Universitat Autònoma de Barcelona

# Duration-Based Measures as an Alternative to Studying Union Formation and Fertility

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# Duration-Based Measures as an Alternative to Studying Union Formation and Fertility


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# Preface

This thesis is an assortment of my three published and accepted articles and a book chapter completed during my PhD program. All three studies in this thesis have a common idea; bringing an alternative perspective, *duration*, to measure family demographic trends. The famous phrase from Forrest Gump said “Life was like a box of chocolates. You never know what you’re gonna get.” My PhD life was just like a box of chocolate in this sense with several lucky encounters. In the winter of 2015, I was a first-year Master student in Japan when I searched for information about completing a PhD in Spain. Although I found some options, I was not able to find clear information, e.g. exam dates, the tuition fee, salary, and so on. Thus, I sent an email to my senior colleague, Setsuya Fukuda, to ask if he knew some Spanish demographers. He helped arrange a meeting with Dr. Albert Esteve in the Population Association of America 2015 in San Diego. From there, the story went rapidly, carrying me to a PhD position at the Centre d’Estudis Demográfics (CED) with a generous support from the Spanish Ministry of Economy and Competitiveness.

The first year of my PhD started with European Doctoral School of Demography (EDSD) in the Max Planck Institute for Demographic Research (Rostock, Germany) and La Sapienza University (Rome, Italy). I had many lucky encounters there as well, meeting accomplished demographers, other colleagues, and fellow EDSD students. Learning various demographic topics was a great way of starting a career in demography. Especially, I was inspired by Prof. Vladimir Canudas-Romo. He taught us basic math and showed how math is useful in demography. I asked so many questions in his class, at the bar, and by email. The luckiest thing during the EDSD was that I was able to convince him and work my EDSD thesis. Each iteration, during which I updated the draft and he gave detailed comments, was very fruitful. This process functioned as on-the-job training in how to write an academic article. In addition, life in Rome was amazing. Walking through the ruins from 2000 years ago, eating crispy Roman pizzas, cycling to Colosseo at night, and having a gelato are great memories.

During my PhD, I had so much support and encouragement from various people. First, I would like to thank my supervisors, Dr. Albert Esteve and Dr. Diederik Boertien. Second, my "seven-samurai" colleagues and co-authors; Bruno Arpino, Daniele Vignoli, Fumiya Uchikoshi, Jessica Nisén, Michael del Mundo, Peter McDonald, and Vladimir Canudas-Romo. Lastly, I have received much kindness and joy from my friends in Barcelona, Canberra, Japan, and all over the world, as well as from all the CED staff, especially Soco and Eulalia.

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This thesis is just a starting point for my future research. I have several on-going projects and other ideas with larger scopes that I would like to focus on after completing my PhD. Demography is a study of people's life-course; thus, I would be grateful if my research contributes to making people's lives a little bit better in future. I am going to work hard to do so.

Before closing this preface, I would like to thank my wife, Chie, and my family. Chie quit her successful job in Japan and followed me, supporting my career. Thanks to her efforts, support, and smile, I have been able to keep working hard. A big "Thank you," Chie.

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# Overview

# Chapter 1

# Introduction

This is a dissertation thesis about formal family demography consisting of three published and accepted articles by international journals. This thesis aims to bring an alternative perspective to the study of family demography by measuring the *duration* of union and fertility events and to develop three new indexes for a better understanding of current family dynamics. In the next section, I explain the conventional indexes used in family demography, which are (1) the quantum and timing indexes and (2) the period and cohort indexes, and the advantages and disadvantages of these conventional indexes. Then, I describe the duration indexes and what they can add to the study of family demography. After showing data sources in this thesis, the current changes in first marriage and fertility in high-income countries are shown. The final section outlines each chapter: Chapter 2 titled "Expected Years Ever-Married", Chapter 3 titled "Decomposing changes in first birth trends: Quantum, timing, or variance", and Chapter 4 titled "Cross-sectional average length of life by parity: Illustration for the US cohorts in reproductive ages in 2015."

## 1.1 Formal demography in family demography

Demography is the study of changes in the size, growth rate, and structure of the human population (Preston et al. 2001). Thus, indexes to measure those changes are significantly important in this field. The study of depicting and measuring demographic changes is called formal demography. Another sub-group in demography examines population compositions and changes from an interdisciplinary point of view, such as sociological, economic, biological, or anthropological. This is called social demography or population studies (Xie 2000). Using these definitions, this thesis falls within formal demography.

There are a series of great research developing demographic measures (e.g. for mortality, Life Years Lost: Andersen et al. (2013), Cross-sectional Average Length of Life: Brouard (1986) and Guillot (2003) and for fertility, Tempo-adjusted Total Fertility Rate: Bongaarts and Feeney (1998), Multistate Life Table Methods: Schoen (1988)); however, the approach of developing measures is different between mortality research and fertility research. Mortality is relatively simple to conceptualize in mathematical equations, i.e. everyone dies at some point and only once; on the other hand, union and fertility events can occur multiple times during a lifetime or not all. Besides, in family demography, survey data including rich information on life history is often available. Therefore, researchers tend to develop and use regression-type methods to examine the association of explanatory variables with an outcome controlling away the other important variables instead of developing demographic measures. For example, researchers use a regression model to ask "How is education associated with first marriages?" (e.g. Raymo (2003)) or "How does having children relate to parents' subjective well-being?" (e.g. Myrskylä and Margolis (2014)).

This thesis tackles two common methodological issues especially in family demography using a classical demographic method. The first issue is about quantum and timing indexes. In family demography, indexes to measure quantum and timing have been used frequently. The quantum index measures the number of people who experienced a certain event at (or by) a certain time—for example, the share of childless women at the end of the reproductive period, the proportion of never-married people, and total fertility rate (TFR). The timing index specifies when a certain event occurred and the following are commonly used: the mean age of first marriage and the mean age of first birth. They are frequently used in both media and scientific articles. For example, the proportion of cohort childless women increases from 11.2 % in 1953 cohort to 27.9 % in 1973 cohort in Japan and from 9.5 % in 1960 cohort to 21.6 % in 1972 cohort in Spain. The mean age at first birth increased from 26.0 years in 1953 cohort to 29.0 years in 1976 cohort in Japan and from 25.8 years in 1960 cohort to 30.0 years in 1976 cohort in Spain (Human Fertility Database 2018). Based on these data, people increasingly remain childless and postpone their first birth to a later age. These results prompt many questions: Why is this? Can we find a similar trend in other countries? Will this increasing trend continued?

These traditional union and fertility indexes are undoubtedly useful for studying either quantum changes or timing changes. However, the weakness of these indexes is that they are incapable of taking into account both changes; the proportion of childless people does not consider timing changes (e.g. the same proportion of childless people can be reached with different mean age at first childbirth) and the mean age at first birth per se does not show quantum changes. This feature is especially problematic in high-income countries, where both quantum and timing of union and fertility events have changed. One of the most well-known problems is the tempo-distortion effect in the total fertility rate (Bongaarts and Feeney 1998). The changes in the timing of births influence to the level of quantum—total fertility rate. Moreover, given the current union and fertility changes in high-income countries, the meaning of timing indexes may differ over time. Assume a hypothetical situation: the mean age at first birth has been the same, age 25, from 1950 to 2010 in a country, while the share of childless people has increased from 0 % in 1950 to 50 % in 2010. In this scenario, the mean age at first birth in 1950 is calculated for 100 % of the population (everyone has had at least one child) but in 2010 it is calculated among only half of the population. Even if the mean age at first birth has not changed, its meaning may change due to the changes in the share of childless people. Thus, the changes in quantum indexes effect on the timing indexes as well. Prior research also has assessed the interaction between timing changes and quantum changes. For example, Kneale and Joshi (2008) explore the extent of first birth postponement and project its impact on eventual levels of childlessness in the UK. Goldstein and Kenney (2001) try to separate nonmarriage (quantum changes) and delayed marriage (timing changes) by estimating the cumulative proportion of women who ever marry.

A comprehensive measure that (1) captures both timing and quantum and (2) covers the entire population both conceptually and mathematically may provide additional insight regarding the current dynamics in union formation and fertility. Therefore, this study brings an alternative perspective, *duration*, into the study on union and fertility.

The second issue relates to the well-known limitations of period and cohort indexes. The period perspective (using a synthetic cohort approach) combines information from many different cohorts; thus, it does not necessarily reflect the experience of any real birth cohort (Bongaarts and Sobotka 2012; Luy 2011), while the cohort indexes provide an outdated picture of fertility as they are based on information on populations who are past their reproductive age. Therefore, Chapter 4 introduces an alternative period measure including all the cohort information at reproductive ages present at a given time. As such, it provides a period measure that is informative not only of the childbearing behavior of a given period but also of fertility behavior of female cohorts who are at present at reproductive ages in this given period.

This thesis introduces three alternative indexes to measure the duration of ever marrying (Expected Years Ever-Married; EYEM, Chapter 2), remaining childless (Expected Years Without Children; EYWC, Chapter 3), and spending in each parity (Cross-sectional Average Length of Life by Parity; CALP, Chapter 4). These three indexes give a more intuitive interpretation: how many years the entire population spends in these family-life statuses on average. They are macro-level indexes, summarizing the changes in individual-level behavior for an aggregate-level process (Preston et al. 2001). As Schoen (2019) mentioned, the changes in family life call for new forms of analysis, and researchers are still at an early stage of developing tools to analyze modern family demography. Alternative indexes may improve our understanding of union and fertility trends in high-income countries where both quantum and timing have changed. Union formation and fertility decision depend highly on age, especially as women have limited window for childbirth, called the "biological fertility clock." Therefore, measuring the duration of family-life events provides a detailed picture of family building.

## 1.2 Data and studied countries

To apply these alternative indexes, I employed national censuses, vital statistics/register data, and national surveys. Due to data availability, Chapter 2 used different data sources than Chapters 3 and 4. Specifically, Chapter 2 used population counts by sex, age, and marital status from national statistical offices and the United Nations database. National statistical offices normally publish population counts with single-age intervals; hence, those are the most accurate databases. The United Nations offers only population counts by five-year age groups; however, for some countries, this was the only information available to construct a historical series. Chapters 3 and 4 used age- and parity-specific female population counts and birth counts by birth order and the mother's cohort from the Human Fertility Database (HFD). The HFD is a joint project of the Max Planck Institute for Demographic Research (MPIDR) in Rostock, Germany and the Vienna Institute of Demography (VID) in Vienna, Austria and is based at MPIDR. Detailed information can be found on its website (http://www.humanfertility.org/cgi-bin/main.php). The HFD is an open-access database containing 28 countries from Europe, North and South America, and Asia. The data is heavily scrutinized for quality and only countries that have comprehensive, high-quality information are included in the database.

To compare changes in union and fertility using three indexes as many countries as possible, all high-income countries that satisfied data criteria were selected. Chapter 2 included Austria, Belgium, Canada, Czechia, Denmark, France, Germany, Greece, Ireland, Italy, the Netherlands, Spain, Sweden, Switzerland, and the UK; Chapter 3 included Canada, Czechia, Japan, the Netherlands, Norway, Portugal, Sweden, and the US; and Chapter 4 analyzed the US as an example to show the applicability of method.

The major weakness of these data is that they do not include information on socioeconomic status, which has a significant impact on first-marriage timing and decision (e.g. Blossfeld et al. (2005)) and fertility timing and decision (e.g. Beaujouan et al. (2016), Kneale and Joshi (2008), and Kravdal and Rindfuss (2008)).

## 1.3 Trends in first marriage

Marriage and childbirth have been strongly connected in the past; however, accompanying an increase in cohabitation, the nonmarital childbirth rate has increased in many high-income countries. Those family changes are interpreted in the framework of the Second Demographic Transition (Lesthaeghe 2010). Due to the changes in union and fertility behaviors, previous research suggests that marriage has decoupled from childbearing (Kiernan 2001; Lesthaeghe 2010; Raley 2001; Smock and Greenland 2010; van De Kaa 2001). However, there is contradictory research reporting that approximately 60 % of women get married before or after their first birth (Perelli-Harris et al. 2012). This suggests that the order of union and fertility events—i.e. cohabitation, childbirth, and marriage—have been more flexible in those countries, but marriage as a union formation plays an important role in childbearing and childrearing. Contrary to those countries, the nonmarital childbirth rate is still low in East Asia. For example, Japan and South Korea have a low nonmarital childbirth rate, i.e. 2.2 % in 2017 (National Institute of Population and Social Security Research 2019) and 1.9 % in 2014 (OECD 2018), respectively. In these countries, marriage is the dominant type of union to have children.

Although marriage keeps its importance as a type of union formation, the never-married rate at the end of reproductive life has been increasing in most high-income countries. In

2015, the never-married rate is 29.5 % in Sweden and 16.1 % in Japan (National Institute of Population and Social Security Research 2019), which is an increase from 4.5 % in 1980 in Japan. Thus, even in a country where marriage and childbearing are strongly tied, the proportion of never-married women has risen significantly. Changes in marriage behavior are seen not only in the quantum but also the timing. For example, the female mean age at first marriage has a remarkable increase from 1980 to 2017 from 26.4 years to 33.8 years in Sweden (United Nations Economic Commission for Europe 2019; Eurostat 2019) and from 25.2 years to 29.4 years in Japan (National Institute of Population and Social Security Research 2019), due possibly to the prevalence of cohabitation, a retreat from marriage, or marriage postponement.

## 1.4 Trends in fertility

In each decade since the 1970s, the mean age of first birth has increased on average by about one year in high-income countries (Mills et al. 2011). The country variations are substantial; women in the US and Eastern European countries continue to have the first child at a particularly low age, while women in Southern Europe and East Asia enter motherhood generally at later age, i.e. the current average age of the first birth is above the 30. The postponement of parenthood places women at a higher risk of remaining childless (Kneale and Joshi 2008; Schmidt et al. 2012; Toulemon 1996).

The phenomenon of childlessness has received significant attention as the share of childless people has increased in high-income countries in recent decades. This share has been steadily increasing in Japan, Spain, and Taiwan for decades and has recently begun increasing in Northern, Central, and Eastern European countries after being stable for decades. In Austria, Finland, and Spain, approximately 20 % of women are childless and 28 % of women in Japan remain childless at the end of their reproductive period among the recent cohort (Human Fertility Database 2018). In addition, the proportion of childless women is projected to continue increasing (Sobotka 2017). According to the Human Fertility Database, among high-income countries, only Canada and the US are projected to have decreasing the level of eventual childless.

Increasing the level of eventual childlessness influences the total fertility rate. In German-speaking countries, Southern Europe, and East Asia, the increase in the share of childless women is the most influential factor in changing completed cohort fertility rates between 1955 and 1970 cohorts, compared to changes in other parity behaviors (Zeman et al. 2018). Besides its influence at the country level, becoming a mother or remaining childless influences various aspects of an individual's life. For example, remaining childless has an impact on income (Budig et al. 2012), health (Kendig et al. 2007), old-age wellbeing (Huijts et al. 2013), and support networks (Albertini and Kohli 2009). Therefore, examining childlessness and the timing of the first birth is significantly important.

The arrival at the current low TFR can be viewed as a result of changes in parity-specific behavior over time. The parity progression ratios (PPR) from parity two-to-higher parities dropped significantly from the 1930 birth cohort to the 1965 birth cohort and from parity zero-to-one and from one-to-two started decreasing in more recent cohorts in most European countries (Frejka 2008). For instance, the parity progression ratio from parity 0 to parity 1 (denoted as PPR 0-1) decreased from 0.89 (1953 cohort) to 0.75 (1967) in Japan and from 0.91 (1960) to 0.86 (1966) in Spain, and the other parity ratios decreased as well: PPR 1-2, from 0.87 to 0.76 and from 0.70 to 0.66; PPR 2-3, from 0.26 to 0.20 and from 0.37 to 0.31 during the same period in Japan and Spain, respectively (Human Fertility Database 2018).

## 1.5 Outline of the study

Accompanying the increase in cohabitation and nonmarital childbirth, women increasingly enter motherhood through different paths (Lesthaeghe 2010; Perelli-Harris et al. 2012; Raley 2001). However, more than 60 % of women are married at some point around their first birth (Perelli-Harris et al. 2012). This indicates that marriage continues to have an important role in having and raising children. Therefore, for a better understanding of first marriage behavior, Chapter 2 "Expected Years Ever-Married" illustrates changes in first marriage behavior using an alternative index, Expected Years Ever-Married (EYEM). Population counts by sex, age, and marital status of 15 countries from national statistical offices and the United Nations database are used. This work is from the peer-reviewed article with Vladimir Canudas-Romo (School of Demography, Australian National University) published in *Demographic Research* Volume 38, Article 47 in 2018. I was the corresponding author of this study and took the lead outlining the manuscript, preparing the data, and analysing. Prof. Canudas-Romo contributed to develop the methodology and gave comments throughout all stages of the preparation of the manuscript.

Chapter 3 "Decomposing changes in first birth trends: Quantum, timing, or variance" introduces an alternative index, Expected Years Without Children (EYWC) to quantify changes in first birth behavior. Using the Human Fertility Database, EYWC is calculated to show time trends for eight countries: Canada, Czechia, Japan, the Netherlands, Norway, Portugal, Sweden, and the US. Those countries were selected because of their sufficiently long enough histories of first birth data. The time trends of EYWC show that women born in the latest cohorts observed in Canada, Japan, and the Netherlands spent half of their reproduction periods without any child. Furthermore, we decompose the changes in EYWC over time into three effects: remaining childless, postponing first birth, and expansion of the standard deviation of mean age at first birth. Results of the decomposition show that postponement is the most influential factor on EYWC changes in North America and Northern Europe while remaining childless is the main contributor in Japan and Portugal. This work is from the peer-reviewed article with Michael Dominic del Mundo (Population Institute, University of the Philippines) accepted in *Vienna Yearbook of Population Research*. I was the corresponding author of this study and took the lead outlining the manuscript, preparing the data, and analysing. del Mundo contributed to develop the methodology and gave comments throughout all stages of the preparation of the manuscript.

Finally, in Chapter 4 "Cross-sectional average length of life by parity: Illustration for the US cohorts in reproductive ages in 2015," we investigate the analysis of parity-specific behaviors from parity 0 to parity 5 and over. The Cross-sectional Average Length of Life by Parity (CALP) is introduced as an alternative way of understanding fertility trends. CALP shows the length of time women spent in each parity during reproductive ages and is a period measure including all the cohort fertility information of reproductive-aged women at a given time. Selecting the US data from the Human Fertility Database for illustration, CALP is calculated using a hierarchical multistate life table model. CALP for 2015 shows that women in the US spent 47 % (17.91/38 years) of reproductive years (ages 12 to 50) in childlessness, followed by 16 %, 19 % and 11 % in parities 1, 2, and 3, respectively. This work is from the peer-reviewed book chapter written with Vladimir Canudas-Romo in a book titled "*Analyzing Contemporary Fertility*" edited by Robert Schoen (Population Research Institute, Penn State University). I was the corresponding author of this study and took the lead outlining the manuscript, preparing the data, and analysing. Prof. Canudas-Romo contributed to develop of the methodology and gave comments throughout all stages of the preparation of the manuscript. Dr. Schoen reviewed the manuscript and gave comments.

Lastly, Chapter 5 summarizes the aims, the main findings and implications of this study. In addition, it mentions the major limitations of this study and suggests ideas for future research.

# Part II

# Alternative indexes in family demography

# Chapter 2

# Expected Years Ever-Married

## 2.1 Introduction

Nuptiality behaviour has changed remarkably in many countries since the middle of the twentieth century. This change is often described as the "second demographic transition" (Lesthaeghe 1983; Van de Kaa 1987). The main characteristics of this change are a tendency for people not to get married (non-marriage), and to postpone their marriage (delayed marriage), which creats a wide variability in first marriage age across countries (Winkler-Dworak and Engelhardt 2004; Elzinga and Liefbroer 2007; European Commission 2015). So far, research has focused on analysing the determinants of those nuptiality changes. However, work to clearly disentangle whether people tend not to get married or tend to postpone marriage is missing. This long overdue explanation (Oppenheimer 1994) is the main purpose of this article.

Theoretically, non-marriage and delayed marriage are clearly separated phenomena (Becker 1981; Oppenheimer 1988, 1994). While Becker's theory predicts a rise in non-marriage, this

This chapter is a published article; Mogi, R. and Canudas-Romo, V. (2018). Expected Years Ever-Married. *Demographic Research* 48(47):1423-1456.

is not supported by empirical analyses (Oppenheimer 1994; Goldstein and Kenney 2001; Winkler-Dworak and Engelhardt 2004). Research on the topic has worked on separating non-marriage and delayed marriage. For example, Goldstein and Kenney (2001) estimated the cumulative proportion of women ever-marrying using the Coale and McNeil (1972) model (CM model) and the Hernes model. They concluded that delayed marriage was the main component of the changes of proportions ever-marrying in US female cohorts in the 1950s and 1960s, because the proportion of marriages decreased only slightly by birth cohort. In addition, the change from 1965 to 1980 for non-Hispanic white American female cohorts was also explained by delayed marriage (Oppenheimer 1994). While those studies focused on survival functions and cumulative proportions, Wu (2003) suggested distinguishing non-marriage and delayed marriage, by checking the shape of the hazard rate of first marriage. He showed how this hazard rate would change if pure delayed marriage was occurring (Wu 2003). However, an analytical disentanglement of the components and the quantification of the effects of non-marriage and delayed marriage remains to be done.

An additional component in the changes observed in first marriage is the variance in first marriage age, which is increasing over time. Elzinga and Liefbroer (2007) compared the life course trajectories of young cohorts in 19 countries, and concluded that those life trajectories into marriage varied more than for older cohorts. Winkler-Dworak and Engelhardt (2004) explained the significance of variance in marriage timing and highlighted that most research has ignored the changes in this component. Hence, the change in the standard deviation of age at first marriage, which we call an "expansion effect", remains to be investigated. Besides non-marriage and delayed marriage, we also examine the effect of variance in age at first marriage on nuptiality changes.

Our research is different to studies that develop tempo-adjusted indices. The proportion of those who ever marry and the mean age at marriage are often used as quantum and timing indices, respectively. However, these period indices are influenced by tempo distortions and the majority of the research has focused on adjusting them (Winkler-Dworak and Engelhardt

2004; Schoen and Canudas-Romo 2005; Bongaarts and Feeney 2006). The purpose of the tempo-adjusted indices is to have more accurate results at each given time, while our interest is in quantifying changes over time and disentangling the contribution of each component: non-marriage, delayed marriage, and expansion of first marriage timing.

This article has two aims. Firstly, we introduce "expected years ever-married" (EYEM) as a new alternative index to describe the transition from never-married to ever-married status. Secondly, the changes over time in EYEM are decomposed into three effects: scale (the changes in the proportion of never-married population, or non-marriage), location (the changes in timing of first marriage, or delayed marriage) and variance (the changes in the standard deviation of first marriage age, or expansion). The decomposition method reveals the impact on the change in marriage behaviours by each of these components. We illustrate the new measure and its decomposition by looking at historical trends and comparing those effects across countries.

This article is divided into four sections, with this introduction as the first section. In the second section, we introduce the new measure and method of decomposition as well as the data used. The third section illustrates the use of the new index and its decomposition in long-term nuptiality changes, comparing 15 countries for period data and six countries for cohort data. A discussion, limitations, future developments and conclusion are found in the final section.

## 2.2 Methods and data

### 2.2.1 Expected years ever-married (EYEM)

EYEM is an alternative index to interpret nuptiality changes over time using classical demographic methods. As pointed out above, previous research that separated non-marriage and delayed marriage inspected this graphically (Oppenheimer 1994; Goldstein and Kenney 2001). For example, the two lines in Figure 2.1 represent the probability of remaining never-married ($l_{x,t}$) by age among a cohort of never-married female, 15-year-olds exposed to the marriage probabilities of Sweden in 1970 and 2015.

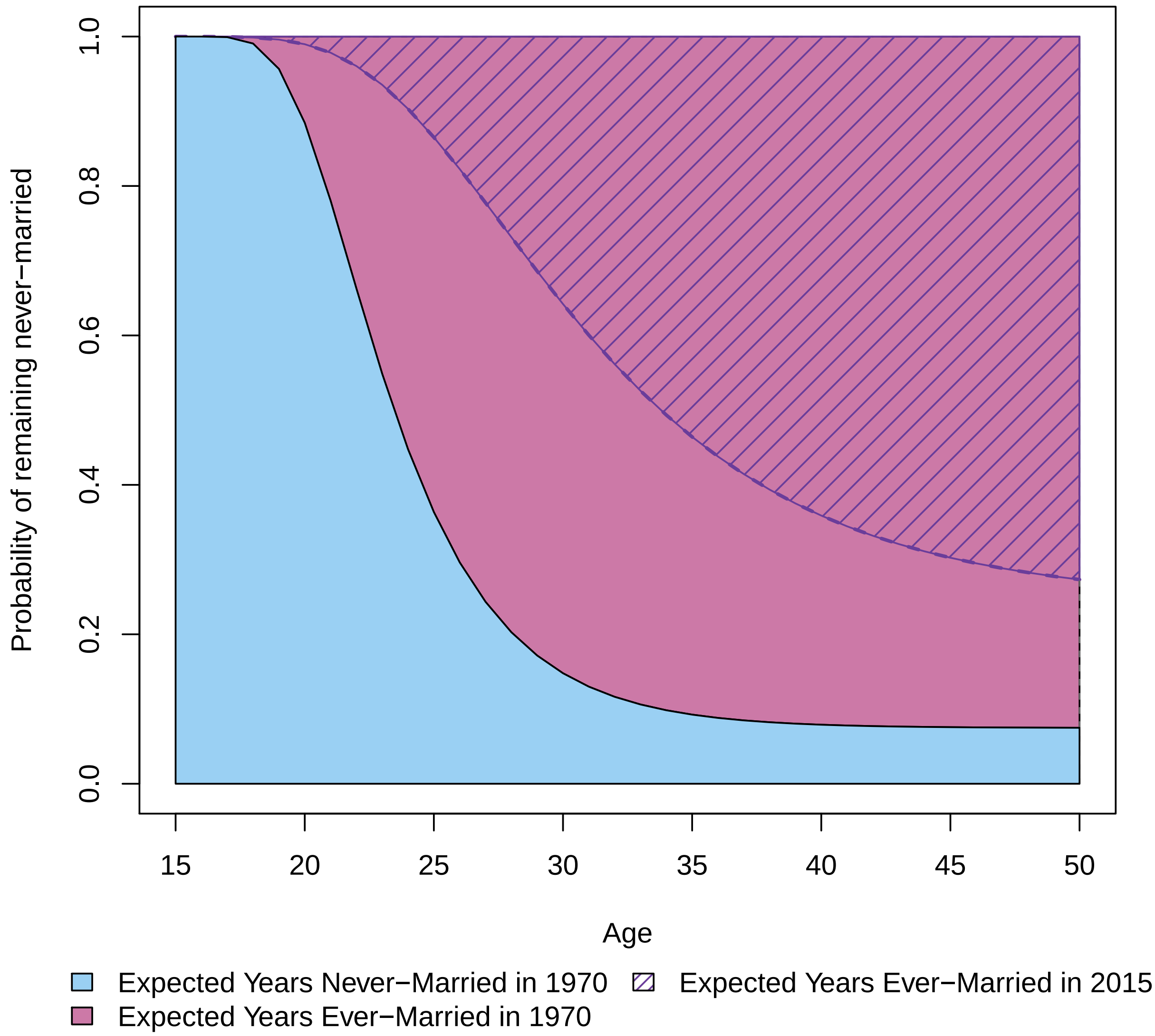



Figure 2.1: Probability of remaining never-married by age among a synthetic cohort of never-married 15 year olds exposed to the marriage probabilities of Swedish females in 1970 and 2015

*Note*: Each probability of remaining never-married is estimated using the Rodríguez and Trussell's parametrisation (Rodríguez and Trussell 1980), explained in section 2.2.3. The parameters of the probabilities of remaining never-married are $C = 0.925$, $\mu = 24.429$, and $\sigma = 4.044$ for 1970, and $C = 0.757$, $\mu = 32.712$, and $\sigma = 8.492$ for 2015.

*Source*: Authors' calculations, using Swedish female data described in Table 2.1.

In classical life table methods, life expectancy between two ages, say 0 and $X$, can geometrically be seen as the area below a survival function from age 0 to that fixed age $X$. This is interpreted as the average number of years people live between these ages (Preston, Heuveline and Guillot 2001). The area above the survival function between age 0 and age $X$ is called life years lost (Andersen, Canudas-Romo and Keiding 2013). This index shows the average years lost due to death in this age interval. In the marriage context, the transition of interest is from never-married to marriage. In addition, we set the minimum legal age for marriage as age 15[1]. One of demographers' focus on marriage is its relation with fertility and as noted by Perelli-Harris (2014), this relation is still important today, particularly for second births. Since age 50 is the last fecundity age for the vast majority of women, this age can be regarded as the upper age of interest. For the rest of the analysis we assumed that mortality is not present in this age-interval, since mainly low mortality countries were studied. Therefore, the expected number of years of never-married (EYNM) from age 15 to age 50, denoted as ${}_{35}e_{N,15}$, is calculated as ${}_{35}e_{N,15}(t) = \int_{15}^{50} l_{x,t}dx$. It corresponds to the lower-left shaded area in Figure 2.1. The complement is the expected years ever-married between the ages of 15 and 50, denoted as ${}_{35}e_{M,15}$ and calculated as ${}_{35}e_{M,15}(t) = \int_{15}^{50} 1 - l_{x,t}dx$, and shown in the two upper areas in Figure 2.1 for the years 1970 and 2015, respectively. Further advantage of the complementarity of EYNM and EYEM is that they add to the total 35 years at all times,

$$ {}_{35}e_{N,15}(t) + {}_{35}e_{M,15}(t) = 35. $$

In this life table approach to marriage, the measure EYEM is calculated from the probabilities of remaining never-married ($l_x$), which are computed from a set of age-specific marriage rates. One advantage of using EYEM to describe nuptiality change is that it has a simple and meaningful demographic interpretation, namely the number of years ever-married. Thus, it

[1]For most European countries, the minimum legal age at which marriage can take place without parental consent is 18. However, if they have parental consent, they are allowed to get married at a younger age than 18 (United Nations 2016). In this study, we assigned the minimum legal age for marriage as 15, as this is the lowest age found in the data used with a marriage rate above zero.

allows us to numerically compare transitions to marriage at different times. For instance, in 1970, EYNM between age 15 and age 50 was 11.3 years and EYEM was 23.7 years for Swedish females - shown as the filled upper area in Figure 2.1. Those expectations reversed to 21.7 years for EYNM, and 13.3 years for EYEM in 2015 (lined area in Figure 2.1).

The expected years ever-married measure has a close relationship to an index that is commonly used in nuptiality research, namely the age-specific proportion ever-marrying ($\text{PEM}_{x,\,t}$), since,

$$\text{PEM}_{x,\,t} = 1 - l_{x,\,t}, \tag{2.1}$$

where $l_{x,\,t}$ is, as before, the probability of remaining never-married at age $x$ at time $t$ and the proportion ever-marrying at age 50 is also denoted as $C_t = \text{PEM}_{50,\,t}$. The EYEM can then be calculated as

$$_{35}e_{M,\,15}(t) = \int_{15}^{50} \text{PEM}_{x,\,t}\, dx. \tag{2.2}$$

In this study, we focus on EYEM as a main index to describe nuptiality changes and compare it over time.

### 2.2.2 Decomposition method

Let the age-specific probability of first marriage rates at time $t$ be denoted as $f_{x,\,t} = f_x(C_t, \mu_t, \sigma_t)$, and be a function of three parameters: scale (the proportion of the cohort eventually marrying), location (the mean age at first marriage), and variance (the standard deviation of age at first marriage). We decompose the changes in EYEM over time, denoted as $_{35}\dot{e}_{M,\,15}(t)$, into the contribution of those three parameters as

$$_{35}\dot{e}_{M,\,15}(t) = \frac{\partial\, _{35}e_{M,\,15}(t)}{\partial C_t}\dot{C}_t + \frac{\partial\, _{35}e_{M,\,15}(t)}{\partial \mu_t}\dot{\mu}_t + \frac{\partial\, _{35}e_{M,\,15}(t)}{\partial \sigma_t}\dot{\sigma}_t, \tag{2.3}$$

where each term is the change in $_{35}\dot{e}_{M,\,15}(t)$ resulting from changes in the scale, location, and variance, respectively. The succinct notation of a dot on top of a variable, used here, indicates the derivative with respect to time, which is shown to simplify equations and aid in the development of new methodology (Vaupel and Canudas-Romo 2003; Bergeron-Boucher, Ebeling and Canudas-Romo 2015). When the change in scale factor ($\frac{\partial\, {}_{35}e_{M,\,15}(t)}{\partial C_t}\dot{C}_t$) is the biggest value among the three components, it means that the changes in EYEM are mainly caused by non-marriage. Likewise, when the location ($\frac{\partial\, {}_{35}e_{M,\,15}(t)}{\partial \mu_t}\dot{\mu}_t$) or variance ($\frac{\partial\, {}_{35}e_{M,\,15}(t)}{\partial \sigma_t}\dot{\sigma}_t$) factor is the biggest, this corresponds to delayed marriage and expansion respectively. This decomposition is inspired by research that separates transitions in life expectancy into change due to compression and shifting effects (Bergeron-Boucher, Ebeling and Canudas-Romo 2015).

Figure 2.2 illustrates four different age-patterns of first marriage distributions for Swedish females. The solid black line is the probability distribution of first marriage in Sweden 1970, and the solid purple line is the one in 2015. The other dashed lines are the simulated distributions when only one component changes from 1970 to 2015. The dashed orange line demonstrates a hypothetical marriage distribution in 2015, if only the parameter $C$ (the proportion ever marrying) had changed from 1970 to its value attained in 2015. When a pure non-marriage occurs (i.e. only $C$ decreases), the probability is just compressed with the same average age at marriage (in Figure 2.2, the orange arrow). Pure delayed marriage is represented by the change of only $\mu$. As people tend to marry later (i.e. only $\mu$ increases), the probability slides to the right (the black solid line to the dotted blue line in Figure 2.2) but the sizes below the probability distribution are the same. Lastly, if people's first marriage timing becomes more varied (i.e. $\sigma$ increases), as shown by the green arrow in Figure 2.2, the maximum value of the probability declines and its shape is widened. The decomposition in equation (2.3) allows us to perfectly disentangle the contribution of these three components to the time change in EYEM.

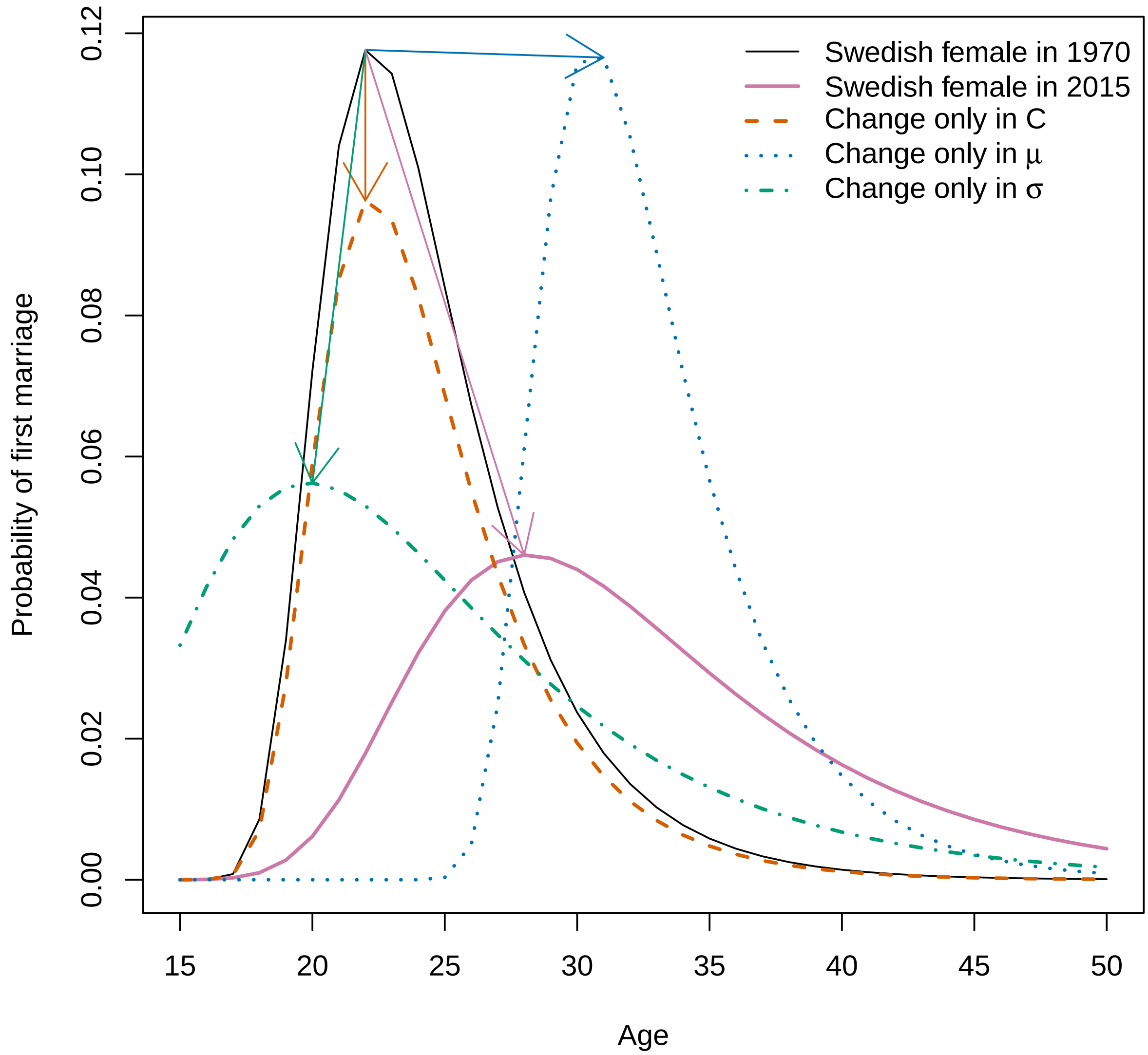



Figure 2.2: Changes in the probability of first marriage: Swedish females from 1970 to 2015
*Note*: The parameters used are the same as noted in Figure 2.1. The other lines reflect changing only one of the components at a time to its value in 2015 and keeping the rest as per those in 1970.
*Source*: Authors' calculations, using Swedish females data described in Table 2.1.

### 2.2.3 Parametric models of first marriage

The Coale-McNeil model (CM model) (Coale and McNeil 1972) is widely used for estimating the probability of first marriage (Rodríguez and Trussell 1980; Bloom and Bennett

1990; Goldstein and Kenney 2001; Kaneko 2003; Peristera and Kostaki 2015). To calculate EYEM and apply it to the decomposition equation, we use a standardised version of the CM model, namely Rodríguez and Trussell's parametrisation (Rodríguez and Trussell 1980) of the probability density function of first marriage, which we refer to as the RT parametrisation hereafter. This probability of first marriage at age $x$ and time $t$, denoted as $f_{x,t}$, is expressed in the RT parametrisation as a function of the proportion of the cohort eventually marrying at time $t$ ($C_t$), the mean age at first marriage ($\mu_t$), and the standard deviation of age at first marriage ($\sigma_t$):

$$f_{x,t} = C_t \frac{1}{\sigma_t} a_1 \exp\Big[a_2(\frac{x-\mu_t}{\sigma_t} + a_3) - \exp\{-a_4(\frac{x-\mu_t}{\sigma_t} + a_3)\}\Big], \tag{2.4}$$

where the usual values for the constants are $a_1 = 1.281$, $a_2 = -1.145$, $a_3 = 0.805$, and $a_4 = 1.896$. Equation (2.4) can be concisely formulated as:

$$f_{x,t} = C_t \frac{1}{\sigma_t} f_0(\frac{x-\mu_t}{\sigma_t}), \tag{2.5}$$

where $f_0$ is the density function derived from equation (2.4) where values of the mean age ($\mu_t$) and the standard deviation ($\sigma_t$) are the vital input information to standardize it. Its cumulative density function is written as

$$F_{x,t} = C_t F_0(\frac{x-\mu_t}{\sigma_t}), \tag{2.6}$$

where $F_0$ is the cumulative schedule of values of the density function $f_0$ starting at age 15 until age $X$. The parameter $C_t$ is the proportion ever-married at age 50 and $\mu_t$ can be interpreted as the singulate mean age at marriage (SMAM) (Rodríguez and Trussell 1980).

While the CM model is commonly used to parametrise first marriage, there are some opposing opinions to its application. Kaneko (2003) applied the RT parametrisation to Japanese female cohorts (1953-1960) and explained that the standardised CM model might be inappropriate for some countries and times, because the model does not fit well to the observed data. This limitation of the model is also seen in European countries (Peristera and Kostaki 2015). Therefore, Kaneko (2003) suggested using an extended version of the CM model, namely the generalised log gamma distribution model, and Peristera and Kostaki (2015) recommended using a mixture model. The reason that the CM model does not fit well to the observed data from those countries is mainly because of the mixture of marriage types, whose timings are distinctively different (e.g., arranged marriage and love marriage in the Japanese case, migration, religion, or the other socio-economic status for European countries) (Kaneko 2003; Peristera and Kostaki 2015). While those mixture models fit better than a series of the CM model, it is difficult to interpret and decompose those models. We use the parsimonious RT parametrisation for this study because its three parameters have meaningful demographic interpretation and it is a simple model, although we recognise the limitations of the model[2].

To quantify the effects of scale, location, and variance in the changes of EYEM over time, firstly, the cumulative density distribution in equation (2.6) is substituted in the definition of EYEM as

$$ {}_{35}e_{M,15}(t) = \int_{15}^{50} F_{x,t}\, dx. \tag{2.7} $$

Secondly, the derivative with respect to time is studied. Detail derivations of these equations and the calculations of EYEM are found in Appendix B.

Each parameter is estimated by the maximum likelihood estimation method suggested by Rodríguez and Trussell (1980). Our method can be applied to discrete data by estimating the functions at their midpoint over time (Preston, Heuveline and Guillot 2001; Vaupel and Canudas-Romo 2003). The detailed procedures involved in applying the decomposition to discrete data are found in Appendix C. For example, we used a linear approximation in the interval for the change over time of EYEM. Further sensitivity analysis was carried out using

[2] We compared the observed age-specific first marriage rate with the estimated based on the RT parametrisation. The RT parametrisation generally estimates quite well our selected data especially countries that have single age-groups, even though the RT parametrisation tends to underestimate the maximum value. The figures showing how the model fits can be seen in Appendix A.

exponential change instead, without any changes in the main results and conclusions.

### 2.2.4 Data

In order to quantify the scale, location, and variance of the first marriage using the decomposition method, we used population counts by sex, age, and marital status. Coale and McNeil (1972) applied their parametric model to cohort data, and other researchers, such as Goldstein and Kenney (2001) and Kaneko (2003) used cohort data for their analyses. However, other studies applied the CM model to period data as well (Rodríguez and Trussell 1980; Peristera and Kostaki 2015), with the purpose of examining the current trends. It is well-known that period and cohort data have strengths and weaknesses. The period data can describe current trends, while it mixes behaviours of different cohorts. The cohort data avoids the tempo-distortions, however, birth cohorts only refer to one group of people present at a given time. Taking into consideration those advantages and disadvantages, in this study, we present results from both period and cohort data.

Table 2.1 presents the details of the data used for the 15 selected countries. We used data from national statistical offices as the first choice when available; otherwise, the data were taken from the United Nations database. National statistical offices normally publish population counts with single age intervals; hence, those are the most accurate databases. The United Nations offers only population counts by five year age groups; however, for some countries, this was the only information available to construct a historical series. In addition, Denmark, the Netherlands, and Sweden also have registered partnership information. For the purposes of this study, we counted them as married. The cohort data was constructed from all the above period information. Due to data constraints, cohort data was built for six countries out of 15 countries in Table 2.1. When single age group was available, the cohort data was reconstructed from the period data incrementing over age and time: e.g. age 15 in 1940, age 16 in 1941, and so forth. Similarly, when the age group was five years, we used increments of five year age-groups every five calendar years: e.g. ages 15-19 in 1960, ages

20-24 in 1965, and so forth. Only completed cohorts, which contained data until age 49 were selected. The information of cohort data can also be seen in Table 2.1.

Table 2.1: Countries included in the analysis, and analysed years, birth cohorts, age group, and the data source

| Country | Year | Cohort | Age group | Source |
|---|---|---|---|---|
| Austria | 1951 - 2011 | | 5 | United Nations (UN) |
| Belgium | 1961 - 2011 | | 5 | UN |
| Canada | 1951 - 2014 | 1936 - 1966 | 5 | UN |
| Czech | 1960 - 2015 | 1945 - 1970 | 5 | Czech Statistical Office |
| Denmark | 1948 - 1970 | | 5 | UN |
| | 1971 - 2017 | 1956 - 1968 | 1 | Statistics Denmark |
| France | 1952 - 2013 | | 5 | UN |
| Germany | 1972 - 2015 | 1960 - 1970 | 5 | Federal Statistical Office (GENESIS) |
| Greece | 1951 - 2011 | | 5 | UN |
| Ireland | 1926 - 2011 | | 1 | Central Statistics Office |
| Italy | 1951 - 2014 | | 5 | UN |
| Netherlands | 1950 - 2015 | 1935 - 1966 | 1 | Statistics Netherlands |
| Spain | 1900 - 1981 | | 5 | National Statistic Institute (INE) |
| | 1991 - 2011 | | 5 | UN |
| Sweden | 1949 - 1967 | | 5 | UN |
| | 1968 - 2015 | 1953 - 1966 | 1 | Statistics Sweden |
| Switzerland | 1950 - 2015 | | 5 | UN |
| UK | 1971 - 2001 | | 1 | Office for National Statistics |
| | 2011 | | 5 | UN |

*Source*: Czech Republic: population and housing census (https://www.czso.cz/csu/czso/home)
Denmark: population register (http://www.statbank.dk)
Germany: microcensus (https://www-genesis.destatis.de/genesis/online)
Ireland: decennial census (http://www.cso.ie/en/databases)
The Netherlands: population register (https://opendata.cbs.nl/dataportaal#/CBS/nl/)
Spain: decennial census (http://www.ine.es/en/welcome.shtml)
Sweden: population register (http://www.scb.se)
The UK: estimation from decennial census (https://www.ons.gov.uk)
UN (http://data.un.org)

## 2.3 Illustration of EYEM

### 2.3.1 The results of period data

Figure 2.3 presents the time trends of period EYEM for the selected countries.

The changes in period EYEM show similar patterns for females and males, albeit with lower levels for males. In the remainder of this article we focus on the results for females, but results for males are available in Appendix D. There are three patterns in terms of the timing of reduction in period EYEM. The first group, which contains Canada, Denmark, and Sweden, experienced a decrease in their period EYEM by 1970. This group can be categorised as the North European and Canadian pattern. Austria, Belgium, France, Germany, the Netherlands, Switzerland, and the UK belong to the second group, which started reducing between 1970 and 1980, and can be categorised as the West European pattern. Finally, the Czech Republic, Greece, Ireland, Italy, and Spain constitute the South-East European pattern with a declining period EYEM starting after 1980. Nevertheless, the variability from country to country is present in all groups. For example, in recent years, females from Denmark, France, Germany, Ireland, the Netherlands, and Sweden have less than 15 years of period EYEM, while the other countries have more than 17 years. As seen in Figure 2.3, period EYEM started decreasing in the 1970s. Hence, we decompose period EYEM from 1970, and the results are presented in Figure 2.4.

Overall, location is the most influential factor in the changes in period EYEM. This shows that delayed marriage is the main contributor to nuptiality changes in most countries and periods. The scale factor also has an important role in the changes in period EYEM. Sweden had a negative effect (contributing to the decline) of the scale component from 1985; later, Denmark, France, Germany, and Switzerland had it from 1990, the Netherlands from 1995, and Italy and the UK from 2000. A negative effect of the scale factor means that the decline in proportion of marriages contributed to the decline in period EYEM. In Sweden, the decline of period EYEM is 28.1% due to non-marriage and 71.9% due to delayed marriage, from 1990

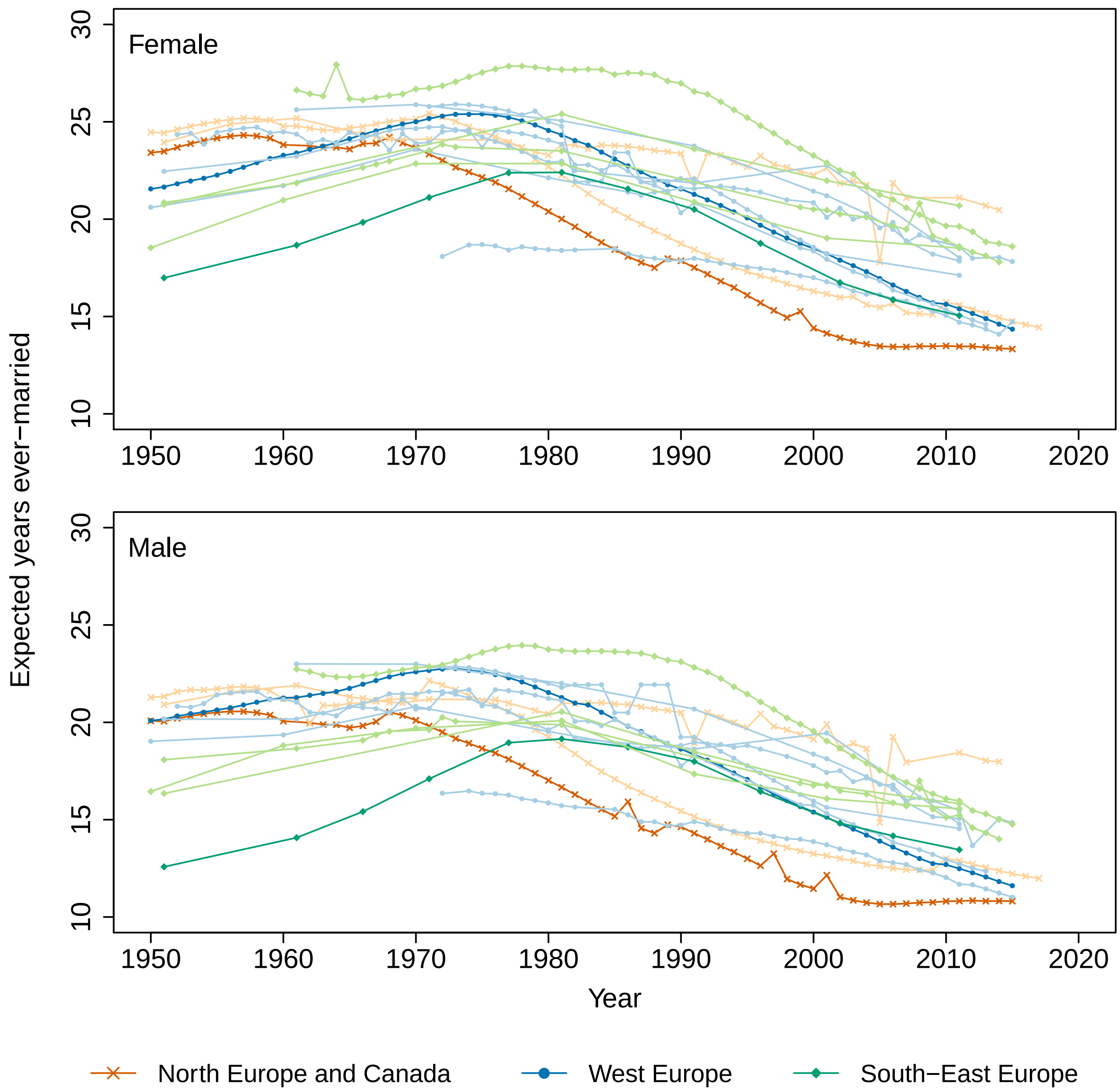



Figure 2.3: Time trends in period expected years ever-married in 15 countries
*Note*: North Europe and Canada comprises Canada, Denmark, and Sweden (highlighted). West Europe includes Austria, Belgium, France, Germany, the Netherlands (highlighted), Switzerland, and the UK. South-East Europe represents the Czech Republic, Greece, Ireland (highlighted), Italy, and Spain.
*Source*: Authors' calculations, using data described in Table 2.1.

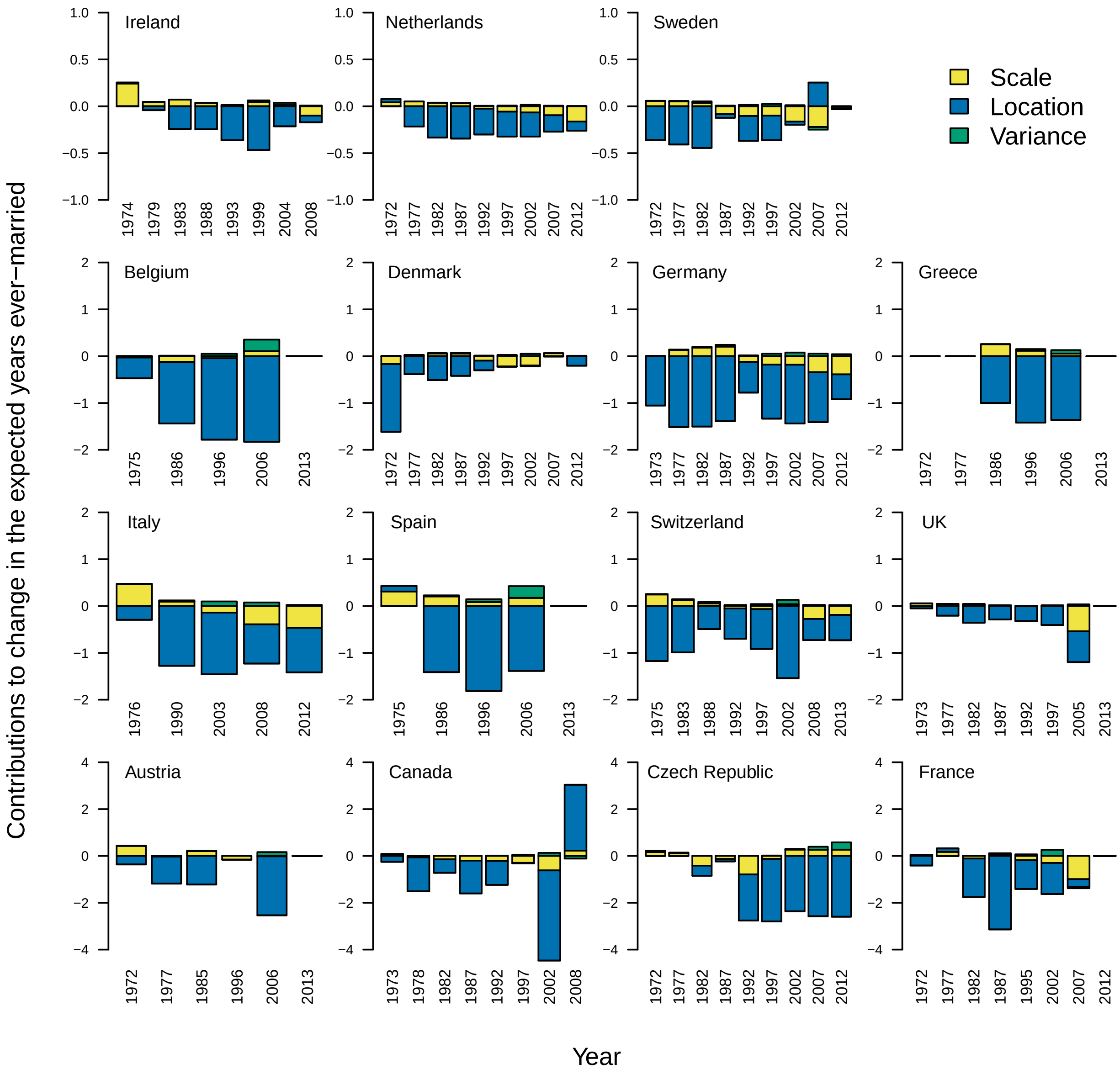



Figure 2.4: Decomposition of the change over time in female period expected years ever-married in 15 countries

*Note*: The year presented corresponds to the mid-year between two points in times. For example, for the changes in period EYEM from 1970 to 1975, it is written as 1972. Detailed information can be found in Appendix D.

*Source*: Authors' calculations, using data described in Table 2.1.

to 1995[3]. However, it reversed from 2000 to 2005, when non-marriage contributed 82.8% to the decline of period EYEM and delayed marriage contributed 17.2% (See Table 2.2). In the period 2005 to 2010, the two components have opposing contributions. While most of the North and West European countries and Canada had negative scale and location effects, the scale factor has not started contributing enough to this decline in Austria, Belgium, Greece, and Spain. This shows that, in the latter group of countries, the main nuptiality change was delayed marriage. Lastly, the variance has not had much impact on the changes in period EYEM.

Table 2.2: Contribution of scale, location, and variance to the change in females' period expected years ever-married (${}_{35}\dot{e}_{M,15}(t)$) in Sweden, 1970 to 2015

| Year | Mid-year | ${}_{35}\dot{e}_{M,15}(t)$ | Scale | Location | Variance | Sum of all components |
|---|---|---|---|---|---|---|
| 1970 - 1975 | 1972 | −0.302 | 0.057 | −0.361 | 0.001 | −0.303 |
| 1975 - 1980 | 1977 | −0.351 | 0.052 | −0.408 | 0.005 | −0.351 |
| 1980 - 1985 | 1982 | −0.391 | 0.037 | −0.445 | 0.015 | −0.392 |
| 1985 - 1990 | 1987 | −0.115 | −0.084 | −0.039 | 0.008 | −0.115 |
| 1990 - 1995 | 1992 | −0.354 | −0.104 | −0.266 | 0.015 | −0.355 |
| 1995 - 2000 | 1997 | −0.338 | −0.100 | −0.263 | 0.025 | −0.338 |
| 2000 - 2005 | 2002 | −0.186 | −0.164 | −0.034 | 0.012 | −0.186 |
| 2005 - 2010 | 2007 | 0.004 | −0.221 | 0.253 | −0.028 | 0.004 |
| 2010 - 2015 | 2012 | −0.032 | −0.020 | 0.000 | −0.012 | −0.032 |

*Note*: The sum of all components (Scale, Location, and Variance) varies slightly from the difference in the expected years ever-married (${}_{35}\dot{e}_{M,15}(t)$), due to rounding the numbers to the third decimal point in the table.
*Source*: Authors' calculations, using data described in Table 2.1.

However, caution is warranted in the interpretation of the results. Similar to period life expectancy, which corresponds to the mortality experience of a synthetic cohort, period EYEM is also an index combining the information of many cohorts. As previous research has stated, a period index is biased by tempo effects (Winkler-Dworak and Engelhardt 2004; Schoen and Canudas-Romo 2005; Bongaarts and Feeney 2006), and period EYEM could also be affected. Thus, the next section presents the changes in cohort EYEM over time.

[3]The percentages are calculated among negative values. For instance, the percentage of the contribution of scale from 1990 to 1995 (28.1%) is computed as $0.104/(0.104 + 0.266)$

### 2.3.2 The results of cohort data

As the results for period data, the changes in cohort EYEM present similar trends for females and males (Figure 2.5). For all countries, males have smaller cohort EYEM, which means that males spend relatively longer periods in never-married status. North Europe and Canada, which comprises Canada, Denmark, and Sweden, have a declining trend in all cohorts analysed. The Netherlands increased its cohort EYEM until the late 1940s birth cohort and decreased thereafter, while the Czech Republic shows an almost stagnating high EYEM trend.

Figure 2.6 presents the results of decomposing the changes over time in the female cohort EYEM. Compared to the results for the period data, the cohort results illustrate more diversity in trends. The decline of cohort EYEM in Canada was mainly a non-marriage effect until the 1955 birth cohort. Then delayed marriage became the main factor. The most recent Canadian cohort had a positive scale factor. It means that the female 1965 birth cohort got married more than the 1960 birth cohort, although the delayed process more than offset this. Similarly, the scale factor had a positive effect for the youngest Danish cohort, although the location factor was the main effect. For Sweden, the scale factor made a relatively large contribution to the decline in cohort EYEM compared with to the recent cohorts of other countries. West European countries followed a similar pattern, which the location factor reduced cohort EYEM mainly while the Netherlands had a positive location effect between 1935 and 1950 birth cohort (earlier marriage) and the scale factor became the main contributor from 1945 to 1955 birth cohort. One thing should be mentioned from the cohort results to the period results. The large location effect in the cohort EYEM implies that the period results may be affected more by tempo distortions.

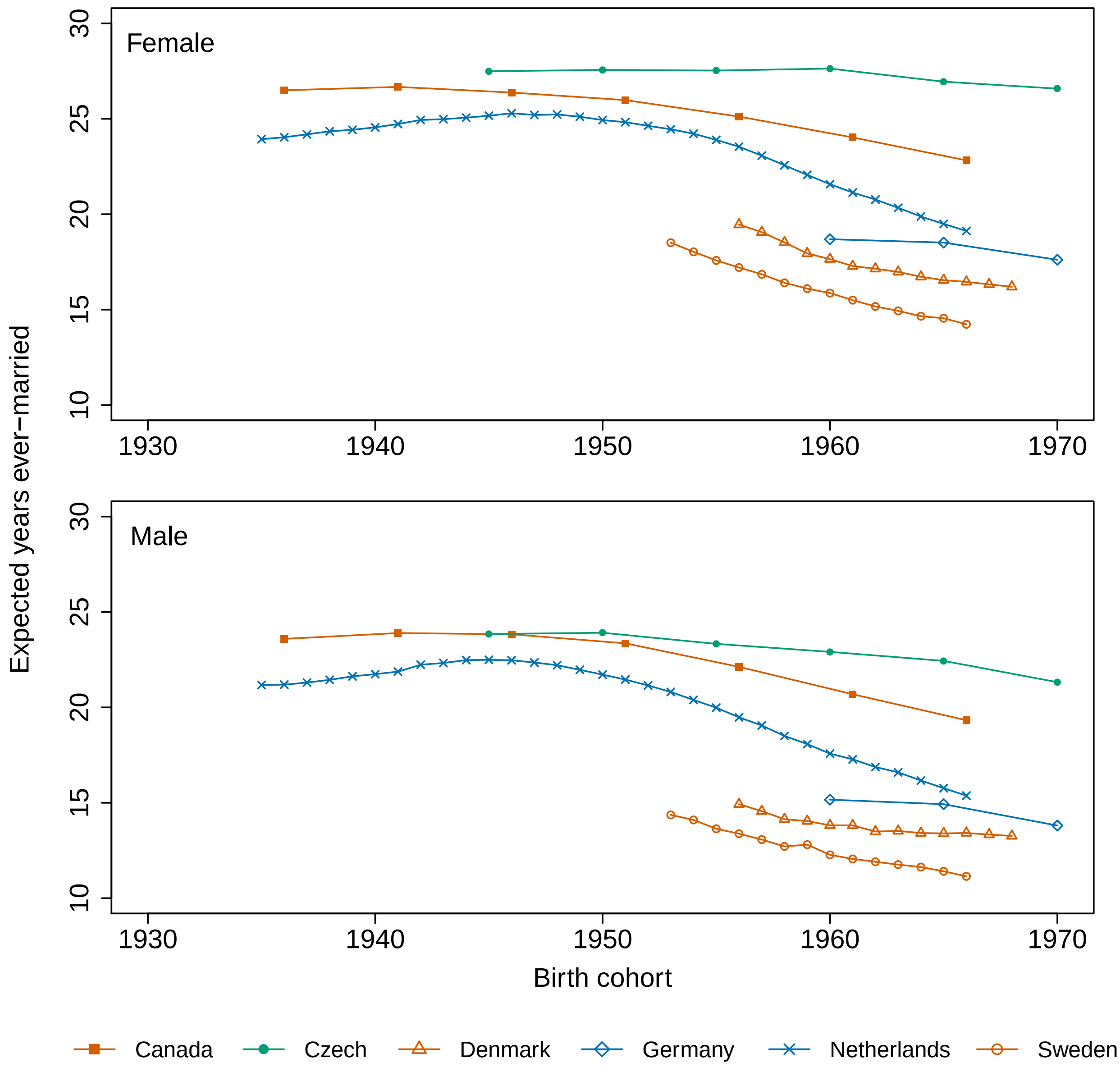



Figure 2.5: Time trends in cohort expected years ever-married in six countries
*Source*: Authors' calculations, using data described in Table 2.1.

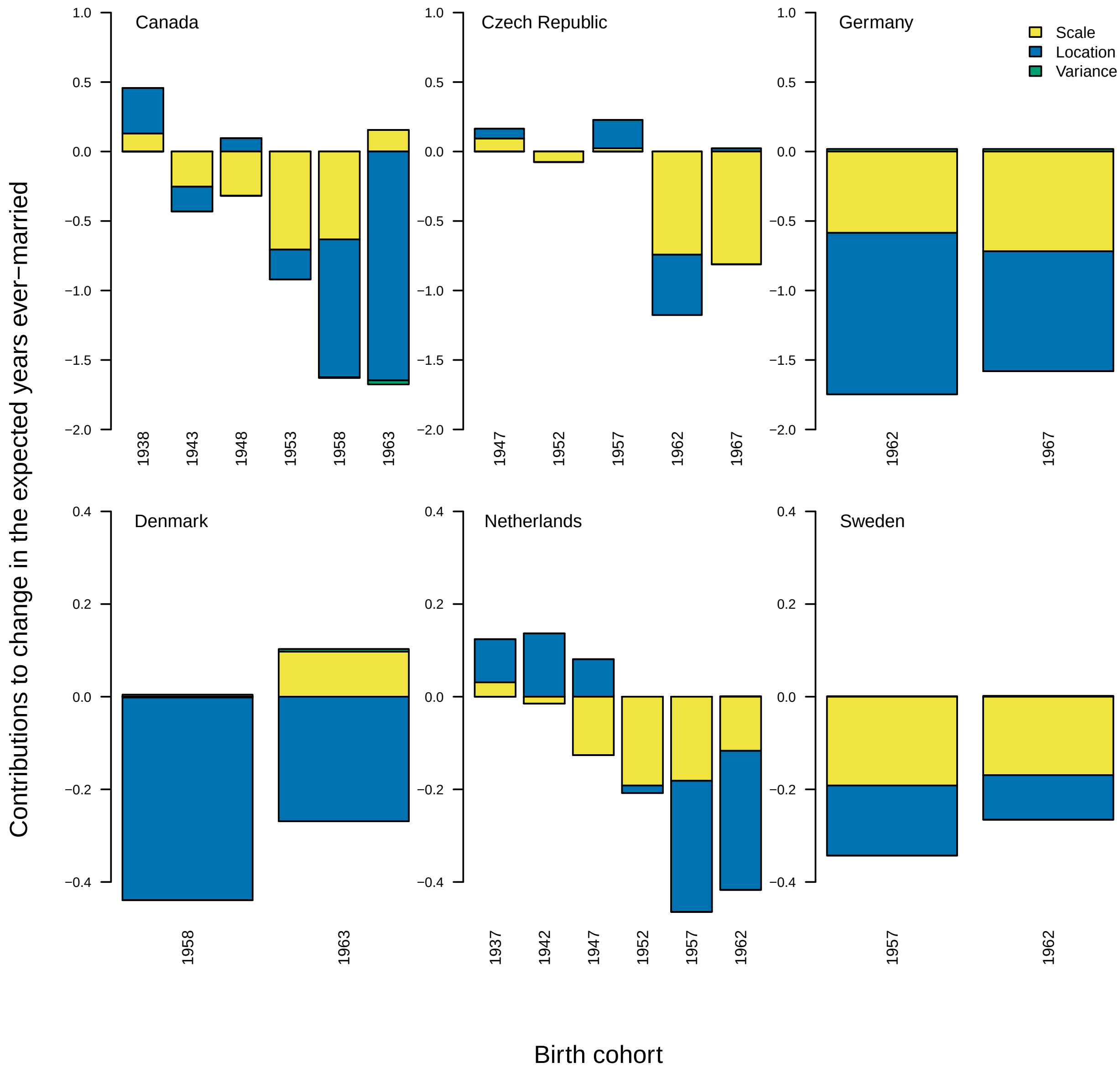



Figure 2.6: Decomposition of the change over time in female cohort expected years ever-married in six countries

*Note*: The birth cohort presented corresponds to the mid-year between two points in birth cohorts. For example, for the changes in EYEM from 1950 to 1955, it is written as 1952. Detailed information can be found in Appendix D.

*Source*: Authors' calculations, using data described in Table 2.1.

## 2.4 Discussion and conclusion

Non-marriage, delayed marriage and expansion of first marriage timing are well reported and described as changes that happened in the second half of the twentieth century (Lesthaeghe 1983; Van de Kaa 1987; Winkler-Dworak and Engelhardt 2004; Elzinga and Liefbroer 2007; European Commission 2015). In this article, we used the expected years ever-married (EYEM) as a new alternative index to quantify nuptiality change and propose its decomposition into the three aforementioned components. Examining both period and cohort data allows us to study the changes in EYEM from both complementary perspectives. Period EYEM decreased from 1970, and the trends of changes of period EYEM are similar for males and females in the studied countries. Our results suggest that, in most countries and time periods, the decline in period EYEM is mainly due to delayed marriage. This result is consistent with other research that has analysed the US trend (Oppenheimer 1994; Goldstein and Kenney 2001). However, new trends can be seen in our selected countries, with the non-marriage component influencing recently in Northern Europe, Canada and in most West European countries. The expansion effect has practically no influence on the changes in EYEM. Similar to the period EYEM trends, the trends of males' cohort EYEM are similar to those observed for females, but with different scales. The decline of the current cohort EYEM in Canada, Denmark, Germany, and the Netherlands is mainly due to delayed marriage, while non-marriage was the main factor in Canada and the Netherlands in older cohorts. On the other hand, non-marriage influenced just over half of the changes in cohort EYEM of Sweden.

Period measures are an aggregation of different cohorts and are affected by the changes in cohorts measures. This is also the case in EYEM and our results highlight some of the cohort effects in the periods results. Hence, the recent increase in non-marriage component in period EYEM may be partially explained by the delayed marriage effects in cohort EYEM, especially observed in the Netherlands. Quantifying how much the decomposition results of period EYEM are affected by cohort EYEM is beyond the scope of the present study.

However, this suggests a new area of research on how the decomposition of period measures and the decomposition of cohort measures interconnect.

The limitations of this study should be mentioned. First, our data does not include cohabiting couples' information nor socio-economic status, such as educational level. The latter has an important impact on marriage decision and its timing (e.g. Blossfeld et al. 2005). One could speculate that the rise in the scale factor contribution in recent years indirectly shows the increase in cohabitation. It is possible to hypothesise that people have tended to choose cohabitation as their style of union formation, and that is the reason for the recent negative contribution of non-marriage in the Northern Europe and Canada and in most of the West European countries. This is also found in cohort analysis in Germany and Sweden. However, due to data limitations, this study could not test this hypothesis. The second limitation corresponds to the well-known problem of fitting observed data to the Coale-McNeil model. This issue is particularly seen in some countries and times when the population consists of subpopulations whose first marriage timings are distinctly different from each other (Kaneko 2003; Peristera and Kostaki 2015). However, if data for those subpopulations were available, our decomposition method could be extended to also cover these cases. Hence, subpopulation analysis could increase the preciseness of nuptiality modeling, and future research might benefit from looking at the effects of scale, location, and variance on the changes in EYEM by subpopulations.

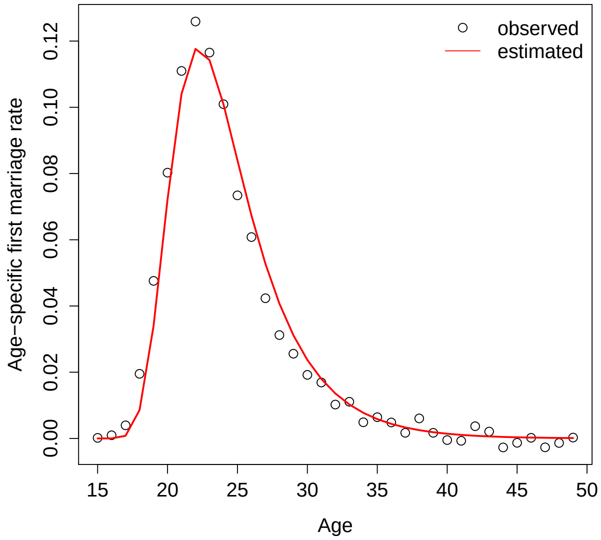
Finally, EYEM measures the expected number of years after first marriage. Therefore, it does not take into account exits from marriage (i.e. divorce/separation, widowhood, or death). This study, however, focuses on the transition from never-married to ever-married status. For this reason, we introduced EYEM as an alternative index to study nuptiality changes. If the interest is to quantify the duration of first marriage until divorce, widowhood, or death, such as seen in Schoen and Nelson (1974) and Philipov and Jasilioniene (2008), one must consider exits from first marriage.

Which of the three effects, non-marriage, delayed marriage, or expansion, has the most impact on nuptiality changes? How does the most influential factor differ by time periods, birth cohort and countries? This study approaches those questions by introducing a new index and decomposing its change into the contribution of each of those three components. By examining both period and cohort data, we present a full view of the changes in first marriage behaviours through Europe and Canada. The decomposition steps presented in equations (2.1) to (2.7) offer an open possibility for more elaborated parametric marriage models. Nuptiality dynamics keep evolving and researchers would benefit from analysing future changes by using the methods developed here.

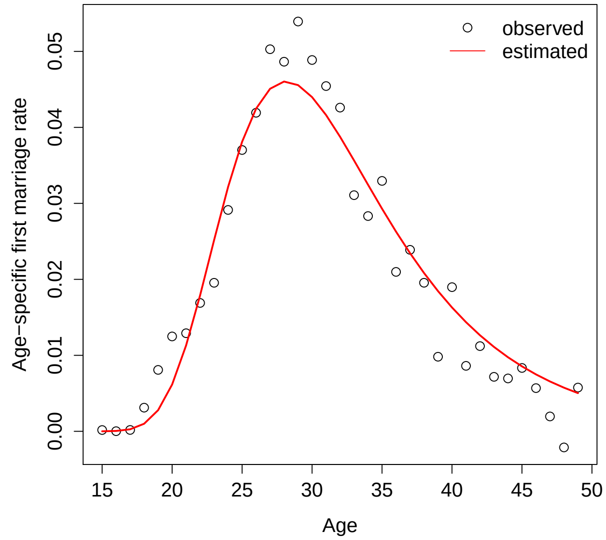
# Appendices

## Appendix 2.A The comparison between the observed and the estimated age-specific first marriage rate

As Kaneko (2003) and Peristera and Kostaki (2015) pointed out, CM model may not fit well to some countries and some time periods. If CM model can not capture the observed rate, the presented results will be misleading. Thus, we compared the observed age-specific first marriage rate with the estimated one. CM model generally estimates quite well to our selected data especially countries that have single age-groups, even though CM model tends to underestimate the maximum value. For the countries that do not have single age data, CM model does not fit as well as for the other countries. As Figure 2.A.1 and 2.A.2 show, the estimated rates have only slightly different scale and location from the observed data, which would not make our conclusion deviate from the findings presented here. Furthermore, as mentioned earlier in the main text, our methodology can adapt to other parametric formulations of the age-patterns of marriage.

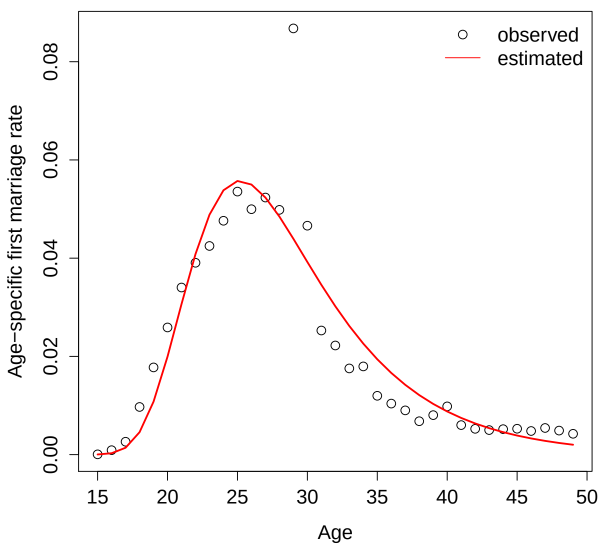



(a) Year 1970

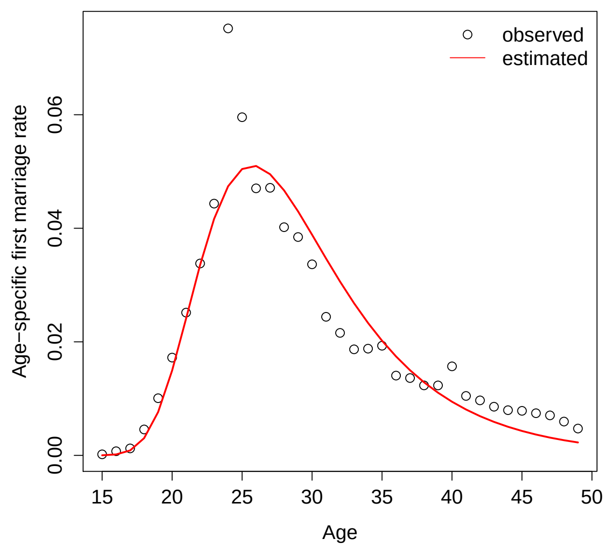



(b) Year 2015

Figure 2.A.1: Comparison between female's observed and estimated period age-specific first marriage rates for Sweden
*Source*: Authors' calculations using data described in Table 1 of the main text.




(a) 1960 birth cohort




(b) 1965 birth cohort

Figure 2.A.2: Comparison between female's observed and estimated cohort age-specific first marriage rates for Sweden
*Note*: There is a heap because cohort data is constructed from period data without smoothing.
*Source*: Authors' calculations using data described in Table 1 of the main text.

# Appendix 2.B Calculation process: Expected years ever-married

We denote ${}_{35}e_{M,15}(t)$ as the expected years ever-married from age 15 to age 50. It is formulated as:

$$\begin{aligned} {}_{35}e_{M,15}(t) &= \int_{15}^{50} (1 - l_{x,t})\, dx \\ &= \int_{15}^{50} F_{x,t}\, dx \end{aligned} \tag{2.8}$$

where $l_x$ is a probability of remaining never-married and $F_x$ is its cumulative probability function. We use Rodríguez and Trussell's parametrisation (Rodríguez and Trussell 1980) for the density function:

$$f_{x,t} = C_t \frac{1}{\sigma_t} a_1 \exp\Big[a_2(\frac{x-\mu_t}{\sigma_t} + a_3) - \exp\{-a_4(\frac{x-\mu_t}{\sigma_t} + a_3)\}\Big] \tag{2.9}$$

$$f_{x,t} = C_t \frac{1}{\sigma_t} f_0(\frac{x-\mu_t}{\sigma_t}),$$

where the usual values for the constants are $a_1 = 1.281$, $a_2 = -1.145$, $a_3 = 0.805$, and $a_4 = 1.896$. $f_0$ is the density function defined from equation (2.9) as

$$f_0(x) = a_1 \exp\Big[a_2(x + a_3) - \exp\{-a_4(x + a_3)\}\Big]. \tag{2.10}$$

Its cumulative density function is written as

$$F_{x,t} = C_t F_0(\frac{x-\mu_t}{\sigma_t}), \tag{2.11}$$

and substituting equation (2.11) in equation (2.8) results in an expression of EYEM that depends on our three variables of interest (scale, location, and variance) as

$$ {}_{35}e_{M,\,15}(t) \;=\; \int_{15}^{50} C_t F_0(\frac{x-\mu_t}{\sigma_t})\, dx. \tag{2.12} $$

To quantify the effects of scale, location, and variance in the changes of EYEM over time, the partial derivative respect to time of the probability distribution in equation (2.12) is studied. Let a dot on top of a variable denote its partial derivative respect to time. The change over time in EYEM, or , ${}_{35}\dot{e}_{M,\,15}(t)$, is decomposed as:

$$ {}_{35}\dot{e}_{M,\,15}(t) = \frac{\partial\, {}_{35}e_{M,\,15}(t)}{\partial C_t}\dot{C}_t + \frac{\partial\, {}_{35}e_{M,\,15}(t)}{\partial \mu_t}\dot{\mu}_t + \frac{\partial\, {}_{35}e_{M,\,15}(t)}{\partial \sigma_t}\dot{\sigma}_t, \tag{2.13} $$

where each term is the change in ${}_{35}\dot{e}_{M,\,15}(t)$ resulting from changes in the scale, location, and variance, respectively.

The derivative of $F_0(\frac{x-\mu_t}{\sigma_t})$ respect to time $t$ is

$$ \begin{aligned} \dot{F}_0(\frac{x-\mu_t}{\sigma_t}) &= f_0(\frac{x-\mu_t}{\sigma_t})(\frac{\frac{d}{dt}(x-\mu_t)\sigma_t - (x-\mu_t)\dot{\sigma}_t}{\sigma_t^2}) \\ &= -\frac{1}{\sigma_t} f_0(\frac{x-\mu_t}{\sigma_t})\dot{\mu}_t - f_0(\frac{x-\mu_t}{\sigma_t})\frac{(x-\mu_t)}{\sigma_t^2}\dot{\sigma}_t, \end{aligned} $$

substituting this in equation (2.12) helps obtaining the time derivative of EYEM as

$$ {}_{35}\dot{e}_{M,\,15}(t) = \dot{C}_t \int_{15}^{50} F_0(\frac{x-\mu_t}{\sigma_t})\, dx - \dot{\mu}_t \int_{15}^{50} C_t \frac{1}{\sigma_t}\, f_0(\frac{x-\mu_t}{\sigma_t})\, dx - \dot{\sigma}_t \int_{15}^{50} C_t\, f_0(\frac{x-\mu_t}{\sigma_t})\frac{x-\mu_t}{\sigma_t^2}\, dx. \tag{2.14} $$

Therefore, the changes of each factor is expressed as

$$ \frac{\partial\, {}_{35}e_{M,\,15}(t)}{\partial C_t}\dot{C}_t = \dot{C}_t \int_{15}^{50} F_0(\frac{x-\mu_t}{\sigma_t})\, dx, \tag{2.15} $$

for declines (increases) in the proportion ever marrying, or scale effect which contributes to the decline (increase) in the overall EYEM. The second term is

$$\begin{aligned}
\frac{\partial\, {}_{35}e_{M,15}(t)}{\partial \mu_t}\dot{\mu}_t &= -\dot{\mu}_t \int_{15}^{50} C_t \frac{1}{\sigma_t}\, f_0(\frac{x-\mu_t}{\sigma_t})\, dx \\
&= -\dot{\mu}_t [F_{50,\,t} - F_{15,\,t}] \\
&= -\dot{\mu}_t F_{50,\,t} \\
&= -\dot{\mu}_t C_t, \qquad (2.16)
\end{aligned}$$

corresponding to the changes in the mean age at first marriage between ages 15 and 50. For all the cases when the mean age at first marriage has been increasing over time this term contributes negatively to the overall change in EYEM. Finally, the contribution of the standard deviation term is

$$\begin{aligned}
\frac{\partial\, {}_{35}e_{M,15}(t)}{\partial \sigma_t}\dot{\sigma}_t &= -\dot{\sigma}_t \int_{15}^{50} C_t\, f_0(\frac{x-\mu_t}{\sigma_t})\frac{x-\mu_t}{\sigma_t^2}\, dx \\
&= -\dot{\sigma}_t \int_{15}^{50} f_{x,\,t} \frac{x-\mu_t}{\sigma_t}\, dx, \qquad (2.17)
\end{aligned}$$

which has negligible contribution in the cases studied here and presented in Tables 2.2 to 2.D.3.

# Appendix 2.C The decomposition to discrete data

The three parameters of $f_x$ are estimated using Maximum Likelihood Estimation method as suggested by Rodríguez and Trussell (1980).

$$\text{lnLH} = \sum_{15}^{50}(\text{Mar}_x \log[F_{(x+0.5)}] + \text{NMar}_x \log[1 - F_{(x+0.5)}]), \tag{2.18}$$

where $\text{Mar}_x$ is ever-married population at age $x$ and $\text{NMar}_x$ is never-married population at age $x$, and $F_x$ is the cumulative probability function at age $x$. We checked the validity of this estimation method to five year age-group data. The sensitivity analysis consisted on changing the single age-groups to five year age-groups and showed that the parameters were still well estimated, although at different levels, but the time trends were preserved. Our assessment confirmed this and our results might be overestimated for the countries that have only five year age-groups (see Figure 2.C.1 below). As age-groups do not influence the components' time trends and their relative contribution to change in EYEM, it is likely that age-group did not affect our overall conclusions.

We followed Vaupel and Canudas-Romo (2003) of applying the continuous decomposition equation to discrete time data. To apply our decomposition method to discrete time data, each function is estimated at their midpoint over a time interval (Preston, Heuveline and Guillot 2001). For the functions except EYEM, an exponential change assumption is used.

$$\upsilon_{x,\,t+\frac{h}{2}} = \upsilon_{x,\,t}(\frac{\upsilon_{x,\,t+h}}{\upsilon_{x,\,t}})^{0.5} \tag{2.19}$$

The derivative of the function $\upsilon_{x,\,t+\frac{h}{2}}$ was estimated by

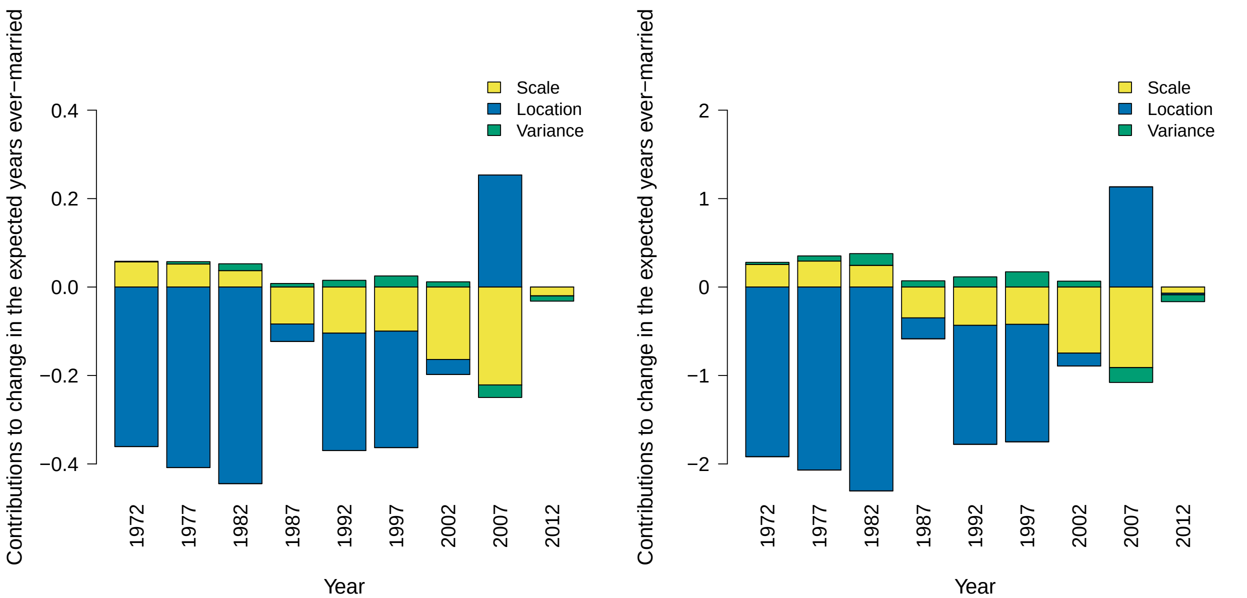
$$\dot{\upsilon}_{x,\,t+\frac{h}{2}} = \upsilon_{x,\,t+\frac{h}{2}}(\frac{\log[\frac{\upsilon_{x,\,t+h}}{\upsilon_{x,\,t}}]}{h}). \tag{2.20}$$

EYEM was assumed to have a linear change in the interval and its midpoint was calculated as

$$\upsilon_{x,\,t+\frac{h}{2}} = \frac{\upsilon_{x,\,t+h} + \upsilon_{x,\,t}}{2}, \tag{2.21}$$

and

$$\dot{\upsilon}_{x,\,t+\frac{h}{2}} = \frac{\upsilon_{x,\,t+h} - \upsilon_{x,\,t}}{h}. \tag{2.22}$$




(a) Single age-groups (b) Five year age-groups

Figure 2.C.1: Comparison between the decomposition results using single age-groups and five year age-groups for Sweden
*Source*: Authors' calculations using data described in Table 1 of the main text.

# Appendix 2.D The results of decomposition for males

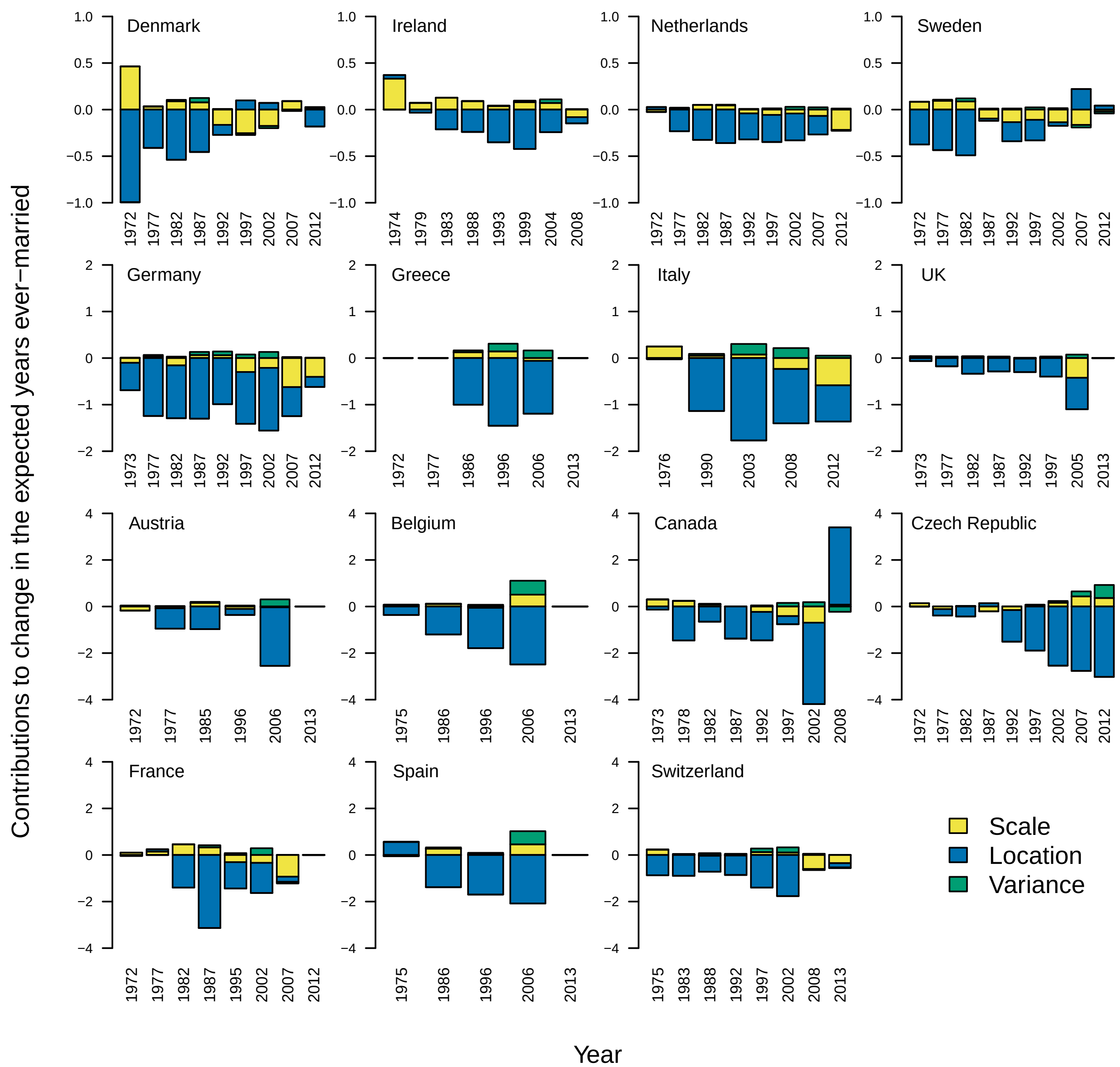



Figure 2.D.1: Decomposition of the change over time in the male period expected years ever-married in 15 countries

*Note*: The year presented corresponds to the mid-year between two times. For example, for the changes in period EYEM from 1970 to 1975, it is written as 1972. The detail information can be seen in Table 2.D.1 in this supplemental material.

*Source*: Authors' calculations using data described in Table 1 of the main text.

Table 2.D.1: The contribution of each component to the time change in the female period expected years ever-married

| Country | Year | Mid-year | Scale | Location | Variance | Country | Year | Mid-year | Scale | Location | Variance |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Austria | 1970-1975 | 1972 | 0.429 | −0.364 | 0.000 | Ireland | 1971-1977 | 1974 | 0.242 | 0.012 | 0.000 |
| | 1975-1980 | 1977 | −0.030 | −1.153 | 0.005 | | 1977-1981 | 1979 | 0.047 | −0.042 | 0.000 |
| | 1980-1991 | 1985 | 0.208 | −1.222 | 0.021 | | 1981-1986 | 1983 | 0.071 | −0.243 | 0.000 |
| | 1991-2001 | 1996 | −0.156 | −0.004 | 0.005 | | 1986-1991 | 1988 | 0.036 | −0.246 | 0.001 |
| | 2001-2011 | 2006 | −0.018 | −2.518 | 0.165 | | 1991-1996 | 1993 | 0.011 | −0.363 | 0.003 |
| | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 | | 1996-2002 | 1999 | 0.048 | −0.467 | 0.015 |
| Belguim | 1970-1981 | 1975 | −0.030 | −0.444 | 0.001 | | 2002-2006 | 2004 | 0.013 | −0.214 | 0.024 |
| | 1981-1991 | 1986 | −0.123 | −1.315 | 0.011 | | 2006-2011 | 2008 | −0.100 | −0.071 | 0.008 |
| | 1991-2001 | 1996 | −0.045 | −1.739 | 0.051 | | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 |
| | 2001-2011 | 2006 | 0.109 | −1.827 | 0.245 | Italy | 1971-1981 | 1976 | 0.471 | −0.294 | 0.001 |
| | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 | | 1981-2000 | 1990 | 0.095 | −1.276 | 0.026 |
| Canada | 1971-1976 | 1973 | 0.083 | −0.252 | 0.000 | | 2000-2006 | 2003 | −0.140 | −1.316 | 0.098 |
| | 1976-1980 | 1978 | −0.064 | −1.449 | 0.004 | | 2006-2010 | 2008 | −0.391 | −0.838 | 0.077 |
| | 1980-1985 | 1982 | −0.143 | −0.578 | 0.008 | | 2010-2014 | 2012 | −0.465 | −0.951 | 0.024 |
| | 1985-1990 | 1987 | −0.201 | −1.401 | 0.004 | Netherlands | 1970-1975 | 1972 | 0.042 | 0.037 | 0.000 |
| | 1990-1995 | 1992 | −0.215 | −1.019 | 0.015 | | 1975-1980 | 1977 | 0.051 | −0.217 | 0.000 |
| | 1995-2000 | 1997 | −0.295 | −0.022 | 0.054 | | 1980-1985 | 1982 | 0.038 | −0.334 | 0.001 |
| | 2000-2005 | 2002 | −0.619 | −3.850 | 0.130 | | 1985-1990 | 1987 | 0.033 | −0.345 | 0.004 |
| | 2005-2011 | 2008 | 0.224 | 2.816 | −0.114 | | 1990-1995 | 1992 | −0.026 | −0.276 | 0.005 |
| Czech Republic | 1970-1975 | 1972 | 0.168 | 0.059 | 0.000 | | 1995-2000 | 1997 | −0.058 | −0.266 | 0.008 |
| | 1975-1980 | 1977 | 0.107 | 0.032 | 0.000 | | 2000-2005 | 2002 | −0.067 | −0.257 | 0.017 |
| | 1980-1985 | 1982 | −0.416 | −0.431 | 0.000 | | 2005-2010 | 2007 | −0.096 | −0.175 | 0.007 |
| | 1985-1990 | 1987 | −0.128 | −0.107 | 0.000 | | 2010-2015 | 2012 | −0.163 | −0.098 | 0.002 |
| | 1990-1995 | 1992 | −0.792 | −1.965 | 0.005 | Spain | 1970-1981 | 1975 | 0.310 | 0.120 | 0.001 |
| | 1995-2000 | 1997 | −0.127 | −2.671 | 0.017 | | 1981-1991 | 1986 | 0.206 | −1.412 | 0.022 |
| | 2000-2005 | 2002 | 0.272 | −2.363 | 0.031 | | 1991-2001 | 1996 | 0.087 | −1.815 | 0.058 |
| | 2005-2010 | 2007 | 0.263 | −2.575 | 0.131 | | 2001-2011 | 2006 | 0.175 | −1.385 | 0.250 |
| | 2010-2015 | 2012 | 0.270 | −2.596 | 0.311 | | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 |
| Denmark | 1970-1975 | 1972 | −0.169 | −1.447 | 0.005 | Sweden | 1970-1975 | 1972 | 0.057 | −0.361 | 0.001 |
| | 1975-1980 | 1977 | 0.023 | −0.385 | 0.001 | | 1975-1980 | 1977 | 0.052 | −0.408 | 0.005 |
| | 1980-1985 | 1982 | 0.058 | −0.511 | 0.006 | | 1980-1985 | 1982 | 0.037 | −0.445 | 0.015 |
| | 1985-1990 | 1987 | 0.053 | −0.420 | 0.022 | | 1985-1990 | 1987 | −0.084 | −0.039 | 0.008 |
| | 1990-1995 | 1992 | −0.096 | −0.205 | 0.013 | | 1990-1995 | 1992 | −0.104 | −0.266 | 0.015 |
| | 1995-2000 | 1997 | −0.221 | 0.025 | −0.005 | | 1995-2000 | 1997 | −0.100 | −0.263 | 0.025 |
| | 2000-2005 | 2002 | −0.199 | 0.050 | −0.017 | | 2000-2005 | 2002 | −0.164 | −0.034 | 0.012 |
| | 2005-2010 | 2007 | 0.063 | −0.006 | −0.005 | | 2005-2010 | 2007 | −0.221 | 0.253 | −0.028 |
| | 2010-2015 | 2012 | −0.001 | −0.203 | 0.008 | | 2010-2015 | 2012 | −0.020 | 0.000 | −0.012 |
| France | 1970-1975 | 1972 | 0.046 | −0.410 | 0.001 | Switzerland | 1970-1980 | 1975 | 0.247 | −1.174 | 0.010 |
| | 1975-1980 | 1977 | 0.172 | 0.147 | 0.001 | | 1980-1986 | 1983 | 0.129 | −0.991 | 0.018 |
| | 1980-1985 | 1982 | −0.111 | −1.645 | 0.012 | | 1986-1990 | 1988 | 0.054 | −0.493 | 0.038 |
| | 1985-1990 | 1987 | 0.065 | −3.139 | 0.051 | | 1990-1995 | 1992 | −0.051 | −0.648 | 0.023 |
| | 1990-2000 | 1995 | −0.184 | −1.226 | 0.068 | | 1995-2000 | 1997 | −0.066 | −0.850 | 0.043 |
| | 2000-2005 | 2002 | −0.300 | −1.325 | 0.263 | | 2000-2005 | 2002 | 0.041 | −1.540 | 0.092 |
| | 2005-2010 | 2007 | −0.991 | −0.327 | −0.061 | | 2005-2011 | 2008 | −0.276 | −0.451 | 0.025 |
| | 2010-2015 | 2012 | 0.000 | 0.000 | 0.000 | | 2011-2015 | 2013 | −0.190 | −0.540 | 0.022 |
| Germany | 1972-1975 | 1973 | 0.002 | −1.054 | 0.002 | UK | 1971-1975 | 1973 | 0.057 | −0.050 | 0.000 |
| | 1975-1980 | 1977 | 0.134 | −1.514 | 0.007 | | 1975-1980 | 1977 | 0.044 | −0.205 | 0.000 |
| | 1980-1985 | 1982 | 0.184 | −1.503 | 0.017 | | 1980-1985 | 1982 | 0.041 | −0.357 | 0.001 |
| | 1985-1990 | 1987 | 0.205 | −1.389 | 0.037 | | 1985-1990 | 1987 | 0.016 | −0.288 | 0.003 |
| | 1990-1995 | 1992 | −0.122 | −0.657 | 0.018 | | 1990-1995 | 1992 | −0.006 | −0.313 | 0.005 |
| | 1995-2000 | 1997 | −0.181 | −1.154 | 0.053 | | 1995-2000 | 1997 | −0.001 | −0.404 | 0.017 |
| | 2000-2005 | 2002 | −0.184 | −1.255 | 0.077 | | 2000-2011 | 2005 | −0.538 | −0.659 | 0.037 |
| | 2005-2010 | 2007 | −0.342 | −1.067 | 0.055 | | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 |
| | 2010-2015 | 2012 | −0.389 | −0.530 | 0.042 | | | | | | |
| Greece | 1970-1975 | 1972 | 0.000 | 0.000 | 0.000 | | | | | | |
| | 1975-1980 | 1977 | 0.000 | 0.000 | 0.000 | | | | | | |
| | 1981-1991 | 1986 | 0.256 | −0.997 | −0.006 | | | | | | |
| | 1991-2001 | 1996 | 0.117 | −1.420 | 0.035 | | | | | | |
| | 2001-2011 | 2006 | 0.055 | −1.363 | 0.075 | | | | | | |
| | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 | | | | | | |

*Source*: Authors' calculations using data described in Table 1 of the main text.

Table 2.D.2: The contribution of each component to the time change in the male period expected years ever-married

| Country | Year | Mid-year | Scale | Location | Variance | Country | Year | Mid-year | Scale | Location | Variance |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Austria | 1970-1975 | 1972 | −0.173 | 0.039 | −0.006 | Ireland | 1971-1977 | 1974 | 0.332 | 0.039 | 0.000 |
| | 1975-1980 | 1977 | −0.076 | −0.874 | 0.021 | | 1977-1981 | 1979 | 0.072 | −0.033 | 0.000 |
| | 1980-1991 | 1985 | 0.157 | −0.972 | 0.043 | | 1981-1986 | 1983 | 0.127 | −0.212 | 0.000 |
| | 1991-2001 | 1996 | −0.103 | −0.258 | 0.042 | | 1986-1991 | 1988 | 0.091 | −0.240 | 0.001 |
| | 2001-2011 | 2006 | −0.034 | −2.513 | 0.307 | | 1991-1996 | 1993 | 0.038 | −0.351 | 0.005 |
| | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 | | 1996-2002 | 1999 | 0.080 | −0.422 | 0.016 |
| Belguim | 1970-1981 | 1975 | 0.073 | −0.362 | 0.002 | | 2002-2006 | 2004 | 0.072 | −0.242 | 0.038 |
| | 1981-1991 | 1986 | 0.111 | −1.199 | 0.015 | | 2006-2011 | 2008 | −0.081 | −0.066 | 0.006 |
| | 1991-2001 | 1996 | −0.056 | −1.731 | 0.076 | | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 |
| | 2001-2011 | 2006 | 0.511 | −2.487 | 0.598 | Italy | 1971-1981 | 1976 | 0.249 | −0.022 | −0.001 |
| | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 | | 1981-2000 | 1990 | 0.050 | −1.136 | 0.040 |
| Canada | 1971-1976 | 1973 | 0.305 | −0.132 | 0.004 | | 2000-2006 | 2003 | 0.077 | −1.769 | 0.226 |
| | 1976-1980 | 1978 | 0.241 | −1.458 | 0.009 | | 2006-2010 | 2008 | −0.234 | −1.166 | 0.215 |
| | 1980-1985 | 1982 | 0.090 | −0.651 | 0.029 | | 2010-2014 | 2012 | −0.586 | −0.777 | 0.053 |
| | 1985-1990 | 1987 | 0.000 | −1.374 | −0.011 | Netherlands | 1970-1975 | 1972 | −0.025 | 0.028 | 0.000 |
| | 1990-1995 | 1992 | −0.232 | −1.220 | 0.049 | | 1975-1980 | 1977 | 0.018 | −0.232 | 0.000 |
| | 1995-2000 | 1997 | −0.411 | −0.350 | 0.156 | | 1980-1985 | 1982 | 0.050 | −0.325 | 0.002 |
| | 2000-2005 | 2002 | −0.694 | −3.497 | 0.188 | | 1985-1990 | 1987 | 0.045 | −0.359 | 0.008 |
| | 2005-2011 | 2008 | 0.082 | 3.318 | −0.227 | | 1990-1995 | 1992 | −0.041 | −0.279 | 0.008 |
| Czech Republic | 1970-1975 | 1972 | 0.140 | −0.003 | 0.001 | | 1995-2000 | 1997 | −0.057 | −0.291 | 0.013 |
| | 1975-1980 | 1977 | −0.110 | −0.278 | 0.002 | | 2000-2005 | 2002 | −0.042 | −0.289 | 0.030 |
| | 1980-1985 | 1982 | 0.020 | −0.429 | 0.008 | | 2005-2010 | 2007 | −0.068 | −0.199 | 0.025 |
| | 1985-1990 | 1987 | −0.205 | 0.143 | 0.001 | | 2010-2015 | 2012 | −0.217 | 0.011 | −0.010 |
| | 1990-1995 | 1992 | −0.153 | −1.356 | 0.009 | Spain | 1970-1981 | 1975 | −0.047 | 0.559 | 0.009 |
| | 1995-2000 | 1997 | 0.070 | −1.894 | 0.005 | | 1981-1991 | 1986 | 0.273 | −1.385 | 0.053 |
| | 2000-2005 | 2002 | 0.161 | −2.540 | 0.077 | | 1991-2001 | 1996 | 0.022 | −1.701 | 0.071 |
| | 2005-2010 | 2007 | 0.437 | −2.764 | 0.214 | | 2001-2011 | 2006 | 0.459 | −2.082 | 0.563 |
| | 2010-2015 | 2012 | 0.368 | −3.024 | 0.558 | | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 |
| Denmark | 1970-1975 | 1972 | 0.463 | −0.995 | −0.001 | Sweden | 1970-1975 | 1972 | 0.083 | −0.375 | 0.002 |
| | 1975-1980 | 1977 | 0.033 | −0.411 | 0.003 | | 1975-1980 | 1977 | 0.096 | −0.435 | 0.011 |
| | 1980-1985 | 1982 | 0.089 | −0.539 | 0.016 | | 1980-1985 | 1982 | 0.088 | −0.490 | 0.032 |
| | 1985-1990 | 1987 | 0.078 | −0.455 | 0.047 | | 1985-1990 | 1987 | −0.097 | −0.024 | 0.012 |
| | 1990-1995 | 1992 | −0.164 | −0.107 | 0.007 | | 1990-1995 | 1992 | −0.135 | −0.205 | 0.012 |
| | 1995-2000 | 1997 | −0.255 | 0.098 | −0.018 | | 1995-2000 | 1997 | −0.110 | −0.221 | 0.024 |
| | 2000-2005 | 2002 | −0.177 | 0.072 | −0.023 | | 2000-2005 | 2002 | −0.137 | −0.037 | 0.015 |
| | 2005-2010 | 2007 | 0.091 | −0.013 | −0.002 | | 2005-2010 | 2007 | −0.165 | 0.221 | −0.027 |
| | 2010-2015 | 2012 | 0.016 | −0.182 | 0.012 | | 2010-2015 | 2012 | −0.021 | 0.043 | −0.021 |
| France | 1970-1975 | 1972 | 0.102 | −0.036 | −0.009 | Switzerland | 1970-1980 | 1975 | 0.230 | −0.870 | 0.013 |
| | 1975-1980 | 1977 | 0.147 | 0.097 | 0.003 | | 1980-1986 | 1983 | 0.020 | −0.896 | 0.026 |
| | 1980-1985 | 1982 | 0.457 | −1.399 | 0.008 | | 1986-1990 | 1988 | −0.033 | −0.686 | 0.079 |
| | 1985-1990 | 1987 | 0.329 | −3.135 | 0.090 | | 1990-1995 | 1992 | −0.023 | −0.836 | 0.051 |
| | 1990-2000 | 1995 | −0.309 | −1.130 | 0.081 | | 1995-2000 | 1997 | 0.127 | −1.399 | 0.153 |
| | 2000-2005 | 2002 | −0.338 | −1.292 | 0.291 | | 2000-2005 | 2002 | 0.110 | −1.765 | 0.218 |
| | 2005-2010 | 2007 | −0.930 | −0.225 | −0.068 | | 2005-2011 | 2008 | −0.606 | 0.051 | −0.041 |
| | 2010-2015 | 2012 | 0.000 | 0.000 | 0.000 | | 2011-2015 | 2013 | −0.351 | −0.183 | −0.031 |
| Germany | 1972-1975 | 1973 | −0.102 | −0.590 | 0.010 | UK | 1971-1975 | 1973 | 0.038 | −0.063 | 0.000 |
| | 1975-1980 | 1977 | 0.039 | −1.243 | 0.029 | | 1975-1980 | 1977 | 0.032 | −0.177 | 0.000 |
| | 1980-1985 | 1982 | −0.158 | −1.134 | 0.032 | | 1980-1985 | 1982 | 0.035 | −0.336 | 0.002 |
| | 1985-1990 | 1987 | 0.066 | −1.301 | 0.066 | | 1985-1990 | 1987 | 0.028 | −0.287 | 0.006 |
| | 1990-1995 | 1992 | 0.062 | −0.990 | 0.080 | | 1990-1995 | 1992 | −0.011 | −0.291 | 0.007 |
| | 1995-2000 | 1997 | −0.299 | −1.113 | 0.077 | | 1995-2000 | 1997 | 0.009 | −0.399 | 0.025 |
| | 2000-2005 | 2002 | −0.210 | −1.346 | 0.132 | | 2000-2011 | 2005 | −0.423 | −0.675 | 0.075 |
| | 2005-2010 | 2007 | −0.624 | −0.624 | 0.023 | | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 |
| | 2010-2015 | 2012 | −0.404 | −0.216 | 0.010 | | | | | | |
| Greece | 1970-1975 | 1972 | 0.000 | 0.000 | 0.000 | | | | | | |
| | 1975-1980 | 1977 | 0.000 | 0.000 | 0.000 | | | | | | |
| | 1981-1991 | 1986 | 0.126 | −1.002 | 0.039 | | | | | | |
| | 1991-2001 | 1996 | 0.144 | −1.454 | 0.165 | | | | | | |
| | 2001-2011 | 2006 | −0.061 | −1.134 | 0.164 | | | | | | |
| | 2011-2015 | 2013 | 0.000 | 0.000 | 0.000 | | | | | | |

*Source*: Authors' calculations using data described in Table 1 of the main text.

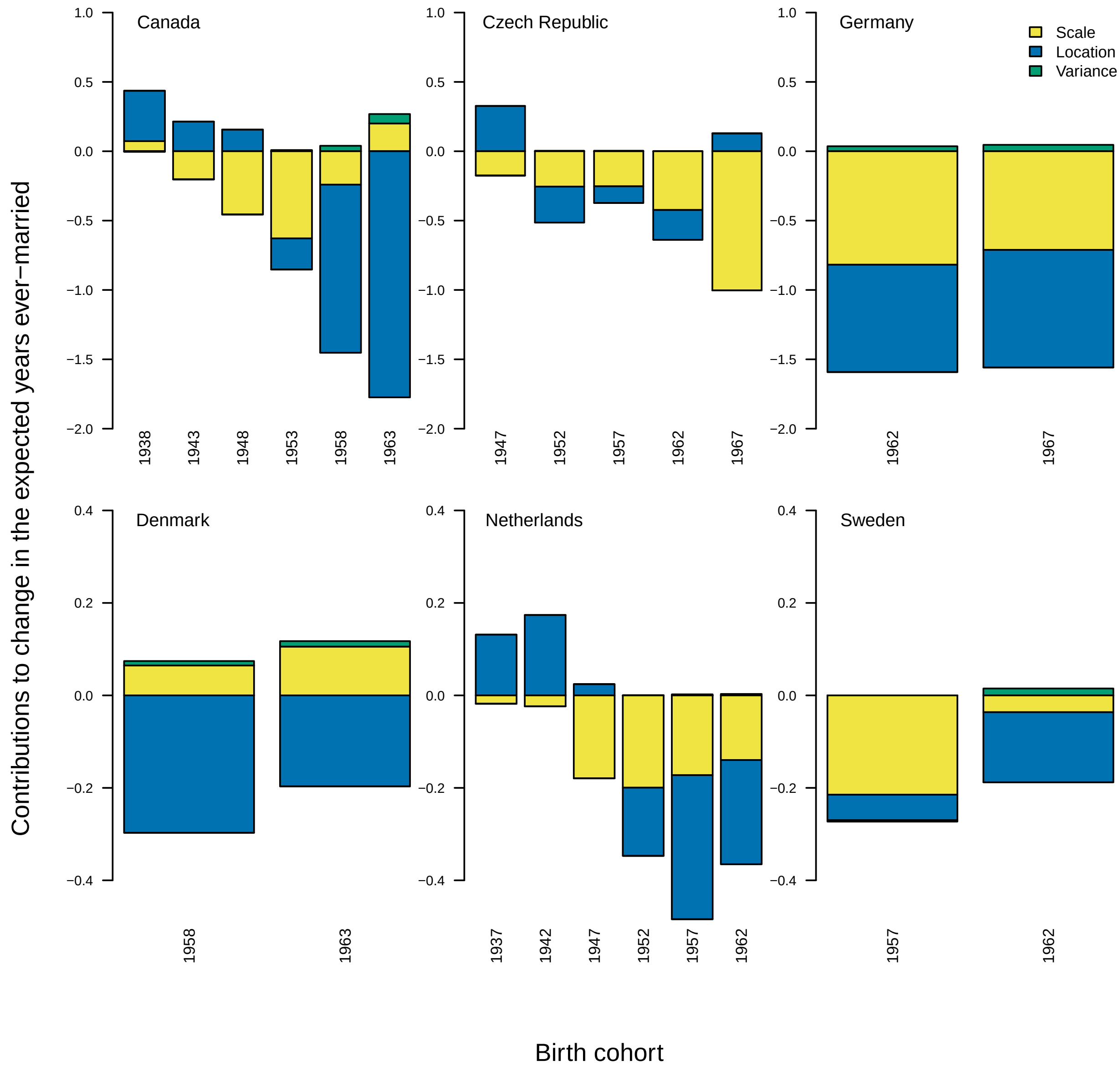



Figure 2.D.2: Decomposition of the change over time in the male cohort expected years ever-married in six countries

*Note*: The birth cohort presented corresponds to the mid-year between two birth cohorts. For example, for the changes in cohort EYEM from 1950 to 1955, it is written as 1952. The detail information can be seen in Table 2.D.1 in this supplemental material.

*Source*: Authors' calculations using data described in Table 1 of the main text.

Table 2.D.3: The contribution of each component to the time change in the cohort expected years ever-married

| | | | Female | | | Male | | |
|---|---|---|---|---|---|---|---|---|
| Country | Year | Mid-year | Scale | Location | Variance | Scale | Location | Variance |
| Canada | 1936-1941 | 1938 | 0.130 | 0.327 | 0.000 | 0.070 | 0.363 | −0.005 |
| | 1941-1946 | 1943 | −0.253 | −0.178 | 0.000 | −0.193 | 0.213 | −0.003 |
| | 1946-1951 | 1948 | −0.318 | 0.097 | 0.000 | −0.436 | 0.156 | −0.001 |
| | 1951-1956 | 1953 | −0.706 | −0.214 | 0.000 | −0.600 | −0.224 | 0.010 |
| | 1956-1961 | 1958 | −0.632 | −0.992 | −0.006 | −0.230 | −1.209 | 0.046 |
| | 1961-1966 | 1963 | 0.155 | −1.647 | −0.029 | 0.191 | −1.767 | 0.077 |
| Czech Republic | 1945-1950 | 1947 | 0.094 | 0.071 | 0.000 | −0.175 | 0.327 | −0.001 |
| | 1950-1955 | 1952 | −0.076 | 0.002 | 0.000 | −0.255 | −0.259 | 0.004 |
| | 1955-1960 | 1957 | 0.024 | 0.204 | 0.000 | −0.253 | −0.120 | 0.004 |
| | 1960-1965 | 1962 | −0.742 | −0.434 | 0.000 | −0.423 | −0.215 | 0.001 |
| | 1965-1970 | 1967 | −0.811 | 0.023 | 0.000 | −1.003 | 0.128 | 0.002 |
| Denmark | 1956-1961 | 1958 | −0.002 | −0.438 | 0.004 | 0.065 | −0.297 | 0.010 |
| | 1961-1966 | 1963 | 0.098 | −0.269 | 0.005 | 0.106 | −0.197 | 0.012 |
| Germany | 1960-1965 | 1962 | −0.559 | −1.160 | 0.022 | −0.776 | −0.771 | 0.041 |
| | 1965-1970 | 1967 | −0.684 | −0.861 | 0.021 | −0.673 | −0.843 | 0.051 |
| Netherlands | 1935-1940 | 1937 | 0.031 | 0.093 | 0.000 | −0.018 | 0.131 | 0.000 |
| | 1940-1945 | 1942 | −0.015 | 0.137 | 0.000 | −0.024 | 0.174 | 0.000 |
| | 1945-1950 | 1947 | −0.126 | 0.081 | 0.000 | −0.179 | 0.024 | 0.000 |
| | 1950-1955 | 1952 | −0.192 | −0.016 | 0.000 | −0.200 | −0.148 | 0.001 |
| | 1955-1960 | 1957 | −0.182 | −0.283 | 0.000 | −0.172 | −0.312 | 0.002 |
| | 1960-1965 | 1962 | −0.117 | −0.300 | 0.001 | −0.140 | −0.225 | 0.003 |
| Sweden | 1955-1960 | 1957 | −0.192 | −0.151 | 0.001 | −0.215 | −0.055 | −0.003 |
| | 1960-1965 | 1962 | −0.169 | −0.096 | 0.002 | −0.036 | −0.152 | 0.015 |

*Source*: Authors' calculations using data described in Table 1 of the main text.

# Chapter 3

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

## 3.1 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 of their reproductive period since 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 was also reported that the future level of childlessness is projected to increase with the divergence of countries (Sobotka 2017). Becoming a parent or remaining childless influences various aspects of a female's and male's life. It affects income (Budig et al. 2012), health (Kendig et al. 2007), old-age wellbeing (Huijts et al. 2013; Zhang and Hayward 2001), and

This chapter is a forthcoming article; Mogi, R. and del Mundo, M. (forthcoming). Decomposing changes in first birth trends: Quantum, timing, or variance. *Vienna Yearbook of Population Research.*

support networks (Albertini and Kohli 2009). Therefore, it is worthwhile to have a deep understanding of the trend of childlessness. Simultaneous with the increase in the number of childless women, the mean age at first birth has increased. For instance, the mean age at having the first child has increased by about one year each decade since the 1970s in the OECD countries (Mills et al. 2011). Therefore, in these decades, both phenomena of childlessness (*quantum* changes) and postponement of first birth (*timing* changes) have taken place.

The most common indices: the proportion of childless women at age 50 and the mean age at first birth are useful to study either quantum changes or timing changes. This is simply because the proportion of childless women does not consider the timing changes (e.g., the same proportion of childless women can be reached with different mean age at first birth) and the mean age at first birth per se does not show quantum changes. Previous research investigates how changes in the timing of first birth affect the level of childlessness. For instance, Kneale and Joshi (2008) explore the extent of childbirth postponement and project its impact on eventual levels of childlessness in the UK. Similarly, te Velde et al. (2012) also estimated the effect of postponement of entering motherhood on permanent involuntary childlessness in six European countries.

This study, on the other hand, suggests an alternative index that takes into account both quantum and timing changes altogether in one index: Expected Years Without Children (EYWC). This index measures the average length of life women spent without children at their reproductive period assessing both life years by women who remained childless at the end of reproductive period (quantum) and life years before having a child (timing). The concept of EYWC has been already introduced by Bongaarts and Feeney (2006) in fertility research and employed by Andersson et al. (2017) to describe the contemporary family changes using a life table method. However, this research did not focus on the index per se nor depict its trends. Therefore, the first aim of this study is to present the trend of EYWC to show how first birth trends have changed considering both quantum and timing changes.

Our second aim is to quantify which of two effects, remaining permanently childless or postponing 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, which is the variance in first birth age, is investigated. With an increase in the mean age at first birth, the variance of the people's first birth timing gets larger (Kohler and Philipov 2001). Partially because of this expansion of the first birth timing, the female's family life-course becomes more diverge across birth cohorts (Elzinga and Liefbroer 2007). Therefore, we quantify the effects of remaining permanently childless, postponement of first birth, and the expansion of first birth timing in the changes of 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 to describe the trend of first birth (EYWC) and the decomposition method, as well as the data used. Then, the third section presents the trend of EYWC and the results of its decomposition. In the final section, we mention the discussion, limitations, future developments and conclusion of this study.

## 3.2 Data and methods

### 3.2.1 Data

This study used data from the Human Fertility Database (HFD) (Human Fertility Database 2018). Specifically, this study used the female population counts by age and birth cohort and birth counts by birth order and mother's cohort. Only completed cohort fertility data were used to avoid problems arising from truncation and censoring bias. Based on these criteria, data from eight countries were selected: Canada (1929 - 1962 birth cohort), 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).

### 3.2.2 Coale-McNeil model for first birth

The parametric model for overall fertility has been developed well (see, for example, Kostaki and Paraskevi (2007)). In contrast to this trend, there are few models for a 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 to quantify the effects of remaining permanently childless and postponement on fertility behaviours.

The CM model was developed to estimate the age patterns of the first marriage of a birth cohort (Coale and Trussell 1978). Extended from the original use, the CM model has often been applied to the first birth distribution by age as well(Bloom 1982$a$,$b$; Bloom and Trussell 1984; Henz and Huinink 1999; Rao 1987; Trussell and Bloom 1983). Previous studies applied the CM model to various countries, e.g., Canada, Columbia, Finland, Germany, Italy, and the US, with much success. The advantage of the CM model is that its parameters have clear demographic meanings as follows. The probability of first birth at age $x$ and time $t$, denoted as $f_{x,t}$, is expressed as:

$$f_{x,t} = C_t \frac{1}{\sigma_t} a_1 \exp\Big[a_2(\frac{x-\mu_t}{\sigma_t} + a_3) - \exp\{-a_4(\frac{x-\mu_t}{\sigma_t} + a_3)\}\Big], \tag{3.1}$$

where $C_t$ pertains to the proportion of the cohort eventually having a child by age 50 at time $t$, $\mu_t$ refers to the mean age at first birth at time $t$, $\sigma_t$ 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 adequately applicable to our study. First, for one feature of the CM model (the convolution structure), an additional term was proposed, which is 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 one with the waiting term. Thus, previous research applied the CM model to first birth without using the additional waiting time segment (Bloom 1982*a*,*b*; Bloom and Trussell 1984; Henz and Huinink 1999; Rao 1987; Trussell and Bloom 1983). In addition, this is not theoretically appropriate to the current data. As a nonmarital childbirth rate 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 after marriage to childbirth is not important to our data studied. The second model is a log-logistic function (LL model). Henz and Huinink (1999) employed it for the first birth in the German data. However, it does not have a scale parameter, which indicates the proportion of the population that never had children. Hence, it is not applicable to our aim and decomposition method.

Despite the wide use of the CM model to first birth, it has not been statistically examined whether the CM model fits well with the observed age pattern of first birth. The goodness-of-fit using the Kolmogorov-Smirnov tests shows that the CM model estimates the data for all the countries and years well (see Appendix 3.A). For these reasons, we use the CM model to explore the changes in first birth behaviours and the expected years without children (EYWC) over time.

However, we have to mention the CM model's limitations. There are statistical issues when the CM model is used to estimate the mean and standard deviation of age at first birth (Bloom and Trussell 1984): (1) if the available sample for estimation is restricted to women who have become mothers on or before the survey date, there will be a truncation bias problem; (2) if the data used in estimation are for 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 using completed cohort fertility data from the HFD.

### 3.2.3 Expected years without children (EYWC)

We use life expectancy to measure fertility behaviours. Life expectancy between two ages (e.g., 0 and $X$), in classical life table methods, represents the area below a survival function from age 0 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, life expectancy is interpreted as the expected years without children, as "death" can be taken as a first birth. The minimum age should be 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 Fig 3.1 for the 1940 and 1962 birth cohorts. The use of EYWC has two primary advantages. First, this index is used to capture fertility trends in one simple value; thus, it is possible to numerically compare fertility trends at different times and in different countries. Second, it can take into account both the population having children and that without children in one index. While the most common indices to show first birth trends are the mean age at first birth and the proportion of childless women, EYWC can illustrate those two indices at the same time. This is very useful in the current world, in which both indices have an increasing trend.

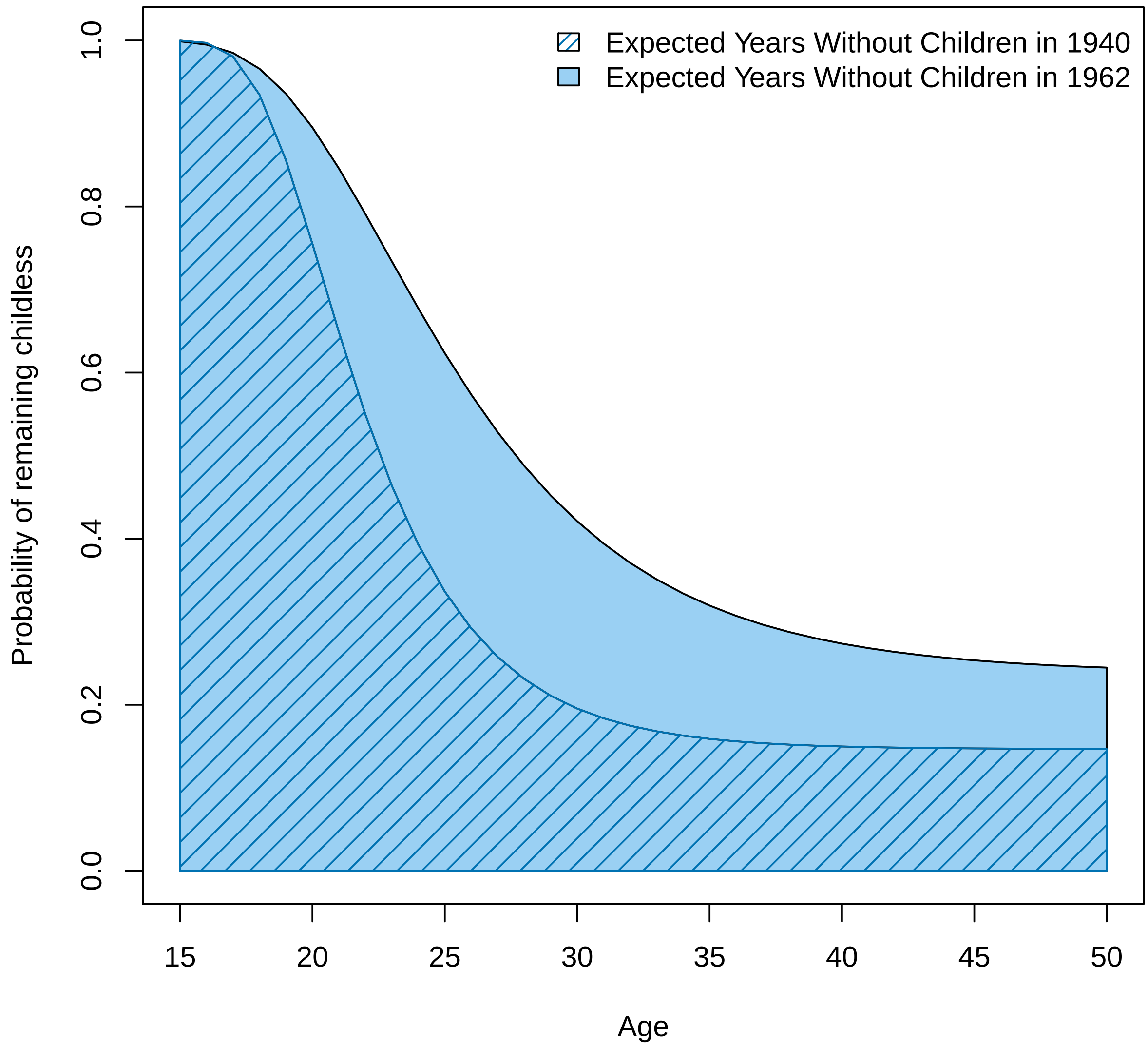



Figure 3.1: Probability of remaining childless by age among a Canadian female birth cohort in 1940 and 1962

*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.

### 3.2.4 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

$$ {}_{35}\dot{e}_{15}(t) = \frac{\partial\, {}_{35}e_{15}(t)}{\partial C_t}\dot{C}_t + \frac{\partial\, {}_{35}e_{15}(t)}{\partial \mu_t}\dot{\mu}_t + \frac{\partial\, {}_{35}e_{15}(t)}{\partial \sigma_t}\dot{\sigma}_t, \tag{3.2} $$

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 3.B explains the detailed method of estimating the three parameters of the CM model and applying the decomposition equation to discrete data.

## 3.3 Results

### 3.3.1 Cross-country trend analysis of EYWC

The trends of EYWC for the selected countries are presented in Fig 3.2.

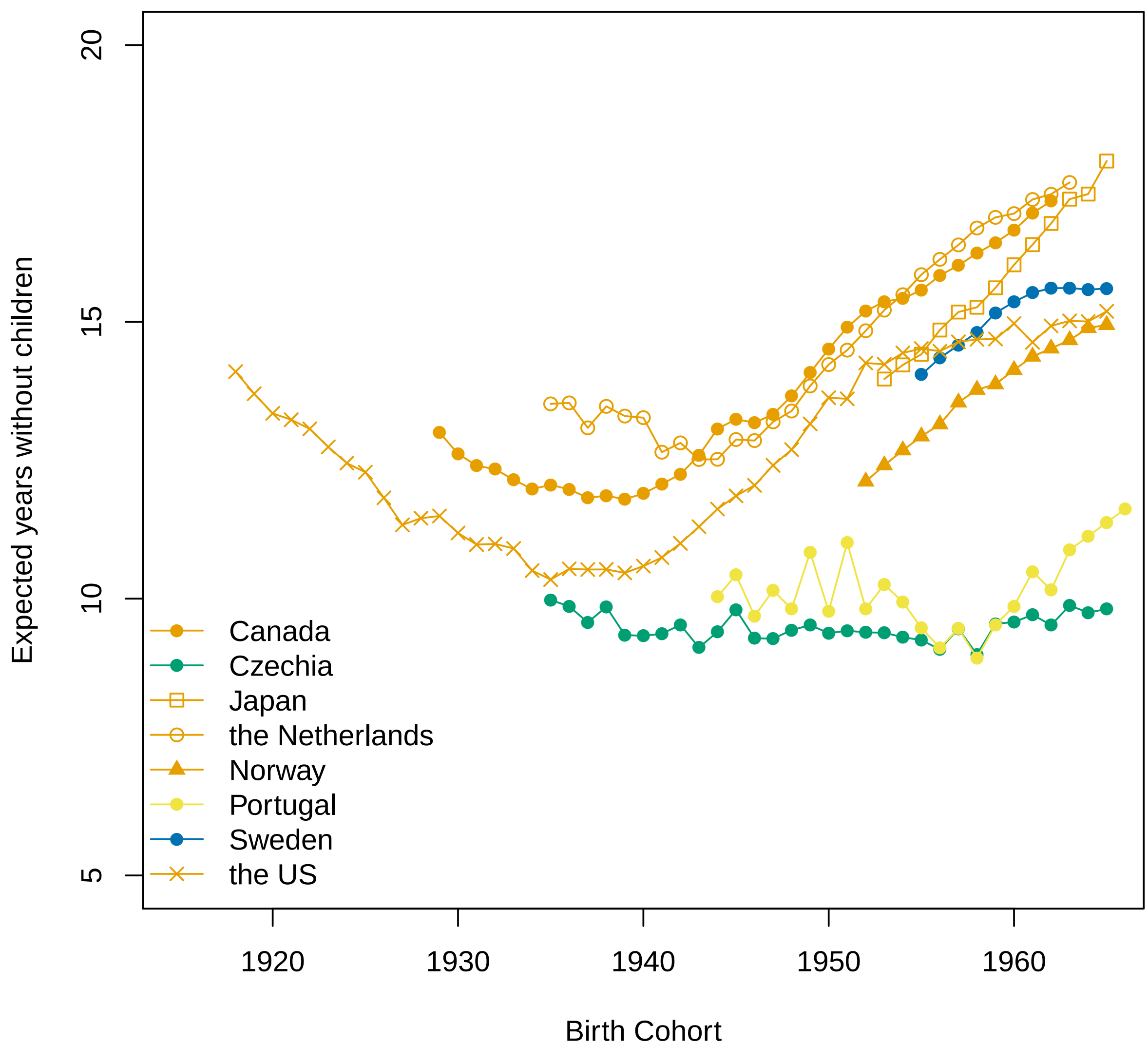



Figure 3.2: Time trends in expected years without children in eight countries
*Source*: Authors' calculations using the Human Fertility Database.

Figure 3.2 shows that most of the countries in the study had an increasing trend in EYWC starting from 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 increased to 16.6 years for the 1960 cohort. As EYWC is 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 it as 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 a similar increasing trend to Canada until the 1950s birth cohorts. The EYWC in the US strongly increased from the 1940 birth cohort (10.5 years) to the 1950 birth cohort (13.6 years). This increase, however, slowed after the 1950 birth cohort. The EYWC for the US only increased by 1.3 years from the 1950 to 1960 birth cohorts. The Netherlands also displayed an increase in EYWC from the 1945 birth cohort. The EYWC for the 1945 birth cohort was 13 years. This figure increased to 17 years for the 1960 birth cohort. The EYWC in other countries, such as Japan and Norway, also showed an increasing trend for all the observable periods. Japan had an EYWC of 14 years for the 1953 birth cohort, which increased to 18 years for the 1965 birth cohort. In the case of Norway, the 1952 cohort had 12 years of EYWC that increased to 15 years by the 1965 cohort.

In contrast to the trends of the countries discussed above, Portugal, Sweden, and the Czech Republic have different trends. The EYWC for Portugal fluctuated for approximately 10 years between the 1940s and 1950s birth cohorts. However, by the 1960 birth cohort, Portugal's EYWC increased and it is currently estimated to be at 11.6 years for the latest observed birth cohort. The Czech Republic, on the other hand, is different, as its EYWC does not have an increasing or decreasing trend. It has plateaued at 10 years for all observable birth cohorts from 1935 to 1965. Interestingly, the EYWC trend for Sweden also plateaued in recent cohorts. It initially increased from the 1955 birth cohort (14.4 years) but started to plateau by the 1960 birth cohort at 16 years.

In the latest birth cohorts that we can observe, Canada, Japan, and the Netherlands had approximately 17.5 years of EYWC, while Norway, Sweden, and the US had approximately 15 years. Following these countries, Portugal had 11.5 years, and the Czech Republic had fewer than 10 years. Small values for EYWC suggest that women gave birth to their first child at an early age. The latest cohort observed in Canada, Japan, and the Netherlands spent half of their reproduction periods without any child (the expected age at first birth is $15 + 17.5 = 32.5$). In other words, they only have 17.5 years left in their reproductive periods on average for any subsequent child.

The main objective of this study is to investigate whether the changes in EYWC, as seen in Fig 3.2, are due to remaining permanently childless, postponing childbirth, or the expansion effect. We decompose EYWC from 1940, and the results are presented in Fig 3.3.

### 3.3.2 Decomposition of EYWC

For 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 3.1 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 observed for the US; 68% of the increase in EYWC from 1958 to 1963 can be attributed to the increase in the average age at first birth. Likewise, the increase in the average age at first birth was responsible for 81% of the increase in the EYWC of Norway from 1957 to 1962. The Netherlands and Sweden also exhibited very similar results. This suggests that more people of the current birth cohorts in these countries have postponed childbirth rather than remaining permanently childless. The changes in the scale parameter or the proportion of women having children play a less influential role in the current increase in the EYWC of these countries.

In addition, the Netherlands (1955 - 1960) and Sweden (1960 - 1965) show a negative value for the scale parameter in the most recent birth cohort, which implies that the proportion of

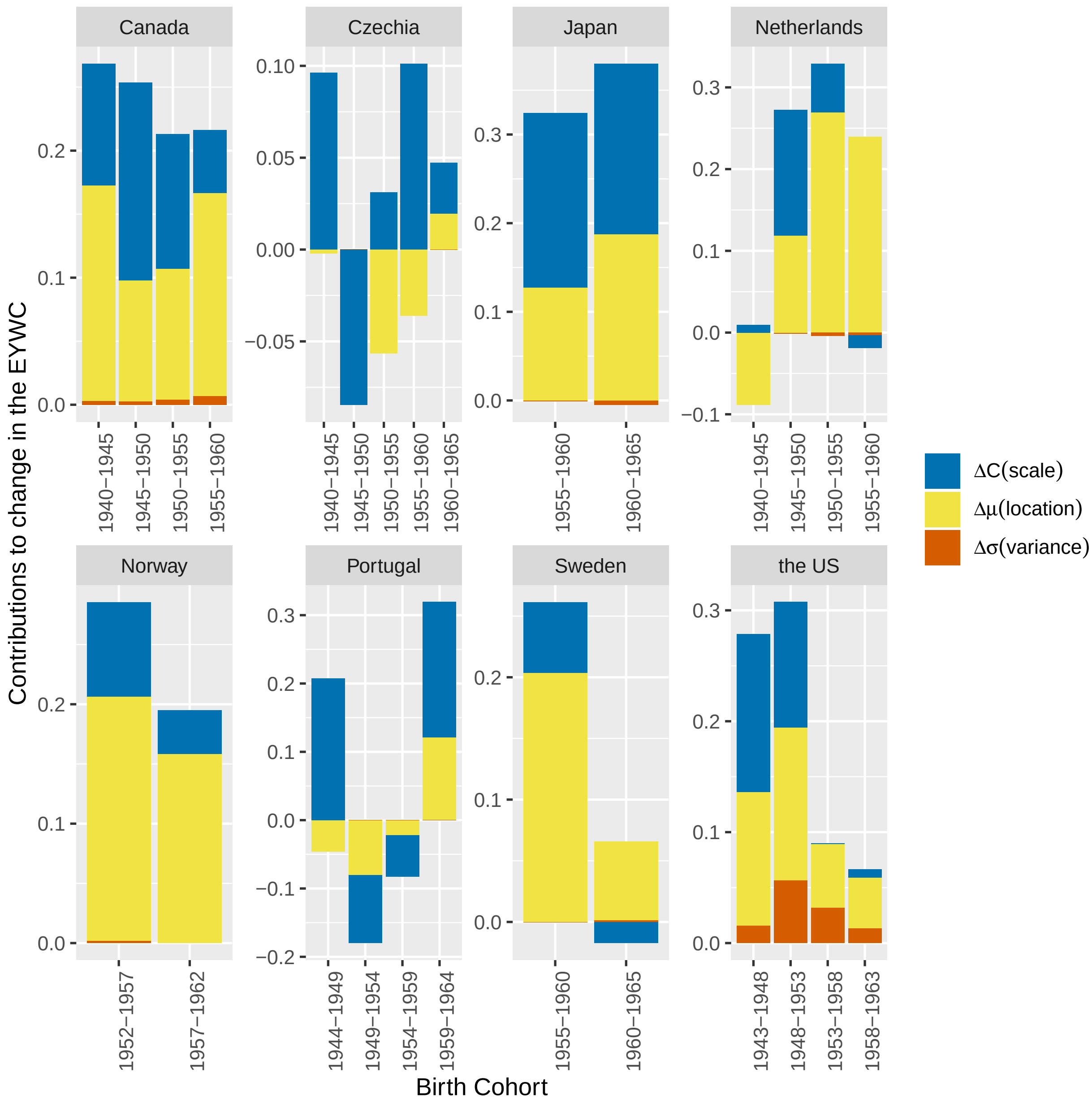



Figure 3.3: Decomposition of the change over time in females' expected years without children in eight countries

*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.

women having a first child increased in those periods. The negative value of the scale factor in Sweden is coherent to the current decrease in the proportion of childless women (Persson 2010; Miettinen et al. 2015). The higher educational category has a more distinct decreasing trend (Persson 2010). Therefore, it might be that this negative value is mainly driven by a higher educational group.

Japan and Portugal, on the other hand, show different trends from the other countries. The scale parameter, which indicates the proportion of women eventually having a child, is the most influential factor for the changes in their EYWC after 1955 and 1959, respectively. For instance, Table 1 shows that 62% of the change in the EYWC of Portugal from 1959 to 1964 can be attributed to the scale parameter, which means that remaining permanently childless was the main occurrence. Likewise, in Japan, the increase in EYWC was mainly due to the changes in scale parameter, even though the location parameter shared only approximately half as much impact on the latest change. This result indicates that more women in these countries remained permanently childless rather than postponed their first birth to a later age. Regarding Japan, the strong linkage between marriage and childbearing may be a key to understanding the large impact of the scale factor. As the nonmarital birth rate is still at a low level, i.e., 2.29% in 2015 (National Institute of Population and Social Security Research 2017), most births occur in marital unions. Hence, the large influence of the scale factor for Japan may present an increase in the never-married population in Japan. Indeed, the never-married population at age 50 increased between 1980 and 2015 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). Never-married people are almost equal to childless people in Japan, which is why the scale factor in Japan is so influential compared to the other countries' trends. The interpretation of the Czech Republic is difficult because the trend of its EYWC fluctuated.

For all countries, changes in the variance parameter were observed not to be influential in the changes of EYWC except the US. The relatively huge variance effect in the US may

Table 3.1: 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 (${}_{35}\dot{e}_{15}(t)$) in selected eight countries

| Country | Birth cohort | ${}_{35}\dot{e}_{15}(t)$ | Scale | Location | Variance | Sum of all components |
|---|---|---|---|---|---|---|
| Canada | 1940 - 1945 | 0.2685 | 0.0957 | 0.1698 | 0.0029 | 0.2684 |
| | | | (35.66) | (63.26) | (1.08) | |
| | 1945 - 1950 | 0.2535 | 0.1558 | 0.0951 | 0.0026 | 0.2535 |
| | | | (61.46) | (37.51) | (1.03) | |
| | 1950 - 1955 | 0.2129 | 0.1059 | 0.1027 | 0.0043 | 0.2129 |
| | | | (49.74) | (48.24) | (2.02) | |
| | 1955 - 1960 | 0.2164 | 0.0497 | 0.1597 | 0.0070 | 0.2164 |
| | | | (22.97) | (73.80) | (3.23) | |
| 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 | 0.0195 | 0.0000 | 0.0473 |
| | | | (58.77) | (41.23) | (0.00) | |
| 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 |
| the 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 | 0.2045 | 0.0021 | 0.2856 |
| | | | (27.66) | (71.60) | (0.74) | |
| | 1957 - 1962 | 0.1951 | 0.0368 | 0.1582 | 0.0002 | 0.1952 |
| | | | (18.85) | (81.05) | (0.10) | |
| 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 | 0.1205 | 0.0007 | 0.3192 |
| | | | (62.03) | (37.75) | (0.22) | |
| 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 |
| the USA | 1943 - 1948 | 0.2785 | 0.1422 | 0.1202 | 0.0160 | 0.2784 |
| | | | (51.08) | (43.18) | (5.75) | |
| | 1948 - 1953 | 0.3082 | 0.1131 | 0.1376 | 0.0568 | 0.3075 |
| | | | (36.78) | (44.75) | (18.47) | |
| | 1953 - 1958 | 0.0902 | 0.0007 | 0.0578 | 0.0318 | 0.0902 |
| | | | (0.78) | (64.08) | (35.25) | |
| | 1958 - 1963 | 0.0667 | 0.0077 | 0.0455 | 0.0136 | 0.0667 |
| | | | (11.54) | (68.22) | (20.39) | |

*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 (${}_{35}\dot{e}_{15}(t)$), due to rounding the numbers to the third decimal point in the table.

*Source*: Authors' calculations using the Human Fertility Database.

greatly come from the strong differentials in the first birth behaviours by socioeconomic statuses, such as educational level, union status, and race/ethnicity. Rendall et al. (2010) found persistence in early first births among low-educated women in the US and a shift towards later first births form women at higher education levels. Chandola et al. (2002) also suggested that the heterogeneity is related to marital status, with an early bulge linked to extra-marital births, often among solo mothers. They also indicated that ethnic differences play an important role in explaining the variance of first birth trend in the US. Likewise, Kostaki and Paraskevi (2007) also linked the observed heterogeneity in first birth patterns to a range of fertility determinants including the rise of migrant populations together with racial and ethnic differences in the US.

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

## 3.4 Conclusion

The increases in the proportion of childless women, the mean age at first birth, and the variance of the timing of first birth have been observed in high-income countries (Kohler and Philipov 2001; Kreyenfeld and Konietzka 2017; Mills et al. 2011). These increases have raised an important question: Which of three effects, remaining permanently childless, postponing first birth, or expansion of the standard deviation of mean age at first birth has the most impact on the changes in 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. Analysis of the trends shows that the EYWC of

Canada, the US, the Netherlands, Japan and Norway steadily increased through cohorts. EYWC in Sweden increased until the 1960 birth cohort and then remained at the same level. In Portugal, EYWC increased from the late 1950s cohort after fluctuating. EYWC in the Czech Republic bore no strong trends throughout cohorts that can 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. Findings of the decomposition analysis show that more people postpone first birth to a late age rather than remaining permanently childless in Canada, the US, the Netherlands, Sweden and Norway. The EYWC for these countries has an increasing trend mainly because their female populations have delayed childbirth. This could be a result of improvements in female education, labour force participation and better access to effective contraceptive methods. On the other hand, more women remain permanently childless rather than postpone childbirth in Japan and Portugal. Regarding Japan, the strong tie between marriage and childbirth may be an important key to understand the strong impact of scale factor (National Institute of Population and Social Security Research 2017). The variance factor did not have an important impact on the changes in EYWC for all periods and countries analyzed except the US. This might be due to the relatively large heterogeneity of first birth trend by socioeconomic statuses in the US compared to other countries (Chandola et al. 2002; Kostaki and Paraskevi 2007; Rendall et al. 2010).

Finally, for future research, it is worthwhile to apply this decomposition method to subpopulations, such as educational groups. It is well known that the timing of first birth and the share of childless women significantly differ by education (Beaujouan et al. 2016; Kneale and Joshi 2008). An application to union status (e.g., single, cohabitation, married) and race/ethnicity will also be beneficial.

# Appendices

## Appendix 3.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 3.A.1 shows, the CM model estimated the observed data statistically well for all countries and birth cohorts.

Table 3.A.1: Goodness-of-fit of the CM model using 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 |
| | 1955 | 0.1714 | 0.6902 | Portugal | 1944 | 0.2000 | 0.4916 |
| | 1960 | 0.1714 | 0.6902 | | 1949 | 0.1429 | 0.8674 |
| Czech Republic | 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 |
| | 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 |
| Japan | 1955 | 0.1143 | 0.9794 | the 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 |
| the Netherlands | 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 3.B Calculation process: Expected years without children

The life expectancy from age 0 to age $X$ is shown as

$$_{X}e_0(t) = \int_0^X l_{x,t}\, dx.$$

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 3.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).

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

where $\text{With}_x$ is the female population with children at age $x$, $\text{Without}_x$ 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 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).

# Chapter 4

# Cross-sectional average length of life by parity: Illustration of US cohorts of reproductive age in 2015

## 4.1 Introduction

In developed countries, the cohort total fertility rate (CTFR) has declined below the replacement fertility level beginning with the 1940s birth cohorts (Frejka and Calot 2001; Myrskyla et al. 2013; Sobotka et al. 2015). Although Nordic countries, France and the US have higher CTFRs, the CTFRs of Southern Europe and East Asia are far below the replacement level, with 1.5 children per woman on average in Italy and Japan based on the 1970 birth cohort; these cohorts have practically concluded their reproductive ages (Frejka and Calot 2001; Human Fertility Database 2019; Sobotka et al. 2015). This decline in the total fertility rate is often accompanied by changes in the parity progression ratios. The parity progression

This chapter is a forthcoming book chapter article; Mogi, R. and Canudas-Romo, V. (forthcoming). Cross-sectional average length of life by parity: Illustration of US cohorts of reproductive age in 2015. In Schoen, R (Ed.), *Analyzing Contemporary Fertility*. Springer.

ratios from parity two to higher parities dropped significantly from the 1930 to 1965 birth cohorts, and the progression ratios from parity zero to one and from parity one to two started decreasing in more recent cohorts in most European countries (Frejka 2008). In addition, the age of entering motherhood has increased since the 1970s by approximately one year each decade on average in high-income countries (Mills et al. 2011), although age disparities can be observed. For example, women in Eastern European countries tend to enter motherhood at a relatively young age, while women in Southern Europe and Eastern Asia show late entrance into motherhood, above the age of 30 on average (Kneale and Joshi 2008; Schmidt et al. 2012; Toulemon 1996). The postponement of motherhood increases the risk of remaining childless and leaves less time for further births.

Traditional fertility indexes are effective in describing current childbirth patterns by capturing either the quantum or timing of childbirth. For example, the cohort total fertility rate exclusively shows quantum changes, while the mean age at birth indicates changes in the timing of childbearing. In addition, parity progression ratios are useful to show transitions between parities; however, it is difficult to observe the quantum of parity structure based on these ratios. Although these indexes have undoubtedly strengthened our understanding of quantum and timing changes in fertility, we suggest that a comprehensive index to measure a women's fertility life history during her reproductive ages can offer an alternative perspective to study changes in fertility.

Moreover, the period and cohort indexes have well-known limitations. The period perspective (using a synthetic cohort approach) does not necessarily reflect the experience of any real birth cohort (Bongaarts and Sobotka 2012; Luy 2011), while the cohort indexes, based on information on populations that have passed their reproductive ages, provide an outdated picture of fertility. This study introduces an alternative measure, the Cross-sectional Average Length of Life by Parity ($CALP$). This index utilizes all the age-specific information in the birth histories of cohorts currently at childbearing ages. The concept of $CALP$ builds on an existing index developed in mortality research as an alternative measure of life expectancy, namely, the Cross-sectional Average Length of Life, CAL (Brouard 1986; Guillot 2003).

This study discusses how the *CALP* can provide an alternative perspective in fertility research. The *CALP* is valuable because it complements the existing period and cohort measures and shows the duration or average number of years spent in each parity during reproductive ages. For example, the $\underline{CALP}$(2015) summarizes all the fertility histories of cohorts from 1966 to 2003 in parities 0 to 5+ into one index and comprehensively describes that year's fertility situation. As an illustration, the *CALP* is estimated for the US fertility series. As opposed to the fertility trends of other high-income countries, the US maintains a high level of CTFR, even for cohorts that have recently completed their reproductive ages (Frejka and Calot 2001; Human Fertility Database 2019; Myrskyla et al. 2013). This high fertility level is a result of higher progression ratios after the first birth (Frejka and Sardon 2007; Zeman et al. 2018).In addition, the birth schedule by parity in the US has remained practically unchanged from the 1960s-1980s birth cohorts (Frejka and Sardon 2007). The stable fertility levels and parity-specific birth schedules of cohorts make the US an optimal case country to demonstrate the usefulness of the *CALP* as a measure of the fertility experience of the many cohorts present at a given time. The average years spent in each parity during reproductive years among all women aged 12 to 50 in 2015 were estimated using a hierarchical multistate life table model.

## 4.2 Data and Methods

### 4.2.1 Data

We used the Human Fertility Database (HFD) to obtain age- and parity-specific fertility rates by cohorts of US women. The HFD is an open-access database that includes strong quality control measures, and only countries that have comprehensive, high-quality information are included in the database. For the measures proposed here, the US long-span birth cohort data in the HFD were selected. The data cover individuals aged 12 to 50 and the birth cohorts from 1966 to 2003, thus enabling the calculation of the *CALP* in 2015.

### 4.2.2 Methods

**The Cross-sectional Average Length of Life by Parity (*CALP*)**

The $\underline{CALP}(t)$ is an alternative period measure to interpret fertility behaviors. The $\underline{CALP}(t)$ is a column vector including the duration spent in each parity $i$ at a given time $t$ for reproductive-aged women between 12 and 50 years. Here, underlining a variable indicates a matrix. In contrast to traditional period indexes that use a synthetic cohort approach, the $\underline{CALP}(t)$ uses real cohort data, including all the cohort age-specific fertility rates (occurrence-exposure rates) by parity for all female cohorts at reproductive ages at time $t$. The $\underline{CALP}(t)$ is defined as

$$\underline{CALP}(t)' = (CALP_0(t), CALP_1(t), CALP_2(t), CALP_3(t), CALP_4(t), CALP_{5+}(t))$$

and

$$CALP_i(t) = \int_{12}^{50} l_i^c(x, t-x)dx, \tag{4.1}$$

where $CALP_i(t)$ is the average number of years spent in parity $i$ by women of reproductive age in year $t$ and $l_i^c(x, t-x)$ is the survival function for persons reaching age $x$ and parity $i$ (0, 1, 2, 3, 4, 5+) at time $t$ who were born in year $t-x$. As the studied reproductive age range is 12 to 50 (i.e. 38 years) then $\sum_{i=0}^{5+} CALP_i(t) = 38$. Figure 4.1 depicts the elements included in equation (4.1) using a Lexis diagram. Each diagonal dashed line corresponds to the age-specific cohort survival function for persons who were born in year $t-x$ and who reached age $x$ and parity $i$ in 2015. For example, $CALP_i(2015)$ sums the 1966 birth cohort survivors at parity $i$ and age 50, the 1967 birth cohort survivors at parity $i$ up to age 49, and so on until the 2003 birth cohort survivors at parity $i$ who reached age 12 in 2015.

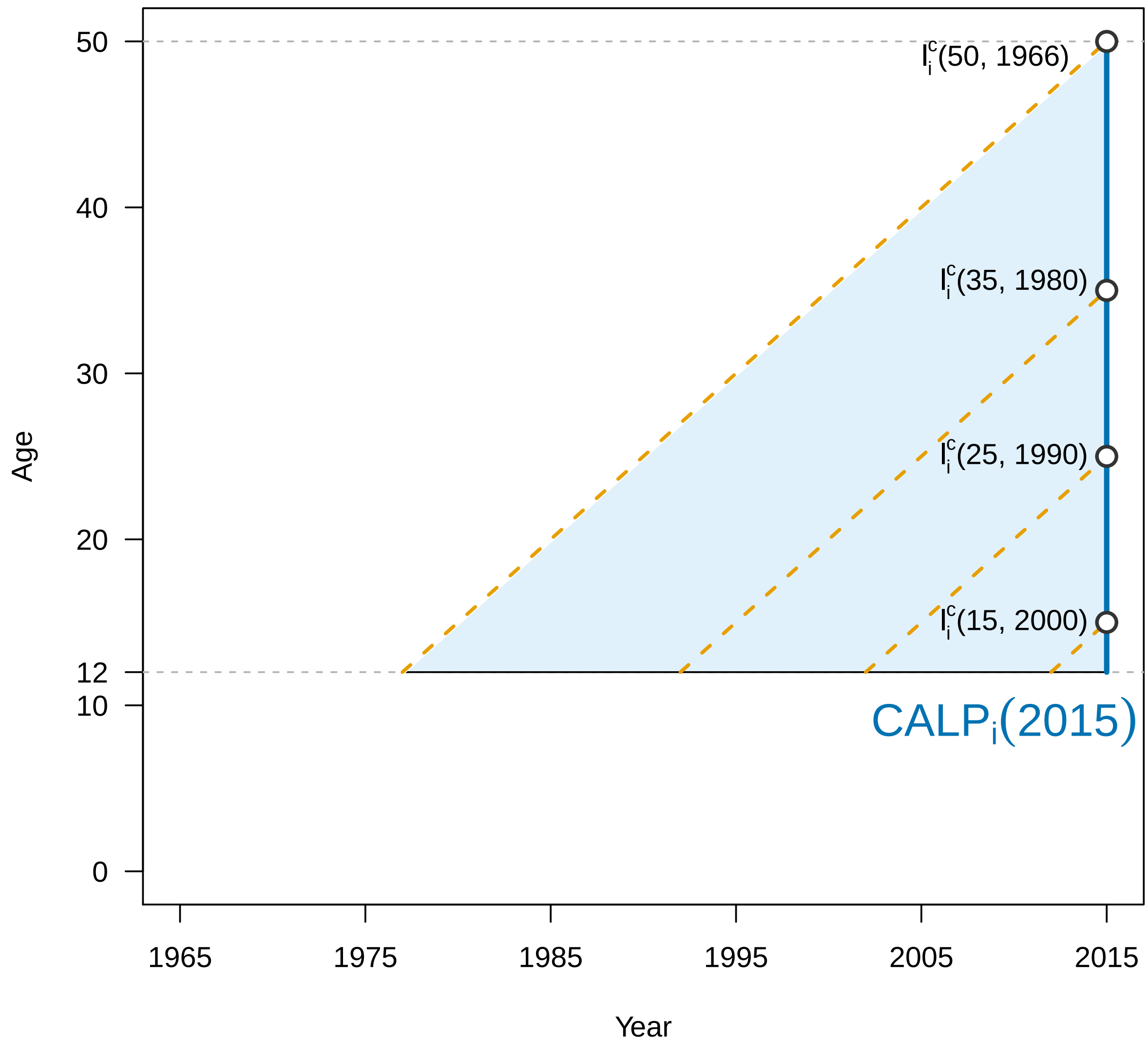



Figure 4.1: Illustration of the structure of the cross-sectional length of life for parity $i$ in 2015, $CALP_i(2015)$, in the Lexis diagram

*Note*: $l_i^c(x, t-x)$ represents the survival function for persons reaching age $x$ and parity $i$ (0, 1, 2, 3, 4, 5+) at time $t$ who were born in year $t-x$.

The cross-sectional average length of life for individuals with a parity of 0 ($CALP_0$) was previously determined by Mogi et al. (2019) as the Cross-sectional Average Length of Life Childless (CALC), which was inspired by an analogous measure developed in mortality research (CAL by (Brouard 1986; Guillot 2003). The index $CALP$ defined in equation (4.1) considers transfers only between parity states under the assumption of no other source of attrition. In high-income countries, mortality among reproductive-aged women is very low, and sensitivity analysis conducted on $CALP_0$ (or CALC) shows that mortality had a minor influence (Mogi et al. 2019). Therefore, mortality is not expected to make a significant difference in the results presented here. This assumption was also used in Schoen (2016).

### Hierarchical multistate life tables

Hierarchical multistate life tables were used to estimate the cohort survival function by parity ($l_i^c(x, t-x)$, used in equation (4.1)). Hierarchical multistate models are a particular case of the general multistate model where the transitions between states can only happen in one way, e.g. it is possible to transition from parity 0 to parity 1 but not from parity 1 to parity 0 (Schoen 2016; Schoen and Canudas-Romo 2006). For example, the transitions between states (here, parity status) are shown in Figure 4.2, with transition rates $m_{ij}$ relating the movement from parity $i$ to parity $j$ ($j = i+1$); for example $m_{23}$ corresponds to the transition from parity 2 to parity 3.

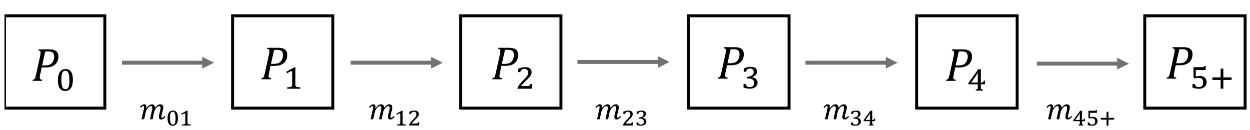



Figure 4.2: Diagram of the parity hierarchical multistate model
*Note*: $m_{ij}$ is the transition rate from parity $i$ to parity $j$ ($j = i+1$).

For this hierarchical model, the corresponding matrix of transitions at age $a$ for women born in year $t-x$ is defined as

$$\underline{m}(a,t-x) = \begin{pmatrix} m_{01} & -m_{01} & 0 & 0 & 0 & 0 \\ 0 & m_{12} & -m_{12} & 0 & 0 & 0 \\ 0 & 0 & m_{23} & -m_{23} & 0 & 0 \\ 0 & 0 & 0 & m_{34} & -m_{34} & 0 \\ 0 & 0 & 0 & 0 & m_{45+} & -m_{45+} \\ 0 & 0 & 0 & 0 & 0 & 0 \end{pmatrix},$$

where the elements of the matrix are the age- and cohort-specific transition rates, $m_{ij}(a, t-x)$, and age $a$ can range from age 12 to age $x$, achieved by the cohort in year $t$. For example, the transition matrix for the 1966 cohort of women in the US at age 30, who turn exactly 50 years old in 2015 ($a = 30, x = 49, t = 2015$) is

$$\underline{m}(30,1966) = \begin{pmatrix} 0.1053 & -0.1053 & 0 & 0 & 0 & 0 \\ 0 & 0.1479 & -0.1479 & 0 & 0 & 0 \\ 0 & 0 & 0.0733 & -0.0733 & 0 & 0 \\ 0 & 0 & 0 & 0.0580 & -0.0580 & 0 \\ 0 & 0 & 0 & 0 & 0.1271 & -0.1271 \\ 0 & 0 & 0 & 0 & 0 & 0 \end{pmatrix}.$$

The transition matrix is used to calculate the number of persons in the cohort life table with each parity $i$ at exact age $x$ (Schoen 2016). Thus, we have

$$\underline{l^c}(x,t-x)' = \underline{l^c}(x-1,t-x)' \left[I - \frac{1}{2}\underline{m}(x-1,t-x)\right]\left[I + \frac{1}{2}\underline{m}(x-1,t-x)\right]^{-1},$$

where $\underline{l_i^c}(x, t-x)$ is the survivorship vector at age $x$ for the life table following the transition rates of the cohort born in year $t-x$; its elements are the parity-specific number of persons, $\underline{l}^c(x,t-x)' = (l_0^c, l_1^c, l_2^c, l_3^c, l_4^c, l_{5+}^c)$; and $I$ is the $6\times 6$ identity matrix. For example, the US 1966 cohort of women aged 50 years in 2015 has the following survivorship vector:

$$\underline{l}^c(50,1966)' = (l_0^c = 0.14, l_1^c = 0.18, l_2^c = 0.34, l_3^c = 0.21, l_4^c = 0.05, l_{5+}^c = 0.08),$$

corresponding to the distribution of the cohort across parities by age 50: 14% remained childless, 18% were in parity 1, 34% were in parity 2, 21% were in parity 3, and 5% and 8% were in parities 4 and 5 or more, respectively. For all cohorts present in year t, we assume that all women start at parity 0 at age 12; thus, the radix of $\underline{l}^c(a,t-x)$ is $\underline{l}^c(12,t-x)' = (1,0,0,0,0,0)$.

The presentation of the hierarchical model above focuses on the $CALP$ measures. However, a similar methodology can be used to calculate the expected number of years spent in each parity in a synthetic cohort (using the period information of a given year) from ages 12 to 50, e.g. for 2015 denoted as ${}_{38}\underline{e}_{12}^p(2015)$ or using the actual information of a cohort, e.g. for the 1966 cohort, denoted ${}_{38}\underline{e}_{12}^c(1966)$. The expected years in each parity from the period perspective are defined as

$${}_{38}\underline{e}_{12}^p(t)' = ({}_{38}e_{12,0}^p(t), {}_{38}e_{12,1}^p(t), {}_{38}e_{12,2}^p(t), {}_{38}e_{12,3}^p(t), {}_{38}e_{12,4}^p(t), {}_{38}e_{12,5+}^p(t))$$

and

$${}_{38}e_{12,i}^p(t) = \int_{12}^{50} l_i^p(x,t)dx, \qquad (4.2)$$

where ${}_{38}e_{12,i}^p(t)$ is the average number of years spent in parity $i$ between the ages of 12 and 50 in the synthetic cohort of women of reproductive age following the parity transitions observed in year $t$ and $l_i^p(x,t)$ is the survival function for persons reaching age $x$ and parity $i$ (0, 1, 2, 3, 4, 5+) at time $t$. Similar equations are used for the cohort life expectancies by parity or elements of ${}_{38}\underline{e}_{12}^c(t)$. Because the studied reproductive age range is 12 to 50 (i.e. 38

years), both sets of durations by parity for period and cohort life expectancies ensure that $\sum_{i=0}^{5+} {}_{38}e^{p}_{12,i}(t) = \sum_{i=0}^{5+} {}_{38}e^{c}_{12,i}(t) = 38$, similar to the *CALP*. In the illustrations presented here, all three perspectives—period, cohort and *CALP*—are compared.

## 4.3 Results

### 4.3.1 Parity transition rates

The US female cohort age-specific transition rates for the first three parities are shown in Figure 4.3. Each diagonal line represents the age-specific transition rate from parity $i$ to $i+1$ of a cohort of reproductive-aged women present in 2015. The light colors in the diagrams correspond to lower transition rates and the darker color to higher rates. The same scale is used across all three transitions. Thus, it is possible to compare them: the likelihood of having a second child is higher than that of the transition from childlessness to parity 1. The transition rates from parity 1 to 2 are high, especially among individuals in their early 20s, and remain high until the late 30s. This means that there is a high probability of having a second child in American cohorts of reproductive age in 2015. However, the transition rates from parity 2 to parity 3 concentrate around the 20s; these are women who entered motherhood at an early age and have a high propensity toward a third birth at an early age.

### 4.3.2 Survival function and duration by parity

Figure 4.4 shows the survival function of the number of persons in each parity $i$ at exact age $x$ estimated with a hierarchical multistate model based on the three perspectives: period ${}_{38}\underline{e}^{p}_{12}(2015)$, cohort ${}_{38}\underline{e}^{c}_{12}(1966)$ and $\underline{CALP}(2015)$. Parities are indicated in Figure 4.4 by labels $P_i$ corresponding to the parity order next to the lines, and line types indicate the type of index. The dashed line shows the period values, the dotted line represents the cohort perspective, and the solid line corresponds to the *CALP* survival function. At each age and for each perspective, all women in the cohort are distributed in the different parities; thus,

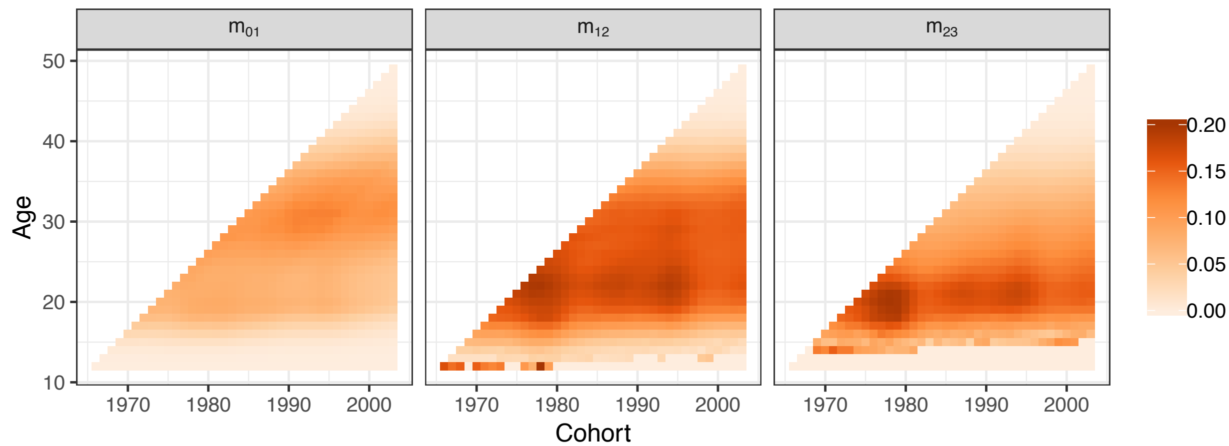



Figure 4.3: US 1966 to 2003 cohort age-specific transition rates for the first three parities
*Source*: Authors' calculations using the HFD.

the number of survivors in the different parities totals 100%.

Several interactions between the three perspectives can be observed in Figure 4.4, although for a major portion of the age range, the survival functions of $\underline{CALP}(2015)$ are located approximately between the other two indexes. For parity 0, $CALP_0$ at young ages is near the period measure, but at older ages it resembles the cohort number of women remaining childless. However, for some cohorts in their late 30s and 40s, the number of women transitioning to parity 1 is even greater than that for the cohort of 1966. Consequently, a crossover occurs, and there are fewer women remaining childless (i.e. in $CALP_0$ at these ages. The conventional period and cohort survival functions of ${}_{38}\underline{e}^{p}_{12}(2015)$ and ${}_{38}\underline{e}^{c}_{12}(1966)$ include only one cohort and show a monotonic decreasing trend in this parity. However, the probability of remaining in parity 0 for $\underline{CALP}(2015)$ increases after age 40, which occurs because $CALP$ is an aggregated index of several cohorts.

Similar transitions can be observed for other parities, where the trends of $CALP$ survivors move from resembling the period trends at young ages to match the 1966 cohort at old ages. The extreme case of crossover occurs at $P_1$. The cohort measure shows newcomers to this parity at an early age peaking around age 29 with a total of 25% of women in this parity, which subsequently decreased, reaching 18% in $P_1$ by age 50. In contrast, the period trends of $P_1$ show late arrival to this parity, reaching a peak at age 33 with a total of 26% and decreasing to 24% by the end of reproductive age.

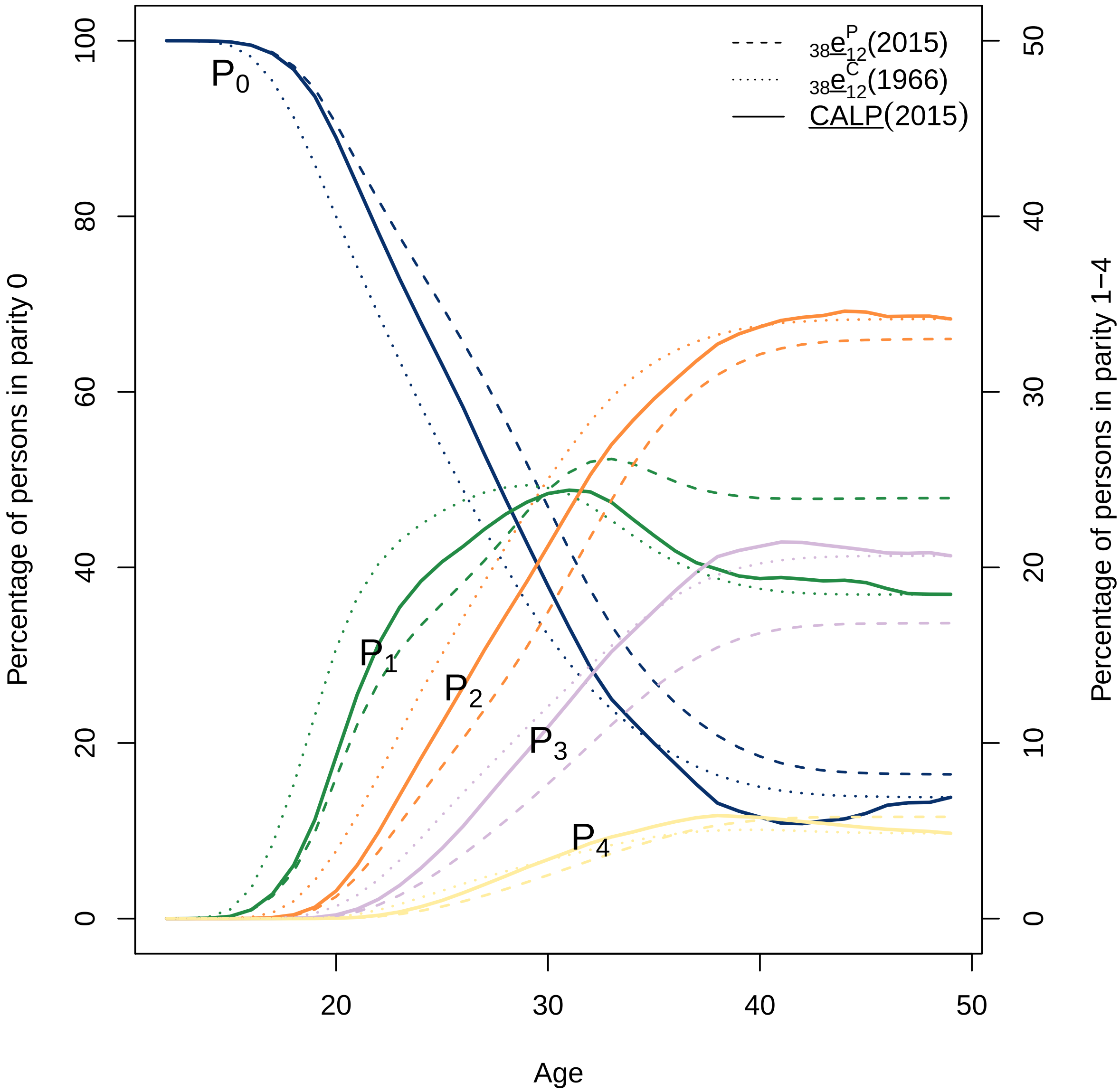



Figure 4.4: Survival functions for parities from 0 to 4 by age in the US for three different perspectives: period, cohort and CALP
*Source*: Authors' calculations using the HFD.

Table 4.1 shows the results of three different indexes measuring the expected duration in each parity for period data in 2015, the cohort born in 1966 and the *CALP* in 2015. In many cases, the values of $\underline{CALP}$(2015) are between the other two perspectives but are

Table 4.1: Measures of duration in each parity in the US: period, cohort and CALP

| Parity | ${}_{38}\underline{e}^{p}_{12}(2015)$ | ${}_{38}\underline{e}^{c}_{12}(1966)$ | $\underline{CALP}(2015)$ |
|---|---|---|---|
| $P_0$ | 19.78(52.1) | 17.00(44.7) | 17.91(47.1) |
| $P_1$ | 6.67(17.6) | 6.46(17.0) | 6.08(16.0) |
| $P_2$ | 6.66(17.5) | 7.86(20.7) | 7.28(19.2) |
| $P_3$ | 3.18(8.4) | 4.27(11.2) | 4.17(11.0) |
| $P_4$ | 1.07(2.8) | 1.08(2.8) | 1.13(3.0) |
| $P_{5+}$ | 0.64(1.7) | 1.35(3.6) | 1.43(3.8) |

*Source*: Authors' calculations using the HFD.
*Note*: All three indexes add to 38 across parities.
The values in parentheses are the percentages of years spent in parity $i$ between the ages of 12 and 50.

relatively closer to the cohort estimations than the period estimations. The large disparities from the period perspective show that the synthetic cohort does not represent the different parity-specific fertility histories of the real birth cohorts. The gap in $CALP$ from the cohort perspective simply shows that younger cohorts have trends different from the selected cohort. For instance, the three values of parity 1 have relatively large gaps, and the value of $CALP_1(2015)$ is smaller than that of the other two indexes. Figure 4.4 helps to explain this phenomenon. The period expected duration ${}_{38}e^{p}_{12,1}(2015)$ (the dashed line) starts close to $CALP_1(2015)$ as young cohorts have similar trends as those observed in 2015; however, from age 30, it stagnates at a higher level than the other two lines. In contrast, the cohort expected duration ${}_{38}e^{c}_{12,1}(1966)$ (the dotted line) increases earlier than the other two lines, and $CALP_1(2015)$ is closer to that line after age 30, with low levels of parity 1. This means that younger cohorts transitioned to parity 1 later than the 1966 birth cohort. The period measure ${}_{38}e^{p}_{12,1}(2015)$ does not take into account transitions occurring in the past, which can be substantial for older cohorts. This helps to explain the deviation of the synthetic cohort's fertility pattern from the real birth cohort's pattern, especially at older ages. $CALP$ can be seen as a complementary index integrating all the real cohort data on reproductive ages in 2015. In summary, $\underline{CALP}(2015)$ shows that women in the US spend 47% (17.91/38 years) of their reproductive years in childlessness, followed by 16%, 19% and 11% in parities 1, 2, and 3, respectively.

### 4.3.3 Understanding the age patterns in the three indexes

A simulation helps to explain the relationship among the three indexes ${}_{38}\underline{e}^{p}_{12}(2015)$, ${}_{38}\underline{e}^{c}_{12}(1966)$, and $\underline{CALP}(2015)$. To explain the crossovers seen in Figure 4.4, we assume that (1) cohorts from 1966 to 1985 have the same age-specific fertility rates by parity (ASFR) as those of the cohort from 1966, and (2) the ASFRs of cohorts from 1986 to 2003 are 10% lower than the ASFR of the 1966 cohort. In other words, these settings assume that the younger cohorts (1986-2003) uniformly have 10% lower fertility trends than the 1966 cohort, while the older cohorts have exactly the same trend. The birth schedule by parity in the US has remained practically unchanged for the 1960s-1980s birth cohorts (Frejka and Sardon 2007), justifying our assumption of fixed cohort trends described in (1) above. With these assumptions, the survival functions of the percentage of persons in parities 0 and 1 by age were calculated and are shown in Figure 4.5. The survival functions of ${}_{38}\underline{e}^{c}_{12}(1966)$ are the same as those in Figure 4.4. The simulated survival functions of $CALP_0(2015)$ (and $CALP_1(2015)$) in Figure 4.5 are higher (lower) than those of cohort ${}_{38}e^{c}_{12,0}(1966)$ (and ${}_{38}e^{c}_{12,1}(1966)$) until the age of 29 and then become equal to them. In contrast, the simulated survival functions of $CALP_0(2015)$ and $CALP_1(2015)$ begin to deviate from period ${}_{38}e^{p}_{12,0}(2015)$ and ${}_{38}e^{p}_{12,1}(2015)$ after age 29. The latter age corresponds to the 1986 cohort or the cohort experiencing the 10% reduction in ASFR in assumption (2) above. While the period index only shows an acceleration in ASFR after age 29, the $CALP$ measure captures the entire cohort shift from assumptions (1) and (2). This simulation shows how $CALP$ adapts from a period perspective to a cohort perspective because it is a period measure including the fertility information of women of reproductive age in all the cohorts.

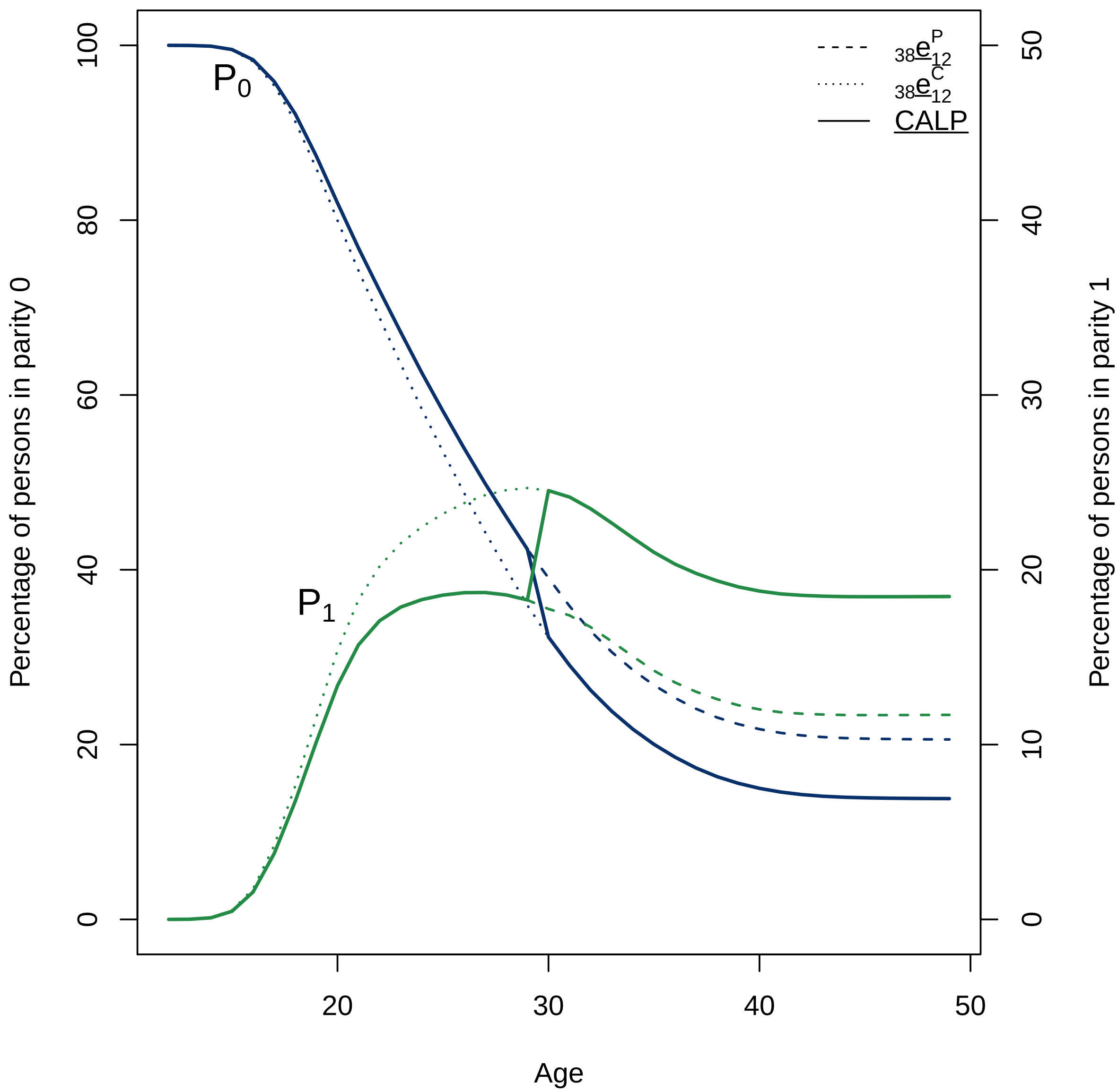



Figure 4.5: Simulated survival function for parities 0 and 1 by age for three different perspectives: period, cohort and CALP

### 4.3.4 Time trends of three indexes

Figure 4.6 presents the time trends from 1970 to 2015 of the average number of years women in the US spent in parities 0, 1, and 2 comparing the three measures: period (dashed lines), cohort (dotted lines), and *CALP* (solid lines).

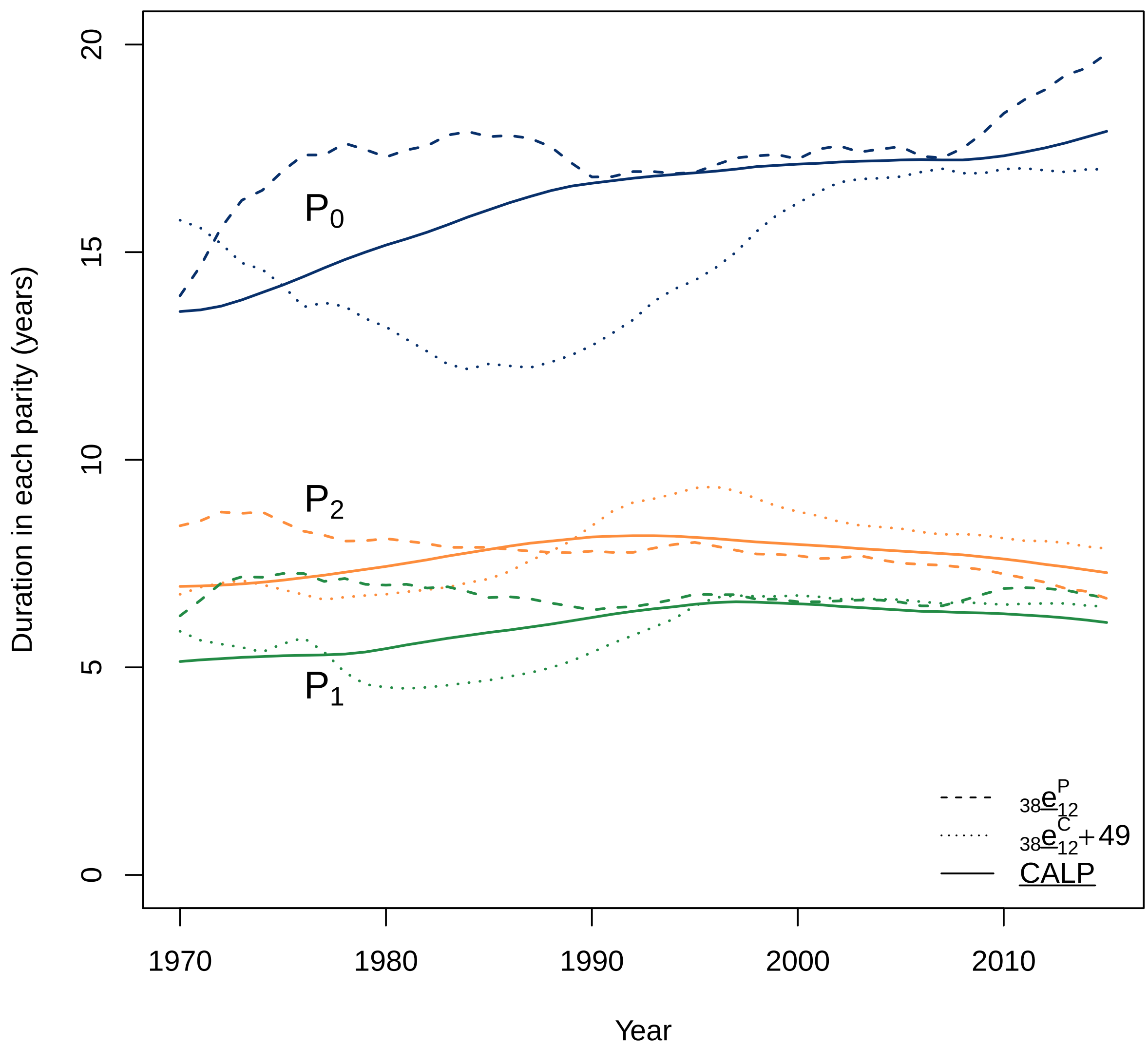



Figure 4.6: Time trends of duration in parities 0, 1, and 2 in the US from 1970 to 2015 for three different perspectives: period, cohort and CALP
*Source*: Authors' calculations using the HFD.

At parity 0, the period index has a hump between 1970 and 1990, while the cohort index has a U-shape. In the early 2000s, all three measures coincide, but in recent years, the period duration in childlessness has increased, followed by $CALP_0$. The average duration that women in the US spent in parity 0 in $CALP$ increased from 13.6 years to 17.9 years and had the smoothest time trends of the three measures. These trends in all three indexes can be generally seen in parities 1 and 2, albeit at different levels. An exception to this is the cohort index of parity 2, which starts at a lower level than the other two indexes but passes these indexes and remains higher after 1990. Again, $CALP_1$ and $CALP_2$ show the smoothest time trends as they integrate the cohort and period perspectives.

## 4.4 Discussion

This study aims to introduce an alternative index to estimate the duration women spend in each parity during their reproductive ages. The measure, Cross-sectional Average Length of Life by Parity ($CALP$), is a period measure including all the cohort age- and parity-specific birth information from age 12 to 50 at a given time. The US data from the Human Fertility Database were selected as an example to illustrate the use of $CALP$ to explain the fertility heterogeneity in a population of women of reproductive age. The $CALP$ for the year 2015 shows that women in the US spend 47% (17.91/38 years) of their reproductive years from ages 12 to 50 in childlessness, followed by 16%, 19% and 11% in parities 1, 2, and 3, respectively. $\underline{CALP}(2015)$ complements the traditional period and cohort indexes, and includes the parity-specific fertility information of all 38 cohorts (from the 1966 to 2003 birth cohorts and from parity 0 to parity 5+) in women of reproductive age in 2015. This clearly presents $CALP$ as a period fertility measure that includes historical cohort information.

Limitations of the proposed measure should be mentioned. A long history of birth data that extends several decades into the past is essential for the calculation of $CALP$, and the use of this measure for countries that do not have this amount of information is limited.

However, *CALP* can also be calculated as a truncated measure using incomplete cohort birth information, similar to the Truncated Cross-sectional Average Length of Life (TCAL) developed and used in mortality research (Canudas-Romo and Guillot 2015). Therefore, the truncated version of *CALP* could be more widely used in middle- or low-income countries where a detailed long history of fertility data is less often available.

Demographers have measured fertility phenomena using quantum (e.g. TFR), timing (e.g. mean age at birth), and rate (e.g. age-specific birth rate) indexes. This study introduced an alternative perspective, *duration*, to understand fertility changes over time. Reproductive decisions and timing depend significantly on age because humans have a time limit to give birth, called the "biological clock". Therefore, measuring the duration of parity statuses provides a detailed picture of family building. Future research could benefit by including the duration perspective as part of the analytical tools. For example, a multiple-country comparison of durations across parities could provide insight into the historical fertility paths followed by women of reproductive age. Additionally, researchers could ask about duration in states other than parity that influence fertility. For example, including states of union formation (e.g. cohabitation, marriage) in the model could show broader differences in the fertility behavior of reproductive-aged women that are not perceived by studying only parity. Traditional period indexes can present the current trend but ignore the past trajectories that define the fertility levels attained in a given year. Cohort measures can be considered outdated since they are mainly calculated for cohorts that have completed their reproductive life. In real life, people from different birth cohorts interact and influence each other on decisions regarding reproduction. Our measure, $\underline{CALP}(t)$, summarizes the entire fertility history of all reproductive-aged women present at that time. As such, *CALP* portrays the cohort dynamics of fertility in a comprehensive way.

# Appendix

The age- and cohort-specific transition rates from parity $i$ to parity $j$ ($j = i+1$) are denoted as $m_{ij}(a, t-x)$ and are calculated following the HFD protocol:

$$
\begin{aligned}
m_{01}(a, t-x) &= \frac{b_1^c(a, t-x)}{l_0^c(a, t-x) - 0.5b_1^c(a, t-x)}, \\
m_{hh+1}(a, t-x) &= \frac{b_{h+1}^c(a, t-x)}{l_h^c(a, t-x) - 0.5b_{h+1}^c(a, t-x) + 0.5b_h^c(a, t-x)}, \\
m_{45+}(a, t-x) &= \frac{b_{5+}^c(a, t-x)}{l_4^c(a, t-x) + 0.5b_4^c(a, t-x)},
\end{aligned}
$$

where $h = 1, 2$ or 3, $b_i^c(a, t-x)$ corresponds to the cohort life table function of the birth rate of women of parity $i$ at exact age $a$ born in year $t-x$, and $l_i^c(a, t-x)$ estimates the parity-specific number of persons at age $a$ from the cohort born in year $t-x$.

# Part III

# Conclusion

# Chapter 5

# Conclusion

This thesis aimed to bring an alternative perspective to the study of family demography by measuring the *duration* of union and fertility events and to develop new indexes for a better understanding of current family dynamics. Focused on first marriage, first birth, and parity-specific fertility behavior, I introduced three indexes: Expected Years Ever Married (EYEM) in Chapter 2, Expected Years Without Children (EYWC) in Chapter 3, and Cross-sectional Average Length of Life by Parity (CALP) in Chapter 4. Each chapter has already undergone a peer review by international journals: Chapter 2 "Expected Years Ever-Married" was published in *Demographic Research* (co-authored with Vladimir Canudas-Romo); Chapter 3 "Decomposing changes in first birth trends: Quantum, timing, or variance" was accepted in *Vienna Yearbook of Population Research* and is in press (co-authored with Michael del Mundo); and Chapter 4 "Cross-sectional average length of life by parity: Illustration for the US cohorts in reproductive ages in 2015" was accepted as a chapter in the book "Analyzing Contemporary Fertility" and is in press (co-authored with Vladimir Canudas-Romo). These three indexes measure the average duration of time that people spend in certain family-life statuses during their reproductive life: ever-married, remaining childless, and in each parity status, respectively.

## 5.1 Implications of this thesis

The introduction of these three alternative indexes may have implication in three respects: First, they are comprehensive measures; while conventional indexes are valuable in examining either quantum or timing of an event, these new indexes have a value in capturing changes in both. As both quantum and timing of union and fertility events have changed in most high-income countries, a comprehensive measure capturing both quantum and timing may provide additional insight regarding the current dynamics in union formation and fertility. Using the concept of life expectancy, which is seldom used in family demography (Andersson et al. 2017; Bongaarts and Feeney 2006), the level of these indexes indicate the total amount of life spent in each family-life status and take into account both individuals who have and have not experienced the event by the end of the reproductive period. In this way, these new indexes complement existing indexes of quantum and timing of union and fertility events, e.g. mean age at first marriage, the proportion of remaining never-married, mean age at first birth, and the proportion of remaining childless.

The second reason these indexes may have implication is concerned with the interpretation of family dynamics at a country level. Age is a key factor for the decision and timing of forming a union and giving birth especially because women have a limited window for childbirth, called the "biological fertility clock." Therefore, measuring the duration of the first marriage, first birth, and each parity-specific behavior provides an alternative perspective in family demography and a detailed picture of family building.

Finally, traditional period indexes can present the current trend, but ignore past trajectories, which define the fertility levels attained at a given year. Cohort measures can be seen to be outdated because they are mainly calculated for cohorts that have completed their reproductive life. In real life, people from different birth cohorts interact and influence each others' reproductive decisions. Using a measure developed in mortality research, Cross-sectional Average Length of Life, the index introduced in Chapter 4 summarizes, with a period perspective, the entire fertility history of all reproductive-aged women present at that time. As such, it portrays the cohort dynamics of fertility comprehensively.

## 5.2 Key findings of each chapter

Chapter 2 presented that women in Denmark, France, Germany, the Netherlands, and Sweden and men in Austria, Czechia, Denmark, France, Germany, Italy, Ireland, the Netherlands, the UK, and Sweden are never married during half of their reproductive lives. The changes in EYEM over time was decomposed into three components: scale (the changes in the proportion of never-married, i.e. nonmarriage), location (the changes in the timing of first marriage, i.e. delayed marriage), and variance (the changes in the standard deviation of first marriage age, i.e. expansion). The result shows that in most countries and periods the decline in period EYEM is mainly due to delayed marriage. However, new trends are seen in selected countries, with the nonmarriage component influencing results in Northern Europe, Canada and in most West European countries. The expansion effect has practically no influence on the changes in EYEM.

In Chapter 3, we found that women born in the latest cohorts observed in Canada, Japan, and the Netherlands spent half of their reproductive periods without any children. Furthermore, we decompose the changes in EYWC over time into three effects: remaining childlessness, postponing first birth, and expansion of the standard deviation of mean age at first birth. Results of the decomposition show that postponement is the most influential factor on EYWC's changes in North America and Northern Europe while remaining childless is the main contributor in Japan and Portugal. In Chapter 4, CALP for the year of 2015 shows that women in the US spent 47 % (17.91/38 years) of reproductive years (ages 12 to 50) in childlessness, followed by 16 %, 19 % and 11 % in parities 1, 2, and 3, respectively.

## 5.3 Limitations and future research

Limitations of this study should be mentioned. First, Chapter 2 does not include information about cohabiting couples. Thus, the decrease in EYEM is likely due to the increase in cohabitation; this speculation can be applied especially to the Northern Europe and Canada and in most of the West European countries in which cohabitation prevails. However, cohabitation can be included in the method once data with information on cohabitation is available. Another limitation is the accuracy of fitting observed data to the Coale-McNeil model, which was used for the estimation of the age-specific first marriage and first birth patterns. Several parametric models give better estimation (e.g. Kaneko (2003) and Peristera and Kostaki (2015)). However, both EYEM and EYWC can be calculated using the life table without using the Coale-McNeil model. Thus, a sensitivity analysis using the life-table method will be helpful to justify the usage of the Coale-McNeil model. Finally, the calculation of CALP requires a long history of fertility data, limiting its applicability in countries with limited data, such as middle or low-income countries. However, CALP can also be calculated as a truncated measure using incomplete cohort birth information, similar to the Truncated Cross-sectional Average length of Life (TCAL) already developed and used in mortality research (Canudas-Romo and Guillot 2015).

This thesis suggests several possibilities for future research. First, it will be worthwhile to calculate the three indexes by subpopulations, such as by educational and race/ethnicity groups. Several studies reported that union and fertility behaviors differ by education (see Raymo (2003) or Beaujouan et al. (2016)). Thus, the three indexes by subpopulations will show more detailed changes in union and fertility events. Secondly, these indexes can be applied to other union and fertility events, such as divorce or emancipation. One can ask "How many years do people spend in their first marriage before getting a divorce?" or "How many years do people spend in a family home before leaving for the first time?" respectively. Moreover, it can be important to ask about the duration in other states that influence fertility. For example, including the states of union formation (e.g. cohabitation, marriage, etc.) into the model may reveal broader differences in fertility behavior of reproductive-aged women that are not perceived when only parity included.

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