Red Herring in the Vistula River: Time-to-Death and Health Care Expenditure

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Abstract

The aim of the article is to quantify the role of time-to-death in health care costs which increase with age. By combining a detailed break-down of population-wide data about publicly financed health care expenditure with the population and mortality data we are able to estimate a flexible model of health care expenditure. Poland is a country that has universal health insurance coverage. We show that introducing time to death into the model enables the health care costs model to fit data perfectly. We show the estimated paths of mean cohort expenditure, depending on the age of death. Our results confirm that the cumulative costs of health care over the life cycle hardly vary based on the age of death. As a result, the effect of rising life expectancy has a minor impact on HCE. However, the transitional effect of the changing age structure of the population should affect the aggregate health care expenditure.

Keywords: healthcare expenditure, ageing, red herring

JEL Classification Numbers: H51, I12, I18, J14

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1 Introduction

The article proposes a novel approach to modelling the role of time-to-death as a determinant of health care expenditure. The idea is to estimate the mean cohort path of health expenditure (HCE) using the data about the expenditure per survivor and per decedent combined with the data about population and mortality. The specification is based on the evidence of an exponential increase in health care more than one year before death (Yang et al., 2003; Zweifel et al., 2004; Weaver et al., 2009; Atella and Conti, 2013). We propose a formal econometric model of health expenditure exponentially increasing before death with stable and varying parameters.

Our estimates provide a detailed insight into the health expenditure pattern after including time-to-death as an explanatory variable. Firstly, we show the life-cycle paths of health care expenditure obtained with different models. Subsequently, the cumulative HCE depending on the age of death is calculated. The comparison of the cumulative HCE conditional on the age of death provides a clear answer to the non-trivial dependence of HCE on age.

We start with a careful description of the demographic data and health expenditure data. Then the basic and extended econometric specification is presented. The model is estimated using multiple forms and different data to ensure that the findings are robust.

2 Literature

As pointed out by Michel and Robine (2004) the observed increase in life expectancy in developed countries is due to a reduction in mortality at later ages. There are three main hypotheses about changes to the morbidity pattern associated with the rise in life-expectancy: expansion, compression and postponement (Fries, 2002; Kramer, 1980; Payne et al., 2007). The first assumes that rising longevity would cause the duration of the illness or disability to be prolonged. The compression hypothesis states that the duration would shrink, while the final hypothesis states that the disability or illness would merely be moved forward to a later period of life. Recent evidence has delivered a mild argument in favour of the compression hypothesis (de Meijer et al., 2013; Martin et al., 2010; Christensen et al., 2009; Lafortune and Balestat, 2013; Parker and

Thorslund, 2007).

There are two main approaches to empirically investigate the stability of the relationship between age and health care expenditure as life expectancy rises. One includes variables associated with the age structure, life-expectancy and death-rates into the macro-panel regressions based on the data aggregated at country level. The results are mixed because a lot of information is lost at the aggregation stage and the results suffer from the common problems of omitted variables, endogeneity and a high autocorrelation of demographic variables(Lago-Peñas et al., 2013; Michel and Robine, 2004). Karlsson and Klohn (2011) use county-level data to show that the mortality rate and age structure of the population have a strong effect on local health expenditure.

More precise results are obtained with the second approach i.e. using micro-data models. Based on individual data from Switzerland, Zweifel et al. (1999) have contested the hypothesis of the stability of the relationship between age and HCE. They postulated that it is a false relationship and therefore called it a *red herring*. They concluded that HCE did not rise with age per se, but that it is an effect of the rising death rates and shortening time-to-death at older age groups. As a result, it is time-to-death and not age that is the main driver of HCE. They showed that when time-to-death is included (as quarterly dummies), age loses its significance as a regressors on health care expenditure. However, the idea of stronger influence on HCE from time-to-death compared to time-from-birth stems from Fuchs (1984). Zweifel et al. (2004), Felder et al. (2010), and Werblow et al. (2007) have confirmed the findings with a more robust methodological approach. Their results about the consequences of population ageing on HCE are strong: the increasing average life expectancy has no effect on HCE.

Recent evidence has supported the *red herring* hypothesis, at least to some extent. Most studies from various countries agreed with the view that both age and time-to-death play a role in explaining health expenditure. Yang et al. (2003) have delivered an in-depth descriptive analysis of US health expenditure data. They have shown that spending rises significantly in the last 6 months of life. The proximity to death is a good predictor of inpatient care, although age better explains the use of long-term care. Karlsson and Klohn (2011) have provided evidence about long-term care based on Swedish administrative data. They have found strong support for *red herring* but also the important effects of age. Weaver et al. (2009) have shown that the proximity to death is the main driver of long-term care expenditure in the US. For Italy, Atella and Conti (2013) have provided evidence that the time to death is a good predictor of outpatient care, but does not eliminate the influence of age. Seshamani and Gray (2004a) show that the difference in inpatient care expenditure remains significant, even 13 years before death, but that age is also important.

By widening the analytical approach, Shang and Goldman (2008) have tried to deal with the issue by calculating life expectancy based on socio-economic characteristics. They have confirmed that both age and proximity to death are important drivers of HCE in the US. de Meijer et al. (2011) elaborated in detail about health care costs after surviving specific diseases in the Netherlands. They have analysed the main causes of LTC, especially the histories of specific diseases, disability and co-residence. They found that that when checking disability and co-residence, the time to death (TTD) loses its significance for LTC while age remains significant. Polder et al. (2006) have offered a new insight into understanding death determinants in the Netherlands. They found that the range of different causes of death within age groups is much larger than between groups. This contradicts the view that inter-disease differences drive the age dynamics of costs. In particular, the most expensive death cause is cancer and cheapest are heart diseases (Lubitz et al., 2003). Wong et al. (2011) were determined by analysing hospital expenditure based on the type of disease. They have found that in some cases what matters is age and in others the proximity to death. Werblow et al. (2007) looked into the role of the proximity to death of the various components of HCE in Switzerland. Apart from LTC they found little evidence of age being significant.

There is, however, evidence against *red herring*. Colombier and Weber (2011), Breyer and Felder (2004) and Westerhout (2006) point out that morbidity outweighs mortality as a cost driver and that the reduction of HCE due to TTD is of lesser importance. Dow and Norton (2002) and Salas and Raftery (2001) show some methodological issues concerning endogeneity, collinearity and selectivity which can lead to spurious estimates in the models of individual data. Most of their points were taken into account in the recent evidence without leading to major changes in the results. By weighting the importance of the *red herring* literature for ageing, Westerhout (2006) noticed that the difference between age and TTD models is only important for selected types of HCE. What is more, Breyer and Felder (2004) have shown that including death related expenditure does not influence the HCE projections, as demographics are not the main driver of HCE, but rather the technological progress in medicine. The increasingly intensive concentration of expenditure in older age groups is called the steepening effect. The hypothesis of steepening has received some support in data (Meara et al., 2004; Goldman et al., 2005; Dormont et al., 2006; Christensen et al., 2009; Breyer et al., 2010). The only available international comparison of the role of time to death is the health care cost projections for OECD and the EU countries (OECD, 2006; Przywara, 2010). The former presents the relationship of expenditure per decedent and survivor by age and sex. Poland is included in this comparison. The implications are not studied in detail and are just taken as given for the HCE projections. According to the data, the difference in expenditure per decedent and survivor narrows from a ratio of 30-50 in age group 0-10 to 1.3-3 in the age group 90+ in European countries. We follow the recent strand of literature to look more precisely into the structure of HCE, find the basis for an analysis of 1-year age groups and translate our results into consequences for cumulated life-cycle expenditure.

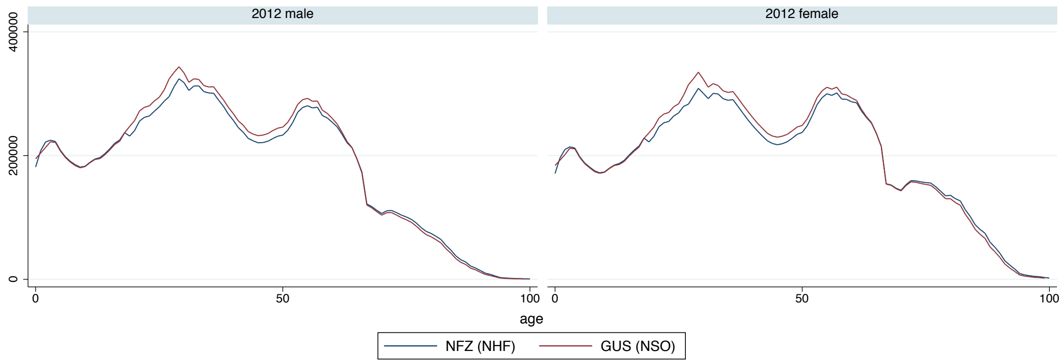
Most research papers with an econometric estimation of the importance of time-to-death do not take into account the falling costs of death for older people (Werblow et al. (2007); Breyer and Felder (2004)). They just estimate the mean effect of time to death for all ages. The following parts show why it is important to take into account the changes in death-dependent costs with age.

3 Data

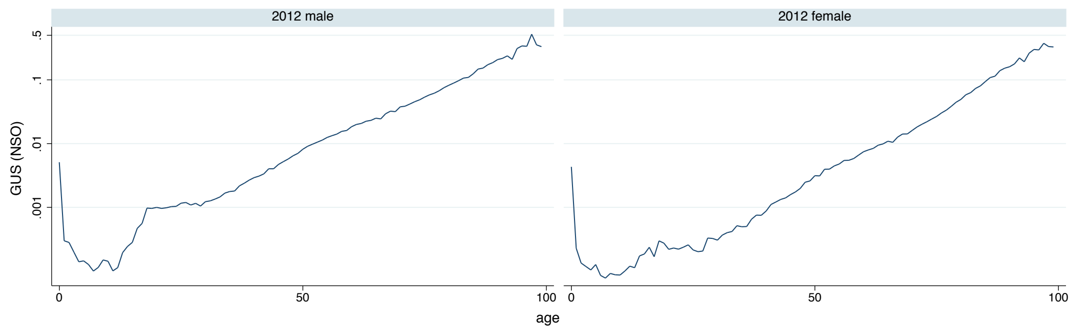
We have combined data from the National Health Fund (NFZ) and the Central Statistical Office (GUS). The data from the NFZ includes health spending with breakdowns for sex, age, type (e. g. inpatient, outpatient, drugs) and information about whether the person died in the given year. The NFZ finances about 60% of HCE in Poland, with an additional 10% financed by other public sources and 30% from out-of-pocket payments or private insurance (stats.oecd.org).

Figure 1: Population and death rate by age and sex in 2012

(a) Population



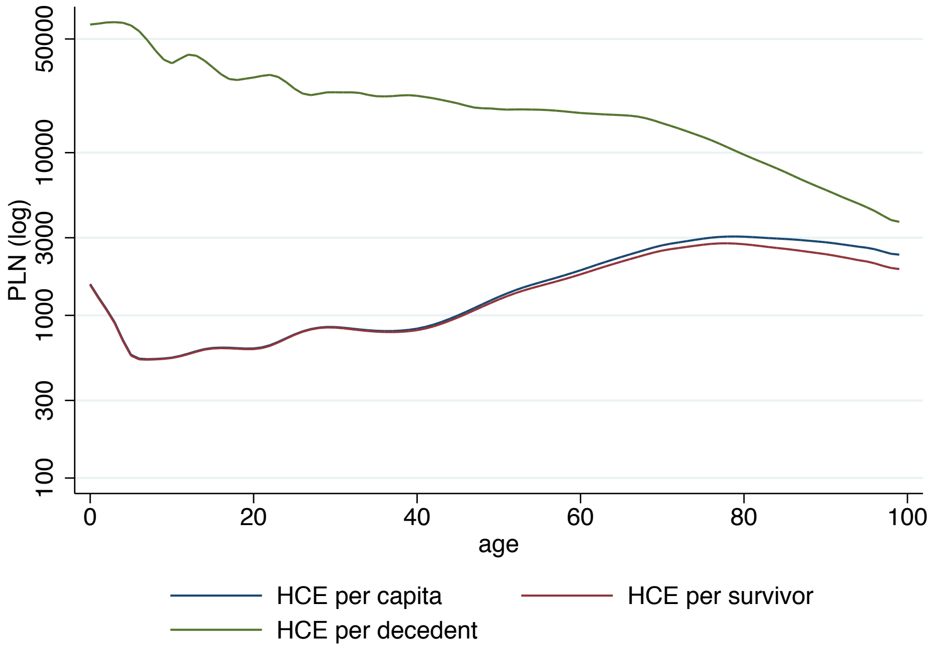
(b) Mortality rate



Source: Own calculation.

The dataset covers almost all NFZ expenditure, only excluding primary health care, which can hardly be attributed to age. In 2012 NFZ expenditure totalled 59.9 billion PLN, while 50.2 billion PLN was ascribed to age. This covers around 38.5 million people. It means that almost 85% of expenditure is covered by the analysis. As a result, most of the public spending and the majority of all spending are covered. Apart from private expenditure, our database also lacks expenditure for long-term care financed by social aid. However, some nursing costs for the elderly and hospices are covered. The GUS data on mortality rates only include mortality rates by age and sex.

Figure 2: Distribution of HCE by age and sex



Source: Own calculation.

The dependence of HCE on age looks trivial when the HCE distribution by age is studied. This distribution forms a well-known inverted S-shape. High expenditure in the first years of life precedes a period of low expenditure which gradually rises after the age of 40. A more careful examination reveals that death-related costs drop during the lifespan. Therefore the rising expenditure with age is a combination of the rising costs for survivors and the rising mortality rates (Figure 2). By looking at gender differences we can conclude that the higher expenditure for females is a result of pregnancy related costs and the fact that in general they pay more attention to their health. The per capita expenditure equals the mean spending of a 56-year old person (1700 PLN, 400 EURO, 2012). They vary 6-fold from the most to the least expensive age groups. Due to the dominating share of survivors they mainly influence the shape of the age-profile.

4 Methods

In order to show the importance of the distinction between the health expenditure models that are age-dependent and focus on the proximity-to-death, we present the accumulated life-cycle health expenditure profile, which is dependent on the age of death. In the case of an age dependent model the profile can be calculated easily from the data. Contrary to this, in the case of the *red herring* model, some assumptions are needed to calculate these profiles.

The profile of HCE per survivor (see Figure **??**) is a product of three phenomena: (1) change of costs independent from the time of death, (2) the change of death-related costs and (3) the distribution of time-to-death in the population. We only observe the last one directly, through the age-dependent mortality rates. However, taking into account the fact that death-related costs rise exponentially as the proximity to death increases, as shown by Atella and Conti (2013), Colombier and Weber (2011), Breyer and Felder (2004), we can formulate the following model of health cost per survivor h^s_a:

$$\begin{aligned} h^s_a &= S_{a,T}\,\alpha + \sum_{t=1}^{T}\left(D_{a,t}\,c_{a+t}\,\beta e^{\gamma(t-1)}\right) + \epsilon_a \\ &\epsilon_a \sim N(0,\mu) \end{aligned} \tag{1}$$

The c_a denotes the health costs incurred in the calendar year of death at age a, $D_{a,t}$ is the share of cohorts a dying at the age of $a+t$, subject to them surviving to the age of a and given the mortality rates, α stands for health care costs incurred each year before T years preceding death, β is a multiplier of the cost in the calendar year prior to the year of death, γ is a parameter of the speed of the increase in costs before death and T states how many years before death the death-associated costs disappear.

In the base specification, we assume that the costs independent of death α remain constant throughout the lifetime. The parameters β and γ are crucial for the results and their interpretation is important. The values of $\beta > 1$ mean that expenditure in the calendar year previous to the year of death are higher than in the calendar year of death. This situation does not contradict the assumption that costs decline monotonically before death as the expected time-to-death in the last calendar year of life is just 6 months. The lower values of parameter γ mean that the death-related costs are more concentrated before death. For example, $C = -0.05$ means that the expenditure five years before death is equal to 0.78β of the expenditure in the year of death. Declining γ to -0.2 reduces the costs 5 years before death by more than half, to 0.37β.

H^s_a, $D_{a,t}$ and c_{a+t} are calculated directly from the data. The model is left with 4 estimation parameters α, β, γ, T. The parameters α, β, γ are estimated using non-linear least squares for values of $T \in \{1, 2, 3, .., 23\}$. For T the value which minimises the residual term is chosen.

Morbidity patterns show that for those dying at an older age, the cost structure may change smoothly. More importantly however, the costs independent of death may also change with age. We therefore allow for the dependency of parameter α on age. The logistic transformation function does allow for this dependence:

$$\alpha(a) = \alpha_1 + \alpha_2 / \left(1 + e^{\alpha_4(a-\alpha_3)}\right) \tag{2}$$

Additionally, to exclude the possibility of a jump in costs before death, we link two functions in the period T before death using the restriction:

$$c_{a+T} \beta e^{\gamma T} = \alpha_T \tag{3}$$

In effect we achieve a restriction on γ

$$\ln(\gamma) = \ln \alpha_T - \ln c_T - \ln \beta_T - T \tag{4}$$

The second model, which allows the survivor costs to be age-dependent, is estimated using the equations 1, 2 and 4.

The last two models allow us to relax the assumptions on the constant relationship of the HCE one year before death to the expenditure in the year of death through the logistic transition of the β with age. Model no 3 therefore looks like this:

$$\begin{aligned} h_a^s \quad &= \quad S_{a,T}\,\alpha + \sum_{t=1}^{T} \left(D_{a,t}\, c_{a+t}\, \beta e^{\gamma(t-1)}\right) + \epsilon_a \\ \alpha(a) &= \alpha_1 + \alpha_2 / \left(1 + e^{\alpha_4(a-\alpha_3)}\right) \\ \beta(A_d) &= \beta_1 + \beta_2 / \left(1 + e^{\beta_4(A_d-\beta_3)}\right) \\ \ln(\gamma) &= \ln \alpha_T - \ln \beta_T - \ln c_{a+t} - T \\ \epsilon_a &\sim N(0, \mu) \end{aligned} \tag{5}$$

Finally, in order to check the relevance of including death-related costs in the equation

c_{a+t}, we unlink the *beta* parameter from the expenditure in the last year of life. Model no 4 is specified as follows:

$$
\begin{aligned}
h_a^s \quad &= \quad S_{a,T}\,\alpha + \sum_{t=1}^{T}\left(D_{a,t}\,\beta e^{\gamma(t-1)}\right) + \epsilon_a \\
\alpha(a) &= \alpha_1 + \alpha_2/\left(1 + e^{\alpha_4(a-\alpha_3)}\right) \\
\beta(A_d) &= \beta_1 + \beta_2/\left(1 + e^{\beta_4(A_d-\beta_3)}\right) \\
\ln(\gamma) &= \ln\alpha_T - \ln\beta_T - T \\
\epsilon_a &\sim N(0,\mu)
\end{aligned} \tag{6}
$$

In models 3 and 4 the number of estimated parameters rises from 5 to 10 in order to make the relationship between the costs of death and current expenditure more flexible. The exponential model is nested in the model with parameter transition. For instance, parameter values $\alpha_2 = \beta_2 = 2, \alpha_4 = \beta_4 = 0$ simplify the transition model to the exponential case.

All the models are estimated for cohorts aged 12-100 in 2012, which results in 89 observations. The younger cohorts are excluded due to there being a different pattern of HCE among children which is hardly related to ageing, whereas cohorts older than 100 are excluded due to the small number of observations. We use non-linear least squares to estimate the models. In order to check the robustness we also report bootstrapped and jack-knifed standard errors.[1] The bootstrap and jack-knife methods are simulations, non-parametric statistical methods that allow standard errors to be calculated with no distributional assumptions. They also provide a check of the stability of estimates in terms of outliers and numerical stability (Efron, 1981; Efron and Gong, 1983; MacKinnon, 2006). The models are fit to GUS data on deaths and population. All models are estimated for each $T \in \{1, 2, 3, .., 23\}$, and the value with the best fit in terms of mean square error (MSE) is chosen. MSE is chosen as a criterion as the number of parameters do not change with T, only the functional specification changes. Consequently, Bayesian or Akaike information criterion would lead to the same decision on model selection.

[1] Bootstrap is run for 500 trials and jack-knife for all 89 observations

5 Results

We have estimated four models in order to explain the evolution of costs per survivor by age. Consecutive models allow a less strict assumption about the relationship between age, proximity to death and HCE. The models differ in terms of the specification and number of parameters. The first model is specified with HCE exponentially rising before death (see equation 1), the second by the costs per survivor transition with age (see Equations 1, 2 and 4), while the third one allows for the transition of the relationship between expenditure in the last year of life and the year before (see Equation 5). The final model is more flexible in terms of the costs in the year of death (see Equations 6). For each model we report the parameters' estimates, models fit, transition logistic functions of the parameters, life-cycle paths of HCE and the cumulated life-span costs dependent on the age of death.

Table 1: Models estimates

	Model 1	Model 2	Model 3	Model 4
β	0.862**	0.790**		
	(-0.014;-0.03;-0.03)	(-0.053;-0.105;-0.119)		
α	-0.120**			
	(-0.005;-0.008;-0.008)			
γ	486**			
	(-19.0;-13.1;-12.7)			
β_1			1.029**	40,835**
			(-0.042;-0.037;-0.043)	(8,961;5,141;4,831)
β_2			-0.471**	-35,742**
			(-0.053;-0.045;-0.051)	(8,710;4,901;4,642)
β_3			98.164**	79.4**
			(-0.48;-0.4;-0.46)	(-0.66;-0.32;-0.36)
β_4			0.337**	-0.236**
			(-0.037;-0.026;-0.027)	(-0.023;-0.012;-0.012)
α_1		2,875**	2,344**	2,062**
		(-70;-106;-119)	(-94;-72;-72)	(-229;-154;-141)
α_2		-2,294**	-1,796**	-1,506**
		(-76;-108;-122)	(-96.9;-77.3;-77.9)	(-227.3;-154.9;-140.8)
α_3		60**	71.5**	59.9**
		(-0.463;-0.496;-0.545)	(-0.621;-0.467;-0.479)	(-0.484;-0.433;-0.46)
α_4		0.159**	0.132**	0.156**
		(-0.009;-0.006;-0.007)	(-0.005;-0.005;-0.005)	(-0.008;-0.007;-0.007)
R^2	1	1	1	1
Adjusted R^2	1	1	1	1
Root MSE	110.85	90.16	25.96	29.6
Deviance	1,087	1,048	823	847
N	89	89	89	89
T	16	12	11	1

* $p < 0.05$; ** $p < 0.01$
standard errors in parenthesis,
consecutively: classical; bootstrapped; jackknifed.

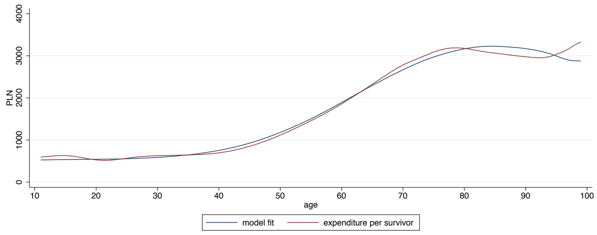
Estimates of all the models are presented in Table 1. Although we have estimated the models for time-to-death from 1 to 23 years for each specification, only the best-fit is reported. In addition to classical errors, bootstrap and jack-knifed standard errors are also reported consecutively in the parenthesis. The simulation-based standard errors also serve as checks for outlier-sensitivity and the numerical stability of the estimates. The model fit for age-groups is presented graphically (see Figures 3). All interpretable parameters (α and β), which are calculated using primary estimates, are also presented graphically with 95% levels of confidence (see Figures 4 and 5). The estimates allow the age-of-death dependent HCE paths and accumulated life-cycle HCE to be calculated (see Figures 6 and 7). These are the most important results as they allow a direct comparison of the consequences of ageing on HCE through the lens of estimated models. The results for each and every model are described briefly below.

The exponential model (1) enables us to replicate the general patterns of costs per survivor (see Table 1 and Figure 4a). The results show quite similar costs one year before and in the year of death (86% of the previous year's expenditure) and per survivor expenditure of 486 PLN. The γ parameter of 0.12 means that the expenditure 10 years before death constitutes 26% of the previous year's expenditure. The life-cycle paths and cumulated life-cycle HCE drop strongly after the age of sixty (see Figures 7a and 7). According to these results, dying at an older age even leads to a significant reduction in costs. These results are driven by the fact that the costs of survivors do not change with age and death-related costs start to appear a long time before death (16 years). These strong results are not supported by evidence from other research (Breyer and Felder, 2004; Werblow et al., 2007; Atella and Conti, 2013). We therefore present the results of the models with a more flexible functional form. They allow for more realistic assumptions as well as data which is a better fit.

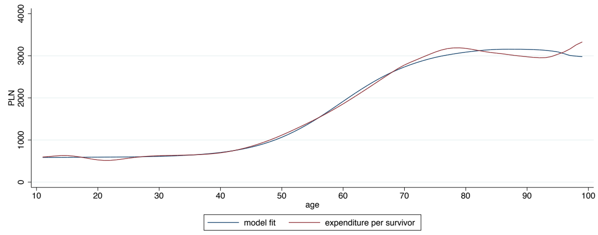
Relaxing the assumption that the costs for survivors (Model 2) are constant significantly improves the fit-to-data (see Figure 4b and Table 1). The mean square error drops by 20% compared to the exponential model. The best fit considers data for 12 years before death. The β estimate is of similar magnitude (0.79) as before. However, the transition of α is significant. It rises from 500 PLN before the age of 40 to almost 3000 PLN by the age of 80 (see Figure 5a). This is the most important feature in terms of the main results of accumulated life-cycle health expenditure, which are dependent on the age of death. In the case of the model which only uses variable costs per survival, we obtain almost no time-to-death dependence (see Figure 7). The whole rise in costs with age is due to the rise of costs which are independent of the time-to-death. This provides the only argument against the *red herring* hypothesis.

Figure 3: Models fit

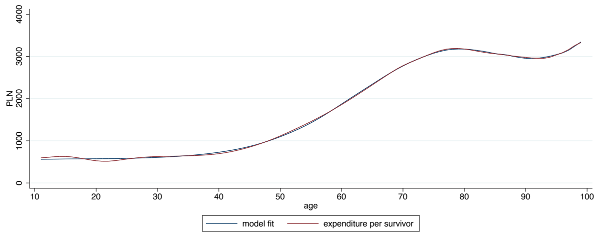
(a) Exponential models (1)



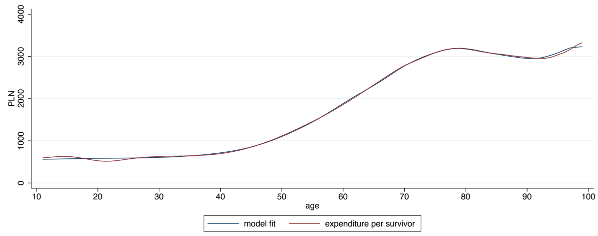
(b) Model with transitional α (2)



(c) Model with transitional α and β (3)



(d) Model with transitional α and β and unbounded death costs (4)



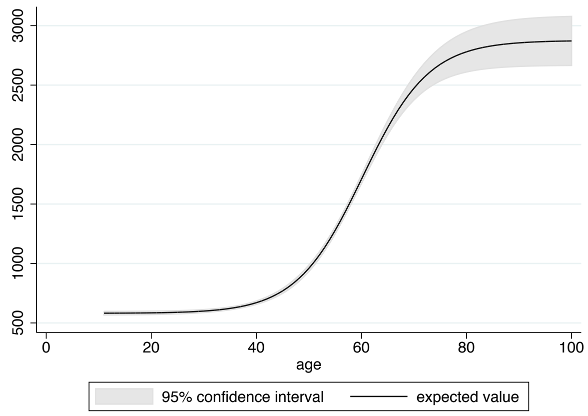
Source: Own calculation.

Finally we allow for variability in the HCE in the year of death and the year before (Model 3). A smooth logistic change is also allowed. As a result, the root mean square error drops by a factor of 3.6 (see Table 1 and Figure 4c). An almost perfect fit is achieved with costs starting to rise 11 years before death. The β parameter slows down the drop in costs per decedent and flattens the cost profiles for people over the age of 80 (see Figure 6a). Simultaneously, the rise in costs per survivor moves to a later point in life (see Figure 5b). The cost per survivor at the age of 90 drops to 2300 PLN compared to 2800 PLN in the previous model. The model with best fit-to-data shows stability of the cumulated life cycle HCE for those dying between the age of 70 and 90 (see Figure 7).

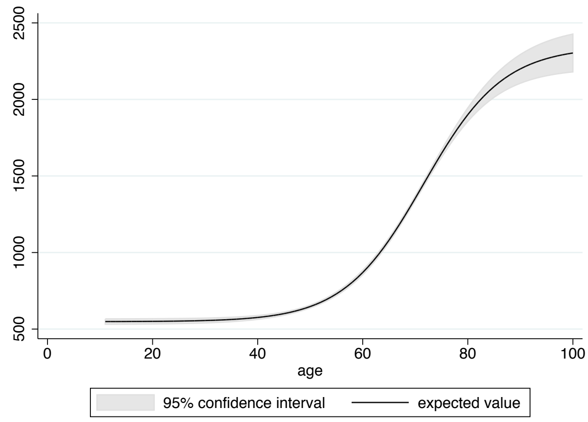
Disconnecting the evolution of HCE before death from the costs incurred in the last year of life (Model 4) worsens the model fit slightly (see Table 1 and Figure 4d). The evolution of costs per survivor hardly changes and drops from 2300 PLN to 2000 PLN for 90 year olds (see Figure 5c). However, the paths of costs change significantly - with the rise in costs before death (up to 1 year) and the sharp rise before death being signif-

Figure 4: Parameter α (per survivor expenditure) transition with age of death

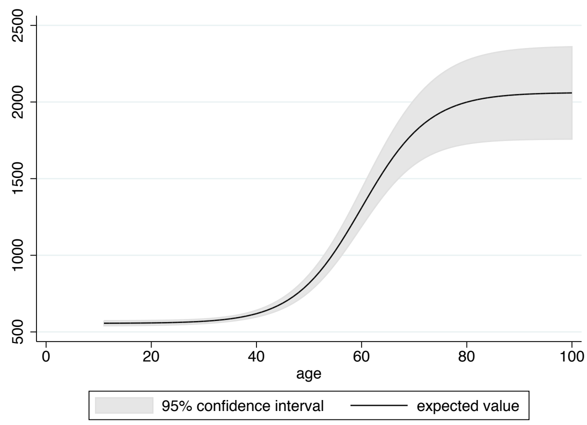
(a) model with transitional α (2)



(b) model with transitional α and β (3)



(c) model with transitional α and β and unbounded death costs (4)

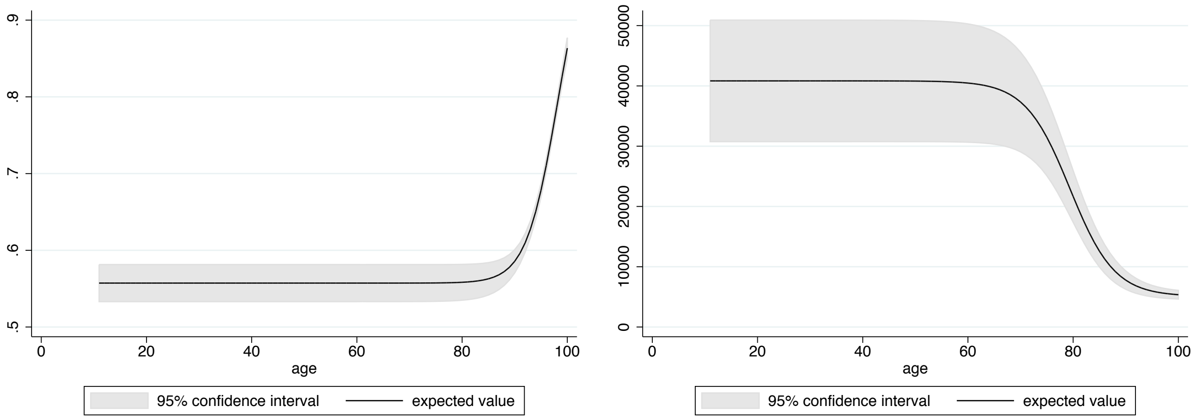


Source: Own calculation.

Figure 5: Parameter β (two-years before death) transition with age of death

(a) model with transitional α and β (3)

(b) model with transitional α and β and unbounded death costs (4)



Source: Own calculation.

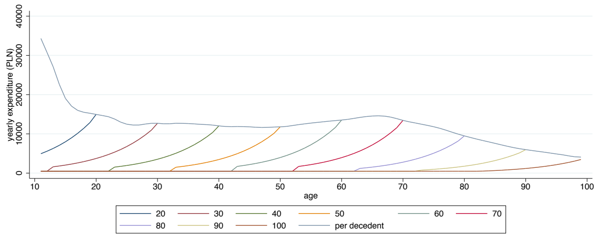
icantly reduced, especially for people under the age of 80 (see Figures 7d, 6b). Despite the different paths and the estimations of the period of the increase in costs before death, the cumulated life-time expenditure of the model with HCE that was linked and unlinked to the costs incurred in last year of life hardly differs for the period 70-90 years old (see Figure 7). This is an argument in favour of the weak version of the red-herring hypothesis: for the period 70-90, the change in population structure does not increase the aggregated HCE. However, an increase in the number of people over the age of 90 would lead to slight increase in HCE.

The joint analysis of all models leads to the conclusion that fitting the model to leads is probably consistent with the literature life-cycle paths of HCE and those cumulated over the whole lifespan. The models with the most flexible form show support for the *red herring* hypothesis – that the HCE is strongly related to the age of death, and ageing should have much smaller effects on HCE than those estimated by the naive model of age-specific HCE.

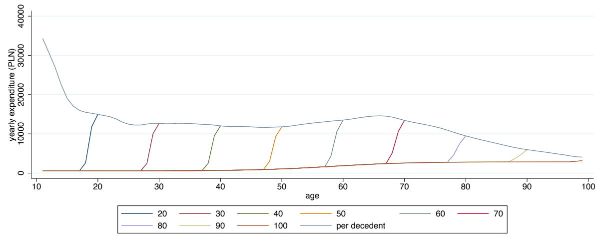
All models which include the proximity-to-death show that the difference in lifetime spending on people dying at the age of 20 and 100 is 2-3-fold and not 10-30-fold as suggested by the naive model of age-specific costs. Therefore the inter-generational and inter-age group transfers through health systems are much smaller than would be expected. The models with the most flexible assumptions and based on more reliable demographic data from GUS show that ageing would increase the HCE, but to about half the level that would have been expected in the naive model of age-dependent costs.

Figure 6: Life-cycle paths of costs depending on age of death

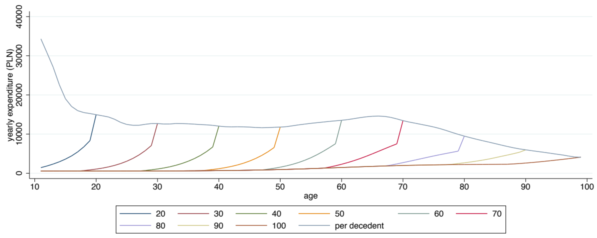
(a) exponential model (1)



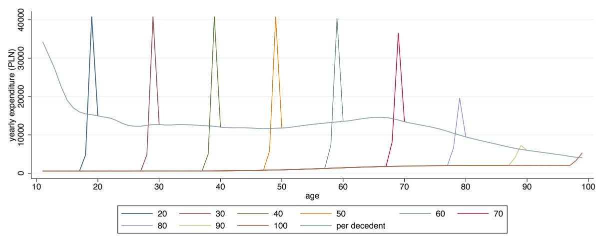
(b) model with transitional α (2)



(c) models with transitional α and β (3)



(d) model with transitional α and β with unbounded death costs (4)

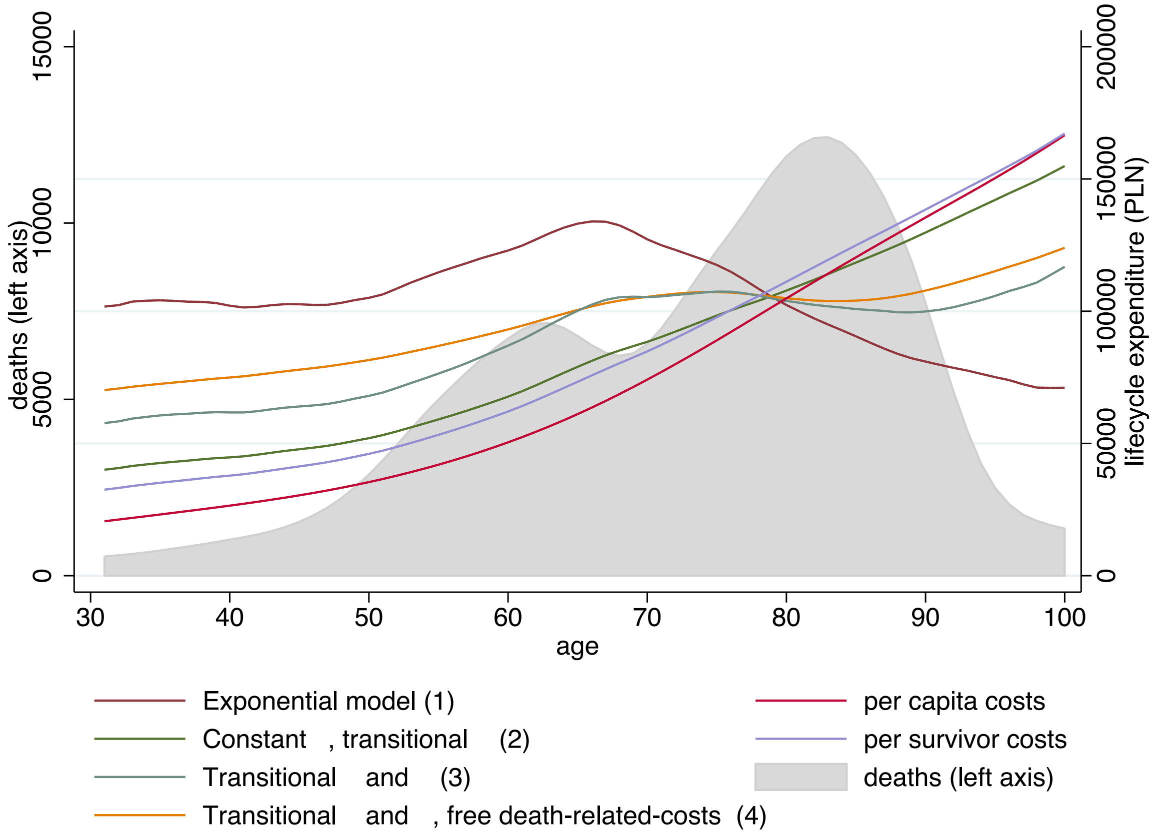


Source: Own calculation.

Model 3 shows the best fit and therefore delivers the most reliable HCE-age profiles and cumulated life-cycle expenditure. It shows that HCE starts rising 12 years before death. This is consistent with the results from studies from other countries (Seshamani and Gray, 2004b; Wong et al., 2011). Other models show a shorter and longer period of rise of HCE before death (1-16).

It is important to underline that the population data matters a lot, the model selection strongly influences the results and the best-fit models show evidence in favour of *red herring*: the costs would increase slower than when purely driven by age, with a huge chance of increasing life expectancy to be cost neutral. The costs of death are mostly driven by people younger than 60, for whom the death-related-costs dominate all other costs during their life-time. From the age of 70 the death related costs start to drop and for those at the age of 100 they do no differ significantly.

Figure 7: Cumulated life-cycle costs depending on age of death



Source: Own calculation.

The obtained results sharpen the findings of (Spillman and Lubitz, 2000) for the USA. They found that cumulative acute health care expenditure financed by Medicare for those over the age of 65 do not rise significantly with age. It is the long-term costs that are strongly affected by ageing. The results of the NFZ data reinforce these conclusions since the whole-life health care expenditure can be examined in Poland due to the near universal coverage.

6 Conclusions

People who live longer generate less costs when approaching death. The cumulative life-cycle health care expenditure only rises slightly with the age of death. If people who live longer were to follow the age-pattern in morbidity of the current generations, the effect of ageing would only cause a slight increase in the aggregated health care costs. As morbidity is a dynamic process, the observed relationship between health care costs, age and the proximity to death may change in the near future.

A new method has been applied to the analysis of aggregated data of health care expenditure. More precise data is needed to perform a robustness check. In particular, data about expenditure 2-5 years before death would enable the crucial assumptions on the stability of functional form of cost acceleration before death to be tested. This data is, however, unavailable in Poland to date. A careful analysis of morbidity and the costs of treatment also remain an open field for research.

The impact of ageing on public finances through health care expenditure is a great challenge for public finance, as HCE is financed publicly to a large degree and is dependent on age. For example, the estimates obtained using generational accounts (Auerbach et al., 1991; Jablonowski et al., 2011) or national transfers accounts (Mason et al., 2006) expect HCE to explode, in line with literature postulating that the strong relationship of HCE to age is a *red herring*. As I have shown, it is the combined effect of age and proximity-to-death that drive health expenditure. The results obtained are in line with recent evidence that supports the weak version of the *red herring* hypothesis (Breyer et al., 2010; Spillman and Lubitz, 2000), i.e. that the pressure on health care expenditure other than long-term care due to ageing is rather low, and much lower than previously feared. The fact that people are working longer and the mild increase in health care expenditure could counterbalance the expected negative effects of ageing on public finances.

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