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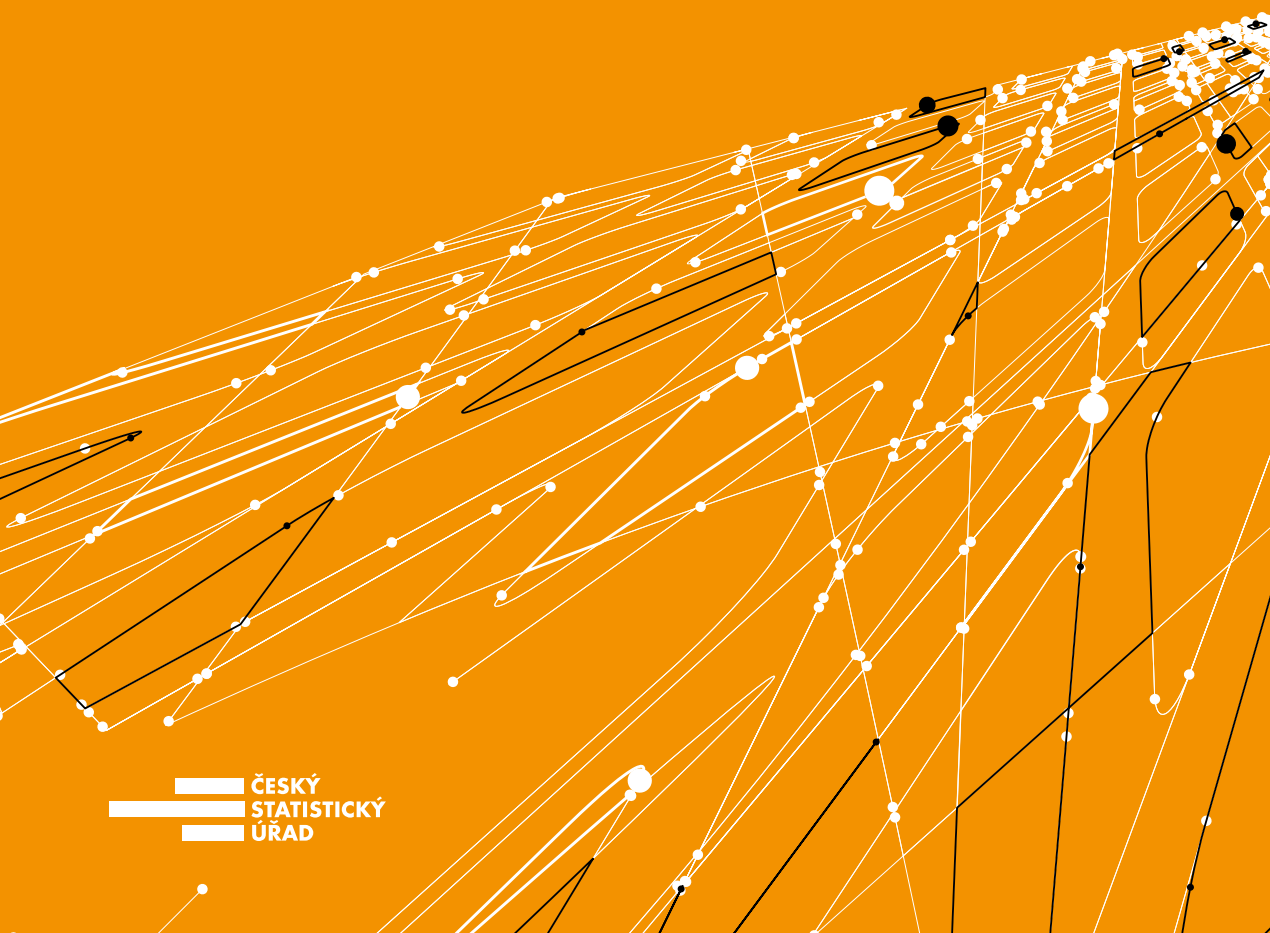

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A COMPARATIVE ANALYSIS OF DIVORTIALITY PATTERNS AND ATTITUDES TOWARDS DIVORCE IN GREECE, ITALY, SPAIN AND FRANCE

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Abstract

Since the mid-1970s, family patterns in the European South have undergone substantial change, with some increasing similarity in aggregate demographic indicators but also persistent cross-national diversity. In Mediterranean countries, declining marriage and fertility rates began in the 1970s, followed by an increase in divorce since the 1980s, although the timing and pace of change differ across countries. Against this background, the present study examines changes in divortiality in Greece, Italy, Spain, and France – a country that shares some geographic and cultural proximity with Southern Europe but also differs in important institutional and historical respects. The paper begins with a demographic analysis of divorce since the 1950s (period and cohort analysis). The second part explores attitudes toward divorce. Traditional demographic analysis, descriptive statistics, chi-square tests, binary logistic regression, and several machine learning algorithms (XGBoost, CatBoost, LightGBM and Random Forest) were employed. Our findings indicate that divorce has been more prevalent since the 2010s in France and Spain, while Greece and Italy remain at comparatively lower levels and experienced increases at a later stage. Attitudes toward divorce are also more positive in France and Spain. Age, sex, religiosity, marital status, and views regarding the deinstitutionalisation of marriage are associated with attitudes toward divorce. The machine learning models further indicate that age and religiosity are the most influential features contributing to the prediction of unconditional acceptance of divorce. Finally, we show that legal reforms have played an important role in shaping the observed trends in divortiality.

Keywords: divortiality, attitudes towards divorce, family law, religion, Machine Learning Algorithms
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INTRODUCTION

Since the 1960s and 1970s, divorce rates have increased across many European societies, although the timing, pace, and institutional contexts of change have varied substantially. Legal reforms facilitated marital dissolution in several countries, reducing state interference in private life and removing formal barriers to divorce (*Festy – Prioux,* 1975; *Sardon,* 1986). While some attributed rising divorce to a desire to remarry (*Berger,* 1980), remarriage rates increased more slowly, revealing more complex societal shifts (*Roussel,* 1987). Divorce rates in many countries began to rise even before major legislative reforms, suggesting that legal changes often followed–rather than initiated–broader societal transformations (*Festy,* 1983; *Sardon,* 1986). Nonetheless, specific law reforms did at times produce marked short-term fluctuations. For example, the shortening of divorce proceedings in Sweden in 1974 led to a temporary surge in divorce registrations, whereas in the Federal Republic of Germany the introduction of a new divorce law in 1977 initially produced a decline due to an accumulation of applications and legal ambiguities (*Monnier,* 2006). These examples illustrate how legislation interacts with underlying social change rather than operating in isolation.

In Southern Europe, divorce historically remained comparatively less prevalent, reflecting strong family norms, higher levels of religiosity, and later institutional recognition of divorce. Greece, Italy, and Spain shared a pattern of relatively late legal liberalisation and enduring cultural attachment to marriage. Over recent decades, however, divorce intensity has increased in all three countries. These developments raise questions about whether Southern European societies are undergoing similar processes of transformation or maintaining distinct national patterns shaped by institutional and cultural contexts.

France provides a useful contrast in this setting. Although part of France lies in the Mediterranean, it is commonly classified as Western European due to its distinct historical, demographic, and institutional developments. France is not included as a representative of Western Europe as a whole, nor as a normative benchmark, but as a theoretically relevant contrast case. Its earlier secularisation, long-standing divorce legislation, and specific Church–state relations provide an institutional configuration that differs from Southern European settings. The comparison therefore does not aim to generalise to "Western Europe," but to contrast Southern European patterns with this distinct case in order to assess how institutional and cultural contexts shape divorce behaviour and divorce attitudes.

The transformation of divorce patterns has been closely linked to broader theoretical debates in demography and sociology. Gender Equality Theory posits that more egalitarian gender roles increase relationship stability and fertility (*Esping-Andersen,* 2009; *Esping-Andersen – Billari,* 2015; *Goldscheider et al.,* 2010). The Second Demographic Transition (SDT) theory further explains changing family forms through shifts in values such as individualisation and secularisation (*Lesthaeghe,* 2010, 2014; *Surkyn – Lesthaeghe,* 2004; *Van de Kaa,* 1987; 2001). Scholars also highlight evolving gender roles and economic uncertainty as key drivers of demographic change (*Bernhardt,* 2004; *Perelli-Harris – Gassen,* 2010). From the 1960s onward, marriage became increasingly fragile as individualistic values reshaped expectations, with individuals remaining married only insofar as their personal goals and aspirations were fulfilled. Within this context, rising female employment has often been linked to greater economic independence and, consequently, higher divorce risks (*Becker,* 1993), although empirical evidence is not uniform. For instance, women's employment is positively associated with divorce in Poland but not in Hungary (*Vignoli et al.,* 2018). While Becker argued that women's labour-force participation reduces the "utility" of marriage, subsequent studies point to more complex dynamics, suggesting that in some settings women's improved education and income may instead support the formation of unions with partners who share similar aspirations and values (*Blossfeld,* 2009; *Oppenheimer et al.,* 1997).

Finally, growing union instability influences fertility patterns–both by reducing the number of children married couples may have, as it shortens the duration of marital life available for family formation and increasing them through re-partnering (*Oláh,* 2015; *Thomson,* 2004). However, this phenomenon will not be explored in detail in the current study. Overall, marital instability has contributed to the diversification of family forms across Europe, including single-parent and stepfamilies, within and beyond marriage.

1. Attitudes towards divorce

Attitudes toward divorce constitute an essential normative dimension of these processes. Normative acceptance both reflects and potentially reinforces behavioural change.

This study is grounded in theoretical perspectives associated with the Second Demographic Transition and the deinstitutionalization of marriage, which emphasize the growing role of individual values, secularization, and autonomy in shaping family behaviors. While changing values are often conceptualized as drivers of demographic change–such as rising divorce rates and earlier marital dissolution–the relationship between values and behavior is likely to be dynamic and reciprocal. Increasing divorce prevalence may itself contribute to the normalization of marital dissolution, reducing social stigma and reinforcing more permissive attitudes toward divorce. From this perspective, demographic behavior and attitudinal change are best understood as mutually reinforcing processes rather than as a simple unidirectional causal chain. Accordingly, the combined analysis of divorce behaviors and attitudes toward divorce, using both traditional demographic indicators and individual-level modeling approaches, allows us to examine these intertwined processes from complementary analytical perspectives. The inclusion of machine learning techniques further enables the identification of complex, nonlinear relationships and interaction patterns that may characterize value–behavior dynamics across different social and institutional contexts.

Against this background, the present study examines both divorce behaviour and attitudes toward divorce in Greece, Italy, Spain, and France.

The study addresses five key questions:

1. To what extent do divorce patterns differ across Greece, Italy, Spain, and France in terms of level, timing, and pace of change?
2. Have divorce intensity and timing changed since the 1970s in these countries?
3. Are recent marriage cohorts divorcing earlier or more often than older ones?
4. How has legislation influenced divorce trends?
5. How do attitudes toward divorce vary across these countries?

To assess attitudes, we use data from the European Values Study (EVS 2017), based on the item "Do you justify divorce?" Respondents answered on a 1 ("never") to 10 ("always") scale. We define unconditional acceptance as selecting the value "10," which represents a substantial share of respondents in each country.

We investigate factors associated with these attitudes. In addition to country, we examine age, gender, marital status, religiosity, and views regarding the deinstitutionalisation of marriage. The belief that marriage is outdated draws on *Cherlin's* (2004) discussion of the deinstitutionalisation of marriage and has been linked to greater support for divorce (*Kaufman et al.,* 2024). Religiosity is included given its established association with marital stability and lower tolerance of divorce (*Guarneri et al.,* 2021; *Kapinus – Pellerin,* 2008; *Stokes – Ellison,* 2010; *Vermeulen et al.,* 2023). However, it remains unclear whether the same relationship holds for couples and whether the effect of couple-level religiosity depends on the broader contextual level of religiosity (i.e., a cross-level interaction effect).

We adopt a mixed-methods quantitative design: first, we document long-run divorce dynamics using demographic period and cohort indicators; second, we examine individual-level attitudes toward divorce using regression models for adjusted associations and machine learning algorithms to assess predictive structure and variable importance.

2.1. Demographic analysis of divortiality (period and cohort/longitudinal analysis)

To analyse the intensity (the frequency of the dissolution of marriages) and timing (the median duration of marriage at its dissolution) of divorce, we used data on divorces by reached duration of marriage and the total number of marriages from which these divorces originated. Divorce data were available for France (1952–2016), Italy (1971–2022), Greece (1972–2023), and Spain (1981–2022). French data post-2016 became incomplete following the introduction of out-of-court divorces (*Breton et al.,* 2022). Marriage data were available from the 1890s in France, the 1910s in Italy and Greece, and the 1920s in Spain.

Data sources included the European Demographic Observatory (*EDO,* 2011), *EUROSTAT* (2023), *LADS* (2018), and national statistical agencies/institutes (*ELSTAT,* 2023; *INE,* 2023; *INED,* 2023; *INSEE,* 2023a; *ISTAT,* 2023). We applied classical methods of the French School of Demography, notably the Total

Divorce Rate (TDR), which estimates the number of divorces per 10,000 marriages based on observed divorce rates by duration of marriage:

$$TDRi = \sum_{k=1}^{60+} divorce\_rate,$$

where i denotes the calendar year and k the duration of marriage. The TDR provides an indicator of the period-specific intensity of divorce, while the median duration of marriage at dissolution captures the timing of divorce within a given year.

Figure 1 illustrates how duration-specific divorce rates are constructed. Each parallelogram in the Lexis diagram (e.g. E, D, C) represents the number of divorces occurring in a given calendar year at a specific reached duration of marriage. The horizontal segments labelled MC and MD show the marriages from which these divorces originate (i.e. the marriage cohorts that have reached the relevant durations in the observed year) in 2010. Thus, parallelogram C shows the divorces occurring at a duration of two years. To obtain the divorce rate at duration two, the number of divorces in parallelogram C is related to the marriages represented by MC, corresponding to marriages formed in 2008 (the cohort that reached two years of marriage in 2010). Similarly, parallelogram D shows divorces at a duration of three years, with the corresponding marriages at risk represented by MD (marriages formed in 2007).

The same logic applies to all other durations: each parallelogram contains divorces observed at duration *k*, and each horizontal segment identifies the marriages from which those divorces originate. After determining all duration-specific rates for 2010, these rates are summed to obtain the Total Divorce Rate for that year. The TDR thus represents the intensity of divorce that would be observed in a hypothetical marriage cohort if it were exposed to the duration-specific divorce rates recorded in that calendar year, in the absence of mortality and migration.

Figure 1 **Lexis Diagram (analysis of divortiality)**

The median duration of marriage at divorce was calculated in the same way as a conventional median derived from a set of numerical values. In this case, the values correspond to the duration-specific divorce rates observed in a given calendar year (e.g. rates at durations 0, 1, 2, 3, …, 60+ in 2010). These rates were ordered by duration, and the median was identified as the duration at which half of the cumulative divorce

intensity was observed at shorter durations and half at longer durations.

Alongside period analysis, we conducted cohort-based longitudinal analyses. For each marriage cohort (e.g. 1972), divorce rates were calculated for successive durations of marriage (e.g. first year, second year, etc.). This approach allows the estimation of cohort-specific patterns in the intensity and timing of divorce for marriage cohorts observed over time: 1972–1994 (Greece), 1971–1993 (Italy), 1981–1993 (Spain), and 1952–1987 (France). These analyses cover the first 30 years of marriage, during which the majority of divorces occur—94% in Greece and France, 85.9% in Italy, and 88.9% in Spain. Divorce rates for durations beyond 30 years were imputed using information from more recent observation periods (e.g. 2014–2022 for Greece).

2. 2. Examination of the attitudes towards divorce

Beyond divorce trends and timing, we analysed attitudes toward divorce using the European Values Study (EVS 2017). For France, Italy, and Spain we rely on the Joint EVS/WVS 2017–2021 release (EVS/WVS, 2021), while Greece is drawn from the EVS 2017 Greece country file (EVS, 2020), which employed snowball sampling rather than the multistage or simple random sampling used in the other countries. Accordingly, country-specific inferential results for Greece should be interpreted with caution, and the pooled models are understood as capturing associations within the combined sample rather than providing population-level estimates for Greece. Greece was nevertheless retained in the comparative analysis due to the dataset's substantial size and geographic coverage, including both urban and rural areas across the country.

Descriptive statistics, chi-square tests, and binary logistic regression were applied to examine cross-national variation in attitudes toward divorce and to identify socio-demographic factors associated with these attitudes. To complement these traditional statistical techniques, several machine learning (ML) algorithms were also employed to assess the relative importance of predictors in shaping highly positive attitudes toward divorce, allowing for potential non-linearities and interactions among covariates. Specifically, we implemented XGBoost (Extreme Gradient Boosting) (*Chen et al.,* 2024; *Chen – Guestrin,* 2016; *Liu – Just,* 2019; *Lundberg – Lee,* 2017), LightGBM (Light Gradient Boosting Machine) (*Ke et al.,* 2017), CatBoost (Categorical Boosting) (*Dorogush – Ershov – Gulin,* 2018; *Prokhorenkova et al.,* 2018), and Random Forests (*Breiman,* 2001; *Wright – Ziegler,* 2017), all of which are ensemble methods based on decision trees and are widely used in classification tasks. Detailed descriptions of the data sources and analytical procedures are provided in the sections devoted to the empirical analysis.

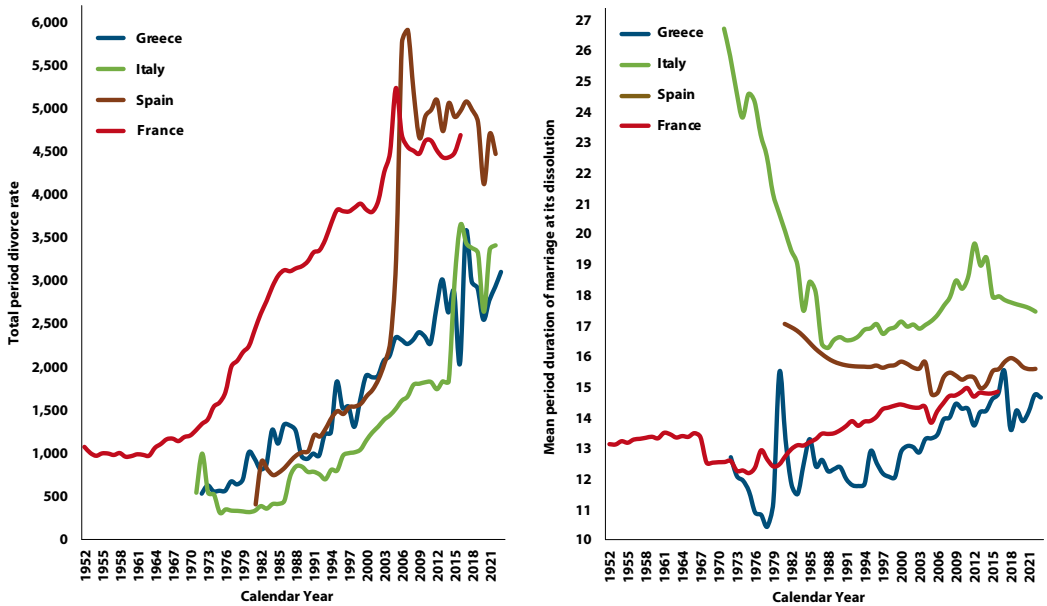
The decision to focus on unconditional acceptance of divorce rather than modelling the full ten-point scale reflects a deliberate analytical choice rather than a methodological constraint. Our primary interest lies in identifying the socio-demographic correlates of unequivocal normative approval of divorce, understood here as a distinct attitudinal position rather than as the upper end of a continuous scale. While alternative modelling strategies using the full scale (e.g. ordinal or continuous specifications) are possible, they address related but conceptually different questions concerning gradations of approval. Accordingly, we acknowledge that this operationalisation entails a loss of information and does not capture heterogeneity within intermediate response categories, a limitation that should be borne in mind when interpreting the results. Finally, additional details on data preparation and model implementation are provided in the relevant analytical sections.

RESULTS

3.1. The dissolution of marriages: demographic analysis of divorce and the impact of legislation

Divorce became widespread earlier in France than in Greece, Italy, or Spain (Fig. 2). By the mid-1980s, approximately 30% of French marriages ended in divorce (TDR ≈ 3,000 divorces per 10,000 marriages), while divorce intensity in Southern European countries remained below 13%. Spain experienced a pronounced increase in divorce intensity from the mid-2000s onward. After 2010, divorce affected around 46% of marriages in France, 49% in Spain, 27% in Italy, and 28% in Greece. The onset of the COVID-19 pandemic in 2020 led to a temporary decline in divorce rates across Southern Europe, likely reflecting the suspension

Figure 2 **Total divorce rate (number of divorces per 10,000 initial marriages, period 1952–2023) – left, median marriage duration at its dissolution – right**



Source: EDO, 2011; ELSTAT, 2023; EUROSTAT, 2023a; INE, 2023; INSEE, 2023b; ISTAT, 2023; LADS, 2018; own calculations.

of judicial proceedings rather than changes in marital behaviour.

Unlike fertility, divorce intensity is closely linked to changes in family law (*Kontogiannis*, 2022). In Greece, legislation prior to 1979 imposed strong restrictions on marital dissolution. Law 868/1979 represented a major institutional change, allowing divorce after six years of separation without proof of fault. Following its introduction, the Total Divorce Rate rose from 631 per 10,000 marriages in 1978 to 1,000 in 1980. Over the same period, the median duration of marriage at divorce increased from 6.7 years to 12 years, reflecting the exit of long-standing marriages that had previously been legally constrained. The effect of this reform was temporary, as the law remained in force for only six months, although many divorces initiated during that period were finalised in 1980 or 1981.

Greece: legislative shifts and rising divorce intensity

Although divorce rates declined slightly in 1982–1983, they remained above 1970s levels, suggesting broader changes in social norms. Law 1329/1983 introduced consensual divorce, leading to a renewed increase in divorce intensity (exceeding 1,000 divorces per 10,000 marriages between 1984 and 1988) and a rise in the median duration of marriage at divorce to 9.6 years by 1985. Since the late 1980s, divorce intensity in Greece has followed a generally upward trajectory, interrupted by short-term fluctuations, such as the sharp decline in 2016 caused by a temporary abstention of lawyers from court proceedings.

Subsequent legal reforms, including Laws 3719/2008 and 4055/2012, reduced the minimum duration of marriage required for divorce. Law 4509/2017 further simplified consensual divorce procedures by allowing their completion through notaries, substantially reducing administrative barriers.

Before 1983, mutual-consent divorce was not legally permitted in Greece. Couples seeking amicable separation often resorted to formally contested divorces by attributing fault–a practice commonly described as "pseudo-consensual" divorce *(crypto-synainetika)*. Although systematic data are unavailable, media reports suggest an increase in so-called "white divorces" *(leyka diazygia)* during the post-2009 economic crisis,

whereby couples formally divorced while continuing to cohabit in order to address financial or legal constraints.

France: historical evolution and the shift to consensual divorce

In France, judicial separation was permitted in the seventeenth and eighteenth centuries, while divorce itself was prohibited under Canon Law (*Sardon,* 2005). Divorce was legalised in 1792 during the French Revolution, allowing for both consensual and incompatibility-based separation ("incompatibilité d'humeur") (*von Bóné,* 2014) abolished in 1816 during the Bourbon Restoration (*Carella et al.,* 2014), and reintroduced in 1884 in a fault-based form. Subsequent reforms progressively expanded access to divorce, culminating in the 1975 law (effective 1976), which reintroduced consensual divorce and broadened the grounds for marital dissolution (*Sardon,* 2005). Further simplifications were introduced in 2004, and since 2017 consensual divorce may proceed without judicial involvement (*Breton et al.,* 2018).

Sardon (2005) argued that the 1976 reform accelerated an already ongoing rise in divorce rates (Fig. 2). Divorce intensity continued to increase through the 1980s and 1990s, with further changes observed after 2005 (*Bellamy,* 2016). Simplification of procedures coincided with a decline in fault-based divorces, which fell from 42.6% in 1999 to 7% in 2015, indicating a shift in both legal practice and social norms surrounding marital dissolution (*Breton et al.,* 2017).

Spain: from prohibition to the 2005 "express divorce" reform

In Spain, divorce was briefly legal between 1934 and 1936 and was reintroduced in 1981 (*Flaquer – Garriga,* 2009; *Spijker – Solsona,* 2012). Prior to 1981, marriage was largely indissoluble, with judicial separation or ecclesiastical annulment as the only available options (*Bernardi – Martínez-Pastor,* 2011; *Carella et al.,* 2014). The 1981 law legalised divorce but imposed procedural and financial constraints that limited its immediate diffusion (*Roigé,* 2013).

A major institutional change occurred in 2005 with the introduction of so-called "express divorce" (*Duato – Jódar,* 2013; *González-Val – Marcén,* 2017), which eliminated the requirement for prior judicial separation and allowed consensual divorce after three months of marriage (*Carella et al.,* 2014). This reform substantially altered the institutional context of marital dissolution. Following its implementation, divorce intensity in Spain increased markedly, reaching levels comparable to those observed in other Western and Northern European countries (*Roigé,* 2013). Although the Catholic Church continued to oppose divorce, its influence on family behaviour declined over time (*Dominguez-Folgueras – Castro-Martin,* 2013).

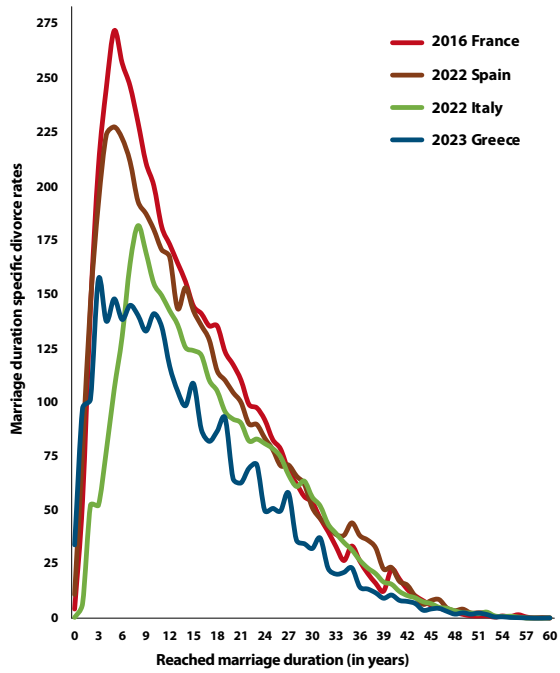
Italy: gradual liberalisation and late divorce timing

In Italy, marriage was legally indissoluble until 1970, when divorce was introduced through Law 898 (*Castiglioni – Dalla Zuanna,* 2008). This law established separate procedures for judicial separation and divorce, with a mandatory five-year separation period (*Tamanza et al.,* 2013). Subsequent reforms in 1975, 1978, and 1987 expanded the grounds for divorce and reduced the required separation period to three years, contributing to a noticeable increase in divorce intensity between 1987 and 1991 (*Carella et al.,* 2014).

Guarneri et al. (2021) noted that judicial separation, often the first step toward divorce in Italy, was sometimes the only one, as couples avoided divorce due to its financial and administrative costs. Divorce became necessary mainly when remarriage was desired. *Salvini and Vignoli* (2011) found that only 60% of couples separated in 1995 had divorced within ten years. Further reforms in 2014 and 2015 simplified procedures and reduced the separation period to six months for consensual divorces and twelve months for contentious divorces (*Colonnelli,* 2015). Following these changes, divorce intensity increased, reflecting both institutional adjustments and changing social attitudes toward marital dissolution.

Since 2010, the median duration of marriage at divorce has been broadly similar across Greece, France, and Spain (11.3, 11.6, and 12.1 years, respectively), while remaining higher in Italy (15.4 years). As shown in Figure 3, divorce intensity across countries is generally low during the first year of marriage, peaks between the fourth and ninth year (fourth year in Greece, sixth in France and Spain, ninth in Italy), and declines thereafter.

Figure 3 **Distribution of marriage duration specific divorce rates, Greece (2023), Italy (2022), Spain (2022) and France (2016)**



Source: EDO, 2011; ELSTAT, 2023; EUROSTAT, 2023a; INE, 2023; INSEE, 2023b; ISTAT, 2023; LADS, 2018; own calculations.

An analysis of divorce intensity by marriage duration (1952–2023) shows that, across all countries examined, divorces are most frequently observed between the sixth and tenth year of marriage, occur less often between the eleventh and fifteenth year, and are relatively rare thereafter (Fig. 4). The first five years of marriage (durations 0–4) display notable cross-national variation, which appears closely related to differences in legal frameworks and procedural requirements.

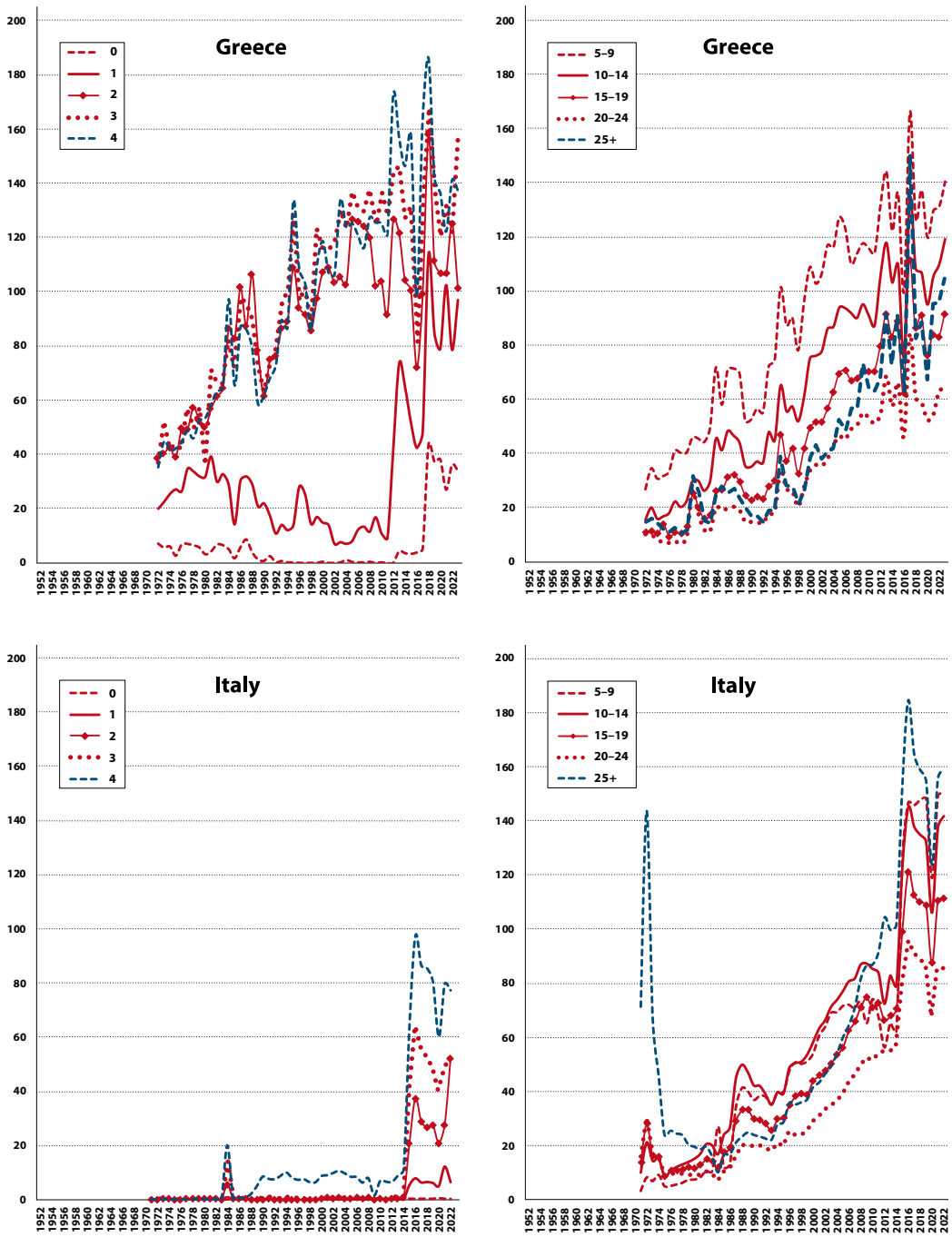
Italy exhibits comparatively low divorce intensity during the early years of marriage, a pattern consistent with the historically restrictive legal context and the widespread use of judicial separation as an intermediate or alternative arrangement (*Guarneri et al.,* 2021). This institutional configuration is also reflected in a higher median duration of marriage at divorce. Following the 2015 reform, which shortened the required separation period, early divorce rates increased modestly, suggesting a gradual adjustment to the revised legal framework.

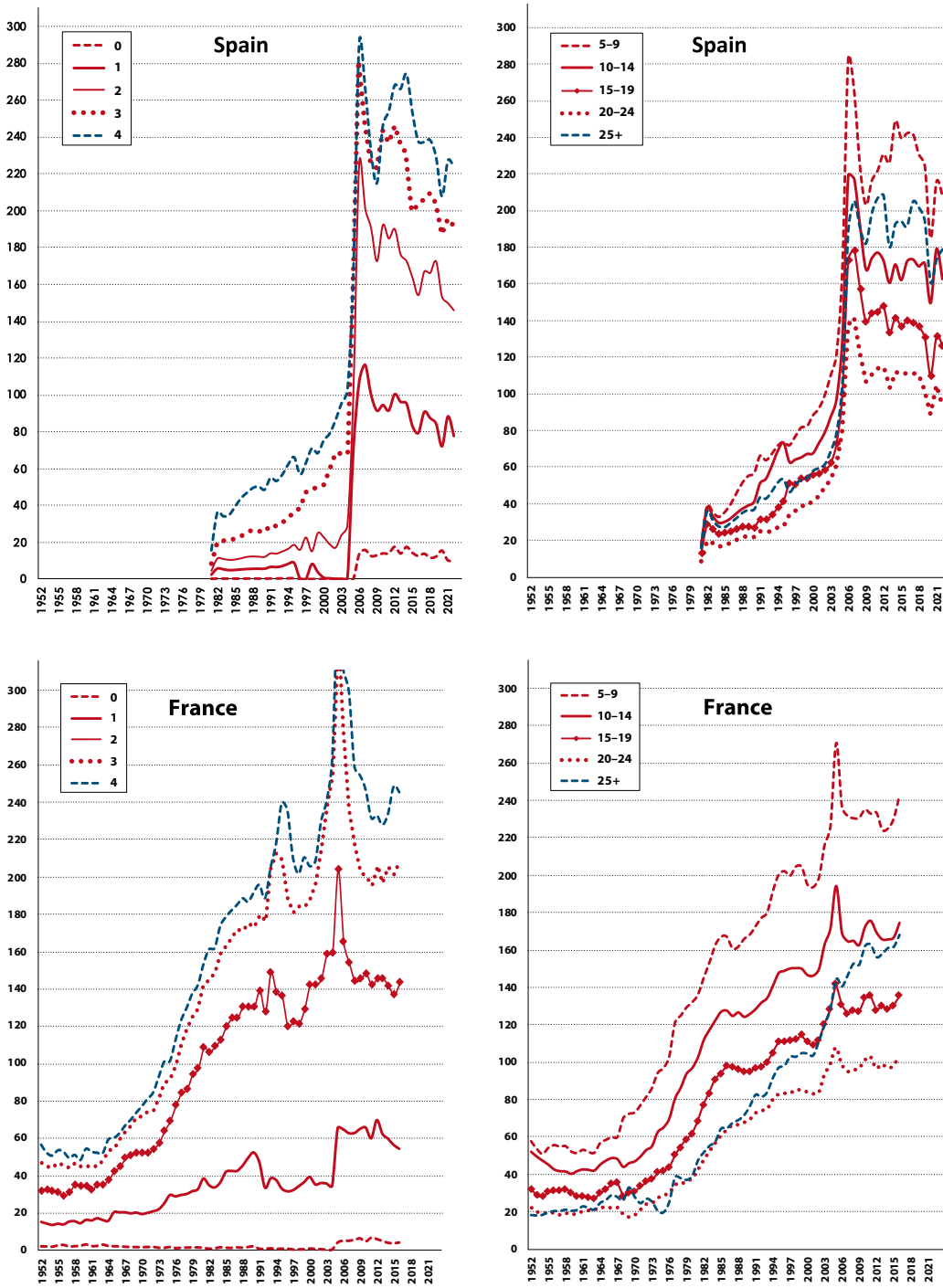
In Greece, divorce intensity during the early years of marriage–particularly around the second year–rose following Law 4055/2012, which reduced the minimum duration of marriage required for divorce. A more pronounced increase in divorces within the first two years of marriage is observed after 2018, coinciding with the implementation of Law 4509/2017, which further simplified divorce procedures. Unfortunately, comparable post-2016 French data are not available, preventing a direct assessment of the effects of the 2017 reform on early divorce timing in France.

An analysis of marriage cohorts[5)] formed between 1952 and 1994 highlights systematic differences in di-

5) A marriage cohort or cohort of marriages consists of all marriages (first, second, third order marriages etc.) contracted during a year.

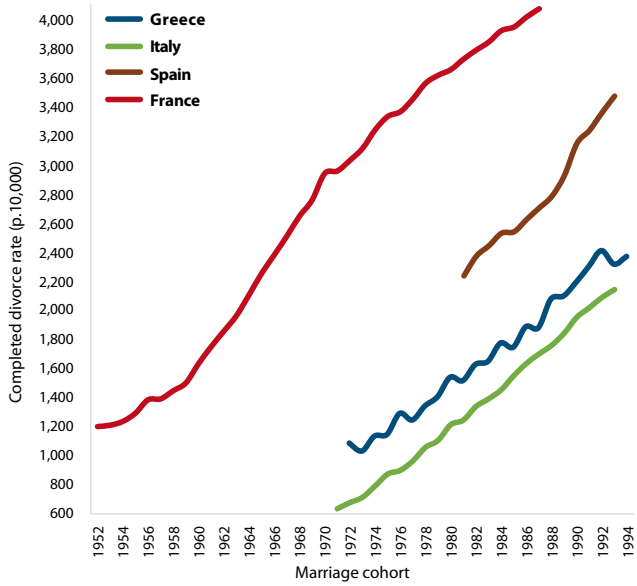
Figure 4 **Divorce rates by duration of marriage (years of marriage: 0, 1, 2, 3, 4 and 5–9, 10–14, 15–19, 20–24, 25+), 1952–2023, in Greece, Italy, Spain, and France**





Note: The 5–9 duration category is constructed as the mean of durations 15, 16, 17, 18, and 19. An identical averaging procedure was applied to all other duration categories (10–14, 15–19, 20–24, and 25+).
Source: EDO, 2011; ELSTAT, 2023; EUROSTAT, 2023a; INE, 2023; INSEE, 2023b; ISTAT, 2023; LADS, 2018; own calculations.

Figure 5 **Completed divorce rate (number of divorces per 10,000 marriages, cohorts of marriages 1952-1994), Greece, Italy, Spain and France**



Note: Sum of marriage duration specific divorce rates, cohort view.
Source: EDO, 2011; ELSTAT, 2023; EUROSTAT, 2023a; INE, 2023; INSEE, 2023b; ISTAT, 2023; LADS, 2018; own calculations.

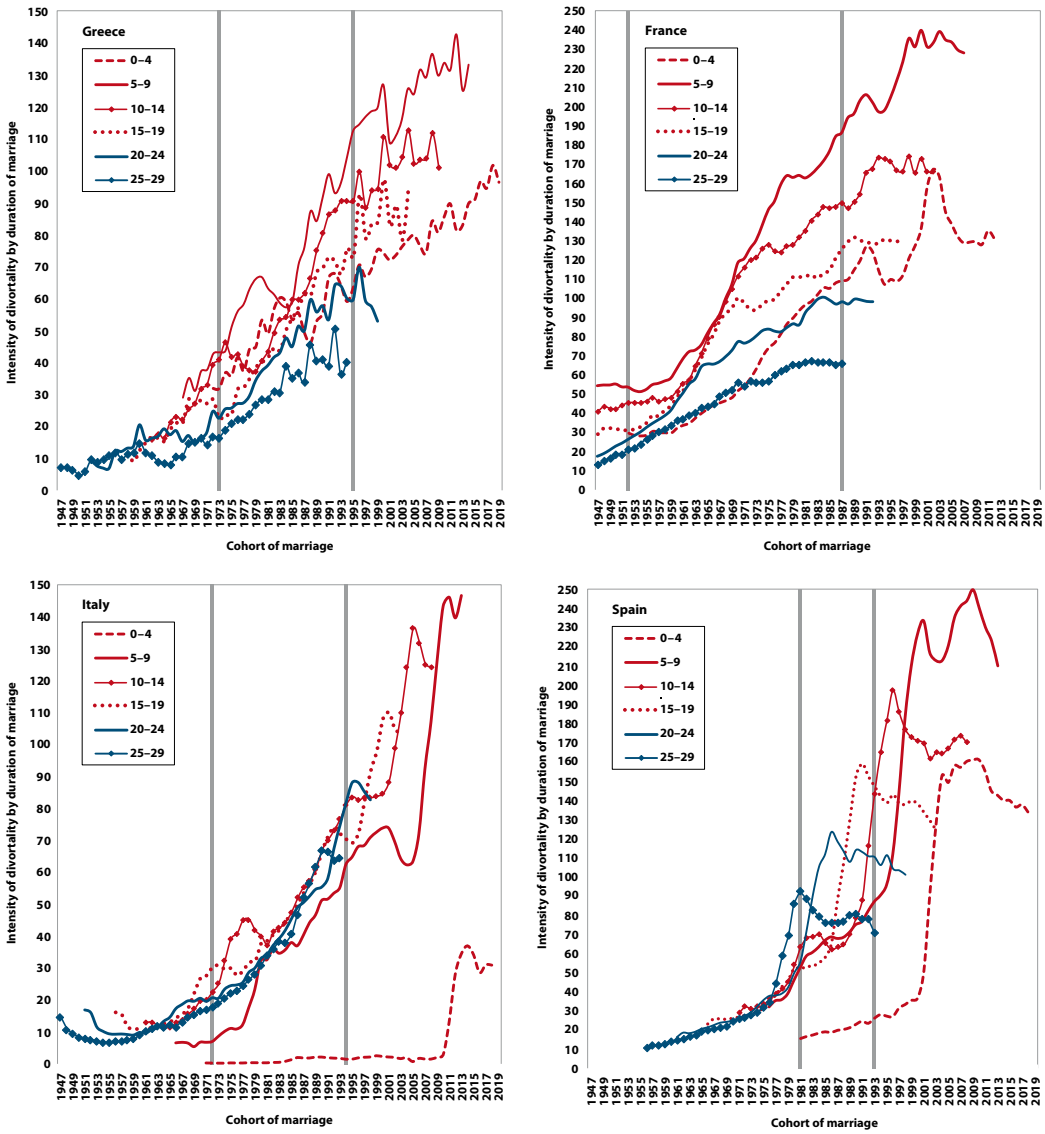
vorce outcomes across cohorts and countries. Among couples married in 1987, approximately 41% had divorced in France, compared with 27% in Spain, 17% in Italy, and 18.7% in Greece (Fig. 5). For more recent cohorts, divorce intensity is higher in all four countries, although the magnitude of change varies cross-nationally. By the 1993 cohort, 34.7% of Spanish couples, 21.4% of Italian couples, and 23.2% of Greek couples had experienced marital dissolution, with divorce intensity in Greece reaching 23.8% in the 1994 cohort. Overall, these cohort comparisons indicate that divorce has become more prevalent among marriages formed in later periods, while maintaining persistent differences in levels across countries.

Figure 6 presents divorce rates by marriage cohort across all countries. For Greece, data availability varies by cohort and duration: for cohorts married between 1947 and 1971, information is available only for later durations (e.g. 5–9 and 10–14 years), as divorces occurring during the earliest years of marriage (0–4) predate 1972, the first year for which divorce data are available. Conversely, for cohorts married between 1995 and 2019, data are available for early durations but not for later ones (25+ years), as these durations extend beyond 2023, the most recent year of observation.

For Greece, cohorts married between 1947 and 1971 display lower divorce intensity at the observed durations than cohorts married from 1972 onward. Comparable cohort differences are observed in Spain (1965–1980 versus post-1981 cohorts), Italy (1947–1970 versus early 1970s cohorts), and France (1947–1951 versus early 1950s cohorts). Final divorce intensities were calculated only for cohorts with at least 30 years of observation: Greece (1972–1994), Italy (1971–1993), Spain (1981–1993), and France (1952–1987).

Although data limitations prevent the calculation of completed divorce intensities for more recent cohorts, the available evidence allows for preliminary observations regarding cohort differences in divorce timing. In particular, cohorts married after 1994 in Greece, after 1993 in Italy and Spain, and after 1987 in France exhibit higher divorce rates at early dura-

Figure 6 **Divorce rates by duration of marriage (1st, 2nd, …,21st duration), cohorts of marriage 1947–2019, Greece, Italy, Spain and France**

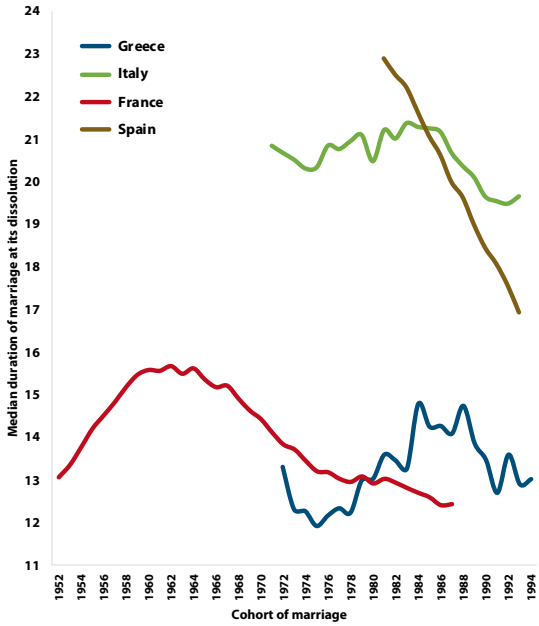


Source: EDO, 2011; ELSTAT, 2023; EUROSTAT, 2023a; INE, 2023; INSEE, 2023b; ISTAT, 2023; LADS, 2018; own calculations.

tions compared with older cohorts. These patterns are consistent with changes in legal frameworks and broader social contexts over recent decades, although further observation is required to assess their long--term implications.

Couples in Italy and Spain tend to experience marital dissolution at later durations than those in Greece and France (Fig. 7), a pattern that is consistent with differences in institutional contexts and the timing of legal reforms. In Italy and Spain, divorce legisla-

Figure 7 **Median duration of marriage at its dissolution, cohorts of marriages 1952–1994, Greece, Italy, Spain and France**



Source: EDO, 2011; ELSTAT, 2023; EUROSTAT, 2023a; INE, 2023; INSEE, 2023b; ISTAT, 2023; LADS, 2018; own calculations.

tion remained comparatively restrictive until 2015 and 2005, respectively, and judicial separation often preceded divorce, particularly in Italy. As a result, divorces among couples married in the 1970s and 1980s were more frequently observed after longer marriage durations.

For more recent cohorts, divorce occurrences are observed at earlier durations in both countries compared with older cohorts. While completed divorce trajectories for these cohorts cannot yet be assessed, available evidence indicates a concentration of divorces at shorter durations than in earlier cohorts. These patterns coincide with changes in legal frameworks and broader shifts in social attitudes toward divorce, although further longitudinal observation is required to determine their long-term implications.

3.2 Attitudes towards divorce

As outlined in the Data and Methods section, we used data from the European Values Study (EVS 2017) to examine factors associated with attitudes toward divorce, resulting in a final pooled sample of 8,494 individuals after excluding observations with missing values (Tab. 1). Greece accounts for 41.5% of the sample, followed by Italy (24.6%), France (20.6%), and Spain (13.3%). The Greek subsample is younger on average, reflecting the sampling procedure, which began with university students. Accordingly, the median age in the Greek sample is 33, compared with 52 in Italy, 49 in Spain, and 53 in France.

Women constitute 55.8% of the total sample, with higher female representation in Greece (60%), Spain (55.4%), and France (54.4%), while the Italian subsample displays a more balanced gender composition. With respect to marital status, 39.2% of respondents are single, 42.1% married, 9.5% divorced or separated, 7.1% widowed, and 2.1% report being in a registered partnership (RP). The Greek subsample contains a higher share of single respondents (54%), a pattern consistent with its younger age structure, whereas the proportion of singles ranges between 27% and 31% in the other countries. Widowed individuals are more

Table 1 **Variables describing the four samples (Greece, Italy, Spain, France)**

	Greece	Italy	Spain	France	Total
Sample Size					
Sample Size	3,527 (41.5%)	2,092 (24.6%)	1,129 (13.3%)	1,746 (20.6%)	8,494 (100.0%)
Median Age					
Median Age	33	52	49	53	46
Sex					
Females	2,116 (60.0%)	1,045 (50.0%)	626 (55.4%)	949 (54.4%)	4,736 (55.8%)
Males	1,411 (40.0%)	1,047 (50.0%)	503 (44.6%)	797 (45.6%)	3,758 (44.2%)
Marital Status					
Single	1,906 (54.0%)	580 (27.7%)	299 (26.5%)	545 (31.2%)	3,330 (39.2%)
In Registered Partnership	40 (1.1%)	0 (0.0%)	51 (4.5%)	87 (5.0%)	178 (2.1%)
Married	1,261 (35.8%)	1,179 (56.4%)	516 (45.7%)	621 (35.6%)	3,577 (42.1%)
Widowed	117 (3.3%)	191 (9.1%)	110 (9.7%)	183 (10.5%)	601 (7.1%)
Separated	39 (1.1%)	74 (3.5%)	63 (5.6%)	81 (4.6%)	257 (3.0%)
Divorced	164 (4.6%)	68 (3.3%)	90 (8.0%)	229 (13.1%)	551 (6.5%)
Is marriage an outdated institution?					
Yes	905 (25.7%)	680 (32.5%)	408 (36.1%)	631 (36.1%)	2,624 (30.9%)
No	2,622 (74.3%)	1,412 (67.5%)	721 (63.9%)	1,115 (63.9%)	5,870 (69.1%)
Is religion important for your life?					
Yes	2,372 (67.3%)	1,401 (67.0%)	453 (40.1%)	627 (35.9%)	4,853 (57.1%)
No	1,155 (32.7%)	691 (33.0%)	676 (59.9%)	1,119 (64.1%)	3,641 (42.9%)
I justify divorce					
Always	901 (25.5%)	489 (23.4%)	371 (32.9%)	637 (36.5%)	2,398 (28.2%)
Sometimes/Never	2,626 (74.5%)	1,603 (76.6%)	758 (67.1%)	1,109 (63.5%)	6,096 (71.8%)

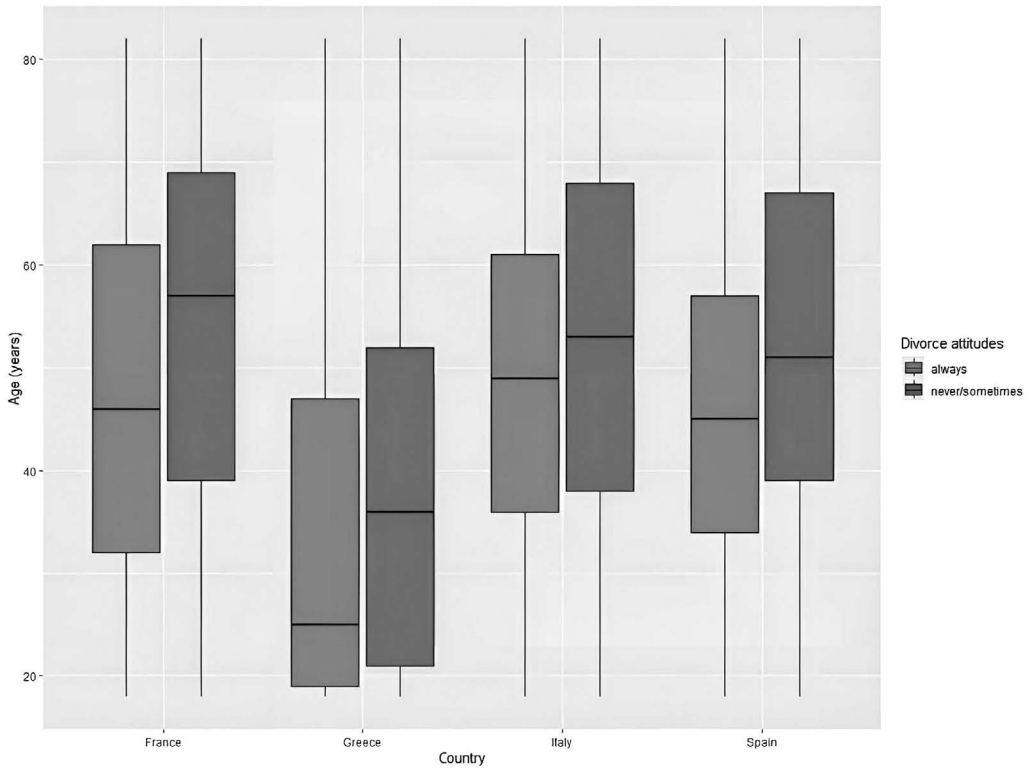
Source: EVS/WVS, 2021; EVS, 2020; own calculations.

prevalent in the Italian, Spanish, and French subsamples (approximately 10%) than in Greece.

The prevalence of registered partnerships varies across countries in line with differences in legal frameworks and timing of institutional recognition. No respondents in Italy report an RP, reflecting the relatively recent introduction of this legal status in 2016. Registered partnerships remain uncommon in Greece, where legal recognition was introduced in 2008 for heterosexual couples and extended to same-sex couples in 2015, while higher proportions are observed in Spain and France. Divorce and separation are more frequently reported in the Spanish and French subsamples, a pattern consistent with the higher divorce intensity observed at the population level in these countries.

Regarding values, attitudes toward marriage and religion display marked cross-national variation. In Greece and Italy, a smaller proportion of respondents consider marriage to be an outdated institution (25.7% and 32.5%, respectively), compared with Spain and

Figure 8 **Divorce attitudes versus age (Greece, Italy, Spain, France)**



Source: EVS/WVS, 2021; EVS, 2020; Plot created using ggplot2 package in R by Wickham (2016).

France (both 36.1%). Similarly, religion is reported as important by a larger share of respondents in Greece and Italy (approximately 67%) than in Spain (40.1%) or France (35.9%). These patterns are consistent with previous findings highlighting stronger attachment to traditional family norms and higher religiosity in Southern European contexts (*Kontogiannis,* 2024a; *Kontogiannis et al.,* 2025; *Rontos et al.,* 2025).

As noted in Section 2, the analysis focuses on the most positive attitudes toward divorce, defined as respondents who report that they always justify divorce. Accordingly, the binary dependent variable equals 1 for respondents who always justify divorce and 0 for those who report that they never or sometimes justify it. The proportion of respondents who always justify divorce is lowest in Greece (25.5%) and Italy (23.4%), while higher shares are observed in Spain (32.9%) and France (36.5%). These cross-national differences in attitudes correspond to the demographic patterns described earlier, indicating that countries with higher divorce intensity also display higher levels of unconditional acceptance of divorce.

A central objective of this study is to identify socio-demographic factors associated with these highly positive attitudes toward divorce. Age is examined first. As shown in Figure 8, younger individuals report more supportive views on divorce across all countries, while unconditional acceptance declines with increasing age. We next examined the association between sex and attitudes toward divorce. Tables 2a and 2b report chi-square tests examining the association between unconditional acceptance of divorce and each categorical explanatory variable. Regarding sex differences, statistically significant associations are observed pri-

Table 2a **Variables associated with more positive attitudes towards divorce in the four samples (Greece, Italy, Spain, France)**

	Do you justify divorce?	Greece	Italy	Spain	France	Total
			Sex			
Females	*Always*	604 (28.5%)	232 (22.2%)	216 (34.5%)	354 (37.3%)	1,406 (29.7%)
	Never/Sometimes	1,512 (71.5%)	813 (77.8%)	410 (65.5%)	595 (62.7%)	3,33 (70.3%)
Males	*Always*	297 (21.0%)	257 (24.5%)	155 (30.8%)	283 (35.5%)	992 (26.4%)
	Never/Sometimes	1,114 (79.0%)	790 (75.5%)	348 (69.2%)	514 (64.5%)	2,766 (73.6%)
			Marital Status			
Divorced or Separated	*Always*	52 (25.6%)	63 (44.4%)	60 (39.2%)	118 (38.1%)	293 (36.3%)
	Never/Sometimes	151 (74.4%)	79 (55.6%)	93 (60.8%)	192 (61.9%)	515 (63.7%)
Single or in a RP	*Always*	610 (31.3%)	158 (27.2%)	165 (47.1%)	287 (45.4%)	1,22 (34.8%)
	Never/Sometimes	1,336 (68.7%)	422 (72.8%)	185 (52.9%)	345 (54.6%)	2,288 (65.2%)
Married or Widowed	*Always*	239 (17.3%)	268 (19.6%)	146 (23.3%)	232 (28.9%)	885 (21.2%)
	Never/Sometimes	1,139 (82.7%)	1,102 (80.4%)	480 (76.7%)	572 (71.1%)	3,293 (78.8%)
			Is marriage an outdated institution?			
Yes	*Always*	357 (39.4%)	205 (30.1%)	168 (41.2%)	253 (40.1%)	983 (37.5%)
	Never/ ometimes	548 (60.6%)	475 (69.9%)	240 (58.8%)	378 (59.9%)	1,641 (62.5%)
No	*Always*	544 (20.7%)	284 (20.1%)	203 (28.2%)	384 (34.4%)	1,415 (24.1%)
	Never/Sometimes	2,078 (79.3%)	1,128 (79.9%)	518 (71.8%)	731 (65.6%)	4,455 (75.9%)
			Is religion important for your life?			
Yes	*Always*	411 (17.3%)	233 (16.6%)	95 (21.0%)	160 (25.5%)	899 (18.5%)
	Never/Sometimes	1,961 (82.7%)	1,168 (83.4%)	358 (79.0%)	467 (74.5%)	3,954 (81.5%)
No	*Always*	490 (42.4%)	256 (37.0%)	276 (40.8%)	477 (42.6%)	1,499 (41.2%)
	Never/Sometimes	665 (57.6%)	435 (63.0%)	400 (59.2%)	642 (57.4%)	2,142 (58.8%)

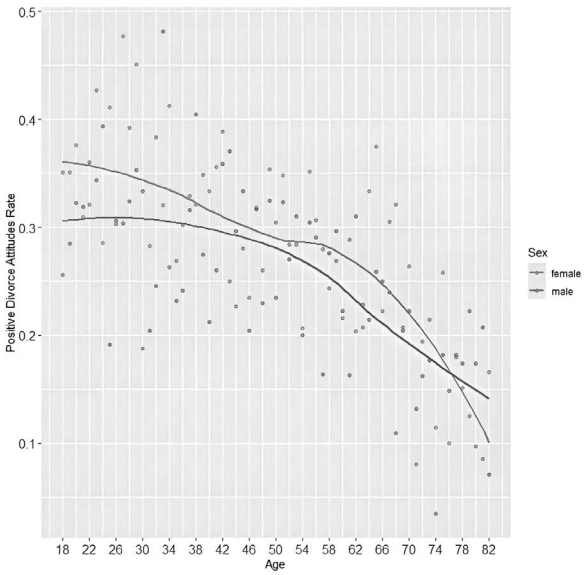
Source: EVS/WVS, 2021; EVS, 2020; own calculations.

Table 2b **Variables associated with more positive attitudes towards divorce in the four samples (Greece, Italy, Spain, France)**

		Greece	Italy	Spain	France	Total
Sex	Chi-square value	24.613	1.478	1.558	0.527	11.035
	Degrees of freedom	1**	1	1	1	1**
Marital Status	Chi-square value	83.163	50.895	60.973	42.260	202.380
	Degrees of freedom	2**	2**	2**	2**	2**
Is marriage an outdated institution?	Chi-square value	122.710	25.241	19.438	5.321	159.000
	Degrees of freedom	1**	1**	1**	1*	1**
Is religion important for your life?	Chi-square value	255.920	106.56	47.580	50.024	525.400
	Degrees of freedom	1**	1**	1**	1**	1**

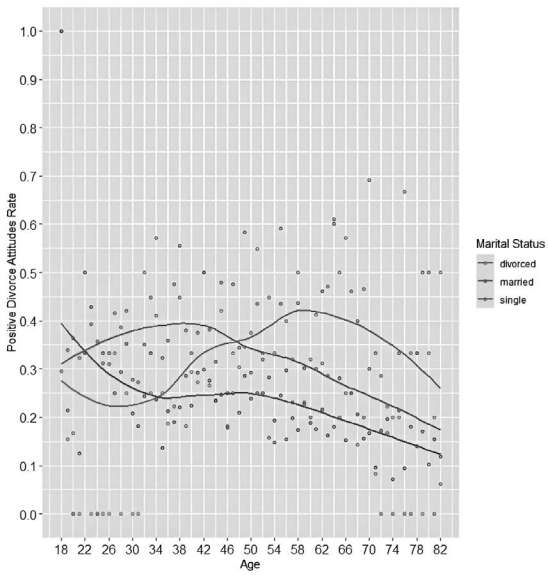
*P< 0.05, **P < 0.01.
Source: EVS/WVS, 2021; EVS, 2020; own calculations.

Figure 9 **Positive Divorce Attitudes Rate versus age and sex (Greece, Italy, Spain, France)**



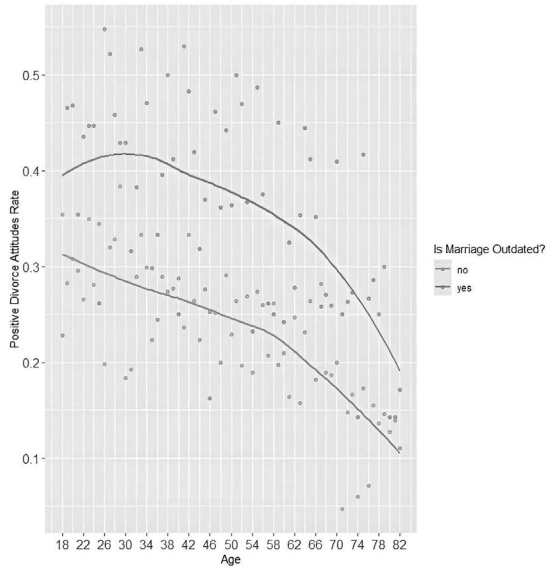
Note: Positive Divorce Attitudes Rate is the proportion of respondents who always justify divorce (value 10 on the EVS 2017 scale), calculated for each age (single year) and subgroup displayed in the figure.
Source: EVS/WVS, 2021; EVS, 2020; Plot created using ggplot2 package in R by Wickham (2016).

Figure 10 **Positive Divorce Attitudes Rate versus age and marital status (Greece, Italy, Spain, France)**



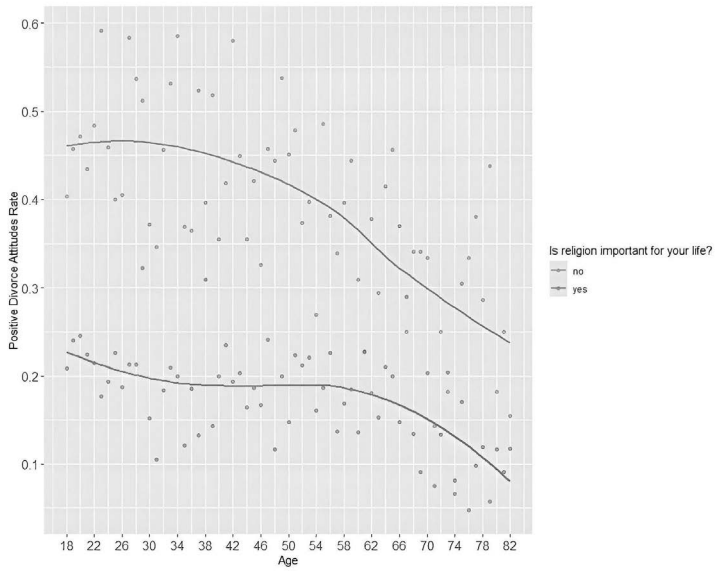
Note: Positive Divorce Attitudes Rate is the proportion of respondents who always justify divorce (value 10 on the EVS 2017 scale), calculated for each age (single year) and subgroup displayed in the figure.
Source: EVS/WVS, 2021; EVS, 2020; Plot created using ggplot2 package in R by Wickham (2016).

Figure 11 **Positive Divorce Attitudes Rate versus age and 'Is marriage outdated?' (Greece, Italy, Spain, France)**



Note: Positive Divorce Attitudes Rate is the proportion of respondents who always justify divorce (value 10 on the EVS 2017 scale), calculated for each age (single year) and subgroup displayed in the figure.
Source: EVS/WVS, 2021; EVS, 2020; Plot created using ggplot2 package in R by Wickham (2016).

Figure 12 **Positive Divorce Attitudes Rate versus age and 'Is religion important for your life?' (Greece, Italy, Spain, France)**



Note: Positive Divorce Attitudes Rate is the proportion of respondents who always justify divorce (value 10 on the EVS 2017 scale), calculated for each age (single year) and subgroup displayed in the figure.
Source: EVS/WVS, 2021; EVS, 2020; Plot created using ggplot2 package in R by Wickham (2016).

marily in Greece and in the pooled sample, a pattern that is likely influenced by the larger sample size of the Greek subsample.

Across countries, women generally report more positive attitudes toward divorce than men, with Italy representing an exception where differences are minimal. The largest gender gap is observed in Greece. Figure 9 further illustrates that, overall, women tend to express higher levels of support for divorce than men across most age groups, with the exception of the oldest respondents (aged 75 and above).

Marital status is also associated with attitudes toward divorce (Tables 2a and 2b). Divorced or separated respondents and single individuals (including those in a registered partnership) tend to report more positive attitudes than married or widowed respondents. In most countries, singles display higher levels of unconditional acceptance of divorce than divorced individuals, with Italy again representing an exception to this pattern.

Figure 10 shows that singles consistently report more positive attitudes toward divorce than married respondents across nearly all age groups. Among divorced individuals, attitudes vary by age: younger divorced respondents tend to express less unconditional acceptance of divorce, whereas older divorced individuals (aged 50 and above) display comparatively higher levels of support. While a detailed interpretation of these age-specific patterns lies beyond the scope of the present analysis, they point to heterogeneity within marital status groups that warrants further investigation.

Attitudes toward the institution of marriage are also associated with divorce attitudes. Respondents who consider marriage to be an outdated institution tend to express higher levels of unconditional acceptance of divorce (Tables 2a and 2b). In France, however, this association is weaker than in the other countries: support for divorce is high regardless of views on marriage, with 40.1% of respondents who view marriage as outdated and 34.4% of those who do not report always justifying divorce (chi-square test, $p = 0.021$). Notably, respondents in France who do not regard marriage as outdated display more positive attitudes toward divorce than respondents in Italy who do consider marriage outdated (34.4% versus 30.1%). Figure 11 confirms that across age groups, respondents holding more critical views of marriage tend to express higher levels of support for divorce.

Finally, religiosity remains associated with attitudes toward divorce. Across all countries, respondents for whom religion is not considered important report more favorable attitudes toward divorce (Tables 2a and 2b). Figure 12 illustrates that this difference is particularly pronounced among younger respondents (under the age of 50).

3.2.1 Modelling attitudes towards divorce via Binary Logistic Regression

Model selection was guided by standard information criteria and diagnostic checks. Comparisons based on the Akaike and Bayesian Information Criteria indicated that the fully specified binary logistic regression model–including age, sex, country, religiosity, marital status, and attitudes toward marriage–provided an appropriate balance between parsimony and explanatory capacity. More specifically, the transition from the intercept-only to the fully specified model results in substantial reductions in AIC (10,112.04 to 9,372.746) and BIC (10,119.09 to 9,443.217), indicating a markedly better model fit. Diagnostic checks further indicated no problematic multicollinearity among the explanatory variables and a satisfactory overall model fit, based on commonly used procedures in regression analysis (*Fox et al.,* 2021; *Fox – Monette,* 1992; *Hosmer – Lemeshow,* 1980; *Melo,* 2013). Detailed diagnostic results are available upon request.

We estimated a binary logistic regression model in which the dependent variable equals 1 if respondents report *always* justifying divorce (value 10 on the EVS 2017 scale) and 0 if they report that they *sometimes* or *never* justify divorce (values 1–9). The model estimates the probability that an individual reports unconditional acceptance of divorce as a function of age, sex, country, religiosity, marital status, and attitudes toward the institution of marriage. All reported coefficients and odds ratios represent adjusted associations, controlling for all other variables included in the model. The results of the final model are presented in Table 3.

Results from the final model (Table 3) indicate that age is negatively associated with unconditional acceptance of divorce: with each additional year of age, the odds of always justifying divorce decrease

Table 3 **Results of the binary logistic regression model**

Variables	Estimate	Standard Error	Exp(Estimate)
Intercept	0.370**	0.117	1.448
Age-18	–0.011**	0.002	0.989
Sex (male vs female {ref.})	–0.278**	0.052	0.757
Country (Greece vs France {ref.})	–0.422**	0.073	0.656
Country (Italy vs France {ref.})	–0.303**	0.077	0.739
Country (Spain vs France {ref.})	–0.133	0.084	0.876
Is religion important for your life? ('Yes' vs 'No' {ref.})	–0.925**	0.055	0.397
Is marriage an outdated institution? ('Yes' vs 'No' {ref.})	+0.300**	0.055	1.350
Marital Status (married or widowed vs divorced or separated {ref.})	–0.426**	0.087	0.653
Marital status (Single or in a RP vs divorced or separated {ref.})	–0.147	0.094	0.863

*P< 0.05, **P < 0.01
Source: EVS/WVS, 2021; EVS, 2020; own calculations.

by approximately 1.1%. Men display lower odds of unconditional acceptance compared with women. Relative to France, respondents in Greece and Italy exhibit lower odds of always justifying divorce, while differences between Spain and France are not statistically significant. Religiosity is strongly associated with divorce attitudes, with individuals for whom religion is important displaying substantially lower odds of unconditional acceptance. This association should be interpreted in light of the outcome definition: religious respondents may consider divorce acceptable under specific circumstances without endorsing it unconditionally, and may therefore select high but non-extreme values on the response scale (e.g. 8 or 9) rather than the maximum value of 10. In contrast, respondents who regard marriage as an outdated institution exhibit higher odds of unconditional acceptance. Finally, married or widowed respondents show lower odds compared with divorced or separated individuals, while single respondents and those in a registered partnership do not differ significantly from divorced or separated individuals.

3.2.2 The Application of ML Algorithms in the study of attitudes towards divorce

While binary logistic regression enables statistical inference and the estimation of adjusted associations between covariates and attitudes toward divorce, it relies on parametric assumptions and linear effects on the log-odds scale. To complement this inferential approach, we also apply several machine learning (ML) algorithms. The purpose of the ML analysis is not causal inference, but predictive assessment and the identification of variables that contribute most strongly to the classification of highly positive attitudes toward divorce, while allowing for potential non-linearities and complex interactions among predictors. Accordingly, ML results are interpreted in terms of predictive contribution and feature importance rather than marginal effects or causal relationships, and should therefore be viewed as complementary to–rather than a replacement for–the regression-based findings.

Table 4 reports the classification performance of the machine learning models considered. Overall, all models display a comparable ability to distinguish respondents who always justify divorce from those who do not. Among them, the Random Forest model performs particularly well when recall and F1-score are prioritised, indicating a strong capacity to correctly identify respondents with highly positive attitudes toward divorce in the presence of class imbalance. Boosting-based models (XGBoost and CatBoost) achieve slightly higher overall discrimination as measured by the area under the ROC curve, suggesting similar predictive performance across alternative modelling strategies.

Table 4 **Evaluation Metrics of Machine Learning Algorithms on Attitudes Toward Divorce**

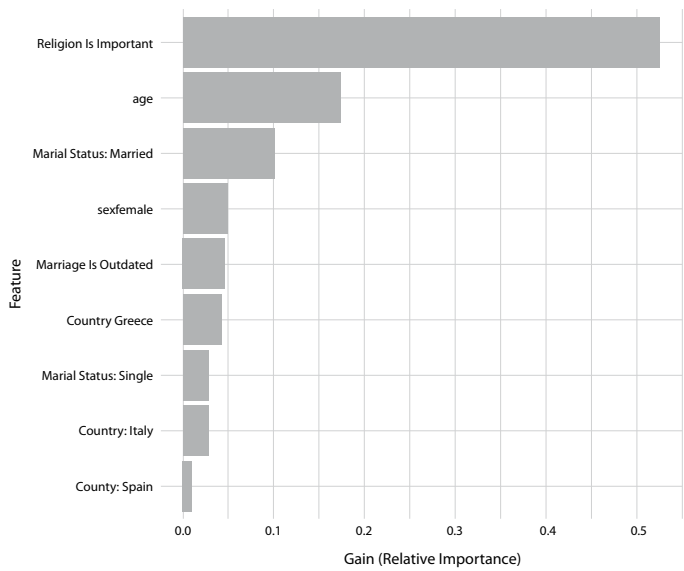
Model	Optimal Threshold	Accuracy	Precision	Recall	F1	AUC	AUPRC
XG Boost	0.51	0.666	0.446	0.651	0.529	0.707	0.468
Random Forest	0.38	0.653	0.607	0.870	0.715	0.734	0.714
Light GBM (with SMOTE)	0.42	0.627	0.414	0.712	0.524	0.699	0.470
Cat Boost (with SMOTE)	0.48	0.663	0.444	0.663	0.532	0.709	0.479

Source: EVS/WVS, 2021; EVS, 2020; own calculations.

To explore the variables most relevant for model predictions, we examine feature importance measures derived from the two best-performing models–Random Forests and XGBoost (Fig. 13 and 14). Differences in the presentation of feature importance across these figures reflect methodological differences between the algorithms. In the XGBoost model, categorical predictors are represented through one-hot encoding, and feature importance is therefore displayed at the level of individual categories. In contrast, the Random Forest model reports importance aggregated at the level of the original variables. Although the magnitude of importance scores is not directly comparable across models, both approaches consistently identify age and the perceived importance of religion as the most influential predictors, with clearly higher importance than the remaining covariates.

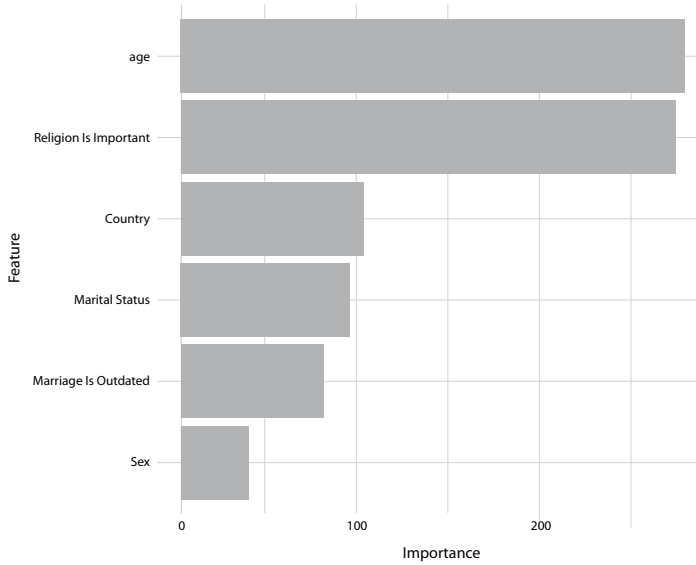
Additional insight into the direction and contribution of specific features is provided by the SHAP summary plot based on the XGBoost model (Fig. 15). Positive SHAP values indicate an increased predicted likelihood of always justifying divorce, while negative values indicate a reduced likelihood rela-

Figure 13 **XGBoost Top Feature Importance (Gain) plot**



Source: EVS/WVS, 2021; EVS, 2020; own calculations.

Figure 14 **Random Forests' Top Feature Importance plot**



Source: EVS/WVS, 2021; EVS, 2020; own calculations.

tive to the model's baseline prediction. The results show that respondents for whom religion is not considered important are associated with positive SHAP values, whereas those who report religion as important tend to exhibit negative SHAP values, corresponding to a lower predicted probability of unconditional acceptance of divorce. A similar pattern is observed for marital status, with married or widowed respondents displaying predominantly negative SHAP values. Higher age values are also mainly associated with negative SHAP values, indicating a reduced predicted likelihood of always justifying divorce among older respondents. Importantly, negative SHAP values do not imply a negative outcome in absolute terms, but rather a downward contribution of specific feature values to the predicted probability.

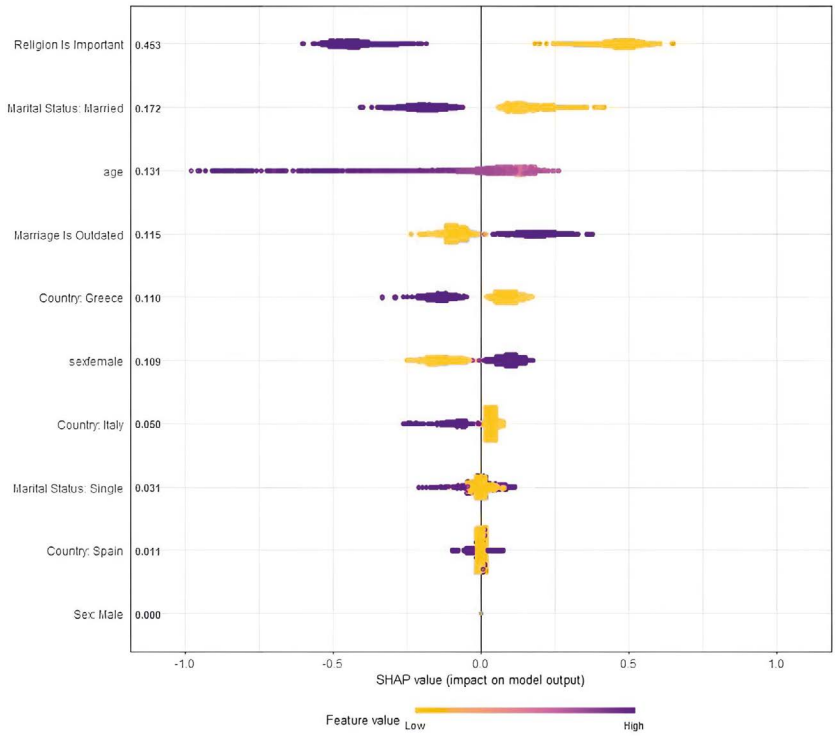
Overall, the machine learning results are consistent with the regression findings. Across both analytical frameworks, age and religiosity emerge as the most influential dimensions associated with attitudes toward divorce. This convergence suggests that these factors are not only statistically relevant in regression models, but also play a central role in shaping the predictive structure of divorce attitudes across countries.

4. DISCUSSION

This study set out to examine both divorce behaviour and attitudes toward divorce in Greece, Italy, Spain, and France, combining long-term demographic indicators with individual-level attitudinal data. The results reveal both shared patterns and persistent cross-national differences, highlighting the importance of institutional context, religiosity, and cohort dynamics in shaping divorce trends and normative orientations.

From a behavioural perspective, divorce intensity increased in all four countries over the past decades, although at different points in time and with varying magnitude. France and Spain reached comparatively higher levels earlier, whereas Greece and Italy exhibited lower divorce intensity and later increases. These differences do not suggest a simple linear process of uniform transformation, but rather reflect context-specific configurations shaped by legal frameworks,

Figure 15 **SHAP (SHapley Additive exPlanations) values plot [XGBoost]**



Source: EVS/WVS, 2021; EVS, 2020; own calculations.

religious influence, and broader social change. Legislative reforms facilitated divorce in each country, yet the evidence indicates that institutional change interacted with underlying value shifts and demographic dynamics rather than acting as an isolated trigger.

The cohort analysis further demonstrates that more recent marriage cohorts tend to experience higher divorce risks and shorter marital durations before dissolution. This pattern is consistent with broader transformations in marital expectations, where individual fulfilment and personal autonomy increasingly shape the evaluation of intimate relationships. However, the pace and extent of these changes differ across countries, reinforcing the importance of institutional and cultural context.

The attitudinal analysis provides complementary insight. Younger respondents and women are more likely to express unconditional acceptance of divorce, consistent with prior research linking generational change and gender equality to more permissive family norms. Religiosity emerges as a central factor across analytical approaches, with higher levels of religious commitment associated with lower unconditional acceptance of divorce. Importantly, the outcome variable captures unconditional endorsement ("always acceptable"), rather than general permissiveness. Religious respondents may regard divorce as acceptable under specific circumstances without endorsing it unconditionally, which helps explain why high but non-extreme responses may coexist with lower selection of the most permissive category. The negative association should therefore be interpreted as reflecting resistance to absolute endorsement rather than categorical rejection of divorce.

The convergence between regression and machine learning results strengthens confidence in these as-

sociations. Across modelling strategies, age and religiosity consistently appear as the most influential predictors, suggesting that these dimensions structure divorce attitudes across diverse institutional settings. The inclusion of machine learning techniques does not alter the substantive conclusions, but highlights the robustness of key relationships and allows the identification of potential non-linearities and interaction patterns.

The comparison between Southern Europe and France illustrates how divorce behaviour and attitudes are embedded within distinct institutional and cultural configurations. France's earlier secularisation and long-standing divorce legislation coincide with comparatively higher divorce intensity and more permissive attitudes. However, the findings do not support the idea of a single, uniform pathway of family change. Instead, the results suggest that similar macro-level transformations–such as increased individualisation and shifting gender roles–are mediated by country-specific institutional arrangements and religious contexts.

Theoretically, these findings contribute to debates surrounding the Second Demographic Transition and the deinstitutionalisation of marriage. While value change is often conceptualised as a driver of behavioural transformation, the evidence here supports a more reciprocal interpretation. Divorce prevalence and normative acceptance appear to reinforce one another over time, but within boundaries shaped by institutional structures and cultural traditions. Rather than a simple unidirectional process, the relationship between behaviour and attitudes emerges as context-dependent and mutually reinforcing.

Beyond academic debate, the findings have broader social implications. Rising divorce intensity and increasing acceptance of marital dissolution affect family structures, intergenerational relationships, and social policy considerations. Countries characterised by stronger religious influence or later legal liberalisation may experience distinct policy challenges compared to contexts where divorce has been institutionalised for decades. Understanding these differences is essential for designing family policies that respond to diverse demographic realities.

Several limitations should be acknowledged. The attitudinal analysis relies on cross-sectional data, limiting causal inference regarding the directionality between behaviour and values. Additionally, sampling differences–particularly in the Greek survey–require cautious interpretation of country-specific estimates. Future research could extend the comparative framework to additional European contexts or examine other dimensions of family change, such as cohabitation or fertility behaviour, in conjunction with divorce.

Overall, the combined analysis of behavioural and attitudinal dimensions demonstrates that divorce change in Europe cannot be understood through a single explanatory lens. Institutional context, religiosity, generational change, and gender dynamics interact in shaping both marital instability and normative orientations. By focusing on Southern Europe and contrasting it with a historically distinct European case, this study provides a structured assessment of how demographic and value transformations unfold within differentiated institutional environments.

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SUMMARY

The paper examines long-term divorce patterns and attitudes toward divorce in Greece, Italy, Spain, and France, combining demographic period and cohort analysis with attitudinal analysis from the European Values Study (EVS 2017). France is used as a contrast case rather than as a benchmark because of its earlier secularisation and longer history of divorce legislation. The study applies descriptive statistics, chi-square tests, binary logistic regression, and machine learning models including Random Forest, XGBoost, LightGBM, and CatBoost.
The analysis shows that divorce spread earlier and reached higher levels in France and, later, in Spain, while Greece and Italy followed a slower and more delayed path. After 2010, divorce affected roughly half of marriages in Spain and nearly half in France, compared with clearly lower levels in Greece and Italy. Cohort analysis further indicates that more recent marriage cohorts tend to divorce more often and at earlier durations than older cohorts, although the pace of change differs across countries.
A central finding is that legal reforms strongly shaped divorce trends, although they interacted with broader cultural and demographic changes rather than acting alone. Major reforms in Greece, Spain, and Italy progressively lowered institutional barriers and contributed to rising divorce levels, while France's longer legal tradition of divorce helped explain its earlier and higher prevalence.
The attitudinal analysis shows that unconditional acceptance of divorce is more common in France and Spain than in Greece and Italy. Younger individuals, women, divorced or single respondents, and those who consider marriage an outdated institution are more likely to justify divorce, whereas religiosity is consistently associated with lower acceptance.
Logistic regression and machine learning models confirm these patterns. Age and religiosity emerge as the strongest predictors of attitudes toward divorce, while marital status and beliefs about marriage also play an important role. Overall, the findings highlight the interaction between institutional change and broader cultural transformations, showing that although Southern European countries share some common trends, their divorce trajectories remain shaped by distinct legal, religious, and social contexts.

LUNG CANCER MORTALITY AND TOBACCO CONTROL POLICIES IN CZECHIA, HUNGARY AND POLAND: A RETROSPECTIVE ANALYSIS

Vitalie Stirba – Ivana Kulhánová

Abstract

Long-term tobacco smoking is the primary cause of lung cancer and a major driver of avoidable mortality, but the impact of tobacco control policies on lung cancer mortality is delayed by 20–30 years. This study analyses period and cohort trends in lung cancer mortality in Czechia, Hungary, and Poland from 1950 to 2023 within the framework of the smoking epidemic model and the timing of tobacco control policy implementation. Using harmonised cause-of-death data and population exposure, we examined age-standardised mortality rates and cohort patterns. The results revealed pronounced cohort effects and substantial gender disparities, with earlier peaks and subsequent declines in male mortality, particularly in Czechia, and later increases or plateauing trends among females across all three countries, consistent with different stages of the smoking epidemic model. Comparisons with policy timing indicate that lung cancer mortality declines cannot be attributed to tobacco control policies alone. In Czechia, declines began before major interventions, while in Hungary and Poland, they occurred later but still reflect earlier smoking histories. Lung cancer mortality trends are thus primarily driven by cohort-specific smoking histories, with policies influencing outcomes indirectly over the long term.

Keywords: tobacco control policies, lung cancer mortality, cohort mortality, Central Europe
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INTRODUCTION

Lung cancer is one of the leading causes of cancer mortality worldwide, accounting for 18.7% of all cancer deaths in 2022, mainly due to long-term tobacco smoking, particularly in males (*Bray et al.*, 2024). Due to typically late diagnosis and low survival rates, lung cancer mortality closely reflects cumulative smoking exposure at the population level (*de Groot – Wu – Carter – Munden*, 2018). Although other risk factors such as air pollution, second-hand smoke, and genetic predisposition contribute to disease incidence, lung cancer remains largely preventable through reductions in tobacco consumption and effective public health interventions (*OECD/Eurostat*, 2022).

The relationship between smoking behaviour and lung cancer mortality has been evident since the 1950s (*DeVita – Rosenberg*, 2012), when a significant proportion of the population was exposed to smoking, and is best understood within the framework of the smoking epidemic model (*Lopez – Collishaw – Piha*, 1994). This model describes the evolution of smoking and its health consequences across successive stages, characterised by rising and subsequently declining smoking prevalence, followed after a substantial delay by corresponding changes in mortality. A key feature

of this framework is the long lag, typically 20 to 30 years (*Islami – Torre – Jemal*, 2015), between smoking exposure and lung cancer mortality, implying that observed mortality trends reflect historical smoking patterns rather than contemporaneous behaviours.

Tobacco control policies, including taxation, advertising bans, and smoke-free legislation, have been widely implemented to reduce smoking prevalence (*Pierce – White – Emery*, 2012). Public health campaigns highlighting the risk of lung cancer from long-term smoking have also contributed to declining social acceptance of smoking (*Cummings – Proctor*, 2014). Among these measures, increases in tobacco taxation are particularly effective in reducing smoking prevalence, especially among adolescents and young adults who are more sensitive to price changes (*Chaloupka*, 1999). These policies primarily affect smoking initiation and early cessation, shaping the smoking behaviour of younger cohorts. However, their impact on lung cancer mortality is indirect and substantially delayed due to the long latency period between smoking exposure and lung cancer mortality (*Thun – Peto – Boreham – Lopez*, 2012). As a result, changes in lung cancer mortality trends can rarely be attributed to specific policy interventions, since observed declines typically reflect behavioural shifts that occurred decades earlier (*González-Marrón et al.*, 2019). Consistent with this lag structure, modelling studies project a continued decline in lung cancer mortality in Europe, with persistent gender differences reflecting the later uptake of smoking among females (*Gredner – Mons – Niedermaier – Brenner – Soerjomataram*, 2021). Cross-national differences in lung cancer mortality may therefore reflect not only variation in policy timing and intensity, but also differences in the stage of the smoking epidemic and cohort-specific smoking histories.

Trends in cigarette smoking prevalence and lung cancer mortality reveal pronounced differentiations between males and females and between countries with the same level of development (*Barta – Powell – Wisnivesky*, 2019). Countries in Central and Eastern Europe provide a particularly relevant context for examining these dynamics. During the socialist period, smoking prevalence was high and tobacco consumption was subject to limited regulation, followed by the introduction of comprehensive tobacco control policies during the post-1990 transition (*Neuberger*, 2019). This study focuses on Czechia, Hungary, and Poland, which share similar historical and socioeconomic backgrounds but exhibit different trajectories in lung cancer mortality, particularly in the timing of peak mortality among males and subsequent declines. These differences allow for a comparative assessment of how cohort-specific smoking patterns and the timing of policy implementation are reflected in long-term mortality trends.

The aim of this study is to analyse period and cohort trends in lung cancer mortality in Czechia, Hungary, and Poland from 1950 to 2023, and to interpret these trends within the smoking epidemic framework and the timing of tobacco control policies. By combining age-standardised mortality rates with cohort-based analysis, the study seeks to clarify the extent to which observed lung cancer mortality changes are driven by long-term smoking behaviours versus policy interventions, while explicitly accounting for the lag between exposure and outcomes.

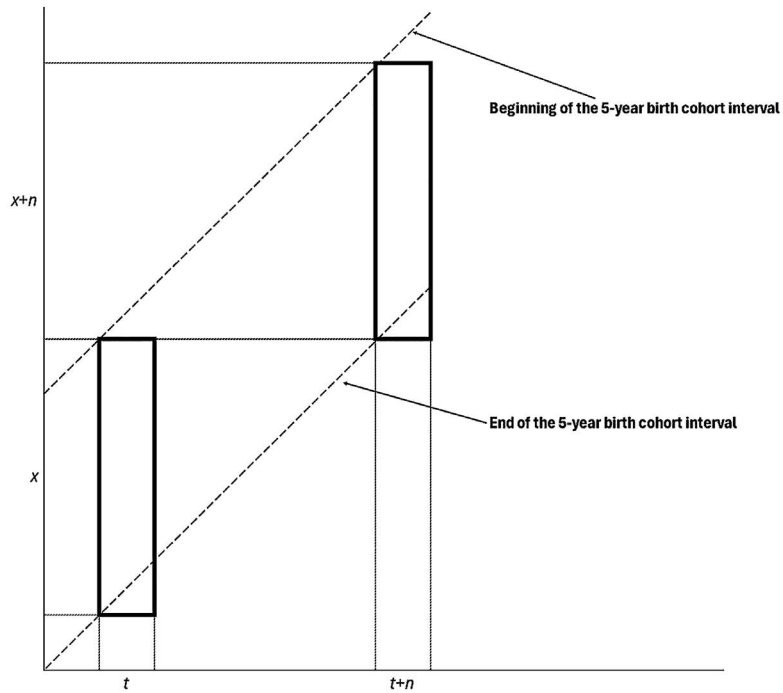
DATA AND METHODS

The research is designed to assess the changes in trachea, bronchus and lung cancer mortality (further referred to as lung cancer) in Czechia, Hungary, and Poland in alignment with previously implemented tobacco control policies and legislation amendments on tobacco products. For this, the perspectives of changes in mortality patterns within calendar years (a period approach) and changes in death rates between 5-year birth cohorts were considered. Thus, to assess the period changes in the age profile of lung cancer, age-specific mortality and age-standardised death rates were calculated. Mortality standardisation was performed by using the direct method, for which the New European Standard Population of the 2013 revision was used (*Eurostat*, 2013). To establish cohort changes in age- and cause-specific mortality, period data were transformed into cohort data. Figure 1 presents the algorithm for transforming the period age- cause-specific mortality rates into a cause-specific mortality vector according to the following formula:

$$m_c = \{m_{x,t+x}\}_{x=0}^{\omega};$$

where: m_c is the mortality vector for the birth cohort c, x is the age, from 0, to the last open age group ω, $m_{x,t+x}$ is the mortality rate at age x within the period $t+x$.

Figure 1 **The algorithm for transforming the period mortality rates into a cohort mortality vector displayed on a Lexis diagram**



Source: Authors' processing.

Data on death counts for lung cancer were retrieved from the World Health Organisation Mortality Database (*WHO Mortality Database*) and the Czech Statistical Office (*Český statistický úřad*), and harmonised according to the International Classification of Diseases (ICD) of the 7th–9th (code 162), and 10th (codes C33–C34) editions. Data on population by age and sex were obtained from the Human Mortality Database (*Human Mortality Database*).

Information on tobacco control policies and laws for tobacco products by country was collected from the national legislative framework. The tobacco control policies implemented in Czechia, Hungary and Poland are presented in Appendix.

RESULTS

Figure 2 presents age-standardised death rates (SDR) for trachea, bronchus, and lung cancer from the 1950s to the most recent available years. Lung cancer mortality increased markedly across all analysed countries from the 1950s onwards, reaching peak levels among males in Czechia in 1980 and in Hungary and Poland in 1998.

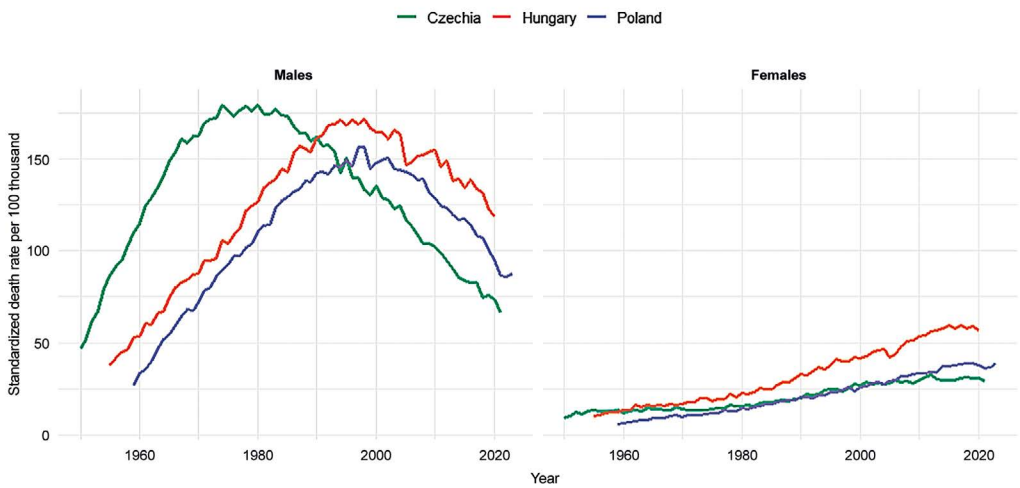
In Czechia, male lung cancer mortality rose rapidly from 47 deaths per 100 thousand in 1950 to a peak of 179 deaths per 100 thousand in 1980, followed by a sustained decline to approximately 65 deaths per 100 thousand in recent years. Hungary and Poland exhibited a similar pattern, with male lung cancer mortality peaking at 171 and 156 deaths per 100 thousand, respectively, in 1998, after which rates gradually decreased.

Among females, lung cancer mortality remained substantially lower throughout the period. Increases were more gradual and continued into recent decades, reaching peak values of 32 deaths per 100 thousand in Czechia in 2012, and 40 and 60 deaths

per 100 thousand in Poland and Hungary, respectively, in 2019.

Overall, female lung cancer mortality rose during a period when male mortality was already declining, reflecting a pronounced temporal lag between sexes. The steepest increase in female lung cancer mortality occurred from the 1980s onwards, with peak or near-peak levels observed approximately 15–20 years later than in males across all analysed countries.

Figure 2 **Age-standardised mortality rate for trachea, bronchus and lung cancer for Czechia, Hungary, and Poland for 1950–2023, by sex**



Source: The WHO Mortality Database, the Human Cause-of-Death Data series.

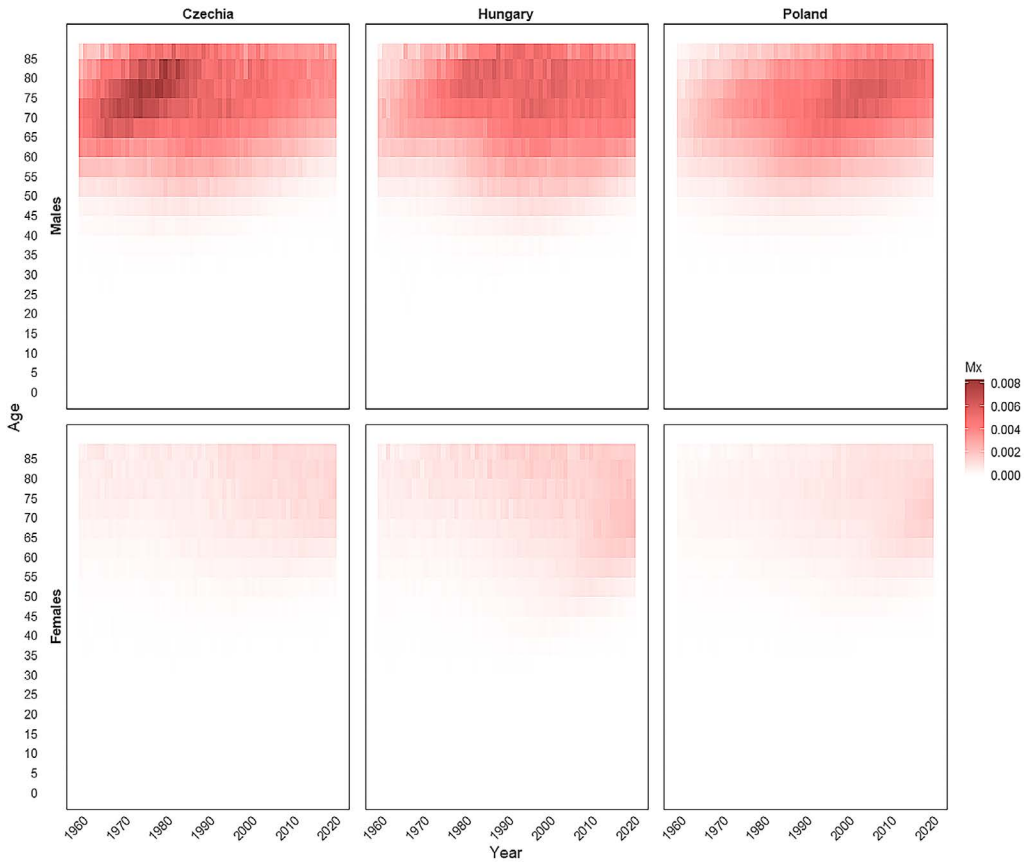
Figure 3 displays annual age-specific lung cancer mortality rates by 5-year age groups from the 1950s to the most recent years, highlighting that lung cancer mortality was concentrated in older age groups. This reflects the long latency period associated with cumulative smoking exposure. Variation in age-specific rates over time also reflects cohort effects, distinguishing generations with different levels of smoking intensity and contributing to the overall trends observed in Figure 2.

In males, the initial increase in lung cancer mortality from the 1950s was driven primarily by rising death rates among those aged 55–74, while younger and older age groups initially showed lower levels. In subsequent decades, the increase shifted towards older ages, alongside declining rates in the 50–69 age group. From the 1980s onwards, high lung cancer mortality rates were observed across all age groups above 50. This pattern persisted until the early 2000s in Czechia and Poland, and until the 2010s in Hungary, after which declines became evident, particularly among those aged 40–59.

In females, age-specific lung cancer mortality rates were consistently lower than in males across all countries and throughout the study period. From the 1950s to the late 1980s, relatively higher female lung cancer mortality was concentrated in older age groups, while rates among females under 70 remained low. From the 1990s onwards, however, all three countries experienced increasing female lung cancer mortality in age groups above 50, reflecting a broader upward shift across older age categories.

Figure 4 presents age-specific lung cancer mortality rates reorganised from period-based data into cohort mortality profiles. Across all three countries, male cohorts born before the early 1900s showed low lung cancer mortality even at older ages, while successive cohorts exhibited substantially higher

Figure 3 **Annual changes in age-specific mortality rates for 5-year age groups for trachea, bronchus, and lung cancer, displayed on Lexis diagram in Czechia, Hungary, and Poland, for males and females, 1960–2020**



Source: The WHO Mortality Database, the Human Cause-of-Death Data series.

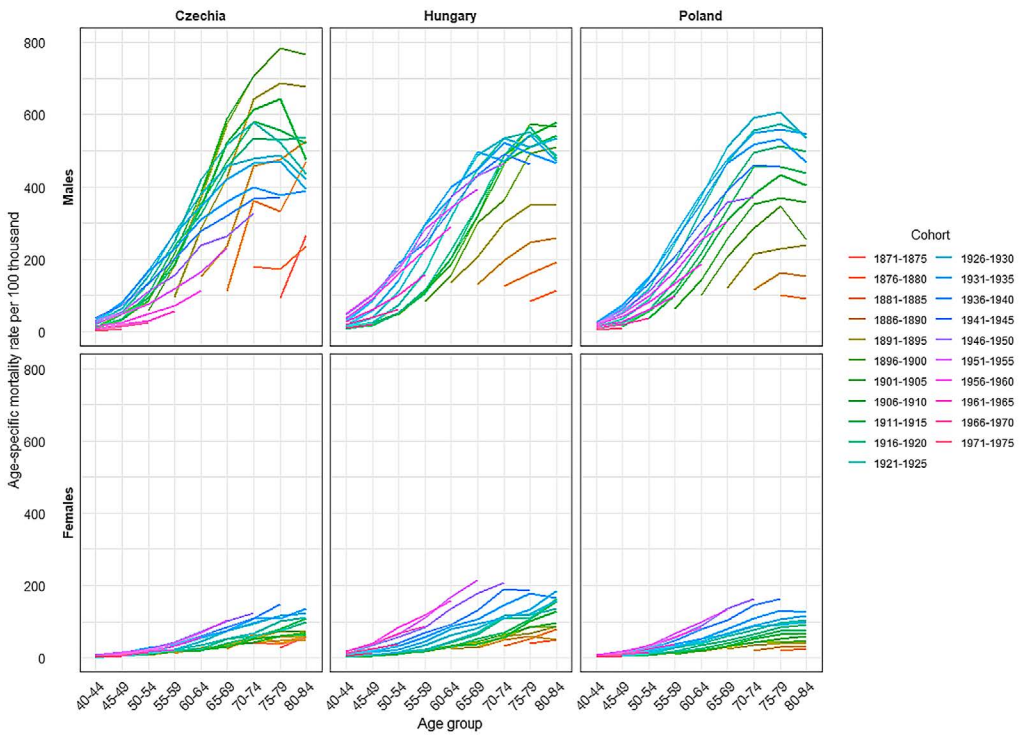
mortality levels. Among Czech males, the highest lung cancer mortality was observed in cohorts born between 1896 and 1925, followed by a marked decline in subsequent generations, particularly among those born after 1950. In contrast, Hungarian and Polish males showed a more continuous cohort increase, with each successive cohort experiencing higher lung cancer mortality than the previous one. A reversal of this trend was only evident in the youngest cohorts born after the 1950s.

Age patterns were broadly similar across countries. In all cases, lung cancer mortality was negligible before age 35, rose sharply thereafter, and peaked in the 70–74 age group, followed by a slight decline at older ages. In Czechia, cohorts born between 1886 and 1940 displayed relatively similar mortality levels between ages 35 and 64, while the largest cohort differences emerged at ages 65 and above. In Hungary and Poland, cohorts born between 1926 and 1950 showed substantially higher lung cancer mortality between ages 35 and 69 compared with earlier cohorts born between 1896 and 1915.

Compared with males, female lung cancer mortality remained considerably lower across all countries, cohorts and age groups. Nevertheless, successive female cohorts showed increasing lung cancer mortality, with a gradual shift towards younger ages, although this pattern appeared to improve in more recent genera-

tions. Intercohort differences were less pronounced in Czech females than in Hungary and Poland.
The highest female lung cancer mortality was observed among cohorts born in the 1930s and 1940s. Cohorts born after the 1950s continued to exhibit increasing lung cancer mortality, especially at ages above 60, although this upward trend appeared to stabilise in the youngest generations.

Figure 4 **Age-specific mortality rates for trachea, bronchus, and lung cancer displayed for 5-year cohorts in Czechia, Hungary, and Poland, for males and females**



Source: The WHO Mortality Database, the Human Cause-of-Death Data series.

DISCUSSIONS

This study analyses long-term trends in lung cancer mortality in Czechia, Hungary, and Poland using period and cohort perspectives, in relation to the timing of implemented tobacco control policies. The results revealed a pronounced cohort effect, with substantial gender differences, highlighting the historical evolution of smoking behaviour across generations. Although smoking prevalence was high during the socialist era in all these countries, lung cancer trajectories diverged, particularly in the timing of increases and declines among males and in the continuing increase and recent plateauing among females.
Observed trends in lung cancer mortality can be interpreted within the framework of the smoking epidemic, which conceptualises tobacco-attributable mortality as progressing through four stages (*Lopez – Collishaw – Piha*, 1994). In Czechia, males experienced an earlier transition to the later stages of the epidemic than in Hungary and Poland, with a peak in lung cancer mortality in the late 1970s to early 1980s, followed by a subsequent decline. In contrast, Hungarian and

Polish men experienced an increase in lung cancer mortality until the mid-1990s, emphasising a delayed progression through the stages of the smoking epidemic. Females in all three countries exhibited patterns consistent with earlier stages of the smoking epidemic, characterised by later smoking uptake and a continuous increase in lung cancer mortality. These gender differences were also evident in the cohort mortality analysis, where the male cohorts with the highest lung cancer mortality rate were born in the early 20th century, while the most affected female cohorts were born several decades later.

The findings emphasised the substantial lag between smoking exposure, tobacco control policies implementation, and lung cancer mortality outcomes. Given the long latency of smoking-attributable lung cancer, current mortality trends largely reflect the smoking behaviour from previous decades (*Thun – Peto – Boreham – Lopez*, 2012). This time lag challenges the direct attribution of lung cancer mortality declines to specific tobacco control policies. For instance, lung cancer mortality among Czech males began to decline prior to the widespread introduction of tobacco control measures, possibly reflecting earlier shifts in social norms and lifestyle (*Hoek – Edwards – Waa*, 2022) that reduced smoking prevalence among young and middle-aged males in the mid 1980s (*Škodová et al.*, 2000; *Spilková – Dzúrová – Pikhart*, 2011). In Hungary and Poland, the later decline in lung cancer mortality among males aligns more closely with implemented tobacco control policies, even though a decline in smoking prevalence was observed decades before (*Zatoński et al.*, 2017).

In addition to tobacco control policies, other factors have likely contributed to shaping lung cancer mortality trends (*Mackenbach – Karanikolos – McKee*, 2013). Changes in cigarette composition, particularly reductions in tar and nicotine levels, and the introduction of cigarette filters, may have reduced the disease risk, although their overall impact remains debated (*Lee*, 2001; *Lee – Sanders*, 2004). Broader socioeconomic transformations that affected smoking behaviour and access to healthcare, as well as the improvements in lung cancer diagnosis and treatment, have also likely contributed to reduced risk exposure and improved survival.

An important consideration in interpreting these results is the role of competing risks among smokers. High cardiovascular mortality in earlier decades may have reduced the number of individuals surviving to older ages, where lung cancer is more prevalent. Accounting for this dynamic is therefore important when assessing long-term trends in lung cancer mortality.

This study has several limitations. First, the analysis relied on aggregated cause-specific mortality data, which limited the ability to establish a causal relationship between tobacco control policies and lung cancer mortality outcomes. Second, despite efforts to harmonise causes of death across the ICD revisions, some inconsistencies may persist. Third, cohort mortality vectors were estimated by aligning annual age-specific mortality rates with corresponding 5-year birth cohorts, which does not fully account for within-period variation. Finally, the study does not explicitly quantify the impact of competing risks on lung cancer mortality.

CONCLUSIONS

The results of this study suggest that lung cancer mortality trends in Czechia, Hungary, and Poland are primarily driven by a cohort-specific history of tobacco smoking, where tobacco control policies have important but indirect and delayed influence. The observed decline in lung cancer mortality reflects the previous reductions in smoking prevalence, particularly among males. The recent plateau in female lung cancer mortality is consistent with the later diffusion of smoking and the ongoing progression of the smoking epidemic among females. These findings underscore the importance of accounting for cohort effects and lag structures when assessing the impact of tobacco control policies, which is crucial for the development of effective public health strategies. Future lung cancer mortality trends will largely depend on smoking prevalence in the population, smoking initiation among younger generations, as well as the countries' ability to enforce effective tobacco control policies and the capacity of their healthcare systems to provide preventive measures and timely, efficient treatment.

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SUMMARY

Long-term tobacco smoking is the primary cause of lung cancer development, and prevention efforts are therefore focused mainly on tobacco control policies. In this regard, most developed countries banned smoking advertising and increased excise taxes on tobacco products, ultimately prohibiting public smoking.

This study analyses period and cohort changes in lung cancer mortality in Czechia, Hungary, and Poland, with a focus on tobacco control policies and amendments.

Data from the WHO Mortality Database on trachea, bronchus, and lung cancer were harmonised according to the International Classification of Diseases to ensure a consistent data series since the 1950s.

For decades, Czechia, Hungary, and Poland have shared a similar legislative framework for tobacco control, while recording some of the highest smoking rates and lung cancer mortality among European nations. However, since the 1990s, all three countries have enacted a series of legislative amendments to discourage cigarette smoking. The further lung cancer mortality decline, predominantly observed in males, is rather a result of the shift in the population structure, as the generations that experienced a smaller tobacco smoking exposure were ageing, while the cohorts of heavy smokers still registered high death rates. In females, a decline in smoking prevalence began later than in males, with lung cancer mortality rates plateauing only in recent years.

APPENDIX

Table 1. **Chronology of implemented tobacco control legislation in Czechia, Hungary, and Poland**

Czechia
Act No. 40/1995 on the Regulation of Advertising and on Amendments to Act No. 468/1991, on radio and television broadcasting.
Act No. 379/2005 on measures to protect against damage caused by tobacco products, alcohol and other addictive substances, and amending related laws.
Decree No. 344/2003 on Stipulating the requirements for tobacco products.
Act No. 132/2010 on on-demand Audiovisual Media Services.
1 Jun 2012 Rectification of the WHO Framework Convention on Tobacco Control.
Decree No. 261/2016 on tobacco products.
Decree No. 37/2017 on Electronic Cigarettes, Refill Containers, and Herbal Products for Smoking.
Act No. 65/2017 on Protection of Health Against the Harmful Effects of Addictive Substances.
Hungary
Act 1 of 1996 on Radio and Television Broadcasting.
Act XLII of 1999 on the Protection of Non-Smokers and Certain Regulations on the Consumption and Distribution of Tobacco Products.
Act CXXVII of 2003 on the Excise Tax and Excise Tax Rules in the Marketing of Products.
7 Apr 2004 Rectification of the WHO Framework Convention on Tobacco Control.
Act XLVIII of 2008 on the Basic Requirements of and Certain Restrictions on Commercial Advertising Activities.
Decree No. 291/2011 on the Regulation of Tobacco Products Labelling.
Act CXXXIV of 2012 on Reducing Smoking Prevalence Among Minors and on the Retail of Tobacco Products.
Government Decree No. 39/2013 on the Production, Placing on the Market and Control of Tobacco Products, on Combined Warnings, and the Detailed Provisions on the Application of Healthcare Penalties.
Government regulation No. 239/2016 on the amendment of government regulation 39/2013 on the detailed rules of production, distribution and control of tobacco products, the combined warnings and the application of health protection fines.
Government Decree No. 120/2024 of 10 June 2024 amending Government Decree No. 39/2013 of 14 February 2013 on the production, placing on the market and control of tobacco products, on combined warnings, and the detailed provisions on the application of healthcare penalties.
Poland
Act of November 9, 1995, on the protection of health against the effects of using tobacco and tobacco products.
Decree of the Minister of Health of February 24, 2004, on testing the amount of certain substances in cigarette smoke and the notifications and warnings placed on tobacco product packaging.
15 Sep 2006 Rectification of the WHO Framework Convention on Tobacco Control.
Act of May 21, 2025, on the amendment of the Act on protection of health against the consequences of the use of tobacco and tobacco products.

55. konference České demografické společnosti

Ve dnech 20. až 22. května 2026 se v Jihlavě, v prostorách Krajského úřadu Kraje Vysočina, uskutečnila 55. konference České demografické společnosti (ČDS). Této významné demografické události, konané za finanční podpory Rady vědeckých společností ČR, z. s., se zúčastnilo téměř sto odborníků. Program byl rozdělen do tří dnů, během nichž zazněly příspěvky pokrývající témata od reprodukčního chování přes zdravotní stav populace až po pokročilé modelování příčin smrti.

Předseda ČDS *Luděk Šídlo* (PřF UK) přivítal účastníky konference v Jihlavě. Hejtman Kraje Vysočina *Martin Kukla* zdůraznil využívání demografických dat při plánování kapacit škol, nemocnic a zařízení sociálních služeb. *Petra Kuncová* (ČSÚ) vyzdvihla využívání dat z ČSÚ a zmínila začátek přípravy Sčítání lidu 2031.

Odborný program zahájil *Ladislav Rabušic* (FSS MU) a v příspěvku **Bezdětnost jako nezamýšlený důsledek druhé modernizace?** otevřel diskusi o hodnotových proměnách mladých generací. Podle něj jsou současná nízká plodnost a rostoucí podíl bezdětných žen makro-strukturním důsledkem úspěšné modernizace, která upřednostňuje individuální svobodu a materiální komfort před tradičním modelem rodiny. Zároveň zdůraznil roli nejistot. Následně vystoupil *Petr Jirásek* (KÚ Kraje Vysočina; PřF UK) s příspěvkem **Demografický vývoj statutárního města Jihlavy a prognóza jeho dalšího populačního vývoje**. Věnoval se v něm komplexní analýze Jihlavy jako případové studie středně velkého města čelícího stárnutí populace a zdůraznil potřebu propojení demografických dat s plánováním veřejné infrastruktury a bydlení. Střední varianta prognózy předpokládá mírný pokles počtu obyvatel.

Zdravotní tématiku do úvodu vnesli *Benjamin Petruželka* a *Kryštof Hanzlík* (MPSV ČR) s analýzou **Zdravotní stav příjemců příspěvku na péči: analýza komorbidit na základě propojených administrativních a zdravotnických dat**. Jejich práce založená na datech projektu SZ DATA za roky 2014–2024 ukázala na silnou zdravotní zátěž příjemců příspěvků, u nichž dominují kardiovaskulární, neurologická a metabolická onemocnění, a poukázala na výraznou nadúmrtnost této skupiny během pandemických let. Metodologické aspekty migrace rozebral *Radek Havel* (ČSÚ) v příspěvku **Úskalí tvorby a hodnocení statistiky stěhování**, kde varoval před nekritickým využíváním a interpretací administrativních dat bez znalosti jejich technických specifik, zejména na regionální úrovni. Středeční dopoledne uzavřela *Michaela Němečková* (ČSÚ) prezentací **Osoby s dočasnou ochranou – stavy, toky a věková struktura,** v níž podrobně zmapovala dynamiku uprchlické populace v Česku mezi lety 2023 až 2026 na základě dat z Cizineckého informačního systému.

Odborná sekce věnovaná současné české rodině se opírala o data mezinárodního programu Generation and Gender Programme (GGP). *Anna Šťastná*, *Jitka Slabá* (PřF UK) a *Eva Beaujouan* (University of Vienna) představily srovnání **První reprodukční zkušenosti: porovnání dat výběrového šetření se zdravotnickými registry**, které odhalilo přidanou hodnotu propojení otázek na první reprodukční zkušenosti z GGP s daty Národního registru reprodukčního zdraví spravovaného ÚZIS ČR. *Darina Kmentová* (FSS MU) se v příspěvku **Partnerské dynamiky v souvislosti s prvními reprodukčními zkušenostmi** zaměřila na stabilitu vztahů a zjistila, že páry, u nichž nedošlo k početí, čelí vyššímu riziku rozpadu vztahu. *Eva Waldaufová* (PřF UK) v práci **Od záměru až k narození dítěte: role formy partnerství v reprodukčním procesu** ukázala, že ačkoliv typ vztahu ovlivňuje záměr mít dítě, na samotnou realizaci má po zohlednění jiných faktorů menší vliv. Sekci zakončila *Sylva Höhne* (RILSA) tématem **Nízkopříjmové rodiny s dětmi pohledem administrativních dat**, kde na datech o příjemcích přídavků na dítě popsala životní podmínky těchto rodin v roce 2024.

Druhý konferenční den otevřely příspěvky zaměřené na zdraví a úmrtnost. Kolektiv autorek

Michala Lustigová, *Klára Hulíková Tesárková*, *Ivana Kulhánová*, *Pavlína Netrdová* a *Dagmar Dzúrová* (PřF UK) prezentoval příspěvek **Nerovnosti ve zdraví v kontextu socioekonomického znevýhodnění**, který doložil výrazné rozdíly ve střední délce života (ve zdraví) mezi obyvateli oblastí s rozdílnou mírou chudoby. Stejný tým v další práci **Zdraví, kontext a úmrtnost: sociálně-zdravotní paradox v po-pandemickém období v Česku** zkoumal vliv sociální deprivace, sociálního postavení, využívání zdravotní péče a komorbidit na riziko úmrtí pomocí logistické regrese. Největší riziko našly autorky u osob, kde se spojuje špatný zdravotní stav, sociální znevýhodnění a slabší kontakt se zdravotním systémem. *Terezie Štyglerová* a *Markéta Šafusová* (ČSÚ) následně představily **Novou etapu sledování úrovně úmrtnosti podle vzdělání**, díky níž se po přechodu na nové zdroje dat podařilo výrazně snížit podíl nezjištěných údajů o nejvyšším dosaženém vzdělání zemřelých na pouhých 9 %. Autorky také prezentovaly úmrtnostní tabulky a standardizované úmrtnosti příčin smrti podle vzdělání. *Jaromír Běláček* a *Jan Schubert* (MENDELU) v příspěvku **Standardizace příčin smrti pro ČR v časových řadách** kriticky zhodnotili publikované evropské statistiky a poukázali na vliv volby standardu na výsledné trendy.

Posterová sekce nabídla pestrou škálu témat ve stručnější formě. *Jana Vrabcová*, *Jitka Langhamrová* a *Tomáš Fiala* (VŠE) uvedli poster **Demografie a udržitelnost ve výuce** o nové vedlejší specializaci na VŠE, a také spolu s *Terezií Frömmelovou* inovativní metodu **LEGO® Serious Play: Demografie v kostkách** pro popularizaci oboru, která umožňuje převést demografické koncepty do lépe uchopitelné podoby. *Sylvie Chrtková* a *Jiřina Kocourková* (PřF UK) analyzovaly **Plánované rodičovství v Jižní Asii** zejména z pohledu „neuspokojené potřeby" antikoncepce, zatímco *Tereza Hrabcová*, *Ivana Kulhánová* a *Michala Lustigová* (PřF UK) doložily vliv nadměrného užívání léků na mobilitu v posteru **Polyfarmacie a vznik potíží s chůzí u starších osob v Evropě**. *Adéla Volejníková* a *Jiřina Kocourková* (PřF UK) zhodnotily **Postavení Česka mezi hlavními cílovými zeměmi přeshraniční reprodukční péče v Evropě** v kontextu asistované reprodukce. *Filip Čábela*, *Tereza Patáková* a *Luděk Šídlo* (PřF UK) představili aplikaci prospektivního přístupu k měření stárnutí na zdravotnickém personálu v Česku v příspěvku **Přizpůsobení se stárnoucí pracovní síle ve zdravotnictví**. *Jana Rozmarinová* a *Simona Činčalová* (VŠPJ) zkoumaly **Využití demografických dat při plánování zdravotní péče** na základě krajských strategických dokumentů. *Marie Hanková* a *Michala Lustigová* (PřF UK) uzavřely posterovou sekci longitudinální analýzou well-beingu s názvem **Změny subjektivní psychické pohody v čase v populaci Česka**.

Odpolední program se věnoval populačnímu vývoji a projekcím. *Branislav Bleha* (PrF UK v Bratislavě) a *Boris Vaňo* (Infostat) prezentovali **Demografickou projekciu okresov Slovenska do roku 2050**, v níž predikovali nezvratné stárnutí všech slovenských regionů. *Tomáš Fiala*, *Jitka Langhamrová* a *Jana Vrabcová* (VŠE) v příspěvku **Může zahraniční migrace zajistit finanční udržitelnost důchodového systému?** vytvořili vlastní populační projekci Česka a modelovali nezbytné migrační objemy pro zachování stability systému. *Pavol Hurbánek* a *Ilona Svobodová* (ÚZEI) v navazujících příspěvcích rozebrali **Vztah velikosti obcí s vybranými charakteristikami obyvatelstva v Česku** a **Vztah velikosti obcí se zaměstnaností v zemědělství a s nezaměstnaností v Česku**.

Závěrečný konferenční den byl otevřen příspěvky na téma plodnost a aplikace. *Kryštof Zeman* (Austrian Academy of Sciences) v analýze **Zrychlení poklesu plodnosti a odložená rekuperace** popsal aktuální pokles plodnosti ve vyspělých zemích na základě dat z Human Fertility Database. Výsledky ukazují plošný pokles plodnosti a zastavení růstu plodnosti ve vyšším věku. Zahájil také diskusi na téma vymírání populace. *Magdalena Baštecká* (ČSÚ) pak v příspěvku **Změny plodnosti v Česku v letech 2015–2025** kvantifikovala vliv intenzity plodnosti, věkové struktury, zahraniční migrace a revize populačních dat na českou plodnost. Od roku 2022 byl nejvýraznější efekt intenzity plodnosti. **Ondřej Nývlt** (VŠE) prezentoval **Možnosti sociodemografických analýz prostřednictvím vizualizačních nástrojů** jako jsou ArcGIS (aplikace na plánování kapacit mateřských škol), Power BI (analýza nerovnosti ve vzdělávání) a R Shiny (průvodce světem trhu práce a vzdělávání).

Celou konferenci uzavřela odborná sekce věnovaná vícečetným příčinám smrti v Česku. *Šárka Daňková* (ÚZIS ČR) a *Bety Ukolova* (PřF UK) se ptaly **Jak vznikají data o vícečetných příčinách smrti a jaká**

je jejich kvalita? Popsaly proces sběru, zpracování a validace těchto dat. Nechyběl ani pohled na jejich kvalitu a příklady z praxe. *Bety Ukolova* (PřF UK), *Lukáš Kahoun* (PřF UK), *Magdalena Baštecká* (ČSÚ) a *Šárka Daňková* (ÚZIS ČR) představili aplikaci algoritmu RiCoDa, který na základě vztahů mezi vícečetnými příčinami smrti dokáže určit typ procesu vedoucího k úmrtí. V příspěvku **Úmrtnost podle typu procesu vedoucího ke smrti v Česku a jeho regionech** se zaměřili na aplikaci tohoto algoritmu na data v Česku v letech 2022–2023. Stejní autoři prezentovali výzkum **Přesah nerovností ve využívání zdravotní péče až do úmrtnostních statistik**, který prokázal, že kontakt s poskytovateli zdravotní péče před smrtí zásadně ovlivňuje podrobnost zápisu v Listu o prohlídce zemřelého.

Konference opět nabídla účastníkům nejen odborný, ale i doprovodný program, který letos zahrnoval např. prohlídky Horácké multifunkční arény, města či jihlavského podzemí. Nechyběl ani tradiční společenský večer. Ústní příspěvky i postery z konference jsou dostupné na webových stránkách České demografické společnosti – https://www.czechdemography.cz/akce/konference/konference-cds-2026/prispevky-z-konference/.

Redakce časopisu Demografie

Kulatý stůl České demografické společnosti: Význam demografie v životním pojištění

Dne 28. dubna 2026 uspořádala Česká demografická společnost (ČDS) ve spolupráci s Českým statistickým úřadem (ČSÚ) a Českou společností aktuárů (ČSpA) odborný Kulatý stůl na téma „Význam demografie v životním pojištění". Setkání, které se uskutečnilo v prostorách Českého statistického úřadu, mělo za cíl diskutovat o průsečících demografického výzkumu a aktuárské (pojistněmatematické) praxe, především v souvislosti se stárnutím populace, a to u příležitosti navázání neformální spolupráce mezi oběma společnostmi. Jako panelisté vystoupili Jan Šváb (předseda České společnosti aktuárů), Jan Kořistka (Generali Česká pojišťovna) a Patrik Balla (Kooperativa).

V úvodním příspěvku představil *Jan Šváb* profesi pojistného matematika (aktuára) jako disciplínu zaměřenou na finanční dopady rizik a nejistoty. Zdůraznil, že klíčovým pilířem této práce je aplikace odborného úsudku při volbě dat a modelů, což s sebou nese vysokou profesní odpovědnost a nutnost transparentnosti. V závěru poukázal na globální výzvy, které propojují demografii s pojišťovnictvím – zejména fenomén dlouhověkosti, rizika pandemií a s nimi spojenou potřebu inovací v penzijních a zdravotních systémech.

Následující části se ujal *Jan Kořistka*, který přítomné seznámil s metodologií modelování úmrtnosti v komerčních pojišťovnách. Po krátkém historickém exkurzu k počátkům úmrtnostních tabulek vysvětlil rozdíl v přístupu ke dvěma hlavním rizikům: riziku úmrtnosti a riziku dlouhověkosti a popsal využívání Lee-Carterova modelu pro projekce úmrtnosti. Upozornil, že pojišťovny musí populační data upravovat o specifické selekční faktory, neboť pojištění klienti vykazují jinou úmrtnost než celá populace. V mezinárodním srovnání pak otevřel otázku dalšího vývoje a možného zpomalování trendu zlepšování úmrtnosti v Česku.

Závěrečný expertní pohled přinesl *Patrik Balla*, který se věnoval budoucím výzvám stárnutí populace a dlouhodobé péče v ČR v horizontu do roku 2050. Aktuální demografické odhady jasně ukazují, že dramatický nárůst podílu seniorů, a zejména nejstarší věkové skupiny nad 85 let, zásadně zvýší potřebu ošetřovatelské a lůžkové péče. Z pohledu pojistné matematiky představuje nesoběstačnost zásadní pojistné riziko. V rámci diskuse zaznělo, že bude nutné řešit rostoucí finanční nároky na sociální a zdravotní systém, přičemž jedním z řešení je aktivní zapojení klientů prostřednictvím komerčních produktů kombinujících pojištění se spořením.

Závěrečná diskuse potvrdila, že akademická demografie a komerční aktuárská praxe jsou úzce propojené disciplíny. Zatímco demografové dodávají robustní populační prognózy a projekce, pojišťovny je transformují do finančních modelů pro řízení rizik.

Příspěvky z akce jsou dostupné na stránkách www.czechdemography.cz.

HV ČDS

LONG-TERM CARE DEPENDENCY IN THE CZECH REPUBLIC: A POPULATION-BASED MODEL FOR AN AGEING SOCIETY

Jiří Vopátek[1] – Vladislav Bína[1]

Abstract
Population ageing will increase long-term care needs in the Czech Republic, where support relies heavily on family caregiving and a public care allowance. Using administrative counts of allowance recipients (2014–2024, collected by Ministry of Labour and Social Affairs and available at Czech Statistical Office, 2025a) and the Czech Statistical Office population projection (2025–2080, Czech Statistical Office, 2024), we project age–sex patterns of care need with a mixed-effects multinomial logit model. For 2025–2040, prevalence rises in most groups, particularly at ages 85+.

Keywords: long-term care, informal caregiving, care dependence, social care policy, population ageing, Czech Republic, projection modelling
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INTRODUCTION

Population ageing is increasingly understood not only as a macro-demographic shift but also as the cumulative outcome of life-course processes shaping health, functional capacity, and care needs. Care dependency in later life reflects accumulated (dis)advantage, cohort-specific exposures, and institutional contexts (*Barrett – Barbee,* 2022). Differences in education, labour-market trajectories, family formation, health behaviours, and access to care compound over the life span and become most visible at older ages, when functional limitations and the need for assistance are more prevalent.

Long-term care (LTC) dependency is a downstream outcome of life-course inequalities rather than a sudden event of 'old age'. Its timing, intensity, and duration are shaped by cohort replacement, gendered trajectories, and the organisation of health and social care. Formally recognised dependency does not directly mirror underlying need. It is mediated by assessment rules, benefit design, and service availability, which in turn influence how care is distributed between families and formal provision.

The Czech Republic provides an important and interesting case. Official statistics show a sustained rise in the proportion of the population aged 65 and over (from 17.8% in 2014 to 20.7% in 2024), and projections indicate further growth to around 27.2% by 2044 and 30.7% by 2059 (*Czech Statistical Office,* 2024 and

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2025b). This trend is driven by the ageing of large cohorts born in the 1940s and 1950s, as well as strong cohorts born in the 1970s. Czech scholarship has highlighted the implications for care provision and support for families (*Wija,* 2015). Internationally, there is extensive evidence of rising LTC needs in ageing societies (*Chen et al.,* 2016; *OECD,* 2020; *Kallestrup-Lamb et al.,* 2024) and an active debate on projections, sustainability, and the long-term commitments embedded in LTC entitlements (*Börsch-Supan et al.,* 2013; *Bahnsen et al.,* 2020; *Spitzer – Shaikh,* 2022; *Hu et al.,* 2025). Moreover, both the role of modelling assumptions in LTC projections and the importance of unpaid care in shaping future care arrangements (*Comas-Herrera et al.,* 2006; *Pickard,* 2015).

Ageing directly increases the number of care allowance recipients and thus the population formally recognised as care dependent. Prevalence rises steeply with age, implying further shifts towards older ages where severe limitations are concentrated. Growing dependency also heightens labour-market pressures—both via demand for formal care workers and through stronger work–care reconciliation constraints within families, particularly for the 'sandwich generation' (*Rubin – White-Means,* 2009).

This article focuses on the care allowance regulated by Act No. 108/2006 Coll. (*Czech Republic,* 2006), a key instrument for supporting individuals with reduced self-sufficiency. In the Czech policy and academic debate, projections have often relied on applying current prevalence rates to future population counts. What remains limited is a population-based projection that systematically integrates publicly available administrative information on care allowance recipients with official demographic projections, while distinguishing outcomes by age, sex, and degree of dependency in a transparent manner.

The objectives of this article are threefold: first, to develop a population-based projection model using publicly available data from the Ministry of Labour and Social Affairs and the Czech Statistical Office to quantify expected changes in the number of care allowance recipients under projected demographic ageing; second, to provide disaggregated projections by age, sex and dependency level that can inform planning and debate about the organisation of support; and third, to assess the feasibility of using publicly available administrative and demographic data to construct a reproducible projection framework for the Czech Republic.

CARE FOR PERSONS DEPENDENT ON ASSISTANCE AND THE ROLE OF INFORMAL CAREGIVERS

Long-term care for persons dependent on the assistence in a home setting affects family life and has material consequences for caregivers' economic circumstances and labour market participation, including decisions about early retirement when disability pension eligibility is not met. These effects concern both caregivers and care recipients of working age, and many caregivers therefore combine unpaid care with paid work or self-employment.

Svobodová (2010) characterises caregiving as 'labour of love' (a term used by several authors), a moral obligation and duty that may strain family relationships yet is considered a legitimate expectation of relatives. In the event of loss of self-sufficiency, most respondents would prefer to remain in their original homes with support from their children or, where necessary, professional institutions, but future home-based care is increasingly 'threatened' by a declining number of potential family caregivers.

A substantial literature addresses informal care, especially work–care reconciliation, often using survey and qualitative evidence (*Horrell et al.,* 2014). *Lilly et al.* (2007) conclude that labour-market withdrawal is more likely among intensive caregivers, while moderate caregiving may remain compatible with employment; recent longitudinal evidence similarly documents changes in work hours, labour-market exit and related outcomes for both men and women (*Josten et al.,* 2024).

The formal–informal care distinction is central (*Bauer – Sousa-Poza,* 2015; *Francová – Novotný,* 2008; *Mikanová,* 2018; *Triantafillou et al.,* 2010). Informal care imposes implicit costs, particularly foregone earnings, and caregivers are less likely to be employed and, when employed, tend to earn less than non-caregivers (*Carmichael – Charles,* 2003; *Raiber et al.,* 2022). Long-term caregiving may also generate wage penalties via statistical discrimination and reduced perceived reliability (*Heitmueller – Inglis,* 2007), and

may disadvantage caregivers through skill loss, reduced work experience and confidence. In line with modern European concepts, residential services should primarily serve highly dependent persons (mainly levels III–IV), while levels I–II should form only a small share of institutional clients (*Zimmelová et al.,* 2010). A Czech-focused discussion is provided by *Kotrusová et al.* (2013).

For the Czech Republic, *Triantafillou et al.* (2010) report average family caregiving duration of about four to five years and that most care for older persons is provided by working-age women, many of whom are simultaneously employed; care is most often provided by adult children, followed by spouses, other relatives and friends, with women forming the majority of caregivers. *Holmerová* (2004) estimates 400,000–500,000 family caregivers. *Geissler* (2016) emphasises the long-term nature of caregiving, reporting average duration of 10.4 years (median six). Caregiving can therefore reduce labour supply among the working-age population (when not provided mainly by retirees) and generate broader social costs through reduced participation and low caregiver incomes, including reliance on allowances, pensions and other transfers. Work–care reconciliation remains difficult, and the institutional context includes recognition of caregiving in pension entitlements and care allowances; with increasing life expectancy, these issues are likely to grow in importance (*Geissler,* 2016).

Dependency concerns not only older persons but also individuals of working and pre-working age (*Ministry of Labour and Social Affairs of the Czech Republic,* 2025; *Czech Statistical Office,* 2025a). Publicly available Czech sources include the Statistical Yearbook of Labour and Social Affairs (*Ministry of Labour and Social Affairs of the Czech Republic,* 2025), the Structure of Care Allowance Recipients (*Czech Statistical Office,* 2025a) and Selected Data on Social Security (*Czech Statistical Office,* 2022). In international comparison, *Wija* (2012) reports that LTC recipients in the Czech Republic account for about 2.8% of the population (2.4% home-based, 0.4% institutional), and that roughly 70% of LTC recipients across comparable countries are supported at home, with the Czech Republic exceeding 80%.

The Ageing Report: Economic & Budgetary Projections for the EU Member States, 2022–2070 (*European Commission,* 2024) highlights increasing life expectancy at birth and at age 65 and emphasises that these trends will have substantial implications for LTC through growth in the number of older persons requiring assistance. Based on our own analysis of publicly available data (*Ministry of Labour and Social Affairs of the Czech Republic,* 2025 and Czech Statistical Office, 2025b), the share of persons dependent on care relative to the total population remained relatively stable between 2014 and 2023, despite increasing absolute numbers. Dependent persons can be grouped by care setting into formal (institutional) and informal (home/household) care. Formal care comprises residential social service facilities, while informal care is predominantly provided in private households by family or other non-professional caregivers. Using *Czech Statistical Office data* (2023) and our calculations, the share of dependent persons in formal residential facilities ranged from 18.3% (2020) to 19.9% (2022) during 2014–2023; these figures exclude persons hospitalised in LTC beds.

A comparison of 2014 with the current situation in 2024 (the most recent available period) reveals a stable yet strongly age and sexdifferentiated structure in the proportion of persons dependent on care. In the 0–17 age group, the proportion of caredependent individuals in 2014 amounted to 1.80% among men and 1.17% among women. In the current period, this share increased to approximately 2.05% among men, while reaching 1.19% among women. In the workingage population aged 18–64 years, the proportion among men rose from 1.28% in 2014 to approximately 1.40% in 2024, whereas among women it remains virtually stable (about 1.05% in 2014 compared with approximately 1.09% in the current period). In the younger oldage group aged 65–74 years, the proportion of caredependent persons in 2014 reached 4.29% among men and 4.20% among women; by 2024, it increased to approximately 4.64% among men, while slightly declining to 4.12% among women. From the age of 75 onwards, the sexspecific pattern reverses markedly: in the 75–84 age group, the proportion among men decreased from 11.24% in 2014 to 10.57% in 2024, while among women it declined from 18.98% to 15.13%. In the 85–94 age group, the proportion in 2014 amounted to 32.79% among men and 54.19% among women, decreasing by 2024 to approximately

30.30% among men and 49.13% among women. In the oldest age cohort aged 95 years and over, care dependency already reached extreme levels in 2014 (60.82% among men and 84.05% among women) and remained very high in 2024, at approximately 65.97% among men and 82.22% among women. A comparative analysis of 2014 and 2024 thus confirms that women consistently constitute the majority of caredependent persons in older age groups, reflecting their higher life expectancy and the cumulative burden of healthrelated and functional limitations at very advanced ages.

It can therefore be expected that population ageing reinforced by increasing life expectancy will raise the absolute number of dependent persons. International evidence is consistent: *Spitzer – Shaikh* (2022) discuss mechanisms linking behaviour and health misperceptions to later diagnosis and subsequent care needs; *Kallestrup-Lamb et al.* (2024) identify LTC as particularly expenditure-sensitive under rapid ageing; *Wittenberg et al.* (2018) project substantial growth in demand driven by age structure and prevalence profiles. The *European Commission* (2024) likewise concludes that growth of the 80+ population will be a principal determinant of future LTC needs, underscoring the need for reliable projections in the Czech context.

In summary, increasing life expectancy is a key determinant of care dependency, and the ageing of large cohorts implies a progressive rise in the number of persons requiring care at older ages. This process may be intensified if health status deteriorates, increasing both the intensity and scope of care needs.

DATA SOURCES AND METHODOLOGICAL APPROACH

This article draws on an analysis of publicly available administrative data on individuals dependent on LTC, aggregated for the Czech Republic. We link these data by age group, level of care dependency, and sex to demographic statistics published by the *Czech Statistical Office* (2024). We use demographic data (population stocks as of 31 December of the given year and population projections) and data on persons dependent on care (*Czech Statistical Office,* 2025a), likewise measured as of 31 December of the given year and disaggregated (in the publicly available form) into the following age groups: 0–17, 18–64, 65–74, 75–84, 85–94 and 95+ years.

We aggregate LTC allowance recipients by sex, age group, and level of dependency (Levels I–IV) for the period 2014–2024 at the regional level (regions and municipalities) and link them to the population projection released by the Czech Statistical Office in December 2024, categorised into the corresponding age groups.

Methodological limitations of the data

Our conclusions are subject to several limitations. First, the analysis relies on a specific reference date: in both data sources, observations refer to 31 December of the given year. This 'stock' approach does not capture seasonal fluctuations or within-year dynamics; however, it is consistent with the Czech Statistical Office population stocks and enables an unambiguous linkage between the two data sources.

Second, the administration of the LTC allowance agenda was technically unstable in 2012–2013, when repeated data migrations took place from the OK NOUZE information system to the new AIS system. These transitions may have resulted in inaccuracies, incompleteness, or duplicates. For this reason, we restrict the analysis to the period from 2014 onwards, when the Ministry of Labour and Social Affairs data are stably available and methodologically consistent.

Third, revisions in assigned dependency levels may affect year-to-year comparability. For example, an individual may be granted a given dependency level in December of a given year (e.g., Level II), and a backdated change may be made in January of the following year (e.g., to Level III). In such a case, the individual is recorded as Level II in the December report for the first year and appears as Level III in the December report for the subsequent year, assuming no further change or death occurs during the year. The same principle applies in the opposite direction when a higher dependency level recorded in December is subsequently reduced retroactively. Changes that occur in months other than December do not affect counts reported as of 31 December. Crucially, each care-dependent individual is counted only once in a given year.

We further note that the published Ministry of Labour and Social Affairs data include all individuals entitled

to the LTC allowance in the given month, regardless of whether care is provided in home settings or in institutional facilities. The public data, therefore, capture the total number of care-dependent persons (across formal and informal care), but they are not explicitly disaggregated into these two groups.

Within the projection horizon, we acknowledge substantial uncertainty, especially beyond 2040. Long-term projections are inherently sensitive to assumptions about demographic developments and the institutional framework (benefit levels and eligibility rules, population health, and related factors). Similar to studies assessing the fiscal consequences of ageing in the EU (*Kluge et al.,* 2019), our results depend strongly on the demographic projection, which we treat as an exogenous input: we do not explicitly propagate the uncertainty of the population projection, but we discuss it in interpreting the results. In later parts of the projection, the model also operates with relatively small population groups (e.g., the oldest age groups), for which historical information is limited. This may increase the variability of estimates in these segments and calls for caution in interpretation. It is also one of the reasons why we divide the long-term projection into two periods.

Policy changes

The care allowance was repeatedly increased in nominal terms over the period 2014–2024. The development of the allowance's nominal and real value is analysed in detail by Vopátek (2025). It documents a long-term and substantial decline in the real value of the benefit. The most pronounced drop was identified among persons aged 18+ in dependency level I over the broader period 2007–2024, driven in particular by the reduction effective 1 January 2011 from CZK 2,000 to CZK 800 per month. Our analysed period (2014–2024) lies after this episode and focuses on care-dependent persons as captured in publicly available administrative data. Nominal increases of the care allowance also occurred during 2014–2024. At the same time, the benefit levels in dependency levels III and IV were differentiated during the period by the form of care provision (informal care and institutional care), and the corresponding rate changes were not monotonic.

Methodological note: In principle, differentiating the care allowance by the form of care provision could affect the composition of allowance recipients. However, the available administrative data do not allow recipients to be disaggregated by care setting, and this mechanism therefore cannot be analysed directly. We thus work with the aggregate total number of care allowance recipients without distinguishing the setting in which care is provided. Time-series patterns in recipient counts by dependency level, age, and sex do not indicate pronounced discontinuities or structural breaks in 2014–2024 that could be unambiguously attributed to changes in benefit levels or to differentiation by care setting. Trends (especially in grades III and IV) are smooth and consistent with the long-run demographic dynamics of population ageing. Given the data limitations, we cannot demonstrate that differentiation of benefit levels by care setting had a substantial impact on the distribution of recipients across dependency levels during the study period.

Comparison with existing projection approaches in the Czech Republic: the Průša (2018) model

In his second projection, *Průša* (2018) employs two approaches, static and dynamic, based on data from 2010–2016. The static (stationary) model holds the shares of LTC allowance recipients in each age and other category at the level observed in 2016. The dynamic model reflects the empirically observed trend over 2010–2016 insofar as it uses the average annual difference from that period for projection. Under these models, the number of LTC allowance recipients is expected to increase from 348,761 persons in 2016 to 469,115 (static) and 489,142 persons in 2030 (dynamic). The largest projected increase concerns dependency levels III and IV, and already in 2018, the author warned that service capacities would be insufficient by 2030.

A comparison with observed data for 2020 shows a natural and methodologically expected deviation: *Průša* (2018) projected 368,609 (static) and 377,334 (dynamic) persons, whereas the observed number was 359,518. The difference illustrates that projection uncertainty increases with horizon length and that projections are sensitive to policy changes and

changes in population health status, which the author explicitly notes.

Our approach differs from *Průša* (2018) in both the coverage of input data (we use the longer series 2014–2024) and the modelling framework. Rather than relying on scenario-based models built on aggregated shares, we apply a multinomial mixed-effects model that models the population distribution across dependency levels as a function of age, sex, and time. The longer time series (including the period after 2016) improves the stability of trend estimation, particularly for the early years of the projection horizon.

Comparison with existing projection approaches in the Czech Republic: the National Health Information Portal (2025) model

A second relevant benchmark is the model published on the National Health Information Portal (NZIP) (2025) within the project Data support for projecting needs for social and socio-health services. NZIP uses a different indicator definition: it counts persons who received the LTC allowance for at least one month in a given year and, when classifying by dependency level, records only the highest level attained by each person during that year. Under this definition, NZIP reports 425,778 persons for 2024, whereas the Ministry of Labour and Social Affairs records 374,127 persons in levels I–IV as of 31 December 2024. The approximately 13.8% difference, therefore, primarily stems from the differing definition of the outcome (annual incidence versus a stock measured as of 31 December).

NZIP fixes age-specific prevalence at the 2024 level and assumes it remains unchanged in the future. Values are then carried forward analogously to the static model in *Průša* (2018) by re-weighting according to the projected demographic structure. This static approach does not allow endogenous trend changes in the prevalence of dependency levels to be captured (e.g., effects of improving health status or, conversely, increasing multimorbidity). Consequently, the model does not exploit the full potential of available data and does not incorporate historical trends.

By contrast, our projection model estimates time trends in the prevalence of individual dependency levels using the full 2014–2024 series and extrapolates these trends into the future. In this respect, our approach is closer to international studies that combine demographic projections with statistical models of prevalence or care use (e.g., *Comas-Herrera et al.,* 2006 for LTC in Germany, Spain, Italy, and the United Kingdom; *Hu et al.,* 2025 for microsimulation models of LTC).

CONSTRUCTION OF THE PROJECTION MODEL

Data inputs and model cells

The projection model combines administrative data on LTC allowance recipients from the Ministry of Labour and Social Affairs available from *Czech Statistical Office* (2025a) with demographic data from the *Czech Statistical Office* (2024; 2025b), including the population projection by sex and age. The basic modelling unit ('cell') is defined by the combination of calendar year, sex, age group, and dependency level. For 2014–2024, we observe realised numbers of LTC allowance recipients (by dependency level) and the corresponding population stocks as of 31 December. For projections to 2080, we use the Czech Statistical Office population projection as an exogenous input. In this framework, demography determines the size of each year–sex–age group, while the statistical model determines the distribution of individuals across dependency levels (*cf. Kluge et al.,* 2019).

We aggregate counts of care-dependent persons by year–sex–age–dependency-level cells and, in parallel, aggregate population counts into year–sex–age cells. For each year–sex–age group, we then compute a 'no care' category as the difference between the total population and the sum of persons in dependency levels I–IV (with a non-negativity check and truncation at zero to prevent negative values). As a result, for each year–sex–age group, we obtain a five-category vector (no care, levels I–IV) whose sum equals the total population of that demographic group.

Mixed-effects multinomial model

For each future year and each age × sex group in the Czech Statistical Office population projection, we first compute predicted category probabilities from the multinomial model. We then obtain projected counts by multiplying these probabilities by the projected population size of the corresponding year and age × sex group. Because all categories are modelled joint-

ly within a multinomial framework, predicted shares are bounded between 0 and 1 and sum to one in each cell, and projected counts therefore sum exactly to the projected population.

To model the five categories simultaneously (no care and dependency levels I–IV), we use a mixed-effects multinomial logistic model. Mixed-effects multinomial models are commonly used for multi-category outcomes in clustered and longitudinal settings (*Hartzel et al.,* 2001; *Hedeker,* 2003). Related approaches have also been applied to study older individuals' behaviour and health-related service use (*Spitzer – Shaikh,* 2022). For each demographic cell i (a combination of year, age group, and sex) and each care category $k \in \{$'I', 'II', 'III', 'IV', 0$\}$ (where '0' denotes the category with no provided care), we consider the random count vector $Y_i = (Y_{i,\mathrm{I}}, Y_{i,\mathrm{II}}, Y_{i,\mathrm{III}}, Y_{i,\mathrm{IV}}, Y_{i,0})$, which we model as multinomial with size N_i (the total population in the group) and probabilities $\pi_{i,\mathrm{I}},\ldots, \pi_{i,0}$):

$$Y_i \sim \mathrm{Multinom}(N_i, \pi_i), \quad \sum_k \pi_{i,k} = 1, \quad \pi_{i,k} \in (0,1),$$

We select category 'I' as the reference category. For the remaining categories, we model log-odds ratios relative to this reference using a linear mixed model:

$$\log \frac{\pi_{i,k}}{\pi_{i,I}} = \beta_{0,k} + \beta_{1,k}\mathrm{year}_i + \gamma_k(\mathrm{group}_i) + b_{0,\mathrm{group}_i,k} + b_{1,\mathrm{group}_i,k}\mathrm{year}_i,$$

where year_i is the standardised year, $\gamma_k(\mathrm{group}_i)$ captures fixed differences across age–sex groups, and $b_{0,.,k}$ and $b_{1,.,k}$ are random intercepts and random time slopes for each age × sex combination ('group'). The dependent variable is the categorical care status with five mutually exclusive outcomes in each year–age–sex cell: no care allowance and dependency grades I–IV. Fixed effects thus include a linear time trend (standardised year) and a categorical factor for groups of age × sex combination. To allow for heterogeneity in baseline levels and trends across demographic groups, we include random intercepts and random time slopes for standardized years by groups of age × sex combination. Reference categories are care dependency level I and male × 0–17 years.

We estimate the model using the *mblogit()* function from the mclogit package, which implements multinomial logit models with random effects (*Elff,* 2025). Random effects are estimated via penalised quasi-likelihood (PQL) based on a Laplace approximation to the marginal likelihood, a common approximation in generalised linear mixed models (*Breslow – Clayton,* 1993).

A key feature of this approach is that we model the entire distribution of care categories in a single multinomial mixed effects regression rather than estimating each category separately. This ensures that the projected counts in dependency levels I–IV and in the 'no care' category sum exactly to the projected population in each demographic cell, and that no category can exceed 100% of the population. This is a major advantage over alternative approaches (e.g., Poisson generalized linear mixed models with a offset), where separately modelled categories may yield projected incidences whose sum exceeds the population size.

Beyond generalised regression approaches, one can consider microsimulation and Markov models of LTC (e.g., *Hu et al.,* 2020; *Vanella et al.,* 2022), which explicitly simulate transitions between dependency states using individual longitudinal data. However, these approaches require data sources that are not available in the Czech context when relying on aggregated administrative data from the Ministry of Labour and Social Affairs of the Czech Republic and the Czech Statistical Office. We therefore opt for a statistically robust and less data-sensitive mixed-effects multinomial model.

Projecting care needs by dependency level

The model is estimated on the 2014–2024 period, for which stable administrative time series from the Ministry of Labour and Social Affairs of the Czech Republic are available (*Czech Statistical Office,* 2025a). For each future year and demographic group (age–sex) from the *Czech Statistical Office* (2024) population projection, we compute the standardised year relative to the training period; the model then yields predicted probabilities for the 'no care' category and for dependency levels I–IV.

Projected counts in each dependency level are obtained by scaling with the projected group population:

$$\hat{Y}_{i,k} = \hat{\pi}_{i,k} \cdot N_i,$$

where N_i is the projected number of persons in the relevant year–age–sex group from the demographic

data. This guarantees full consistency between the population projection and the projected number of persons in each dependency category.

RATIONALE FOR THE MODELLING CHOICE AND METHODOLOGICAL LIMITATIONS

The chosen mixed-effects multinomial model, implemented in R (*R Core Team,* 2025) using the mclogit package (*Elff,* 2025), offers several key advantages:

- *Consistency with demographic constraints:* Category shares in each cell always sum to the projected population and remain within (0,1).
- *Heterogeneity:* Random intercepts and time slopes by age–sex group allow partial pooling across groups while capturing group-specific trajectories, consistent with approaches used in international studies in the economics of ageing (*Kluge et al.,* 2019; *Spitzer – Shaikh,* 2022).
- *Replicability and transparency:* The model is implemented in open-source software with relatively simple syntax and can be replicated and extended using publicly available data.

At the same time, we acknowledge the following methodological limitations:

- PQL approximation: Penalised quasi-likelihood with a Laplace approximation may underestimate random-effects variances and the absolute magnitude of fixed effects for rare categories (*Breslow – Clayton,* 1993). In our case, we model relatively large age–sex groups at the national level, which reduces the risk of substantial bias. Nevertheless, we interpret results for extreme projection years with caution.
- *Linear time trend:* The model assumes linear changes in log-odds ratios over time (in standardised-year units). While this assumption is plausible for 2014–2024, it may idealise dynamics in the extended projection period (after 2040). For this reason, we distinguish between a baseline and an extended projection period in interpreting the results.
- *No explicit mortality or transition model:* We work with repeated cross-sections by age and sex rather than individual trajectories. Transitions between dependency levels and within-year mortality are therefore captured only implicitly through the historical data.
- *Exogenous demography and institutional stability:* The model assumes relative stability of the institutional framework governing the LTC allowance and treats the Czech Statistical Office population projection as given. Major policy reforms, changes in population health, or organisational shifts in LTC may lead to deviations between projected and realised outcomes, especially after 2040.

Despite these limitations, we consider the mixed-effects multinomial model a suitable tool for analysing and projecting the number of care-dependent persons in the Czech Republic using publicly available administrative and demographic data. It enables a detailed decomposition of future care dependency by age, sex, and dependency level while fully respecting the demographic structure of the population, thereby providing a quantitative basis for debates in social gerontology on the organisation of support, the division of care between families and services, and the planning of LTC.

Assumptions of the projection model

The projections presented in this paper are constructed under a ceteris paribus principle, assuming that other key determinants remain relatively stable over time. Similar to international projections of population ageing and LTC (*European Commission,* 2024; *Wittenberg et al.,* 2018), our approach implicitly assumes that the institutional framework does not change fundamentally over the projection horizon and that the demographic projection is an appropriate exogenous input. The core assumptions can be summarised as follows:

a) No legislative changes occur that would affect the eligibility for granting of, or withdrawal of, the LTC allowance.
b) The definition of care dependency and the assessment system for awarding the LTC allowance remain unchanged.
c) The financing arrangement of the LTC allowance system remains unchanged.
d) No major or extreme (rather, largely random) shocks occur to the population age structure.
e) Trends observed in the reference period remain valid, and demographic changes are broadly predictable.
f) The model incorporates the demographic trend of population ageing, reflected in rising numbers of

persons in selected age cohorts, in particular due to the gradual shift of large birth cohorts into older ages.

g) No major changes occur in the uptake and deployment of modern technologies (e.g., AI, robotics) that could reduce care needs among care-dependent persons.

These assumptions are central to projection robustness. Substantial violations may lead to deviations from projected values in either direction. Stability in the legislative framework and the definition of dependency are essential for the consistency of the input data. Demographic assumptions determine the baseline population structure, and technological change may, in the longer run, alter the relationship between age and dependency on assistance.

At the same time, the model (a) does not provide a deterministic 'forecast' of the future, (b) does not attempt to predict future policies or patterns of care provision, and (c) produces projections conditional on explicit assumptions about demographic trends and observed trends in the prevalence of care dependency.

DESCRIPTIVE ANALYSIS OF THE INPUT DATA

The descriptive characteristics of the input data provide a bridge between the methodological framework and the projection model. Their purpose is to make transparent the main empirical patterns that the model subsequently quantifies. In line with international work emphasising the need to link demographic trends to the prevalence of functional limitations and LTC (*Chen et al.,* 2016; *Kallestrup-Lamb et al.,* 2024), we summarise below the key findings from the 2014–2024 data.

Trends in the number of care-dependent persons (2014–2024)

According to official data from the Ministry of Labour and Social Affairs of the Czech Republic, the total number of persons dependent on care increased steadily over 2014–2024.

The distribution by dependency level indicates that the largest numbers of recipients draw the allowance in levels II and I, while the fastest growth is observed in levels IV and III, among those requiring the highest intensity of care. Higher dependency levels typically require more intensive and more costly forms of support. Recent international evidence also suggests that LTC systems can be sensitive to structural demand pressures and broader economic conditions, with implications for costs and the mix of care arrangements (*Geyer et al.,* 2025).

In 2024, within the 18–64 age group, 46,869 men were recorded compared with 35,202 women. The higher male occurrence may reflect differences in occupational exposures, risk-taking behaviour, and a higher incidence of injuries and disability. Among working-age adults, men are more often represented in the higher dependency levels (III and IV), suggesting a higher prevalence of severe limitations.

Whereas among older adults the key issue is high prevalence at very advanced ages, for younger cohorts, it is also necessary to consider the availability of rehabilitation, special education, and support services, as well as the need for gender-sensitive social policy.

Demographic trends in 2014–2024 and the projection to 2080

The 2024 Czech Statistical Office population projection constitutes the main exogenous input to our model. As in international projection studies (*European Commission,* 2024; *Hu et al.,* 2025), we assume that demography is the primary driver of the future number of persons experiencing loss of self-sufficiency. In the Czech Republic, this is particularly related to the gradual ageing of the large birth cohorts of the 1970s, which will enter age groups with the highest prevalence of care dependency in the coming decades.

Combining care-dependency data for 2014–2024 with the demographic projection through 2080 provides the core reference framework for the mixed-effects multinomial model. This enables the model to trace how future changes in population structure by age and sex translate into the prevalence of individual dependency levels.

PRESENTATION OF MODEL PARAMETERS

The full parameterisation of the mixed-effects multinomial model comprises several dozen coefficients. We

provide in text only a concise summary for interpretation. Model coefficients are presented in Appendix in Tables A.7 and A.8. The fixed effects show the expected pattern: the relative odds of higher dependency levels increase markedly with age. The time component is positive in all equations, indicating a modest strengthening over time of the relative odds of higher dependency levels compared with level I. Consistent with international evidence (*Kallestrup-Lamb et al.,* 2024), the model indicates strong age gradients in LTC needs, with a pronounced concentration of higher dependency levels at advanced ages. *Wittenberg et al.* (2018) further show that demand for care among individuals with more severe limitations is expected to grow faster than among those with milder levels of dependency, which aligns with the projections produced by our model. More specifically:

- *Strong age gradient of dependency:* for example, the coefficient for the 95+ age group (women) reaches 1.67 relative to the reference age group 0–17 in the equation for level III versus level I, corresponding to an odds ratio of approximately 5.3; for men, the corresponding coefficient is 1.25 (odds ratio ≈ 3.5).
- *Positive time trend:* for higher dependency levels (e.g., 0.126 for level III versus level I), indicating a gradual strengthening of the relative prevalence of more severe dependency over time.
- *Pronounced sex differences:* in younger age groups (0–17 and 18–64), women exhibit lower probabilities across dependency levels, whereas at advanced ages, the coefficients for women aged 95+ are systematically higher than those for men.

Both random intercepts and random time slopes indicate substantial heterogeneity across age–sex groups. Given the breadth of the parameter structure, we focus on the most important differences and trends and proceed to the projected values for individual categories, including the graphical presentation for advanced age groups (see Figure 1).

PRESENTATION OF THE MODEL'S PROJECTION OUTPUTS

The projections are presented in two temporal blocks reflecting different degrees of predictive reliability and the availability of empirical data. Tabular summaries include observed values for 2014 (the first year of the series) and 2024 (the last year for which stable administrative data are available). The model is estimated on the full 2014–2024 series and combined with the Czech Statistical Office (2024) population projection through 2080. Outputs are therefore organised into two blocks:

1. *Baseline projection period 2025–2040*
   This interval is treated as methodologically more reliable, as it closely follows empirical patterns observed over 2014–2024 and does not extend beyond roughly one generation (see Appendix, Tables A.1, A.2, and A.3).
2. *Extended projection period 2041–2080*
   Estimates beyond 2040 are presented as an informative, no-surprises outlook only and are interpreted with considerable caution. They substantially exceed the observed time series and are more sensitive to structural population changes and potential institutional reforms; accordingly, they provide an illustrative, scenario-consistent trajectory rather than a forecast (see Appendix, Tables A.4, A.5, and A.6).

The tables are organised horizontally by calendar year and vertically by dependency level (I–IV), six age groups, and sex, and they also report aggregate totals. The 'no care' category represents the residual share required to sum to 100% and is contained in tables only implicitly. For each group, we report:

- *Relative prevalence of dependency (share of the respective age cohort).*
- *Corresponding absolute numbers.*

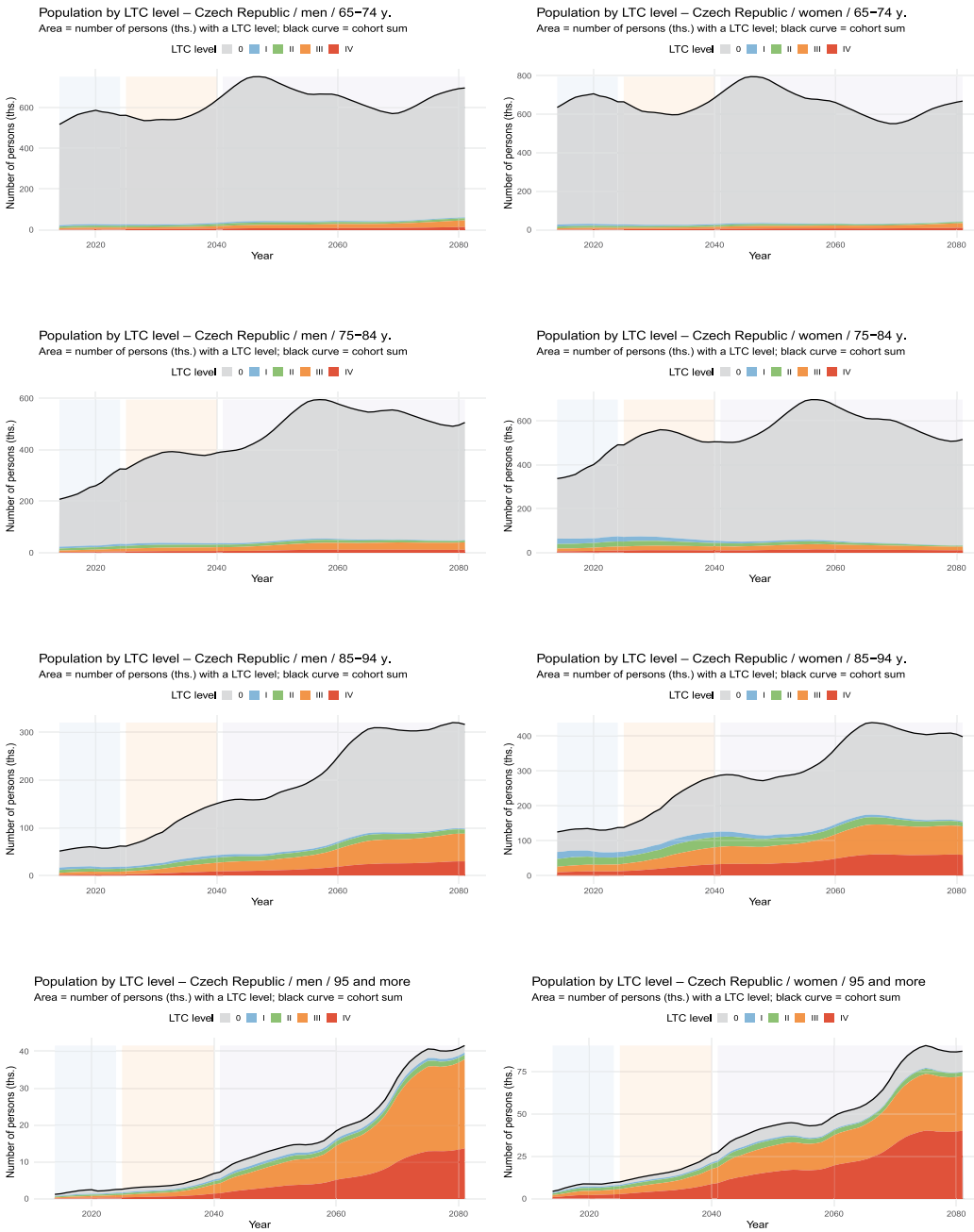
Internally consistent totals across all categories, whose sum equals the projected population in the respective cell (a property guaranteed by the construction of the multinomial model).

Men

Projections for the Czech Republic indicate a steady increase in the prevalence of care dependency among men across age groups, driven primarily by demographic ageing and cohort size. Over 2026–2040, the overall share of men dependent on care rises from 2.87% to 3.69%. The highest values are observed in the oldest age groups:

- 95+ years: an increase from 65.00% to 75.71%.
- 85–94 years: a slight decline in relative terms from 30.19% to 28.38%, alongside a substantial increase in absolute numbers.

Figure 1 **Trends in the number of LTC–dependent persons by dependency level, age group, and sex**



Note: Trends in the number of LTC–dependent persons by dependency level (different colours), age group (top to bottom), and sex (men on the left, women on the right). Background shading divides the timeline into three periods: observed data (2014–2024), the baseline projection period (2025–2040), and the extended projection period (2041–2080).

Source: Czech Statistical Office (2025a) for 2014–2024; authors' calculations for 2025–2080 (multinomial mixed-effects model output).

In the extended projection to 2080, these patterns intensify: the overall share of men dependent on care reaches 7.32%, while the 95+ group exceeds 95%, implying near-universal dependency at the very highest ages under the model assumptions. Trends for ages 65+ are shown in Figure 1; full results are reported in the Appendix (Tables A.1 and A.4).

Women

Among women, projections show the same overall direction, with particularly pronounced increases at the highest ages, consistent with higher life expectancy and the concentration of women in very old age. Over 2026–2040, the overall share of women dependent on care increases from 4.13% to 5.16%, with two notable features:

- 95+ years: persistently very high values (82% and above).
- A marked increase in prevalence above age 80 across all degrees of dependency.

By 2080, the extended projection indicates a further rise in the overall share of women dependent on care to 7.29%, while the 95+ group exceeds 86%. Trends for ages 65+ are shown in Figure 1; full results are reported in the Appendix (Tables A.2 and A.5).

Men and women combined

When both sexes are aggregated, projections again indicate rising care dependency across age groups. During 2026–2040, the total share of persons dependent on care increases from 3.51% to 4.43%, with the 95+ group reaching 78.79% to 81.00%.

In the extended projection to 2080, the overall share reaches 7.30%, while the 95+ group attains values of 88% and above. Full results are reported in the Appendix (Tables A.3 and A.6).

COMPARISON OF RESULTS WITH PREDICTIONS OF NZIP AND PRŮŠA MODELS

Tables 1 and 2 summarise selected predictions of the number of care allowance recipients from different sources varying in their methodological approach, input data, and interpretative framework. *Průša* (2018) uses a projection approach based either on the recipient structure observed in 2016 or on an extrapolation of trends observed over 2011–2016. By contrast, the *NZIP* (2025) figures represent an administrative prediction based on observed recipient data, where a 'recipient' is defined as any person who received the allowance for at least one month in a given year; when stratifying by dependency level, NZIP assigns each person the highest grade attained during the year. The results from both sources are compared with outputs of a multinomial mixed-effects model presented in this paper. This approach differs from the above sources both in the unit of observation and in how it captures changes in the distribution across dependency levels over time.

For these reasons, the reported projected values should not be interpreted as directly comparable across sources. Differences between estimates reflect not only different baseline data and time horizons, but above all distinct methodological and interpretative approaches to modelling the future number of care allowance recipients. Accordingly, the table is not intended to test agreement between estimates, but to illustrate that alternative analytical approaches and slightly different data sources from varying periods can yield materially different estimate of the future scale of demand for long-term care benefits. Further details are provided in the respective sources.

Let us denote, that the NZIP (2025) source does not report figures for 2025. For comparison, only the year 2030 is used. The figures are not fully comparable, as NZIP counts persons who received the care allowance for at least one month during the year and assigns each person the highest dependency grade attained in that year. Age breakdowns are not available. Total sums differ by one person.

DISCUSSION AND IMPLICATIONS FOR SOCIAL POLICY AND PUBLIC FINANCES

From a social gerontological perspective, the projected increase in LTC dependency reflects not only population ageing, but also the cumulative effects of socio-economic and health trajectories over the life span. Elevated care needs at older ages are shaped by accumulated disadvantage, cohort-specific exposures, and unequal access to resources across earlier stages of life. Educational attainment, employment histories, family trajectories, and health shocks interact over

Table 1 **Absolute number of care allowance recipients by age group**

Age group	Průša (2018)				NZIP (2025)	Mixed-effects mult. model (authors)	
	Projected number of care allowance recipients based on the 2016 status quo		Projected number of care allowance recipients based on trends observed in 2011–2016		Predicted number of care allowance recipients	Predicted number of care allowance recipients	
	2025	2030	2025	2030	2030	2025	2030
0–17	28,498	25,714	36,673	37,191	–	33,508	31,184
18–64	75,105	74,778	91,525	100,189	–	84,014	86,439
65–74	55,446	51,818	59,661	58,207	–	54,547	52,062
75–84	129,254	155,245	121,371	141,646	–	105,498	110,751
85+	123,281	161,561	118,831	151,909	–	96,895	125,303
Total	**411,584**	**469,116**	**428,061**	**489,142**	**493,990**	**374,462**	**405,739**

Sources: Průša (2018; Table 3, p. 56; for comparison, figures are available only for 2025 and 2030); NZIP (2025) and authors' calculations based on model output.

Table 2 **Absolute number of care allowance recipients by dependency level**

Dependency level	Průša (2018)				NZIP (2025)	Mixed-effects mult. model (authors)	
	Projected number of care allowance recipients based on the 2016 status quo		Projected number of care allowance recipients based on trends observed in 2011–2016		Predicted number of care allowance recipients	Predicted number of care allowance recipients	
	2025	2030	2025	2030	2030	2025	2030
Level I	127,827	144,937	105,300	98,675	123,478	96,136	91,035
Level II	134,322	153,510	143,949	167,705	147,395	114,726	118,417
Level III	94,814	108,563	116,800	147,236	136,309	102,576	122,402
Level IV	54,622	62,105	62,012	75,527	86,809	61,024	73,885
Total	**411,585**	**469,115**	**428,061**	**489,143**	**493,991**	**374,462**	**405,739**

Sources: Průša (2018; Table 3, p. 56; for comparison, figures are available only for 2025 and 2030); NZIP (2025) and authors' calculations based on model output.

time and influence both the timing and severity of care dependency in later life.

In this sense, the strong age gradient identified by the model can be read as more than a purely biological effect of ageing: it is consistent with long-term processes that sort individuals into different pathways of health and support. Life-course research highlights how discontinuous employment histories, prolonged caregiving responsibilities, and adverse health events during midlife increase the risk of functional limitations and care dependency in later life. The role of social networks and relationships is also crucial; observed age gradients in care dependency likely reflect cumulative exposure to health shocks and unequal access to social support (*Weiss et al.*, 2022). Gendered life trajectories may further amplify these patterns, as women are more likely to combine paid work with unpaid caregiving across the life course, and women are also concentrated at very advanced ages.

The Czech Republic illustrates how such processes operate within an institutional context characterised by strong reliance on informal care and relatively late transitions into formal LTC. This has direct relevance for debates about the organisation of support

and the future balance between families and services. The future care trajectories depend not only on demand but also on the availability of unpaid care, which may not keep pace with rising needs (*Pickard,* 2015), and inequalities in LTC use and unmet need remain central to understanding who benefits from existing arrangements (*Hu et al.,* 2025). As large cohorts born in the 1970s move into older age, these institutional and social dynamics are likely to become more important, beyond what can be inferred from agestructure alone.

From a macro-structural perspective, the model:

- Indicates a positive time trend in the relative odds of higher dependency levels compared with level I.
- However, this trend is modest relative to the pronounced shifts implied by changes in the population age structure. Consequently, demographic development remains the primary driver of growth in the absolute number of care-dependent persons.
- This pattern is consistent with the international work suggesting that LTC demand is strongly shaped by demographic structure and is less tightly linked to short-run fluctuations in health status or economic conditions (*Wittenberg et al.,* 2018; *Kallestrup-Lamb et al.,* 2024; *European Commission,* 2024). In addition, the wider international evidence base emphasises that projections are conditional on assumptions about future institutional arrangements and care-mix dynamics (*Comas-Herrera et al.,* 2006).

The identified trends indicate that the Czech Republic is likely to face substantial and sustained pressures on LTC, with increasing demands on institutional capacity as well as on both informal and formal care provision. Beyond growth in the number of care-dependent individuals, the projections point to a changing profile of care needs, with an increasing concentration in the oldest age groups. These developments are closely linked to the rising share of the population aged 65 and over and increasing life expectancy, with women's longer survival than men and persistently low fertility shaping the age–sex composition of later life. From a policy perspective, LTC services will therefore need to respond to the higher representation of women among clients, particularly in institutional care, given that older women are more likely to live alone. In some years, the model also signals a decline in the relative share of care-dependent persons, which may reflect improvements in health status on the one hand and postponement of dependency to higher ages on the other.

Based on the model results, we identify four key implications:

1. F*inancial pressure:* A substantial increase in public expenditure on LTC can be expected. This underscores the need for a long-term indexation strategy for the care allowance and for linking benefit levels to the actual costs of care provision, including the costs associated with expanding social services (both residential and community-based). Financial protection against LTC costs remains a priority, as board and lodging in institutions can be several times higher than the cost of nursing care alone (*Colombo et al.,* 2011).
2. *Capacity pressure:* Rapid growth in the number of individuals aged 85+ and 95+ will require a significant expansion of care capacity, including pronounced development of home-based services, as existing institutional care capacities will be insufficient.
3. *Labour market:* Demand for care workers is likely to grow faster than the number of economically active individuals, which is expected to decline. Staffing availability is therefore likely to become a binding constraint.
4. *Active ageing policy:* Prevention of functional decline, education and training, assistive technologies, support for community-based services, and the digitalisation of care represent key factors that may help mitigate pressure on the system, as discussed in part by *Spitzer – Shaikh* (2022) and *Geyer et al.* (2025).

From a comparative European perspective, the Czech Republic represents a typical Central European ageing trajectory, making the results relevant beyond the national context. Without adequate policy responses, there is a risk that care needs will outstrip effective access to support, particularly because the number of individuals with care needs may exceed the number of recipients of the care allowance (as not all persons in need of care apply for, or receive, the benefit).

POSSIBLE DIRECTIONS AND RECOMMENDATIONS FOR FURTHER RESEARCH

The transparency and interpretability of the projection results could be strengthened by incorporating a for-

mal scenario analysis. This would allow a systematic assessment of how projected numbers of individuals dependent on LTC respond to violations of individual model assumptions. In LTC, where projections often extend decades into the future, assessing robustness to changes in key parameters is particularly important (*Wittenberg et al.,* 2018; *European Commission,* 2024). Relevant areas include:

- *Legislative scenarios:* Changes in the definition of dependency, assessment procedures, or eligibility criteria.
- *Demographic variants:* Alternative rates of population ageing, including mortality and migration variants based on scenarios produced by the Czech Statistical Office.
- *Technological innovations:* Wider adoption of assistive technologies, telecare, robotic care, or digitalisation that may reduce the intensity of personal care needs.
- *Financial and institutional scenarios:* Alternative financing arrangements, changes in the availability of formal services, shifts in family caregiving behaviour, and adjustments in benefit levels across dependency levels (or, conversely, inadequate indexation).

The outcome should be an explicit characterisation of uncertainty and identification of parameters with the greatest influence on projections. Such sensitivity and scenario work is a standard component of LTC projection studies and supports strategic decision-making (*Wittenberg et al.*, 2018). At the same time, technological and policy uncertainty implies that distant-horizon projections are likely to be associated with wide uncertainty intervals, which is why we interpret the extended projection primarily as an informative trend under a 'no-surprises' assumption.

A comprehensive scenario analysis lies beyond the scope of the present paper, whose primary aim was to present a transparent, population-based projection model built on publicly available administrative data. It should therefore be addressed in a subsequent, dedicated study.

CONCLUSION

This paper presents a population-based projection of LTC dependency in the Czech Republic, disaggregated by age, sex, and degree of dependency, by integrating administrative data from the Ministry of Labour and Social Affairs with demographic projections from the Czech Statistical Office. To the best of our knowledge, multinomial mixed-effects modelling has not previously been applied for this purpose in Central and Eastern Europe. The results indicate that LTC needs will rise substantially in the coming decades, driven primarily by demographic ageing and growth in the older and very old population. In particular, strong increases are projected for dependency levels III and IV, which typically require the most intensive forms of care.

The projected growth in the number of care-dependent persons will place pressure on both pillars of the care system: informal care, which currently covers around 80% of needs (especially grades I and II), and formal care, where capacity is limited and primarily oriented towards grades III–IV. Informal care is also shaped by longer-term socio-demographic change, including smaller family sizes, higher geographic mobility, and increasing labour market participation of women, which reduces the pool of available family caregivers. As a result, demand for formal services is likely to grow faster than the supply of caregivers. This pattern is consistent with evidence reported for OECD countries more broadly (*OECD,* 2023; *Rocard – Llena-Nozal,* 2022; *Carrino et al.,* 2023).

From a social policy perspective, it will therefore be essential to expand formal LTC capacity while supporting the sustainability of informal care through measures that facilitate the reconciliation of work and care, strengthen caregiver training, develop community-based services, and deploy assistive technologies. Without such measures, unmet needs and declining access to care may increase. Our findings are consistent with *Průša* (2018), who argued that 'society is not prepared for the increase in the number of benefit recipients' and that strengthened financing of social services and support for home-based care will be necessary. We add that the projected growth in care dependency calls for strategic planning across social policy, healthcare, the pension system, and the labour market, with implications for the long-term sustainability of public finances.

More broadly, the findings underline that future care dependency is shaped by cumulative processes across the life span as well as by institutional arrangements that translate need into entitlement and support. The projected increase in care needs reflects not on-

ly demographic ageing, but also the long-term consequences of cohort-specific employment histories, health trajectories, and family arrangements shaped within a given institutional context. The results underline that policies aimed at LTC sustainability cannot be confined to old-age interventions alone, but should also address earlier life-course stages where inequalities in health, work, and caregiving responsibilities accumulate. Future work that links projection models to individual-level longitudinal data could provide a more detailed assessment of these mechanisms and their implications for inequality in ageing societies.

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APPENDIX

Table A1 **Baseline projection period (2025–2040, 16 years), share of men dependent on care in the total male population.**

Age group	Dependency level	Observed values		Baseline projection period (2025–2040)							
		2014	2024	2026	2028	2030	2032	2034	2036	2038	2040
0–17	I	0.71%	0.70%	0.68%	0.67%	0.66%	0.65%	0.64%	0.63%	0.63%	0.62%
	II	0.47%	0.60%	0.64%	0.67%	0.70%	0.73%	0.77%	0.80%	0.84%	0.87%
	III	0.33%	0.44%	0.46%	0.49%	0.52%	0.55%	0.59%	0.62%	0.66%	0.70%
	IV	0.29%	0.31%	0.30%	0.31%	0.31%	0.32%	0.32%	0.33%	0.33%	0.34%
	Total	**1.80%**	**2.05%**	**2.08%**	**2.14%**	**2.19%**	**2.25%**	**2.32%**	**2.38%**	**2.45%**	**2.53%**
Abs. num. of inhabitants – men		**960,722**	**1,062,349**	**1,030,908**	**976,245**	**943,168**	**916,877**	**888,526**	**858,958**	**833,562**	**812,098**
18–64	I	0.35%	0.36%	0.35%	0.35%	0.36%	0.36%	0.36%	0.36%	0.36%	0.36%
	II	0.43%	0.42%	0.42%	0.42%	0.41%	0.41%	0.41%	0.41%	0.40%	0.40%
	III	0.31%	0.37%	0.40%	0.41%	0.42%	0.44%	0.45%	0.47%	0.48%	0.50%
	IV	0.19%	0.26%	0.28%	0.30%	0.32%	0.34%	0.36%	0.38%	0.41%	0.43%
	Total	**1.28%**	**1.40%**	**1.45%**	**1.48%**	**1.51%**	**1.55%**	**1.58%**	**1.62%**	**1.66%**	**1.70%**
Abs. num. of inhabitants – men		**3,438,180**	**3,340,069**	**3,330,565**	**3,321,241**	**3,326,386**	**3,332,627**	**3,345,229**	**3,351,384**	**3,342,752**	**3,306,889**
65–74	I	1.32%	1.23%	1.17%	1.14%	1.12%	1.09%	1.07%	1.05%	1.02%	1.00%
	II	1.50%	1.55%	1.58%	1.58%	1.58%	1.58%	1.58%	1.58%	1.57%	1.57%
	III	1.02%	1.30%	1.41%	1.47%	1.54%	1.61%	1.68%	1.76%	1.84%	1.92%
	IV	0.44%	0.56%	0.60%	0.63%	0.66%	0.69%	0.72%	0.75%	0.78%	0.81%
	Total	**4.29%**	**4.64%**	**4.77%**	**4.83%**	**4.90%**	**4.97%**	**5.05%**	**5.13%**	**5.22%**	**5.31%**
Abs. num. of inhabitants – men		**517,837**	**562,084**	**553,251**	**536,407**	**540,751**	**540,990**	**544,584**	**565,705**	**595,583**	**637,140**
75–84	I	3.40%	2.59%	2.27%	2.14%	2.01%	1.89%	1.78%	1.67%	1.57%	1.47%
	II	3.81%	3.39%	3.23%	3.13%	3.03%	2.94%	2.85%	2.76%	2.68%	2.59%
	III	2.74%	3.14%	3.24%	3.31%	3.38%	3.46%	3.53%	3.61%	3.69%	3.77%
	IV	1.29%	1.45%	1.52%	1.54%	1.56%	1.59%	1.61%	1.64%	1.66%	1.69%
	Total	**11.24%**	**10.57%**	**10.25%**	**10.12%**	**9.99%**	**9.88%**	**9.77%**	**9.68%**	**9.60%**	**9.52%**
Abs. num. of inhabitants – men		**207,586**	**325,576**	**338,250**	**363,794**	**381,025**	**392,237**	**390,282**	**382,692**	**377,932**	**388,194**
85–94	I	9.80%	6.91%	6.10%	5.65%	5.23%	4.84%	4.47%	4.13%	3.81%	3.52%
	II	11.50%	9.82%	9.30%	8.96%	8.62%	8.29%	7.97%	7.66%	7.35%	7.06%
	III	7.89%	9.20%	10.03%	10.31%	10.59%	10.88%	11.16%	11.45%	11.74%	12.02%
	IV	3.60%	4.38%	4.75%	4.90%	5.05%	5.21%	5.36%	5.51%	5.67%	5.83%
	Total	**32.79%**	**30.30%**	**30.19%**	**29.83%**	**29.50%**	**29.21%**	**28.96%**	**28.75%**	**28.57%**	**28.43%**
Abs. num. of inhabitants – men		**51,438**	**62,009**	**65,564**	**73,811**	**86,241**	**100,266**	**117,985**	**130,467**	**142,176**	**150,905**
95+	I	11.60%	10.87%	9.73%	9.35%	8.97%	8.58%	8.18%	7.79%	7.39%	7.00%
	II	21.39%	17.54%	16.14%	15.63%	15.10%	14.54%	13.98%	13.40%	12.81%	12.22%
	III	18.13%	24.76%	25.50%	27.06%	28.64%	30.24%	31.84%	33.45%	35.05%	36.63%
	IV	9.71%	12.80%	13.63%	14.49%	15.37%	16.26%	17.15%	18.06%	18.96%	19.85%
	Total	**60.82%**	**65.97%**	**65.00%**	**66.53%**	**68.07%**	**69.62%**	**71.16%**	**72.69%**	**74.21%**	**75.71%**
Abs. num. of inhabitants – men		**1,164**	**2,492**	**2,801**	**3,097**	**3,251**	**3,441**	**3,797**	**4,587**	**5,503**	**6,828**
Total absolute population – men		**5,176,927**	**5,354,579**	**5,321,340**	**5,274,596**	**5,280,823**	**5,286,438**	**5,290,404**	**5,293,793**	**5,297,508**	**5,302,053**
Tot. num. of dependent on care – men of which in:		**2.40%**	**2.79%**	**2.87%**	**2.97%**	**3.09%**	**3.21%**	**3.33%**	**3.43%**	**3.55%**	**3.69%**
Level I		0.73%	0.73%	0.70%	0.70%	0.69%	0.69%	0.68%	0.67%	0.66%	0.66%
Level II		0.80%	0.87%	0.88%	0.90%	0.92%	0.93%	0.95%	0.95%	0.96%	0.98%
Level III		0.56%	0.76%	0.83%	0.89%	0.95%	1.02%	1.09%	1.16%	1.23%	1.32%
Level IV		0.31%	0.43%	0.46%	0.49%	0.53%	0.57%	0.61%	0.65%	0.69%	0.74%
Tot. num. of dependent on care – men (levels I–IV)		**124,308**	**149,592**	**152,489**	**156,833**	**163,191**	**169,460**	**175,929**	**181,546**	**187,823**	**195,479**

Source: Czech Statistical Office (2025a) for 2014–2024; authors' calculations for 2025–2040 (multinomial mixed-effects model output).

Table A2 **Baseline projection period (2025–2040, 16 years), share of women dependent on care in the total female population**

Age group	Dependency level	Observed values		Baseline projection period (2025–2040)							
		2014	2024	2026	2028	2030	2032	2034	2036	2038	2040
0–17	I	0.53%	0.48%	0.45%	0.44%	0.42%	0.41%	0.40%	0.39%	0.38%	0.37%
	II	0.26%	0.32%	0.33%	0.34%	0.36%	0.37%	0.38%	0.39%	0.40%	0.41%
	III	0.18%	0.21%	0.21%	0.21%	0.22%	0.23%	0.23%	0.24%	0.24%	0.25%
	IV	0.19%	0.18%	0.17%	0.17%	0.17%	0.17%	0.16%	0.16%	0.16%	0.16%
	Total	**1.17%**	**1.19%**	**1.16%**	**1.17%**	**1.17%**	**1.17%**	**1.17%**	**1.18%**	**1.18%**	**1.19%**
Abs. num. of inhabitants – women		**911,975**	**1,008,584**	**981,252**	**929,394**	**898,622**	**873,463**	**847,125**	**819,631**	**795,537**	**775,040**
18–64	I	0.31%	0.29%	0.29%	0.28%	0.28%	0.28%	0.27%	0.27%	0.26%	0.26%
	II	0.35%	0.33%	0.33%	0.33%	0.32%	0.32%	0.31%	0.31%	0.31%	0.30%
	III	0.24%	0.28%	0.30%	0.31%	0.31%	0.32%	0.33%	0.34%	0.35%	0.36%
	IV	0.15%	0.19%	0.20%	0.21%	0.22%	0.23%	0.25%	0.26%	0.27%	0.29%
	Total	**1.05%**	**1.09%**	**1.12%**	**1.13%**	**1.14%**	**1.15%**	**1.17%**	**1.18%**	**1.20%**	**1.21%**
Abs. num. of inhabitants – women		**3,346,992**	**3,242,611**	**3,220,927**	**3,176,285**	**3,169,043**	**3,164,546**	**3,166,553**	**3,162,557**	**3,145,120**	**3,101,896**
65–74	I	1.64%	1.31%	1.24%	1.18%	1.13%	1.07%	1.02%	0.97%	0.93%	0.88%
	II	1.40%	1.37%	1.38%	1.37%	1.35%	1.34%	1.32%	1.31%	1.30%	1.28%
	III	0.79%	0.98%	1.08%	1.13%	1.18%	1.23%	1.29%	1.35%	1.41%	1.47%
	IV	0.37%	0.46%	0.50%	0.52%	0.54%	0.56%	0.59%	0.61%	0.64%	0.66%
	Total	**4.20%**	**4.12%**	**4.20%**	**4.19%**	**4.20%**	**4.21%**	**4.22%**	**4.24%**	**4.27%**	**4.30%**
Abs. num. of inhabitants – women		**635,257**	**664,876**	**647,608**	**616,113**	**609,481**	**600,407**	**597,867**	**617,807**	**646,322**	**684,848**
75–84	I	7.35%	4.76%	4.14%	3.76%	3.41%	3.09%	2.80%	2.53%	2.29%	2.08%
	II	6.10%	4.76%	4.52%	4.28%	4.05%	3.83%	3.62%	3.42%	3.23%	3.05%
	III	3.61%	3.59%	3.63%	3.62%	3.61%	3.61%	3.60%	3.58%	3.57%	3.56%
	IV	1.92%	2.03%	2.05%	2.06%	2.07%	2.08%	2.09%	2.10%	2.11%	2.12%
	Total	**18.98%**	**15.13%**	**14.34%**	**13.72%**	**13.14%**	**12.61%**	**12.11%**	**11.64%**	**11.21%**	**10.80%**
Abs. num. of inhabitants – women		**337,627**	**491,411**	**507,248**	**537,687**	**553,007**	**558,492**	**545,455**	**523,625**	**504,857**	**505,161**
85–94	I	15.80%	11.15%	9.89%	9.11%	8.38%	7.70%	7.07%,	6.49%	5.94%	5.44%
	II	17.99%	14.87%	14.32%	13.67%	13.04%	12.43%	11.83%	11.25%	10.69%	10.14%
	III	12.63%	14.19%	15.12%	15.44%	15.75%	16.05%	16.34%	16.61%	16.87%	17.12%
	IV	7.77%	8.94%	9.72%	9.97%	10.22%	10.46%	10.70%	10.94%	11.16%	11.39%
	Total	**54.19%**	**49.13%**	**49.05%**	**48.20%**	**47.40%**	**46.65%**	**45.94%**	**45.28%**	**44.67%**	**44.09%**
Abs. num. of inhabitants – women		**125,195**	**137,862**	**145,391**	**160,045**	**182,140**	**206,866**	**235,610**	**255,660**	**273,562**	**283,831**
95+	I	12.02%	7.80%	7.15%	6.54%	5.97%	5.44%	4.95%	4.50%	4.09%	3,71%
	II	23.52%	18.60%	17.43%	16.48%	15.57%	14.68%	13.83%	13.01%	12.23%	11.48%
	III	27.17%	29.89%	30.42%	31.03%	31.61%	32.15%	32.66%	33.14%	33.59%	34.00%
	IV	21.34%	25.93%	27.39%	28.26%	29.13%	29.98%	30.81%	31.63%	32.43%	33.21%
	Total	**84.05%**	**82.22%**	**82.39%**	**82.32%**	**82.27%**	**82.25%**	**82.26%**	**82.28%**	**82.33%**	**82.40%**
Abs. num. of inhabitants – women		**4,302**	**9,577**	**10,733**	**12,375**	**13,755**	**14,929**	**16,393**	**18,848**	**21,725**	**25,977**
Total absolute population – women		**5,361,348**	**5,554,921**	**5,513,158**	**5,431,899**	**5,426,048**	**5,418,704**	**5,409,003**	**5,398,127**	**5,387,124**	**5,376,754**
Tot. num. of dependent on care – women of which in:		**3.88 %**	**4.04%**	**4.13%**	**4.30%**	**4.47%**	**4.63%**	**4.80%**	**4.92%**	**5.03%**	**5.16%**
Level I		1.32%	1.13%	1.05%	1.03%	1.00%	0.97%	0.94%	0.90%	0.85%	0.82%
Level II		1.25%	1.24%	1.24%	1.27%	1.29%	1.30%	1.31%	1.30%	1.29%	1.28%
Level III		0.82%	1.04%	1.13%	1.23%	1.33%	1.43%	1.55%	1.64%	1.74%	1.83%
Level IV		0.49%	0.64%	0.71%	0.77%	0.85%	0.92%	1.01%	1.08%	1.15%	1.23%
Tot. num. of dependent on care – women (levels I–IV)		**207,903**	**224,535**	**227,505**	**233,628**	**242,548**	**251,139**	**259,862**	**265,409**	**271,241**	**277,368**

Source: Czech Statistical Office (2025a) for 2014–2024; authors' calculations for 2025–2040 (multinomial mixed-effects model output).

Table A3 **Baseline projection period (2025–2040, 16 years), share of men and women dependent on care in the total population (male and female)**

Age group	Dependency level	Observed values		Baseline projection period (2025–2040)							
		2014	2024	2026	2028	2030	2032	2034	2036	2038	2040
0–17 y. persons dependent on care	I	0.62%	0.59%	0.57%	0.55%	0.54%	0.53%	0.52%	0.51%	0.51%	0.50%
	II	0.37%	0.46%	0.49%	0.51%	0.53%	0.55%	0.58%	0.60%	0.62%	0.65%
	III	0.26%	0.33%	0.34%	0.36%	0.37%	0.39%	0.41%	0.43%	0.46%	0.48%
	IV	0.24%	0.25%	0.24%	0.24%	0.24%	0.24%	0.24%	0.25%	0.25%	0.25%
	Total	**1.49%**	**1.63%**	**1.64%**	**1.66%**	**1.69%**	**1.72%**	**1.76%**	**1.79%**	**1.83%**	**1.87%**
Abs. num. of inhabitants – men and women		**1,872,697**	**2,070,933**	**2,012,160**	**1,905,639**	**1,841,790**	**1,790,340**	**1,735,651**	**1,678,589**	**1,629,099**	**1,587,138**
18–64 y. persons dependent on care	I	0.33%	0.33%	0.32%	0.32%	0.32%	0.32%	0.32%	0.31%	0.31%	0.31%
	II	0.39%	0.37%	0.38%	0.37%	0.37%	0.37%	0.36%	0.36%	0.36%	0.35%
	III	0.28%	0.32%	0.35%	0.36%	0.37%	0.38%	0.39%	0.41%	0.42%	0.43%
	IV	0.17%	0.22%	0.24%	0.26%	0.27%	0.29%	0.31%	0.32%	0.34%	0.36%
	Total	**1.16%**	**1.25%**	**1.29%**	**1.31%**	**1.33%**	**1.35%**	**1.38%**	**1.40%**	**1.43%**	**1.46%**
Abs. num. of inhabitants – men and women		**6,785,172**	**6,582,680**	**6,551,492**	**6,497,526**	**6,495,430**	**6,497,174**	**6,511,782**	**6,513,941**	**6,487,872**	**6,408,785**
65–74 y. persons dependent on care	I	1.50%	1.27%	1.21%	1.17%	1.12%	1.08%	1.05%	1.01%	0.97%	0.94%
	II	1.45%	1.45%	1.47%	1.47%	1.46%	1.45%	1.45%	1.44%	1.43%	1.42%
	III	0.89%	1.12%	1.23%	1.29%	1.35%	1.41%	1.48%	1.54%	1.62%	1.69%
	IV	0.40%	0.50%	0.55%	0.57%	0.60%	0.62%	0.65%	0.68%	0.70%	0.74%
	Total	**4.24%**	**4.36%**	**4.46%**	**4.49%**	**4.53%**	**4.57%**	**4.62%**	**4.67%**	**4.72%**	**4.79%**
Abs. num. of inhabitants – men and women		**1,153,094**	**1,226,960**	**1,200,859**	**1,152,520**	**1,150,232**	**1,141,397**	**1,142,450**	**1,183,512**	**1,241,905**	**1,321,988**
75–84 y. persons dependent on care	I	5.85%	3.90%	3.39%	3.10%	2.84%	2.59%	2.37%	2.17%	1.98%	1.81%
	II	5.23%	4.22%	4.00%	3.82%	3.63%	3.46%	3.30%	3.14%	2.99%	2.85%
	III	3.28%	3.41%	3.47%	3.49%	3.52%	3.54%	3.57%	3.59%	3.62%	3.65%
	IV	1.68%	1.80%	1.84%	1.85%	1.87%	1.88%	1.89%	1.91%	1.92%	1.93%
	Total	16.03%	13.32%	12.70%	12.27%	11.86%	11.48%	11.13%	10.81%	10.52%	10.25%
Abs. num. of inhabitants – men and women		**545,213**	**816,987**	**845,497**	**901,480**	**934,032**	**950,728**	**935,737**	**906,316**	**882,789**	**893,355**
85–94 y. persons dependent on care	I	14.05%	9.83%	8.71%	8.02%	7.37%	6.77%	6.20%	5.69%	5.22%	4.77%
	II	16.10%	13.30%	12.76%	12.19%	11.62%	11.08%	10.54%	10.04%	9.55%	9.07%
	III	11.25%	12.64%	13.54%	13.82%	14.09%	14.36%	14.61%	14.87%	15.12%	15.35%
	IV	6.55%	7.52%	8.17%	8.37%	8.56%	8.75%	8.92%	9.10%	9.29%	9.46%
	Total	**47.95%**	**43.29%**	**43.19%**	**42.40%**	**41.65%**	**40.96%**	**40.28%**	**39.70%**	**39.16%**	**38.66%**
Abs. num. of inhabitants – men and women		**176,633**	**199,871**	**210,955**	**233,856**	**268,382**	**307,132**	**353,595**	**386,127**	**415,738**	**434,736**
95+ y. persons dependent on care	I	11.93%	8.43%	7.68%	7.10%	6.54%	6.03%	5.56%	5.15%	4.76%	4.39%
	II	23.07%	18.38%	17.16%	16.31%	15.48%	14.66%	13.86%	13.09%	12.35%	11.63%
	III	25.25%	28.83%	29.40%	30.24%	31.04%	31.79%	32.51%	33.20%	33.88%	34.55%
	IV	18.86%	23.22%	24.54%	25.51%	26.50%	27.41%	28.24%	28.97%	29.70%	30.43%
	Total	**79.11%**	**78.86%**	**78.79%**	**79.16%**	**79.56%**	**79.89%**	**80.17%**	**80.41%**	**80.69%**	**81.00%**
Abs. num. of inhabitants – men and women		**5,466**	**12,069**	**13,534**	**15,473**	**17,006**	**18,371**	**20,191**	**23,434**	**27,228**	**32,805**
Total absolute population – men and women		**10,538,275**	**10,909,500**	**10,834,497**	**10,706,495**	**10,706,871**	**10,705,142**	**10,699,407**	**10,691,920**	**10,684,632**	**10,678,807**
Tot. num. of dependent on care – men and women of which in:		**3.15%**	**3.43%**	**3.51%**	**3.65%**	**3.79%**	**3.93%**	**4.07%**	**4.18%**	**4.30%**	**4.43%**
Level I		1.03%	0.93%	0.88%	0.87%	0.85%	0.83%	0.81%	0.78%	0.76%	0.74%
Level II		1.03%	1.06%	1.07%	1.09%	1.11%	1.12%	1.13%	1.13%	1.13%	1.13%
Level III		0.69%	0.90%	0.98%	1.06%	1.14%	1.23%	1.32%	1.40%	1.49%	1.58%
Level IV		0.40%	0.54%	0.58%	0.64%	0.69%	0.75%	0.81%	0.86%	0.92%	0.99%
Tot. num. of dependent on care – men and women (levels I–IV)		**332,211**	**374,127**	**379,994**	**390,461**	**405,739**	**420,599**	**435,791**	**446,955**	**459,065**	**472,847**

Source: Czech Statistical Office (2025a) for 2014–2024; authors' calculations for 2025–2040 (multinomial mixed-effects model output).

Table A4 **Extended projection period (2041–2080, 40 years), share of men dependent on care in the total male population**

Age group	Dependency level	Extended projection period (2041–2080)								
		2041	2045	2050	2055	2060	2065	2070	2075	2080
0–17 y. persons dependent on care	I	0.61%	0.60%	0.58%	0.56%	0.54%	0.53%	0.51%	0.49%	0.47%
	II	0.89%	0.98%	1.09%	1.21%	1.35%	1.51%	1.68%	1.87%	2.08%
	III	0.72%	0.81%	0.93%	1.08%	1.25%	1.44%	1.66%	1.91%	2.21%
	IV	0.34%	0.35%	0.36%	0.38%	0.39%	0.40%	0.42%	0.43%	0.45%
	Total	**2.57%**	**2.73%**	**2.96%**	**3.23%**	**3.53%**	**3.88%**	**4.27%**	**4.71%**	**5.21%**
Abs. num. of inhabitants – men		**807,275**	**826,480**	**852,404**	**857,239**	**830,894**	**790,046**	**758,289**	**749,577**	**758,244**
18–64 y. persons dependent on care	I	0.36%	0.36%	0.37%	0.37%	0.37%	0.38%	0.38%	0.38%	0.38%
	II	0.40%	0.39%	0.38%	0.37%	0.37%	0.36%	0.35%	0.34%	0.34%
	III	0.51%	0.55%	0.59%	0.65%	0.70%	0.77%	0.83%	0.90%	0.98%
	IV	0.45%	0.51%	0.59%	0.69%	0.81%	0.94%	1.10%	1.28%	1.49%
	Total	**1.72%**	**1.81%**	**1.94%**	**2.08%**	**2.25%**	**2.44%**	**2.66%**	**2.91%**	**3.19%**
Abs. num. of inhabitants – men		**3,282,281**	**3,163,585**	**3,076,974**	**3,007,565**	**2,960,143**	**2,976,803**	**2,972,604**	**2,900,835**	**2,821,651**
65–74 y. persons dependent on care	I	0.99%	0.94%	0.89%	0.84%	0.79%	0.75%	0.71%	0.66%	0.63%
	II	1.57%	1.57%	1.56%	1.55%	1.55%	1.54%	1.53%	1.52%	1.51%
	III	1.97%	2.15%	2.40%	2.68%	3.00%	3.34%	3.73%	4.16%	4.63%
	IV	0.83%	0.90%	1.00%	1.11%	1.23%	1.37%	1.52%	1.68%	1.86%
	Total	**5.36%**	**5.57%**	**5.86%**	**6.19%**	**6.57%**	**7.00%**	**7.48%**	**8.02%**	**8.62%**
Abs. num. of inhabitants – men		**660,788**	**744,711**	**724,491**	**667,179**	**660,666**	**604,771**	**573,558**	**644,936**	**693,770**
75–84 y. persons dependent on care	I	1.43%	1.26%	1.08%	0.92%	0.79%	0.67%	0.57%	0.49%	0.42%
	II	2.55%	2.40%	2.21%	2.04%	1.88%	1.73%	1.60%	1.47%	1.35%
	III	3.81%	3.97%	4.19%	4.41%	4.64%	4.88%	5.14%	5.40%	5.67%
	IV	1.70%	1.75%	1.82%	1.88%	1.95%	2.02%	2.09%	2.16%	2.24%
	Total	**9.49%**	**9.38%**	**9.29%**	**9.25%**	**9.26%**	**9.31%**	**9.40%**	**9.52%**	**9.68%**
Abs. num. of inhabitants – men		**392,272**	**412,310**	**490,443**	**586,484**	**578,614**	**546,585**	**551,719**	**512,043**	**495,752**
85–94 y. persons dependent on care	I	3.38%	2.88%	2.34%	1.90%	1.54%	1.25%	1.01%	0.81%	0.66%
	II	6.91%	6.35%	5.71%	5.12%	4.57%	4.08%	3.64%	3.24%	2.87%
	III	12.17%	12.75%	13.49%	14.23%	14.99%	15.76%	16.53%	17.31%	18.11%
	IV	5.91%	6.23%	6.65%	7.07%	7.50%	7.95%	8.41%	8.88%	9.36%
	Total	**28.37%**	**28.21%**	**28.18%**	**28.32%**	**28.61%**	**29.04%**	**29.59%**	**30.25%**	**31.00%**
Abs. num. of inhabitants – men		**154,848**	**158,982**	**172,085**	**194,296**	**248,707**	**306,736**	**304,900**	**305,362**	**319,812**
95+ y. persons dependent on care	I	6.81%	6.04%	5.14%	4.31%	3.57%	2.93%	2.38%	1.92%	1.54%
	II	11.92%	10.74%	9.30%	7.95%	6.70%	5.60%	4.63%	3.80%	3.10%
	III	37.42%	40.47%	44.06%	47.32%	50.19%	52.68%	54.77%	56.50%	57.91%
	IV	20.30%	22.04%	24.11%	26.03%	27.74%	29.26%	30.57%	31.70%	32.65%
	Total	**76.44%**	**79.30%**	**82.62%**	**85.60%**	**88.22%**	**90.46%**	**92.35%**	**93.92%**	**95.19%**
Abs. num. of inhabitants – men		**7,208**	**10,577**	**13,344**	**14,609**	**18,393**	**22,084**	**32,730**	**40,647**	**40,763**
Total absolute population – men		**5,304,673**	**5,316,646**	**5,329,742**	**5,327,373**	**5,297,417**	**5,247,025**	**5,193,801**	**5,153,401**	**5,129,991**
Tot. num. of dependent on care – men of which in:		**3.75%**	**4.01%**	**4.36%**	**4.76%**	**5.29%**	**5.82%**	**6.29%**	**6.80%**	**7.32%**
Level I		0.65%	0.64%	0.61%	0.59%	0.56%	0.53%	0.50%	0.48%	0.46%
Level II		0.98%	1.00%	1.02%	1.03%	1.05%	1.05%	1.03%	1.02%	1.03%
Level III		1.36%	1.52%	1.75%	2.01%	2.35%	2.69%	2.99%	3.32%	3.63%
Level IV		0.76%	0.85%	0.98%	1.13%	1.33%	1.55%	1.77%	1.98%	2.20%
Tot. num. of dependent on care – men (levels I–IV)		**199,154**	**213,145**	**232,396**	**253,457**	**280,345**	**305,557**	**326,596**	**350,680**	**375,359**

Source: Authors' calculations (multinomial mixed-effects model output).

Table A5 **Extended projection period (2041–2080, 40 years), share of women dependent on care in the total female population**

Age group	Dependency level	Extended projection period (2041–2080)								
		2041	2045	2050	2055	2060	2065	2070	2075	2080
0–17	I	0.36%	0.34%	0.32%	0.30%	0.27%	0.26%	0.24%	0.22%	0.21%
	II	0.42%	0.45%	0.48%	0.52%	0.56%	0.61%	0.65%	0.71%	0.76%
	III	0.25%	0.27%	0.28%	0.30%	0.32%	0.34%	0.37%	0.39%	0.42%
	IV	0.16%	0.15%	0.15%	0.14%	0.14%	0.13%	0.13%	0.13%	0.12%
	Total	**1.19%**	**1.21%**	**1.23%**	**1.26%**	**1.30%**	**1.34%**	**1.39%**	**1.45%**	**1.51%**
Abs. num. of inhabitants – women		**770,958**	**790,470**	**816,088**	**821,598**	**797,348**	**759,085**	**729,166**	**721,025**	**729,378**
18–64	I	0.26%	0.25%	0.24%	0.23%	0.22%	0.22%	0.21%	0.20%	0.19%
	II	0.30%	0.29%	0.28%	0.28%	0.27%	0.26%	0.25%	0.24%	0.23%
	III	0.37%	0.39%	0.42%	0.45%	0.48%	0.52%	0.56%	0.60%	0.64%
	IV	0.29%	0.32%	0.36%	0.41%	0.46%	0.52%	0.59%	0.67%	0.75%
	Total	**1.22%**	**1.26%**	**1.31%**	**1.37%**	**1.44%**	**1.52%**	**1.61%**	**1.71%**	**1.82%**
Abs. num. of inhabitants – women		**3,072,631**	**2,942,199**	**2,848,278**	**2,778,256**	**2,739,886**	**2,769,477**	**2,780,071**	**2,729,670**	**2,665,032**
65–74	I	0.86%	0.78%	0.69%	0.61%	0.54%	0.48%	0.42%	0.37%	0.33%
	II	1.28%	1.25%	1.22%	1.19%	1.16%	1.12%	1.09%	1.06%	1.03%
	III	1.50%	1.64%	1.83%	2.05%	2.28%	2.55%	2.84%	3.16%	3.52%
	IV	0.68%	0.74%	0.81%	0.90%	1.00%	1.10%	1.22%	1.35%	1.49%
	Total	**4.32%**	**4.41%**	**4.56%**	**4.75%**	**4.98%**	**5.25%**	**5.57%**	**5.94%**	**6.37%**
Abs. num. of inhabitants – women		**708,176**	**790,192**	**759,214**	**683,601**	**660,839**	**590,860**	**551,302**	**614,194**	**662,776**
75–84	I	1.97%	1.61%	1.25%	0.97%	0.75%	0.58%	0.45%	0.35%	0.27%
	II	2.96%	2.64%	2.28%	1.96%	1.69%	1.46%	1.25%	1.08%	0.92%
	III	3.55%	3.52%	3.47%	3.42%	3.37%	3.32%	3.26%	3.20%	3.14%
	IV	2.12%	2.13%	2.14%	2.15%	2.16%	2.16%	2.17%	2.17%	2.17%
	Total	**10.61%**	**9.90%**	**9.15%**	**8.51%**	**7.98%**	**7.52%**	**7.13%**	**6.80%**	**6.51%**
Abs. num. of inhabitants – women		**504,748**	**514,161**	**592,268**	**692,222**	**669,271**	**611,981**	**597,988**	**538,554**	**508,900**
85–94	I	5.20%	4.35%	3.46%	2.74%	2.16%	1.70%	1.34%	1.05%,	0.82%
	II	9.88%	8.87%	7.72%	6.70%	5.79%	4.99%	4.29%	3.68%	3.16%
	III	17.25%	17.71%	18.22%	18.69%	19.10%	19.47%	19.79%	20.09%	20.35%
	IV	11.49%	11.92%	12.42%	12.89%	13.34%	13.76%	14.17%	14.55%	14.93%
	Total	**43.82%**	**42.84%**	**41.82%**	**41.01%**	**40.39%**	**39.92%**	**39.59%**	**39.38%**	**39.26%**
Abs. num. of inhabitants – women		**288,074**	**281,631**	**280,916**	**300,520**	**364,940**	**436,462**	**423,349**	**404,212**	**405,318**
95+	I	3.53%	2.89%	2.25%	1.73%	1.33%	1.02%	0.78%	0.59%	0.45%
	II	11.11%	9.75%	8.24%	6.93%	5.80%	4.84%	4.02%	3.34%	2.76%
	III	34.19%	34.90%	35.61%	36.17%	36.58%	36.86%	37.03%	37.10%	37.09%
	IV	33.60%	35.10%	36.88%	38.56%	40.15%	41.65%	43.08%	44.44%	45.74%
	Total	**82.44%**	**82.64%**	**82.97%**	**83.39%**	**83.86%**	**84.37%**	**84.91%**	**85.47%**	**86.05%**
Abs. num. of inhabitants – women		**27,345**	**36,845**	**43,053**	**43,481**	**48,765**	**55,328**	**75,514**	**90,366**	**86,620**
Total absolute population – women		**5,371,933**	**5,355,497**	**5,339,817**	**5,319,678**	**5,281,049**	**5,223,192**	**5,157,389**	**5,098,021**	**5,058,023**
Tot. num. of dependent on care – women of which in:		**5.20%**	**5.29%**	**5.42%**	**5.63%**	**6.14%**	**6.70%**	**6.98%**	**7.19%**	**7.29%**
Level I		0.80%	0.71%	0.62%	0.54%	0.48%	0.43%	0.36%	0.31%	0.28%
Level II		1.27%	1.20%	1.12%	1.07%	1.04%	0.99%	0.90%	0.82%	0.76%
Level III		1.88%	2.00%	2.16%	2.34%	2.67%	3.02%	3.20%	3.34%	3.44%
Level IV		1.27%	1.38%	1.52%	1.68%,	1.95%	2.27%	2.51%	2.71%	2.81%
Tot. num. of dependent on care – women (levels I–IV)		**279,602**	**283,391**	**289,346**	**299,310**	**324,326**	**350,174**	**359,903**	**366,577**	**368,585**

Source: Authors' calculations (multinomial mixed-effects model output).

Table A6 **Extended projection period (2041–2080, 40 years), share of men and women dependent on care in the total population (male and female).**

Age group	Dependency level	Extended projection period (2041–2080)								
		2041	2045	2050	2055	2060	2065	2070	2075	2080
0–17	I	0.49%	0.47%	0.45%	0.43%	0.41%	0.39%	0.38%	0.36%	0.34%
	II	0.66%	0.72%	0.79%	0.87%	0.97%	1.07%	1.18%	1.30%	1.43%
	III	0.49%	0.54%	0.62%	0.70%	0.79%	0.90%	1.03%	1.17%	1.33%
	IV	0.25%	0.25%	0.26%	0.26%	0.27%	0.27%	0.28%	0.28%	0.29%
	Total	**1.90%**	**1.99%**	**2.12%**	**2.27%**	**2.44%**	**2.63%**	**2.86%**	**3.11%**	**3.39%**
Abs. num. of inhabitants – men and women		**1,578,233**	**1,616,950**	**1,668,492**	**1,678,836**	**1,628,242**	**1,549,131**	**1,487,455**	**1,470,602**	**1,487,622**
18–64	I	0.31%	0.31%	0.31%	0.30%	0.30%	0.30%	0.30%	0.29%	0.29%
	II	0.35%	0.34%	0.34%	0.33%	0.32%	0.31%	0.30%	0.29%	0.29%
	III	0.44%	0.47%	0.51%	0.55%	0.60%	0.65%	0.70%	0.76%	0.82%
	IV	0.37%	0.42%	0.48%	0.56%	0.64%	0.74%	0.85%	0.98%	1.13%
	Total	**1.48%**	**1.54%**	**1.63%**	**1.74%**	**1.86%**	**2.00%**	**2.15%**	**2.33%**	**2.53%**
Abs. num. of inhabitants – men and women		**6,354,912**	**6,105,784**	**5,925,251**	**5,785,821**	**5,700,028**	**5,746,280**	**5,752,675**	**5,630,505**	**5,486,682**
65–74	I	0.92%	0.86%	0.79%	0.73%	0.67%	0.61%	0.57%	0.52%	0.48%
	II	1.42%	1.40%	1.39%	1.37%	1.35%	1.33%	1.32%	1.30%	1.28%
	III	1.73%	1.89%	2.11%	2.36%	2.64%	2.95%	3.29%	3.67%	4.09%
	IV	0.75%	0.82%	0.91%	1.01%	1.12%	1.24%	1.37%	1.52%	1.68%
	Total	**4.82%**	**4.97%**	**5.19%**	**5.46%**	**5.77%**	**6.14%**	**6.55%**	**7.01%**	**7.52%**
Abs. num. of inhabitants – men and women		**1,368,964**	**1,534,903**	**1,483,705**	**1,350,780**	**1,321,505**	**1,195,631**	**1,124,859**	**1,259,130**	**1,356,547**
75–84	I	1.74%	1.46%	1.17%	0.95%	0.77%	0.63%	0.51%	0.42%	0.34%
	II	2.78%	2.53%	2.25%	2.00%	1.78%	1.59%	1.42%	1.27%	1.14%
	III	3.66%	3.72%	3.80%	3.87%	3.96%	4.06%	4.16%	4.27%	4.39%
	IV	1.94%	1.96%	2.00%	2.03%	2.06%	2.10%	2.13%	2.17%	2.20%
	Total	**10.12%**	**9.67%**	**9.21%**	**8.85%**	**8.57%**	**8.36%**	**8.22%**	**8.13%**	**8.08%**
Abs. num. of inhabitants – men and women		**897,021**	**926,471**	**1,082,711**	**1,278,706**	**1,247,885**	**1,158,566**	**1,149,707**	**1,050,597**	**1,004,651**
85–94	I	4.57%	3.82%	3.03%	2.41%	1.91%	1.52%	1.20%	0.95%	0.75%
	II	8.84%	7.96%	6.96%	6.08%	5.30%	4.62%	4.02%	3.49%	3.03%
	III	15.47%	15.92%	16.43%	16.94%	17.43%	17.93%	18.43%	18.90%	19.36%
	IV	9.54%	9.87%	10.23%	10.60%	10.97%	11.36%	11.76%	12.11%	12.47%
	Total	**38.42%**	**37.56%**	**36.64%**	**36.03%**	**35.62%**	**35.43%**	**35.40%**	**35.45%**	**35.62%**
Abs. num. of inhabitants – men and women		**442,922**	**440,613**	**453,002**	**494,817**	**613,647**	**743,198**	**728,249**	**709,575**	**725,129**
95+	I	4.21%	3.60%	2.93%	2.38%	1.95%	1.57%	1.26%	1.01%	0.80%
	II	11.28%	9.97%	8.49%	7.18%	6.05%	5.06%	4.21%	3.48%	2.87%
	III	34.87%	36.14%	37.61%	38.97%	40.31%	41.37%	42.39%	43.12%	43.75%
	IV	30.82%	32.18%	33.86%	35.41%	36.75%	38.12%	39.30%	40.49%	41.55%
	Total	**81.18%**	**81.89%**	**82.89%**	**83.94%**	**85.05%**	**86.11%**	**87.16%**	**88.09%**	**88.97%**
Abs. num. of inhabitants – men and women		**34,553**	**47,422**	**56,396**	**58,091**	**67,159**	**77,412**	**108,244**	**131,013**	**127,383**
Total absolute population – men and women		**10,676,606**	**10,672,143**	**10,669,558**	**10,647,051**	**10,578,466**	**10,470,218**	**10,351,190**	**10,251,422**	**10,188,014**
Tot. num. of dependent on care – men and women of which in:		**4.48%**	**4.65%**	**4.89%**	**5.19%**	**5.72%**	**6.26%**	**6.63%**	**7.00%**	**7.30%**
Level I		0.73%	0.67%	0.61%	0.56%	0.52%	0.48%	0.43%	0.40%	0.37%
Level II		1.13%	1.10%	1.07%	1.05%	1.04%	1.02%	0.96%	0.92%	0.90%
Level III		1.62%	1.76%	1.95%	2.18%	2.51%	2.85%	3.10%	3.33%	3.54%
Level IV		1.01%	1.12%	1.25%	1.40%	1.64%	1.91%	2.14%	2.34%	2.50%
Tot. num. of dependent on care – men and women (levels I–IV)		**478,755**	**496,537**	**521,742**	**552,767**	**604,671**	**655,731**	**686,499**	**717,257**	**743,944**

Source: Authors' calculations (multinomial mixed-effects model output).

Table A7 **Mixed-effects multinomial model fixed effects (reference category: dependency level I and male, 0–17 years)**

	Term	Estimate	Std. Error	z	P
Panel: II vs I					
	intercept	–0.238***	0.007	–35.28	<0.001
	year (std.)	0.047***	0.010	4.88	<0.001
	male, 18–64	0.459***	0.008	55.17	<0.001
	male, 65–74	0.467***	0.009	52.79	<0.001
	male, 75–84	0.478***	0.009	54.19	<0.001
	male, 85–94	0.518***	0.010	54.42	<0.001
	male, 95+	0.713***	0.027	26.20	<0.001
	female, 0–17	–0.256***	0.010	–24.87	<0.001
	female, 18–64	0.374***	0.009	43.91	<0.001
	female, 65–74	0.207***	0.009	24.02	<0.001
	female, 75–84	0.174***	0.008	21.85	<0.001
	female, 85–94	0.474***	0.008	58.99	<0.001
	female, 95+	1.001***	0.015	64.67	<0.001
Panel: III vs I					
	Term	**Estimate**	**Std. Error**	**z**	**P**
	intercept	–0.617***	0.007	–86.29	<0.001
	year (std.)	0.126***	0.012	10.74	<0.001
	male, 18–64	0.632***	0.008	76.42	<0.001
	male, 65–74	0.569***	0.009	63.81	<0.001
	male, 75–84	0.671***	0.009	75.90	<0.001
	male, 85–94	0.742***	0.010	77.96	<0.001
	male, 95+	1.251***	0.027	46.59	<0.001
	female, 0–17	–0.344***	0.011	–31.16	<0.001
	female, 18–64	0.501***	0.009	58.68	<0.001
	female, 65–74	0.150***	0.009	17.15	<0.001
	female, 75–84	0.136***	0.008	17.12	<0.001
	female, 85–94	0.669***	0.008	84.11	<0.001
	female, 95+	1.673***	0.015	110.67	<0.001
Panel: IV vs I					
	Term	**Estimate**	**Std. Error**	**z**	**P**
	intercept	–0.906***	0.008	–116.45	<0.001
	year (std.)	0.124***	0.015	8.44	<0.001
	male, 18–64	0.490***	0.009	54.13	<0.001
	male, 65–74	0.023*	0.010	2.27	0.023
	male, 75–84	0.225***	0.010	22.51	<0.001
	male, 85–94	0.273***	0.011	24.98	<0.001
	male, 95+	0.904***	0.030	29.78	<0.001
	female, 0–17	–0.117***	0.012	–10.07	<0.001
	female, 18–64	0.347***	0.009	36.94	<0.001
	female, 65–74	–0.327***	0.010	–32.22	<0.001
	female, 75–84	–0.168***	0.009	–19.12	<0.001
	female, 85–94	0.499***	0.009	57.52	<0.001
	female, 95+	1.817***	0.016	116.55	<0.001
Panel: No care vs I					
	Term	**Estimate**	**Std. Error**	**z**	**P**
	intercept	4.935***	0.005	986.94	<0.001
	year (std.)	0.074***	0.018	4.11	<0.001
	male, 18–64	0.727***	0.006	126.33	<0.001
	male, 65–74	–0.608***	0.006	–97.74	<0.001
	male, 75–84	–1.484***	0.006	–237.59	<0.001
	male, 85–94	–2.793***	0.007	–403.98	<0.001
	male, 95+	–3.633***	0.025	–148.26	<0.001
	female, 0–17	0.357***	0.007	52.83	<0.001
	female, 18–64	0.866***	0.006	147.06	<0.001
	female, 65–74	–0.762***	0.006	–129.46	<0.001
	female, 75–84	–2.279***	0.005	–420.37	<0.001
	female, 85–94	–3.650***	0.006	–648.02	<0.001
	female, 95+	–4.370***	0.015	–285.17	<0.001

Notes: Estimates are log-odds from mclogit::mblogit (PQL). Significance: * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
Source: Authors' calculations (multinomial mixed-effects model parameters).

Table A8 **Mixed-effects multinomial model random effects (reference category: dependency level I)**

Random effects: Variance–covariance matrix estimates

	II: intercept	III: intercept	IV: intercept	no care: intercept	II: year(std.)	III: year(std.)	IV: year(std.)	no care: year(std.)
II: intercept	$5.201 \cdot 10^{-5}$							
III: intercept	$3.113 \cdot 10^{-5}$	$8.251 \cdot 10^{-5}$						
IV: intercept	$2.348 \cdot 10^{-5}$	$7.882 \cdot 10^{-5}$	$9.304 \cdot 10^{-5}$					
no care: intercept	$3.151 \cdot 10^{-5}$	$6.245 \cdot 10^{-5}$	$7.545 \cdot 10^{-5}$	$7.255 \cdot 10^{-5}$				
II: year(std.)	$-4.504 \cdot 10^{-5}$	$1.547 \cdot 10^{-4}$	$2.397 \cdot 10^{-4}$	$1.880 \cdot 10^{-4}$	$1.059 \cdot 10^{-3}$			
III: year(std.)	$1.910 \cdot 10^{-4}$	$1.825 \cdot 10^{-4}$	$2.417 \cdot 10^{-4}$	$2.887 \cdot 10^{-4}$	$5.217 \cdot 10^{-4}$	$1.586 \cdot 10^{-3}$		
IV: year(std.)	$3.005 \cdot 10^{-4}$	$1.860 \cdot 10^{-4}$	$1.408 \cdot 10^{-4}$	$2.260 \cdot 10^{-4}$	$-2.659 \cdot 10^{-4}$	$1.533 \cdot 10^{-3}$	$2.526 \cdot 10^{-3}$	
no care: year(std.)	$2.299 \cdot 10^{-4}$	$5.126 \cdot 10^{-4}$	$4.976 \cdot 10^{-4}$	$4.406 \cdot 10^{-4}$	$9.596 \cdot 10^{-4}$	$1.723 \cdot 10^{-3}$	$2.007 \cdot 10^{-3}$	$3.817 \cdot 10^{-3}$

Source: Authors' calculations (multinomial mixed-effects model parameters).

PODKLADY

Redakce přijímá rukopisy v elektronické podobě.

ROZSAH PŘÍSPĚVKU:

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1. Pro zpracování grafů je kromě požadovaného typu (sloupcový, spojnicový, bodový apod.) nutné připojit zdrojová data v programu Excel.
2. Všechny obrazové soubory – např. mapy, fotografie ukládejte mimo textový soubor samostatně ve formátech *.tif, *.jpg, *.eps s odkazem v textu (graf 1, schéma 1 apod.).
3. Pro další technologické zpracování je důležité, aby bitmapové soubory měly ve velikosti 1:1 rozlišení 300 dpi.

C. PRAVIDLA CITACÍ A POPISKY

Pokud má dokument přiděleno doi, musí být v citaci uvedeno.

Příklady základních druhů citací:

Monografie

Roubíček, V. 1997. *Úvod do demografie*. Praha: Codex Bohemia. (U publikace s více než třemi autory se uvádí jen příjmení prvního autora, za ním následuje zkratka a kol., u zahraničních publikací et al.)

Hantrais, L. (ed.). 2000. *Gendered Policies in Europe. Reconciling Employment and Family Life*. London: Macmillan Press.

Potraty. 2005. Praha: Ústav zdravotnických informací a statistiky.

Články v časopisech

Dudová, R. – Vohlídalová, M. 2018. Muži a ženy pečující o seniory v rodině. Sociologický časopis, 54(2), s. 219–252. https://doi.org/10.13060/00380288.2018.54.2.400.

Články ve sbornících

Daly, M. 2004. Rodinná politika v evropských zemích. In *Perspektivy rodinné politiky v ČR*, s. 62–71. Praha: MPSV ČR.

Elektronické dokumenty

Je třeba uvést:

1. Specifikaci média (on-line, databáze, datový soubor)
2. Datum stažení (cit. 29. 10. 2005)
3. Webovou adresu (dostupné z: http://www.csu.gov.cz)

Přednášky z konferencí

Maur, E. *Problémy studia migrací v českých zemích v raném novověku.* Příspěvek přednesený na konferenci Dějiny migrací v českých zemích v novověku. Praha, 14. 10. 2005.

Seznam literatury a odkazy

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Příklad:

Syrovátka, A. 1962a. Úrazy v domácnosti. *Česká pediatrie*, 17, s. 750–753.

Syrovátka, A. 1962b. Úmrtnost dětí v českých zemích na dopravní úrazy. *Časopis lékařů českých*, 101, s. 1513–1517.

Odkazy v textu na seznam literatury

(Srb, 2004*)*; *(Srb,* 2004: 36–37*)*; *(Syrovátka a kol.,* 1984*)*; (*Dudová – Vohlídalová*, 2018)

Popisky tabulek a grafů (dodat v češtině a angličtině)

Tab. 1: Pohyb obyvatelstva, 1990–2010; Population and vital statistics, 1990–2010

Graf 1: Relativní věková struktura cizinců a obyvatelstva ČR celkem, 31. 12. 2009; Relative age distribution of foreigners and total population of CR, 31 Dec 2009

D. DOPORUČENÁ STRUKTURA ČASOPISU

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