Educational and Gender Differences in the Disability Life Expectancy for the Elderly: Brazil, 1998 and 2003

Gilvan Ramalho Guedes[♦]
Mirela Castro Santos Camargos[♣]
Carla Jorge Machado[♥]

Key-words: Elderly, Educational Level, Disability, Disability-free Life Expectancy

Summary: This study estimated the educational inequalities of the disability life expectancy by age and sex for the Brazilian elderly and introduced a time comparison, generating a decomposition analysis. We used data from the National Household Survey and Brazilian Institute of Geography and Statistics, for 1998 and 2003. To estimate the expected number of years and the proportion of these years lived in absence and presence of disability, we applied the Sullivan method. The results revealed that life expectancy increased for both sexes, more for men and concentrated in the oldest old. Women tended to live longer with and without disability, but a small percentage disability-free than men. The reduction in percentage of DLE_x was bigger for the lower educated in the period. The decomposition revealed that the prevalence effect was responsible for the total reduction in proportions of DLE_x and the mortality effect acting in the opposite direction.

[♦] Economist, Doctoral Student, Department of Demography, CEDEPLAR/UFMG, Minas Gerais, Brazil. Email: grguedes@cedeplar.ufmg.br.

[♣] Physiotherapist, Doctoral Student, Department of Demography, CEDEPLAR/UFMG, Minas Gerais, Brazil. Email: mirela@cedeplar.ufmg.br.

[♥] Professor, Department of Demography, CEDEPLAR/UFMG, Minas Gerais, Brazil. Email: machado@cedeplar.ufmg.br.

Educational and Gender Differences in the Disability Life Expectancy for the Elderly: Brazil, 1998 and 2003

Gilvan Ramalho Guedes
Mirela Castro Santos Camargos
Carla Jorge Machado

INTRODUCTION

Life expectancy in Brazil has consistently increased over the last two decades. In recent years, the increase in life expectancy at later ages has contributed to a change in the population age structure beyond the contribution of fertility decline (Duarte *et al.*, 2002; Finger, 2003). The total increase in life expectancy has led us to question the quality of the added years. The health life expectancy of an is fundamental to policy makers because of the impact of morbidity on public budgets (especially in the public health sector) and on the efficiency and productivity of this group in the labor market (Luft, 1975; Noronha & Andrade, 2005).

Mortality decline has been documented and discussed as a result of better socioeconomic status of the new cohorts, the adoption of healthier behaviors among new cohorts (e.g., decline in tobacco consumption, increase in physical activity), and improvement in public hygiene and nutrition education (Manton *et al.*, 1997; Manton & Stallard, 1996). The lack of high-quality information about both mortality and morbidity, combined with the speed of decline and the relative recentness of the process (when compared to developed countries) results in the lack of existing evidence on trends in mortality and morbidity decline in Brazil (with the exception of Campos, 2003).

Despite the absence of historical trend data, recent cross-sectional health surveys provide a unique opportunity to estimate healthy life expectancy, a more refined health indicator, which combines mortality and morbidity information in a single index. In spite of the existence of these surveys, studies about health and disability-free life-expectancy and their socioeconomic correlates are not numerous because of the recency of robust data about health and mortality (Baptista, 2003; Camargos *et al.*, 2006; Guedes *et al.*, 2006).

The purpose of this paper is to estimate the educational and gender differences in the number of years with presence and absence of functional disability among the elderly in Brazil in two isolated points of time: 1998 and 2003. Our aim is to estimate socioeconomic and gender inequalities using a nationally representative dataset. To our knowledge, the only past study of educational differences in functional disability in Brazil is only representative of São Paulo City in 2000 (Camargos *et al.*, 2006). No past studies have examined these differences over time.

We use education as a proxy for socioeconomic status. Education is a robust SES proxy and shows a strong and significant correlation with health status, especially among the elderly (Manton *et al.*, 1997). Although the link between mortality and morbidity improves our estimates of inequalities in health among the population, stratifying the estimates by education refines the analysis, since a large share of health differences can be explained by educational attainment (Camargos *et al.*, 2006; Manton *et al*., 1997; Sihvonen *et al*., 1998). Education has important externalities for the health of the elderly because of the accumulated effects of the relationship between educational attainment and income, and because of the reciprocal relationship between income and health during adulthood. These life-cycle experiences carry over into the advanced ages, influencing the well-being of the elderly (Baptista, 2003; Manton et al., 1997).

Freedman and Martin (1999) use longitudinal data to estimate the role of education in functional limitations among older Americans. Other studies similarly focus on the relation between education and active life expectancy for different countries, including Asian countries (Ofstedal *et al.*, 2002), Austria (Doblhammer & Kytir, 2001), Belgium (Bossuyt *et al*. 2004), Cambodia (Zimmer, 2005), Finland and Norway (Sihvonen *et al.*, 1998; Valkonen *et al*., 1997) and the United States (Manton *et al*., 1997). Some studies in specific cities have found significant variation in disability-free life expectancy by education, over the entire life cycle (e.g., in Madrid and Barcelona, Martinez_Sánchez *et al.*, 2001) and for the elderly (e.g., in São Paulo, Camargos *et al.*, 2006). These different healthy life expectancy measures, however, are not fully comparable because of differences in the methodology and the definition of morbidity (the only study here documented totally comparable is the one made by Sihvonen *et al*., 1998, for Finland and Norway).

MATERIAL AND METHODS

METHOD

To estimate the number of years in absence and presence of functional disability, we use the Sullivan Method (Jagger, 1999; Sullivan, 1971). The advantages of this method are the simplicity of the procedures to obtain the final estimate and the small amount of data required (prevalence of activities of daily living [ADL] by age, education and sex). The limitations of the Sullivan method, on the other hand, are the sensitivity to changes in the measurement of disability [though, in the absence of considerable variation, this is a reasonable technique (Mathers & Robine, 1997)] and the impossibility of capturing transitions among different states of health. As a typical single-decrement technique, a person is supposed to change from absence to presence of disability, without considering recoveries, estimated to be on the order of 20% or more among the elderly (Katz *et al.*, 1983; Rogers *et al*., 1992; Rogers *et al*., 1989). Moreover, the assumption of the stationary mortality and morbidity functions for the cohorts cannot be realistic in many places. We avoid this assumption by using simulation methods (decomposition) (Robine *et al.*, 1993).

Despite the limitations of the method, our relatively small time window and the absence of longitudinal data require its use. Many studies have been published worldwide using the same method (Camargos *et al.*, 2006; Sihvonen *et al*., 1998), supporting our selection of this method. Research questions involving competing risks are not possible using this method, given its limitations; this is not our case, however.

To estimate the disability life expectancy, we applied prevalence (of disability) rates by quinquennial intervals beginning at 60 years of age; the use of single year prevalence rates is not desirable because of small number problems and the influence of age misreporting (Romero *et al.*, 2005). The disability-free life expectancy ($DFLE_a$) and the disability life expectancy (DLE_a) are estimated according to the following formulae:

$$DFLE_a = \frac{\sum_{a=60}^{80} \left({}_n\pi_a * {}_nL_a \right)}{l_a}; \; DLE_a = \frac{\sum_{a=60}^{80} \left[\left(1 - {}_n\pi_a\right) * {}_nL_a \right]}{l_a}$$

where ${}_n\pi_a$ is the prevalence rate of disability-free among the elderly; ($1-{}_n\pi_a$) is the prevalence rate of functional disability among the elderly; l_a is the number of survivors at exact age a; $\sum({}_n\pi_a * {}_nL_a)$ is the total number of years lived without functional disability between the ages a and a+n by a cohort ; $\sum\left[\left(1-{}_n\pi_a\right)*{}_nL_a\right]$ is the total number of years lived with functional disability in the age group (a, a+n) by a cohort. It is clear in the formulae above that $DFLE_a$ and DLE_a have two independent components: mortality (${}_nL_a$) and morbidity ($1-{}_n\pi_a$). The most direct interpretation of the numerator in the second formula is the number of years lost by a cohort to functional disability between age a and a+n.

DATA

Data on mortality was acquired from Brazilian Institute of Geography and Statistics (IBGE, 2006); and the National Household Survey (PNAD/IBGE, 1998 and 2003). The prevalence of functional disability (measured by a question on activities of daily life [ADL]) by age, sex and educational attainment comes from the National Household Survey (PNAD/IBGE, 2006) for 1998 and 2003. PNAD is a nationally representative survey, with the exception of the rural area of the North of Brazil.[1] In both years, the survey collected a health supplement with information about morbidity and physical mobility, and access, utilization, supply and quality of health care.

The survey is not specifically designed to study the elderly; therefore there are not many questions about disability. The ideal index of disability includes its three dimensions, cognitive, emotional and physical, and uses methods like Grade of Membership (GoM), Factor Analysis or Principal Components Analysis, transforming the multiple dimensions of functional disability into a single continuous scale, as in the approach proposed by the World Health Organization in the *International Classification of Functionality, Disability and Health* (ICF) [WHO, 2001]. Unfortunately, the PNAD contains just seven questions about physical ability, one referring to ADL and the others focusing on physical mobility. The ADL measure is the least susceptible of these to changes over time introduced by perception and physical characteristics of residences and neighborhoods, though it retains the influence of medical advances which may influence the early identification of the disability, raising the prevalence in a specific period (Guedes *et al*., 2006). We therefore use only the ADL indicator of disability. This functional disability indicator has been

[1] Since 2004, PNAD incorporated the rural North.

used by other researchers (Crimmins *et al*., 1994; Parahyba *et al*., 2005; Sihvonen *et al*., 1998; Zimmer, 2005).

The activity of daily life is defined in PNAD as "normally, due to health problems, do you have difficulty eating, bathing or going to the bathroom?" in both years. We selected individuals 60 years and older, following the suggestion of the United Nations for defining elderly in developing countries. In 1998, 33 cases were discarded because of non-reported age and six because of missing information on the ADL question. The final 1998 sample includes 28,937 cases. In 2003, 72 cases were lost because of missing information on age and five because of missing information on the ADL question, resulting in a sample size of 35,037.

VARIABLES

The responses to the ADL question in the questionnaire include four categories ranging from the absence of difficulty to impossibility to execute the task. We transformed these into a dummy variable, with 1 indicating the presence of disability (small difficulty + great difficulty + impossibility to do) and 0 indicating the absence of disability.

Schooling is measured using a dichotomous variable: less than 5 years (0) and 5 years and more (1). Camargos *et al*. (2006) used the same categorical strategy for two main reasons: the right-skewed distribution of the number of years of school completed among the elderly in Brazil and to ensure a small number of final life tables. We created abridged life tables by sex and by quinquennial age groups.

There is a wide range of interconnections between education and health. There is a kind of endogeneity during the first period of the life-cycle, but it is reasonable to consider the level of educational stable from 25 years old. Education creates differences in access to information and the ability to handle the acquired knowledge; furthermore, the level of education can be established in a comparable way for the whole population. The mortality information in Brazil for 1998 and 2003, however, cannot be decomposed by level of education because of the high proportion of missing information (estimated to be 40.1% in 2003, according to the Mortality Information System [DATASUS, 2003]).

RESULTS

THE PICTURE IN 1998

The weighted prevalence of functional disability increases with age for both sexes and educational levels. Women showed a higher prevalence of disability than men, as did the lower educated group compared to the higher one.

Women could expect to live longer than men at all selected ages, with the sex differentials reducing with increasing age. At age 60, the sex gap in life expectancy is about 3.4 years, reducing to 1.2 years at age 80. In all selected ages, women are expected to live longer than men in presence *and* in absence of disability.

The $DFLE_a$ for men decreases with age for both levels of education, with the impact of senescence lower for the higher educated group. For example, the ratio of $DFLE_{60}/DFLE_{80}$ is 1.28 among the poorly educated and 1.14 for the more educated group. At all selected ages, the proportion of remaining life that is lived in the absence of disability is greater than 50% for both levels of education. Comparing the proportion of the $DFLE_a$ of the men at the same age in different educational levels, the gap increased with age, moving from 1.07 to 1.20. The percentage of the DLE_a between the two levels of education, however, remained virtually constant across the ages.

The $DFLE_a$ percentage for women, although reducing with increasing age, was lower than for men in all ages and for both educational categories. The senescence effect reduced with the increase in the educational attainment, but in a smaller ratio than for men. As an example, the ratio $DFLE_{60}/DFLE_{80}$ moved from 1.29 (for lower education) to 1.22 (for higher education). Using the proportion of DLE_{60}/DLE_{80}, the picture is opposite: the impact of decline is higher for women. Educational differences for women at the same age increase both ratios.

The educational effect declines with increasing age for both sexes, especially for women. The sex differences for $\%DLE_a$ decline at advanced ages for the lower educational level, but presented widen again at age 80.

Finally, comparing more educated women to less educated men, women have an advantage, living a larger share of the expected number of years without functional disability in comparison with men at same age. Even in this situation, sex differentials reduce with increasing age.

THE PICTURE IN 2003

As in 1998, the weighted prevalence of functional disability increases with age for both sexes, and is higher for women at all age groups. The prevalence of disability is lower in the higher educational group.

The life expectancy continues to be higher for women at all ages, with the sex gap decreasing by more than 2/3rds from age 60 to 80. For example, at age 60, a woman was expected to live, on average, 22.1 years, against 19.1 years for a man at the same age. At age 80, the life expectancy was 9.6 and 8.8 years, respectively.

The educational effect on the senescence was slightly stronger for men in the proportion of $DFLE_a$, but stronger for women in the percentage of DLE_a. To illustrate, while the DLE_a ratio moved from 0.48 to 0.47 among men, it showed a reduction from 0.52 to 0.46 among women. The educational effect was positive for both sexes; while a woman at age 70 with lower educational attainment was expected to live 27% of her remaining life with disability, another woman at the same age with higher education was expected to spend just 17% of her remaining life living with a functional disability.

The educational effect, however, declines slightly with increasing age, especially for women. At age 60, the ratio %DLE_a(Higher Education)/%DLE_a(Lower Education) was 0.63, moving up to 0.64 at age 80, while for a woman these values were 0.62 and 0.71, respectively. The sex differentials for %DLE_a decline with increasing age for the less educated elderly, although vary for the more educated, showing an increase in the sex gap for the oldest group.

An increase in education for women, as in 1998, puts them in a better position than less educated men. For example, a better educated woman at age 60 was expected to live 3% longer without functional disability than a less educated man at the same age; this sex differential increases with age for both percentages of life expectancy (with and without disability).

TIME COMPARISON

Life expectancy increased for both men and women during the period, with the increase being higher for the oldest old. The gains favored men over women. The percentage of the disability life expectancy increased for men up to age 70, independent of the educational attainment, and then decreased for the oldest old, especially in the lower education group. For women, the picture was not so clear, with small oscillations across age and educational category. The reduction, however, concentrated in the oldest old, like in men.

GENDER COMPARISON OVER THE PERIOD

In spite of the fact that the major increase in life expectancy occurred for men and concentrated in the oldest old for both sexes, the sex ratio of the over time variation remained stable across ages. In other words, it means that the mortality reduction in advanced ages during the period reduced the age gap between life expectancies by sex, but preserved the sex differentials in all ages, equally reducing the sex gap for all age groups in relative terms. Additionally, up to age 70, the percentage of $DFLE_a$ decreased for men, especially among the less educated; this pattern, however, is not clear among women. Nevertheless, comparing both sexes, the percentage of $DFLE_a$ variation in the period among less educated women was 9% of the men, while approximately 5 fold higher for the better educated ones among the oldest old. We can say, therefore, that while men reduced their proportion of expected years without disability up to age 70, especially the less educated, women increased it at age 80 more than men among the more educated and the very opposite among the less ones.

APPROXIMATING THE EDUCATIONAL DIFFERENCES IN MORTALITY

DESCRIBING THE INDIRECT DEMOGRAPHIC TECHNIQUES

In the previous section, we used only one life table for both prevalences stratified by educational level. The importance of education in survivorship is well documented, especially among children (Cambois *et alli*, 2000; Christenson and Johnson, 2001 and

Hummer *et alli*, 1998). As the survivorship at advanced ages is a cumulative experience through life course, differences in education are influential in survivorship at older ages. In Brazil, available mortality data are inappropriate when disaggregated by educational level because of the high percentage of missing information on education (DATASUS, 1998 and 2003). We apply indirect demographic techniques in order to estimate survivorship functions by educational attainment using the PNAD data. We divide the population in study into two groups: 0 to 4 years and 5 years or more of education.

We combine two indirect techniques: the Brass Method for child mortality estimation, using information on children ever born and children surviving, and estimation of survivorship to adulthood from birth, applying maternal orphanhood data. The latter technique produces conditional survivorship estimates. In order to obtain the unconditional adult survivorship probability and complete the mortality function for all age groups we applied the linkage method (United Nations, 1983) that uses an estimate of child mortality, obtained from the first method.

THE BRASS METHOD FOR CHILD MORTALITY

The PNAD database is statistically representative for Brazil. Nevertheless, when data are disaggregated, as required for the Brass Method for child mortality estimation, the proportion of children surviving by age and education of mother is unstable. In order to smooth the distributions across mother's age of the proportion of children dead by sex of child and educational attainment of mother, we used the total population distributions of these proportions by sex of child and by educational level.

Once we have the educational difference by age group, for both sexes, and have the proportions of children dead by sex (for both educational groups), we make the assumption that the sex gap in proportions of children dead remains exactly the same for both groups of educational attainment of mother. This is not implausible, since there is no reason to believe the education of mother makes her selective in maximizing the survivorship of children according to their sex. Keeping the same sex gap in mortality, we apply the distance to the central proportions of children dead to the extremes corresponding to the proportions by educational level in each sex. With this strategy, we have preserved the sex ratio of proportions of children dead and introduced, as a consequence, the educational difference.

The procedure described above was not able to smooth and correct perfectly the instability in age distribution of proportion of children dead, though it made it more reasonable. Because of the persistence of instability, we decided to use the proportion of children dead from mothers between 25 and 29 years old to estimate child mortality. This percentage is used, in conjunction with the Brass multiplier, to generate the q(3) estimate.

In spite of being more flexible and updated, the regression method using Trussell´s coefficients (United Nations, 1983) is not the best choice for our case because they are based in Coale-Demeny standards, which are not the more appropriate models for the Brazilian mortality pattern. For this reason, we applied the original Brass multipliers

(Brass, 1974) that are independent of a mortality pattern, using just parameters of the fertility pattern for interpolation.

THE METHOD OF ORPHANHOOD FOR ADULT MORTALITY AND THE LINKAGE METHOD

The method used to estimate adult unconditional survivorship depends on the assumption that the models used in simulating data represent reality adequately, for it is based on equations fitted by least-squares regressions to simulated data (United Nations, 1983). As explained above, the second strategy of estimation uses, as input, an average estimated value of l(2). Therefore, an additional assumption is required, that the model life tables used to produce the simulated data, from childhood to adulthood, cover the range of actual experience. Because of schedule instability, we averaged l(2), l(3) and l(5) as a proxy for child mortality experience.

To obtain the male adult mortality, we used the estimated female survivorship function and compared with a close implicit female Coale-Demeny model life table. Then, we use the corresponding male Coale-Demeny model life table and apply the same age differences in the estimates between the two female functions (the model and the estimated). Using this strategy, we make the additional assumption that the sex difference in the Brazilian mortality function is identical to the sex difference of survivorship embodied in the model life tables.

A strong argument, valid for Brazilian mortality, is that there is a significant sex differential in mortality, especially due to difference in exposure to risk of death by violent cause and other external reasons. On the other hand, we are not sure if this sex differential, relatively true for the child mortality, can be replicated for adult mortality functions because of the oscillation of sex differences in exposure to risk along the age axis. Therefore, we used the sex differences implied in the model life tables.

Allocation of estimates to specific historical timepoints was done according to Brass and Bamgboye (1981), using a regression method. We assume that both child and adult mortality are changing regularly during the years before the survey (with a linear change on the logit scale). Therefore, the time to which each of the survivorship probabilities refers, denoted by t(n), can be calculated by means of the following equations:

$$u(n) = 0.3333 \ln S(n-5) + Z(M+n-2.5) + 0.0037(27-M)$$

and

$$t(n) = (n-2.5)[1.0 - u(n)]/2.0$$

All the variables above are known, with the exception of functions u(.) and Z(.). According to United Nations (1983), “n-2.5” represents a rough indicator of the mean age of respondents. Finally, Z(.) is a standard function of age whose values in any given case can be obtained by interpolating between values previously tabulated and available in United Nations (1983).

The use of previously tabulated values of Z(.) is sensitive to bias if female adult mortality in the population under analysis has a different pattern from that of the standard population. Fortunately, the impact of deviations from this assumption is unlikely to be large, at least in comparison with the effects on the final estimates of deviations from the assumed linear trend followed by adult mortality changes over time (United Nations, 1983).

THE LINKAGE METHOD

The Brazilian mortality pattern is relatively unique in having a strong mortality difference by sex, especially among young people. Young men more frequently engage in risky behaviors and, in Brazil, the violence and road accidents impact men more than women, especially among the lowest socioeconomic strata, causing a hump in the male mortality function by age between ages 15 and 30. The Coale-Demeny families of model life tables are not the most appropriate standards to be applied to the Brazilian pattern for the lower educated group. The closest approximation of Brazilian mortality pattern is the Chilean pattern, from United Nations (1982). A comparison of Chilean q(a) and q(a) from West Family by Coale-Demeny (1983) reveals the appropriateness of the former pattern to our reality. The proximity of the Chilean mortality pattern can mostly be explained by the very high inequality in the income distribution of the Chilean population (Contreras, 2003).

As we stratify the sample in two groups by the educational level of mother (and respondent, for adult mortality), the higher educated group generated estimates, out of Chilean Model Life Tables range, in the case of women mortality. Even using a second degree polynomial by age across different life expectancy at birth, within a range varying from $61 \le e_0^o \le 75$, covering 15 levels and applying the constraint imposed by the mortality curve (non-negative values), using the logit transformation for each l(a) and, then, generating the observed λ(a)[2], the final estimates of two logit parameters were very far from the original forecasted standard values. For this reason, we tried alternative model life tables and decided on the North model from Coale and Demeny (1983). With this change in mortality standard, we implicitly assume differences in mortality pattern between the two education groups.

To link the estimate of child mortality to the adult conditional adult mortality, we applied the iterative method described in United Nations (1983) in order to obtain an estimate of $l_f(25)$. This adult survivorship probability is used the transform the conditional expectations into unconditional adult survivorship probabilities. The consistency of

---

[2] The following equations represent, analytically, the procedure of polynomial method:

$$\lambda_t(a) = \phi t^2 + \psi t + \tau \quad with \quad t = 1,2,3,\cdots,\infty$$

$$\lambda_t(a) = 0.5\ln\left\{\frac{1-l(a)}{l(a)}\right\}$$

$$l^*(a) = \{1 + \exp[2\alpha + 2\beta\lambda_t(a)]\}^{-1} \quad for \quad t > 15$$

The subscript "t" represents the moment of forecast horizon. The age-specific equations and the correlate R^2 are presented in appendix. The R^2 was 1.00 for all the age groups and this can represent an alternative for other stratification, when necessary.

estimates of adult mortality is sensitive not only to the relationship between child and adult mortality but to the pattern of the mortality schedule between early child ages and 25 as well. Performing the comparison of a logit transformation of some l(a) points from a model life table and our child and adult reasonable points, we derived our alpha and beta estimates and completed the survivorship function for female. As the PNAD questionnaire does not have a question about survivor status of fathers, just female adult mortality can be indirectly estimated. We found the implicit mortality level in the estimated female survivorship and applied it to the corresponding level for males in the model life tables, using linear interpolation.

Decomposition of Change in the Percentage of Disability Life Expectancy

As many assumptions were necessary to generate life tables by educational attainment, we keep the results using the life tables produced by IBGE (2006), not stratified by schooling, and apply our tables only for decomposition.

According to Shorrocks (1999), any decomposition should have two basic properties: additivity and anonymity. Moreover, results must be intuitive (the values must be interpretable). The additive property means that if one sums up the partial decomposed effects, one must obtain the total change in the period. Anonymity implies that the decomposition does not depend on the base or final year of analysis. The total amount of change in disability life expectancy in Brazil between 1998 and 2003 was decomposed into the *prevalence effect* and the *mortality effect*.

The first effect represents how the prevalence rates of disability by age, sex and educational attainment changed among the elderly between 1998 and 2003. The latter represents mortality changes in the period. As reduction in the prevalence of disability increases the healthy life expectancy and a reduction in mortality rates should expose elderly for a longer time to the risk of being disabled, we expect the two effects to work in opposite directions.

The decomposition technique is based on average results of counterfactual simulations. This is to say, we can interpret the results as a counterfactual analysis. For example, among 60 year old lower educated men, the reduction in the proportion in DLE_a was 1.87% in 5 years. If mortality was kept in its 1998 level, the reduction in prevalence would be 2.09%. Mortality reduction, in contrast, contributed to an increase in $\%DLE_a$ of 0.22%. Broadly speaking, the reduction in $\%DLE_a$ was bigger for lower educated elderly, especially for men.

Discussion

The weighted prevalence of disability increases with age and is higher for women in both years. These results are consistent with a higher level of selectivity in mortality among men than women, with men who would have been disabled having died at an earlier age (Alves & Rodrigues, 2005; Case & Pason, 2005; Guedes & Rodrigues, 2006).

Recent literature in Brazil (Berquó, 1996) points out the sex differentials in mortality over the life-cycle. Nonetheless, against the findings on the proportionally higher benefit for women, our study, in accordance to Guedes *et al.* (2006), revealed that sex gap in mortality reduced in recent years (from 1998 to 2003, as showed from IBGE estimates), with the main mortality reductions in the oldest old group.

Despite this decline, women live longer than men and more years in both disability and disability-free states. This has already been documented by Camargos *et al*. (2006) and Baptista (2003) for the elderly in São Paulo, 2000, and by Guedes *et al*. (2006) for Brazilian elderly over the same period analyzed in this paper. However, if we change the focus to the proportion of life expectancy lived in both conditions, the story is different: women live a higher proportion of their expected lifespan in disability. This percentage increases with age but the sex gap reduces (except for age 80 and older), following the recent mortality trend.

The sex gap in morbidity has been explained by three different but complementary arguments: women tend to be more self-conscious about their health and, therefore, declare more morbidity episodes (McDonough & Walters, 2001); another important gender asymmetry is the social construction of health (Berger & Luckmann, 1996; Guedes & Rodrigues, 2006), that embeds the gender roles and modifies the way of dealing with health and illness; excess male mortality during the life-cycle, and a Brazilian peculiarity, an excess of mortality during the young adult ages, creating a selection effect, visible in the elderly as a "stronger" male group, less susceptible to morbidity episodes (Camargos *et al*., 2006).

The social etiology of health is especially relevant for the elderly, because of the differences in image that men and women project in the labor market, the influence of social network support (Auslander & Litwin, 1990; Mor-Barak & Miller, 1991) and in health care utilization. Frailty extends beyond the biological perspective and involves your self-image. Gender differentials in reporting were pointed out by Pinheiro *et al.* (2002), with men tending to report more critical chronic episodes and women concentrating the reporting less severe diseases. Differences in prevalence and severity are another explanation of the gender morbi-mortality apparent paradox and the larger percentage of disability life expectancy for women (Alves & Rodrigues, 2005; Case & Pason, 2005; Guedes & Rodrigues, 2006). This macro-societal dimension of health is directly connected with the behavior of the actors, being modified by the self-interest in health (Verbrugge, 1989, *apud* Pinheiro *et al.*, 2002).

An extra argument that must be emphasized is the distributional asymmetry for educational attainment, especially for older cohorts. Brazilian large scale investment in education is relatively recent. Therefore, low educational position of the older cohorts combined with the strong social forces in the past that molded the gender inequalities in schooling opportunities produced a higher heterogeneity among elderly women. The strong influence of education on income during adulthood and the continuing effect of income on health (and vice versa) over the life cycle (Noronha & Andrade, 2005) can explain the widening of the sex gap in the percentage of disability life expectancy among the oldest old between 1998 and 2003 in Brazil. This occurs not only because of the cohort effect but because of the income outliers concentrated among men.

Educational differentials regarding the disability life expectancy were higher for men, in both years, from age 70. Camargos *et al*. (2006) found an opposite result for the elderly for São Paulo City in 2000. This same opposite finding is documented for other previous studies (Baptista, 2003; Bossuyt *et al.*, 2004; Manton *et al*., 1997; Martinez-Sánchez, *et al*., 2001; Sihvonen *et al*., 1998; Valkonen *et al*., 1997; Zimmer, 2005).

Sex differentials, on the other hand, show no clear pattern over the period. Sex differences in the life expectancy in functional disability were larger for the higher educated group, from age 70, in 1998. An interesting finding is that while the sex gap reduces with age for the poorly educated group, it increases with age in the more educated elderly. This suggests a "boost" in the already heterogeneous female group in higher socioeconomic strata. In 2003, sex differentials were lower for the less educated group (except for the oldest old), the opposite of the result in 1998.

Medical innovation and public assistance diffusion seem to influence the reduction of the percentage of disability life expectancy in Brazil during the 5 year-window studied here. This suggestion is directly connected with the prevalence reduction in the period. However, the changes benefited women more, especially the oldest old. Two non-observable influences might also be present in these results: the barriers in access (Guedes *et al*., 2006, documented that between 1998 and 2003 the barriers in access are stronger for older women than for others, except for poor older men) and the educational asymmetries in mortality. This last influence is not possible to measure because of the lack of reliable education information in the death registers.

This study used a nationally representative survey to estimate life expectancy in presence and absence of disability among the elderly by sex and education, comparing two years, 1998 and 2003. The time comparison revealed a reduction in the prevalence of disability among the elderly, but a small increase in the sex gap in the percentage of the disability life expectancy among the oldest old, even though there was a sex differential compression for the younger old in the period. Many non-observable covariates are likely to be affecting the findings, such as housing amenities, health care, income distribution and hygiene practices. Education had a negative relationship with the percentage of the life expectancy living in disability for both years. As the results suggest, efforts in reducing educational inequalities are important for increasing the well-being of the elderly and reducing the pressure of morbidity on the public health system.

LIMITATIONS

Empirical work on mortality in Brazil must always be viewed carefully because of the lack of information, age misreporting, high percentage of missing information, under registration of deaths. As a consequence, many techniques are necessary to correct or adjust the mortality rates, introducing possible additional bias when the assumptions of the methods are not met.

As discussed previously, the absence of reliable data on mortality by education in Brazil required us to estimate the life table for more and less educated groups using indirect techniques. The stratification of the population by educational attainment led to instability in the functions by age on information about surviving children, and necessitated the application of a smoothing technique. In addition, the same model life table did not fit for both groups, forcing us to use multiple mortality standards. Our lack of the necessary information to estimate male adult mortality implies that the male life table by education does not reflect the real picture. For disability comparison purposes, the ideal was a life table with identical survivorship function for the non-elderly and the differences just for people older than 60 years old. This figure is approximately right for women, with the major gap beginning around 50 years old, but for men this widening gap occurs at about 30 years old. Therefore, the difference in DLE_a at each age is a combination of the educational difference in prevalence and the educational difference in mortality.

Table 1

ADL Prevalences by Sex and Educational Attaiment - Elderly, Brazil, 1998 and 2003

		1998				2003			
Sex	Age group	0 - 4 years of school		5 and more years of school		0 - 4 years of school		5 and more years of school	
		Without	With	Without	With	Without	With	Without	With
Men	60 a 64	90,53	9,47	94,32	5,68	92,91	7,09	96,99	3,01
	65 a 69	88,07	11,93	93,6	6,4	91,72	8,28	93,64	6,36
	70 a 74	86,32	13,68	91,89	8,11	85,86	14,14	93,88	6,12
	75 a 79	80,89	19,11	87,78	12,22	82,91	17,09	85,16	14,84
	80+	66,29	33,71	79,63	20,37	68,64	31,36	79,84	20,16
Women	60 a 64	88,46	11,54	93,79	6,21	91,04	8,96	95,23	4,77
	65 a 69	86,69	13,31	94,08	5,92	88,6	11,4	93,13	6,87
	70 a 74	81,47	18,53	89,88	10,12	84,36	15,64	92,29	7,71
	75 a 79	74,45	25,55	85,39	14,61	77,71	22,29	87,85	12,15
	80+	61,84	38,16	72,01	27,99	62,06	37,94	73,12	26,88

Source: PNAD Database (1998 and 2003)

Table 2

Male and female summarized life tables. Number of years to be lived with disability and disability-free, by sex; percentage of years to be lived with disability and disability-free, by sex and educational attainment, Brazil, 1998 and 2003

						Percentage of life expectancy			
			Life expectancy			$DFLE_x$	DLE_x	$DFLE_x$	DLE_x
Year	Sex	Exact age	Total	$DFLE_x$	DLE_x	Lower education		Higher education	
1998	Males	60	15,9	13,7	2,2	85,1	14,9	91,1	8,9
		70	9,7	7,9	1,8	79,8	20,2	87,6	12,4
		80	5,3	3,6	1,7	66,3	33,7	79,6	20,4
	Females	60	19,2	15,7	3,6	80,1	19,9	88,1	11,9
		70	12,0	9,0	3,1	73,1	26,9	82,8	17,2
		80	6,5	4,1	2,4	61,8	38,2	72,0	28,0
2003	Males	60	19,1	16,4	2,6	84,9	15,1	90,5	9,5
		70	13,1	10,4	2,6	78,4	21,6	86,1	13,9
		80	8,8	6,2	2,6	68,6	31,4	79,8	20,2
	Females	60	22,1	18,0	4,0	80,1	19,9	87,6	12,4
		70	15,0	11,3	3,8	73,0	27,0	82,8	17,2
		80	9,6	6,2	3,5	62,1	37,9	73,1	26,9

Source: PNAD database (1998 and 2003), and IBGE (2006)

Table 3

Differentials in the estimates of the percentages of the disability-free life expectancy and disability life expectancy, Brazil, 1998 and 2003

Educational attainment	Sex	Exact age	Δ e_x	Δ $DFLE_x$	Δ DLE_x	Δ % $DFLE_x$	Δ % DLE_x
Lower		60	3,19	2,69	0,50	-0,15	0,15
	Males	70	3,35	2,49	0,86	-1,39	1,39
		80	3,49	2,52	0,97	2,35	-2,35
		60	2,82	2,27	0,55	0,05	-0,05
	Females	70	3,02	2,19	0,83	-0,09	0,09
		80	3,12	1,95	1,17	0,22	-0,22
Higher		60	3,19	2,78	0,41	-0,67	0,67
	Males	70	3,35	2,73	0,61	-1,53	1,53
		80	3,49	2,80	0,69	0,21	-0,21
		60	2,82	2,38	0,45	-0,50	0,50
	Females	70	3,02	2,51	0,51	0,04	-0,04
		80	3,12	2,35	0,77	1,11	-1,11

Source: PNAD database (1998 and 2003), and IBGE (2006)

Table 4

Male and female summarized life tables. Number of years to be lived with disability and disability-free and percentage of years to be lived with disability and disability-free, by sex and educational attainment, Brazil, 1998 and 2003 - Original Values with estimated life tables by educational level

Educational attainment	Sex	Exact age	e(x)		(nπx)		$DFLE_x$		DLE_x		% $DFLE_x$		% DLE_x	
			1998	2003	1998	2003	1998	2003	1998	2003	1998	2003	1998	2003
Lower		60	16,07	17,52	0,91	0,93	13,61	15,02	2,46	2,49	84,68	85,76	15,32	14,24
	Males	70	10,32	11,74	0,86	0,86	8,16	9,28	2,15	2,46	79,14	79,08	20,86	20,92
		80	6,04	7,92	0,66	0,69	4,01	5,44	2,04	2,48	66,29	68,64	33,71	31,36
		60	17,70	19,81	0,88	0,91	14,27	16,11	3,42	3,70	80,65	81,33	19,35	18,67
	Females	70	11,33	13,24	0,81	0,84	8,32	9,80	3,00	3,44	73,49	74,01	26,51	25,99
		80	6,15	8,63	0,62	0,62	3,80	5,36	2,35	3,28	61,84	62,06	38,16	37,94
Higher		60	17,68	18,85	0,94	0,97	15,98	17,10	1,71	1,75	90,35	90,73	9,65	9,27
	Males	70	11,55	12,56	0,92	0,94	10,01	10,85	1,54	1,71	86,67	86,41	13,33	13,59
		80	7,11	8,19	0,80	0,80	5,66	6,54	1,45	1,65	79,63	79,84	20,37	20,16
		60	21,65	23,12	0,94	0,95	18,76	20,18	2,89	2,94	86,66	87,28	13,34	12,72
	Females	70	14,70	15,93	0,90	0,92	11,93	13,14	2,77	2,78	81,17	82,52	18,83	17,48
		80	9,40	10,42	0,72	0,73	6,77	7,62	2,63	2,80	72,01	73,12	27,99	26,88

Source: PNAD database (1998 and 2003); DATASUS (2006); IBGE/DPE (Population Department ans Social Indicators) and Ensp/Fiocruz/Fensptec (Disease Burden Project) and Preston et alli (2000)

Table 5

Decomposition of the Diference in the Percentage of Disability Life Expectancy by age, sex and educational level - Brazil, 1998 and 2003

Educational attainment	Sex	Exact age	Total Effect	Prevalence Effect	Mortality Effect
Lower	Males	60	-1,87	-2,09	0,22
		70	-0,89	-1,07	0,19
		80	-2,35	-2,35	0,00
	Females	60	-1,68	-2,26	0,58
		70	-1,64	-2,24	0,60
		80	-0,22	-0,22	0,00
Higher	Males	60	-0,76	-0,79	0,03
		70	-0,08	-0,11	0,03
		80	-0,21	-0,21	0,00
	Females	60	-0,55	-1,20	0,65
		70	-1,29	-1,96	0,67
		80	-1,11	-1,11	0,00

Source: PNAD database (1998 and 2003); DATASUS (2006); IBGE/DPE (Population Department ans Social Indicators) and
Ensp/Fiocruz/Fensptec (Disease Burden Project) and Preston et alli (2000)

REFERENCES

ALVES, L. C.; RODRIGUES, R. N. Determinantes da autopercepção de saúde entre idosos do Município de São Paulo, Brasil. **Revista Panamericana de Salud Publica**, v.17, no.5/6, pp.333-341, 2005.

AUSLANDER, G. K.; LITWIN, H. Social support networks and formal help seeking: differences between applicants to social services and a nonapplicant sample. **Journal of Gerentology: Social Sciences**, 1990, 45: 112-119.

BAPTISTA, D. B. D. A. **Idosos no município de São Paulo: expectativa de vida ativa e perfis multidimensionais de incapacidade a partir da SABE**. Masters Thesis, Universidade Federal de Minas Gerais, Belo Horizonte, MG, Brazil, 2003.

BERGER, P. L.; LUCKMANN, T. **A construção social da realidade : tratado de sociologia do conhecimento.** 13. Ed. Petrópolis: Vozes, 1996.

BERQUÓ, E. Algumas considerações demográficas sobre o envelhecimento da população no Brasil. *In* **Seminário Internacional sobre Envelhecimento Populacional: uma Agenda para o Fim do Século**. Ministério da Previdência Social, Brasília, 1996.

BOSSUYT, N.; GADEYNE, S.; DEBOOSERE, P.; VAN OYEN; H. Socio-economic inequalities in health expectancy in Belgium. **Public Health**, no.118, p.3-10, 2004.

BRASS, W.; BAMGBOYE, E. A. The time location of reports of survivorship: estimates for maternal and paternal orphanhood and the ever-widowed. Working Paper No. 81-1, **London School of Hygiene and Tropical Medicine**, Centre for Population Studies, 1981.

CAMARGOS, M. C. S.; MACHADO, C. J.; RODRIGUES, R. N. Disability life expectancy for the elderly, city of São Paulo, Brazil, 2000: gender and educational differences. **Journal of Biosocial Science**, pp.1-9, 2006.

CAMPOS, N. O. B. **O Ritmo de Declínio da Mortalidade dos Idosos nos Estados do Sudeste, 1980/2000**. Masters Thesis, Federal University of Minas Gerais, Belo Horizonte, MG, Brazil, 2003.

CASE, A.; PAXSON, C. Sex Differentials in Morbidity and Mortality. **Demography**, v.42, no.2, pp.189-214, May/2005.

CONTRERAS, D. Poverty and Inequality in a Rapid Growth Economy: Chile 1990-96. **The Journal of Development Studies**, v.39, no.3, pp.181-200, Feb/2003.

DATASUS. **Dados de declaração de óbito do Sistema de Informações de Mortalidade (SIM)** Brasília: Ministério da Saúde. (Available in CDs-ROM, 1998 and 2003)

DOBLHAMMER, G.; KYTIR, J. Compreesion or expansion of morbidity? Trends in healthy-life expectancy in the elderly Austrian population between 1978 and 1998. **Social Science and Medicine**, no.52, p.385-391, 2001.

DUARTE, E. C.; SCHNEIDER, M. C.; PAES-SOUSA, R.; SILVA, J. B.; CASTILLO-SALGADO, C. Life expectancy at birth and mortality in Brazil, 1999: exploratory analysis of regional differences. **Revista Panamericana Salud Publica**, no.12, pp.436-444, 2002.

FINGER, C. Fears grow for Brazil' youth. Increased life-expectancy estimates mask deep-rooted problems with health care. **Lancet**, v.362, pp.2078.

FOGEL, R. Economic growth, population theory, and physiology: The bearing of long-term processes on the making of economic policy. **American Economic Review**, no.84, pp.369-395, 1994.

FREEDMAN, V.A.; MARTIN, L. The role of education in explaining and forecasting trends functional limitations among older Americans. **Demography**, v.36, n.4, p.461–473, Nov./1999.

GUEDES, G. R.; CAMARGOS, M. C. S.; MACHADO, C. J. **Disability life expectancy for the elderly, Brazil, 1998 and 2003: income and gender inequalities**. (mimeo), 2006.

GUEDES, G. R.; RODRIGUES, R. N. **Determinantes da Autopercepção de Saúde e da Restrição de Atividades Rotineiras do Idoso Não-institucionalizado no Brasil**. (mimeo), 2006.

HUMMER, R. A.; ROGERS, R. G.; EBERSTEIN, I. W. Sociodemographic differentials in adult mortality: a review of analytic approaches**. Population and Development Review**, v.24, n.3, p.553-578, Sept./1998.

INSTITUTO BRASILEIRO DE GEOGRAFIA E ESTATÍSTICA. **Pesquisa Nacional por Amostra de Domicílios**, Rio de Janeiro: 1998.

INSTITUTO BRASILEIRO DE GEOGRAFIA E ESTATÍSTICA. **Pesquisa Nacional por Amostra de Domicílios**, Rio de Janeiro: 2003.

INSTITUTO BRASILEIRO DE GEOGRAFIA E ESTATÍSTICA. **Tábuas Completas de Mortalidade – 1998**. Available in: http://www.ibge.gov.br/home/estatistica/populaçao/tabuadevida/default_1998.shtm. Accessed in: Jun/05/2006.

INSTITUTO BRASILEIRO DE GEOGRAFIA E ESTATÍSTICA. **Tábuas Completas de Mortalidade – 2003**. Available in: http://www.ibge.gov.br/home/estatistica/populaçao/tabuadevida/default_2003.shtm. Accessed in: Jun/05/2006.

JAGGER, C. **Health expectancy calculation by the sullivan method: a practical guide**. Nihon University Population Research Institute, Tokyo, 1999 (NUPRI Research Paper, 68).

KATZ, S.; BRANCH, L.G.; BRANSON, M.H.; PASIDERO, J.A.; BECK, J.C.; GREER, D.S. Active life expectancy. **New England Journal of Medicine**, v.309; no.20, pp.1218-1224, 1983.

LUFT, H. S. The impact of poor health on earnings. **The Review of Economics and Statistics**, v.57, no.1, pp.43-57, 1975.

MANTON, K. G.; STALLARD, E. S.; CORDER, L. Education-specific estimates of life expectancy and age-specific disability in the U.S. elderly population: 1982 to 1991. **Journal of Aging and Health**, v.9, no.4, pp.419-450, nov/1997.

MARTINEZ-SÁNCHEZ, E.; GUTIÉRREZ-FISAC, J. L.; GISPERT, R.; REGIDOR, E. Educational differences in health expectancy in Madrid and Barcelona. **Health Policy**, no.55, pp.227-231, 2001.

MATHERS, C. D., ROBINE, J. M. How good is Sullivan's method for monitoring changes in population health expectancies? **Journal of Epidemiology and Community Health**, v.51, no.1, pp.80-86, Feb/1997.

MOR-BARAK, M.; MILLER, L. S. A longitudinal study of the causal relationship between social networks and health of the poor frail elderly. **Journal of Applied Gerentology**, 1991, 10: 293-310.

NORONHA, K. V. M. S.; ANDRADE, M. V. A importância da saúde como um dos determinantes da distribuição de rendimentos e pobreza no Brasil. **Texto para discussão**, CEDEPLAR/UFMG, 2005).

OFSTEDAL, M. B.; ZIMMER, Z.; CRUZ, C.; CHAN, A.; LIN, Y. Self-assessed Health Expectancy Among Older Asians: A Comparison of Sullivan and Multistate Life Table Methods. Report No. 03-60, Research Reports, **Population Studies Center**, Institute for Social Research, University of Michigan, 2002.

PINHEIRO, R. S.; VIACAVA, F.; TRAVASSOS, C.; BRITO, A. S. Gênero, morbidade, acesso e utilização de serviços de saúde no Brasil. **Ciência & Saúde Coletiva**, v.7, no.4, pp.687-707, 2002.

PRESTON, S. H.; HEUVELINE, P.; GUILLOT, M. **Demography: measuring and modeling population processes.** Malden, MA: Blackwell Publishers, 2000. 291p.

ROBINE, J.M.; MATHERS, C.; BROUARD, N. Trends and differentials in disability-free life expectancy. In: CONFERENCE ON HEALTH AND MORTALITY TRENDS AMONG ELDERLY POPULATIONS: DETERMINANTS AND IMPLICATIONS. Sendai, Japan. **Papers**. Liege: IUSSP, 1993.v.1, section 2.

ROGERS, R. G.; ROGERS, A.; BELANGER, A. Active life among the elderly in the United States: multistate life-table estimates and population projections. **The Milbank Quarterly**, v.67, no.3-4, pp.370-411, 1989.

ROGERS, R. G.; ROGERS, A.; BELANGER, A. Disability-free life among the elderly in the U.S.: sociodemographic correlates of functional health. **Journal of Aging and Health**, v.4, no.1, pp.19-42, 1992.

ROMERO, D. E.; LEITE, I. C.; SZWARCWALD, C. L. Healthy life expectancy in Brazil: applying the Sullivan method. **Cadernos de Saúde Pública**, v.21, sup.:S7-S18, 2005.

SHORROCKS, A. F. Decomposition Procedures for Distributional Analysis: A Unified Framework Based on the Shapley Value. Mimeo. Uk: University of Essex, 1999.

SIHVONEN, A.; KUNST, A. E.; LAHELMA, E.; VALKONEN, T.; MACKENBACH, J. P. Socioeconomic inequalities in health expectancy in Finland and Norway in the late 1980s. **Social Science and Medicine**, v.47, no.3, p.303-315, 1998.

STERN, Y.; GURLAND, B.; TATEMICHI, T. K., TANG, M. X., WILDER, D., MAYEUX, R. Influence of education and occupation on the incidence of Alzheimer's disease. **Journal of the American Medical Association**, v.271, pp.1004-1010, 1994.

SULLIVAN, D. F**.** A single index of mortality and morbidity**. HSMHA Health Reports**, 86: 347-54, 1971.

UNITED NATIONS. **Indirect techniques for demographic estimation**. New York: United Nations, 1983. 304p.

UPPAL, S. Impact of the timing, type and severity of disability on the subjective well-being of individuals with disabilities. **Social Science & Medicine**, v.63, pp.525–539, 2006.

VALKONEN, T.; SIHVONEN, A; LAHELMA, E. Health expectancy by level of education in Finland. **Social Science and Medicine**, v.44, no.6, p.801-808, 1997.

WORLD HEALTH ORGANIZATION. **International classification of functioning, disability and health: ICF**. Geneva: World Health Organization; 2001.

ZIMMER, Z. Active Life Expectancy and Functional Limitations Among Older Cambodians: Results from a 2004 Survey. Working Paper, no. 201, **Population Council**, New York, 2005.

APPENDIX

Table A1

Polynomial Equations of Second Moment for Projection of Female Chilean Pattern - United Nations Model Life Tables for Developing Countries [Life Expectancy at Birth - Range: 61 to 75 years old]

$\lambda_t[l(a)]$	Polynomial Regression Equation	R^2
1	-0,0007t2 - 0,0220t - 1,1141	0,9999
2	-0,0007t2 - 0,0243t - 1,0487	0,9999
3	-0,0007t2 - 0,0254t - 1,0194	1,0000
5	-0,0007t2 - 0,0264t - 0,9904	1,0000
10	-0,0007t2 - 0,0273t - 0,9633	1,0000
15	-0,0007t2 - 0,0279t - 0,9428	1,0000
20	-0,0007t2 - 0,0290t - 0,9103	1,0000
25	-0,0007t2 - 0,0304t - 0,8665	1,0000
30	-0,0007t2 - 0,0316t - 0,8166	1,0000
35	-0,0007t2 - 0,0327t - 0,7604	1,0000
40	-0,0007t2 - 0,0333t - 0,6973	1,0000
45	-0,0007t2 - 0,0334t - 0,6255	1,0000
50	-0,0007t2 - 0,0330t - 0,5418	1,0000
55	-0,0007t2 - 0,0325t - 0,4427	1,0000
60	-0,0006t2 - 0,0319t - 0,3188	1,0000
65	-0,0005t2 - 0,0312t - 0,1650	1,0000
70	-0,0004t2 - 0,0307t + 0,0341	1,0000
75	-0,0004t2 - 0,0309t + 0,2738	1,0000
80	-0,0003t2 - 0,0325t + 0,5667	1,0000
85	-0,0003t2 - 0,0352t + 0,9346	1,0000

Source: Prepared from United Nations (1982)