

Decomposing changes in first birth trends: Quantum, timing, or variance

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Abstract

In high-income countries, women and men born since the 1940s have delayed the birth of their first child, more of them have remained childless, and the timing of the first birth has become more diverse in these cohorts. The interaction between these three trends makes the research on first birth patterns more complex. This study has two main aims: (1) we introduce an alternative index, Expected Years Without Children (EYWC), to quantify changes in first birth behaviour; and (2) we decompose the changes in EYWC over time into three effects: remaining permanently childless, postponing the first birth, and the expansion of the standard deviation of the mean age at first birth. Using data from the Human Fertility Database, EYWC is calculated to illustrate time trends among women born in the 1910s–1960s in eight countries with longer series of data on cohort first birth trends: Canada, the Czech Republic, Japan, the Netherlands, Norway, Portugal, Sweden, and the United States. Our decomposition shows that the changes in EYWC are mainly attributable to postponement in North America and northern Europe, whereas these changes are largely due to increasing shares of women remaining childless in Japan and Portugal.

Keywords: childlessness; postponement of first birth; decomposition method; Coale-McNeil model; life table

1 Introduction

The share of childless women has steadily increased in recent decades in high-income countries. For example, according to [Kreyenfeld and Konietzka \(2017\)](#), more than 20% of women in German-speaking countries were childless at the end

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of their reproductive period since the 1950s birth cohorts. This increasing trend can be observed throughout Europe (Beaujouan et al. 2016; Miettinen et al. 2015; Kreyenfeld and Konietzka 2017), North America (the US: Frejka 2017, Canada: Ravanera and Beaujot 2014) and Japan (Raymo et al. 2015). It has also been projected that the future level of childlessness will increase as the fertility patterns of countries diverge (Sobotka 2017). Becoming a parent or remaining childless influences various aspects of a man's or a woman's life, including the individual's income (Budig et al. 2012), health (Kendig et al. 2007), old-age well-being (Huijts et al. 2013; Zhang and Hayward 2001), and support networks (Albertini and Kohli 2009). Therefore, it is important that we gain a deeper understanding of the trend of childlessness. The increase in the number of childless women has coincided with an increase in the mean age at first birth. For instance, in the OECD countries, the mean age at first birth has increased by about one year each decade since the 1970s (Mills et al. 2011). Thus, in these decades, levels of childlessness have risen (*quantum* changes) and the first birth has been increasingly postponed (*timing* changes).

The most common indices used to measure first birth trends, the proportion of childless women at age 50 and the mean age at first birth, are useful for studying either quantum changes or timing changes, but not for studying both. This is simply because the proportion of childless women does not take into account the timing changes (e.g., the same proportion of childless women can reflect different mean ages at first birth), and the mean age at first birth per se does not show quantum changes. Previous research has investigated how changes in the timing of the first birth affect the level of childlessness. For instance, Kneale and Joshi (2008) explored the extent of childbirth postponement and projected its impact on eventual levels of childlessness in the UK. Similarly, te Velde et al. (2012) estimated the effect of postponement of entry into motherhood on permanent involuntary childlessness in six European countries.

In this study, by contrast, we suggest an alternative index that takes both quantum and timing changes into account in a single index: Expected Years Without Children (EYWC). This index measures the average length of life women spend without children during their reproductive period by estimating both the life years of women who remained childless at the end of their reproductive period (quantum) and the life years of women before they had a child (timing). The concept of EYWC was previously introduced by Bongaarts and Feeney (2006) in fertility research, and was employed by Andersson et al. (2017) to describe contemporary family changes using a life table method. However, these authors did not focus on the index per se, nor did they depict its trends. Therefore, the first aim of this study is to present the trend in EYWC in order to show how first birth trends have changed due to both quantum and timing changes.

Our second aim is to quantify which of two effects, remaining permanently childless or postponing the first birth, has the most impact on changes in first birth behaviour. Although those two phenomena are the main focus in this study, an additional component, the variance in the age at first birth, is investigated. When the mean age at first birth increases, the variance in the timing of the first birth in a

given population becomes larger (Kohler and Philipov 2001). In part because of this expansion of the timing of the first birth, the family life courses of women diverge more across birth cohorts (Elzinga and Liefbroer 2007). Therefore, we quantify the effects of remaining permanently childless, postponing the first birth, and the expansion of the timing of the first birth on the changes in EYWC over time using a decomposition method presented by Mogi and Canudas-Romo (2018).

In the following section, we explain our methods – a parametric model for a first birth, the main measure used to describe the first birth trend (EYWC), and the decomposition method – as well as the data we used. In the third section, we present the trend in EYWC and the results of its decomposition. In the final section, we conclude by discussing our findings, the limitations of our study, and future research directions.

2 Data and methods

2.1 Data

In this study, we used female population counts by age and birth cohort and birth counts by birth order and mother's cohort from the Human Fertility Database (HFD) (Human Fertility Database 2018). To avoid problems arising from truncation and censoring bias, we used completed cohort fertility data only. Based on these criteria, data from eight countries were selected: Canada (1929–1962 birth cohort), the Czech Republic (1935–1965 birth cohort), Japan (1953–1965 birth cohort), the Netherlands (1935–1963 birth cohort), Norway (1952–1965 birth cohort), Portugal (1944–1966 birth cohort), Sweden (1955–1965 birth cohort), and the US (1918–1965 birth cohort).

2.2 Coale-McNeil model for first birth

While the parametric model for overall fertility is well developed (see, for example, Kostaki and Paraskevi 2007), there are few models that measure the first birth. In this study, we use the Coale-McNeil model (CM model) to estimate the age-specific first birth rate. The parameters of the CM model have conventional demographic meanings, and fit best with the aim of our study, which is to quantify the effects of remaining permanently childless and postponing fertility.

The CM model was developed to estimate the age at first marriage patterns of a birth cohort (Coale and Trussell 1978). The CM model has since been extended beyond its original purpose, and has often been applied to the first birth distribution by age (Bloom 1982*a,b*; Bloom and Trussell 1984; Henz and Huinink 1999; Rao 1987; Trussell and Bloom 1983). Previous studies successfully applied the CM model to various countries, including Canada, Columbia, Finland, Germany, Italy, and the US. The advantage of using the CM model is that its parameters have clear demographic meanings, which are as follows. The probability of first birth at age x

and time t , denoted as $f_{x,t}$, is expressed as:

$$f_{x,t} = C_t \frac{1}{\sigma_t} a_1 \exp \left[a_2 \left(\frac{x - \mu_t}{\sigma_t} + a_3 \right) - \exp \left\{ -a_4 \left(\frac{x - \mu_t}{\sigma_t} + a_3 \right) \right\} \right], \quad (1)$$

where C_t pertains to the proportion of the cohort eventually having a child by age 50 at time t , μ_t refers to the mean age at first birth at time t , σ_t is a measure of the standard deviation of age at first birth at time t , and the usual values for the constants are $a_1 = 1.281$, $a_2 = -1.145$, $a_3 = 0.805$, and $a_4 = 1.896$. This equation is known as a standardised version of the CM model developed by [Rodríguez and Trussell \(1980\)](#).

Although several models have been used to estimate first birth patterns, these models are not fully applicable to our study. First, for one feature of the CM model (the convolution structure), an additional term has been proposed: namely, an exponentially distributed waiting time segment to account for the time from first marriage to first birth. However, [Trussell and Bloom \(1983\)](#) reported that the original CM model fitted better than the model with the waiting term. Thus, previous research applied the CM model to first birth patterns without using the additional waiting time segment ([Bloom 1982a,b](#); [Bloom and Trussell 1984](#), [Henz and Huinink 1999](#), [Rao 1987](#), [Trussell and Bloom 1983](#)). In addition, this approach is not theoretically appropriate given the current data. As nonmarital childbirth has become common in many high-income countries ([Eurostat 2018](#); [Department of Health and Human Services 2018](#)), the first birth does not always occur after marriage. Thus, the additional term that considers the period from marriage to childbirth is not relevant to the data in our study. The second model is a log-logistic function (LL model). [Henz and Huinink \(1999\)](#) applied it to the first birth in the German data. However, it does not have a scale parameter, which indicates the proportion of the population who never had children. Hence, it is not applicable given our aim and decomposition method.

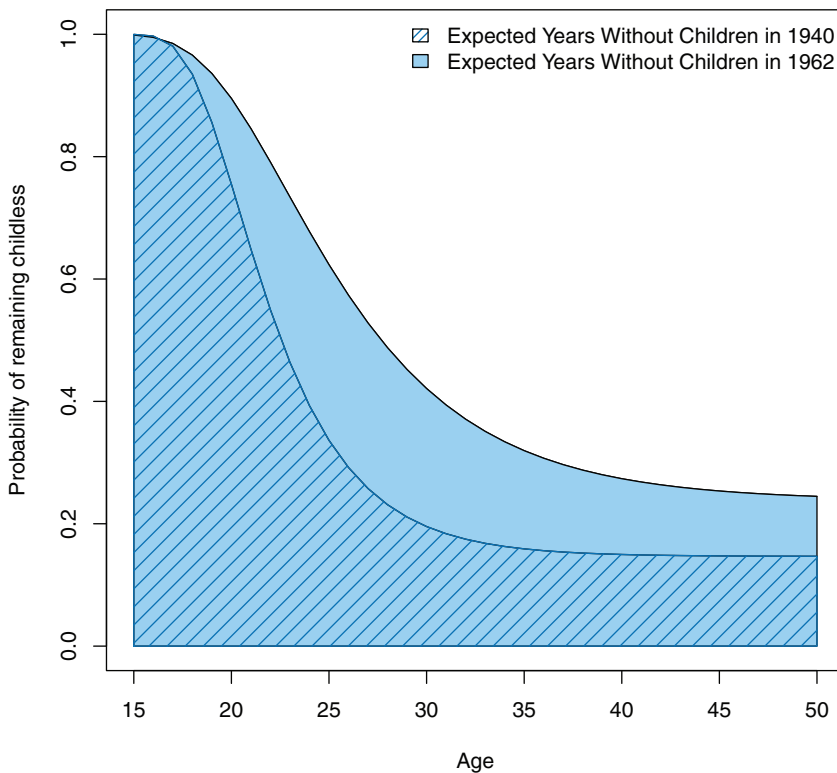
Although the CM model is widely used to study first birth trends, whether the CM model fits well with the observed pattern of the age at first birth has not been statistically examined. The goodness-of-fit using the Kolmogorov-Smirnov tests shows that the CM model estimates well the data for all the countries and years (see [Appendix A](#)). For these reasons, we have chosen to use the CM model to explore the changes over time in first birth behaviours and the expected years without children (EYWC).

However, the limitations of the CM model should be mentioned. There are statistical issues when the CM model is used to estimate the mean and the standard deviation of age at first birth ([Bloom and Trussell 1984](#)): (1) if the sample available for estimation is restricted to women who have become mothers on or before the survey date, a truncation bias problem will arise; and (2) if the data used for estimation refer to a sample of all women, there will be a censoring problem if any of the women who will ultimately have a first birth have not done so by the time of the survey ([Bloom and Trussell 1984](#)). We avoid those problems by using completed cohort fertility data from the HFD.

2.3 Expected years without children (EYWC)

We use life expectancy to measure fertility behaviours. In classical life table methods, life expectancy between two ages (e.g., 0 and X) represents the area below a survival function from age zero to the fixed age X . This is interpreted as the average number of years people live between these ages (Preston et al. 2001). In this research analysing first birth trends, life expectancy is interpreted as the expected years without children, with a first birth representing a “death”. The minimum age is the age of first menstruation and the maximum age is the age of menopause. Therefore, the EYWC from age 15 to age 50, denoted as ${}_{35}e_{15}$, is calculated as ${}_{35}e_{15}(t) = \int_{15}^{50} l_{x,t} dx$. It corresponds to the two shaded areas in Figure 1 for the

Figure 1:
Probability of remaining childless by age among a Canadian female birth cohort in 1940 and 1962



Note: The parameters of the probabilities of remaining childless are $C = 0.85$, $\mu = 23.10$, and $\sigma = 4.14$ for the 1940 birth cohort and $C = 0.76$, $\mu = 26.98$, and $\sigma = 6.88$ for the 1962 birth cohort.

Source: Authors' calculations using the Human Fertility Database.

1940 and the 1962 birth cohorts. The use of EYWC has two primary advantages. First, as this index is able to capture fertility trends in one simple value, it can be used to numerically compare fertility trends at different times and in different countries. Second, it can take into account both the population having children and the population without children. While the most common indices used to study first birth trends are the mean age at first birth and the proportion of childless women, the EYWC index can reflect those two indices simultaneously. These features are very useful in the current context, in which both indices have an increasing trend.

2.4 Decomposition method

We applied the decomposition method developed by [Mogi and Canudas-Romo \(2018\)](#). A detailed explanation can be found in their study ([Mogi and Canudas-Romo 2018](#)). The changes in EYWC over time, denoted as ${}_{35}\dot{e}_{15}(t)$, are decomposed into three parameters: scale (the proportion of the cohort eventually having a child), location (the mean age at first birth), and variance (the standard deviation of age at first birth). The decomposition of ${}_{35}\dot{e}_{15}(t)$ can be formulated as

$${}_{35}\dot{e}_{15}(t) = \frac{\partial {}_{35}e_{15}(t)}{\partial C_t} \dot{C}_t + \frac{\partial {}_{35}e_{15}(t)}{\partial \mu_t} \dot{\mu}_t + \frac{\partial {}_{35}e_{15}(t)}{\partial \sigma_t} \dot{\sigma}_t, \quad (2)$$

where each term is the change in ${}_{35}\dot{e}_{15}(t)$ resulting from changes in the scale, location, and variance. A dot on top of a variable indicates the derivative with respect to time. The largest value among these three components shows that the changes in EYWC are mainly caused by that factor; i.e., scale factor ($\frac{\partial {}_{35}e_{15}(t)}{\partial C_t} \dot{C}_t$): remaining permanently childless, location ($\frac{\partial {}_{35}e_{15}(t)}{\partial \mu_t} \dot{\mu}_t$): postponing first birth, or variance ($\frac{\partial {}_{35}e_{15}(t)}{\partial \sigma_t} \dot{\sigma}_t$): expansion effect. Appendix C explains the detailed method for estimating the three parameters of the CM model and applying the decomposition equation to discrete data.

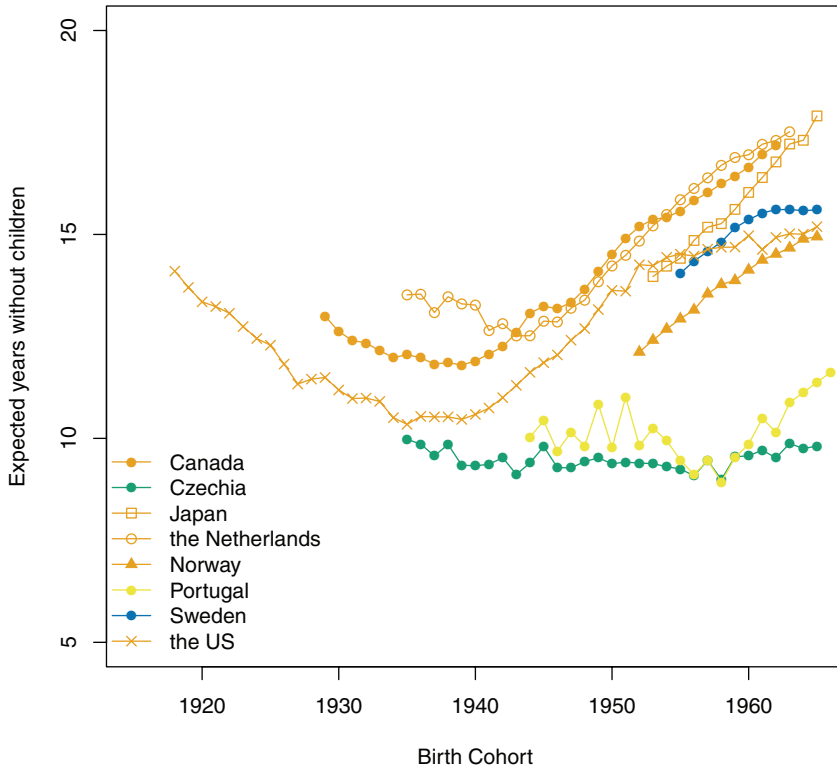
3 Results

3.1 Cross-country trend analysis of EYWC

The trends in EYWC for the selected countries are presented in Figure 2.

Figure 2 shows that most of the countries in the study had an increasing trend in EYWC starting with the 1940s birth cohorts. The EYWC in Canada declined until the 1940s birth cohorts, and then started steeply increasing. The 1940 birth cohort in Canada had 12 years of EYWC. This value had increased to 16.6 years for the 1960 cohort. As EYWC represents the expected number of years without children starting from age 15, each number signifies that the expected age at first birth in the specific cohort in a country is 15+ EYWC. Thus, we can interpret these

Figure 2:
Time trends in expected years without children in eight countries



Source: Authors' calculations using the Human Fertility Database.

findings as showing that the expected age at first birth increased from 27 (15 + 12) in the 1940 birth cohort to 31.6 (15 + 16.6) in the 1960 birth cohort. The US had an increasing trend similar to that of Canada until the 1950s birth cohorts. In the US, the EYWC increased sharply from the 1940 birth cohort (10.5 years) to the 1950 birth cohort (13.6 years). However, this increasing trend slowed after the 1950 birth cohort. The EYWC in the US increased by only 1.3 years from the 1950 to the 1960 birth cohort. In the Netherlands, the EYWC increased starting with the 1945 birth cohort, from 13 years for the 1945 birth cohort to 17 years for the 1960 birth cohort. Moreover, in several other countries, such as in Japan and Norway, the EYWC displayed an increasing trend for all the observable periods. In Japan, the EYWC rose from 14 years for the 1953 birth cohort to 18 years for the 1965 birth cohort. In Norway, the EYWC increased from 12 years for the 1952 cohort to 15 years for the 1965 cohort.

The EYWC trends in Portugal, Sweden, and the Czech Republic differ from those in the countries discussed above. In Portugal, the EYWC fluctuated by

approximately 10 years between the 1940s and 1950s birth cohorts. However, by the 1960 birth cohort, the EYWC in Portugal had increased, and it is currently estimated at 11.6 years for the latest observed birth cohort. By contrast, in the Czech Republic, the EYWC trend has been neither increasing nor decreasing, as it has plateaued at 10 years for all observable birth cohorts from 1935 to 1965. Interestingly, in Sweden, the EYWC trend has also plateaued in recent cohorts. After increasing starting with the 1955 birth cohort (14.4 years), the EYWC trend had plateaued at 16 years by the 1960 birth cohort.

Among the latest birth cohorts we can observe, the EYWC was approximately 17.5 years in Canada, Japan, and the Netherlands; and was approximately 15 years in Norway, Sweden, and the US. The EYWC values in these countries were followed by those in Portugal, at 11.5 years; and in the Czech Republic, at less than 10 years. Lower EYWC values indicate that, on average, the women in the population gave birth to their first child at an early age. The latest cohort observed in Canada, Japan, and the Netherlands spent half of their reproductive period with no children (the expected age at first birth is $15 + 17.5 = 32.5$). In other words, these women have only have 17.5 years left in their reproductive period, on average, to have any subsequent children.

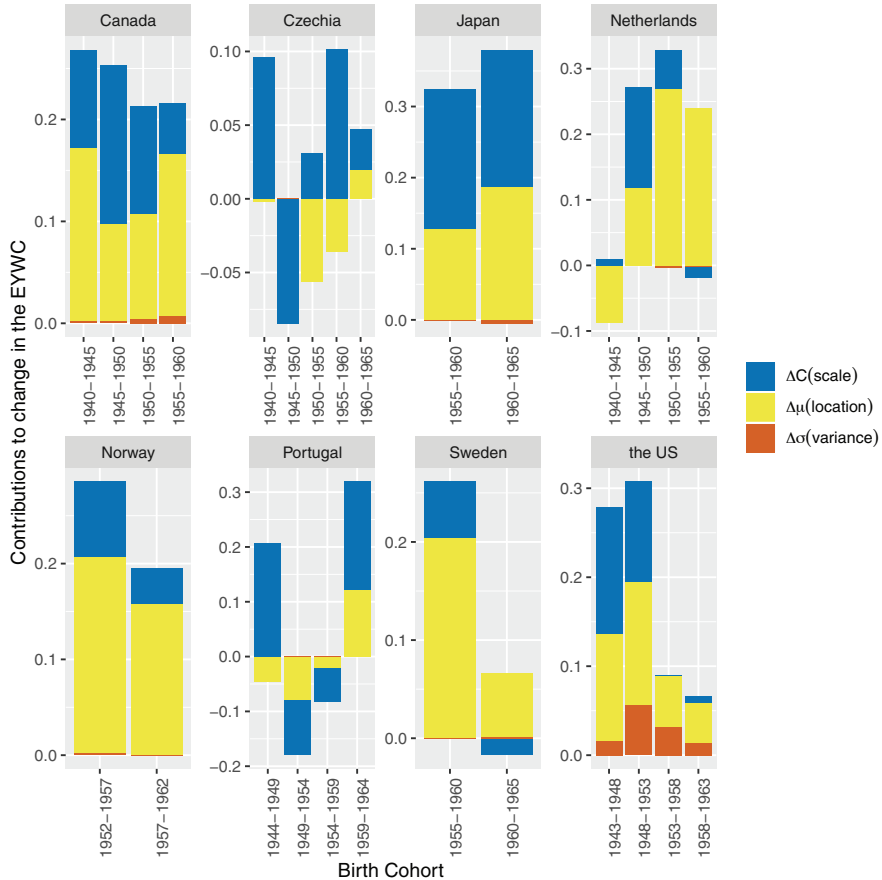
The main objective of this study is to investigate whether the changes in EYWC, as shown in Figure 2, are attributable to women remaining permanently childless or postponing childbirth, or to the expansion effect. We decompose EYWC from 1940 onwards; the results are presented in Figure 3.

3.2 Decomposition of EYWC

In Canada, the US, the Netherlands, Sweden, and Norway, the location parameter (the timing of first birth) was the most influential factor in the changes in EYWC after 1950. Table 1 shows that 74% of the increase in the EYWC in Canada from 1955 to 1960 was due to the increase in the average age at first birth. Similar results can be reported for the US: 68% of the increase in the EYWC from 1958 to 1963 in the US can be attributed to the increase in the average age at first birth. Likewise, in Norway, the increase in the average age at first birth was responsible for 81% of the increase in the EYWC from 1957 to 1962. The results for the Netherlands and Sweden were very similar. These findings suggest that larger shares of the current birth cohorts in these countries have postponed childbirth rather than remaining permanently childless. The changes in the scale parameter and the proportion of women having children influenced the current increase in the EYWC of these countries to a lesser extent.

In addition, in the Netherlands (1955–1960) and Sweden (1960–1965), the value for the scale parameter in the most recent birth cohort was negative, which implies that the proportions of women having a first child increased in those periods. The negative value of the scale factor in Sweden is consistent with the current decrease in the proportion of childless women (Persson 2010; Miettinen et al. 2015). The higher

Figure 3:
Decomposition of the change over time in females' expected years without children in eight countries



Note: scale: the proportion of the cohort eventually having a child; location: the mean age at first birth; and variance: the standard deviation of age at first birth.

Source: Authors' calculations using the Human Fertility Database.

educational category has a more distinct decreasing trend (Persson 2010). Therefore, this negative value could be mainly driven by women in a higher educational group.

However, the trends in Japan and Portugal differ from those in the other countries. The scale parameter, which indicates the proportion of women who eventually have a child, is the most influential factor in the changes in the EYWC in these countries after 1955 and 1959, respectively. For instance, Table 1 shows that 62% of the change in the EYWC in Portugal from 1959 to 1964 can be attributed to the scale parameter, which means that the shift was mainly attributable to women remaining

permanently childless. Likewise, in Japan, the increase in EYWC was mainly due to the changes in the scale parameter, even though the location parameter had only approximately half as much impact on the latest change. This result indicates that in these countries, more women remained permanently childless than postponed

Table 1:
The contribution of three parameters (scale: the proportion of the cohort eventually having a child, location: the mean age at first birth, and variance: the standard deviation of age at first birth) to the changes in females' expected years without children over time (${}_{35}e_{15}(t)$) in eight selected countries

Country	Birth cohort	${}_{35}e_{15}(t)$	Scale	Location	Variance	Sum of all components
Canada	1940–1945	0.2685	0.0957 (35.66)	0.1698 (63.26)	0.0029 (1.08)	0.2684
	1945–1950	0.2535	0.1558 (61.46)	0.0951 (37.51)	0.0026 (1.03)	0.2535
	1950–1955	0.2129	0.1059 (49.74)	0.1027 (48.24)	0.0043 (2.02)	0.2129
	1955–1960	0.2164	0.0497 (22.97)	0.1597 (73.80)	0.0070 (3.23)	0.2164
Czech Republic	1940–1945	0.0943	0.0963	-0.0020	0.0000	0.0943
	1945–1950	-0.0845	-0.0844	-0.0001	0.0000	0.0845
	1950–1955	-0.0253	0.0311	-0.0565	0.0000	-0.0254
	1955–1960	0.0651	0.1012	-0.0361	0.0000	0.0651
	1960–1965	0.0473	0.0278 (58.77)	0.0195 (41.23)	0.0000 (0.00)	0.0473
Japan	1955–1960	0.3232	0.1968	0.1276	-0.0011	0.3232
	1960–1965	0.3749	0.1923	0.1875	-0.0048	0.3750
Netherlands	1940–1945	-0.0789	0.0094	-0.0883	0.0000	-0.0789
	1945–1950	0.2709	0.1536	0.1186	-0.0013	0.2709
	1950–1955	0.3247	0.0594	0.2695	-0.0041	0.3249
	1955–1960	0.2205	-0.0158	0.2395	-0.0031	0.2206
Norway	1952–1957	0.2856	0.0790 (27.66)	0.2045 (71.60)	0.0021 (0.74)	0.2856
	1957–1962	0.1951	0.0368 (18.85)	0.1582 (81.05)	0.0002 (0.10)	0.1952
Portugal	1944–1949	0.1611	0.2075	-0.0462	0.0003	0.1615
	1949–1954	-0.1796	-0.0996	-0.0803	0.0003	-0.1796
	1954–1959	-0.0823	-0.0606	-0.0216	0.0001	-0.0824
	1959–1964	0.3192	0.1980 (62.03)	0.1205 (37.75)	0.0007 (0.22)	0.3192
Sweden	1955–1960	0.2611	0.0578	0.2037	-0.0002	0.2613
	1960–1965	0.0485	-0.0173	0.0643	0.0015	0.0485

Continued

Table 1:
Continued

Country	Birth cohort	${}_{35}e_{15}(t)$	Scale	Location	Variance	Sum of all components
US	1943–1948	0.2785	0.1422 (51.08)	0.1202 (43.18)	0.0160 (5.75)	0.2784
	1948–1953	0.3082	0.1131 (36.78)	0.1376 (44.75)	0.0568 (18.47)	0.3075
	1953–1958	0.0902	0.0007 (0.78)	0.0578 (64.08)	0.0318 (35.25)	0.0902
	1958–1963	0.0667	0.0077 (11.54)	0.0455 (68.22)	0.0136 (20.39)	0.0667

Note: Percentages are presented under each value in parentheses and are calculated only when all terms go in the same direction. The sum of all components (scale, location, and variance) varies slightly from the difference in the expected years without children (${}_{35}e_{15}(t)$) due to rounding the numbers to the third decimal point in the table.

Source: Authors' calculations using the Human Fertility Database.

their first birth to a later age. For Japan, the strong linkage between marriage and childbearing may be the key to understanding the large impact of the scale factor. As the nonmarital birth rate in Japan is still at a low level – i.e., 2.29% in 2015 ([National Institute of Population and Social Security Research 2017](#)) – most births in that country occur in marital unions. Hence, the large influence of the scale factor in Japan may indicate that the never-married population in Japan is increasing. Indeed, between 1980 and 2015, the never-married population at age 50 increased from 2.6% to 23.4% for males and from 4.5% to 14.1% for females ([National Institute of Population and Social Security Research 2017](#)). In Japan, the shares of never-married people are almost equal to the shares of childless people. This explains why the scale factor plays a larger role in Japan than in the other countries. Interpreting the findings for the Czech Republic is difficult because its EYWC trend fluctuated.

The changes in the variance parameter were not found to be influential in the changes in the EYWC in any of the countries in our sample except the US. The much larger variance effect observed in the US may be attributable to the strong differentials in first birth behaviours by indicators of socio-economic status, such as educational level, union status, and race/ethnicity. [Rendall et al. \(2010\)](#) found that in the US, there has been a persistent trend of early first births among low-educated women, and a shift towards later first births among women with higher educational levels. Moreover, [Chandola et al. \(2002\)](#) suggested that this heterogeneity is related to marital status, with an early bulge linked to extra-marital births, often among solo mothers. They also argued that ethnic differences play an important role in explaining the variance in the first birth trends in the US. Likewise, [Kostaki and Paraskevi \(2007\)](#) linked the observed heterogeneity in first birth patterns in the US

to a range of fertility determinants, including the rise of migrant populations in conjunction with racial and ethnic differences.

Overall, the results strongly indicate that changes in EYWC are mainly due to two factors: remaining permanently childless and postponing childbirth. Except in the US, the variability in the timing of childbirth is generally not found to be a factor in the changes in EYWC.

4 Conclusion

Increases in the proportion of childless women, the mean age at first birth, and the variance in the timing of the first birth have been observed in high-income countries (Kohler and Philipov 2001; Kreyenfeld and Konietzka 2017; Mills et al. 2011). These increases have led researchers to ask an important question: Which of three effects – remaining permanently childless, postponing the first birth, or the expansion of the standard deviation of the mean age at first birth – has the most impact on the changes in the first birth trend?

We used cohort data for eight high-income countries from the Human Fertility Database (HFD) and applied the decomposition method developed by Mogi and Canudas-Romo (2018) to quantify the effect of the aforementioned components on the changes in Expected Years Without Children (EYWC) over time. Our analysis of the trends showed that the EYWC in Canada, the US, the Netherlands, Japan, and Norway increased steadily across cohorts. In Sweden, EYWC increased until the 1960 birth cohort, and then remained at the same level. In Portugal, the EYWC increased starting with the late 1950s cohorts after fluctuating. In the Czech Republic, no clear EYWC trends across cohorts could be observed. The decomposition results strongly indicate that changes in EYWC are mainly due to two factors: remaining permanently childless and the postponement of childbirth. The findings of the decomposition analysis show that in Canada, the US, the Netherlands, Sweden, and Norway, more people postpone the first birth to a higher age than remain permanently childless. In these countries, the EYWC has an increasing trend mainly because their female populations have delayed childbirth. This could be the result of improvements in female education and labour force participation and better access to effective contraceptive methods. By contrast, in Japan and Portugal, more women remain permanently childless than postpone childbirth. The strong link between marriage and childbirth in Japan may be an important key to understanding the large impact of the scale factor in that country (National Institute of Population and Social Security Research 2017). For all periods and for all of the countries analysed except the US, the variance factor did not have a large impact on the changes in EYWC. This might be due to the greater heterogeneity in the first birth trend by socio-economic status in the US compared to in other countries (Chandola et al. 2002; Kostaki and Paraskevi 2007; Rendall et al. 2010).

Finally, in future research, this decomposition method could be applied to subpopulations, such as educational groups. It is well known that the timing of first

birth and the share of childless women differ significantly by education (Beaujouan et al. 2016; Kneale and Joshi 2008). Applying this method to union status (e.g., single, cohabitation, married) and race/ethnicity would also be beneficial.

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Appendix A: The statistical test of the CM model

We conducted a statistical examination of the CM model using the Kolmogorov-Smirnov test (KS test). The KS test is a nonparametric test to check the goodness-of-fit of the observed distribution and the estimated one by the CM model. The test statistic D of the KS test quantifies the supremum distance between the empirical distribution function of the data and the cumulative distribution function of the CM distribution. The null distribution of this statistic is calculated under the null hypothesis that the data follow the CM distribution. Thus, a p -value greater than $\alpha = 0.05$ indicates that the data and the CM model have a good fit. As Table A.1 shows, the CM model estimated the observed data statistically well for all countries and birth cohorts.

Table A.1:
Goodness-of-fit of the CM model using the Kolmogorov-Smirnov test

Country	Birth cohort	D	P-value	Country	Birth cohort	D	P-value	
Canada	1940	0.1429	0.8745	Norway	1952	0.1429	0.8674	
	1945	0.1714	0.6902		1957	0.2000	0.4858	
	1950	0.1714	0.6902		1962	0.1714	0.6902	
	Czech Republic	1955	0.1714	0.6902	Portugal	1944	0.2000	0.4916
		1960	0.1714	0.6902		1949	0.1429	0.8674
1940		0.1143	0.9794	1954	0.1143	0.9763		
1945		0.1429	0.8674	1959	0.1143	0.9794		
1950		0.1143	0.9763	1964	0.1714	0.6902		
Japan	1955	0.0857	0.9995	Sweden	1955	0.1714	0.6902	
	1960	0.1143	0.9763		1960	0.1714	0.6902	
	1965	0.1143	0.9763		1965	0.1429	0.8745	
	Netherlands	1955	0.1143	0.9794	US	1943	0.1143	0.9794
		1960	0.1143	0.9794		1948	0.1429	0.8745
1965		0.1714	0.6902	1953		0.1714	0.6902	
1940	0.1714	0.6826	1958	0.1714		0.6902		
1945	0.1143	0.9763	1963	0.1714		0.6902		
	1950	0.1429	0.8745					
	1955	0.2000	0.4916					
	1960	0.1714	0.6826					

Source: Authors' calculations using the Human Fertility Database.

Appendix B: Calculation process: Expected years without children

The life expectancy from age zero to age X is shown as

$${}_xe_0(t) = \int_0^X l_{x,t} dx.$$

We call the life expectancy between age 15 and age 50 the expected years without children (denoted ${}_{35}e_{15}(t)$). It is formulated as follows:

$$\begin{aligned} {}_{35}e_{15}(t) &= \int_{15}^{50} l_{x,t} dx \\ &= 35 - \int_{15}^{50} F_{x,t} dx, \end{aligned}$$

where l_x is the probability of remaining childless and F_x is its cumulative probability function. The detailed calculation procedure can be found in [Mogi and Canudas-Romo \(2018\)](#).

Appendix C: The decomposition to discrete data

Three parameters of f_x are estimated using the maximum likelihood estimation method, as suggested by [Rodríguez and Trussell \(1980\)](#):

$$\ln LH = \sum_{15}^{49} (\text{With}_x \log[F_{(x+0.5)}] + \text{Without}_x \log[1 - F_{(x+0.5)}]),$$

where With_x is the female population with children at age x , Without_x is the female population without children at age x , and F_x is the cumulative probability function at age x .

[Vaupel and Canudas-Romo \(2003\)](#) and [Bergeron-Boucher et al. \(2015\)](#) applied the continuous decomposition equation to discrete time data, and we followed their method. To estimate each function by applying our decomposition method to discrete time data, we use the midpoint over a time interval ([Preston et al. 2001](#)). As [Mogi and Canudas-Romo \(2018\)](#) assumed for the nuptiality decomposition, an exponential change assumption is used for the functions except EYWC, and the midpoint of EYWC is assumed to be a linear change in the interval. The details can be found in [Mogi and Canudas-Romo \(2018\)](#).

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