How does retirement affect healthcare expenditures? Evidence from a change in the retirement age

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Abstract

Using individual-level administrative panel data from Hungary, we estimate causal effects of retirement on outpatient and inpatient health care expenditures, and pharmaceutical expenditures. Our identification strategy is based on an increase in the official early retirement age of women, using that the majority of women retire upon reaching that age.

According to our descriptive results, people who are working before the early retirement age have substantially lower health care expenditures than non-workers, but the expenditure gap declines after retirement. Our causal estimates from a two-part (hurdle) model show that the shares of women with positive outpatient care, inpatient care and pharmaceutical expenditures, respectively, decrease by 3.0, 1.4 and 1.3 percentage points in the short run due to retirement. These results are driven by the relatively healthy, by those who spent some time on sick leave and by the less educated. The effect of retirement on the size of positive health care expenditures is generally not significant.

Keywords: administrative panel data, health care expenditures, hurdle model, retirement

JEL Classification: C23, I10, I12, J26

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1 Introduction

With increasing longevity and decreasing fertility rates, governments tend to raise the retirement age in order to maintain the sustainability of pension systems (OECD [2015]). If public health care expenditures change as a result of retirement, then raising the retirement age may have indirect effects on public finances through this channel. In this paper we use individual-level administrative data from Hungary, a post-communist EU member state with approximately 10 million inhabitants, to identify the short run causal effect of retirement on outpatient and inpatient care expenditures and expenditures on prescribed medications.

Such a causal analysis requires administrative panel data on health care expenditures, and also an exogeneous element of retirement, which is the (changing) statutory early retirement age in our setting. Our baseline identification strategy exploits a pension reform that increased the early retirement age for females in Hungary. We make use of the fact that the majority of women retire at the statutory early retirement age. We also give a thorough descriptive analysis of health care expenditure patterns of workers and non-workers around the early retirement age.

In line with the selection hypothesis to non-employment based on adverse health status, we find that health care expenditures of workers are substantially lower than that of non-workers a few years before the typical retirement age, but this gap generally narrows around and after retirement. Meanwhile, according to our causal estimates from two-part (hurdle) models, retirement reduces health care expenditures on the extensive margin: the shares of women with positive outpatient care, inpatient care and prescribed pharmaceutical expenditures decrease by 1-3 percentage points in the short run after retirement. Since our focus is mostly on public health care expenditures, these results imply that retirement itself may ease the burden on the public health care budget in the short run. The effect on the size of positive health care expenditures (i.e. on the intensive margin) is statistically not significant. There is some evidence for negative effects on the extensive margin for men as well, although the chosen model has limitations because it relies on the constant early retirement age in the examined period.

Among women, the effects of retirement are driven by the healthier, by those who spent some time on sick leave, and by the less educated. These heterogeneity results are mostly in line with the previous theroretical literature on the relationship between retirement and health care use (e.g. Kuhn et al. [2015] or Bíró [2016]), but we explore an additional, previously not investigated mechanism, the role of sick leave.

To our knowledge, estimating the causal effect of retirement on health care expenditures is unique in the literature, although there are related articles on the effects on health care use or health status. Earlier empirical studies find little evidence for a significant effect of retirement on health care use (Boaz and Muller [1989], Shapiro and Roos [1982], Soghikian et al. [1991]). In contrast, Bíró [2016] finds that conditional on health, being retired implies 3-9% more outpatient visits in Europe, and the results are similar for the U.S. However, if retirement has a positive effect on health, then the total effect of retirement on health care use (without conditioning on health) may even be negative. This is documented by Eibich [2015], who finds on German data that outpatient care use decreases after retirement if health is not controlled for. We contribute to this line of the literature with focusing on health care *expenditures* rather than indicators of health care utilisation. Based on administrative data, we can capture how retirement affects the public health care budget.

Since we have information only on health care expenditures, but not on health status, and we look at short- and medium-run changes in expenditures after retirement, we cannot directly contribute to the literature on health effects (although we point out that our results do not suggest considerable negative health effects of retirement). Here we only mention that there is no consensus on the sign of the health effects. The systematic literature review by van der Heide et al. [2013] documents that retirement has a beneficial effect on mental health, while the effect on physical health is contradictory (see Coe and Zamarro [2011] and Insler [2014] for positive and Behncke [2012] for negative published effects). Similarly, there is no clear empirical evidence in the literature regarding the effect of *retirement age* on health. For instance, Hernaes et al. [2013] find no effect on mortality, while Calvo et al. [2013] show that earlier retirement has negative impact on health.

The rest of the paper is organised as follows. In section 2 we introduce the theoretical mechanisms and in section 3 we provide an overview of the Hungarian health care and pension systems. We introduce the data in section 4 and the models in section 5. Results are presented in section 6. Section 7 concludes.

2 Theoretical framework

Theoretically, retirement may influence health care use through various channels. To analyse this, Bíró [2016] presents a simple static model with Cobb-Douglas utility of consumption and health, where health is allowed to depend on retirement status and is a linear function of health care use, and the budget constraints are different while working vs. while being retired. On the one hand, retirement may decrease the optimal level of health care in this model because financial incentives to remain healthy and thus to receive higher wages disappear after retirement. On the other hand, decreasing time costs and possible reductions in the out-of-pocket costs of health care point to an increase in health care use after retirement. Changing individual health preferences may also affect health care utilisation according to the model.

Kuhn et al. [2015] develop a life-cycle framework which explicitly models why investment into health care should change after retirement. (See e.g. Bloom et al. [2014] and Galama et al. [2013] for related papers.) Their model allows health to affect morbidity, longevity, earnings and the disutility of labour, as well. Their analysis implies (see Corollary 1 in Kuhn et al. [2015]) that a deferral of the retirement age may lead to a reduction of *post*-retirement health care "if there is a strong impact of health on earnings and/or the disutility of labor," but otherwise it may lead to a reduction of *pre*-retirement health care. Hence, theoretically, the effect of retirement on health care use is ambiguous but may differ e.g. across education level. Health shocks seem to have a greater relative negative impact on earnings of low-skilled individuals (Lundborg et al. [2015]), and also the disutility of labour arising from poor health may be higher for people in physically demanding occupations (who have lower education on average). Thus the above model suggests a stronger negative impact of retirement on health care use of people with low education level than of the general population.

We consider an additional, work-related influencing mechanism, which was previously not investigated in the literature but can be regarded as a special case of the above channels. Employees have to contact their general practitioner (GP) if they are sick, so as to claim sick leave. The GP then might prescribe medications or refer the patient further to specialist care, or even to inpatient care if the condition is more serious. Indeed, as Puhani and Sonderhof

[2010] show, a reduction in sick pay reduces demand for both inpatient and outpatient care. Once retired, the option of claiming sick leave benefits no longer works as an incentive for health care utilisation. This channel suggests a stronger negative relation between retirement and health care use for those who were on sick leave in the years before retirement, especially on the extensive margin.

3 Institutional background

3.1 Health care system in Hungary

The following overview of the health care system in Hungary is based on Gaál et al. [2011]. It is a single-payer system, where the vast majority of individuals are insured, and services are financed from contributions and state subsidies, administered by the National Health Insurance Fund Administration (NHIFA).

The majority of health care services, including both outpatient and inpatient care, do not require co-payments, although informal payments are common for a wide range of services.[1] The NHIFA finances most outpatient services with fee-for-service points (depending on the complexity and resource requirements of the procedures) and most inpatient services in a HDG-based system (homogenous disease groups, depending on the type and severity of the medical case). People may opt for using private care (which is common only in certain specialties, e.g. in dental care or gynecology), when they have to pay fee for the services.

User fees of medications depend on the amount of subsidies from the NHIFA, which

[1] There was an exceptional period between February 2007 – March 2008, when a small amount of flat co-payment for outpatient care and per diem payment for inpatient care were required. These co-payments were abolished as a consequence of a referendum held on the issue.

greatly varies across substances, and may also vary within the same substance, depending on who prescribed the medication (GP or specialist). No out-of-pocket payments are required for medicines received during inpatient care.

Hungary spent 7-8% of GDP on health care in 2005-2009, of which around 70% were public expenditures (Table 3.1 of Gaál et al. [2011]). In our empirical analysis, we focus on public outpatient and inpatient care expenditures, and public and private (out-of-pocket – OOP) expenditures on prescribed medications between 2003 and 2011. Altogether we analyse around 60% of total health care expenditures in Hungary. The most important excluded items are primary care services, and OOP payments on non-prescribed medications and on medical services.

Although GPs are meant to act as gatekeepers in Hungary, the gatekeeping system is not particularly effective. First, patients have direct access to a range of specialist services (dermatology, gynecology, ear, throat and nose diseases, among others). Secondly, GPs have no incentives to avoid referrals to specialists (Gaál et al. [2011]). These institutional circumstances imply a relatively high usage rate of specialist outpatient care (Elek et al. [2015]), which is also reflected in the data we use (we observe positive specialist care expenditures for the vast majority of people aged 54-63). Therefore, the effect of retirement on total outpatient care expenditures is likely to be captured well with our measure despite the fact that we do not observe expenditures on GP services.

As noted earlier, retirement may influence health care use through the sick leave channel, hence its effect on health care expenditures is likely to be related to the generosity of sick leave. If the sick pay system is more generous then retirement may have larger negative effect on health care expenditures, as the incentive to use health care so as to claim sick pay is no longer present after retirement. In Hungary, the amount of sick leave benefit is 70% for the first 15 days of sick leave, and 60% or 50% (depending on the length of work history) afterwards for up to one year. The amount of sick leave benefit was capped at 400% of the minimum wage up to May 2011, since then the cap is 200% of the minimum wage. In terms of the generosity of sick leave policy (e.g. the replacement level), Hungary is in the middle range of the countries of the European Union (Spasova et al. [2016]). To claim sick leave benefits, all countries of the European Union, including Hungary, require a declaration of incapacity for work, issued by a doctor (GP, specialist or hospital doctor) (MISSOC [2017]).

3.2 Pension system in Hungary

Hungary has a mandatory, pay-as-you go system, where pension benefits are based on earnings before retirement, with a minimum pension. Eligibility is conditional on 20 years of service. The government sets the standard retirement age, which has been increasing since 1998. However, under some mild conditions, earlier retirement is possible. First, during the analysed period, an early retirement age existed both for women and men.[2] Retiring at the early retirement age implies a reduction in benefits, the amount of which changed slightly during the period. Still, as we show in Figure 1, people tend to retire at the early, rather than at the full retirement age. Second, people working in jobs arduous to health (e.g. miners, people of the armed forces) and some other professional groups (artists, members of the parliament) could retire early. These eligibilities have been restricted since January 2012 (OECD [2013]).

[2]This was abolished in January 2012 but as a new rule, since 2011, women can retire after 40 years of service, irrespective of age.

Table 1 displays the official retirement ages for the cohorts in our analysis. (For later cohorts, the full retirement age is increasing further, up to age 65.) The increase of women's early retirement age from 57 to 59 years (i.e. cohort 1951 could retire already in 2008, while cohort 1952 only in 2011) is of particular interest because it provides a sudden and largely unexpected shift that can be regarded as an exogenous policy change. The corresponding law was enacted in December 2007, implying that people born in 1952 or later had only about a year to adjust to the change.

Table 1: Official early and full retirement age by cohorts

	Women		**Men**	
Year of birth	Early retirement age	Full retirement age	Early retirement age	Full retirement age
1945	55	60	60	62
1946	56	61	60	62
1947–1951	**57**	62	60	62
1952	**59**	62.5	60	62.5
1953	-	63	-	63

Source: Cseres-Gergely [2015] and OECD [2011].

Disabled people may opt for disability pension before reaching the statutory retirement age. In this paper, we do not focus on the effects of disability pension on health care expenditures (strong correlations are expected), but our focus is rather on the causal effects of age-based retirement.

4 Data and descriptive analysis

The empirical analysis is based on a unique administrative data set from Hungary. The data cover years 2003–2011, the sample is the half of the 15-74 years old population in 2003. The data set was created by linking administrative data from the Hungarian tax authority, the pension and the health authorities, among others.[3]

The core labour market data – such as various categories of income and employment – are observed monthly. When constructing the monthly indicators of labour force status, we categorise an individual as working (as opposed to being retired) if the monthly sum of her employment income is larger than the monthly pension benefit received. We then collapse the whole data set to annual frequency because health care expenditures are observed only on the annual level in our database. The annual indicators of labour force status (working or retired) are defined as the median of the monthly status indicators. The year of observed retirement is obtained as the first year when an individual receives old-age pension benefits.

To focus on genuine transitions from employment to full-time retirement, we make two sample restrictions in our main analysis. First, we include only people who were working three years before the official early retirement age (i.e. at 54 years for women and 57 years for men) and earned at least 90% of the minimum wage.[4] In our dataset, 57% of women and 49% of men were classified as working at the age of 54 and 57 years, respectively, which is comparable to the average employment rates of the corresponding cohorts in the European Union.[5] Second, we exclude those who remained employed (i.e. received some labour income) after getting retired, which affects around 25% of women of ages 58-59, who worked at age

[3]The linked data set is under the ownership of the Central Administration of National Pension Insurance, the National Health Insurance Fund Administration, the Educational Authority, the National Tax and Customs Administration, the National Labour Office, and the Pension Payment Directorate of Hungary. The data used were processed by the Institute of Economics, Centre for Economic and Regional Studies of the Hungarian Academy of Sciences.

[4]The 3-year window is chosen because, on one hand, direct preparations for retirement are unlikely to have started and employment rates are fairly stable at that age, but on the other hand, not too many observations are lost due to missing lagged values. By the minimum wage restriction we concentrate on the core segment of the labour market, i.e. exclude most employees in part-time jobs and short job spells.

[5]According to the slightly different terminology of the EU Labour Force Survey, women aged 50-54 years had an employment rate of 66% in Hungary and 63% in the EU-27, while men aged 55-59 years had an employment rate of 57% in Hungary and 65% in the EU-27 in 2005 (Eurostat [2017]).

54. As demonstrated later, these sample restrictions do not change the econometric results.

Figure 1 shows that 60-70% of those individuals who worked at age 54 (women) or 57 (men) retire at or before the age of the early retirement age (57 years for women born in 1947-1951, 59 years for women born in 1952 and 60 years for men), and 40-50% at exactly this age. Although women whose official early retirement age is 59 are observed up to age 59 only, it is clear from the graph that retirement rates at ages 57-58 are substantially lower for them than for the older cohorts with lower early retirement age. Compliance with the early retirement age facilitates our identification strategy, i.e. using the early retirement age as an exogenous determinant of retirement status.

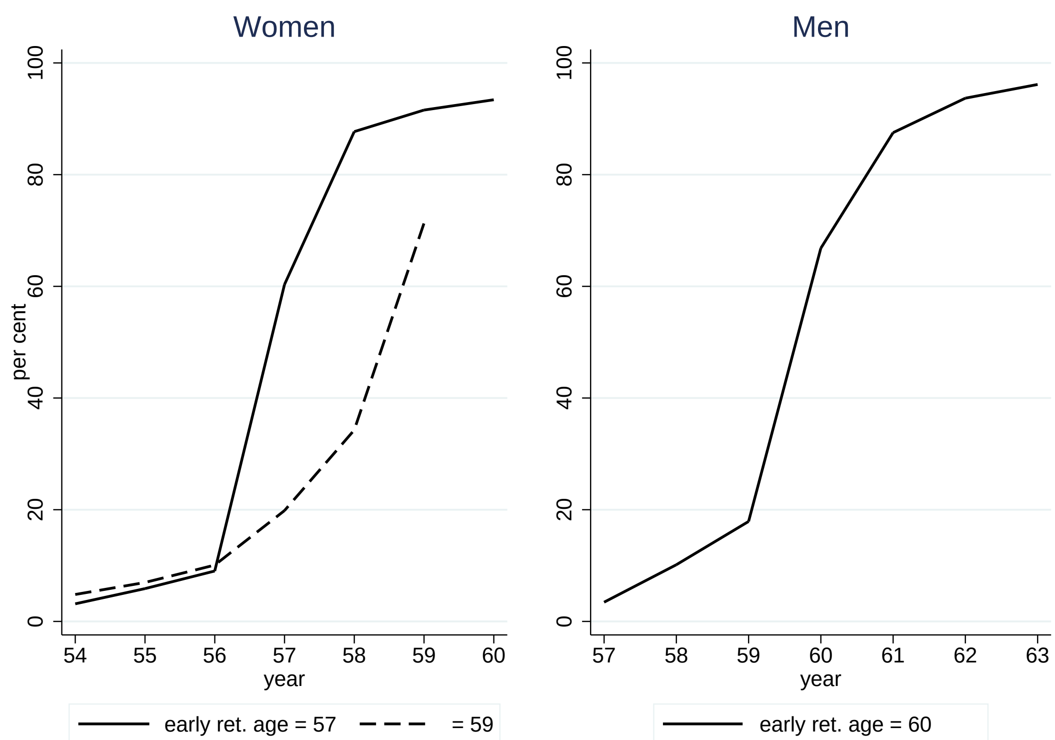

Figure 1: Percentage of retired individuals by age, according to early retirement age
Sample: working at age of 54 years (women) or 57 years (men)

Figures 2–3 display the age patterns of health care expenditures around the early retirement age (54–60 years for women and 57–63 years for men). We present the fractions of positive values and the logarithms of positive values for the three categories of care. All expenditure graphs control for calendar year effects with simple OLS regressions, using the grand mean across years as the baseline.

Each plot shows groups according to the employment status three years before the early retirement age. The first group is formed by those who worked at age 54 (women) or 57 (men), and earned at least 90% of the annual minimum wage at that age, thus for whom the early retirement age may be relevant. For females, this group is split further according to the official early retirement age of the cohort (57 or 59 years).[6] The other group contains those who neither worked at all nor went into old-age retirement any time from age 54 (women) or 57 (men). The (early) retirement age is irrelevant for this group. To exclude very ill patients with high end-of-life health care expenditures, the expenditure graphs of males are conditional on survival up to three years after the early retirement age.[7]

People who have worked prior to the early retirement age form a relatively healthy subgroup of the population in terms of both mortality and health care expenditures, which is in line with the literature (see e.g. Disney et al. [2006]). According to the mortality data (not shown here in detail), non-working men experience a drastic 1.8% annual mortality rate between ages 57 and 63, while this figure is only 0.5% for those men who were working at age 57. Similar mortality differences (0.5% vs. 0.2% annually) are present for women in the 54-60 age range.

[6] Older cohorts whose early retirement age was lower than 57 years are automatically excluded because their employment status at age 54 is not observed.

[7] This constraint does not apply for women because the cohort with official early retirement age equal to 59 years is observed only until this age.

According to the upper panels of figures 2–3, only around 10% of workers but 20-30% of non-workers have positive inpatient expenditures at age 54 / 57. Meanwhile, the vast majority (75–90%) of both workers and non-workers have positive outpatient and pharmaceutical expenditures. Looking at the logarithm of the positive values, the lower panels show that working people have on average much lower medical expenditures than the non-working population. The gap in the average expenditures of workers vs. non-workers tends to narrow down substantially along the six-year long observed interval.

Finally, it is worth looking at the two groups of women whose official early retirement age is 57 and 59 years, respectively. After calendar year adjustment, the cohort with retirement age 59 is slightly more likely to have positive expenditures in all three categories (especially in outpatient care and pharmaceuticals), and, more importantly, the gaps become more substantial above age 56, when women in the older cohorts typically started to retire. This suggests a negative effect of retirement on the extensive margin, and will be reflected in our estimation results.

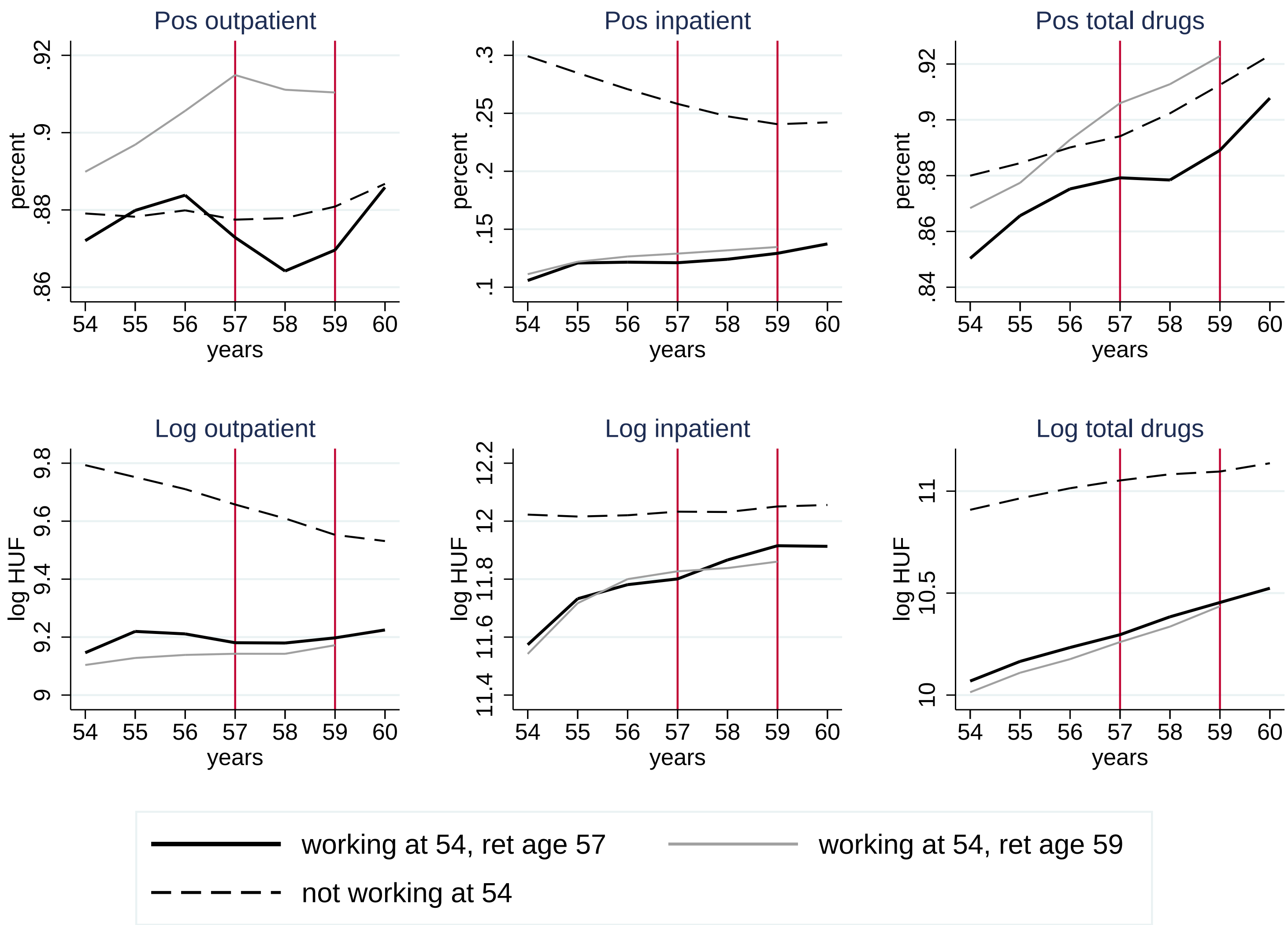

Figure 2: Age-specific (and calendar year-adjusted) health care expenditures of women by employment status and official early retirement age (57 or 59). The fractions of positive values and the logarithms of positive values are shown.

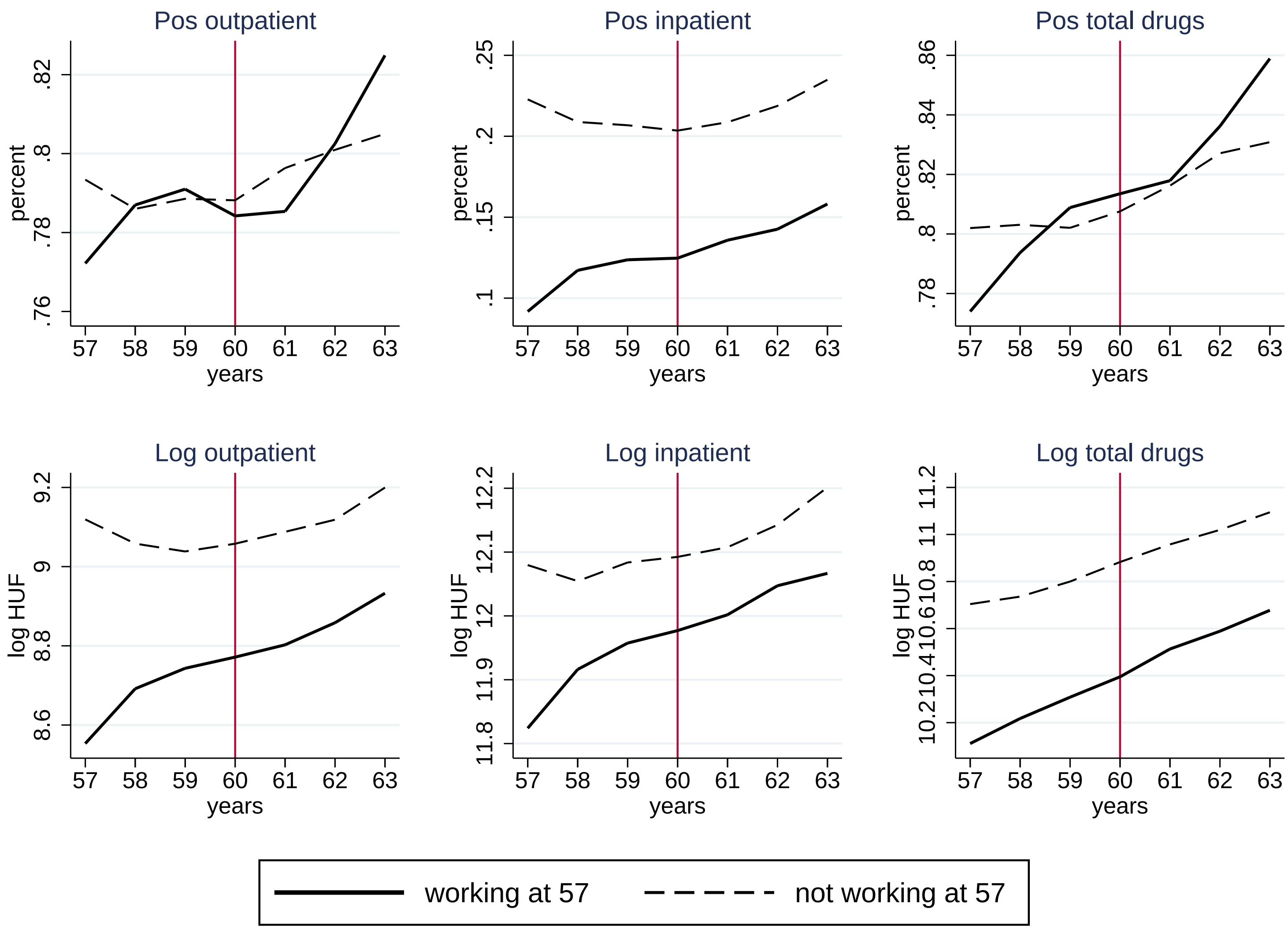

Figure 3: Age-specific (and calendar year-adjusted) health care expenditures of men by employment status (official early retirement age is 60). The fractions of positive values and the logarithms of positive values are shown.

5 Methods

Consider the following model of health care expenditures:

$$y_{it} = \alpha_0 + R_{it}\alpha_r + h_{\alpha_z}(Z_{it}) + D_{it}^{year}\alpha_d + v_i + u_{it}, \tag{1}$$

where i stands for individuals, t for years. Here, expenditures are determined by retirement status R_{it}, by calendar year dummies D_{it}^{year} (capturing changes in prices and regulatory environment), and by the continuous age[8] of the individual Z_{it}. We model the age-expenditure relationship with a parametric function $h_{\alpha_z}(Z_{it})$. Individual fixed effects (such as time-invariant health care use preferences, supply- and demand-side factors in the living area etc.) are captured by v_i.

The outcome variable (y) is an indicator of outpatient, inpatient or pharmaceutical expenditures. For all three categories, we use the binary indicator of positive expenditures (extensive margin) and the logarithm of positive expenditures (intensive margin) as the dependent variable, hence we estimate two-part (hurdle) models. The two-part model with logarithmic transformation has two main advantages here: first, it allows heterogeneous effects in the two margins and second, it provides more efficient estimates than regressions on levels do because of the skewness and non-normality of the expenditure distributions. (For the use of two-part and other models of health care expenditures, see e.g. Mihaylova et al. [2011]. See also Figure 4 in Appendix A for the histograms of health care expenditures.)

Since health status may affect the retirement decision of individuals, R_{it} may be endogenous in the model. We instrument R_{it} with $I\{Z_{it} > E_i\}$, where $I\{.\}$ is the indicator

[8] We define Z_{it} as a continuous variable by the age of the individual in July of a given year.

function, and E_i denotes the official early retirement age. Since in Hungary the official early retirement age for women increased from 57 to 59 years in 2009, this instrumenting strategy still allows a flexible parametric form for $h_{\alpha_z}(Z_{it})$. In our baseline specification, we use a quadratic function for the age effect and estimate the IV model for women at ages 56–59, born between 1949–1953 (i.e. with official early retirement age 57 or 59 years), whose observed retirement age is between 57–59, and who were working and earning at least 90% of the annual minimum wage at age 54. These sample restrictions are driven by the aim of focusing on the cohorts and ages that were directly affected by the increasing statutory early retirement age, and by data limitations (i.e. the cohort whose early retirement age is 59 years was observed only until that age). The IV model is estimated with fixed-effects (FE) two-stage least squares, and standard errors are robust to heteroscedasticity and serial correlation.[9] Since the linear age effects and the calendar year dummies are perfectly collinear and hence cannot be identified separately in the model with individual fixed effects, the age and year coefficients will not be reported in the tables of results.

We also perform robustness checks. First, results on the full sample (including people not working at age 54 and people working after retirement) are shown. Second, FE OLS results without the instrumenting strategy are displayed. Third, pooled bivariate probit models are estimated on the binary variables of having positive health care expenditures, where retirement is a binary endogenous regressor.

To investigate heterogeneity in the effect of retirement on health care expenditures, we estimate a version of equation (1) by interacting the retirement status (and the instrument) with certain individual characteristics capturing health status, previous history of sick leave,

[9] The estimation is carried out with the xtivreg2 Stata command, written by Schaffer [2010].

education and sector of employment. In these models, we interact all control variables (age and calendar year effects) with the heterogeneity variables.

Finally, to give supplementary evidence and to exploit the available data for men and for women in the period of their unchanged early retirement age, we also estimate models for both genders, in which we use the *constant* early retirement age as instrumental variable for retirement. This is the usual way in the literature to overcome endogeneity (see Neuman [2008], Coe and Zamarro [2011], Bonsang et al. [2012], Eibich [2015], Bíró [2016]). Compared to the baseline IV model, this specification may exploit a wider age range before and after retirement. We use three years before and after the official early retirement age, i.e. 54–60 for women and 57–63 for men, and restrict the sample to those who worked at the beginning of these age ranges and earned at least 90% of the annual minimum wage. However, if the official retirement age is time-constant, this identification strategy requires a severe restriction on the age-expenditure relationship in a single-country setting because a flexible functional form for $h_{\alpha_z}(Z_{it})$ would not allow the separation of the age and the retirement effect. Similarly to the regression discontinuity design specifications in the literature, we define this as a linear function with a possible trend break at the early retirement age.

6 Results

6.1 Baseline specification

Table 2 displays the estimation results of equation (1) for women aged 56-59 years, exploiting the exogenous increase in their early retirement age from 57 to 59 years. (Descriptive statistics of the variables, along with those used in the heterogeneity analysis, are shown in Appendix A.) In line with Figure 1, the first-stage equation indicates a strong compliance with the official early retirement age. Being above this age increases the probability of being retired by more than 40 %points. This effect is larger than similar studies suggest (Bonsang et al. [2012], Coe and Zamarro [2011], Neuman [2008]) mostly because here we focus on those women who worked at the age of 54 years. By removing all the work- and retirement-related sample restrictions, the first-stage coefficient halves, while the second stage results do not change substantially (see Table 6 in Appendix A).

Turning to the main findings, the probability of having positive outpatient care expenditures, inpatient care expenditures and pharmaceutical expenditures fall statistically significantly due to retirement (by 3.0, 1.4 and 1.3 %points, respectively). The effects on the intensive margin (i.e. on the logarithm of positive expenditures) are also negative, although statistically insignificant at the 5% level. The results are qualitatively robust to the choice between the linear FE IV model and the probit model with endogenous retirement. According to Table 8 in Appendix A, the probit marginal effects are slightly larger in magnitude for outpatient care and pharmaceuticals and somewhat lower for inpatient care (and become statistically insignificant in the latter case). The FE OLS model without the instrumenting strategy gives moderately smaller estimates in absolute value, which reflects the selective nature of retirement due to health (see Table 7 of Appendix A).

In line with the theoretical considerations of section 2, there are multiple possible explanations for the negative effects on the extensive margin. First, it can be due to demanding sick leave while working. If one wishes to claim sick leave benefits, she has to visit her GP, who may refer her to outpatient or inpatient care, or prescribe medications. After retire-

Table 2: Effect of retirement on health care expenditures, FE IV results (women, aged 56-59)

Dependent variable: expenditure category	prob. of positive outpatient	log of positive outpatient	prob. of positive inpatient	log of positive inpatient	prob. of positive drugs	log of positive drugs
retired	-0.030***	-0.027	-0.014**	-0.074	-0.013**	-0.035*
	[0.006]	[0.026]	[0.007]	[0.094]	[0.006]	[0.019]
Number of obs.	186,296	157,637	186,296	8,789	186,296	159,248
First stage equation: dependent variable = retired						
Above early retirement age	0.439***	0.435***	0.439***	0.457***	0.439***	0.438***
	[0.003]	[0.003]	[0.003]	[0.016]	[0.003]	[0.003]
First stage F-statistic	19,198	15,534	19,198	835	19,198	16,117

Robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
Individual fixed effects are included. Controls: age, squared age and calendar year dummies
First stage F-statistics refer to the relevance of the instrument.
Sample: women born in 1949-1953, aged 56-59 years and working at age 54, with MW restriction

ment, women may be less willing to see a GP and then a specialist due to minor illnesses, leading to the observed negative effect. Second, following Kuhn et al. [2015], people may be willing to invest more in their health while working so as to achieve higher earning – and these mainly affect diagnostic and prevention procedures, and less likely inpatient care. Third, income factors may also play a role in e.g. pharmaceutical spending. Finally, it is unlikely that direct health effects are substantial here because of the short time span after retirement.

6.2 Heterogeneity

6.2.1 Health and sick leave status

We first check how health status before retirement affects the estimated impact of retirement on health care expenditures. Since our data set does not contain direct information on health, we use total drug expenditures at age 54 as its proxy. This expenditure measure was selected because it has non-negligible variation across individuals, and has a substantial effect on observed mortality in six years, as shown in Appendix B. We create a binary indicator which equals one for a woman if her drug expenditures at age 54 were below the calendar year specific median of the expenditures of women of the same age.

We also examine the heterogeneous effect of retirement across whether a woman was on sick leave any time during the calendar year at age 54. This sick leave dummy measures a different, work-related component of health than the drug expenditure dummy above: 14.3% of women with higher than median drug expenditures spent some time on sick leave at the age of 54, while this ratio was still 6.4% for women with lower drug expenditures.

According to the first panel of Table 3, retirement has essentially no effect on the health care expenditures of the less healthy women (who had higher drug expenditures and thus greater need for care), and a substantially negative effect on the probability of positive outpatient and positive pharmaceutical expenditures for the relatively healthy women (who had lower drug expenditures at age 54). The interaction term of retirement and low drug expenditures is not statistically significantly different from zero for the four other dependent variables.

The first panel of the table also shows that the probability of having positive outpatient expenditures decreases much more strongly (by 4.4 %points) due to retirement for those who spent some time on sick leave at age 54, which yields evidence for the sick leave hypothesis. There is also a negative, although statistically insignificant effect on the extensive margin of pharmaceutical expenditures. We note that the heterogenous retirement effects by health status above may also be explained by the sick leave hypothesis: patients with higher drug expenditures (most of whom have chronic illnesses) maintain their health care use after retirement, while healthy patients with low drug expenditures at 54 years might catch acute minor illnesses approximately randomly. This latter group may visit the doctor for sick leave while working but stay out of health care with such illnesses after retirement.[10]

6.2.2 Socio-economic status

Second, we analyse how the effect of retirement varies with the level of education. Although the level of education is not included in the administrative database, we can approximate it by using the four-digit occupational code (ISCO) of the individual, which is recorded for employees (but not for the self-employed) in the sample. More precisely, we impute the level of education with the median education level of employees of the same occupational code between ages 50-59 years as observed in the Labour Force Survey (LFS). We differentiate four categories: primary, lower secondary, upper secondary, and tertiary education. According to the results in the second panel of Table 3, retirement effects are mainly present for the less educated. For people with primary education (who on average earn less), retirement substantially reduces the probability of positive outpatient and positive pharmaceutical expenditures. These effects are somewhat weaker, but still significant among the lower secondary educated. At the same time, the parameters for the tertiary educated are never statistically significant.

The heterogenous effect of retirement by the level of education points to the mechanism outlined in Kuhn et al. [2015]: if earnings and the disutility of work are more strongly affected by health for the lower educated then their health care expenditures drop more after retirement than in the case of the higher educated. Moreover, as Eibich [2015] points out, the lower educated might experience a relief from arduous working conditions with retirement, implying improved health for them.

[10]Both the health status and the sick leave indicator correspond to data 3–5 years before retirement, hence mean reversion of the expenditure categories is not likely to drive the econometric results.

Finally, we check whether the effect of retirement on health care expenditures is heterogeneous across sectors of employment. Different effects may be expected not only due to varying income and education by sector, but possibly also due to varying time constraints and access to care while working. The results of the third panel of Table 3 indicate that the negative baseline effects of retirement on the extensive margin are mostly driven by women employed in the private and public sector and less markedly by women working as entrepreneurs at age 54.[11] This finding is in line with the hypothesis that demand for outpatient care so as to claim sick leave benefits while working may have been more prevalent in the former employment groups.

Remarkably, retirement is estimated to have a substantial negative effect on the size of pharmaceutical expenditures among women working as entrepreneurs at age 54. In line with Kuhn et al. [2015], a possible explanation is that entrepreneurs spend more on medications (as opposed to the sick leave channel) while working so as to maintain their capacity to work. The earnings-based motivation to maintain good health disappears after retirement.

Overall, the heterogeneity analyses yield evidence that the baseline negative effects on the extensive margin are mostly driven by the relatively healthy, those having been on sick leave, and the less educated.

6.3 Specification with constant early retirement age

According to Table 4, the specifications with constant early retirement age (i.e. equation (1), where R_{it} is instrumented with the constant early retirement age and $h_{\alpha_z}(Z_{it})$ is a linear

[11]Other employment categories are not included in this heterogeneity analysis due to their small fraction in our estimation sample.

Table 3: Heterogeneity of the retirement effect, FE IV results (women, aged 56-59)

Heterogenous effect of retirement	prob. of positive outpatient	log of positive outpatient	prob. of positive inpatient	log of positive inpatient	prob. of positive drugs	log of positive drugs
Model: interaction with below median drug expenditures and positive sick leave status at 54 years						
retired (baseline)	-0.0033	-0.035	-0.016	0.0064	0.0057	-0.015
	[0.0085]	[0.038]	[0.012]	[0.137]	[0.0048]	[0.023]
retired × low drug expend at 54	-0.040***	0.047	0.0023	-0.295	-0.028***	-0.024
	[0.018]	[0.054]	[0.014]	[0.209]	[0.011]	[0.040]
retired × sick leave at 54 years	-0.044**	-0.112	0.0046	-3.9E-0.5	-0.021	-0.075
	[0.018]	[0.078]	[0.024]	[0.250]	[0.015]	[0.058]
Number of obs.	186,296	157,637	186,296	8,789	186,296	159,248
Model: interaction with education						
retired × primary	-0.072***	0.037	0.00016	-0.144	-0.032**	-0.024
	[0.016]	[0.067]	[0.016]	[0.237]	[0.014]	[0.046]
retired × lower secondary	-0.044***	0.013	0.0017	-0.036	-0.028**	-0.033
	[0.013]	[0.052]	[0.013]	[0.189]	[0.012]	[0.037]
retired × upper secondary	-0.021**	-0.049	-0.032**	-0.014	-0.0099	-0.0080
	[0.010]	[0.040]	[0.011]	[0.153]	[0.0088]	[0.030]
retired × tertiary	-0.0017	-0.016	-0.016	-0.150	0.0096	-0.051
	[0.018]	[0.075]	[0.020]	[0.232]	[0.016]	[0.057]
Number of obs.	174,652	148,122	174,652	8,294	174,652	149,581
Model: interaction with sector of employment at 54 years						
retired × private sector	-0.026***	-0.021	-0.021***	-0.018	-0.016***	-0.055**
	[0.008]	[0.032]	[0.008]	[0.012]	[0.007]	[0.024]
retired × public sector	-0.041***	0.00077	-0.0054	-0.108	-0.017	0.027
	[0.012]	[0.050]	[0.013]	[0.165]	[0.011]	[0.037]
retired × entrepreneur	-0.0075	-0.183	-0.0063	-0.414	0.024	-0.120*
	[0.022]	[0.083]	[0.021]	[0.392]	[0.019]	[0.065]
Number of obs.	185,851	157,267	185,851	8,759	185,851	158,867

Robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
Individual fixed effects are included. All models include age, squared age and calendar year dummies, interacted with the given explanatory variable (drug expenditures, sick leave, education, sector).
Sample: see Table 2

function with a trend break) yield negative effects of retirement on the probability of having positive outpatient, inpatient and pharmaceutical expenditures for both genders (estimates vary from -1.7 to -4.1 %points), hence they reinforce the findings of the baseline model in Table 2. On the other hand, the results on the intensive margin are slightly different from the baseline specification: there is a substantial negative effect on the logarithm of outpatient expenditures for both genders.[12]

Since this model is based on a restrictive assumption on the age-expenditure relationship, we present a set of specification (placebo) checks in Appendix C such as using cut-off points other than the early retirement age or estimating models for the non-working populations. The results of the checks are satisfactory.

Table 4: Models with constant early retirement age, FE IV results

	prob. of positive outpatient	log of positive outpatient	prob. of positive inpatient	log of positive inpatient	prob. of positive drugs	log of positive drugs
Model for women, aged 54-60, E.R.A. = 57						
retired	-0.041***	-0.097***	-0.018***	-0.074	-0.024***	-0.012
	[0.005]	[0.019]	[0.005]	[0.060]	[0.004]	[0.015]
age minus E.R.A.	0.0044**	0.0062	0.0080***	-0.010	0.0000	0.011*
× above E.R.A.	[0.0018]	[0.0073]	[0.0018]	[0.025]	[0.0016]	[0.006]
Number of obs.	293,192	256,506	293,192	20,063	293,192	256,873
Model for men, aged 57-63, E.R.A. = 60						
retired	-0.038***	-0.128***	-0.019***	0.100*	-0.017***	-0.021
	[0.005]	[0.020]	[0.004]	[0.057]	[0.004]	[0.014]
age minus E.R.A.	0.0036***	0.0003	-0.0011	-0.019	-0.0046***	0.0044
× above E.R.A.	[0.0014]	[0.0055]	[0.0012]	[0.016]	[0.0011]	[0.0040]
Number of obs.	412,887	313,169	412,887	25,107	412,887	326,131

Robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
E.R.A. stands for official early retirement age.
The term with "above E.R.A." denotes the trend break at E.R.A.
Individual fixed effects are included. Controls: age minus E.R.A., calendar year dummies
Sample: women and men in the given age ranges and E.R.A.,
who were working three years before this age, with MW restriction

[12] Again, the linear age and the calendar year effects are not identified separately in this FE model, hence Table 4 does not display the coefficients of these control variables.

7 Discussion and conclusions

While there is a rich literature on the health effects of retirement, little is known about the effects of retirement on health care expenditures. Such an analysis requires administrative panel data, and an exogeneous variation in retirement status. In this paper, we took the first steps in estimating the causal effect of retirement on health care expenditures.

Descriptive analysis showed that looking at a couple of years before the early retirement age, the health care expenditures among non-working people tend to be substantially higher than that of workers. This gap gradually narrows around and after the early retirement age. Such differences are in line with the health related nature of labour market status. Nevertheless, when we used the increase of the early retirement age to estimate the causal effect of retirement, we found negative short run effects on the extensive margin for women. These results are mostly driven by the relatively healthy, by those having been on sick leave and by the less educated, which is in line with the theoretical considerations.

We also estimated models with constant early retirement age for women and men, and obtained similar results to the causal estimates on the extensive margin. Although placebo checks did not indicate problems with this specification, the causal model that uses an exogenous increase in the early retirement age has more validity.

Our findings suggest that retirement tends to decrease public health care expenditures, at least in the short run. In this sense, retirement might decrease the burden on public health, while raising the retirement age might lead to somewhat higher public health care expenditures.

Our analysis is subject to some limitations. Due to the institutional setting, we could derive a causal effect of retirement on health care expenditures only for women. Also, due to data censoring, we could estimate only short run effects of retirement. The administrative data set we use does not include indicators of health, thus we could only speculate about how health effects of retirement might influence our results.

References

S. Behncke. Does retirement trigger ill health? *Health Economics*, 21(3):282–300, 2012.

A. Bíró. Outpatient visits after retirement in Europe and the US. *International Journal of Health Economics and Management*, 16(4):363–385, 2016.

D. E. Bloom, D. Canning, and M. Moore. Optimal retirement with increasing longevity. *The Scandinavian Journal of Economics*, 116(3):838–858, 2014.

R. F. Boaz and C. F. Muller. Does having more time after retirement change the demand for physician services? *Medical Care*, 27(1):1–15, 1989.

E. Bonsang, S. Adam, and S. Perelman. Does retirement affect cognitive functioning? *Journal of Health Economics*, 31(3):490–501, 2012.

E. Calvo, N. Sarkisian, and C. R. Tamborini. Causal effects of retirement timing on subjective physical and emotional health. *The Journals of Gerontology Series B: Psychological Sciences and Social Sciences*, 68(1):73–84, 2013.

N. B. Coe and G. Zamarro. Retirement effects on health in Europe. *Journal of Health Economics*, 30(1):77–86, 2011.

Z. Cseres-Gergely. A 2000-es évek magyarországi nyugdíjkorhatár-emeléseinek azonnali hatása az érintett nők munkavállalására [The effect raising the retirement age has on the employment rate of older women. Empirical evidence from retirement-age increases in Hungary in the 2000s]. *Közgazdasági Szemle (Economic Review - monthly of the Hungarian Academy of Sciences)*, 62(6):652–673, 2015.

R. Disney, C. Emmerson, and M. Wakefield. Ill health and retirement in Britain: A panel data-based analysis. *Journal of Health Economics*, 25(4):621–649, 2006.

P. Eibich. Understanding the effect of retirement on health: mechanisms and heterogeneity. *Journal of Health Economics*, 43:1–12, 2015.

P. Elek, B. Váradi, and M. Varga. Effects of geographical accessibility on the use of outpatient care services: Quasi-experimental evidence from panel count data. *Health Economics*, 24 (9):1131–1146, 2015.

Eurostat. Employment and unemployment (LFS). `http://ec.europa.eu/eurostat/web/lfs/data/database`, 2017. Accessed: 2017-10-20.

P. Gaál, S. Szigeti, M. Csere, M. Gaskins, and D. Panteli. Hungary: Health system review. *Health Systems in Transition*, 13(5):1–266, 2011.

T. Galama, A. Kapteyn, R. Fonseca, and P.-C. Michaud. A health production model with endogenous retirement. *Health Economics*, 22(8):883–902, 2013.

E. Hernaes, S. Markussen, J. Piggott, and O. L. Vestad. Does retirement age impact mortality? *Journal of Health Economics*, 32(3):586–598, 2013.

G. W. Imbens and T. Lemieux. Regression discontinuity designs: A guide to practice. *Journal of Econometrics*, 142(2):615–635, 2008.

M. Insler. The health consequences of retirement. *Journal of Human Resources*, 49(1): 195–233, 2014.

M. Kuhn, S. Wrzaczek, A. Prskawetz, and G. Feichtinger. Optimal choice of health and retirement in a life-cycle model. *Journal of Economic Theory*, 158:186–212, 2015.

P. Lundborg, M. Nilsson, and J. Vikström. Heterogeneity in the impact of health shocks on labour outcomes: evidence from Swedish workers. *Oxford Economic Papers*, 67(3): 715–739, 2015.

B. Mihaylova, A. Briggs, A. O'Hagan, and S. G. Thomson. Review of statistical methods for analysing healthcare resources and costs. *Health Economics*, 20(8):897–916, 2011.

MISSOC. Mutual information system on social protection. Comparative tables database, sickness – cash benefits, conditions. http://www.missoc.org/, 2017. Accessed: 2017-09-28.

K. Neuman. Quit your job and get healthier? The effect of retirement on health. *Journal of Labor Research*, 29(2):177–201, 2008.

OECD. *Pensions at a Glance 2011: Retirement-Income Systems in OECD and G20 Countries*. OECD Publishing, Paris, 2011.

OECD. *Pensions at a Glance 2013: OECD and G20 Indicators*. OECD Publishing, Paris, 2013.

OECD. *Pensions at a Glance 2015: OECD and G20 Indicators*. OECD Publishing, Paris, 2015.

P. A. Puhani and K. Sonderhof. The effects of a sick pay reform on absence and on health-related outcomes. *Journal of Health Economics*, 29(2):285–302, 2010.

M. E. Schaffer. xtivreg2: Stata module to perform extended iv/2sls, gmm and ac/hac, liml and k-class regression for panel data models. http://ideas.repec.org/c/boc/bocode/s456501.html, 2010. Accessed: 2017-08-28.

E. Shapiro and N. P. Roos. Retired and employed elderly persons: Their utilization of health care services. *The Gerontologist*, 22(2):187–193, 1982.

K. Soghikian, L. T. Midanik, M. R. Polen, and L. J. Ransom. The effect of retirement on health services utilization: the Kaiser Permanente Retirement Study. *Journal of Gerontology*, 46(6):S358–S360, 1991.

S. Spasova, D. Bouget, and B. Vanhercke. Sick pay and sickness benefit schemes in the European Union. Background report for the Social Protection Committees In-depth Review on sickness benefits (17 October 2016), European Social Policy Network (ESPN), Brussels, European Commission, 2016.

I. van der Heide, R. M. van Rijn, S. J. Robroek, A. Burdorf, and K. I. Proper. Is retirement good for your health? A systematic review of longitudinal studies. *BMC Public Health*, 13(1):1180, 2013.

Appendix

Appendix A: Descriptive statistics and robustness checks of the baseline model

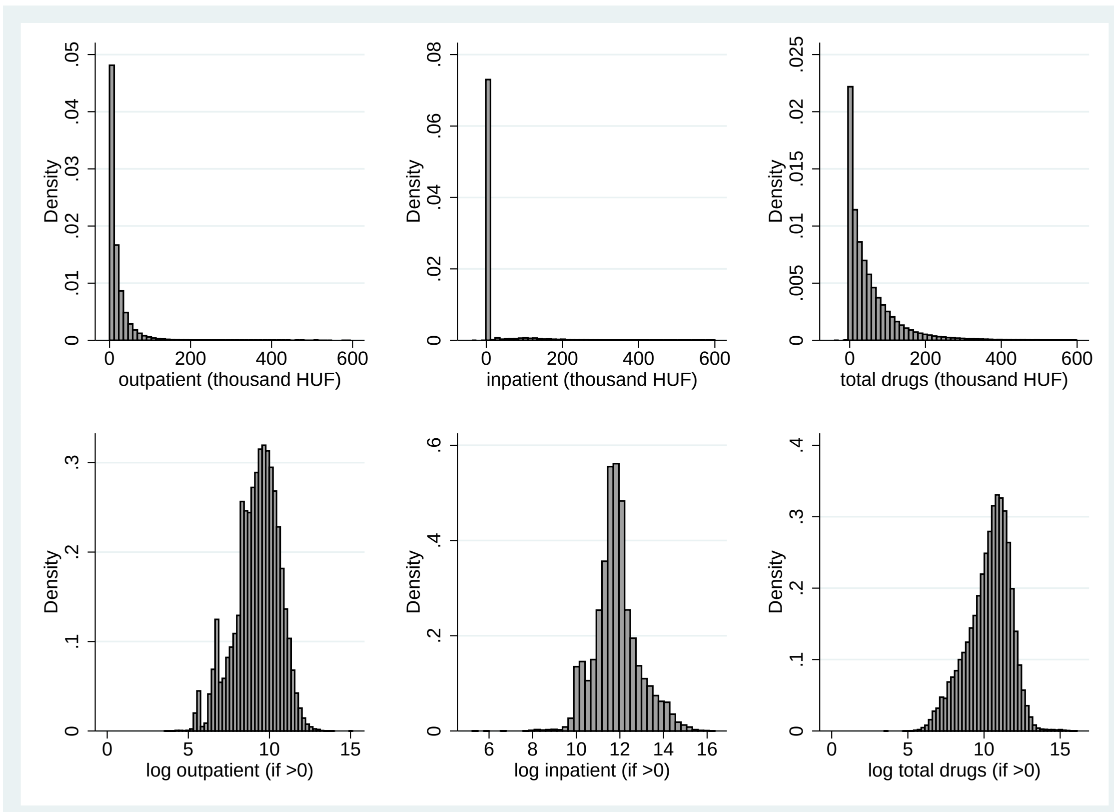

Figure 4: Distribution of health care expenditures and of the logarithm of positive values
Sample: see Table 2

Table 5: Descriptive statistics of the variables in the baseline and heterogeneity models

	mean	sd	N
outpatient care (1000 HUF)	18.949	32.118	191,484
positive outpatient care	0.874	0.332	191,484
log outpatient care	9.249	1.336	167,277
inpatient care (1000 HUF)	29.332	176.079	191,484
positive inpatient care	0.114	0.318	191,484
log inpatient care	11.839	1.032	21,840
drugs (1000 HUF)	62.105	168.828	191,484
positive drugs	0.880	0.325	191,484
log drugs	10.332	1.405	168,523
age	57.375	1.109	191,484
retired	0.394	0.489	191,484
observed retirement age	57.769	0.765	191,484
dummy for sick leave in calendar year at age 54	0.098	0.298	191,484
sector of employment at age 54			
private sector	0.547	0.498	191,484
public sector	0.383	0.486	191,484
entrepreneur	0.068	0.251	191,484
education			
primary	0.172	0.378	179,551
lower secondary	0.202	0.402	179,551
upper secondary	0.419	0.493	179,551
tertiary	0.206	0.404	179,551

Sample: see Table 2
1000 HUF was worth on average 3.6 EUR in the examined period.

Table 6: Effect of retirement on health care expenditures: FE IV results on the unrestricted sample (women, aged 56-59)

Dependent variable: expenditure category	prob. of positive outpatient	log of positive outpatient	prob. of positive inpatient	log of positive inpatient	prob. of positive drugs	log of positive drugs
retired	-0.034***	-0.060**	-0.0099	-0.118	-0.016***	-0.029
	[0.006]	[0.026]	[0.0075]	[0.137]	[0.006]	[0.019]
Number of obs	645,748	544,148	645,748	64,633	645,748	551,053
First stage equation: dependent variable = retired						
Above early retirement age	0.229***	0.226***	0.229***	0.107***	0.229***	0.227***
	[0.002]	[0.002]	[0.002]	[0.005]	[0.002]	[0.002]
First stage F-statistic	17,702	14,093	17,702	440	17,702	14,652

Robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
Individual fixed effects are included. Controls: age, squared age and calendar year dummies
First stage F-statistics refer to the relevance of the instrument.
Sample: women born in 1949-1953, aged 56-59 years

Table 7: Effect of retirement on health care expenditures: FE OLS results without instrumental variable (women, aged 56-59)

Effect of retirement	outpatient	inpatient	drugs
on probability of positive expenditures	-0.023***	-0.0061**	-0.016***
	[0.003]	[0.0027]	[0.0023]
Number of obs.	191,484	191,484	191,484
on logarithm of positive expenditures	-0.015	-0.037	-0.011
	[0.011]	[0.041]	[0.008]
Number of obs.	167,277	21,840	168,523

Robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
Individual fixed effects are included.
Control variables: age and its square, calendar year dummies.
Sample: see Table 2

Table 8: Average marginal effect of retirement on health care expenditures: pooled bivariate probit results (women, aged 56-59)

Effect of retirement	outpatient	inpatient	drugs
on probability of positive expenditures	-0.041***	-0.0060	-0.028***
	[0.005]	[0.0043]	[0.005]
Number of obs.	191,484	191,484	191,484

Robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
Control variables: age and its square, calendar year dummies
Sample: see Table 2

Appendix B: Health care expenditures and survival probability

While we observe health care expenditures, our data set does not include indicators of health, other than the survival. We demonstrate here the relationship between the three categories of health care expenditures and 6-year survival probability.

For women at age of 54 and for all three expenditure categories, Table 9 displays the conditional probabilities of having above median expenditure by 6-year survival status. It shows that above median pharmaceutical expenditures and above median inpatient expenditures (i.e. positive inpatient expenditures because of the presence of excess zeros) predict later mortality substantially better than above median outpatient expenditures do. For instance, 66% of those not alive at age 60 had above median pharmaceutical expenditures 6 years earlier, while this figure was only 59% for above median outpatient expenditures. This result coincides with Table 10, where we present logit models predicting 6-year survival probability with the linear and quadratic functions of all three expenditure categories. The models show that both inpatient and pharmaceutical expenditures are negatively related to the survival probability. At the same time, the estimated ceteris paribus effect of outpatient care expenditures are positive: at given inpatient and pharmaceutical expenditures, a higher use of outpatient care implies a higher probability of survival. All in all, we use above median pharmaceutical expenditures as the health status variable in our analysis because it gives more stable ceteris paribus effects in the logit model even in the presence of quadratic terms.

Table 9: Conditional probabilities of having above median health expenditures by 6-year survival status, sample of women aged 54

	outpatient	inpatient	drugs
not alive at age of 60	58.7%	39.7%	66.5%
alive at age of 60	49.7%	19.5%	49.4%

Table 10: Average marginal effects on being alive at age 60, logit models, women aged 54

	alive at 60	alive at 60
log(outpatient expenditures + 1)	0.0020***	0.0036***
	[0.00022]	[0.00036]
squared log(outpatient expenditures + 1)		0.00030***
		[5.9e-05]
log(inpatient expenditures + 1)	-0.0030***	0.010***
	[0.00012]	[0.00061]
squared log(inpatient expenditures + 1)		-0.0016***
		[7.7e-05]
log(total drug expenditures + 1)	-0.0011***	-0.0057***
	[0.00021]	[0.00032]
squared log(total drug expenditures + 1)		-0.00088***
		[4.8e-05]
Number of obs.	100,768	100,768

Standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
Health expenditures are cleaned from calendar year effects.
Control variables: calendar year dummies and living area indicators

Appendix C: Specification checks with constant early retirement age

Since the model with constant early retirement age is based on a restrictive linearity assumption, we present a set of specification (placebo) checks. First, following Imbens and Lemieux [2008], we examine whether the reduced form estimates of equation (1) as used in section 6.3 yield jumps in the age-expenditure curves at cut-off points other than the early retirement age. Second, we examine whether the reduced form models estimate jumps at the early retirement age for the non-working populations.

The first panel of Table 11 displays the reduced form coefficient estimates of being above the official early retirement age in our models with constant early retirement age. As expected, most parameter estimates are statistically significant at the 1% level. The second panel shows reduced form estimates of being aged above alternative cut-off points (four years before and after the early retirement age) in models with possible jumps and trend breaks at these age cut-offs. As in the original specification, we always look at +/-3 years of age around these points, and for cut-offs at 53 (women) and 56 (men) we restrict the sample to those who worked three years earlier. Due to the retirement patterns, we do not make such restrictions for the cut-offs at 61 (women) and 64 (men). None out of the 24 parameter estimates in the second panel are statistically significant at the 1% level, which is in favour of our specification, as it indicates that the estimated jumps at the early retirement age in the first panel are likely to stem from retirement itself.

Finally, the third panel of Table 11 presents reduced form estimates of being above the early retirement age in the same models as in the first panel, but estimated on the non-working populations (i.e. who were not working three years before the early retirement age and did not change work or retirement status in the following six years). Again, we look at six-year wide age windows. Only two out of the 12 estimates are significant at the 1% level, suggesting that the non-working population does not experience jumps in health expenditures at the early retirement age.

Table 11: Reduced form estimates of being aged above different cut-off points and for working and non-working populations

Working	Gender	Cut-off	prob. of positive outpatient	log of positive outpatient	prob. of positive inpatient	log of positive inpatient	prob. of positive drugs	log of positive drugs
Y	F	57	-0.021***	-0.050***	-0.0093**	-0.042	-0.012***	-0.0061
			[0.0024]	[0.0098]	[0.0025]	[0.034]	[0.0021]	[0.0075]
Y	M	60	-0.019***	-0.061***	-0.0089***	0.048*	-0.0082***	-0.010
			[0.0024]	[0.0093]	[0.0021]	[0.027]	[0.0020]	[0.0065]
Y	F	53	0.0010	-0.018	0.0029	-0.017	-0.0024	-0.0089
			[0.0029]	[0.014]	[0.0035]	[0.042]	[0.0030]	[0.0102]
Y/N	F	61	-0.0013	0.0016	-0.0006	-0.035*	0.0016	0.0067
			[0.0018]	[0.0067]	[0.0020]	[0.0192]	[0.0015]	[0.0051]
Y	M	56	0.0026	-0.016**	-0.0028**	0.0323	0.0012	-0.0068
			[0.0018]	[0.0072]	[0.0014]	[0.0217]	[0.0016]	[0.0050]
Y/N	M	64	-0.0007	0.0056	0.0003	-0.0069	-0.0003	0.0065
			[0.0015]	[0.0063]	[0.0016]	[0.013]	[0.0013]	[0.0043]
N	F	57	-0.0071***	-0.032***	-0.0052	-0.024	-0.0028	-0.0079
			[0.0022]	[0.0092]	[0.0033]	[0.017]	[0.0018]	[0.0069]
N	M	60	0.0025	0.019**	-0.0004	0.012	0.0032*	0.012**
			[0.0020]	[0.0083]	[0.0024]	[0.015]	[0.0017]	[0.0060]

Robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1
Working (Y/N): worked three years before the cut-off age
Gender: female (F) or male (M)
Sample: see in text. Model: see at Table 4