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Jochen Hartwig

What Drives Health Care Expenditure?
Baumol's Model of 'Unbalanced Growth'
Revisited

What drives health care expenditure? – Baumol’s model of ‘unbalanced growth’ revisited

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Abstract

The share of health care expenditure in GDP rises rapidly in virtually all OECD countries, causing increasing concern among politicians and the general public. Yet, economists have to date failed to reach an agreement on what the main determinants of this development are. This paper revisits Baumol’s (1967) model of ‘unbalanced growth’, showing that the latter offers a ready explanation for the observed inexorable rise in health care expenditure. The main implication of Baumol’s model in this context is that health care expenditure is driven by wage increases in excess of productivity growth. This hypothesis is tested empirically using data from a panel of 19 OECD countries. Our tests yield robust evidence in favor of Baumol’s theory.

Key words: Rising health care expenditure, ‘unbalanced growth’, OECD panel

JEL classifications: C12, C22, I10, O41

1 Introduction

The share of current health care expenditure (HCE) in the gross domestic product (GDP) rises rapidly in virtually all developed nations. Figure 1 illustrates this for a couple of countries for which data is available back to 1960. According to the OECD’s health database 2005, only 2 countries (Denmark and Ireland) record similar shares as twenty years ago. In the U.S. on the other hand, the share has risen from 4.8% in 1960 to 14.7% in 2003. This is the highest value of all OECD nations. Switzerland ranks second in this statistic with a share 11.3% in 2003 (up from 4.8% in 1960). An opinion poll carried out in

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this country in early 2006 revealed rising «health care costs» to be the most worrisome topic for Swiss households (cf. NZZ, 2006).

<Insert Figure 1 around here>

In view of this large public concern, it is unfortunate how little we know about the factors that drive the rapid rise in HCE. Back in 1994, HOFFMEYER/MCCARTHY (1994: 67) wrote that “there is just one, very clear, very well-established statistical fact relating to health care expenditure: its correlation with GDP. No other robust and stable correlations have yet been found.” This statement is confirmed by ROBERTS (1999) who dates the starting point of cross-country research into the determinants of HCE back to NEWHOUSE (1977) and then writes: “During this time [the past 20 years] there has been little progress beyond the finding that variations in per capita national income are closely correlated with variations in per capita health spending” (ROBERTS, 1999: 459).

In the literature this judgment is based on we can distinguish two stages. Between the mid-seventies and the mid-nineties, scholars such as KLEIMAN (1974), NEWHOUSE (1977, 1987), CULLIS/WEST (1979), LEU (1986), PARKIN ET AL. (1987), CULYER (1990), MILNE/MOLANA (1991), GETZEN/POULLIER (1991), GERDTHAM/JÖNSSON (1991a, 1991b), GERDTHAM ET AL. (1992), and HITIRIS/POSNETT (1992) provided evidence for a positive correlation between HCE and GDP (mostly) in OECD data. This correlation was found to be robust to varying years covered, estimators and use of conversion factors (such as deflators, exchange rates, or health care purchasing power parities). Other intuitively plausible explanatory variables were normally not found to be statistically significant.¹ An important issue in this first stage of research was the question whether health care is a ‘luxury good’, i.e. whether a larger than proportionate increase in income is spent on health care (cf. GETZEN 2000).

In the more recent stage of research – which started with MURTHY/UKPOLO (1994) and HANSEN/KING (1996) – the time series properties of the variables in question have received more attention than before.² Unit root and cointegration tests were performed

¹ An exception is LEU (1986) who found two parameters relating to the structure of national health systems to be statistically significant explanatory variables, namely the share of public expenditure in total health spending and the presence of a centralized national health system. Later studies mostly failed to confirm these findings.

² Most of the earlier studies were cross section analyses. GERDTHAM (1992) was the first to use pooled cross-section and time series data.

for HCE and GDP. The results of these tests have been somewhat inconclusive and not very robust to the choice of the testing methodology. HANSEN/KING (1996), for instance, found no cointegration between HCE and GDP for all OECD countries except Iceland applying the Engle-Granger (EG) two-step method and testing the residuals from the cointegrating regression with the Augmented Dickey Fuller (ADF) test. BLOMQUIST/CARTER (1997) reversed the result using the Phillips Perron (PP) test for the second step. There is a related controversy on whether the variables are non-stationary in the first place. MCCOSKEY/SELDEN (1998) reject the null hypothesis of unit roots for HCE and GDP while ROBERTS (1999) finds both variables to be non-stationary. Based on country-by-country as well as panel cointegration tests, GERDTHAM/LÖTHGREN (2000) confirm that both health expenditure and GDP have a unit root and that they are cointegrated. Yet, as the latest twist in this debate come the contributions by JEWELL ET AL. (2003) and CARRION-I-SILVESTRE (2005) who, for the first time in this body of literature, consider the possibility of structural breaks in the time series. Using data from 20 OECD countries, they find both HCE and GDP to be stationary around one or two structural breaks.

Reviewing the literature, we cannot help but admit that despite intense research efforts, we don't really know the degree of integration of health expenditure variables. This is unfortunate, since the issue of whether HCE is stationary or not has important consequences for modeling it. If it is integrated of order one ($I(1)$) – and if GDP is in fact cointegrated with it – then the preferred modeling strategy would be to set up an error-correction model (cf. FREEMAN, 2003, HERWARTZ/THEILEN, 2003).³ If, on the other hand, HCE and GDP are stationary variables, then an error-correction model doesn't make much sense, since there cannot be long-run cointegrating relationships between stationary variables. Unfortunately, given that the available time series are rather short, which lowers the power of the tests, and that the number of competing tests is huge (and growing), some uncertainty is likely to remain with respect to the properties of the time series analyzed in this field of research. Of course, uncertainty about the degree of integration of health expenditure and GDP impairs attempts to appropriately model any theorized connection between them.

A new approach might be attractive not only for methodological, but also for theoretical reasons. Over the past 30 years, research into the determinants of health

³ ROBERTS (1999) prefers an autoregressive distributed lag model transformed according to the suggestions by WICKENS/BREUSCH (1988).

care expenditure has concentrated on – and to some degree even confined itself to – evaluating the connection between national health expenditure and GDP. As already mentioned, attempts to detect other explanatory variables – or explanatory variables proper, since a correlation between health care expenditure and GDP does not *explain* much in terms of causal relations, unless one adopts a crude version of Keynesianism – have been sporadic and largely unsuccessful.⁴ Not even such obvious candidates as population shares above certain age thresholds (e.g., 65 years or 75 years) have been found to contribute to the explanation of health care expenditure – except in a few studies, e.g., HITIRIS/POSNETT (1992), DI MATTEO/DI MATTEO (1998), OKUNADE ET AL. (2004). ZWEIFEL ET AL. (1999) show that ‘proximity to death’ rather than ageing drives health care expenditure. Evidently, this variable is contemporaneously unknown and hence inoperative for models intending to forecast HCE.⁵

“Economists have not developed a formal theory to explain or to predict the per capita medical care expenditure of a nation”, concludes WILSON (1999: 160). He continues: “In the absence of a theory, empirical work in this field has necessarily been based on ad-hoc reasoning and data availability” (*ibid.*).

Of course, Wilson’s claim could be disputed on the grounds that GROSSMAN (1972) has introduced a micro-founded model of the demand for health. But the Grossman model is concerned with the individual’s demand for health, while Wilson bemoans the absence of a theory that would explain *aggregate* medical care expenditure. To reduce the Grossman model to a form that can be tested empirically with aggregate data has turned out to be cumbersome since important explanatory variables (e.g. environmental quality) are unobservable (cf. WAGSTAFF, 1986, ERBSLAND ET AL., 1998, NOCERA/ZWEIFEL, 1998). After all, both ROBERTS (1999) and GERDTHAM/JÖNSSON (2000) call for strengthening the theoretical basis for the macroeconomic analysis of health expenditure. According to ROBERTS (1999: 470), this must be the “primary aim for future work”.

We here intend to meet this demand, revisiting BAUMOL’s (1967) theory of unbalanced growth. Baumol develops a simple neoclassical growth model that allows

⁴ These attempts include ABEL-SMITH (1984), OECD (1987), GERDTHAM ET AL. (1992), and GERDTHAM ET AL. (1998) – in addition to the study by LEU (1986) already mentioned.

⁵ SESHAMANI/GRAY (2004a, 2004b) maintain that besides the dominant expenditure driving effect of the ‘proximity to death’, a slight statistically significant effect of ageing also exists. ZWEIFEL ET AL. (2004: 665) dispute this for “a wide range of ages”.

for explicit predictions of the future course of health expenditure. These predictions can be tested empirically. As far as we see, Baumol's model has so far escaped the attention of health economists.⁶ The next section introduces the model; and section 3 proposes and performs empirical tests of its main implications.

2 Baumol's model of unbalanced growth

It is a hallmark of the work of William J. Baumol that it is concerned with recasting neoclassical models in a more 'realistic' form (cf. ELIASSON/HENREKSON, 2004). Baumol's (1967) *American Economic Review* paper 'Macroeconomics of unbalanced growth: the anatomy of urban crisis' is no exception in this respect. In this paper, Baumol divides the economy into two parts: a 'progressive' and a 'non-progressive' sector. He then makes several assumptions, only one of which he claims to be really essential. This essential assumption states that regular growth in labor productivity can occur only in the 'progressive' sector.

For Baumol, *regular* productivity growth is the result of technological innovation which manifests itself in new capital goods. Capital goods are also the source of economies of scale, being another source of productivity growth. *Regular* productivity growth is thus defined to depend on certain physico-technological requirements which exist, as can easily be seen, in manufacturing only. In the service industries, physical capital cannot be employed on a large scale. Baumol cites repeatedly education and – most important for our subject-matter – health care as examples for industries that will inevitably remain highly labor-intensive. Such industries he relegates to the 'non-progressive' sector. "The bulk of our municipal expenditures", Baumol writes, "is devoted to education which ... offers very limited scope for cumulative increases in productivity. The same is true of police, of hospitals, of social services, and of a variety of inspection services. Despite the use of the computer in medicine ..., there is no substitute for the personal attention of a physician ..." (BAUMOL, 1967: 423). Baumol does not claim that increases in labor productivity are impossible in the 'non-progressive' sector, only that this sector comprises "activities which, by their very nature, permit only sporadic increases in productivity" (BAUMOL, 1967: 416). For simplicity, he abstracts from such sporadic productivity increases over the course of his argument.

Another simplifying – but non-essential – assumption is that labor is the only factor of production. BAUMOL (1967: 417) admits, that this assumption is "patently unrealistic".

⁶ KARATZAS (2000) mentions Baumol's theory, but does not test it empirically.

Still, it can be seen as a consequence of his basic assumption since the ‘non-progressive’ sector uses no capital in the simplest version of the model. The employment of capital in the ‘progressive’ sector Baumol captures – at least over time – by postulating that labor productivity in this sector grows at an exogenous rate.

Next, Baumol assumes, although, of course, not in these words, that nominal wages in both sectors are cointegrated. He simplifies further and assumes subsequently that they are equal. His final assumption is that nominal wages (in both sectors) rise to the same extent as labor productivity in the ‘progressive’ sector. This implies that the price level in the ‘progressive’ sector stays constant, whereas it rises in the ‘non-progressive’ sector in order to keep the level of real wages in line with the productivity level. The workers, regardless in which sector they work, buy goods and services from both sectors so that their respective real wages converge.

Equations (1) and (2) describe the production functions of the two sectors. Labor productivity in the ‘non-progressive’ sector (1) stays constant, whereas it grows in the ‘progressive’ sector (2) at the constant rate r . Thus, output in the two sectors (Y_1 and Y_2) at time t is given by:

$$Y_{1t} = aL_{1t} \quad (1)$$

$$Y_{2t} = bL_{2t}e^{rt} \quad (2)$$

with L_1 and L_2 as quantities of labor employed in the two sectors and a and b as constants.

According to one of the the assumptions just mentioned, the nominal wage (in both sectors) is given by:

$$W_t = We^{rt} \quad (3)$$

with W as an arbitrary starting value.

Equation (3) completes the model of ‘unbalanced growth’ already. This simple model has a couple of interesting implications which Baumol draws out. The consequence of prime importance for our context is that the costs per unit of output in the ‘non-progressive’ sector, which are given by (4), tend towards infinity. Costs per unit of output in the ‘progressive’ sector, on the other hand, stay constant (cf. (5)).

$$C_1 = W_t L_{1t} / Y_{1t} = We^{rt} L_{1t} / aL_{1t} = We^{rt} / a \quad (4)$$

$$C_2 = W_t L_{2t} / Y_{2t} = We^{rt} L_{2t} / bL_{2t}e^{rt} = W / b \quad (5)$$

Relative costs also tend towards infinity ($C_1/C_2 = be^{rt}/a$). Under ‘normal’ circumstances – that is, when prices rise in proportion to costs, and when demand is price-elastic – the

‘non-progressive’ sector will vanish. BAUMOL (1967: 421) invokes craftsmanship, fine restaurants, and theaters as examples of establishments that have either disappeared or retreated to luxury niches as a consequence of customers’ unwillingness to tolerate the price increases that would have been necessary to cover rising costs.

Yet, parts of the ‘non-progressive’ sector produce necessities for which the price elasticity is very low. As already mentioned, Baumol calls attention to education and health care as examples. To show what happens in these industries as a consequence of ‘unbalanced growth’, Baumol assumes that the relation of real output of the two sectors remains unchanged as in (6):

$$(b/a) Y_1/Y_2 = L_1/L_2e^{rt} = K, \quad (6)$$

with $K = const.$ If $L = L_1 + L_2$ is the labor force, it follows:

$$L_1 = (L - L_1)Ke^{rt} \quad \Leftrightarrow \quad L_1 = LKe^{rt}/(1 + Ke^{rt}) \quad (7)$$

$$\text{and} \quad L_2 = L - L_1 = L/(1 + Ke^{rt}) \quad (8)$$

From (7) and (8) we learn that, over the years ($t \rightarrow \infty$), L_1 tends towards L , and L_2 tends towards zero. So, if the real output relation of a ‘progressive’ and a ‘non-progressive’ sector is to be kept constant under conditions of ‘unbalanced growth’, an ever larger share of the labor force must move to the ‘non-progressive’ sector (or into unemployment). Because of (4) and (5), a constant relation of real output means that an ever larger share of *nominal* GDP will be allotted to the ‘non-progressive’ sector. This shift of expenditures into the ‘non-progressive’ sector has been termed ‘Baumol’s (cost) disease’ (cf. BAUMOL/TOWSE, 1997).

How relevant is Baumol’s model of ‘unbalanced growth’? – One way of assessing its relevance is to check whether the model’s assumptions and predictions have empirical grounding. This is done by HARTWIG (2005) using Swiss data. He confirms that labor productivity growth is much higher in manufacturing than in any other industry.⁷ Nominal wages in manufacturing and in the tertiary sector seem to be cointegrated. The manufacturing deflator has not risen between 1990 and 2003, whereas prices outside manufacturing increased by 23% over the same period.⁸ The share of health care expenditure in nominal GDP has risen, as has the share of full-time equivalent employment in the health care sector in total employment. But when measured in real

⁷ The same is true for all OECD countries except Norway during the nineties, cf. WÖLFL (2003: 12).

⁸ Such divergent price movements can also be observed in other OECD countries, cf. WÖLFL (2003: 35-6).

terms, the output relation of the two sectors remains fairly constant.⁹ Overall, the assumptions and predictions of Baumol's model seem to be well in line with reality.

What does this model tell us about the causes of rising health expenditure? – Health care is, as we have seen above, one of Baumol's prime examples for that part of the 'non-progressive' sector that produces necessities (for which the price elasticity is very low). The 'non-progressive' sector receives an ever increasing share of total expenditures because wages in the overall economy grow faster than *overall* labor productivity. Baumol's model states that wage increases in excess of labor productivity growth – averaged across both sectors – drive the rise in health expenditure – and that they drive it in a directly proportional manner. The next section will test this statement empirically.

3 Data, methods, and results

As in most other studies, we will use OECD data, to be more specific, data from the *OECD Health Data 2005* CD-ROM. In addition to detailed health data, this database also contains demographic as well as economic references. As a matter of fact, we can extract all data we need to test Baumol's model from the OECD Health Database.

Contrary to most previous studies, we concentrate on the 'total *current* expenditure on health' – a variable that we will call HCE –, rather than on the 'total expenditure on health' as such. Total investment in medical facilities constitutes the difference between the two, but this is a magnitude Baumol's model has not much to say about. Unfortunately, this choice – motivated by the desire to establish the best possible coherence between theory and data – results in a loss of observations since some countries do not disaggregate their 'total expenditure on health' into the two sub-categories. Our remaining sample comprises 19 OECD countries, namely Australia, Austria, Canada, Denmark, Finland, France, Germany, Iceland, Ireland, Italy, the Netherlands, Norway, Portugal, South Korea, Spain, Sweden, Switzerland, the U.K., and the U.S. The national time series for HCE are of different length, covering all the years between 1960 and 2004 for Canada and Norway, but only the period 1990-2003 (and a couple of scattered earlier years) for France. The panel is thus 'unbalanced'. It consists of 622 observations altogether.

⁹ This confirms KANDER's (2005: 128) observation: "When measured in constant prices the relative shares of secondary and tertiary industry are fairly constant in Sweden and it seems other countries".

Baumol's model predicts that wage increases in excess of productivity growth lead to higher HCE growth in a directly proportional manner. So, if we regress the growth rate of (per capita) HCE on the difference between the growth rates of nominal wages per employee and productivity (both averaged over all sectors), we should get a positive and statistically significant coefficient of around one if Baumol's theory has merits.¹⁰

To specify the regression equation this way has certain advantages over previous attempts to explain HCE and meets certain claims put forward in the literature. In their contribution for the *Handbook of Health Economics*, GERDTHAM/JÖNSSON (2000: 48) call for "more theory of the macroeconomics of health expenditure". The specification proposed here offers an explanation for the development of HCE that builds on Baumol's theory of unbalanced growth. Secondly, GERDTHAM/JÖNSSON (2000: 48) encourage researchers to test 'new' explanatory variables (other than GDP). This demand is also met here by testing the difference between wage and productivity growth. Thirdly, Gerdtham and Jönsson propose to study growth rates of health expenditure instead of levels. As far as we know, this is done here for the third time only in the health expenditure literature.¹¹ Studying the growth rates of HCE has the advantage to avoid the cumbersome and so far unresolved issue of the degree of integration of the HCE time series. (As was pointed out in the introduction, some researchers have found HCE to be stationary, while others claim that they are integrated of order one.) By specifying the regression in growth rates (or log differences, respectively), we are unlikely to produce spurious results since no-one has proposed that HCE are $I(2)$.¹² To be on the safe side, we also tested the growth rates entering our model for the presence of unit roots, applying a battery of recently developed panel unit root tests. Results are given in the appendix. Although no absolutely clear picture emerges from these tests, most of them indicate that the series

¹⁰ Another possibility to test Baumol's model would be to regress HCE growth on the development of the relation between health care prices and prices in the 'progressive' sector, e.g. in manufacturing. However, medical care prices are notoriously difficult to measure (cf. NEWHOUSE, 2001), so that the above-mentioned specification is preferred.

¹¹ The pioneering contributions are BARROS (1998) and OKUNADE ET AL. (2004).

¹² OKUNADE/KARAKUS (2001) find HCE to be $I(2)$ for some countries, but this is probably due to the non-inclusion of a deterministic time trend in their ADF regressions. HANSEN/KING (1998) advocate the inclusion of a time trend since HCE are clearly trended.

are stationary. We will thus specify growth regressions much the same way BARROS (1998) does, which means that we will estimate the linear equation

$$\hat{y}_{it} = \alpha X_{it} + e_{it} \quad (9)$$

where \hat{y}_{it} denotes the growth rate (log difference) of HCE per capita in country i in year t . X is the vector of exogenous variables, including the difference of the growth rates of nominal wages per employee and productivity – we will call this difference occasionally ‘the Baumol variable’, for short –, possibly additional explanatory variables, and a constant term. e_{it} is an error term that follows a i.i.d.(0, σ^2) process. As DEVLIN/HANSEN (2001: 562) point out, this testing strategy involves estimating an error correction model, albeit without an error correction term.

Before we proceed to presenting estimation results, we must pause for a moment to stress another feature that, apart from the choice of explanatory variables, distinguishes our approach from the bulk of the literature. In almost every paper on the subject at hand, HCE and GDP are being deflated – either with the same deflator (the GDP deflator as in, e.g., GERDTHAM, 1992), or with different deflators for HCE and GDP, respectively. Here, however, we will use nominal data. The main reason for not deflating nominal data is that we intend to test Baumol’s model of unbalanced growth, which is in nominal terms. But even more generally, there seem to be no obvious reasons for deflating expenditure magnitudes in this field of research. Economists are trained to think in terms of the physical magnitudes behind the ‘veil of money’. But it is not so much the rise in (real) health care consumption that stirs public concern, but the ‘cost explosion’, in other words the increase in nominal expenditure. Therefore, the latter should be modeled as dependent variable.¹³

Also, we will not convert our data into purchasing power parities (PPPs). PPP conversion is necessary when levels of variables are compared across countries. Here, however, we will model only growth rates. Of course, we could compare growth rates of variables after PPP conversion. This would amount to weighting the components of national GDPs and HCEs with international relative prices instead of national relative prices. But since PPPs are calculated on the basis of smaller price samples than national

¹³ FREEMAN (2003) also uses nominal data. GERDTHAM/JÖNSSON (1991a) find the relative price of health care to be uncorrelated with per-capita income, so that, in their model, it doesn’t matter whether health care is measured in terms of expenditure or quantity.

price indexes, the OECD recommends not to convert variables into PPPs in international comparisons of growth rates.¹⁴ We will follow that recommendation.

An important advantage of working with pooled data is that it allows for richer specifications including country and time-specific effects. These effects can be captured by introducing either country dummy variables (cross-section fixed-effects) or period dummy variables (period fixed-effects) or both. If, however, the country or period-specific effect is itself a random variable, it would be appropriate to estimate a random-effects model. The standard way of choosing between fixed and random-effects models is to run a Hausman test. In our case – contrary to the finding of GERDTHAM (1992) – this test does not reject the null hypothesis that the random effects are uncorrelated with the explanatory variables so that the random-effects estimator is to be preferred. This probably reflects the fact, that our sample of 19 countries is drawn somewhat randomly from the group of OECD countries according to data availability. We will present results of cross-section as well as time period random-effects estimations along with OLS estimates. (As our sample is ‘unbalanced’, it is not possible to estimate two-way random-effects models.)

Tables 1 and 2 summarize our estimation results. We have to consider first that the ‘Baumol variable’ – the difference between nominal wage and productivity growth – can be split into three separate variables, namely per-capita wage growth, real GDP growth and growth in employment. Proceeding in a general-to-specific manner, we first estimate the influence of these three variables separately in order to test whether the restriction of summing them together to one variable is legitimate. The estimation period covers the years from 1971 to 2003.

Table 1 shows that the three variables are statistically different from zero with signs as expected. For all three estimations, a Wald test fails to reject the hypothesis that $C(1) + C(2) - C(3) = 0$ so that we can legitimately combine the three variables into one.

Table 2 shows our results for the ‘Baumol variable’. We find that Baumol’s model of unbalanced growth is strongly supported by the data. In all three specifications, the coefficient of the difference between nominal wage and productivity growth rates is statistically different from zero. As predicted by Baumol’s theory, the value of the coefficient is close to one. Again, the Wald test fails to reject the hypothesis that the

¹⁴ “(I)t is generally the case that for purposes of comparing relative output and productivity *growth*, the comparison based on constant national prices is to be preferred. PPPs should be used when output and productivity *levels* are the object of comparison across countries» (AHMAD ET AL., 2003: 19).

coefficients are in fact equal to one. Whereas GERDTHAM's (1992) results were highly dependent on the choice of the estimation technique, our coefficients remain largely stable across specifications. The model can explain around 75% of the variation of health expenditure in our group of 19 OECD countries.

<Insert Tables 1 and 2 around here>

4 Robustness analysis

We now turn to the question how robust our results are to variations in the modeling framework. We will perform three types of robustness tests. First, we ask whether the inclusion of another explanatory variable alters the results. This additional explanatory variable will, of course, be per-capita GDP, which has emerged as *the only* uncontroversial explanatory variable for per-capita HCE from the literature. Other variables could, of course, be tested also, but it is of prime importance to find out whether the model is robust to the inclusion of GDP. As we intend to explain *nominal* HCE here, we will include nominal GDP as additional right-hand side variable.

Table 3 shows that nominal per-capita GDP indeed adds significantly to the explanation of per-capita HCE. The estimated coefficients of around 0.35 are evidence against the hypothesis that health care is a 'luxury good'. The coefficients for the 'Baumol variable' ($d\log(\text{WSPE}) - d\log(\text{PROD})$) drop to around 0.75, confirming that this variable is positively correlated with GDP growth. This is no surprise since both wages and nominal GDP are driven by inflation. Yet, the collinearity is not so strong as to render both variables individually insignificant. If we replace nominal with real GDP growth – as in (9) and (10) –, the multicollinearity vanishes, and the coefficient of the 'Baumol variable' reassumes a value that is not significantly different from one, according to Wald tests.

<Insert Table 3 around here>

The second robustness test concerns parameter stability, which we will test the same way as GERDTHAM (1992), i.e. by splitting the estimation period into three arbitrary sub-periods, namely 1971-1981, 1982-1992, and 1993-2003. Table 4 shows the results for models (7) and (8), allowing us to evaluate not only the robustness of ($d\log(\text{WSPE}) - d\log(\text{PROD})$), but also of GDP. The coefficient of the 'Baumol variable' is not stable; it fluctuates over time between around 0.6 and around 0.9, without showing a clear trend. One thing we can infer from table 4 is that the descriptive quality of the model – if measured by the adjusted R^2 – is lower for the most recent decade than for the two

decades before. This might suggest that OECD economies have recently become less prone to Baumol's 'cost disease'. Probably, globalization has reduced the leeway for wage increases in excess of productivity growth in the 'non-progressive' sectors of OECD economies.¹⁵ Still, however, the 'Baumol variable' is significantly different from zero at the 1% level in all three sub-periods. GDP growth, on the other hand, is insignificant in the first sub-period.

If models (9) and (10) instead of models (7) and (8) are tested, the coefficients of $(\text{dlog}(\text{WSPE}) - \text{dlog}(\text{PROD}))$ are not only highly significant in all three sub-periods, but also stable. They vary between 0.99 and 1.06 for the cross-section random-effects model and between 0.96 and 1.05 for the time period random-effects model. Again, the models have the lowest explanatory power (as well as the lowest coefficient values) in the most recent decade.

<Insert Table 4 around here>

Finally, we test parameter stability by dropping each of the 19 countries in turn and re-estimating models (7) and (8). Table 5 shows results for model (7). (Replacing cross-section random effects with time period effects does not change the overall picture.¹⁶) Obviously, the model is not very sensitive to the exclusion of single countries. The distribution of the 19 estimated coefficients for the 'Baumol variable' has a mean of 0.767 (which is equal to the estimate for the full sample) with a minimum of 0.732 (when Portugal is excluded) and a maximum of 0.803 (when Iceland is dropped). The standard deviation of the estimates around the mean is as low as 0.015.

The estimated coefficients are somewhat sensitive to the exclusion of Iceland (being a high inflation country for most of the estimation period). This finding is in line with GERDTHAM (1992). Gerdtham also reports a high sensitivity of his estimates to the exclusion of the Netherlands, which our results fail to confirm.

Overall, we conclude that the development of health expenditure in OECD countries since the seventies is in line with Baumol's theory of unbalanced growth. Wage increases in excess of productivity growth are a statistically significant explanatory variable for HCE growth. The coefficient of this variable is close to unity, as predicted by

¹⁵ BARROS (1998) also finds a slowdown in HCE growth over the last decade, but doesn't offer an explanation for it.

¹⁶ Results are available from the author upon request.

theory. This finding is robust to the addition of GDP as explanatory variable, the exclusion of countries from the sample, and the variation of the estimation period.

So far, it has been regarded as a statistical fact – even as the only statistical fact established in this area of research (cf. HOFFMEYER/McCARTHY, 1994, ROBERTS, 1999) – that HCE depend on GDP or national income, respectively. Our results do not challenge the role of GDP as explanatory variable for health expenditure, since GDP remains statistically significant even after accounting for the influence of what we here have called the ‘Baumol variable’. Still, the latter seems to be an explanatory variable which is even more robust than GDP (cf. table 4).

<Insert Table 5 around here>

5 A short note on the implications of our findings for health policy

Our study adds empirical support to Baumol’s model of ‘unbalanced growth’. If this model has identified the major cause of the continuous rise in health expenditure correctly, what are the consequences for health policy? To be more specific, what can health policy do to prevent HCE from rising further?

Unfortunately, the answer must be: *not much*. Baumol’s analysis reveals a fundamental supply-side factor to be the main driver behind health expenditure, namely the fact that *regular* (as opposed to sporadic) growth in labor productivity can occur only in sectors that employ capital goods (machines). Technological progress manifests itself in capital goods, which are also the source of economies of scale. In most of the service industries, and especially in health care, there is little leeway for increasing productivity by substituting capital goods for human labor. Health policy cannot change the basic production techniques in the supply of health care.

Baumol, on his part, is unambiguous in this respect. The rise in the share of health care expenditure in nominal GDP he calls “a trend for which no man and no group should be blamed, for there is nothing that can be done to stop it” (BAUMOL, 1967: 423). Elsewhere (1967: 415), he declares: “(E)fforts to offset these cost increases, while they may succeed temporarily, in the long run are merely palliatives which can have no significant effect on the underlying trends.” And even twenty-five years later, he remains convinced that health care is “an industry whose costs are driven by technological imperatives to rapid rise” (BAUMOL, 1993: 27).

This said, it is clear that the best ‘palliative’ against the continuous rise in health care expenditure is to increase productivity in the health sector. The assumption of a zero

productivity growth rate in the ‘non-progressive’ sector is clearly unwarranted. Productivity can be increased in this sector also, either by employing capital goods (to the extent possible) or by an improved organization. If productivity grows, the rise in the share of health care expenditure in nominal GDP slows down. It has to be stressed, though, that as long as the productivity growth rate in the ‘non-progressive’ sector remains below the respective rate in the ‘progressive’ sector – which must be the case if Baumol’s argument about the physico-technological determinants of productivity growth is correct –, the ‘health care cost explosion’ can only be protracted, but not stopped or even reversed.

6 Conclusion

In their contribution for the *Handbook of Health Economics*, GERDTHAM/JÖNSSON (2000: 48) identify three important issues for future research into the determinants of health care expenditure (HCE). First, they call for “more theory of the macroeconomics of health expenditure”. Second, they encourage researchers to move away from focusing on GDP as the only (uncontroversial) explanatory variable for HCE and to test ‘new’ variables. Finally, Gerdtham and Jönsson propose to study growth rates of health expenditure instead of levels.

This paper confronts all three issues. Baumol’s model of ‘unbalanced growth’ is revisited and recognized as a possible theoretical basis for the explanation of health expenditure. Baumol’s model identifies nominal wage growth in excess of productivity growth as the main determinant of the rise in HCE. This hypothesis is here tested empirically with a regression model that is specified in growth rates using pooled cross-section and time-series data from 19 OECD countries. The ‘Baumol variable’ – the difference between wage and productivity growth – is found to contribute significantly to the explanation of HCE. This finding is robust to the inclusion of GDP as additional explanatory variable as well as to varying the estimation period and the sample of countries. Even the hypothesis that the coefficient for the ‘Baumol variable’ is equal to one – which follows from Baumol’s theory – cannot be rejected. Overall, it seems that Baumol’s model of ‘unbalanced growth’ can lay a theoretical foundation for the research into the determinants of health care expenditure. The perspective for health policy this model offers is dim, however, since the rise in health expenditure is predicted to continue inexorably due to technological reasons.

Appendix: Results from panel unit root test

To check whether the assumption is warranted that the log differences of per capita HCE, wages per employee and productivity are stationary, we apply a battery of recently developed panel unit root tests. Panel unit root tests are preferable to testing the individual country series because exploiting the cross-section dimension of the data increases the power of the test. Table 6 summarizes the results of tests for a unit root in first differences which were obtained using EViews5.1[®]. (More detailed test results are available upon request.)

CARRION-I-SILVESTRE (2005) describes the various panel unit root tests in great detail so that we can leave it at a short note here. All the tests are in principle multiple-series unit root tests. While the LEVIN ET AL. (2002), BREITUNG (2000), and HADRI (1999) tests assume the autoregressive coefficient to be identical across cross-sections, the IM ET AL. (2003), as well as the two Fisher-type tests (MADDALA/WU, 1999 and CHOI, 2001), allow the latter to vary freely across cross-sections.

The Levin et al. and Breitung tests are both based on an ADF regression. Estimates for the autoregressive coefficient are calculated from proxies for the variables in question that are standardized and free of autocorrelation and deterministic components. The two tests differ in the way the standardized proxies are constructed and de-trended. Unlike the other five tests, the Hadri test tests the null hypothesis of ‘no unit root’ against the alternative of a unit root in the data. This test is based on the residuals from OLS regressions of the panel variables (in levels) on a constant, or on a constant and a trend.

The other three tests – which allow the autoregressive coefficient to vary across cross-sections – all combine individual unit root tests to derive a panel-specific result. IM ET AL. (2003) have simulated critical values that can be used to test whether the average of the t-statistics for the autoregressive coefficient from the individual countries’ ADF-regressions is statistically significant. The remaining two tests draw on FISHER’s (1932) result that

$$-2 \sum_{i=1}^N \log(\pi_i) \rightarrow \chi_{2N}^2 \quad (10)$$

where π_i denotes the p-value from any individual unit root test for cross-section i . MADDALA/WU (1999) use the ADF-test for the individual unit root test, whereas CHOI (2001) uses the Phillips-Perron (PP-) test.

Note that the test results in Table 6 refer to the panel of 19 OECD countries listed in section 3. In line with the recommendations by HANSEN/KING (1998), individual effects and individual linear trends are included as exogenous variables. For the first four tests listed in the table, maximum lags are automatically selected based on the Schwarz Information Criterion. The remaining two tests use the Bartlett kernel for the Newey-West bandwidth selection. The probabilities for the Fisher tests are computed using an asymptotic Chi-square distribution. All other tests assume asymptotic normality.

<Insert Table 6 around here>

We observe that all tests which test the null hypothesis that there is a unit root in the data reject this hypothesis at conventional significance levels. (The Breitung test with respect to wages and salaries per employee could be regarded as an exception.) The Hadri test, however, points to unit roots in the data for all three variables.

It has been pointed out before that certainty is unlikely to be gained from unit root tests. We will take the respective growth rates to be stationary, which is in line with the result of most of the tests.

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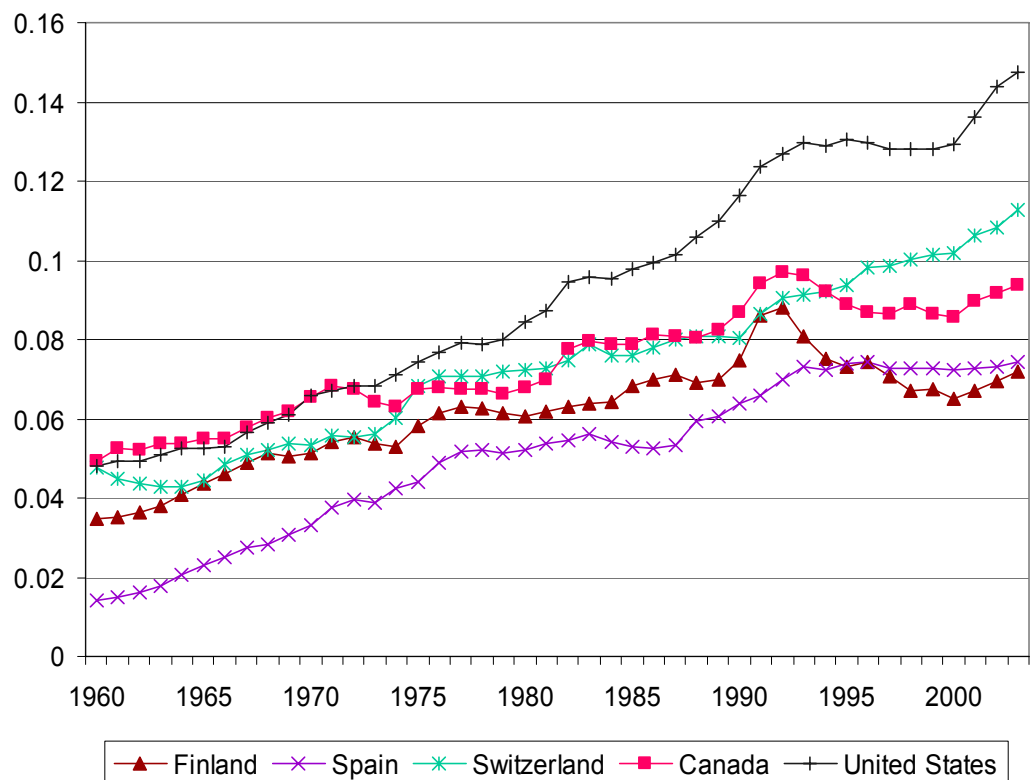


Figure 1: Shares of current health care expenditure in GDP

Source: OECD Health Data 2005 CD ROM

Table 1: Results for growth rate equations – ‘Baumol variable’ split

Log difference of health care expenditure per capita			
	(1)	(2)	(3)
	OLS	Cross-section R.E.	Time period R.E.
dlog(WSPPE)	1.066* (28.557)	1.064* (27.561)	1.059* (27.155)
dlog(GDPR)	-0.339* (-3.951)	-0.351* (-4.049)	-0.308* (-3.571)
dlog(EMP)	0.601* (7.377)	0.599* (7.331)	0.588* (7.511)
Number of obs.	507	507	507
R ² (adj.)	0.810	0.798	0.799
Stand. err. of regr.	0.032	0.031	0.031
D-W	1.830	1.852	1.827
F-stat.	715.308*	667.669*	672.762*

dlog(WSPPE) = log difference of wages and salaries per employee in the overall economy, dlog(GDPR) = log difference of real GDP, dlog(EMP) = log difference of overall employment.

The values shown in parenthesis are t-ratios, based on White’s robust S.E.s. The Swamy-Arora GLS estimator was used to estimate the random effects models, and weighted diagnostic statistics are reported. An asterisk denotes significance at the 1% level. Estimates for constant terms not shown.

Table 2: Results for growth rate equations – ‘Baumol variable’ unsplit

Log difference of health care expenditure per capita			
	(4)	(5)	(6)
	OLS	Cross-section R.E.	Time period R.E.
(dlog(W SPE) – dlog(PROD))	1.033* (34.763)	1.016* (32.763)	1.029* (34.204)
Number of obs.	507	507	507
R ² (adj.)	0.771	0.742	0.762
Stand. err. of regr.	0.035	0.033	0.034
D-W	1.668	1.787	1.663
F-stat.	1699.086*	1456.408*	1623.548*

dlog(W SPE) = log difference of wages and salaries per employee in the overall economy,
dlog(PROD) = log difference of labor productivity (real GDP per employee) in the overall economy.

The values shown in parenthesis are t-ratios, based on White’s robust S.E.s. The Swamy-Arora GLS estimator was used to estimate the random effects models, and weighted diagnostic statistics are reported. An asterisk denotes significance at the 1% level. Estimates for constant terms not shown.

Table 3: Robustness test I – Inclusion of per-capita GDP as explanatory variable

Log difference of health care expenditure per capita				
	(7)	(8)	(9)	(10)
	Cross-section R.E.	Time period R.E.	Cross-section R.E.	Time period R.E.
(dlog(WSPPE) – dlog(PROD))	0.767* (10.649)	0.727* (11.067)	1.056* (28.984)	1.053* (28.629)
dlog(GDPPC)	0.331* (4.594)	0.373* (5.361)		
dlog(GDPRPC)			0.483* (5.778)	0.525* (6.308)
Number of obs.	507	507	507	507
R ² (adj.)	0.787	0.789	0.789	0.789
Stand. err. of regr.	0.032	0.032	0.032	0.032
D-W	1.759	1.731	1.831	1.808
F-stat.	937.377*	946.027*	949.812*	946.094*

dlog(WSPPE) = log difference of wages and salaries per employee, dlog(PROD) = log difference of labor productivity (real GDP per employee), dlog(GDPPC) = log difference of nominal per-capita GDP, dlog(GDPRPC) = log difference of real per-capita GDP

The values shown in parenthesis are t-ratios, based on White's robust S.E.s. The Swamy-Arora GLS estimator was used to estimate the random effects models, and weighted diagnostic statistics are reported. An asterisk denotes significance at the 1% level. Estimates for constant terms not shown.

Table 4: Robustness test II – Inter-temporal parameter stability

Log difference of health care expenditure per capita						
	Cross-section R.E.			Time period R.E.		
	(11)	(12)	(13)	(14)	(15)	(16)
	1971–1981	1982–1992	1993–2003	1971–1981	1982–1992	1993–2003
(dlog(WSPE) – dlog (PROD))	0.917* (13.748)	0.637* (8.488)	0.779* (6.897)	0.884* (15.903)	0.615* (8.834)	0.731* (6.801)
dlog(GDPPC)	0.140 (1.265)	0.454* (4.353)	0.476* (5.312)	0.176 (1.749)	0.465* (4.771)	0.506* (5.452)
Number of obs.	128	185	194	128	185	194
R ² (adj.)	0.712	0.792	0.485	0.716	0.828	0.484
Stand. err. of regr.	0.039	0.031	0.027	0.038	0.032	0.026
D-W	1.778	1.937	1.610	1.788	1.820	1.629
F-stat.	158.302*	351.800*	91.736*	160.979*	443.557*	91.377*

dlog(WSPE) = log difference of wages and salaries per employee, dlog(PROD) = log difference of labor productivity (real GDP per employee), dlog(GDPPC) = log difference of nominal per-capita GDP

The values shown in parenthesis are t-ratios, based on White's robust S.E.s. The Swamy-Arora GLS estimator was used to estimate the random effects models, and weighted diagnostic statistics are reported. An asterisk denotes significance at the 1% level. Estimates for constant terms not shown.

Table 5: Robustness test III – Cross-national stability of parameters of model (9)

	Log difference of health care expenditure per capita											
	Australia	Austria	Canada	Denmark	Finland	France	Germany	Iceland	Ireland	Italy		
(dlog(WSPE) – dlog (PROD))	0.778* (10.489)	0.767* (13.391)	0.757* (9.939)	0.762* (12.529)	0.747* (9.267)	0.767* (10.592)	0.764* (10.627)	0.803* (10.858)	0.788* (10.542)	0.769* (10.617)		
dlog(GDPPC)	0.322* (4.222)	0.325* (5.111)	0.344* (4.572)	0.335* (4.604)	0.352* (4.431)	0.329* (4.518)	0.329* (4.566)	0.311* (4.448)	0.318* (4.299)	0.326* (4.484)		
Number of obs.	477	475	474	484	474	494	476	484	480	492		
R ² (adj.)	0.786	0.805	0.756	0.795	0.783	0.785	0.787	0.718	0.800	0.787		
Stand. err. of regr.	0.033	0.030	0.033	0.033	0.033	0.033	0.033	0.032	0.032	0.032		
D-W	1.767	1.747	1.777	1.756	1.778	1.763	1.768	1.731	1.782	1.765		
F-stat.	876.405*	981.923*	871.581*	938.126*	854.693*	899.734*	877.584*	615.955*	959.017*	908.908*		
	Netherlands	Norway	Portugal	South Korea	Spain	Sweden	Switzerland	U.K.	U.S.			
(dlog(WSPE) – dlog (PROD))	0.768* (10.436)	0.751* (9.409)	0.732* (13.169)	0.776* (11.065)	0.767* (9.874)	0.773* (10.829)	0.765* (10.650)	0.770* (9.589)	0.765* (10.344)			
dlog(GDPPC)	0.328* (4.461)	0.352* (4.389)	0.350* (5.831)	0.323* (4.554)	0.322* (4.261)	0.327* (4.517)	0.332* (4.546)	0.333* (4.465)	0.338* (4.590)			
Number of obs.	480	476	478	487	474	490	475	481	475			
R ² (adj.)	0.785	0.805	0.787	0.797	0.786	0.790	0.785	0.783	0.795			
Stand. err. of regr.	0.033	0.032	0.030	0.031	0.032	0.032	0.033	0.033	0.033			
D-W	1.763	1.735	1.706	1.760	1.771	1.762	1.753	1.742	1.772			
F-stat.	873.365*	980.728*	880.201*	956.538*	869.165*	921.791*	867.194*	868.119*	922.279*			

dlog(WSPE) = log difference of wages and salaries per employee, dlog(PROD) = log difference of labor productivity (real GDP per employee), dlog(GDPPC) = log difference of nominal per-capita GDP
 The values shown in parenthesis are t-ratios, based on White's robust S.E.s. The Swamy-Arora GLS estimator was used to estimate the random effects models, and weighted diagnostic statistics are reported. An asterisk denotes significance at the 1% level. Estimates for constant terms not shown.

Table 6: Panel unit root test results

	log(HCEPC)			log(WSPE)			log(PROD)		
<i>H₀: Unit root in first diff.</i>	Stat.	Prob.	Obs.	Stat.	Prob.	Obs.	Stat.	Prob.	Obs.
Levin, Lin & Chu t*	-7.359	0.000	551	-6.677	0.000	592	-18.162	0.000	596
Breitung t-stat	-2.674	0.004	532	-1.313	0.095	573	-7.739	0.000	577
Im, Pesaran & Shin W-stat	-6.739	0.000	551	-6.024	0.000	592	-16.548	0.000	596
ADF – Fisher Chi-square	127.974	0.000	551	112.242	0.000	592	280.978	0.000	596
PP – Fisher Chi-square	251.240	0.000	569	122.245	0.000	604	362.727	0.000	605
<i>H₀: No unit root in first diff.</i>									
Hadri Z-stat	9.483	0.000	589	4.783	0.000	623	6.580	0.000	624

log(HCEPC) = log of per capita health care expenditure, log(WSPE) = log of wages and salaries per employee,
log(PROD) = log of labor productivity (real GDP per employee)