PE Population and Economics

Population and Economics 10(3): 82–104
DOI 10.3897/popecon.10.e199193

RESEARCH ARTICLE

Gender Differences in Housing Provision in Russia: An Age–Period–Cohort Decomposition, 2003–2023

Aleksandra Ya. Burdyak [1]

1 *RANEPA, Moscow, 119034, Russia*

Received 10 May 2026 ♦ **Accepted** 21 August 2026 ♦ **Published** 30 October 2026

Citation: Burdyak AYa (2026) Gender Differences in Housing Provision in Russia: An Age–Period–Cohort Decomposition, 2003–2023. Population and Economics 10(3):82-104. https://doi.org/10.3897/popecon.10.e199193

Abstract

The well-being of women and men within households is a central concern of gender research in demography and economics, yet it has rarely been examined in relation to housing in Russia using large-scale microdata. Over 2000–2024, Russia's housing provision increased by half against a backdrop of household nuclearisation, a female longevity advantage, and uneven generational access to the post-Soviet housing stock. Whether this expansion reflects macro-level growth or generational replacement – and whether gender differentially shapes the housing trajectory – remains an open question.

We use Rosstat's Household Budget Survey microdata for 2003–2023 (2.61 million observations) and apply an age–period–cohort (APC) decomposition with the Intrinsic Estimator (IE). The pseudo-panel comprises 354 cells across 19 five-year cohorts (1930–2024), 18 age groups (0–4, 5–9, …, 85+) and 21 survey years. The APC decomposition is estimated separately for the combined, male and female samples.

Under the IE identification, age is the largest component of the variance of the per-capita floor area (46%; 37% for men and 51% for women), followed by birth cohort (29%) and survey year (25%). The split between the components and the direction of the intergenerational gradient depend on the identifying restriction, whereas the gender difference in cohort and age effects is robust to it. The within-cohort improvement (+9.4 m^2) is partly offset by a negative compositional shift (−1.8 m^2): older, better-housed cohorts are being replaced by younger, worse-housed ones. The age profile reveals low housing provision at ages 30–44 and a peak at ages 65–69, when the household contracts to a childless couple for men and, more often, to a one-person household for women. For women, the swing between the two stages is nearly twice that for men. Mean housing provision is higher for women (+2.83 m^2 in 2023). The gender gap is significantly positive in most cohorts born in 1935–1974 and significantly changes sign in the cohorts born in 1975–2014, where men are better housed; in the two youngest cohorts (2015–2019 and 2020–2024), observed in only part of the survey waves, the gap is not statistically significant.

The gender gap in housing provision largely reflects the demographically determined composition of households – for example, the prevalence of elderly women living alone (widows, elderly women who

Copyright *Burdyak AYa.* This is an open access article distributed under the terms of the Creative Commons Attribution License (CC-BY 4.0), which permits unrestricted use, distribution, and reproduction in any medium, provided the original author and source are credited.

have never been married) – rather than differences in individual housing rights. The growth in the mean per-capita floor area over 20 years conceals opposing dynamics within and between cohorts, as well as the zigzag pattern of housing provision variation across the stages of the family life cycle.

Keywords

age–period–cohort, gender gap, family life cycle, housing provision, cohort analysis, Intrinsic Estimator, pseudo-panel, Russia, shift-share decomposition

JEL codes: D31, J11, J16, R21

Introduction

Over the last quarter of a century, housing provision for the Russian population has increased by a factor of 1.5 as a result of increased construction and wider access to housing, rising from a mean of 19.2 m^2 per person in 2000 to 29.4 m^2 in 2024 (Rosstat 2025). Alongside the growth in housing stock, the mean household size has decreased; census data show that the proportion of single-person households increased from 22.3 per cent in 2002 to 41.8 per cent in 2021 (Prokofieva and Korchagina 2023). The progress made in the housing sector has affected different cohorts to varying degrees. In the 1990s, Russians entered adulthood amidst a housing shortage. In the 2000s and 2010s, however, housing construction expanded, and mortgage and preferential lending schemes developed, as reported by Kosareva et al. (2015) and Zhirnov (2025). Older cohorts occupying their dwellings at the time of privatisation were the main beneficiaries of the free transfer of the housing stock to its occupants as owners (Zavisca 2012).

The gender dimension of housing adequacy in Russia is of particular interest, given that demographic factors – such as women's higher life expectancy, which results in their predominance among widows and among older adults living alone – create systematic differences in the gender composition of households. International literature consistently highlights housing vulnerability of women, ranging from the challenges faced by single mothers (Nieuwenhuis and Zagel 2023) to the loss of housing wealth experienced by widows (Kapelle and Van Winkle 2024). The Russian context, which combines historically high female employment and a high proportion of female homeowners resulting from equal rights to home ownership and privatisation, as well as a significant proportion of single older women, presents a distinct gender-specific housing profile that has yet to be quantitatively analysed using microdata.

As part of the journal's special issue on gender studies in economics and demography, we examine the differences in housing trajectories between women and men as household members. This study aims to decompose the effects of age, cohort, and survey year on housing provision among the Russian population between 2003 and 2023, providing separate estimates for men and women. The study addresses three key research questions. Firstly, what is the relative importance of the three APC effects – biographical (age), macroeconomic (period) and generational (cohort) – in explaining the variation observed in housing provision? Secondly, what factors have contributed to the observed increase in the mean per-capita housing area over 20 years: improvements within cohorts or a structural shift resulting from generational replacement? Thirdly, are there significant gender differences in the APC effects, and at what ages and across which generations are they most pronounced?

Literature review

During the post-Soviet period, the housing sector in the Russian Federation underwent several stages of institutional development, transitioning from a centralised, socialist system to a market-based one (Zavisca 2012; Kalyukin and Kohl 2020). After 1991, the reform of the sector was accompanied by a lack of investment, inefficient urban planning and the deterioration of the housing stock. The proportion of privately owned housing increased from 27 per cent in the early 1990s to 89 per cent by the mid-2010s, by which time market institutions such as the mortgage and rental markets, as well as the residential building management sector, had begun to rapidly develop, as reported by Kosareva et al. (2015). However, housing wealth in Russia has remained largely a frozen asset, with limited mobility in the housing market, and the main channels for improving housing conditions being intra-family transfers and inheritance (Gerber et al. 2022). Most families' living conditions are determined by housing constructed during the Soviet era. According to the 2020 All-Russian Population Census (conducted in 2021), 75–76 per cent of the population live in housing built before 1995 (Pilipenko 2025). Home ownership became more affordable in Russia between 2005 and 2020 due to reductions in prices and mortgage rates, as well as government support programmes for mortgage lending. This is in contrast to the opposite trend observed in most developed countries (Kosareva and Polidi 2019; 2021). However, Russian estimates reveal significant disparities in housing affordability and provision across regions and among different household types, including families with children (Roshchina and Ilyunkina 2021; Guzanova 2025), as well as intergenerational disparities in housing provision within the population (Burdyak 2015; Bobkov and Odintsova 2021). Apart from regional and intergenerational disparities, gender inequality exists in Russia, too, with a gender pay gap of around 30 per cent (Gerber and Gimpelson 2024), which limits working women's access to mortgage lending and reduces their ability to accumulate housing wealth. The uneven distribution of housing wealth in the country has already been reported (Cherkashina 2025).

International studies highlight the vulnerability of women's housing situation. In European countries[1], single mothers are more likely to experience housing deprivation, including overcrowding, poor housing conditions, high housing costs as a proportion of the household budget, and problems arising from their neighbourhood. In countries with a high proportion of homeowners, single mothers are more likely to experience cramped living conditions and to live in rented accommodation (Nieuwenhuis and Zagel 2023). Employment insecurity impedes entry into first-time homeownership, particularly in the more marketised housing systems of Northern and Western Europe (Lersch and Dewilde 2015). In the US, single women earn 1.5 percentage points less than single men in terms of the annual return on housing investments (Goldsmith-Pinkham and Shue 2023). In France, despite nominally egalitarian property laws, women accumulate less wealth than men due to their property claims being undervalued in divorce and inheritance cases (Bessière and Gollac 2023).

In the international literature, the relationship between demographic composition and housing consumption is examined through the lens of cohorts. A contextually similar study using German Socio-Economic Panel Survey data from 1984 to 2018 showed that, when

1 The two works cited here are both based on an analysis of data from European Union Statistics on Income and Living Conditions (EU-SILC), https://ec.europa.eu/eurostat/web/income-and-living-conditions.

controlling for income level, wealth, household size, and age, later cohorts consume significantly more per-capita floor area (Hein and Kuhnimhof 2026). An analysis of China's data (Fu 2016) revealed cohort differences in housing status attributable to property reforms, as shown by a hierarchical APC model. A study of cohort progression through life cycle stages in the US (Myers et al. 2025) showed that credit restrictions had slowed millennials' transition to home ownership. However, this lag narrowed rapidly after 2016.

In the Russian context, the APC approach is appropriate due to the presence of demographic waves, such as the Soviet cohorts of varying sizes and the demographic trough of the 1990s. The classic application of APC models is the study of cohort differences in mortality (Pustovalov 2015; Maksimov 2025). The APC method has also been used to estimate the effects of cohorts on fertility trends, including by birth order (Vakulenko and Ryzhkina 2024). Differences in the capabilities of older cohorts have been examined (Davis 2018), as well as inequalities between cohorts in terms of consumption (Bryukhanov and Hryshko 2024), including alcohol consumption (Radaev and Roshchina 2019).

The APC method is based on the pseudo-panel technique (Deaton 1985). In the absence of true panel data, cohorts defined by year of birth are tracked across successive cross-sectional surveys and the mean values of these cohorts are used as observations in the panel model. If the cohorts are large enough, the measurement error is small, and the estimate is asymptotically equivalent to the true panel estimate.

The linear relationship between the three variables – age, period and cohort – necessitates the imposition of additional explicit or implicit restrictions on the parameters and their relationships. It is well known that no method provides a perfect solution to the identification problem and that APC modelling results depend partly on model design choice, as the relative contributions of age, period and cohort effects vary according to the restrictions imposed. This supports the argument for viewing the results of APC modelling critically (Luo 2013; Fosse and Winship 2019). The traditional Deaton–Paxson (DP) approach (Deaton and Paxson 1994) resolves the identification problem by assuming that period effects do not exhibit a linear trend. In this model, the period variable captures only cyclical deviations from the trend, while the trend itself is explained by age and cohort rather than the year of observation.

The Intrinsic Estimator (IE) uses a different restriction. From the set of solutions, the one that is orthogonal to the null space of the design matrix – that is, the one with the minimum Euclidean norm – is selected. This solution is provided by the Moore–Penrose pseudoinverse and the linear trend is distributed among three sets of coefficients. The IE is consistent in large samples and is widely used in demographic and sociological APC studies (Yang and Land 2013). Apart from the effect coding (sum-to-zero coding with an intercept) that is canonical for the IE, the minimum-norm solution can also be sought by coding the full set of dummy variables (Pelzer et al. 2015), with the two IE solutions differing precisely by a linear trend, meaning they should be compared to test the robustness of the effects (te Grotenhuis et al. 2016).

The choice between DP and IE is a matter of substantive interpretation. If the linear trend of improving housing conditions is interpreted as intergenerational, with each subsequent generation building more than the previous one, then the DP is preferable. However, if the trend is regarded as macroeconomic, with the housing stock growing simultaneously for all cohorts, then the IE is more appropriate. As Russia's housing stock grew by almost 50 per cent between 2003 and 2023, while the mean per-capita floor area as per the Sample Survey of Household Budgets data increased by 28.7 per cent, the macro-trend of growth in housing

provision is significant. Therefore, the DP's assumption of a zero linear trend for this period appears implausible. This is precisely why we used the IE in this study.

In order to analyse medium-term growth in housing provision, it was broken down into two components below: growth in provision within cohorts and growth resulting from older cohorts being replaced by younger ones (the composition effect) (Kitagawa 1955).

Data and modelling

Data and sample

The data source is Rosstat's microdata from the Sample Survey of Household Budgets for 2003–2023 (21 waves). Rosstat has conducted the survey quarterly since 1996, using a sample of approximately 45,000–50,000 households per year. The data include individual characteristics for each family member. The target variable is the per-capita floor area, which is calculated by dividing the total floor area of the dwelling by the number of residents. This value is the same for all members of the household. A cohort is defined according to the year of birth of each household member. A single household may belong to several cohort cells simultaneously (for example, parents and children) and be included in both gender subsamples. Outliers are removed from the sample: an observation is excluded if the floor area divided by the number of rooms exceeds the median value by more than a factor of four.

As the Sample Survey of Household Budgets is a repeated cross-sectional survey rather than a longitudinal panel, the synthetic/pseudo-panel method (Deaton 1985) was used to model the APC. Individuals were aggregated into 'cohort × survey year' cells across 18 age groups (0–4, 5–9, ..., 85+), 19 five-year cohorts (1930–2024) and 21 periods (2003–2023). Cohorts born between 1930 and 2004 were included in all waves, whereas younger cohorts were only partially observed. For instance, the 2015–2019 cohort was included in nine survey waves (from 2015 to 2023), while the youngest cohort, born between 2020 and 2024, was only partially observed in four waves (from 2020 to 2023).

The sample was split by gender for the gender breakdown, and the APC model was estimated overall for all individuals, as well as for each sex separately, using KVZV weighting[2]. Cells containing fewer than 200 observations were excluded from the analysis (Verbeek and Nijman 1992). This comprises seven cells for those born between 1930 and 1934 in the male subsample for the period 2017–2023. The same seven cells were also excluded from the combined and female samples. The final pseudo-panel comprised 354 cells in each of the three samples (19 five-year cohorts × 21 periods minus seven cells with a small number of observations and minus 38 cells for which the cohort was not yet in existence at the time of the survey). The minimum sample size was 224 observations in 2016 in the male subsample of the 1930–1934 cohort.

The final sample comprised 2,607,832 individual observations[3]. The Sample Survey of Household Budgets is representative of the population of the Russian Federation with KVZV weights.

2 The weighting coefficient KVZV represents the weight assigned to a household in the sample. It is included in the datasets. The weighted total number of respondents equals the population estimate of 140–146 million.

3 Overall, 1.63 per cent of individuals from the 21 waves of the Sample Survey of Household Budgets were excluded from the analysis: 0.03 per cent when adjusting for outliers in the area-per-room ratio; 1.46 per cent when restricting the lower bound of the cohort range to individuals born in 1930 (38,822 individuals aged 73 and over in 2003); and 0.14 per cent when excluding seven cells with a small number of observations.

The APC model and its identification

The model is specified at the level of pseudo-panel cohort cells:

$$\ln\left(\bar{y}_{ct}\right) = \mu + \alpha_a + \pi_t + \gamma_c + \varepsilon_{ct}, \quad (1)$$

where $\bar{y}_{ct}$ – weighted mean per-capita floor area within a cohort cell c in the survey year t; μ – intercept; α_a – age effect ($a = t - c$); π_t – survey year effect; γ_c – cohort effect; ε_{ct} – random error. Age, survey year, and cohort were coded using effect coding. Consequently, the sum of the coefficients within each of the three blocks is zero. Logarithmic transformation stabilised the variance, enabling the coefficients to be interpreted as percentage changes.

Estimation was carried out using the weighted ordinary least squares (OLS) method. Weights (n_{hh}) were assigned based on the number of unique households in a cell, whereby several household members falling within the same cell (peers) were counted as a single observation. This weight was chosen because the per-capita floor area is identical for all members of the same household. Rosstat's household sample weights (KVZV) were applied to the cell at the microdata aggregation stage, and n_{hh} was used at the weighted OLS regression stage.

To address the identification problem, the Intrinsic Estimator was applied in its canonical parametrisation (Yang and Land 2013). Age, survey year, and cohort were coded using effect coding (sum-to-zero coding with an intercept), whereby the effects of each block were summed to zero by construction, and the null space of the design matrix was one-dimensional. The solution with the minimum Euclidean norm was given by the Moore–Penrose pseudoinverse. Unlike the Deaton–Paxson approach, the IE distributes the linear trend across all three sets of coefficients. Age coefficients and cohort effects describe deviations from the macro-trend of improving housing conditions.

Shift-share decomposition

By default, the IE effects are expressed as deviations from the mean (mean = 0). To facilitate interpretation, the effects for each survey year were reparametrised with respect to the base year of 2003: $\pi'_t = \pi_t - \pi_{2003}$, which alters neither the model fit (R^2) nor the estimated age and cohort effects: the shift is absorbed by the intercept. A shift-share decomposition was applied to break down the observed change in mean provision into components attributable to growth within cohorts and changes in cohorts (Kitagawa 1955):

$$\triangle Y = \Sigma_c \triangle w_c \cdot \bar{y}_c + \Sigma_c \bar{w}_c \cdot \triangle y_c, \quad (2)$$

where w_c – proportion of cohort c in the population (weighted by KVZV), y_c – mean housing provision for cohort c, Y – mean housing provision for the entire population (the sum of the products of the proportions and the mean values of the cohorts), Δ – difference between the final and initial years of observation, the overbar denotes the mean of the two end years. The first component is the composition effect and the second is the within-cohort effect. As both factors are symmetrically averaged, the sum of the two components is identically equal to the actual change in the mean. The decomposition was calculated for two ten-year intervals (2003→2013, 2013→2023) and in total for 2003→2023 for the entire population and separately for men and women.

For the gender decomposition, the IE model was estimated independently on three samples: a combined sample of 2,225,549 unique households across all cells, a male sample of 1,109,970, and a female sample of 1,454,702. All three estimates used the same specification

(354 cells) and the same OLS weight structure (n_{hh}), which was calculated for the respective subsample and represents the number of unique households in a cell. The gender gap δ was calculated separately for age and cohort effects as the difference between the female and male IE coefficients (δ = W − M). The standard error (SE) of the difference was calculated as $\sqrt{se_M^2 + se_W^2}$ and serves as an upper-bound estimate, since the exact variance is equal to $Var(\delta) = Var(W) + Var(M) - 2 \cdot Cov(W,M)$, while the male and female estimates within the household are positively correlated (their per-capita floor area is identical). Calculating the standard error using bootstrapping clustered by household and unclustered (by individual) (500 replicates) showed that the analytical SE is approximately one and a half times greater than both estimated values (median ratio 0.63). Clustering by household affected the standard error in both directions; the median ratio of the clustered standard error to the unclustered standard error was 0.91. However, in two out of nineteen cohorts, the ratio exceeded one (1.05 and 1.08). The largest (most conservative) of the three estimates – the analytical one – was used: it exceeded the clustered estimate in all 19 cohorts, and its confidence intervals were wider. If the difference is significant according to these intervals, it is all the more significant according to either of the two bootstrap estimates.

Results

Figure 1 shows the increase in housing provision by cohort between 2003 and 2023. Those born between 1950 and 1969 (+13...+16 m²) led the way, including the 1960–1964 cohort (aged 39–43 in 2003 and 59–63 in 2023), who experienced an increase of +16.0 m² (+93.8 per cent). During the observation period, these cohorts went through the empty nest phase, when their children had grown up and left home. The younger cohorts born between 1990 and 2004 show a high relative growth (42–69 per cent). This generation started with low figures of 12.8–14.4 m² during their childhood, and by the end of the observation period they were leaving their parents' households.

Age-Period-Cohort (APC) modelling using the Intrinsic Estimator on the combined sample explains 99.1 per cent of the variation in the logarithm of the mean per-capita floor area across the pseudo-panel cells. The quality of the fit is also high for both men and women (R^2 = 0.979 and 0.990, respectively). Age is the largest source of variation (45.6 per cent for all respondents, 37.3 per cent for men, and 50.6 per cent for women; Table 1). The birth cohort accounts for 29.5%, 25.8% and 29.0% respectively, while the survey year accounts for 24.9%, 36.9% and 20.4% respectively. Among women, the age component of variation in housing provision is considerably greater, while among men, the survey year has a relatively greater impact.

Cohort effects decline from +0.20 for the 1930–1934 cohort, to −0.27 for the 2015–2019 cohort, and then to −0.22 for the 2020–2024 cohort (see Figure 2; Tables A1–A2 in Appendix for the full sets of coefficients). The total intergenerational reduction amounts to 0.47 log units (approximately 37% in terms of per-capita floor area). Members of the older cohorts (1930–1974; effect >0) lived in small households (1–2 persons) with high per-capita floor area at the time of observation. The middle cohorts (1975–1984, effect approximately equals 0) became homeowners in the 1990s and 2000s, and their level is close to the mean. Members of the younger cohorts (1985–2024, effect <0) either live in young families with children or are children themselves; their households are larger.

The housing space provision for men and women is similar across cohorts and, on the whole, shows a downward trend. In cohorts born between 1935 and 1974, women have

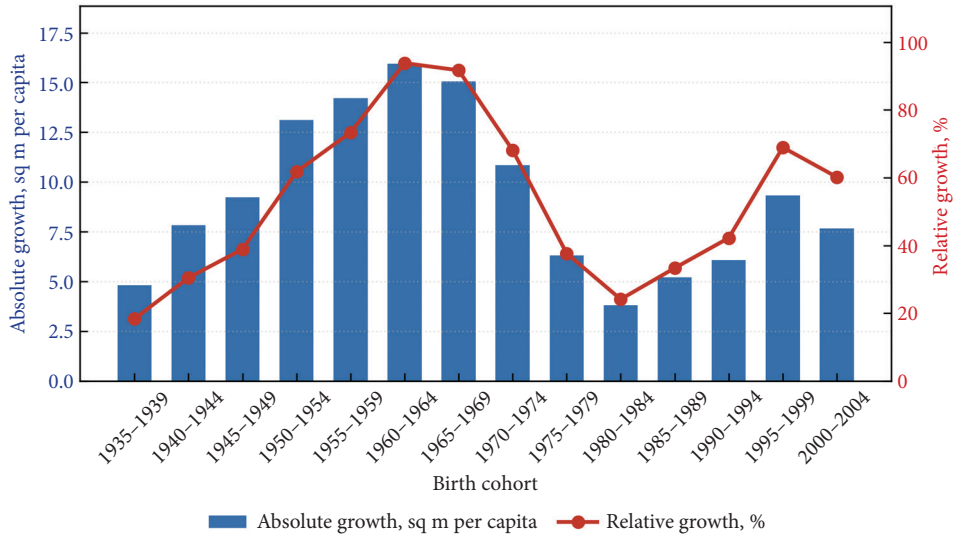

Figure 1. Increase in housing provision by five-year cohorts for 2003–2023: left axis – absolute growth, m² per capita; right axis – relative growth, %. *Source:* Author's calculations based on the Sample Survey of Household Budgets data.

Table 1. Decomposition of the explained variation in the APC model (IE): combined, men, women

Source of variation	Combined	Men	Women
Age	45.6%	37.3%	50.6%
Survey year (period)	24.9%	36.9%	20.4%
Birth cohort	29.5%	25.8%	29.0%
R^2	0.991	0.979	0.990

Source: Author's calculations.

access to better housing (this difference is statistically significant in all cohorts except those born between 1965 and 1969), which is consistent with older women being more likely to live alone or in smaller households. In cohorts born between 1975 and 2014, however, women are more likely than men to live in families with children and share their housing space with other family members. In these cohorts, men's housing provision is higher (the difference is statistically significant in all eight cohorts). In the two youngest cohorts, those born between 2015 and 2019 and between 2020 and 2024, which are observed in only 9 and 4 of the 21 survey waves, the point estimate is negative but not significantly different from zero.

The survey-year effects demonstrate a consistent increase from 0.000 in 2003 to +0.324 in 2023, indicating a +38 per cent improvement over 20 years (Figure 3). This macro-trend corresponds with growth in the housing stock from 2.9 billion to 4.2 billion m², according to Rosstat (2025), during a period of mortgage market development and government support programmes, such as the maternity capital scheme introduced in 2007 and subsidised mortgage introduced in 2020, as well as a growth in real household incomes. Slowdowns in the growth of this effect coincide with macroeconomic shocks in 2008–2009, 2014–2015, and 2020. As the survey-year effect reflects broader economic and demographic factors that

affect everyone regardless of gender, the trajectories for men and women are virtually identical. Discrepancies do not exceed 0.03 log units across all survey waves, with a maximum of 0.026 log units in 2014.

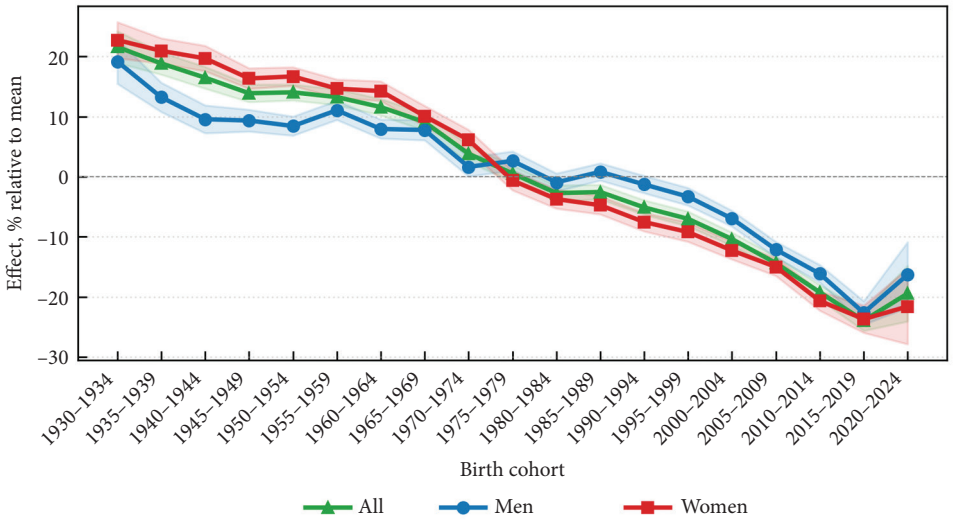

Figure 2. Cohort effect of APC decomposition (IE) as a percentage of the mean. The three lines represent the combined, men and women respectively. The shaded area represents the 95 per cent confidence interval (CI). Conversion from log to percentage: $(\exp(\beta) - 1) \times 100$. *Source:* Author's calculations.

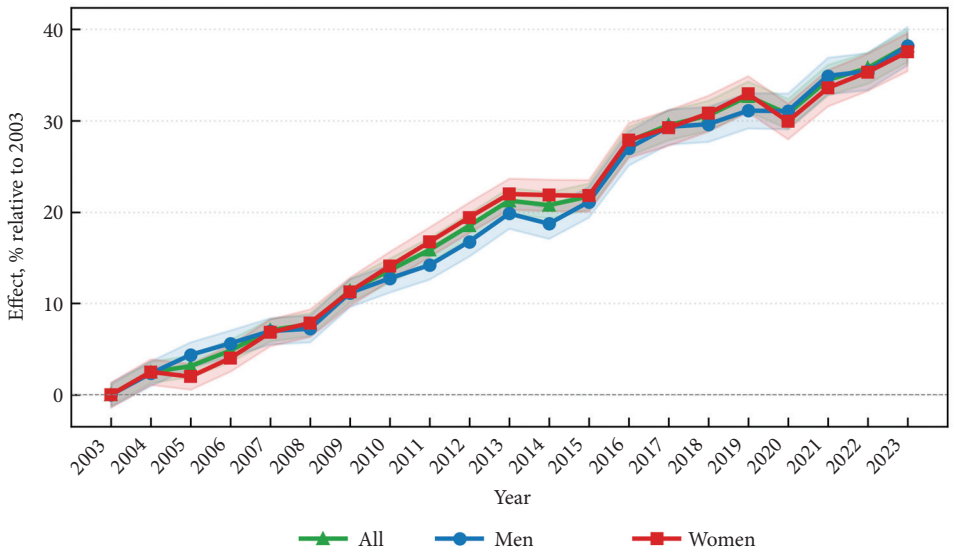

Figure 3. Survey year effect (IE), percentage compared with 2003. The three lines represent the combined, men and women respectively. The shaded area represents the 95 per cent CI. *Source:* Author's calculations.

The age profile is markedly non-linear, with a range of 0.39 log units (48%) in the combined sample (Figure 4). There is a clear four-phase structure: (1) childhood and youth, when people have the least amount of floor area as they live in larger households with their parents; (2) a dip during the family nest phase at ages 30–44, when people form their own families with children; (3) a reduction in household size, peaking at ages 65–69, when the number of residents decreases while the floor area remains the same; and (4) a decline in housing provision after the age of 75, which according to the data corresponds to an increase in the mean household size and more frequent residence in extended and other household types (among those aged 75 and over, 10.5 per cent of women compared with 5.2 per cent of men, Sample Survey of Household Budgets, 2015–2023, Table A3 in Appendix). As the survey only records household composition at the time of the survey and does not track individuals over time, moving in with relatives or downsizing to more compact accommodation remain possible explanations for this shift rather than events that can be observed in the data.

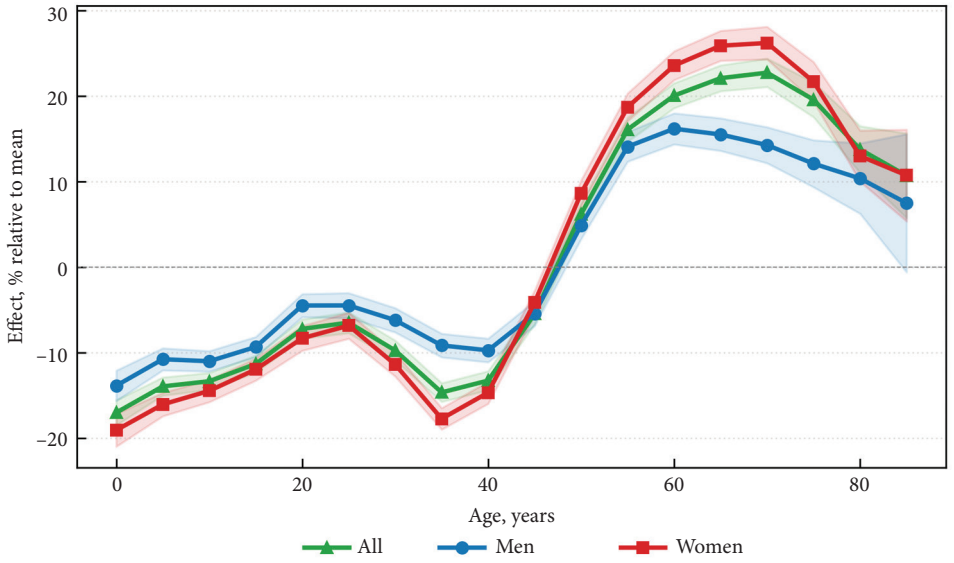

Figure 4. Age effect of APC decomposition (IE) as a percentage of the mean: three lines – combined, men, women. The shaded area represents the 95 per cent CI. *Source:* Author's calculations.

Separate estimates by gender reveal substantial differences in the amplitude of the age profile. The profile is notably steeper for women, with a dip in the family nest phase reaching −0.20 log units at 35 years of age (compared to −0.10 at 40 years of age for men). Meanwhile, the late peak occurs at 70 years of age, equalling +0.23 (compared to +0.15 at 60 years of age for men). The range between these two stages is 0.43 log units for women, compared to 0.25 for men (+54 per cent versus +28 per cent from trough to peak) – almost twice as large. The total amplitude of the profile is 0.44, compared to 0.30 log units. The mechanism behind the late peak also differs. For men, the peak coincides with the empty nest phase: the proportion of men living as a couple without children rises from 19 per cent among 45–49-year-olds to 65 per cent among 65–69-year-olds, while the proportion of men living alone in these age

groups does not exceed seven per cent. For women, the empty nest phase occurs between the ages of 55 and 64. From the age of 65 onwards, living alone becomes the most common pattern: 40 per cent of women aged 65–69 and 45 per cent of women aged 70–74 live alone, compared with 30 per cent and 24 per cent who live as a couple without children (Sample Survey of Household Budgets, 2015–2023). Consequently, the peak for women occurs a decade later than for men (Figure A1 in Appendix). Those living alone at ages 60–79 have a mean per-capita floor area of 50.9 m², compared with 29.1 m² for childless couples. In other words, the late peak in this profile is primarily driven by households consisting of a single person. In contrast, women of reproductive age are more likely to find themselves living in cramped conditions (e.g. families with children or single mothers).

Figure 5 shows the trends in the mean per-capita floor area by age for the six cohorts, presented separately for men (dashed line) and women (solid line). The trajectories for women lie above those for men in the older cohorts and converge in the younger ones. The mean gap is +6.9 and +6.5 m² for the 1940–1944 and 1950–1954 cohorts respectively, +4.4 for the 1960–1964 cohort, +2.2 for the 1970–1974 cohort, close to zero for the 1980–1984 cohort, and −0.3 m² for the 1990–1994 cohort. That is to say, in the younger generations, the sign of the gap reverses, as it does for the gender gap in cohort effects (see Figure 7). All cohorts demonstrate an increase in housing provision with age, with this growth accelerating during retirement years.

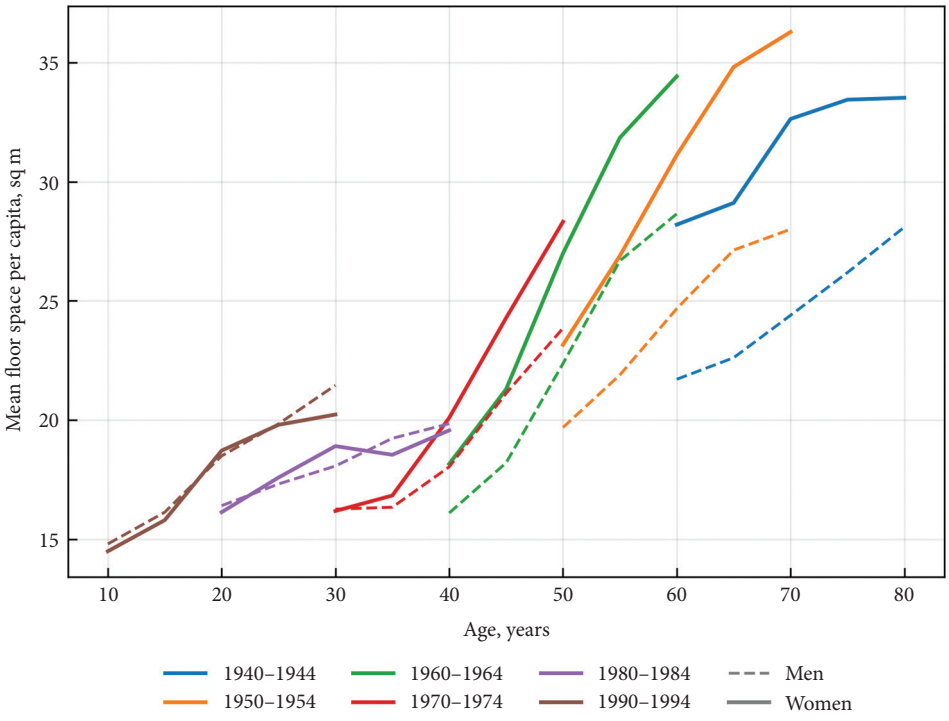

Figure 5. Trajectories for six selected cohorts by gender: mean per-capita floor area, m² (solid line – women, dashed line – men). *Source:* Author's calculations.

Between 2003 and 2023, the mean per-capita floor area across the 14 cohorts observed at the start and end of the twenty-year period increased by 7.56 m² (see Table 2). However, as the calculation only includes these 14 cohorts, the youngest and worst-housed generations are excluded. Consequently, the result of the decomposition exceeds the increase in the mean floor area calculated for the entire sample. This growth can be broken down into two components that move in opposite directions. The within-cohort effect amounts to +9.35 m² (124 per cent of the total change): each cohort substantially improved its housing conditions. The greatest contribution came from the cohorts born between 1940 and 1974, who had passed through the peak of their life cycle. The composition effect was −1.79 m² (−24 per cent of the total change): better-housed older cohorts (born 1930–1954, with per-capita floor areas of 25–35 m²) were replaced by less well-off younger cohorts (born 1990–2004, with per-capita floor areas of 14–20 m²). Due to the natural decrease in older cohorts and growth in smaller cohorts from the 1990s demographic trough, the aggregate indicator is slowing down.

A separate shift-share decomposition revealed a common pattern for both sexes: within-cohort growth dominates structural shift. For men, the total increase is 7.00 m² (+8.12/−1.12), and for women it is 7.95 m² (+10.10/−2.14). Within-cohort growth is more pronounced among women, but the negative structural effect is also stronger: older women who are leaving the sample are characterised by a particularly high per-capita floor area (they are more likely to live alone), while young women entering the sample have a smaller floor area. A breakdown by the decades 2003–2013 and 2013–2023 shows a slowdown in progress. For women, the contribution of within-cohort growth fell from +5.84 m² in the first decade to +4.36 in the second. For men, it fell from +4.29 to +3.87. As the decomposition is calculated for cohorts observed at both ends of each decade (15 and 16 cohorts compared to 14 for 2003–2023), the components of the sub-periods do not add up to the total for the entire period.

The mean per-capita floor area for women was higher than that for men in every year of the observation period: 19.35 m² versus 17.07 m² in 2003 (+2.28 m²) and 24.88 m² versus 22.05 m² in 2023 (+2.83 m²). To analyse the gender gap, the differences in age- and cohort-specific IE effects were calculated.

Table 2. Shift-share decomposition of change in mean housing provision: composition and within-cohort components, m² per person

Period	Sex	Total	Composition	Within-cohort
2003–2023	Combined	+7.56	−1.79	+9.35
2003–2023	Men	+7.00	−1.12	+8.12
2003–2023	Women	+7.95	−2.14	+10.10
2003–2013	Combined	+4.03	−1.21	+5.23
2003–2013	Men	+3.37	−0.92	+4.29
2003–2013	Women	+4.53	−1.31	+5.84
2013–2023	Combined	+2.78	−1.42	+4.20
2013–2023	Men	+2.90	−0.97	+3.87
2013–2023	Women	+2.69	−1.67	+4.36

Source: Author's calculations. *Notes:* The difference between the sum of the components and the Total reflects rounding and does not exceed 0.01 m².

The gender gap in age effects is significant in most age groups and non-monotonic (see Figure 6). Up to the age of 40, women exhibit a lower age effect of between −0.02 and −0.10 log units. At these ages, the majority of women are married and live with their husband and children. Between the ages of 45 and 75, however, the gap reverses to reach +0.10 at age 70, reflecting the effects of widowhood and singlehood among older women. For those aged 80 and over, the confidence interval widens significantly due to a reduction in the number of observations, and the gender gap becomes indistinguishable from zero.

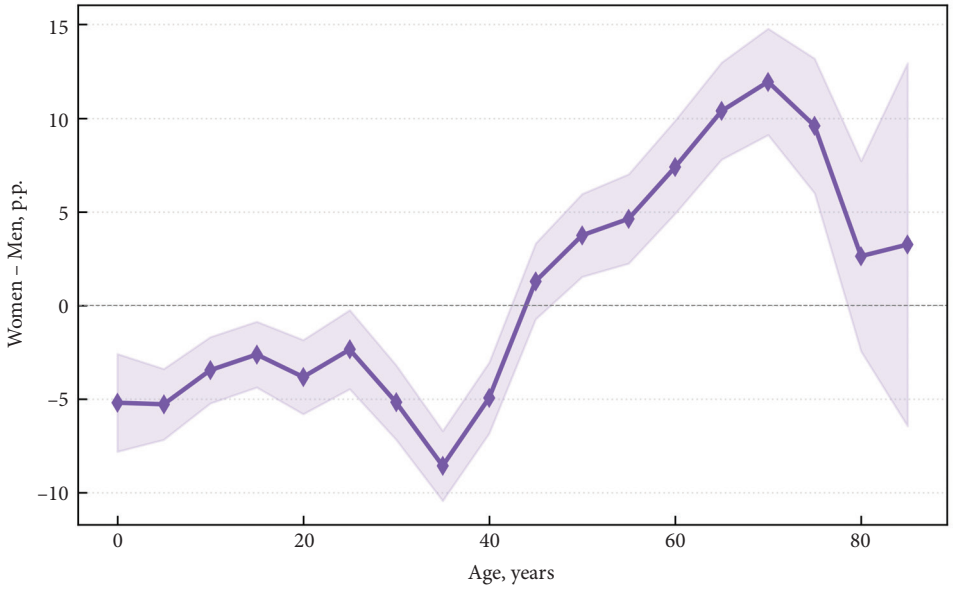

Figure 6. Gender gap (W – M) in the age effect, percentage points. The shaded area represents the 95 per cent CI. *Source:* Author's calculations.

Figure 7 shows that the gender gap across cohorts demonstrates a generational shift. Among those born between 1935 and 1974, women exhibited higher cohort effects (between +0.02 and +0.09 log units). In this group, widows predominate: the proportion of widows among women aged 65 and over is 56.9 per cent, compared with 13.3 per cent among men (Sample Survey of Household Budgets, 2015–2023). Their higher per-capita floor area comes with a small household size. However, housing history is not recorded in the Sample Survey of Household Budgets, so a woman retaining her home following her spouse's death is merely an interpretation of this situation rather than a measured fact. From the 1975–1979 cohorts onwards, the sign of the gap turns negative – initially slightly (at −0.03 for the 1975–1984 cohorts), then steadily (at −0.06 to −0.07 for the 1985–2004 cohorts). Young women are more likely than men to live in households with young children, where the floor area is shared among several family members. The negative gap is statistically significant for cohorts up to and including 2010–2014. However, for the youngest generations, the data do not confirm a male advantage. For the 2015–2019 and 2020–2024 cohorts, the confidence interval includes zero: −0.01 (−0.05; +0.03) and −0.07 (−0.17; +0.04) log units.

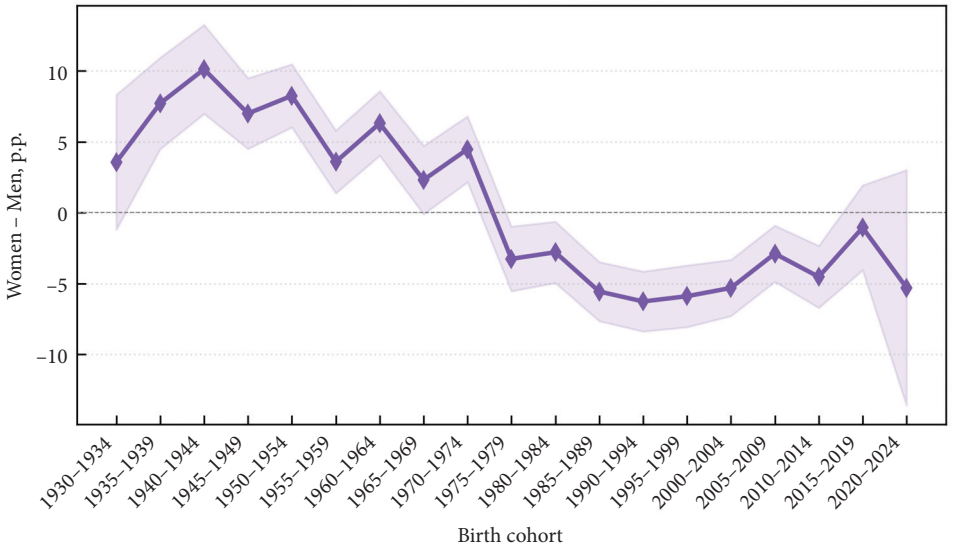

Figure 7. Gender gap (W – M) in the cohort effect, percentage points. The shaded area represents the 95 per cent CI. *Source:* Author's calculations.

Robustness of results

Robustness was evaluated in four distinct ways. Firstly, an alternative Deaton–Paxson APC identification was implemented in which the linear trend relates to age and cohort rather than the survey year. Secondly, an alternative parametrisation of the same minimum-norm IE solution was carried out with projection onto the space of the full set of dummy variables (indicator coding). Thirdly, the reference model was re-estimated using subsamples of people aged 18 and over and 25 and over (excluding children). Fourthly, a cluster bootstrap analysis was performed at the household level (500 replicates) to estimate the standard errors of the gender gap. The findings are presented in Table A4 in Appendix.

Both the variance decomposition and the direction of the intergenerational gradient are sensitive to the choice of restriction. In the Deaton–Paxson identification, the proportion attributable to age rises to 78.5 per cent and the proportion attributable to cohort falls to 21.2 per cent, while the sign of the cohort gradient reverses in this model ($r = -0.96$ with IE estimates). By contrast, in the indicator coding, the cohort accounts for 62.1 per cent and age for 11.2 per cent. Therefore, the declining cohort profile (Figure 2) should be interpreted as the cohort's position relative to the overall trend rather than as absolute deterioration in conditions for younger generations.

By contrast, the gender difference in cohort and age effects is robust to the choice of the APC model. Its sign is consistent across all 19 cohorts and 18 age groups. The correlation is 0.999 when the identification method is changed, and 0.9999 when the parametrisation is changed. Excluding children steepens the cohort gradient without altering its shape (the correlation with the reference specification is 0.999 for the 18+ subsample and 0.994 for the 25+ subsample). Cluster bootstrap analysis shows that, on average, accounting for intra-household dependence does not increase standard errors: the median ratio of cluster SE

to non-cluster SE is 0.91, ranging from 0.76 to 1.08. In two cohorts (2015–2019 and 2020–2024), the ratio exceeds one. The bootstrap estimates are lower than the analytical estimate in all cohorts; therefore, the confidence intervals used in the paper remain conservative.

Discussion of the findings

The key finding of this study is the gender difference in the cohort and age effects on housing provision. This difference remains consistent regardless of the identification method or parametrisation chosen. The declining cohort profile obtained in the reference IE specification of the APC model (from +0.20 to −0.27) does not imply that younger generations are worse off in absolute terms. Under this specification, the effects represent a cohort's position relative to the overall trend while the survey-year effect (+38 per cent over 2003–2023) captures the bulk of the linear trend of macro-level improvement. The age profile (a dip between 30 and 44 years, peaking at 65–69 years) is consistent with that observed in Germany (Hein and Kuhnimhof 2026), where a similar dip is seen between the ages of 30 and 40. The decline after the age of 75 (reaching +0.10 at age 85) is accompanied by an increase in the mean household size compared to the 60–74 age group (see Table A3 in Appendix). Living with relatives is the most plausible explanation for this. However, repeated cross-sectional surveys under the Sample Survey of Household Budgets do not contain information on residential relocation. This pattern contrasts with German data, where households are solely represented by the head of the household and housing provision remains stable in older age groups.

The consistently higher level of housing provision for women can be attributed mainly to the demographic composition of households rather than to differences in individual property rights. In the Sample Survey of Household Budgets, housing adequacy is measured at the household level, while individual per-capita floor area is an aggregate figure that is the same for all family members. The distribution of men and women across household types is asymmetrical, as calculated from the Sample Survey of Household Budgets between 2015 and 2023. Among those aged 65 and over, 41.9 per cent of women live alone compared to 8.6 per cent of men. Meanwhile, men are more likely to remain in complete married households (see Table A3 in Appendix).

This finding is consistent with the results of a study by Vignoli et al. (2016), which found that the disparity in home ownership among older people in Europe is almost entirely explained by household type: women are more likely to live alone or in households without a partner. In the present study, this mechanism manifests as a positive age effect gap among those aged 60–75 (+0.06...+0.10 log units). Kapelle and Van Winkle (2024) found that housing wealth decreased following widowhood in five out of eleven European countries. In our data, the advantage for women disappears at the oldest ages: the gap narrows from +0.08 at age 75 to +0.02 at age 80.

The design of our study follows the methodology of Hein and Kuhnimhof (2026). However, whereas their study defines the cohort by the head of the household, ours defines it individually for each household member. This allows us to consider the gender composition of households.

Shift-share decomposition reveals that the aggregate indicator, mean per-capita floor area, obscures actual trends. A policy focused on the aggregate may underestimate real progress within cohorts and structural challenges resulting from generational replacement. Among women, the increase in housing provision and the negative structural shift are both greater than among men. The gender-specific nature of the shift-share decomposition re-

flects demographic patterns: women are more likely to live alone in older age groups and older cohorts have a higher baseline level.

This study has several limitations. Firstly, the Sample Survey of Household Budgets measures floor area at the household level, meaning the per-capita figure is an aggregate figure identical for all family members. Gender differences arise from differences in the demographic composition of households rather than from differences in individual housing rights, on which the Sample Survey of Household Budgets provides no information. Furthermore, apart from the composition and size of the household at the time of the survey, no information is available on transitions between states, such as widowhood, relocation to live with relatives, or a change of residence. This study offers explanations for these transitions as hypotheses that are consistent with the observed household composition. Direct verification of these hypotheses requires panel data. Secondly, the cohorts born between 2005 and 2024 consist predominantly of children; therefore, their cohort effect does not reflect their own housing-related decisions. Testing on samples of adults aged 18+ and 25+ retains the shape of the cohort profile ($r = 0.999$ and 0.994, respectively, for the combined sample), while the gradient becomes steeper (see Table A4 in Appendix). Thirdly, in the IE model the solution is selected by the minimum Euclidean norm criterion. An alternative identification method (Deaton–Paxson) yields different estimates and the opposite sign for the intergenerational gradient. Therefore, APC effect levels are conditional – only the gender difference remains robust to this choice.

Our findings reveal a specific pattern of housing vulnerability in Russia: while the mean floor area per elderly woman is nominally high, this does not represent housing adequacy. This group is characterised by the combination of having a large dwelling and low income, which may entail high maintenance costs, limited mobility, and financial stress; moreover, life satisfaction among older people living alone in Russia is closely linked to economic conditions (Voronin et al. 2018). Policy instruments – such as programmes that support older people in downsizing, utility-bill subsidies for single older women, and improved access to rental accommodation – must take these characteristics into account. While the younger cohorts born in the 1990s (the demographic trough) remain the target group for preferential mortgage schemes and maternity capital programmes (Vakulenko and Evgrafova 2025), the aggregate data do not show a substantial improvement in their housing provision in recent years. The purchase of a separate home and the birth of children may have formally and temporarily reduced their per-capita housing provision.

Conclusion

This study is the first to apply the APC approach with Intrinsic Estimator identification to housing provision of the Russian population. Microdata from the 2003–2023 Sample Survey of Household Budgets were used to make separate estimates of the age–period–cohort effects for men and women. The following findings were obtained:

1. In the canonical IE identification, age accounts for the largest component of variance in housing provision (approximately 46 per cent across the entire sample, 37 per cent for men and 51 per cent for women). Birth cohort accounts for approximately 29 per cent, and the survey year for approximately 25 per cent. The variance decomposition across components and the direction of the intergenerational gradient depend on the restriction chosen and are not identifiable from the data. Older cohorts (1930–1974) exhibit a positive cohort effect, while younger cohorts (1985–2024) exhibit a negative one. The range is from +0.20 to −0.27 log units.

2. The survey year effect (approximately 25% of the total variance) represents the main macro-trend: a steady increase in the per-capita floor area of 38% between 2003 and 2023. Gender differences in the survey-year effects are negligible (a maximum of 0.03 log units).

3. The age effect (approximately 46% of the total variance) follows a distinctly non-linear pattern: a trough between the ages of 30 and 44, a late peak between 65 and 69, and a subsequent decline towards the end of life. In women, the range between these stages is almost double that in men (0.43 versus 0.25 log units), and the peak occurs a decade later. In men, it coincides with the empty nest phenomenon, while in women it coincides with transitioning to living alone.

4. The increase in the mean housing provision of +7.56 m^2 over 20 years is entirely due to improvements within cohorts (+9.35 m^2), which are partly offset by a compositional shift resulting from generational replacement (−1.79 m^2). Both of these components are more pronounced among women (see Table 2).

5. As seen from Figure 1, the cohorts born between 1950 and 1969 experienced the greatest increase in housing provision (+13...+16 m^2).

6. Women consistently have a higher level of housing provision than men, with the gap between the two increasing from +2.28 m^2 in 2003 to +2.83 m^2 in 2023. Among cohorts born between 1975 and 2014, the gap significantly changes sign (−0.03…−0.07 log units), with young men enjoying a higher level of housing provision than their female peers. However, in the two youngest cohorts (2015–2019 and 2020–2024), the gap is non-significant, and the conclusion regarding the change in sign does not apply.

7. Gender differences primarily represent the different housing conditions experienced by men and women in households of various compositions, rather than differences in their individual housing rights or property resources. Housing policy measures must take these specific circumstances into account, particularly with regard to older women living alone.

References

Bessière C, Gollac S (2023) The gender of capital: How families perpetuate wealth inequality. Harvard University Press, Cambridge, MA, 344 pp.

Bobkov VN, Odintsova EV (2021) Material'noe blagosostoyanie rossiyan: mezhpokolennaya differentsiatsiya [The material well-being of Russians: Intergenerational differentiation]. Mir novoi ekonomiki [The World of New Economy] 15 (2): 16–28. https://doi.org/10.26794/2220-6469-2021-15-2-16-28 [In Russian]

Bryukhanov M, Hryshko D (2024) Inequality in Russia over time and over the life cycle. Scandinavian Journal of Economics 126 (2): 289–319. https://doi.org/10.1111/sjoe.12556

Burdyak AYa (2015) Obespechennost' zhil'em v postsovetskoi Rossii: neravenstvo i problema pokolenii [Housing provision in post-Soviet Russia: inequality and the generational problem]. Zhurnal issledovanii sotsial'noi politiki [The Journal of Social Policy Studies] 13 (2): 273–288. https://jsps.hse.ru/article/view/3330 [In Russian]

Cherkashina TYu (2025) Neravenstvo rossiiskikh domokhozyaistv po zhilishchnomu bogatstvu: uroven' i dinamika [Inequality of Russian households in housing wealth: Level and dynamics]. Sotsiologicheskii zhurnal [Sociological Journal] 31 (2): 51–78. https://doi.org/10.19181/socjour.2025.31.2.3 [In Russian]

Davis CM (2018) The changing capabilities of cohorts of the elderly in Russia during 1990–2020: Measurement using a quantitative index. Journal of Population Ageing 11 (2): 153–208. https://doi.org/10.1007/s12062-017-9179-1

Deaton A (1985) Panel data from time series of cross-sections. Journal of Econometrics 30 (1–2): 109–126. https://doi.org/10.1016/0304-4076(85)90134-4

Deaton A, Paxson C (1994) Intertemporal choice and inequality. Journal of Political Economy 102 (3): 437–467. https://doi.org/10.1086/261941

Fosse E, Winship C (2019) Analyzing age–period–cohort data: A review and critique. Annual Review of Sociology 45: 467–492. https://doi.org/10.1146/annurev-soc-073018-022616

Fu Q (2016) The persistence of power despite the changing meaning of homeownership: An age-period-cohort analysis of urban housing tenure in China, 1989–2011. Urban Studies 53 (6): 1225–1243. https://doi.org/10.1177/0042098015571240

Gerber TP, Gimpelson V (2024) Inequality and social stratification in Russia during the Putin regime: From market transition to war[4] on Ukraine. Sociology Compass 18 (3): e13196. https://doi.org/10.1111/soc4.13196

Gerber TP, Zavisca JR, Wang J (2022) Market and nonmarket pathways to home ownership and social stratification in hybrid housing regimes: Evidence from four post-Soviet countries. American Journal of Sociology 128 (3): 866–913. https://doi.org/10.1086/722927

Goldsmith-Pinkham P, Shue K (2023) The gender gap in housing returns. The Journal of Finance 78 (2): 1097–1145. https://doi.org/10.1111/jofi.13212

Guzanova AK (2025) Zhilishchnaya obespechennost' domokhozyaistv s det'mi v sovremennoi Rossii [Housing provision for households with children in modern Russia]. Uroven' zhizni naseleniya regionov Rossii [Living Standards of the Population in the Regions of Russia] 21 (1): 43–58. https://doi.org/10.52180/1999-9836_2025_21_1_3_43_58 [In Russian]

Hein S, Kuhnimhof T (2026) Unravelling cohort effects in consumption of living space by German households. Housing Studies 41 (3): 569–597. https://doi.org/10.1080/02673037.2025.2473611

Kalyukin A, Kohl S (2020) Continuities and discontinuities of Russian urban housing: The Soviet housing experiment in historical long-term perspective. Urban Studies 57 (8): 1768–1785. https://doi.org/10.1177/0042098019852326

Kapelle N, Van Winkle Z (2024) Changes in household wealth over the process of widowhood across European countries. The Journals of Gerontology: Series B 79 (10): gbae116. https://doi.org/10.1093/geronb/gbae116

Kitagawa EM (1955) Components of a difference between two rates. Journal of the American Statistical Association 50 (272): 1168–1194. https://doi.org/10.1080/01621459.1955.10501299

Kosareva NB, Polidi TD (2019) Dostupnost' zhil'ya v Rossii i za rubezhom [Housing affordability in Russia and foreign countries]. Voprosy Ekonomiki (7): 29–51. https://doi.org/10.32609/0042-8736-2019-7-29-51 [In Russian]

Kosareva N, Polidi T (2021) Housing affordability in Russia. Housing Policy Debate 31 (2): 214–238. https://doi.org/10.1080/10511482.2020.1800778

Kosareva NB, Polidi TD, Puzanov AS (2015) Zhilishchnaya politika i ekonomika v Rossii: rezul'taty i strategiya razvitiya [Housing policy and economics in Russia: results and development strategy]. HSE Publishing House, Moscow, 387 pp. https://urbaneconomics.ru/sites/default/files/ieg_zhilishchnaya_politika_i_ekonomika_v_rossii.pdf [In Russian]

Lersch PM, Dewilde C (2015) Employment insecurity and first-time homeownership: Evidence from twenty-two European countries. Environment and Planning A 47 (3): 607–624. https://doi.org/10.1068/a130358p

Luo L (2013) Assessing validity and application scope of the intrinsic estimator approach to the age-period-cohort problem. Demography 50 (6): 1945–1967. https://doi.org/10.1007/s13524-013-0243-z

Maksimov MA (2025) Regional'nye razlichiya smertnosti v Rossii: kogortnye effekty [Regional mortality differences in Russia: Cohort-specific effects]. Voprosy Statistiki 32 (3): 62–80. https://doi.org/10.34023/2313-6383-2025-32-3-62-80 [In Russian]

4 Special military operation.

Myers D, Lee H, Park J (2025) Misalignment of housing growth and population trends: Cohort size and lagging measurements through recession and recovery. RSF: The Russell Sage Foundation Journal of the Social Sciences 11 (1): 86–109. https://doi.org/10.7758/RSF.2025.11.1.05

Nieuwenhuis R, Zagel H (2023) Housing conditions of single mothers in Europe: The role of housing policies. European Societies 25 (2): 181–207. https://doi.org/10.1080/14616696.2022.2117835

Pelzer B, te Grotenhuis M, Eisinga R, Schmidt-Catran AW (2015) The non-uniqueness property of the intrinsic estimator in APC models. Demography 52 (1): 315–327. https://doi.org/10.1007/s13524-014-0360-3

Pilipenko IV (2025) Zhilishchnoe stroitel'stvo v Rossii za 100 let: dinamika, itogi i sotsial'no-ekonomicheskie problemy [Housing construction in Russia during the 100-year period: Dynamics, results and socio-economic problems]. Voprosy Ekonomiki (1): 134–158. https://doi.org/10.32609/0042-8736-2025-1-134-158 [In Russian]

Prokofieva LM, Korchagina II (2023) Demograficheskaya struktura semei i domokhozyaistv v Rossii, ee dinamika po dannym perepisei naseleniya [The demographic structure of families and households in Russia, its dynamics according to population censuses]. Demograficheskoe obozrenie [Demographic Review] 10 (2): 4–17. https://doi.org/10.17323/demreview.v10i2.17763 [In Russian]

Pustovalov DN (2015) Rossiiskaya spetsifika izbytochnoi smertnosti v molodykh trudosposobnykh vozrastakh [Russian peculiarities of excess mortality among young adults]. Vestnik Moskovskogo universiteta. Seriya 6. Ekonomika [Moscow University Economics Bulletin] (1): 87–107. https://www.econ.msu.ru/sys/raw.php?o=20472&p=attachment [In Russian]

Radaev V, Roshchina Y (2019) Young cohorts of Russians drink less: age–period–cohort modelling of alcohol use prevalence 1994–2016. Addiction 114 (5): 823–835. https://doi.org/10.1111/add.14535

Roshchina I, Ilyunkina N (2021) Impact of government measures to support mortgage lending on housing affordability in Russia: Regional evidence. Russian Journal of Money and Finance 80 (4): 98–123. https://doi.org/10.31477/rjmf.202104.98

Rosstat (2025) Zhilishchnoe khozyaistvo v Rossii. 2025 [Housing economy in Russia. 2025: statistical yearbook]. Rosstat, Moscow, 53 pp. https://rosstat.gov.ru/folder/210/document/13234 [In Russian]

te Grotenhuis M, Pelzer B, Luo L, Schmidt-Catran AW (2016) The intrinsic estimator, alternative estimates, and predictions of mortality trends: a comment on Masters, Hummer, Powers, Beck, Lin, and Finch. Demography 53 (4): 1245–1252. https://doi.org/10.1007/s13524-016-0476-8

Vakulenko ES, Evgrafova RYu (2025) Ipoteka i reproduktivnye namereniya: stimul ili pregrada? [Mortgage and reproductive intentions: A boost or a barrier?]. Voprosy Ekonomiki (6): 137–153. https://doi.org/10.32609/0042-8736-2025-6-137-153 [In Russian]

Vakulenko ES, Ryzhkina AE (2024) APC-modelirovanie svyazi koeffitsientov rozhdaemosti po ocherednosti rozhdeniya s ekonomicheskimi shokami v Rossii [APC modeling of the relationship between birth order fertility rates and economic shocks in Russia]. Zhurnal Novoi ekonomicheskoi assotsiatsii [Journal of the New Economic Association] 3 (64): 84–102. https://doi.org/10.31737/22212264_2024_3_84-102 [In Russian]

Verbeek M, Nijman T (1992) Can cohort data be treated as genuine panel data? Empirical Economics 17: 9–23. https://doi.org/10.1007/BF01192471

Vignoli D, Tanturri ML, Acciai F (2016) Home bitter home? Gender, living arrangements, and the exclusion from homeownership among older Europeans. Genus 72: 9. https://doi.org/10.1186/s41118-016-0014-y

Voronin GL, Zakharov VYa, Kozyreva PM (2018) Odinokie pozhilye: dozhivayut ili aktivno zhivut? [Lonely old aged: surviving or living an active life?]. Sotsiologicheskii zhurnal [Sociological Journal] 24 (3): 32–55. https://doi.org/10.19181/socjour.2018.24.3.5992 [In Russian]

Yang Y, Land KC (2013) Age-period-cohort analysis: New models, methods, and empirical applications. Chapman & Hall/CRC, Boca Raton, 338 pp. https://doi.org/10.1201/b13902

Zavisca JR (2012) Housing the new Russia. Cornell University Press, Ithaca, NY, 264 pp. https://doi.org/10.7591/cornell/9780801450372.001.0001

Zhirnov GA (2025) Massovaya l'gotnaya ipoteka: prodlevat' nel'zya zavershat' [Mass subsidized mortgage: No extension, no termination?]. Voprosy Ekonomiki (1): 115–133. https://doi.org/10.32609/0042-8736-2025-1-115-133 [In Russian]

Funding

The article was written on the basis of the RANEPA state assignment research programme.

Information about the author

■ Aleksandra Yaroslavovna Burdyak – Senior Researcher. Center "Institute of Social Analysis and Forecasting", Institute of Applied Economic Research, Russian Presidential Academy of National Economy and Public Administration. Moscow, 119034, Russia. Email: burdyak-ay@ranepa.ru

Appendix

Table A1. Cohort effects (IE), log units

Cohort	Combined	M	W	W – M
1930–1934	+0.20	+0.18	+0.20	+0.03
1935–1939	+0.17	+0.12	+0.19	+0.07
1940–1944	+0.15	+0.09	+0.18	+0.09
1945–1949	+0.13	+0.09	+0.15	+0.06
1950–1954	+0.13	+0.08	+0.15	+0.07
1955–1959	+0.12	+0.10	+0.14	+0.03
1960–1964	+0.11	+0.08	+0.13	+0.06
1965–1969	+0.09	+0.07	+0.10	+0.02
1970–1974	+0.04	+0.02	+0.06	+0.04
1975–1979	+0.01	+0.03	−0.01	−0.03
1980–1984	−0.03	−0.01	−0.04	−0.03
1985–1989	−0.03	+0.01	−0.05	−0.06
1990–1994	−0.05	−0.01	−0.08	−0.07
1995–1999	−0.07	−0.03	−0.10	−0.06
2000–2004	−0.11	−0.07	−0.13	−0.06
2005–2009	−0.16	−0.13	−0.16	−0.03
2010–2014	−0.21	−0.18	−0.23	−0.06
2015–2019	−0.27	−0.26	−0.27	−0.01
2020–2024	−0.22	−0.18	−0.24	−0.07

Source: Author's calculations. *Notes:* The SE of the IE coefficients is ≤0.03 log units. The W – M column is calculated using unrounded coefficients and may therefore differ by 0.01 from the difference between the rounded values in the W and M columns. According to a conservative analytical SE estimate, the difference between W and M is significant at the 5 per cent level in all cohorts except 1930–1934, 1965–1969, 2015–2019, and 2020–2024. According to cluster bootstrap analysis, however, only the last two are non-significant (see Table A4 in Appendix).

Table A2. Age effects (IE), log units. Selected ages. The W–M difference is significant at the 5 per cent level in all age groups except 45, 80, and 85 years old

Age	Combined	M	W	W – M
0	−0.19	−0.15	−0.21	−0.06
20	−0.08	−0.05	−0.09	−0.04
30	−0.10	−0.06	−0.12	−0.06
35	−0.16	−0.10	−0.20	−0.10
40	−0.14	−0.10	−0.16	−0.06
45	−0.06	−0.06	−0.04	+0.01
50	+0.06	+0.05	+0.08	+0.04
55	+0.15	+0.13	+0.17	+0.04
60	+0.18	+0.15	+0.21	+0.06
65	+0.20	+0.14	+0.23	+0.09
70	+0.20	+0.13	+0.23	+0.10
75	+0.18	+0.11	+0.20	+0.08
80	+0.13	+0.10	+0.12	+0.02
85	+0.10	+0.07	+0.10	+0.03

Source: Author's calculations. *Notes:* Age is given as the lower limit of a five-year age group. The W – M column is calculated using unrounded coefficients and may therefore differ by 0.01 from the difference between the rounded values in the W and M columns.

Table A3. Household size and distribution by household type, and proportion of widowed individuals by sex and age group (% based on the Sample Survey of Household Budgets 2015–2023, weighted by KVZV)

Age	Sex	House-hold size	Living alone	Childless couple	Couple with children	Single-parent family	Other type	Widowed individ-uals
0–17	W	4.0	0.0	0.9	84.4	14.0	0.7	–
0–17	M	4.0	0.0	0.7	84.5	14.1	0.7	–
18–29	W	3.4	5.8	10.5	65.3	16.4	2.0	0.2
18–29	M	3.5	3.7	8.1	69.1	17.6	1.5	0.1
30–44	W	3.4	5.1	8.5	70.1	15.2	1.1	1.4
30–44	M	3.4	6.7	9.3	76.1	7.0	0.8	0.2
45–59	W	2.5	19.1	31.2	34.3	12.9	2.5	10.0
45–59	M	2.9	5.7	35.7	54.5	3.3	0.8	0.9
60–74	W	2.1	38.2	31.3	13.5	13.3	3.7	38.4
60–74	M	2.5	6.4	64.4	24.8	2.8	1.6	5.5
75+	W	2.2	41.6	18.7	13.4	15.8	10.5	76.8
75+	M	2.7	13.1	47.8	22.9	11.0	5.2	29.5

Age	Sex	House-hold size	Living alone	Childless couple	Couple with children	Single-parent family	Other type	Widowed individ-uals
65+ (infor-mational)	W	2.1	41.9	24.6	12.8	14.7	6.2	56.9
65+ (infor-mational)	M	2.6	8.6	60.1	23.4	5.3	2.6	13.3

Source: Author's calculations. *Notes:* The five household type columns are totalled per row to give 100 per cent. The type is determined by the R1V3 code composition (relationship to the head of household): 'couple' – the head's spouse is present; 'with children' – the head's child is present; 'single-parent family' – a child is present, but no spouse; 'other type' – extended and other household compositions. The 'other' proportion is a lower estimate of the prevalence of extended households. For example, an elderly mother living with her daughter's two-parent family would be classified as a 'couple with children' according to the R1V3 codes. 'Widowed individuals' – R1V5 = 3; this proportion is calculated for individuals aged 18 and over. The period 2015–2023 is shorter than the APC model's observation interval: household composition variables (R1V3, R1V5, CHDET and NPS) only appear in the Sample Survey of Household Budgets microdata from 2015 onwards. The 65+ group overlaps with the 60–74 and 75+ groups and is included here because its indicators are cited in the Results and Discussion sections.

Figure A1. Proportions of people living in one- and two-person households by age and sex (%), based on the Sample Survey of Household Budgets (2003–2023), weighted by KVZV. *Source:* Author's calculations.

Table A4. Robustness checks for the APC estimates: explained variance and consistency with the baseline specification

Specification	Shares of variance, % (age / survey year / cohort)	Consistency with the baseline specification
Baseline: IE, effect coding	45.6 / 24.9 / 29.5	–
Deaton–Paxson identification	78.5 / 0.3 / 21.2	cohort effects; r = −0.96; gender difference: sign for 19/19 cohorts and 18/18 ages, r = 0.9986
Indicator coding	11.2 / 26.7 / 62.1	cohort effects; r = 0.996; gender difference: sign for 19/19 cohorts and 18/18 ages, r = 0.9999
18+ subsample	45.2 / 32.5 / 22.3	cohort effects; r = 0.999
25+ subsample	40.6 / 35.5 / 23.9	cohort effects; r = 0.994
Cluster bootstrap, 500 replicates	–	gender difference SE: clustered to non-clustered – median 0.91 (0.76–1.08), ratio > 1 in 2 cohorts out of 19; both bootstrap estimates are below the analytical estimate in all cohorts

Source: Author's calculations. *Notes:* The shares of variance are given for the combined sample. The fits for the baseline specification, the Deaton–Paxson identification and the indicator coding are identical ($R^2 = 0.991$). The only differences lie in the distribution of the linear trend across the components and consequently in the proportions of variance. The significance of the gender difference is as follows: according to the conservative analytical SE, the gap is significant at the 5 per cent level in all cohorts except 1930–1934, 1965–1969, 2015–2019, and 2020–2024, and in all age groups except 45, 80, and 85 years. According to the cluster bootstrap analysis, only the 2015–2019 and 2020–2024 cohorts remain non-significant.