The dynamics of health

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1. The demand for health

Health has long been considered as a fundamental commodity in economic analyses; Michael Grossman (2000) cites Bentham as recognizing that the ‘relief of pain’ is one of the basic arguments in the utility function. Health was viewed both as an investment in human capital and as an output of a household production process by Grossman (1972a & b), the founding father of demand for health models. In the Grossman model, health is both demanded for utility reasons - it is good to feel well - and for investment reasons – to make more healthy time available for market and non-market activities.

Grossman developed a dynamic model for health and solution of the dynamic optimisation problem leads to optimal life-cycle health paths, gross investment in each period, consumption of medical care (which is seen as a derived demand) and time inputs in the gross investment function in each period. By comparing maximum lifetime utility for different lengths of life, it also allows endogenous determination of the length of life. Usually the comparative static and dynamic analyses are performed on sub-models where either the consumption benefits are assumed to equal zero (the investment model), or the investment benefits are assumed to equal zero (the consumption model). We focus on the investment model as sharper predictions are available; this model results in a condition which determines the optimal stock of health in any period and shows that the rate of return on capital (or, marginal efficiency of capital, MEC) must equal the opportunity cost of capital.

Increases in the depreciation rate over time cause the optimal stock of health to decrease, as the opportunity cost of capital increases. However, if the MEC curve is inelastic, gross investment grows over time. Thus the model predicts older people to have more sick time, to consume more medical care and devote more time to investment in health than younger people. An increase in the wage rate shifts the MEC to the right and steepens it simultaneously. Thus the demand for health capital rises. The demand for medical care also rises with the wage, as gross investment increases. However, due to substitution incentives between medical care and time, the elasticity of medical care with respect to the wage is greater than that of health. The more educated would demand more health but, if the MEC elasticity with respect to education is less than one, less medical care.

We note in passing that wage effects and schooling effects on health are ambiguous in the pure consumption or mixed consumption-investments models. Theoretical extensions of the Grossman model have introduced uncertainty of health to the model. Compared to a model with certainty, the optimal level of the stock of health is larger and the optimal quantities of investment are also larger in a pure investment model. A change in initial assets can cause a change in these quantities but the effects depend on how uncertainty influences earnings (see Grossman, 2000).

Some empirical testing of the Grossman model has occurred. Grossman's (1972b) original study, using US data, found positive effects of education and wages on the demand for health. Age had a positive effect on health demand and a negative effect on medical care. However the wage effect was negative for medical care and schooling had a positive effect, although neither of these were statistically significant. Wagstaff (1986) uses a MIMIC model with Danish data, and estimates a conditional demand function for medical care and a health demand equation. His results generally accord with the predictions of the investment model. However, he finds that good health has a large negative effect on physician visits. This may well be due to the assumption of exogeneity. The coefficient on health will be biased downwards if no account is taken of an unobservable initial period depreciation rate and other shocks to health. Erbsland et al. (1995) find similar results to Wagstaff using a MIMIC model with a West German sample. Both Van Doorslaer (1987) and Wagstaff (1993) use longitudinal data to overcome the problems introduced by time-invariant unobservables (such as the initial period depreciation rate) and to allow for costs of adjustment to the desired health stock.

Another major issue in the demand for health literature concerns the impact of schooling on health (see Grossman, 2000). The question is: 'how much of the observed correlation in these variables is due to a causal effect of schooling on health?' Differences in time preference rates and the 'healthy student' and 'anticipated health' effects have been the focal points of discussion. Furthermore, even if the effect is causal, is it due to the impact of education on allocative efficiency, technical efficiency, or preferences (including the rate of time preference)? This is of course relevant for policy, as health-specific education may have large effects under some explanations but perform poorly under others.

2. Reasons for persistence in health

Empirical analysis of health dynamics needs to address the question why some individuals experience persistently good health and others experience persistently poor health. To some extent this may reflect the nature of health problems: some illnesses are inherently chronic and long-lasting. Also a cumulative history of a range of health problems may have a direct influence on current health and on the effectiveness of medical care. These effects can be thought of as pure dynamics, often termed state dependence. But health problems may be persistent for other reasons. As predicted by the demand for health framework, an individual may have individual or socioeconomic characteristics that predispose them to poorer health and that linger over time. Factors such as education, material deprivation, childhood nutrition and environment may have a long-standing influence on an individual's health. Some of these factors may be observable, but others - such as ability, time preference and risk aversion - may be hard to control for.

These issues are exemplified by the debates over the association between health and socioeconomic status (SES): in particular health and education (see e.g., Grossman, 2000, Smith, 2004) and health and income or wealth (see e.g., Smith, 1999, 2004). Evidence of a positive association between health and SES is well-documented across many societies and periods (see e.g. Smith, 1999, Deaton, 2003). But the causal mechanisms underlying this relationship are complex and controversial. There can be a direct causal link from SES to health, for example, through the direct influence of material deprivation on the production of health and on access to health care or of education on the uptake and compliance with medical treatments. There can be a direct causal link from health to SES, for example, through the impact of health shocks on labour market outcomes such as unemployment, early retirement and earnings. But there may also be pathways that link health and SES through "third factors", for example time preference rates, that do not imply any causal link.

A full discussion of the empirical literature on the relationship between health and SES – even that originating within economics - is beyond the scope of this chapter. Readers are referred to Adams et al. (2003) and the associated commentaries for a discussion of the methodological issues involved in identifying causal effects of income on health. The problem facing such analyses is to overcome the biases, caused by reverse causality and selection, by identifying a source of exogenous variation in income. Meer et al. (2003) use inheritance as an instrument for wealth changes using the PSID from 1984-1999. They find a very small effect of changes in wealth requiring a quarter million dollar change to achieve an effect of around 2% points in probability of excellent of good health. Frijters et al. (2003) use the unanticipated shift in permanent income for East German households, following the reunification of Germany, as a source of exogenous variation in income. Using the GSOEP between 1991 and 1999 they find that, after controlling for heterogeneity, there is no evidence for a causal effect of income on health. Of course these findings do not rule out a cumulative and long-term relationship running from SES during childhood and early life to the gradient in adult health (see e.g., Case et al., 2002, Smith, 2004).

3. Empirical analysis of the dynamics of health

Kerkhofs and Linderboom (1997) investigate the impact of changes in labour market status and work history on health. Recognising that health changes with age and that individuals may experience shocks to their health that have permanent consequences for their future health, their empirical model incorporates health transitions by specifying a model in first-differenced form. By conditioning on a variable representing a health shock in the intervening years between the waves of data used, the authors control for previous health experiences. They report that the occurrence of a health shock causes large changes in individual health profiles.

As part of an analysis of inequalities in mental health, Hauck and Rice (2004) investigate health mobility. Concern focuses on the extent to which individuals migrate within the distribution of mental health and how this varies across socio-economic groups. Using 11 waves of data from the British Household Panel Survey (BHPS), two regression based approaches to estimating mobility are employed. The first method is based on the intra-unit correlation coefficient, $\rho = \sigma_{\alpha}^{2} / \left(\sigma_{\alpha}^{2} + \sigma_{\varepsilon}^{2}\right)$ from an error components panel data model where σ_{α}^{2} represents the variance of the unobserved individual effect and σ_{ε}^{2} the variance of the idiosyncratic error component. The coefficient represents the correlation of health scores across periods of observation. Values of ρ close to unity indicate high persistence (low mobility) in health outcomes, whereas values of ρ close to zero are indicative of high random fluctuations resulting in high health mobility (low persistence). The second method is based on the estimated coefficient on lagged health status from a dynamic panel data model. The size of the estimated coefficient is informative about the degree of dependence between previous health and current health status. A coefficient close to zero indicates high mobility, whereas a coefficient close to unity indicates of high persistence.

Hauck and Rice (2004) find evidence of substantial mobility in mental health with the extent differing systematically across socio-economic groups. In general, mental health deteriorates with age and becomes more permanent in nature. Individuals from lower income groups are associated with greater mental ill-health, which exhibits greater persistence over time compared to individuals from higher income groups. Low educational attainment is also found to be associated with a greater reporting of mental health problems and greater persistence than that found for individuals from higher educational groups.

Investigating the association between socio-economic status and health in older populations, by drawing on the health investment framework of Grossman's model, Salas (2002) specifies a dynamic model of the determinants of health which includes previous health status as an exogenous regressor. The model is estimated using two waves (1991 and 1995) of the BHPS and concentrates on a 5-category self-assessed health variable (excellent, good, fair , poor, very poor), together with mortality, as the outcomes of interest. The results indicate a gradient over the categories of lagged self-assessed health status that appears statistically and substantively important – poor health status reported at time *t-1* is a powerful predictive of poor health status at time *t*. The relationship between socio-economic status and survival is less well established leading the author to conclude that while, for the elderly, low socio-economic position has a significant effect over a large range of the latent health variable it might not have sufficient effect to push people past a 'death threshold'.

Buckley et al (2004) also consider the relationship between socio-economic position and health transitions in the older population. Using three consecutive years of data from the Canadian Survey of Labour and Income Dynamics (SLID), and dichotomising an ordered categorical variable on reported health status into good or poor health, the authors estimate a probit model representing the probability of remaining in good health, given that an individual reported good health on entering the study period. Interest focuses on the relationship between income and health and in particular, the impact of long-term income. To this end, the authors construct a measure of household income standardised for age that it is claimed represents an indicator of lifetime income. The probability of remaining in good health for men in the highest quartile for both income and educational attainment is estimated to be 0.15 higher than men in the corresponding lowest quartiles. For women the difference in probabilities is greater at 0.18.

Contoyannis, Jones and Rice (2004a) consider the determinants of a binary indicator for the existence of functional limitations using seven waves (1991–1997) of the BHPS. The model adopted allows for persistence in observed outcomes due to state dependence, unobservable individual effects (heterogeneity), and autocorrelation in the transitory error component. The length of the panel used allows state dependence - the influence of health history on current health – to be investigated with greater rigour than previous studies have allowed. They estimate static and dynamic panel probit models by Maximum Simulated Likelihood (MSL) and consider two approaches to dealing with the problem of initial conditions in models with unobserved effects and lagged dependent variables.

The results suggest that a sufficient parameterization for the error process in such longitudinal models comprises a random effects structure with the addition of a first-order autocorrelated error component. For both men and women they find that, in the models which do not allow for state dependence, the addition of a serially correlated error component reduces the proportion of variance explained by the individual unobserved effect. The magnitude of this reduction (10%) is very similar to the effect of allowing for state dependence. This suggests that the proportion of variance due to time-invariant unobservable factors, and hence the amount of outcome persistence due to these factors, is overestimated in models which do not allow for dynamics. State dependence explains more of the persistence of functional limitations for men than for women.

In a separate study Contoyannis, Jones and Rice (2004b) explore the dynamics of self-assessed health again drawing on data from the BHPS. The variable of interest is the same 5-category ordered measure of self-assessed health used by Salas (2002) and, as with functional limitations, the BHPS reveals evidence of considerable persistence in individual's health status. Two possible sources of this persistence are unobservable heterogeneity and state dependence. Contoyannis, Jones and Rice (2004b) develop an econometric model using a latent variable specification for self-assessed health. The models include measures of socioeconomic status together with the inclusion of lagged health states designed to capture state dependence. To deal with the initial conditions problem an attractively simple approach suggested by Wooldridge (2002) is used.

The sample data used consists of the first 8 waves of the BHPS and care is taken to investigate and control for the consequences of longitudinal attrition of sample respondents: attrition rates are higher among those whose report very poor health at the previous wave of the panel. This health-related attrition may be a source of bias. To control for potential attrition bias inverse probablity weights are applied to the pooled ordered probit estimators. While there is evidence of health-related attrition in the data, the average partial effects of socioeconomic variables and of lagged health status are not influenced by sample attrition. This echoes earlier work by Kerkhofs and Lindeboom (1997), and is reinforced by further analysis of the BHPS and of the European Community Panel (ECHP) in Jones, Koolman and Rice (2004).

Contoyannis, Jones and Rice (2004b) present evidence of persistence in self-assessed health explained in part by state dependence which is stronger amongst men than women, and by individual heterogeneity, with around 30-35% of the unexplained variation accounted for by individual heterogeneity. As with functional limitations, there is evidence of a socioeconomic gradient by education and income.

4. Mobility indices

Health economists have been at the forefront of developing analytic tools for the measurement and explanation of health inequalities. The concentration index of health on income (Wagstaff et al., 1989) is a widely adopted measure of relative income-related health inequality and regression-based decomposition methods for the concentration index, of Wagstaff et al. (2003), have been used in a variety of settings and populations. While these methods have typically been applied to cross sectional information, it is desirable that attention should be paid to the dynamics of health and their relation to socio-economic characteristics as revealed by longitudinal data.

Jones and López-Nicolás (2004) show that there are important features of health inequality that cannot be revealed by cross sectional data. Their starting point is provided by measurement tools from the income distribution literature. In order to approximate a measure of inequality in lifetime income, Shorrocks (1978) considered inequality in the distribution of individual incomes averaged over a sequence of time periods. In particular, he introduced the concept of income mobility to capture the degree to which income inequalities fade as the time interval over which the population is analysed extends. Jones and López-Nicolás (2004) show that this methodology can be a useful empirical tool for the analysis of health inequality over the long-run. For example, a concern with inequalities in individuals' whole lifetime experience of health would mean taking an average over the whole lifespan.

Jones and López Nicolás (2004) show that the long-run concentration index can be written as the sum of a weighted average of short-run concentration indices plus a term that captures the covariance between levels of health (y) and fluctuations in income rank (*R*) over time:

$$CI^T = \sum_t w_t CI^t - \frac{2}{NT\bar{\bar{y}}^T} \sum_i \sum_t \left(y_{it} - \bar{y}^t\right)\left(R_i^t - R_i^T\right) \qquad where \quad w_t = \frac{\bar{y}^t}{T\bar{\bar{y}}^T}$$

This is a key result. Here it is presented in terms of the concentration index for socioeconomic inequalities in health, but the same idea also applies to the Gini coefficient for overall inequality in health or to extended Gini or concentration indices (see Wagstaff, 2002). The result shows that the concentration index (CI^T) for average health after T periods ($\bar{\bar{y}}^T$) can be written down as the sum of two terms. The first term is a weighted sum of the concentration indices (CI^t) for each of the sub-periods (with weights equal to the share of "total" health in each period). If the income ranking remains constant over time, a standard decomposition result for concentration indices implies that the concentration index for the average over time is equal to the (weighted) average of the concentration indices. However income ranks may change over time. The second term captures the difference between period-specific income ranks and ranks for average income over all periods and their relationship to health.

It is useful to measure how much the longitudinal perspective alters the picture that would emerge from a series of cross sections. Jones and López-Nicolás (2004) define an index of *health-related income mobility* as:

$$M^T = 1 - \frac{CI^T}{\sum_t w_t CI^t} = \frac{2}{N \sum_t \bar{y}^t CI^t} \left(\sum_i \sum_t (y_{it} - \bar{y}^t)(R_i^t - R_i^T) \right)$$

This is simply one minus the ratio by which the concentration index for the joint distribution of longitudinal averages differs from the weighted average of the cross sectional concentration indices. An important feature of the mobility index is that it is invariant to linear transformations of the measure of health. This extends the result used in van Doorslaer and Jones (2003) that, if y=α+βx, then $CI(y) = \frac{\alpha}{\alpha+\beta} CI(x)$. Given that $\bar{y} = \alpha + \beta\bar{x}$ this result can be substituted into the expression for the mobility index to show that $M^T(y)=M^T(x)$. The implication of this property is that the problem of imposing the appropriate cardinal scaling, discussed in van Doorslaer and Jones (2003), is less of a restriction for the analysis of mobility indices: the results will be the same for any linear transformation of the health status measure.

Using a concentration index to measure income-related inequalities in health raises the possibility to incorporate an econometric model for health and to decompose inequality into the contributions of each of the regressors (Wagstaff et al., 2003). By analogy, Jones and López-Nicolás (2004) show how health-related income mobility can be decomposed into the contributions of covariates in an econometric model. The methods are illustrated by analysing the dynamics of income and mental health, as measured by the GHQ index of psychological well-being in the first nine waves of the BHPS. The results reveal that, over eight years, adverse mental health becomes more concentrated among the poor. In particular, individual dynamics increase the absolute value of the concentration index of GHQ on income by 15% for men and 5% for women.

Lecluyse (2004) computes the mobility index proposed by Jones and López-Nicolás for a measure of self-assessed health (SAH), using the Panel Study of Belgian Households for 1994-2001. She applies the interval regression approach to the scaling of SAH, as suggested by van Doorslaer and Jones (2003), and finds a mobility index of 10% over the eight years of the panel, suggesting that the short-run measures underestimate inequality by 10% over the period.

To complement these earlier studies we present some new and previously unpublished findings based on the full eight waves, 1994-2001, of the *European Community Household Panel User Database* (ECHP-UDB) which was designed and coordinated by Eurostat. The ECHP is a multi-purpose annual longitudinal survey, based on a standardised questionnaire that involves annual interviewing of a representative panel of households and individuals of 16 years and older in each of the participating EU member states. We use data for the following fourteen member states of the EU for the full number of waves available for each: Austria (waves 2-8), Belgium (1-8), Denmark (1-8), Finland (3-8), France (1-8), Germany (1-3), Greece (1-8), Ireland (1-8), Italy (1-8), Luxembourg (1-3), Netherlands (1-8), Portugal (1-8), Spain (1-8) and the United Kingdom (1-3).

In Table 1 we present comparative evidence on the long-run concentration indices (CI^T) and mobility indices (M^T) for self-reported health limitations. The two binary measures of health limitations have been labelled HAMP1 and HAMP2. The information comes from the variable PH003A: respondents are asked: “Are you hampered in your daily activities by any physical or mental health problem, illness or disability?”, with three possible answers: 1. Yes, severely, 2. Yes, to some extent, 3. No. We use two dummy variables to indicate either any limitation (HAMP1) or a severe limitation (HAMP2). The ECHP income measure is disposable household income per equivalent adult, using the modified OECD equivalence scale, deflated by PPPs and the CPI for each country. Total household income includes all net monetary income received by the household members during the reference year. The estimates are computed using the personal weights supplied with the ECHP-UDB.

The negative values of the long-run concentration indices show evidence of income-related inequalities in health throughout Europe, with health limitations concentrated among those with lower incomes. The mobility indices show that there is greater long-run income related inequality in HAMP1 and HAMP2, than would be suggested by the average of short-run indices, for all countries. For the bulk of countries, that are observed for the full eight waves, health-related income mobility means that the long-run inequality in health is between 5 and 20% greater than the short-run measure over eight years.

The distinction between the short-run and long-run will be of interest to policy makers whose ethical concern is with inequalities in long-run health. For example, the “fair innings” perspective suggests that equity should be defined in terms of a person’s lifetime experience of health (see e.g., Williams and Cookson, 2000). The evidence provided so far, for the relatively short span of eight years, suggests that the gap between long-run and short-run inequalities may be substantial.

Table 1: Long-run concentration indices (CI^T)and mobility indices (M^T) for the ECHP, 1994-2001

	CI^T		M^T	
Country (waves)	Any limitation (HAMP1)	Severe limitation (HAMP2)	Any limitation (HAMP1)	Severe limitation (HAMP2)
Austria (2-8)	-0.154	-0.197	-0.046	-0.076
Belgium (1-8)	-0.182	-0.325	-0.243	-0.186
Denmark (1-8)	-0.201	-0.373	-0.123	-0.029
Finland (3-8)	-0.108	-0.195	-0.061	-0.097
France (1-8)	-0.160	-0.215	-0.074	-0.064
Germany (1-3)	-0.090	-0.155	-0.054	-0.045
Greece (1-8)	-0.223	-0.286	-0.119	-0.112
Ireland (1-8)	-0.279	-0.409	-0.160	-0.167
Italy (1-8)	-0.108	-0.142	-0.092	-0.024
Luxembourg (1-3)	-0.141	-0.209	-0.024	-0.130
Netherlands (1-8)	-0.114	-0.206	-0.116	-0.125
Portugal (1-8)	-0.192	-0.273	-0.136	-0.170
Spain (1-8)	-0.189	-0.216	-0.143	-0.115
UK (1-3)	-0.184	-0.310	-0.059	-0.064

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