higher than without the change in timing. Following Bongaarts and Feeney (1998), the total first marriage rate at time t, $\text{TFMR}(t)$, is composed by the product of the quantum, $\text{TFMR}'(t)$, of first marriage at time t and a factor representing tempo distortions, $(1-r(t))$, i.e.,

$$\text{TFMR}(t) = (1-r(t))\,\text{TFMR}'(t), \tag{7}$$

where $r(t)$ is the annual rate of change of the mean age at first marriage. However, Bongaarts and Feeney (1998) assume that there are no age-period interactions during the derivation of this formula. Kohler and Philipov (2001) overcome this restriction by assuming that the period-specific tempo is also dependent on age, i.e.,

$$r(a,t) := \gamma(t) + \delta(t)\,(a - \bar{a}(t))\,. \tag{8}$$

The first term is similar to Bongaarts and Feeney (1998), where $\gamma(t)$ turns out to be the (linear) rate of change of the mean age $\bar{a}(t)$ of the first marriage schedule in the absence of tempo changes at time t, which corresponds to the adjusted first marriage schedule. Furthermore, it is shown that $\delta(t)$ is the (exponential) rate of change of the standard deviation of the adjusted first marriage schedule. Hence assuming age-period interactions in the tempo effect according to equation (8) implies that changes in the variance of the first marriage schedule are taken into account via $\delta(t)$.

In the case of $\gamma(t) > 0$ and $\delta(t) > 0$, there occurs a general postponement of marrying. Then the tempo changes $r(a,t)$ are less than $\gamma(t)$ for ages $a < \bar{a}(t)$ and $r(a,t)$ exceeds $\gamma(t)$ for ages $a > \bar{a}(t)$. If $\delta(t) < 0$, the tempo changes $r(a,t)$ exceed $\gamma(t)$ for $a < \bar{a}(t)$ and $r(a,t)$ falls below $\gamma(t)$ for older ages $a > \bar{a}(t)$. Hence, the first marriage schedule is affected by tempo effects for different ages in different ways (Kohler and Philipov 2001). Moreover, Kohler and Philipov (2001, p. 8) prove that "the observed total [first marriage] rate does not depend on the extent of variance changes $\delta(t)$", i.e.,

$$\mathrm{TFMR}(t) = (1-\gamma(t))\,\mathrm{TFMR}'(t). \tag{9}$$

However, since $\gamma(t)$ is the rate of change of the mean age of the adjusted rates, the observed mean age has to be freed from variance distortions in order to compute $\gamma(t)$.

Summing up, in order to adjust the total first marriage rate for tempo changes one has to divide the observed total first marriage rate by $(1-r(t))$ if one ignores the age-period interactions (Bongaarts and Feeney 1998), or otherwise by $(1-\gamma(t))$ if they are to be considered (Kohler and Philipov 2001).

In order to adjust the PPEM at time t for tempo effects, we employ the approximation derived by Goldstein (2002), i.e.,

$$\mathrm{PPEM}'(t) \approx 1-(1-\mathrm{PPEM}(t))^{\frac{1}{1-r(t)}}\,. \tag{10}$$

Correcting additionally for variance effects, we replace $r(t)$ by $\gamma(t)$ in formula (10).[6]

[6]Under the framework assumed in Kohler and Ortega (2002), the proof follows straightforward from the proof of their Result 8 on page 141 by setting the integral limits to the lower and upper age limits of the intensity schedule. Moreover, it can be shown that the equality holds in equation (10).

Since time series are subject to random fluctuation, we use for both methods smoothed time series for the adjustment of TFMR and PPEM. Otherwise, large unexplained fluctuations may emerge (see also Kohler and Philipov 2001). In particular, following Kohler and Ortega (2002), we model the time series as Integrated Random Walk (IRW) and apply state-space smoothing (Young 1994, Ng and Young 1990). Applying IRW methods provides an estimate of the slope of the time series which can be used in the case of the mean age of first marriage as an estimate of the tempo $r(t)$ in the sense of Bongaarts and Feeney (1998).[7]

Kohler and Philipov (2001) present an iterative method to estimate $\gamma(t)$ and $\delta(t)$, since both values depend mutually on adjusted age-specific marriage rates, which cannot be observed. In particular, they iteratively adjust the moments of age-specific frequencies, where $\gamma(t)$ is the derivative of the adjusted mean age and $\delta(t)$ is the derivative of the log of the adjusted standard deviation (for details on the iteration refer to Kohler and Philipov 2001).[8] Kohler and Ortega (2002) use a different iteration procedure where they iteratively adjust the intensity schedule itself (for details on the iteration refer to Kohler and Ortega 2002).[9] For adjusting frequencies, Kohler and Ortega (2002) recommend to adjust the corresponding intensities and then transform these adjusted intensities into adjusted frequencies by employing formula (6). Consequently, we adjust the total first marriage rates for changes in mean age and variance according to both algorithms, i.e., Kohler and Philipov (2001) and Kohler and Ortega (2002), in order to highlight also the difference in adjusting the frequencies directly (Kohler and Philipov 2001) and adjusting the

[7] In contrast, Bongaarts and Feeney (1998) measure the annual rate of change of the mean age at first marriage, $\mu(t)$, according to $\hat{r}(t) = (\mu(t+1) - \mu(t-1))/2$.

[8] Kohler and Philipov (2001) recommend to derive $\delta(t)$ from a regression of the observed time series of the logarithm of the standard deviation on a polynomial of time (for further details see Kohler and Philipov 2001). However, we found that the adjusted time series heavily depend on the degree of the polynomial.

[9] During our simulations, we found out that the convergence behaviour depends on the right choice of the noise-variance ratio of the smoothing algorithm. Many thanks to Hans-Peter Kohler, who helped us to find the appropriate value of this smoothing parameter.

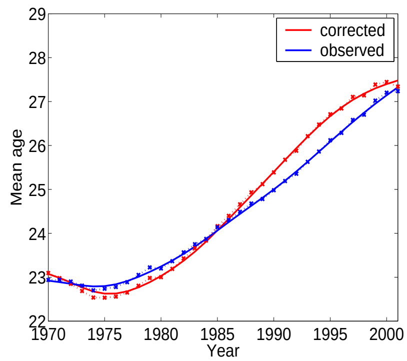
intensities and then transforming into adjusted frequencies (Kohler and Ortega 2002). The period proportion ever married is only adjusted according to the algorithm by Kohler and Ortega (2002). We implement all algorithms in MATLAB (The MathWorks, Inc. 2002).

Rates Adjusted for Tempo and Variance

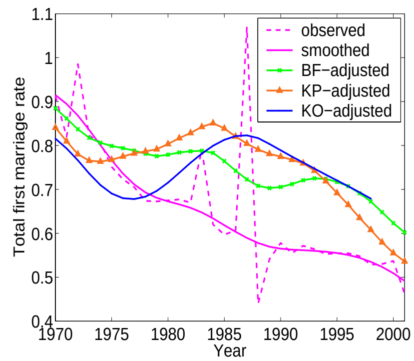
Austria

The observed mean age for first marriage frequencies and its variance-corrected values, which correspond to $\bar{a}(t)$ in equation (8), are depicted in Figure 5(a) for Austria. In the first half of the 1970s the mean age at first marriage showed a slight drop but starting from 1975 it has been increasing continuously. When correcting additionally for variance distortions, the decrease of the mean age at first marriage would have been more pronounced and hence the corrected mean age was lower than the observed figure till the mid-1980s. This is due to the decrease of variance over the same period of time. With the increase of variance in the mid-1980s, the observed mean age was lower than the corrected one, implying a sharper increase of the variance-corrected mean age over time from 1985 onwards. Since the mid-1990s the increase in the variance-corrected mean age at first marriage has slowed down.

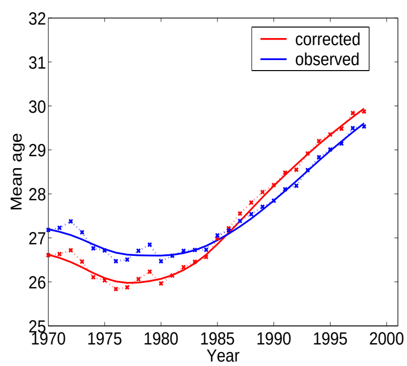
Given the observed changes in tempo, we would expect the Bongaarts-Feeney adjusted first marriage rate to be lower than the observed TFMR in the first half of the 1970s and to lie above it thereafter. Since the drop in the variance-corrected mean age at first marriage was stronger than in the observed one until the second half of the 1970s, the Kohler-Philipov adjusted TFMR should be even less than the Bongaarts-Feeney adjusted TFMR. Afterwards, the mean age corrected for variance changes increased more strongly until about 1995. Therefore the Kohler-Philipov adjusted TFMR should lie above the Bongaarts-Feeney adjustment between the late 1970s and 1995. As described above, the increase of the variance-corrected mean age at first marriage almost ceased in the second half of the 1990s, and therefore the Kohler-Philipov adjusted TFMR should approach the observed TFMR.



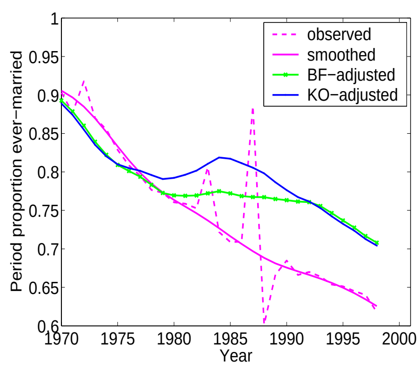
(a) Mean age of first marriage frequencies



(b) Total first marriage rate



(c) Mean age of first marriage intensities



(d) Period proportion ever married

Figure 5 Mean age of female first marriage for Austria, observed and corrected for variance changes, and Austrian period measures of female first marriage, observed, smoothed, adjusted for changes of the mean age according to Bongaarts and Feeney (1998, BF), and adjusted for changes of the mean and variance according to Kohler and Philipov (2001, KP) as well as to Kohler and Ortega (2002, KO).

If the intensities are adjusted according to Kohler and Ortega (2002) and then transformed into frequencies, the adjusted TFMR decreases steeply until the second half of the 1970s and then increases just as steeply again for the next decade. Since the second half of the 1980s the Kohler-Ortega adjusted TFMR has been decreasing again. The adjustment factor of the Kohler-Ortega method is derived from the variance-corrected mean age of intensities (cf. Figure 5(c)). The latter slightly decreases until the second half of the 1970s and then strongly increases until the end of the 1980s. From 1990 onwards, the mean age of the intensities, corrected for variance effects, further rises though at a less pronounced rate. The slight decrease followed by a strong increase of the mean age of the intensities causes therefore the down and up observed for the Kohler-Ortega adjusted TFMR in the 1970s and 1980s. Since, the pace of the increase of the variance-corrected mean age of the intensities slows down in the 1990s, the Kohler-Ortega adjusted TFMR decreases again.

Comparing the variance-corrected mean age of the intensities to the observed one, we find that both time series evolve similarly over time, though at different levels, during the first half of the 1970s and during the 1990s. This explains why the Bongaarts-Feeney adjusted PPEM and the Kohler-Ortega adjusted PPEM are almost the same during these periods. Between 1975 and 1990 the variance-corrected mean age of intensities increases more strongly than the observed one, and therefore the Kohler-Ortega adjusted PPEM lies above the Bongaarts-Feeney adjusted PPEM. Finally, since the variance-corrected as well as the observed mean age of the intensities show a slight decrease until about the second half of the 1970s and an increase afterwards, the adjusted PPEM series lie below the observed PPEM until about the second half of the 1970s and above afterwards.

Adjusted for tempo, we see that the period measures show smaller declines in marriage than typically reported. The Bongaarts-Feeney adjusted TFMR slightly declined from 1970 to 2000 by about 28 percentage points compared to about slightly more than 40 percentage points by the observed TFMR (cf. Figure 5(b)). The adjustment according to Bongaarts and Feeney (1998) has a similar effect on the PPEM measure (cf. Figure 5(d)). Instead of dropping by about 30 percentage points, the adjusted series declines just by about 18 per cent.

Adjusted additionally for variance, the Kohler-Philipov adjusted TFMR and the Kohler-Ortega adjusted TFMR and PPEM even increased slightly till the mid-1980s with a steep decline thereafter (cf. Figures 5(b) and 5(d)). The overall decline of Kohler-Philipov adjusted TFMR only amounts to 30 percentage points compared to about slightly more than 40 percentage points for the observed TFMR. Since the time series of first marriage intensities ends in 1998, the overall decline of the Kohler-Ortega adjusted total first marriage rate by 14 percentage points compares to an overall decline of the observed TFMR from 1970 to 1998 by about 38 percentage points. In the same period, the Kohler-Ortega adjusted PPEM decreases about by 18 percentage points compared to 30 percentage points of the observed PPEM.

Germany

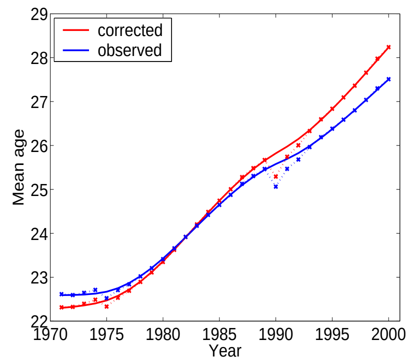
Unfortunately, we were not able to get separate data for former West and East Germany from 1990 onwards. Therefore the results for Germany shown in Figure 6 represent the West German situation in the 1970s and 1980s. From 1990 onwards the results are for West and East Germany combined. From a methodological point of view, the adjustment algorithm should not depend on past values of the time series. However, the smoothing algorithm seems to be sensitive to the initial and terminal value of the corresponding time series. Therefore one should be cautious in interpreting the German results, in particular the values around the unification in 1990.

Other than for Austria, we find a continuously increasing mean age of first marriage frequencies (cf. Figure 6(a)). When correcting for variance distortions, the mean age at first marriage would have been lower until about the early 1980s with only a slight increase. From the early 1980s onwards, the mean age at first marriage corrected for variance effects would have shown a more pronounced increase, which is due to the increase in variance over the same period of time. Summing up, the rate of change of the mean age at first marriage corrected for variance effects over time is greater than the rate of change of the observed mean age at first marriage over the whole investigation period. Hence, the Kohler-Philipov adjusted rates should lie above those adjusted according to Bongaarts and Feeney (1998).

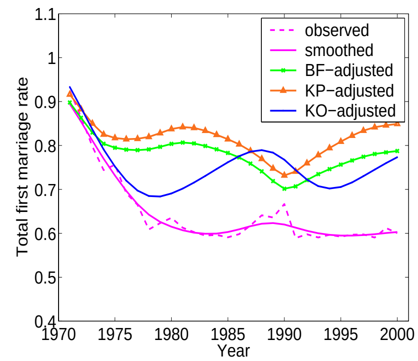
In contrast, the variance-corrected mean age of German first marriage intensities crosses the observed time series of the mean age twice (cf. Figure 6(c)). Contrary to the German first marriage frequencies, the mean age corrected for variance changes according to the Kohler-Ortega method increased at a slower rate than the observed mean age of the intensities during the 1970s. Therefore the Kohler-Ortega adjusted PPEM lies below the Bongaarts-Feeney adjusted PPEM during the 1970s and above it afterwards until the beginning of the 1990s. During the 1990s the adjusted values of the PPEM are almost identical for both methods since the observed and variance-corrected mean age of the intensities increased at nearly the same rate. Moreover, the difference to the frequencies is apparent by comparing the Kohler-Philipov adjusted TFMR to the Kohler-Ortega adjusted TFMR. Since the mean age corrected for variance changes according to the Kohler-Ortega method increased at a small rate during the 1970s, the corresponding adjusted TFMR decreases more steeply during this time. Furthermore, the steep rise in the Kohler-Ortega adjusted TFMR during the 1980s corresponds to an equally steep increase in the variance-corrected mean age of the intensities, while the fall of the Kohler-Ortega adjusted TFMR during the 1990s occurs when the mean age of the intensities corrected for variance changes increases at a lesser rate.

Finally, since both the observed and variance-corrected mean age of the intensities increase over the whole investigation period, the adjusted PPEM values lie above the observed values (cf. Figure 6(d)).

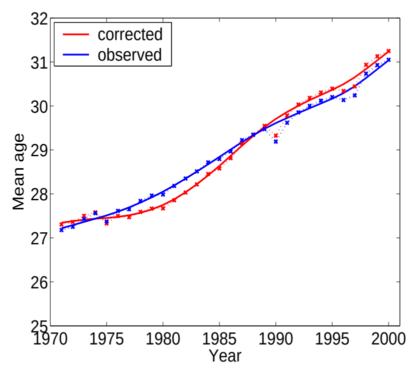
Adjusted for tempo, the period measures of first marriage rates show smaller declines in marriage than reported by official statistics over the 30 years of investigation. Most interestingly, we find that the adjusted total first marriage rates fluctuate heavily, with values ranging between about 0.7 and 0.85 (cf. Figures 6(b)). Therefore the decline over time since the beginning of the 1970s of the adjusted TFMR ranges between about 6 to 25 percentage points compared to about 30 percentage points of the observed TFMR. Simi-



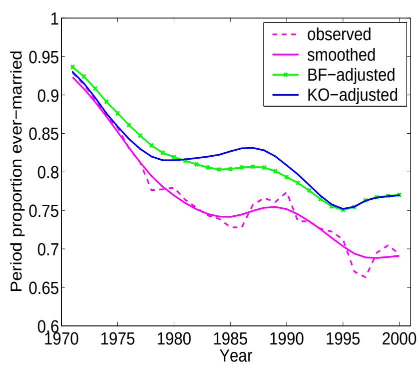
(a) Mean age of first marriage frequencies



(b) Total first marriage rate



(c) Mean age of first marriage intensities



(d) Period proportion ever married

Figure 6 Mean age of female first marriage for Germany, observed and corrected for variance changes, and German female period measures of first marriage, observed, smoothed, adjusted for changes of the mean age according to Bongaarts and Feeney (1998, BF), and adjusted for changes of the mean and variance according to Kohler and Philipov (2001, KP) as well as to Kohler and Ortega (2002, KO).

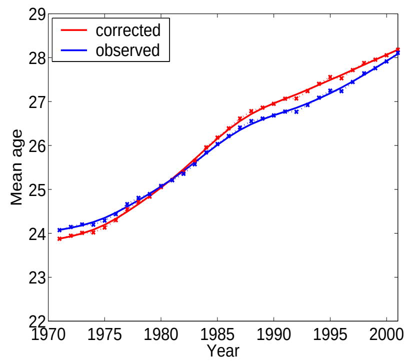
larly, we find fluctuations in the adjusted PPEM series but with a smaller amplitude and with a clear downward trend of the series (cf. 6(d)). Therefore the adjusted period proportion ever married declined by about 16 (Kohler-Ortega) or 17 (Bongaarts-Feeney) percentage points from 1971 to 2000 compared to about 23 percentage points for the observed PPEM series during this period.

Switzerland

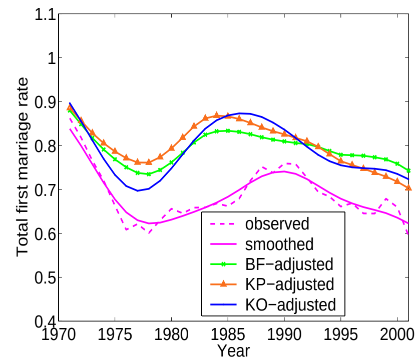
For Swiss first marriage rates, we find, similar to the German case, fluctuations in the adjusted series of TFMR and PPEM (cf. Figure 7(b) and Figure 7(d)). However, unlike the German case, these fluctuations are caused by oscillations in the observed TFMR and PPEM series and merely reinforced by the tempo adjustment, since the observed mean ages and the variance-corrected values of their frequencies and intensities show a steady increase over the whole investigation period (cf. Figure 7(a) and Figure 7(c)). The latter also implies that the adjusted series are above the observed TFMR and PPEM series. Since the increase in the mean ages is more pronounced in the 1980s than in the previous and following decades, there are stronger tempo effects in the 1980s (cf. Figure 7(b) and Figure 7(d)).

Regarding Swiss first marriage frequencies, we additionally find that the variance-corrected mean age of first marriages frequencies increases slightly more than the observed values during the 1970s and 1980s and vice versa during the 1990s. Hence, the Bongaarts-Feeney adjusted TFMR lies below the Kohler-Philipov adjusted TFMR during the first two decades and above it thereafter.

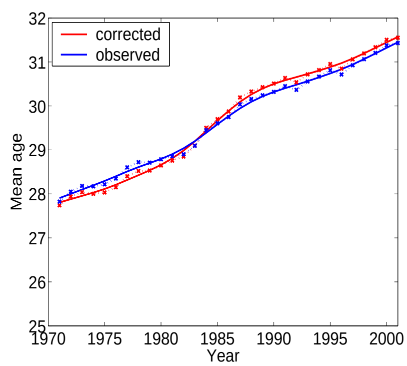
Concerning Swiss first marriage intensities, the observed mean ages and their variance-corrected values of Swiss first marriage intensities evolve almost identically over time in the 1970s and 1990s, though at different levels. Hence, the Bongaarts-Feeney adjusted and the Kohler-Ortega adjusted PPEM series are almost the same in that decades. Consequently, variance effects in the intensities seem to be absent during the 1970s and 1990s.



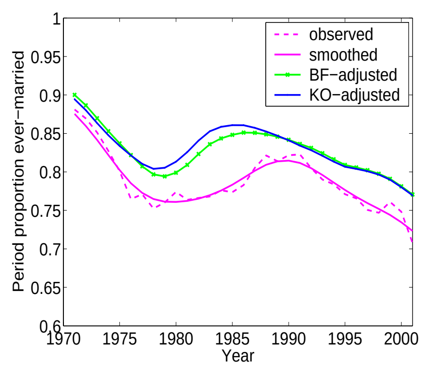
(a) Mean age of first marriage frequencies



(b) Total first marriage rate



(c) Mean age of first marriage intensities



(d) Period proportion ever married

Figure 7 Mean age of female first marriage for Switzerland, observed and corrected for variance changes, and Swiss female period measures of first marriage, observed, smoothed, adjusted for changes of the mean age according to Bongaarts and Feeney (1998, BF), and adjusted for changes of the mean and variance according to Kohler and Philipov (2001, KP) as well as to Kohler and Ortega (2002, KO).

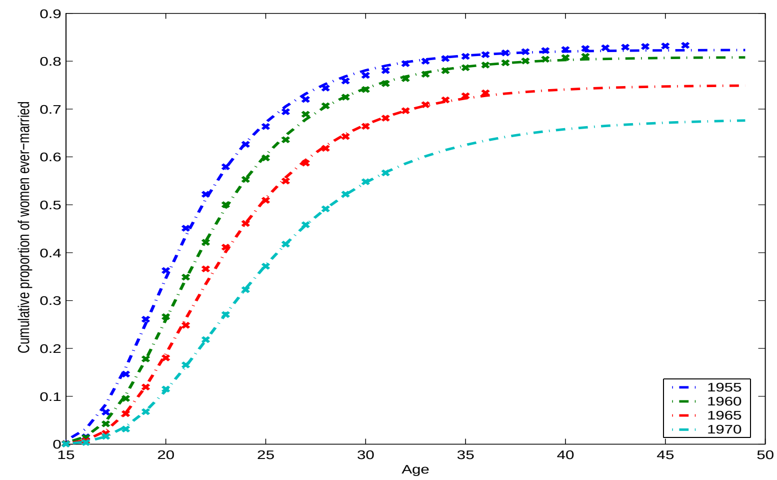
Comparison with Cohort Experience

The most convincing test of the adjustment procedures is perhaps to compare the period measures, adjusted for tempo and variance changes, of first marriage rates with the experience of cohorts, whose marriage proportions are, by definition, unaffected by tempo effects (Goldstein 2002).

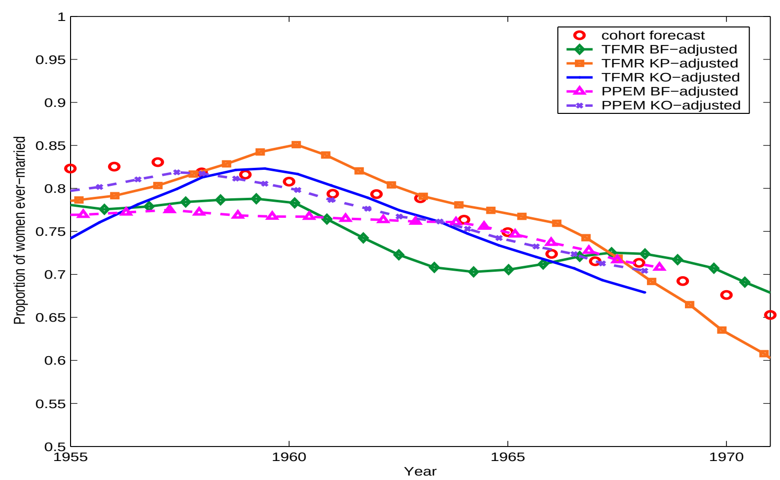
We estimated the completed cohort proportion women who ever married by extrapolating the observed experience of cohorts using the Coale-McNeil model (Coale and McNeil 1972).[10] This model has often been used to forecast cohort marriage rates (Goldstein and Kenney 2001, Bloom and Bennett 1990, Liang 2000). Figure 8(a) shows the cumulative proportions marrying by age of selected cohorts for Austria. The crosses symbolise observed values and the solid line show the fit of the Coale-McNeil model. We see that the predicted proportion marrying by age 50 is falling with each successive cohort. Though, the estimates for the younger cohorts are based on few observed data points. Moreover, it also appears that the Coale-McNeil model may be underestimating late marriage since the model estimates appear to be slightly less than the last observed proportions married for most cohorts.

Figure 8(b) shows the estimated Austrian cohort proportions ever married plotted together with the Austrian period estimates, adjusted for mean and variance changes, that were shown in Figure 5. We have shifted the period measures by their corresponding period mean age of first marriage, where first marriage rates, again adjusted for mean and variance changes, were shifted by the variance-corrected mean age values. The Kohler-Ortega adjusted measures of first marriage are closer to the cohort forecasts than the other adjustments. However, throughout the total 30 years of investigation, the maximum difference for the five period measures to the cohort proportion ever married

[10] In the Coale-McNeil model, the density of age at first marriage is given by $g(x) = .1946\Theta \exp\{-.174\,(x-6.06\kappa) - \exp[-.288\,(x-6.06\kappa)]\}$, where $x = \frac{a-a_0}{\kappa}$ (cf. Coale and McNeil 1972). In particular, a is the age at first marriage, a_0 is the age at which first marriage may start, κ scales the speed of entry into first marriage, and Θ is the proportion of the cohort that eventually marries. The model was estimated with the generalised least square method.



(a) Cumulative proportions of Austrian women ever married, by cohort



(b) Austrian period and cohort measures of first marriage

Figure 8 Comparison of Austrian period and cohort experience of first marriage.

amounts to only about 7 percentage points.

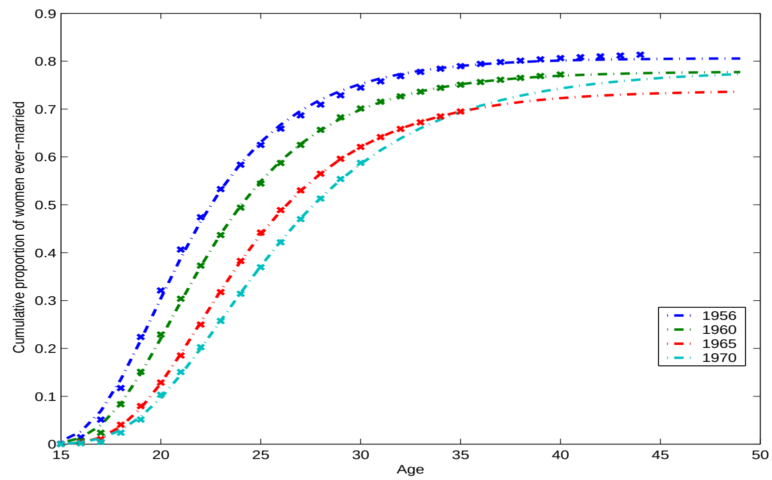
The corresponding German situation can be seen in Figure 9. As in the Austrian case, Figure 9(a) shows that the predicted proportion marrying by age 50 is falling with each successive cohort. Similar to the Austrian case, the Coale-McNeil model seems to underestimate the cumulative proportions of ever married women for the older cohorts, while the estimated proportion ever married by age 50 for the birth cohort 1970 seems to be too high. However, when comparing the estimated cohort proportion ever married to the adjusted period measures, we also find increasing values for the translated period series of the late 1960s birth cohorts. Furthermore, the Bongaarts-Feeney and the Kohler-Ortega adjusted period measures perform well in approximating the cohort forecasts for the earlier birth cohorts.

For the Swiss data, the Coale-McNeil extrapolation is given in Figure 10(a). Similar to the Austrian and German cases, we also find a decreasing cohort proportion ever married by age 50 across cohorts. This decline is well approximated by the Kohler-Ortega adjusted values of the translated total female first marriage rates, except of the early 1950s birth cohorts. For the latter the Bongaarts-Feeney adjusted series perform better. However, for all the adjusted period measures the deviation to the cohort forecasts only amounts to about 6 percentage points.

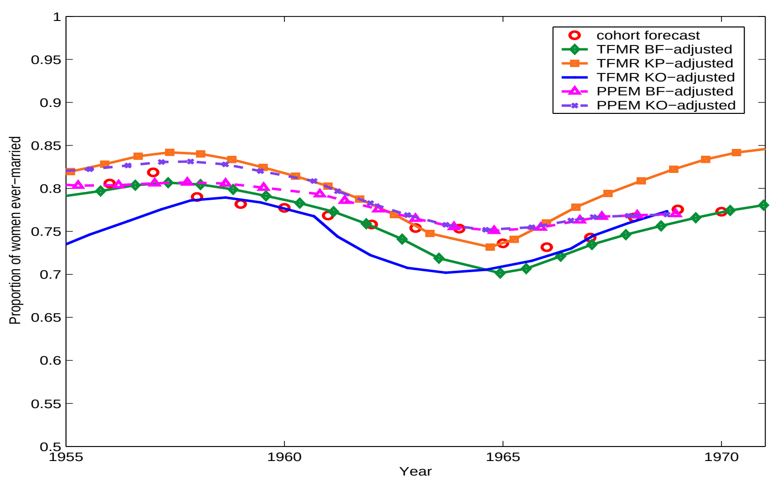
CONCLUSIONS

In this paper, we have quantified the changes in female first marriage rates for Austria, Germany, and Switzerland in terms of tempo and variance effects by applying methods developed for tempo and variance changes in fertility.

We find a uniform pattern of tempo changes across countries for the different decades. For instance, the 1970s were characterised by a slight increase or even a decrease of mean ages of frequencies or intensities regardless of correcting for variance changes. During the 1980s, postponement of first marriages was severe, since the mean ages steeply increased. Therefore the adjusted values of total first marriages rates as well as the period proportion

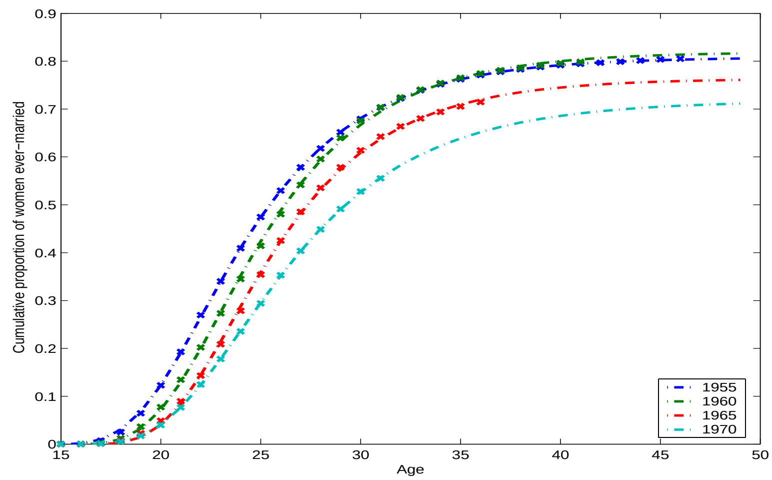


(a) Cumulative proportions of German women ever married, by cohort

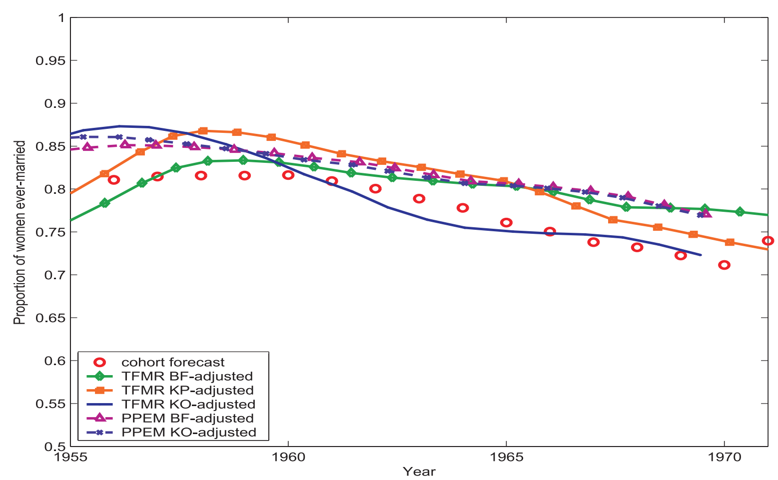


(b) German period and cohort measures of first marriage

Figure 9 Comparison of German period and cohort experience of first marriage.



(a) Cumulative proportions of Swiss women ever married, by cohort



(b) Swiss period and cohort measures of first marriage

Figure 10 Comparison of Swiss period and cohort experience of first marriage.

ever married sharply increase. Finally, the 1990s were characterised by further postponement but at a slower rate than in the previous decade.

Surprisingly, in the 1990s, tempo distortions of first marriage intensities caused by variance changes were absent in all three countries. In fact, the Bongaarts-Feeney adjusted and the Kohler-Ortega adjusted period proportion ever married showed almost the same values in the 1990s. Similarly, in the 1970s variance effects were very small in Austria, Germany, and Switzerland. Only during the 1980s, tempo distortions caused by variance changes were present, though they only amounted to three to four percentage points of women ever married. Sobotka (2003) similarly found that correcting for variance effects did not significantly change the values of the tempo-adjusted total fertility rate and the parity index of total fertility for data from the Czech Republic, Italy, the Netherlands, and Sweden.[11]

Considering variance effects in total first marriage rates, the size of the distortions heavily depends on the method used. This is due to two reasons. First, as stated earlier, the Kohler-Philipov method heavily depends on a technical parameter which can be arbitrarily chosen (see Footnote 8). Secondly, using the Kohler-Ortega method, the corresponding intensities are adjusted and then transformed into frequencies. From cohort comparison, we find that Kohler-Ortega adjusted values approximate the cohort forecasts quite well, though with significant deviations for the late 1950s birth cohorts in all three countries. Further investigations including a sensitivity analysis with respect to technical parameters are needed.

To what extent the results depend on the adjustment method used is also apparent by comparing the share of the overall decline in first marriage rates which is attributable to tempo distortions. In particular, the tempo-adjusted decline in total first marriage rates since 1970 is about 23 to 85 per cent of that published in the vital statistics in Austria, Germany and Switzerland, which implies a share of 15 to 77 per cent of the reduction in first mar-

[11]However, Sobotka (2003) did not iterate for $\gamma(t)$ and $\delta(t)$, which are defined as parameters from the adjusted frequency or intensity schedule, but computed them from observed values.

riage rates due to tempo distortions! Ignoring the Kohler-Philipov adjusted values because of the sensitivity to a technical parameter, the above range of tempo effects decreases to still as much as 20 to 64 per cent of the total decline. However, we also find significant variations of the share of tempo effects across countries, with Switzerland showing the lowest decline due to tempo distortions (20–37%). The tempo effects in the decline of the total first marriage rate in Austria and Germany are significantly higher, i.e., 33–64% for Austria and 45–62% for Germany. However, as evident from Figure 3, Switzerland started in the beginning of the 1970s with a substantially higher mean age at first marriage compared to Austria and Germany.

Since there were almost no variance effects in the 1970s and 1990s in the first marriage intensities, the decline in the tempo-adjusted period proportion ever married is nearly independent of the adjustment method used. In particular, the tempo effect amounts to 15 to 34% of the decline in the period proportion ever married since the 1970s in Austria, Germany and Switzerland. Within the countries, we find again Switzerland exhibiting the lowest tempo (15–18%) and Germany and Austria showing much higher values (28–31% and 34%, respectively). Summing up, the proportion of Austrian, German and Swiss women who will ever marry indeed appears to be declining, but a bit less dramatically than period measures would indicate.

However, the rates declined at different paces in different decades, and there is some evidence that the share of the reduction of first marriage rates which can be attributed either to quantum or tempo effects has changed over the three decades. In particular, for the first half of the 1970s the decline in first marriage rates is mainly attributable to quantum changes since the tempo-adjusted series of first marriage rates shows an almost parallel shrinkage. In the late 1970s and during the 1980s, severe postponement took place, due to increasing tempo-adjusted series of first marriage rates. However, in the 1990s, the observed and tempo-adjusted series of first marriage rates declined again. These decade-specific characterisation of tempo and quantum effects are in line with the observed stages of the diffusion process of pre-marital unions (Kiernan 2001, Prinz 1995). In the first stage, cohabitation is a deviant behaviour practised only by a small group of people. At the second stage, cohabitations represent predominately a probationary period or a prelude to marriage. This was certainly the case in the late 1970s and in the 1980s. Indeed, 45% of the Austrian marriages in 1989, the bride and the groom were co-residing before marriage (Statistik Austria 2002). In the last stages, cohabitation becomes socially acceptable as an alternative to marriage, where cohabitation ends up being almost indistinguishable from marriage (Kiernan 2001). In fact, Kiernan (2000) finds for data from the European Fertility and Family Survey, which was conducted in the mid to late 1990s, that "cohabitation typically initiates first union and about 30-40% of the first unions were cohabitations that had converted into a marriage with the same partner."

Nevertheless, Kiernan (2000) finds for Austrian, German and Swiss FFS data that 52–68% of the 20–39 year old women with a first partnership were married, which goes along with our finding that the majority of women in the German speaking countries are still marrying once in their lifetimes. However, in Austria the forecast proportion choosing never to marry has risen from 18 per cent for cohorts born in 1955 to an estimated 32 per cent for those born in 1970. Germany and Switzerland exhibit somewhat smaller proportions, i.e., 19 per cent of the 1955 birth cohort to an estimated 26 per cent of women born in 1970 in Switzerland, and 19 per cent to 23 for the 1956 and 1970 birth cohorts in Germany, respectively (cf. Figures 8(b), 9(b), and 10(b)).[12] The tempo findings for Austria and Germany are in line with Goldstein's findings for France. Goldstein (2002) reports an increase in the proportion women choosing never to marry from about 10 per cent for cohorts born just after World War II to 20 to 30 per cent for the cohorts born in the 1960s. This compares to a level near 50 per cent, as implied by the observed (unadjusted) period measures of marriage.

Summing up, the extent of the decline in first marriage which is attributable to tempo effects varies from country to country. Furthermore, in

[12] As stated earlier, the Coale-McNeil model seems to overestimate the cohort proportion ever married for the younger birth cohorts in Germany. For instance, the 1966 birth cohort already exhibits a estimated share of 27% choosing never to marry.

none of the analysed countries, the decline in first marriage rates can predominately be explained by postponement. Our empirical results suggest that at the beginning of the 21st century German-speaking Europeans ever get married at least thirty per cent less often than three decades ago. Thus, regarding the theoretical debate in the introduction, we do not find empirical support for the hypothesis that the observed reductions are solely due to tempo distortions. The extent of these distortions may substantially vary by socioeconomic status. Therefore a more differentiated micro-level analysis for various socioeconomic subgroups of the population would be required for a deeper understanding of the quantum and tempo of first marriages.

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