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ORIGINAL PAPER

Ageing and health care expenditure in EU-15

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Abstract The purpose of this paper is to investigate the relationship between ageing and the evolution of health care expenditure per capita in the EU-15 countries. A secondary purpose is to produce estimates that can be used in projections of future health care costs. Explanatory variables include economic, social, demographic and institutional variables as well as variables related to capacity and production technology in the health care sector. The study applies a co-integrated panel data regression approach to derive short-run relationships and furthermore reports long-run relationships between health care expenditure and the explanatory variables. Our findings suggest that there is a positive short-run effect of ageing on health care expenditure, but that the long-run effect of ageing is approximately zero. We find life expectancy to be a more important driver. Although the short-run effect of life expectancy on expenditure is approximately zero, we find that the long-run effect is positive, so that increasing life expectancy leads to a more than proportional, i.e. exponential, increase in health care expenditure.

Keywords Health care expenditure · Ageing · EU-15 countries

JEL Classification H2 · H51 · I1 · J14

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Introduction

The last four decades have witnessed an increasing number of international analyses of health care expenditure. Among these studies, a distinction should be made between analyses based on individuals as unit of observation (micro-level studies) which are usually restricted to a single country, and analyses based on aggregated health care spending with countries as the unit of observation (macro-level studies). Micro-level studies are often used to make predictions of health care spending, based on assumptions on age-specific utilization rates, while macro-level studies are used to find economic, demographic and institutional determinants of expenditure using historical data. While micro-level studies are based on the demand side alone, macro-level studies allow both demand and supply factors to be explored.

Various determinants from both the demand and supply sides have been studied in previous macro-level work. Among them are economic, demographic, social, behavioural and institutional variables. Attention has been paid, in particular, to the impact of Gross Domestic Product (GDP) per capita on aggregate health care expenditure [1, 2]. Of other determinants, the age composition of the population has been used in many studies. The term 'ageing' is used below as synonymous with an increased share of the population being defined as elderly, defined as 65 years or more. Various conclusions have been drawn from these previous macro-level studies with respect to the effect of ageing on health care expenditure per capita when including both demand and supply side variables. Most of these studies that were based on OECD panel data spanning 20–30 years and covering about 20 countries showed almost no effect of ageing or age composition.

While the increasing absolute number of elderly—or any other age group—will inevitably increase total health care expenditure, the expenditure per capita per year will not necessarily increase, although it is a common belief that it does. This consideration should not be mixed up with the question of whether or not an increasing share of elderly citizens will impose an extra "burden" on the younger generations through, e.g., increasing taxation to pay for their health care. Thus, it has been assumed that as individual health care expenditure generally increases by age, health care expenditure *per capita* can be predicted to rise with an ageing population. This common sense type of reasoning has not been supported unanimously in previous studies though for at least three reasons.

First, health care expenditures are determined not only by demographic factors, but also by non-demographic variables, such as technology, economic and institutional characteristics [2], which are influenced by political and managerial decisions [3]. A partial view on the demand pressure may therefore bias and exaggerate the effects of ageing on actual spending.

Secondly, increasing health care costs by increasing age may be explained to a certain extent by the lack of accounting for high cost at the very end of life [4–6]. It has been found that health care costs increase rapidly by increasing proximity to death. Consequently, if death is postponed, the high costs will occur at a later age. The associated increase in the ratio of healthy to unhealthy people may mean that, as a result, health expenditure *per capita* declines.

Thirdly, even if ageing means higher costs for the older population group, total budgets might simply be re-allocated within a certain overall budget ceiling for health care expenditure—at least in countries where there is a strong political constraint on health care expenditure as is the case in most EU countries [7, 8].

Longevity and health scenarios

Various scenarios for longevity and health status and the derived costs of health and long-term care have been formulated. In an optimistic scenario, morbidity is compressed, that is, while life expectancy is assumed to have reached a maximum, the time in good health increases. As a consequence, the time spent in bad health and in need for health care at the end of life becomes shorter. This has been termed the "compression of morbidity" scenario [9]. In contrast, in a pessimistic scenario, life expectancy is assumed to increase but with the age-specific risks of health problems remaining constant. This implies that the time spent in bad health increases with rising life expectancy—the "expansion of morbidity" scenario [10, 11]. Combinations of these two scenarios are possible, of course. Thus, Manton [12] formulated a "dynamic equilibrium hypothesis" implying that longevity increases result in years with good life ("healthy ageing" scenario) [12]. As a consequence of this scenario, the number of years spent in bad health is constant. Obviously, the "dynamic equilibrium" hypothesis is associated with an extreme version of the death-related cost hypothesis as both imply that longevity gains are translated into years of good health [13]. Over time, a country may change from a state of expansion to reduction in morbidity [14]. These scenarios may be used as a frame either for interpreting empirical results or for forecasting purposes. Evidence from seven countries among the EU-15 countries suggests that the "healthy ageing" scenario is valid for these countries [28], and one may hypothesize that this is also true for the rest of the fifteen countries. However, as the healthy ageing scenario is a combination of an optimistic and a pessimistic scenario, there is no unambiguous indication about the net effect of ageing on health care expenditure.

Purpose and hypotheses

It appears from previous studies that health care utilization can be explained by both demand and supply side variables [2]. Hitherto a large number of studies have been performed on a micro level with inadequate consideration of the restrictions on utilization due to supply factors that may take place when demand shifts due to demographic changes. To incorporate supply side variables better, the study needs a design that allows for variation in supply side variables. Such variation exists between each country's health care systems. Several studies have already been made (cf. reviews by [1, 2]), but none based exclusively on EU countries. Our study adds to previous practice in at least two methodological respects. We address specific problems related to stationarity problems, and we treat dynamic aspects of the inter-relationships in the data appropriately.

The primary purpose of the present paper is to investigate the relationship between income, ageing and the development in health care expenditure on a macro level when including economic and institutional variables along with demographic variables. A secondary purpose is to produce estimates that can be used in projections of future health care costs. Projections based on the present study are shown by Khoman and Weale [15].

We explore three hypotheses in this paper.

1. The finding that the income elasticity of demand for health care is unity is robust to an analysis that pays full attention to the dynamic structure of the relationship between health expenditure and GDP.
2. The net long-run effect of ageing on health expenditure is zero as the offsetting influences described in

"Longevity and health scenarios" balance each other out.

3. Health expenditure is influenced by supply factors as well as the demand factors which others have researched. These are discussed further in the data section.

Data

Data used in this paper form a panel data set that covers the old 15 European Union countries (EU-15), Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden and the United Kingdom. The panel spans a time period of 24 years (1980–2003). Attempts to look at a wider range of European countries were frustrated by the incomplete nature of the data available. For the same reason, it was not possible to include a number of variables that *prima facie* were of interest. A full documentation of data can be found in [16]. Definitions and sources of the variables used in the analyses are listed in Table 1.

The dependent variable used in this paper is log total health care expenditure per capita (THCEPC) measured in US dollars in real prices, adjusted for purchasing power parities (PPP). The data are from the OECD Health Data 2004 for OECD countries. The explanatory variables are listed in Table 1. The first group of variables are considered broadly as "demand variables" and includes demographic, economic, behavioural and social variables. The economic variable is log GDP per capita (GDPPC). Our choice of log health care expenditure *per capita* as the dependent variable and log GDP *per capita* as an explanatory variable means that our model is analytically equivalent to one in which the dependent variable is log health care expenditure as a proportion of GDP, and this log proportion is explained by, among other variables, log of GDP *per capita*. With the latter specification, the coefficient on log GDP would simply be equal to the coefficient we estimate less one. Our model includes lags on both log GDP and log health care expenditure *per capita*. The coefficients on these would, of course, be similarly affected.

Demographic variables are included to ensure that the different effects of compression and expansion of morbidity and also mortality-related expenditure can be represented in our model.

Population structure is indicated by the share of the population aged 65–75 (AGE65-74) and the share of the population aged 75 or over (AGE75+). Average life expectancy of men and women at age 65 (AVELE65) offers an alternative demographic variable that reflects planning needs but that also responds to the chances of surviving to extreme old age. The mortality rate (MORTALITY) accommodates the idea that major health

Table 1 Data and data sources

Variable	Description	Data source
THEPC	Total health care expenditure per capita, US$ in nominal prices and adjusted for PPP (in logs)	OECD/WHO
GDPPC	Gross domestic product per capita, US$ in nominal prices and adjusted for PPP (in logs)	OECD
AGE65-74	Population aged 65–74 as a share of the total population	Eurostat
AGE75+	Population aged 75+ as a share of the total population	Eurostat
AVLE65	Average of Life expectancy at aged 65 for males and females	WHO
FLFPR	Female labour force participation rate (% ratio to active population aged 15–65)	OECD
UNEMP	Unemployed individuals as a share of the total labour force	OECD
ALCCON	Alcohol consumption, litres per capita (15+) (in logs)	OECD/WHO
MORTALITY	Number of registered deaths/mid-year population (per 100) (in logs)	WHO
PUHES	Public health expenditure in US$ PPP per capita as a share of the total health expenditure in US$ PPP per capita	OECD/WHO
SALARYGP	Dummy variable for countries with salaried GPs	Christiansen et al. [1]
CAPGP	Dummy variable for countries with capitation payment GPs	Christiansen et al. [1]
CASEHO	Dummy variable for countries with case-based reimbursement of hospitals	Christiansen et al. [1]
COPAYGP	Dummy variable for countries with significant co-payment for GPs	Christiansen et al. [1]
COPAYHO	Dummy variable for countries with significant co-payment for inpatient hospital treatment	Christiansen et al. [1]
FREEGP	Dummy variable for countries with free choice of GP or primary care physician	Christiansen et al. [1]
FREEHO	Dummy variable for countries with free choice of hospital	Christiansen et al. [1]
BEDS	Acute care beds per 1,000 inhabitants (in logs)	OECD/WHO

costs are associated with mortality. To the extent that health expenditure influences mortality, these variables are endogenous. Nevertheless, they are not likely to be greatly affected by short-term movements in health expenditure, and we neglect this possible feedback in our analysis.

The behavioural variable is alcohol consumption *per capita* (ALCCON) while the social variables include female labour force participation rate (FLFPR) and unemployment rate (UNEMP). These variables have been collected from a variety of source including OECD Health Data, WHO and Eurostat. We also looked at the question of including tobacco consumption *per capita,* but the data did not appear to be coherent.

The second group of variables may loosely be labelled "supply variables", which describe characteristics of each country's health care system in the period 1980–2003. This list covers variables that describe institutional factors assumed to affect health care expenditure. These variables are predominantly generated from literature review including the HiT reports from the European Observatory. Changes in the institutional characteristics are traced over time. However, we also use the share of health expenditure which is publicly financed as a variable, since this is also likely to play a role indicating supply conditions.

Institutional variables are included to catch incentives and regulatory factors on the supply side. The variables COPAYGP and COPAYHO indicate whether there is significant co-payment for GP visits or inpatient hospital treatment. It has not been possible to collect data on the exact degree of co-payment, and this qualitative dummy variable is therefore a crude indication for restriction in demand because of co-payment. The dummy variable CASEHO indicate whether countries remunerate their hospital mainly by case-based remuneration. The reference level is countries remunerating their hospitals per diem, fee-for-service and countries with mixed systems. SALARYGP and CAPGP indicate whether countries remunerate their GP mainly by salary (SALARYGP) or by capitation (CAPGP). The reference level is countries remunerating their GPs fee-for-service or with a mix of remuneration schemes. The variables FREEHO and FREEGP indicate whether countries allow patients to make a choice between hospitals (FREEHO) and between GPs (FREEGP). PUHES is share of public finance of health spending.

The last variable in Table 1 is the number of acute care beds per capita which is used as an indicator of the capacity of the health sector to take account of an often found positive relationship between capacity and utilization. We had hoped to include other indicators such as the availability of dialysis machines and nuclear magnetic resonators to supplement indicator for capacity, but coherent EU-wide data were not available. Thus, our model is limited to variables for which we were able to obtain data.

Obviously, the grouping of variables under various headings is not unequivocal; for instance, some variables listed as institutional can also be considered as economic variables. More detailed descriptions of some of the variables are given below. An overview of classification of countries by institutional variables can be seen in [1].

Finally, some dummy variables contained missing values in the beginning (or the end) of the period considered. These were imputed with the first observed (or the last observed) value for the country in question.

As shown in Christiansen et al. [1], more variables characterizing health care systems were collected and included in the initial analysis, but many of the variables were excluded because of strong multicollinearity or omitted because of the country-specific fixed effects models applied. A summary of the key variables is provided by Table 2.

Methods

We use regression methods to explore the relationship between our dependent variable and the various drivers of health care expenditure discussed above. In order to capture dynamic relationships, a first-order autoregressive specification with lagged values of the dependent as well as the explanatory variables is called for. However, as a first step, it is necessary to check whether there are unit roots for health care expenditure and the explanatory variables in our model. We present in Table 3a the results of three tests: the standard augmented Dickey-Fuller test, the Phillps-Perron test and the Im, Pesaran and Shin test [17–19]. The optimal lag length is selected automatically using the Bartlett kernel and the automatic bandwidth parameter suggested by Newey and West [20] with the maximum lag length set to 6. These tests are applied both to the variables in levels, so as to detect the possibility of a unit root, and to the data in first differences, so as to explore the possibility that variables might be stationary only in second differences.

The results show that the tests do not always give the same conclusion. However, there are six variables, THEPC, GDPPC, AGE65-74, AGE75, AVELE65 and MORTALITY for which at least two out of the three tests show p-values of more than 0.05. We treat these variables as being I(1), that is integrated of order 1 [17]. It is no great surprise that the demographic variables as well as the key economic variables turn out to be I(1). Lee and Miller [21] lead us to expect that to be the case. There is no evidence that any of the variables we intend to treat as I(1) are in fact I(2). Table 3b, which explores this, shows that in all cases we give very low probabilities to second-order processes being present.

Table 2 Descriptive statistics of key variables (including interpolates)

	THEPC			GDP			65–74 (%)			75+ (%)		
	1980	1990	2003	1980	1990	2003	1980	1990	2003	1980	1990	2003
Austria	1,592	1,771	2,239	20,897	24,926	28,117	9.61	7.92	8.04	5.88	6.98	7.46
Belgium	1,310	1,765	2,463	20,343	23,732	26,539	8.70	8.16	9.33	5.59	6.65	7.68
Denmark	1,970	2,047	2,495	21,622	24,097	28,745	8.75	8.65	7.79	5.59	6.93	7.02
Finland	1,220	1,863	1,644	19,122	23,754	25,360	7.92	7.69	8.46	3.96	5.61	6.87
France	1,461	2,049	2,563	20,535	23,935	26,770	8.30	7.12	8.56	5.72	6.78	7.73
Germany	1,996	2,278	2,740	23,031	26,830	24,103	9.86	7.67	9.95	5.82	7.23	7.53
Greece	969	1,104	1,818	14,704	14,919	17,797	8.23	7.70	10.79	4.87	5.95	6.73
Ireland	1,068	1,042	2,255	12,733	17,023	32,869	6.88	6.85	6.25	3.83	4.53	4.86
Italy	2,061	1,841	2,096	19,348	22,970	24,664	8.51	8.21	11.54	5.54	6.49	8.02
Luxembourg	1,331	2,020	2,982	22,387	32,997	50,184	8.71	7.33	8.04	4.94	6.05	5.98
Netherlands	1,567	1,870	2,474	20,826	23,335	28,414	6.98	7.36	7.54	4.48	5.42	6.16
Portugal	591	871	1,766	10,640	14,094	17,974	7.44	7.99	9.62	3.72	5.22	7.04
Spain	758	1,139	1,650	14,005	17,094	21,078	6.91	7.90	9.33	3.90	5.52	7.56
Sweden	1,931	2,063	2,342	21,292	24,591	25,987	9.90	9.82	8.28	6.29	7.97	8.86
UK	986	1,287	2,074	17,475	21,386	26,719	9.27	8.77	8.21	5.59	6.90	7.28
Mean	1,129	1,150	1,570	14,365	16,291	19,875	7.72	7.17	8.55	4.55	5.44	6.27
SD	612	706	825	6,007	7,648	9,608	1.69	1.31	1.48	0.96	1.30	1.41

Note: GDP and THEPC were deflated (2000 prices) prior to interpolation

The presence of unit root variables inevitably means that we have to consider how many co-integrating vectors may be present in the model. We do this in the context set out by Breitung [22]. He suggests that the co-integrating vectors can be estimated first of all applying the first stage of Johansen's approach to each member of the panel separately [23]. This allows the data to be purged of dynamic effects, and the co-integrating vectors can be estimated by means of a pooled regression based on exactly the same derived variables as is used in Johansen's approach. He then sets out a test which, given an initial assumption about the number of co-integrating vectors, tests for the significance of at least one extra co-integrating vector. The test is based on orthogonal complements as suggested by Lutkepohl and Saikkonen [24]. It relies on the fact that, if the rank of the co-integration matrix is sufficient, then linear combinations of the co-integration variables calculated using the orthogonal complement of the assumed co-integrating vector should have no explanatory power in the pooled regression.

With six I(1) variables apart from the dummies and a short annual series, there is a pragmatic question about how to explore the rank of the co-integrating space. Breitung presents test statistics for co-integrating spaces of up to rank six against alternatives of lower rank. This gives us six variables with a maximum of five co-integrating relationships.

Breitung's test statistic takes a value of $\lambda = 50.38$. The significance of this is computed by transforming it to a variable which is asymptotically normally distributed with mean 0 and standard deviation 1. With $N = 15$ panel elements, Breitung shows that the transformed variable $\mu = \sqrt{N}\frac{\lambda - 54.33}{9.23}$ has these asymptotic properties, with the constants 54.33 and 9.23 being parameters computed by Breitung for testing the presence of at least one out of a possible five co-integrating relationships. With $N = 15$, the test statistic takes a value of -1.65 which is accepted at a 5% significance level.

We then test whether we can accept the hypothesis that there are at least two co-integrating vectors. In that case, the parameters quoted by Breitung are 35.67 and 7.62, so that the asymptotically normally distributed test statistic is $\mu = \sqrt{N}\frac{\lambda - 35.67}{7.62}$, and since $\lambda = 24.37$ the test statistic is -5.74. We therefore plainly reject the hypothesis that there are two co-integrating vectors against the alternative that there is only one.

With at most one co-integrating vector, we have an alternative test mechanism of estimating an equation with an error correction process in the potentially co-integrated variables, and this is in any case a convenient way of estimating the model. If there is no co-integrating vector present, we will be able to accept the restriction that the coefficient on the error correction term is zero. Or alternatively and equivalently, an equation in the level of log health spending will be found to accept the restriction that the lagged level enters with a coefficient of one and all the other I(1) variables enter only in terms of their differences. The I(0) variables may be present in both levels and differences.

Table 3 Panel unit root tests for the EU-15

Variable	Im et al. [18]	Augmented Dickey-Fuller [17]	Phillips and Perron [18]
a: Variables in levels/log levels			
Log THEPC	4.11742	21.0061	28.5812
p-value	(1.0000)	(0.8877)	(0.5397)
Log GDPPC	3.76288	16.9934	11.5019
p-value	(0.9999)	(0.9727)	(0.9991)
AGE65_74	−0.43112	51.3051	11.0424
p-value	(0.3332)	(0.0090)	(0.9994)
AGE75+	0.76440	25.7248	42.3004
p-value	(0.7777)	(0.6891)	(0.0674)
AVELE65	9.85018	3.55791	4.39433
p-value	(1.0000)	(1.0000)	(1.0000)
FLFPR	−2.22777	54.0744	85.1593
p-value	(0.0129)	(0.0022)	(0.0000)
UNEMP	−1.79099	49.4313	34.6397
p-value	(0.0366)	(0.0142)	(0.2560)
ALCCON	−0.74071	53.0861	86.3141
p-value	(0.2294)	(0.0058)	(0.0000)
PUHES	−1.17358	45.3595	54.7096
p-value	(0.1203)	(0.0357)	(0.0038)
BEDS	−0.36413	68.6014	64.7722
p-value	(0.3579)	(0.0001)	(0.0002)
MORTALITY	2.76183	11.8590	31.8451
p-value	(0.9971)	(0.9987)	(0.3748)
b: Variables in differences/log differences			
Log THEPC	−11.7149	179.922	213.825
p-value	0	0	0
Log GDPPC	−10.7953	159.675	171.730
p-value	0	0	0
AGE65_74	−2.4162	54.3217	34.6698
p-value	0	0	0
AGE75+	−2.41998	46.3724	47.8472
p-value	(0.0078)	(0.0286)	(0.0205)
AVELE65			
p-value			
FLFPR	−17.0325	255.345	352.486
p-value	(0)	(0)	(0)
UNEMP	−6.2482	93.4783	112.983
p-value	(0)	(0)	(0)
ALCCON	−13.3216	203.515	289.492
p-value	(0)	(0)	(0)
PUHES	−14.9786	227.723	481.466
p-value	(0)	(0)	(0)
BEDS	−12.5288	200.926	229.297
p-value	(0)	(0)	(0)
MORTALITY	−20.2074	312.289	822.328
p-value	(0)	(0)	(0)

Note: All tests assume a null hypothesis of a common unit root process

Thus, we consider a simple dynamic panel autoregressive model with lagged values of the dependent as well as the explanatory variables

$$\begin{aligned} &\mathrm{THEPC}_{it} = \delta \mathrm{THEPC}_{i,t-1} + x_{it}'\beta + u_{it} \\ &\quad i = 1,\ldots,N; t = 1,\ldots,T \end{aligned} \tag{1}$$

where i indicates the country in question, and t the time period. x_{it} is a $K \times N$ matrix of covariates (including for ease of notation their lagged variables), β a $1 \times K$ vector of regression slopes, δ a first-order autoregressive parameter, and $u_{it} = \mu_i + \upsilon_{it}$ with $\mu_i \sim \left(0, \sigma_\mu^2\right)$ and $\upsilon_{it} \sim \left(0, \sigma_\upsilon^2\right)$, independent and identically distributed (IID) over the panels. With the hypothesis of a single co-integrating vector accepted, we have provided that the covariates include both current and lagged values of the *I(1)* variables, a satisfactory means of estimating the dynamic relationship. The long-run relationship between *THEPC* and the covariates is given as $\mathrm{THEPC}_i = x_i'\beta/(1-\delta)$ with time subscripts omitted because the long-run does not relate to any particular date.

It has been known for many years that estimation of panel models with lagged dependent variables is subject to biases [25]. One popular means of dealing with this problem is to use the estimation methods described by Arellano and Bond [26] and Arellano and Bover [27] using dynamic generalized method of moments (GMM). However, as Nickell [25] shows the biases fall off rapidly as the sample size increases and are likely to be small with over twenty observations. For our sample size of over twenty observations, it is by no means clear that Arellano and Bond's method is better than more conventional generalized least squares. We did experiment with both methods and found it difficult to identify satisfactory instruments for use with Arellano and Bond's method. Accordingly, we have instead relied on generalized least squares estimation with country fixed effects.

Table 4 Unrestricted and restricted regression results from EU-15

	Unrestricted		Restricted	
	Coefficient	z statistics	Coefficient	z statistics
LOGTHEPC (−1)	0.712597	14.75834	0.7001	22.4300
LOGGDP	0.286783	2.266419	0.3021	10.8121
LOGGDP (−1)	−0.01768	−0.14216	0.0000	Rest*
AGE65_74	0.023154	1.99257	0.0338	5.0139
AGE65_74 (−1)	−0.029777	−2.729757	−0.0338	−5.0139
AGE75+	0.038199	1.904183	0.0311	3.3023
AGE75+ (−1)	−0.049141	−2.26814	−0.0311	−3.3023
AVELE65	−0.008697	−0.582769	−0.0163	−1.8821
AVELE65 (−1)	0.035151	2.510683	0.0288	2.5109
MORTALITY	−0.107016	−0.668941	−0.2128	−2.0307
MORTALITY (−1)	0.168854	1.190521	0.2128	2.0307
FLFPR	0.005525	1.982653	0.0084	7.3622
FLFPR (−1)	0.001753	0.720535	0.0000	Rest*
UNEMP	−0.002544	−0.917307	−0.0019	−3.2634
UNEMP (−1)	0.000201	0.083806	0.0000	Rest*
ALCCON	−0.012096	−1.768828	−0.0058	−1.1924
ALCCON (−1)	0.011911	1.846001	0.0032	0.6695
PUHES	0.004087	1.831964	0.0028	4.9738
PUHES (−1)	−0.000487	−0.198353	0.0000	Rest*
SALARYGP	0.024331	1.369935	0.0000	Rest*
SALARYGP (−1)	0.010513	0.729624	0.0000	Rest*
CAPGP	0.036716	3.136739	0.0220	4.5991
CAPGP (−1)	0.021145	1.849732	0.0201	5.6771
CASEHO	−0.038563	−3.436662	−0.0496	−6.9427
CASEHO (−1)	0.040823	3.760246	0.0447	5.6206
COPAYGP	0.038339	1.553251	0.0472	4.7847
COPAYGP (−1)	−0.014997	−1.140861	0.0086	1.5641
COPAYHO	−0.024313	−2.015252	−0.0199	−3.2467
COPAYHO (−1)	−0.01619	−0.972756	0.0000	Rest*
FREEGP	0.007112	0.441763	0.0000	Rest*
FREEGP (−1)	0.06352	3.966351	0.0429	3.8931
FREEHO	0.052506	4.974856	0.0541	8.1592
FREEHO (−1)	−0.032686	−2.76244	−0.0329	−4.1439
BEDS	0.011626	0.890551	0.0187	5.8882
BEDS (−1)	−0.001009	−0.096272	0.0000	Rest*
Test of restriction			$\chi^2_{12} = 12.90$	

Results

The first column of Table 4 reports the estimated results for an unrestricted version of the model. Looking at this model, the restriction that the coefficient on the lagged dependent variable is 1, which would be required if there were no co-integration/error correction present is plainly rejected on any reasonable basis. Thus, we accept that there is an error correction process present, a finding consistent with the results of Breitung's test, and work on that basis. The second column of Table 4 shows results with zero restrictions imposed on the coefficients of selected statistically insignificant variables. The restrictions are accepted easily ($\chi^2_{12} = 12.9$) when taken together.

Long-run elasticities for the unrestricted model are shown in the first column of Table 5.

Returning to Table 4, it is seen that age composition shows some interesting results. The level of total health expenditure per capita is increasing in the present level of the proportion of the population aged 65–74 and similarly so for the proportion of the population aged 75+. The past levels of these proportions (AGE65-74(-1); AGE75+(-1)) exert a negative influence which is of a magnitude approximately equal to the positive influence of their

Table 5 Long-Run coefficients with different GDP elasticities

	Unrestricted	Restrictions from Table 4			
GDPPC	0.9363	1.0071	1.0000	1.1000	1.2000
AGE65_74	−0.0230	0.0000	0.0000	0.0000	0.0000
AGE75_	−0.0381	0.0000	0.0000	0.0000	0.0000
AVELE65	0.0920	0.0419	0.0431	0.0244	0.0029
FLFPR	0.0253	0.0281	0.0282	0.0273	0.0264
UNEMP	−0.0082	−0.0065	−0.0066	−0.0059	−0.0059
ALCCON	−0.0006	−0.0087	−0.0089	−0.0068	−0.0055
PUHES	0.0125	0.0094	0.0094	0.0101	0.0104
SALARYGP	0.1212	0.0000	0.0000	0.0000	0.0000
CAPGP	0.2013	0.1405	0.1405	0.1460	0.1628
CASEHO	0.0079	−0.0162	−0.0161	−0.0149	−0.0107
COPAYGP	0.0812	0.1862	0.1856	0.1929	0.1982
COPAYHO	−0.1409	−0.0662	−0.0647	−0.0845	−0.1017
FREEGP	0.2458	0.1429	0.1459	0.1088	0.0805
FREEHO	0.0690	0.0707	0.0702	0.0788	0.0895
BEDS	0.0369	0.0622	0.0628	0.0547	0.0465
MORTALITY	0.2152	0.0000	0.0000	0.0000	0.0000
		$\chi^2_{12} = 12.90$	$\chi^2_{13} = 12.93$	$\chi^2_{13} = 16.42$	$\chi^2_{13} = 24.6$

present levels. This implies that a high proportion of elderly people is not in itself a driver of health care expenditure. Rather, it is a shift in the proportion of the population being elderly which causes a shift in the health care expenditure. Thus, to take the effect of an increase in the proportion of people aged 65–74 as an example, if that rises by 1 percentage point to a new, higher level, then in the first year health expenditure is increased by 0.023 log units. By the next year, with no further increase in the proportion of people in this age group, the impact has fallen to just under 0.01 log units. It then falls further, becoming negative. The long-run elasticity of Table 5 shows that it eventually drops to −0.023 log units. However, the fact that we can easily accept the restriction that the lagged coefficient, AGE65-74(-1) is equal but of opposite sign to the coefficient on AGE65-74 means that this long-run effect is not statistically significant. The parameters of the restricted model imply that the first period effect is a rise in health spending by 0.0338 log units. This drops gradually falling below 0.01 log units after 11 years and declines asymptotically to zero.

Life expectancy and mortality show the expected positive association with expenditure. However, for both of these, there seem to be delays in the reaction, as it is the past value rather than the present value of life expectancy and mortality that shifts health care expenditure, presumably indicating that it takes time to implement shifts in these indicators into policy decisions underlying the subsequent shift in health care expenditure. For the case of life expectancy, the short-run effect is approximately zero. However, the positive lagged effect implies a positive long-run effect. In other words, a linear increase in life expectancy is associated with a more than proportional trend, i.e. an exponential growth in health care expenditure. The causal relationship between life expectancy and expenditure may be a two-way relationship, though, but the present study does not allow us to dig deeper into this.

Since the results are partial effects, the combined effect of an increasing life expectancy and an increasing share of the population being elderly cannot be seen directly from Table 4. A tentative conclusion is that although age exerts a positive short-run effect on expenditure, life expectancy is a more important driver given its long-run effect. Turning to mortality, there is a negative short-run effect more than offset by subsequent positive effects. However, as with the share of elderly people in the population, in our restricted model, we are able to accept the hypothesis that there is no long-run influence of mortality on health spending *per capita.*

As expected, GDP/capita has a strong impact on health care expenditures/capita with an elasticity slightly below 1. This suggests that affordability is an important determinant of health care expenditure *per capita*.

Alcohol consumption (*ALCCON*) has an unexpected negative sign implying that lower levels of alcohol consumption is associated with higher health spending. Female labour force participation rate (FLFPR) has an anticipated positive sign, while the unemployment effect (*UNEMP*) indicates that higher unemployment goes along with lower levels of health care expenditure.

The supply side dummy variables tend to have positive signs apart from case-based remuneration of hospitals (*CASEHO*) having a negative association with health spending. There is a positive sign for public health care expenditure (*PUHES*) suggesting that a higher degree of public funding increases health care expenditure. The dummy variables related to access and user payment show a mixture of positive signs for co-payment for visiting a GP (*COPAYGP*), free choice of GP (*FREEGP*) and free choice of hospitals (*FREEHO*) while a negative sign for co-payment for using hospitals (*COPAYHO*).

The implications of the restrictions for the long run are shown in the second column of Table 5. They yield a GDP elasticity very close to one as well as a more satisfactory pattern for the other demographic parameters. We have already noted that neither the age structure of the population nor mortality has an influence on health spending, although average life expectancy plainly does.

Finally, we present in columns 3–5 of Table 5 the implications of restrictions on the GDP elasticity. Not surprisingly, given the results in the first two columns of Table 4, the restriction that the long-run elasticity is one is easily accepted. An expenditure elasticity of 1.1 can be accepted but one of 1.2 cannot.

Discussion and conclusion

The analysis confirms earlier work on different data sets that, once due account is taken of co-variates, the elasticity of health expenditure with respect to GDP is not very different from one, and we can accept that it in fact equals one. Our first hypothesis is therefore accepted. However, it also draws attention to the fact that there are other factors that play a more relevant influence for the development of health care expenditures and that the structure of the health care system may itself be important.

While it has been shown that health care expenditures increases with age on an individual level, the present study addresses the question whether health care expenditure per capita (including the elderly) will increase with an ageing population. Competing models explain the likely outcome, viz. the compression or the expansion of morbidity scenarios, or combinations of these. In a “healthy ageing” scenario which we take as the most realistic, people live longer in good health, and a longer life postpones the high costs at the end of life.

With respect to the effect of ageing, the results suggest that there is a positive correlation between shifts in ageing and shifts in health care expenditure. However, the effect of ageing arises only in the short run; in the long run, we can accept the hypothesis that there is no effect.

We used the variables ‘life expectancy of the elderly’ and ‘mortality’ as proxies for health care costs during the life years of the elderly and the costs near death. We were able to accept restrictions that of these variables, only life expectancy at age 65 has a long-run bearing on health expenditure, thus confirming our second hypothesis that the age structure of the population has no net effect.

We also demonstrated that health care expenditures are not only influenced by the demand side factors, which other macro studies have explored but also by supply factors. This confirmed our third hypothesis. Among the supply side variables are institutional factors that may easily be changed according to changed demand such as capacity, while other variables, such as institutional characteristics, may be less prone to respond to change in demand. Our regression model is limited to variables for which we were able to obtain data; other potentially relevant variables were inevitably omitted.

Thus, our overall conclusion is that health spending depends on a range of different influences. These all have to be taken into account in an analysis of the long-run determinants of future health expenditure.

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