

Hartwig, Jochen

Working Paper

Can Baumol's model of unbalanced growth contribute to explaining the secular rise in health care expenditure? An alternative test

KOF Working Papers, No. 178

Provided in Cooperation with:

KOF Swiss Economic Institute, ETH Zurich

Suggested Citation: Hartwig, Jochen (2007) : Can Baumol's model of unbalanced growth contribute to explaining the secular rise in health care expenditure? An alternative test, KOF Working Papers, No. 178, ETH Zurich, KOF Swiss Economic Institute, Zurich, <https://doi.org/10.3929/ethz-a-005502972>

This Version is available at:

<https://hdl.handle.net/10419/50451>

Standard-Nutzungsbedingungen:

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

Terms of use:

Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.

KOF Working Papers

Can Baumol's Model of Unbalanced Growth Contribute to
Explaining the Secular Rise in Health Care Expenditure?

An Alternative Test

Jochen Hartwig

KOF

ETH Zurich
KOF Swiss Economic Institute
WEH D 4
Weinbergstrasse 35
8092 Zurich
Switzerland

Phone +41 44 632 42 39
Fax +41 44 632 12 18
www.kof.ethz.ch
kof@kof.ethz.ch

Can Baumol's model of unbalanced growth contribute to explaining the secular rise in health care expenditure?

An alternative test

Jochen Hartwig*

ETH Zurich, KOF – Swiss Economic Institute

Abstract

In a recent paper I argued that Baumol's (1967) model of 'unbalanced growth' offers a ready explanation for the observed secular rise in health care expenditure (HCE) in rich countries (HARTWIG 2006). Baumol's model implies that HCE is driven by wage increases in excess of productivity growth. I tested this hypothesis empirically using data from a panel of 19 OECD countries and found robust evidence in favor of Baumol's theory.

An alternative way to test Baumol's theory is to check whether its implication that variations in the relative price of medical care contribute significantly to explaining variations in health expenditure in the same direction has an empirical grounding. Earlier studies, although mostly not in an explicit attempt to test Baumol's theory, have occasionally rejected this hypothesis. Despite poor data quality of the available medical price indices, I perform the alternative test using data for nine OECD countries. My findings suggest that the relative price of medical care is in fact a statistically significant explanatory variable for health expenditure, thus lending support to Baumol's theory.

Key words: Rising health expenditure, 'unbalanced growth', medical care prices, OECD panel

JEL classifications: C12, C23, E31, I10, O41

* Jochen Hartwig, KOF ETH, WEH E7, Weinbergstr. 35, CH-8092 Zurich, Tel.: ++41/+44/6327331, Fax: ++41/+44/6321218, E-mail: hartwig@kof.ethz.ch.

Introduction

The share of health care expenditure (HCE) in GDP rises inexorably, it would seem, in virtually all OECD countries, causing increasing concern among politicians and the general public. Figure 1 illustrates this for a couple of countries for which data are available back to 1960.

<Insert Figure 1 around here>

Starting in the seventies, health economists have sought to explain the reasons behind the surge in health expenditure; but not much consensus has emerged beyond the finding that aggregate HCE is closely correlated with GDP or National Income, respectively (cf. ROBERTS 1999: 459). Although most studies that include additional explanatory variables, for example variables representing institutional features of health systems, initial levels of health expenditure or population shares above certain age thresholds, find some of them to be statistically significant, these findings are not very robust to varying the testing methodology or the sample of countries or years covered.¹ Furthermore, the choice of explanatory variables has been largely ad hoc (cf. WILSON 1999: 160). Reviewing the achievements in this field of research for the *Handbook of Health Economics*, GERDTHAM/JÖNSSON (2000: 48) conclude that “(w)e need more theory of the macroeconomics of health expenditure, at least relative to the macroeconometrics of health expenditure”.

Yet, a well-known model exists that could offer a theoretical explanation for the constant rise in the share of health expenditure in GDP – Baumol’s model of ‘unbalanced growth’. Although Baumol’s (1967) paper introducing this model is widely regarded as a seminal contribution, health economists have been reluctant to embrace his approach. HARTWIG (2006), in his meta-analysis of research on the determinants of HCE that focuses in the main on English-language top journals, finds only one paper to mention Baumol’s model (KARATZAS 2000) and none to test it.²

¹ Cf. GERDTHAM/JÖNSSON (2000) for a review of the literature up to the late nineties and OKUNADE ET AL. (2004) for a recent contribution.

² Recently, the OECD (2006) as well as NIXON/ULMANN (2006) have considered Baumol’s approach, but, again, without testing it.

In a nutshell,³ Baumol's model states that expenditure shares shift towards the services sector – of which the health care sector forms a part – because of divergent productivity growth between the secondary and the tertiary sectors of the economy. Baumol assumes that productivity growth is the result of technological innovation which manifests itself in new capital goods. It follows that productivity growth is largely confined to the manufacturing industries since, in most service industries, physical capital cannot be employed on a large scale. Furthermore, Baumol assumes that wages in both the 'progressive' and the 'nonprogressive' sectors grow at the same rate so that unit labor costs in the tertiary sector rise relative to those in the secondary sector. If prices are a mark-up on costs they will grow faster in sector 3 than in sector 2. What will then happen is a question of the relevant price elasticities of demand. If the price elasticity is high, then the activity 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. The other possibility is that the sector produces necessities for which the price elasticity is low – as the health care sector does. 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. Then, an ever larger share of the labor force must move to the tertiary sector (or into unemployment). And because of the divergent growth in unit labor costs combined with mark-up pricing, an ever larger share of *nominal* GDP will be allotted to the tertiary sector, even though the relation of *real* output of the two sectors is assumed to remain the same. This shift of expenditures into activities largely financed out of tax money, such as education and health care, has been termed 'Baumol's (cost) Disease' (cf. BAUMOL/TOWSE 1997).

1 Are Baumol's assumptions warranted for the health sector?

It is sometimes argued that Baumol's assumptions are unwarranted for the health sector. Two lines of argument are put forward to support this view. First, it is disputed that the capital intensity of the health sector is low (as Baumol asserts). As is well known, more and more medical appliances are installed in hospitals, and technical progress is substantial in this field. Therefore, the production of health care services might be as capital-intensive as manufacturing activities – or even more so.

³ A technical discussion of Baumol's model can be found in HARTWIG (2006). It need not be repeated here.

This argument misses the point, however. The utilization of medical appliances in hospitals differs fundamentally from the employment of capital in manufacturing. In manufacturing, capital is employed as a substitute for labor or in order to make labor more productive. Medical appliances, on the other hand, do not displace labor. They equip the staff with new instruments.⁴ Also, they are not primarily installed in order to raise labor productivity, but in order to allow for better diagnoses or treatments. If, for instance, a physical examination aided by medical technology takes the doctor more time than a tactile examination, measured labor productivity will decline (cf. also WÖLFEL 2005: 55-6). The quality of the examination might have risen, but since price indices used to deflate medical expenses in order to arrive at measures of real value added are normally not quality-adjusted, those quality improvements will not be recorded as an improvement in productivity. Of course, the depreciation allowances for expensive medical appliances raise health care costs further. This is an aspect that Baumol's model does not include; but the model as such is not compromised by the fact that we witness investments in medical appliances.

A second and closely related objection against applying Baumol's model to analyze the health care sector concerns the measurement of prices and productivity. Already the Boskin Commission (BOSKIN ET AL. 1996) suspected the medical care price index in the U.S. to be upward-biased substantially due to a failure to account for quality improvements in treatments, for example the reduction of pain thanks to minimum invasive surgery. In line with earlier studies (e.g., TRIPLETT 1999, BERNDT ET AL. 2000), NEWHOUSE (2001: 52) suggests "a large constant upward bias" in medical prices, albeit without being able to present precise estimates. Each upward bias in a price index translates into a downward bias in the index of real value added that is calculated by deflating nominal expenditure with that price index. So, probably, productivity growth in health care only *appears* to be lower than in manufacturing due to mis-measurement of price changes. If so, Baumol's theory would be compromised.

In 1997, the U.S. *Bureau of Labor Statistics* (BLS) introduced new indices for hospital services in its producer price index. Instead of reflecting the costs of a day in hospital (as before), the new indices track the costs for treating certain conditions (cf. CATRON/MURPHY 1996). The new indices present a picture of lower medical care price inflation than the old components of the Consumer Price Index (CPI). The BLS

⁴ Baumol affirms this with respect to computers. He writes: "Despite the use of the computer in medicine ..., there is no substitute for the personal attention of a physician..." (BAUMOL 1967: 423).

subsequently introduced a similar method into the CPI. Also, the *Bureau of Economic Analysis* (BEA), which is in charge with the U.S. National Accounts, uses the new PPI hospital care indices as deflators. The rate of labor productivity growth, which had been negative before (−0.5 percentage points per year over the period 1987-1995) turned positive (+0.9 PP p.a. over the period 1995-2001), partly because of the new deflators (cf. TRIPLETT/BOSWORTH 2004: 262-3).

One might argue, though, that medical price indices are still upward biased since, for instance, nothing has been done to remove the quality change bias mentioned above. Obviously, it is very difficult if not impossible for statistical offices to measure certain quality changes in the provision of (health) services that come along with new technologies and procedures. If medical price indices as they exist today are probably not a proper reflection of true price developments – and may never be – the question arises how informative it is to test Baumol's theory with them as has been done in previous studies. Let's have a look at international developments in medical care prices.

2 International developments in medical care prices

Figure 2 shows the development of the price index of medical care in relation to the GDP deflator for 15 countries (for which data are available from the *OECD Health Database 2005*). The picture looks like a fan. In most of the countries, the relative price of medical care is rising – which supports Baumol's theory. As was mentioned above, the model predicts that prices in the tertiary sector grow faster than in the secondary sector and that this results in both high expenditure growth for services and the 'Cost Disease'. On the other hand, there are countries like Spain and Denmark where medical care prices have not increased faster than the overall price level since the mid-seventies. In Australia and France, the relative price of medical care even tends to decrease. This is also the case for Italy since 1990. The oddest example is provided by Turkey. Here, medical care prices have risen almost 40 per cent faster than the overall price level between 1982 and 1990. Afterwards, the relative price index dropped by over 90 per cent within only seven years. Three years later – there is a data gap of two years –, the relative price of medical care had doubled again.

<Insert Figure 2 around here>

How can patterns as divergent as these emerge? Firstly, they can result from differences in measurement methods between countries. AHMAD ET AL. (2003) present deflators for software investment as another example of fan-like price indices. Between

1995 and 2000, the price index for Australia fell by almost 30 per cent, while the price index for Sweden rose by approximately the same amount. Ahmad et al. comment: "This largely reflects the dearth of price information available in this area, and the wide range of alternative proxy deflators that are used to represent price trends for software investment" (2003: 20). This statement naturally carries over to medical care.

Even more importantly, health care price trends differ between countries because, unlike the U.S., most OECD countries do not rely on market provision of health care. "In most OECD member countries, health care is provided largely by the government sector" (BERNDT ET AL. 2000: 171). In countries with a nationalized health system, prices are directly set by the government. In other countries where at least parts of the medical services are provided by the private sector, prices are often agreed upon between the providers' associations and the financing body or bodies, for example associations of health insurances. The government may again exert influence on the price formation. National price trends in health care must thus be expected to be as diverse as national health systems. Most notably, price trends cannot be expected to reflect the true development of health care costs in every country as some governments deliberately subsidize certain health services in order to prevent medical care prices from rising in line with costs. In Switzerland, for instance, approximately half of the total expenditure on hospital services is paid for by the federal states (cantons). The rest is reimbursed by the sickness funds on a fee-for-service base. Obviously, prices for hospital services would be quite different without the public subsidies. In fact, the 'Cost Disease' can affect the health sector without any rise in medical care prices. In an environment of heavily regulated prices, there are ways to shift resources into this sector other than raising prices.

3 Consequences for testing Baumol's model

NEWHOUSE (2001: 13) has argued that medical care prices are needed to test Baumol's 'Cost Disease' theory. The model implies that the relative price of services like health care rises, which will result in increases in health care costs. However, since health expenditure can increase without any rise in the price level, medical care price indices are probably not appropriate variables to test this theory. Therefore, in earlier work (HARTWIG 2006), I tested Baumol's theory in a different way by checking the model's implication that wage increases in excess of aggregate labor productivity growth raise health care expenditure in a directly proportional manner. However, other studies have used relative medical care prices as explanatory variables for aggregate health expenditure. As far as I see, L'HORTY ET AL. (1997) is the only one of these studies that

explicitly links the results to Baumol's theory. L'Horty and colleagues try to explain French real per-capita health expenditure over the period 1970-95 in an error-correction framework. Their set of explanatory variables includes the relative price of medical care, which has a consistently negative, and statistically significant, coefficient in their estimations. Therefore, L'Horty et al. conclude that Baumol's theory was not supported by the data. Of course, Figure 2 shows that France is one of the few countries where the relative price of medical care goes down due to rigorous price controls, as L'HORTY ET AL. (1997: 263) acknowledge themselves.

The other studies mentioned which use the relative price of medical care as an explanatory variable for health expenditure can be regarded as implicitly testing Baumol's model. The evidence produced by these studies is mixed, however. Recently, OKUNADE ET AL. (2004) found a significantly positive influence of the relative price of health care on real HCE growth in OECD data – thus lending support to Baumol's theory. Four years earlier, and with another co-author, Okunade had still expressed the view that “the price variable – i.e., the relative health care price – is dubious, as its proxies are problematic and their associated measurement errors are severe. Thus the price variable is excluded from our proposed model” (MURTHY/OKUNADE 2000: 281). ROBERTS (1999) finds mixed evidence depending on which estimator is used to evaluate the impact of the relative price of health care on real health expenditure in OECD countries. The mean group estimator, in which separate time series regressions are estimated for each group first, and the coefficients are then averaged across groups yields a significantly positive coefficient in the static case, but an insignificant (positive) coefficient in the dynamic case. The same is true for pooled estimations. In a cross-section regression using averages for 1960 to 1993, the coefficient turns out to be negative, but not statistically different from zero. Non-significant negative coefficients are also found by MURTHY/UKPOLO (1994) analyzing US time series data over the period 1960-87 and GERDTHAM ET AL. (1992) in a cross-section of 19 OECD countries in 1987. GERDTHAM/JÖNSSON (1991) find a positive, but insignificant coefficient in a cross-section of 22 OECD countries in 1985.

Overall, no clear picture has emerged so far from explicit and implicit tests of Baumol's model of unbalanced growth on the basis of relative health care prices. Although I share the reservations expressed by Murthy & Okunade in the quote given above – in fact, I've argued throughout that medical care price indices are inappropriate variables to test Baumol's theory, especially in an international setting –, I will perform such a test in the next section. My motivation is to check whether the – from the

model's point of view – predominantly unfavorable empirical evidence persists when the test meets certain criteria for testing explanatory variables for HCE that have been proposed only relatively recently.

4 Testing Baumol's model of unbalanced growth using relative medical care prices

In this section, the implication of Baumol's model of unbalanced growth that the rise in relative medical care prices drives health expenditure will be tested empirically. Basically, I will revisit the approach of GERDTHAM/JÖNSSON (1991) who regress the log of per-capita health care expenditure on the log of per-capita GDP and the log of the relative price of health care. Per-capita GDP (or national income) is *the only* variable that has emerged from the literature of the last thirty years as a robust explanatory variable for per-capita HCE (cf. ROBERTS 1999: 459, GERDTHAM/JÖNSSON 2000: 45-6). The relative price of health care is introduced in order to test Baumol's model. It is convenient to focus on these two explanatory variables.

The GERDTHAM/JÖNSSON (1991) estimation approach will be modified here in several ways in order to consider certain propositions that have been made more recently on how to improve the modeling of health expenditure. First, instead of a cross-section, a panel of OECD data is tested. An important advantage of working with pooled data is that it allows for richer specifications including country and time-specific effects. Secondly, I concentrate on the 'total *current* expenditure on health' 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. Time series for both current expenditure on health and relative medical care prices that start 1980 or earlier are available for nine countries from the *OECD Health Database* 2005. These are Australia, Canada, Denmark, Finland, Iceland, Ireland, Spain, Switzerland, and the U.S. The HCE time series for Italy and France start in 1988 and 1990, respectively. Inclusion of these two countries would add 31 HCE observations, yet the sample would become quite unbalanced. Therefore, I will preliminarily abstain from adding them. The nine-country sample consists of 332 observations for HCE and 295 observations for the relative price of medical care.

A third difference between the approach of GERDTHAM/JÖNSSON (1991) and that of the present paper is that all variables are here modeled in log differences (growth rates) instead of log levels. In doing so, I follow the recommendation of GERDTHAM/JÖNSSON

(2000: 48) in their important contribution for the *Handbook of Health Economics*. Modeling growth rates instead of levels has the major advantage that it avoids the cumbersome and so far unresolved issue of the degree of integration of the various variables. MCCOSKEY/SELDEN (1998), for instance, 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. The upshot of this debate is that we can't be sure about the degree of integration of health expenditure variables. By specifying the regression in growth rates, we are unlikely to produce spurious results since no-one has proposed that HCE are $I(2)$.⁵

As has been mentioned, working with pooled data 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 9 countries is drawn somewhat randomly from the group of OECD countries according to data availability. We will present results of cross-section fixed-effects estimations along with the random-effects estimations, though. The estimation period covers the years from 1971 to 2003.

Table 1 summarizes the estimation results. The coefficient of the relative price variable is positive and significantly different from zero at the 1 per cent level in all

⁵ 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.

specifications, and GDP growth is significant with coefficient values of about 1. To check whether the correlation between HCE and GDP growth is not just picking up purely monetary changes, I deflate both per-capita HCE and GDP by the GDP deflator. Table 2 shows that the coefficient of the GDP growth variable drops markedly compared to the nominal case, and it becomes more unstable across specifications. Nevertheless, real per-capita GDP growth remains a statistically significant explanatory variable for the deflated per-capita HCE growth. The coefficient of the relative price variable remains positive and statistically significant. Compared to the nominal case, it becomes more stable across specifications, and the t-values even rise a little. The adjusted R^2 drops markedly, though, indicating that general price fluctuations that affect both GDP and HCE contribute a lot to the explanatory power of the nominal specification. Overall, the results corroborate the recent finding by OKUNADE ET AL. (2004) as well as, of course, Baumol's model.

<Insert Tables 1 and 2 around here>

5 Robustness analysis

I now turn to the question how robust the results are to variations in the modeling framework. I concentrate on the nominal models since Baumol's theory of unbalanced growth is formulated thoroughly in nominal terms. Also, we have seen that the results concerning the relative price variable are not much affected by the deflation procedure. In testing parameter stability over time, I follow GERDTHAM (1992) by splitting the period of observation into three arbitrary sub-periods, namely 1971-1981, 1982-1992, and 1993-2003. Tables 3a and 3b show that the growth rate of the relative price of health care loses its statistical significance as an explanatory variable for HCE growth in the second sub-period both in the random-effects and fixed-effects models. The estimated coefficient retains a positive sign, though. In the other two periods, the relative price variable remains significant (except in the first period in model 20). If significant, the coefficients are noticeably higher than in the estimation covering the whole observation period.

<Insert Tables 3a and 3b around here>

Per-capita GDP growth is statistically significant mostly at the 1 per cent level. There are somewhat greater differences in the magnitude of the coefficients between the

random-effects and the fixed-effects estimations than for the whole period. Also, the explanatory power of the model varies, being lowest for the most recent sub-period.

The second parameter stability test proposed by GERDTHAM (1992) involves varying the sample of countries. To reproduce this test, I re-estimate the model dropping each of the nine countries in turn. This exercise also offers the opportunity to include the Italian and French data that have not been considered so far. Table 3 reveals some instability of the estimated parameters. In the cross-section random-effects estimation, for instance, the coefficient of the relative price variable varies between 0.425 (when Finland is excluded) and 0.625 (when Ireland is dropped). The instability is less marked than in the former test of inter-temporal stability, however, which is not surprising since the former test was more restrictive – it left out up to three fourths of the complete sample. The exclusion of Iceland (being a high inflation country for most of the estimation period) has the largest impact on the coefficients as well as on the explanatory power of the specification. (This has already been reported by GERDTHAM 1992.) Most notably, the income elasticity of health care expenditure falls below one when Iceland is not part of the sample. The coefficient of the relative price variable is hardly affected, though. No matter which country is dropped, or whether France and Italy are included, the coefficient of relative price growth is positive and significantly different from zero at least at the five per cent level.

<Insert Tables 4a and 4b around here>

6 Summary and conclusion

In their important contribution for the *Handbook of Health Economics*, GERDTHAM/JÖNSSON (2000: 48) call for “more theory of the macroeconomics of health expenditure”. In previous work I argued that Baumol’s (1967) model of unbalanced growth offers a ready explanation for the observed secular rise in health care expenditure (HCE) in rich countries. This model can be tested empirically in various ways. In HARTWIG (2006) I tested the model’s implication that HCE is driven by wage increases in excess of productivity growth using data from a panel of 19 OECD countries. I found robust evidence in favor of Baumol’s theory. Another implication of Baumol’s model is that variations in the relative price of medical care contribute significantly to explaining variations in health expenditure in the same direction. Earlier studies have sometimes included the relative price of medical care in the list of explanatory variables for HCE, although usually not in an explicit attempt to test Baumol’s theory. The empirical evidence from these studies is mixed. Sometimes the coefficient for the relative price

variable is found to be positive and statistically significant – which would support Baumol’s theory –, in other studies the coefficient turns out to be either insignificant or even negative, which would speak against Baumol’s model.

Despite severe reservations against using medical care price indices at all in statistical analyses because of their poor data quality, I re-perform a pooled regression of HCE on (GDP and) relative medical care prices using data for 9 OECD countries for which data on both current HCE and medical care prices are available over a considerable period of time. I specify the regression in growth rates, following an advice given by GERDTHAM/JÖNSSON (2000). I find the growth rate of relative medical care prices to contribute positively, and significantly, to HCE growth. This finding, which corroborates recent results by OKUNADE ET AL. (2004), is reasonably robust to varying the country sample or the sample period, although the relative medical care price variable temporarily loses statistical significance during the eighties. Overall, the empirical evidence in favor of Baumol’s model of unbalanced growth seems to be growing.

References

- Ahmad, N., Lequiller, F., Marianna, P., Pilat, D., Schreyer, P., Wölfl, A. (2003): Comparing labour productivity growth in the OECD area: the role of measurement, OECD Directorate for Science, Technology and Industry Working Paper, No. 14/2003.
- Baumol, W. J. (1967): Macroeconomics of unbalanced growth: the anatomy of urban crisis, *American Economic Review*, 57(3): 415-26.
- Baumol, W. J., Towse, R. (eds.) (1997): *Baumol’s Cost Disease: the Arts and Other Victims*, Aldershot: Edward Elgar.
- Berndt, E. R., Cutler, D. M., Frank, R. G., Griliches, Z., Newhouse, J. P., Triplett, J. E. (2000): Medical care prices and output, in: Culyer, A. J., Newhouse, J. P. (eds.): *Handbook of Health Economics*, Vol. 1A, Amsterdam: Elsevier.
- Boskin, M. J., Dulberger, E. R., Gordon, R. J., Griliches, Z., Jorgenson, D. W. (1996): Toward a More Accurate Measure of the Cost of Living: Final Report to the Senate Finance Committee, Washington, D.C. (<http://www.ssa.gov/history/reports/boskinrpt.html>).
- Carrion-i-Silvestre (2005): Health care expenditure and GDP: Are they broken stationary?, *Journal of Health Economics*, 24 (5): 839-854.
- Catron, B., Murphy, B. (1996): Hospital price inflation: What does the new PPI tell us?, *Monthly Labor Review*, 120(7): 24-31.

- Gerdtham, U.-G. (1992): Pooling international health care expenditure data, *Health Economics*, 1(3): 217-31.
- Gerdtham, U.-G., Jönsson, B. (1991): Price and quantity in international comparisons of health care expenditure, *Applied Economics*, 23(9): 1519-28.
- Gerdtham, U.-G., Jönsson, B. (2000): International comparisons of health expenditure: theory, data and econometric analysis, in: Culyer, A. J., Newhouse, J. P. (eds.): *Handbook of Health Economics*, Vol. 1A, Amsterdam: Elsevier.
- Gerdtham, U.-G., Löthgren, M. (2000): On stationarity and cointegration of international health expenditure and GDP, *Journal of Health Economics*, 19(4): 461-75.
- Gerdtham, U.-G., Soogard, J., Andersson, F., Jönsson, B. (1992): Econometric analysis of health care expenditure: a cross-section study of the OECD countries, *Journal of Health Economics*, 11(1): 63-84.
- Hansen, P., King, A. (1998): Health care expenditure and GDP: panel data unit root test results – comment, *Journal of Health Economics*, 17(3): 377-81.
- Hartwig, J. (2006): What drives health care expenditure? – Baumol's model of 'unbalanced growth' revisited, KOF Working Paper No. 133 (http://www.kof.ch/pdf/wp_133.pdf).
- Jewell, T., Lee, J., Tieslau, M., Strazicich, M. C. (2003): Stationarity of health expenditures and GDP: evidence from panel unit root tests with heterogeneous structural breaks, *Journal of Health Economics*, 22(2): 313-23.
- Karatzas, G. (2000): On the determination of US aggregate health care expenditure, *Applied Economics*, 32(9): 1085-99.
- L'Horty, Y., Quinet, A., Rupprecht, F. (1997): Expliquer la croissance des dépenses de santé: le rôle du niveau de vie et du progrès technique, *Economie et Prévision*, 129-130(3/4): 257-68.
- McCoskey, S. K., Selden, T. M. (1998): Health care expenditures and GDP: panel data unit root test results, *Journal of Health Economics*, 17(3): 369-76.
- Murthy, N. R. V., Okunade, A. A. (2000): Managed care, deficit financing, and aggregate health care expenditure in the United States: A cointegration analysis, *Health Care Management Science*, 3(4): 279-85.
- Murthy, N. R. V., Ukpolo, V. (1994): Aggregate health care expenditure in the United States: evidence from cointegration tests, *Applied Economics*, 26(8): 797-802.
- Newhouse, J. P. (2001): Medical care price indices: problems and opportunities, *Academia Economic Papers*, 29(1): 1-65.

- Nixon, J., Ulmann, P. (2006): The relationship between health care expenditure and health outcomes. Evidence and caveats for a causal link, *European Journal of Health Economics*, 7(1): 7-18.
- OECD (2006): Projecting OECD health and long-term care expenditures: What are the main drivers?, OECD Economics Department Working Paper No. 477.
- Okunade, A. A., Karakus, M. C. (2001): Unit root and cointegration tests: time-series versus panel estimates for international health expenditure models, *Applied Economics*, 33(9): 1131-7.
- Okunade, A. A., Karakus, M. C., Okeke, C. (2004): Determinants of health expenditure growth of the OECD countries: jackknife resampling plan estimates, *Health Care Management Science*, 7(3): 173-83.
- Roberts, J. (1999): Sensitivity of elasticity estimates for OECD health care spending: analysis of a dynamic heterogeneous data field, *Health Economics*, 8(5): 459-72.
- Triplett, J. E. (ed.) (1999): *Measuring the Prices of Medical Treatments*, Washington, D.C.: Brookings Institution Press.
- Triplett, J. E., Bosworth, B. P. (2004): *Productivity in the U.S. Services Sector – New Sources of Economic Growth*, Washington, D. C.: Brookings Institution Press.
- Wilson, R. M. (1999): Medical care expenditures and GDP growth in OECD nations, *American Association of Behavioral and Social Sciences Journal*, 2: 159-71.
- Wölfl, A. (2005): The service economy in OECD countries, OECD – STI Working Paper 2005/3.

Acknowledgements

I wish to thank Sheila D. Smith and Philippe Ulmann for their thoughtful comments on earlier work that have inspired the present paper. This does not imply that they are in any way responsible for possible errors or shortcomings herein.

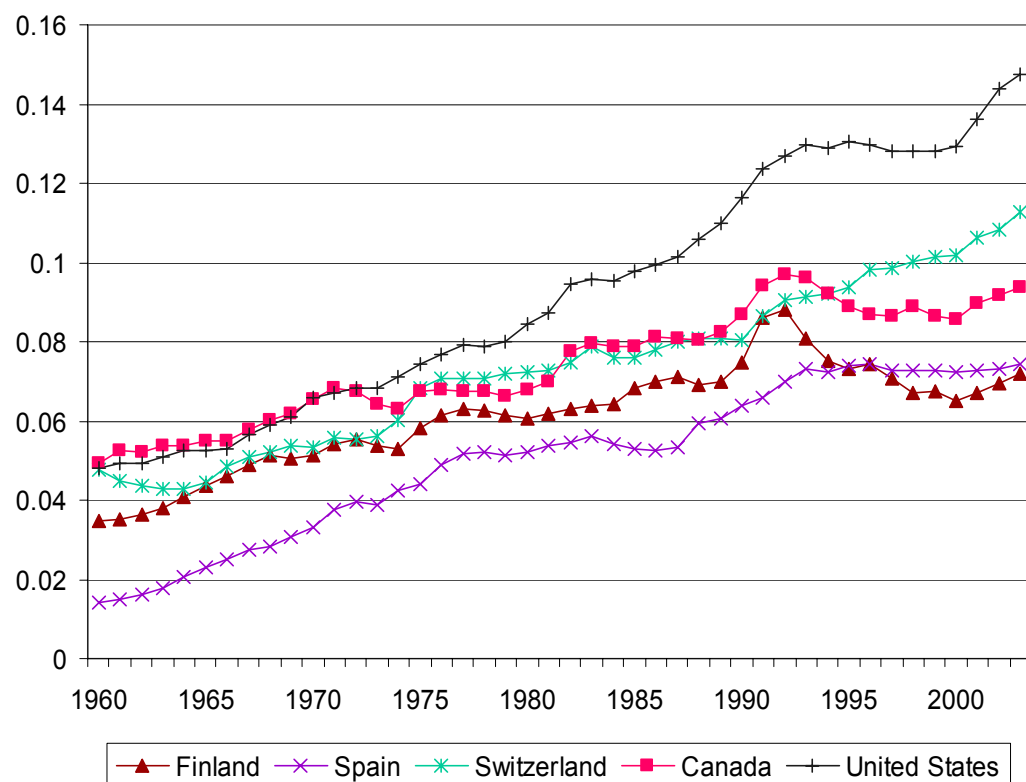


Figure 1: Shares of current health care expenditure in GDP

Source: OECD Health Data 2005 CD ROM

**Price index total expenditure on health relative to GDP,
2000 = 100**

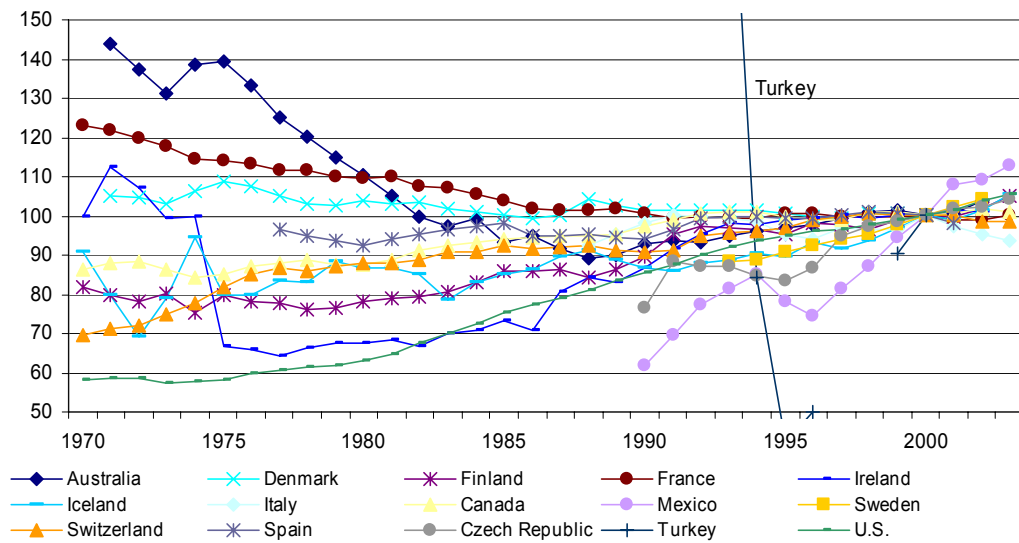


Figure 2: Relative prices of medical care for selected OECD countries

Source: OECD Health Data 2005 CD ROM

Table 1: Results for growth rate equations

Log difference of health care expenditure per capita					
	(1)	(2)	(3)	(4)	(5)
	OLS	Cross-section R.E.	Time period R.E.	Cross-section F.E.	Time period F.E.
dlog(GDPPC)	1.014* (17.515)	1.026* (15.645)	1.013* (18.181)	1.031* (14.785)	1.012* (16.307)
dlog(RELPR)	0.445* (2.834)	0.489* (2.814)	0.405* (3.038)	0.512* (2.794)	0.381* (2.759)
Number of obs.	257	257	257	257	257
R ² (adj.)	0.780	0.766	0.793	0.791	0.839
Stand. err. of regr.	0.035	0.034	0.030	0.034	0.030
D-W	1.414	1.508	1.405	1.560	1.410

dlog(GDPPC) = log difference of nominal per-capita GDP, dlog(RELPR) = log difference of the price of health care relative to the GDP deflator.

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

Log difference of health care expenditure per capita, deflated by the GDP deflator					
	(6)	(7)	(8)	(9)	(10)
	OLS	Cross-section R.E.	Time period R.E.	Cross-section F.E.	Time period F.E.
dlog(GDPRPC)	0.266* (2.953)	0.270* (2.867)	0.337* (4.145)	0.276* (2.701)	0.441* (4.860)
dlog(RELPR)	0.428* (3.615)	0.437* (3.465)	0.430* (3.550)	0.445* (3.313)	0.413* (3.059)
Number of obs.	257	257	257	257	257
R ² (adj.)	0.137	0.132	0.180	0.152	0.270
Stand. err. of regr.	0.030	0.029	0.028	0.029	0.027
D-W	1.211	1.242	1.225	1.279	1.262

dlog(GDPRPC) = log difference of real per-capita GDP, dlog(RELPR) = log difference of the price of health care relative to the GDP deflator.

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 3a: Robustness test I – 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(GDPPC)	0.916** (17.802)	0.984** (9.095)	0.679** (7.127)	0.917** (18.092)	1.038** (14.402)	0.773** (11.744)
dlog(RELPR)	0.606** (2.884)	0.234 (1.131)	0.792** (4.718)	0.594** (2.876)	0.261 (1.420)	0.858** (4.908)
Number of obs.	66	99	92	66	99	92
R ² (adj.)	0.692	0.758	0.295	0.695	0.834	0.524
Stand. err. of regr.	0.035	0.036	0.026	0.034	0.034	0.022
D-W	1.637	1.696	1.001	1.633	1.441	1.097

dlog(GDPPC) = log difference of nominal per-capita GDP, dlog(RELPR) = log difference of the price of health care relative to the GDP deflator.

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. * and ** denote significance at the 5 and 1 per cent level, respectively Estimates for constant terms not shown.

Table 3b: Robustness test I – Inter-temporal parameter stability

Log difference of health care expenditure per capita						
	Cross-section F.E.			Time period F.E.		
	(17)	(18)	(19)	(20)	(21)	(22)
	1971–1981	1982–1992	1993–2003	1971–1981	1982–1992	1993–2003
dlog(GDPPC)	0.677** (3.720)	0.966** (8.323)	0.559* (2.557)	0.943** (20.214)	1.060** (14.720)	0.784** (10.813)
dlog(RELPR)	0.893** (4.242)	0.235 (1.050)	0.710** (3.271)	0.400 (1.936)	0.281 (1.481)	0.853** (4.239)
Number of obs.	66	99	92	66	99	92
R ² (adj.)	0.729	0.833	0.471	0.774	0.851	0.607
Stand. err. of regr.	0.033	0.036	0.026	0.030	0.034	0.022
D-W	1.992	1.822	1.017	1.580	1.411	1.167

dlog(GDPPC) = log difference of nominal per-capita GDP, dlog(RELPR) = log difference of the price of health care relative to the GDP deflator.

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. * and ** denote significance at the 5 and 1 per cent level, respectively Estimates for constant terms not shown.

Table 4a: Robustness test II – Cross-national stability of parameters

Log difference of health care expenditure per capita, Model (2)										
	Italy/France	Australia	Canada	Denmark	Finland	Iceland	Ireland	Spain	Switzerland	U.S.
dlog(GDPPC)	1.023** (16.035)	1.019** (15.187)	1.037** (16.260)	1.023** (15.577)	1.025** (18.926)	0.910** (10.409)	1.048** (15.519)	1.026** (15.183)	1.040** (15.344)	1.031** (16.010)
dlog(RELPR)	0.497** (3.006)	0.501* (2.350)	0.438* (2.505)	0.461** (2.630)	0.360* (2.567)	0.517** (2.704)	0.625** (3.595)	0.489** (2.736)	0.465** (2.657)	0.474** (2.628)
Number of obs.	285	227	224	236	224	234	230	233	224	224
R ² (adj.)	0.763	0.772	0.776	0.770	0.784	0.549	0.808	0.768	0.773	0.772
Stand. err. of regr.	0.033	0.034	0.034	0.035	0.032	0.033	0.032	0.034	0.035	0.035
D-W	1.495	1.516	1.522	1.475	1.607	1.262	1.632	1.509	1.487	1.516

Log difference of health care expenditure per capita, Model (3)										
	Italy/France	Australia	Canada	Denmark	Finland	Iceland	Ireland	Spain	Switzerland	U.S.
dlog(GDPPC)	1.011** (18.107)	1.018** (18.039)	1.024** (19.015)	1.009** (18.513)	1.037** (15.847)	0.820** (13.505)	1.039** (17.643)	1.011** (17.548)	1.032** (16.555)	1.019** (16.600)
dlog(RELPR)	0.428** (3.372)	0.450* (2.345)	0.368** (2.742)	0.373** (2.798)	0.442* (2.348)	0.372* (2.510)	0.545** (4.407)	0.410** (2.983)	0.398** (2.854)	0.345* (2.306)
Number of obs.	285	227	224	236	224	234	230	233	224	224
R ² (adj.)	0.795	0.805	0.805	0.799	0.821	0.531	0.836	0.803	0.791	0.840
Stand. err. of regr.	0.029	0.030	0.031	0.030	0.029	0.028	0.027	0.030	0.031	0.031
D-W	1.422	1.418	1.410	1.375	1.466	1.185	1.618	1.390	1.395	1.394

dlog(GDPPC) = log difference of nominal per-capita GDP, dlog(RELPR) = log difference of the price of health care relative to the GDP deflator.

The values shown in parenthesis are t-ratios, based on White's robust S.E.s. The Swamy-Arora GILS estimator was used to estimate the random effects models, and weighted diagnostic statistics are reported. * and ** denote significance at the 5 and 1 per cent level, respectively. Estimates for constant terms not shown.

Table 4b: Robustness test II – Cross-national stability of parameters

Log difference of health care expenditure per capita, Model (4)										
	Italy/France	Australia	Canada	Denmark	Finland	Iceland	Ireland	Spain	Switzerland	U.S.
dlog(GDPPC)	1.029** (14.978)	1.021** (14.699)	1.043** (15.497)	1.035** (14.625)	1.037** (15.847)	0.929** (9.120)	1.046** (14.510)	1.031** (14.458)	1.042** (14.782)	1.036** (14.973)
dlog(RELPR)	0.523** (3.003)	0.501* (2.325)	0.457* (2.486)	0.501** (2.650)	0.442* (2.348)	0.562** (2.868)	0.639** (3.315)	0.508** (2.709)	0.489** (2.643)	0.507** (2.692)
Number of obs.	285	227	224	236	224	234	230	233	224	224
R ² (adj.)	0.793	0.799	0.804	0.790	0.815	0.584	0.824	0.797	0.796	0.797
Stand. err. of regr.	0.033	0.034	0.034	0.035	0.032	0.033	0.032	0.034	0.035	0.035
D-W	1.553	1.554	1.571	1.538	1.656	1.326	1.676	1.557	1.536	1.576

Log difference of health care expenditure per capita, Model (5)										
	Italy/France	Australia	Canada	Denmark	Finland	Iceland	Ireland	Spain	Switzerland	U.S.
dlog(GDPPC)	1.010** (16.142)	1.022** (16.558)	1.025** (15.497)	1.010** (16.860)	1.028** (17.080)	0.774** (11.474)	1.037** (15.832)	1.010** (15.493)	1.032** (14.158)	1.019** (14.973)
dlog(RELPR)	0.410** (3.151)	0.418* (2.052)	0.343* (2.478)	0.352* (2.529)	0.339* (2.348)	0.330* (2.157)	0.515** (3.793)	0.391** (2.712)	0.378* (2.541)	0.507** (2.692)
Number of obs.	285	227	224	236	224	234	230	233	224	224
R ² (adj.)	0.843	0.844	0.843	0.844	0.855	0.704	0.878	0.847	0.836	0.797
Stand. err. of regr.	0.029	0.030	0.031	0.030	0.029	0.028	0.026	0.030	0.031	0.035
D-W	1.427	1.409	1.432	1.375	1.470	1.120	1.636	1.392	1.536	1.576

dlog(GDPPC) = log difference of nominal per-capita GDP, dlog(RELPR) = log difference of the price of health care relative to the GDP deflator.

The values shown in parenthesis are t-ratios, based on White's robust S.E.s. The Swamy-Arora GILS estimator was used to estimate the random effects models, and weighted diagnostic statistics are reported. * and ** denote significance at the 5 and 1 per cent level, respectively. Estimates for constant terms not shown.