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Equity in health care finance in Palestine: The triple effects revealed

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ABSTRACT

This paper presents an application of the Urban and Lambert "upgraded-AJL Decomposition" approach that was designed to deal with the problem of close-income equals in equity analysis, and as applied to the area of health care finance. Contrary to most previous studies, vertical and horizontal inequities and the triple effects of inter-groups, intra-group and entire-group reranking of various financing schemes are estimated, with statistical significance calculated using the bootstrap method. Application is made on the three financing schemes present in the case of the Occupied Palestinian Territory. Results demonstrate the relative importance of the three forms of reranking in determining overall inequality. The paper offers policy recommendations to limit the existing inequalities in the system and to enhance the capacity of the governmental insurance scheme.

1. Introduction

Assessing the impact of health care finance on income inequality is a relatively new area of analysis in the context of both developing and developed countries (Wagstaff, 2002). Worldwide empirical evidence has already shown that different health care financing schemes may very differently affect the prevailing income distribution of a country, and consequently, the associated level of overall income inequality (van Doorslaer et al., 1999; Wagstaff, 2002). The distributional impact of health care financing can be inferred from measures of vertical inequity; e.g., progressivity analysis. However, as stated by Wagstaff and van Doorslaer (1997): "*depending on the extent of horizontal inequity and reranking involved in health care finance, a progressivity analysis can give a misleading impression about the income redistribution associated with the financing system*". These three dimensions of inequity – relating respectively to violations of the normative principles of "unequal treatment of unequals", "equal treatment of equals" and "proper treatment of unequals" – are generally addressed by decomposing the redistributive effect (*RE*) induced by financing into a vertical, horizontal and reranking effect (Aronson and Lambert, 1994). Applied to health care finance, the vertical effect (*VE*) measures how and to what extent individuals of unequal ability-to-pay are affected by the financing, the horizontal effect (*HE*) captures the extent to which individuals of equal ability-to-pay make unequal contributions to health care, whereas the reranking (*RR*) quantifies changes in ranking-order of individuals (by income) following health payments (O'Donnell et al., 2007).

Empirical studies conducted in the context of developed countries (Wagstaff and van Doorslaer, 1997; van Doorslaer et al., 1999) have demonstrated that different forms of health care financing may indeed be associated with both horizontal and reranking effects. This is even more likely in the context of developing countries, where income protection mechanisms are still far under-

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developed (Sekhri and Savedoff, 2005; Pauly et al., 2006), and where high proportions of health care expenditures are funded by households' direct out-of-pocket payments (Musgrove et al., 2002). Since illness is a stochastic event, the extent of discrepancies in actual payments born by individuals belonging to a similar income group, as well as the extent of changes in income status of individuals due to "catastrophic" health care payments, is likely to be exacerbated in these countries (Wagstaff and van Doorslaer, 2003; Xu et al., 2003; Ichoku, 2005).

A simultaneous assessment of the above three different measures of inequity may therefore be of particular interest to fully reveal the overall *RE* of health care financing in the context of developing countries. Such assessment can indeed help inform the controversial policy debates about the extent to which reforms aimed at increasing the efficiency of health care systems do not simultaneously increase inequities in health care financing. This may in turn help reduce the possible adverse effects of health care financing on the prevailing income inequalities in a country (Kidson, 1999; McPake and Mills, 2000; James et al., 2006).

The literature on public finance offers various methods to quantitatively measure *VE*, *HE* and *RR* (e.g., Kakwani, 1984; Aronson et al., 1994; Ven et al., 2001; Duclos et al., 2003). The approach that has been previously applied in the specific domain of health care financing (Wagstaff and Doorslaer, 1997; van Doorslaer et al., 1999) is the one initially proposed by Aronson et al. (1994)—hereafter the AJL approach. Theoretically, the AJL approach allows to decompose the total *RE* of a financing scheme into *VE*, *HE* and *RR* for a population that is composed of groups of true- or exact-income equals, i.e., a situation where the study sample consists of groups of individuals having exactly the same pre-payment income, and for a distribution where the average post-payment income of each group increases with the respective pre-payment income level, i.e., a payment schedule that is supposed not to produce changes in the groups' ranking-order. However, due to the absence of exact-income equals in real data surveys, empirical implementations of the AJL approach have relied on the principle of close-income equals, i.e., by dividing the study sample into artificial groups of income based on certain definitions of income bandwidths. It has been shown that such practice can lead to misleading results: biases may arise not only due to the arbitrary specification of close-income equals, but also due, in large part, to the possibility of both intra-groups reranking – i.e. the extent to which the payment schedule induces changes in ranking-order of individuals within the specified groups of close-income equals (R_{WG}) – and entire-groups reranking, i.e., the extent to which the payment schedule induces changes in ranking-order of the whole groups of close-income equals (R_{EG}) (Ven et al., 2001).

The need to consider the potential impact of R_{WG} and R_{EG}, as well as the sensitivity of the empirical estimations of *VE* and *HE* to the choice of income bandwidth for close-equals, has been advocated for the assessment of *RE* of tax and transfer systems (Ven et al., 2001; Urban and Lambert, 2008). These aspects may also be relevant with regard to the assessment of *RE* of different health care financing schemes. A methodological extension to the earlier work in tax literature has latterly been provided by Urban and Lambert (2008)—hereafter the UL approach. In contrast to the classical AJL approach and its previous applications, the UL approach reset the measurement system of *VE*, *HE* and *RR* using a conceptual model that is specifically designed to accommodate close-income equals setting. The UL approach presents two complementary advantages: it is able to capture all possible reranking effects, and it provides a more convenient identification of vertical and horizontal inequities by smoothing the actual effect of payments within each close-income equals group. In such approach, the *VE* is measured by allocating to each individual the average payment paid by the respective group of close-income equals, while *HE* is estimated directly based on person-by-person comparisons of actual and counterfactual; i.e., smoothed, post-payment incomes within close-equals groups. Lastly, although there is no consensus in the empirical literature on an optimal procedure to identify the income bandwidth of close-equals (Ven et al., 2001; Duclos et al., 2003), the UL approach, while computationally involves direct estimates of *VE*, *HE*, and *RR* as sample statistics, advocates an assessment of the relative importance of these effects given different choices of income bandwidth. This may, indeed, facilitate an appropriate specification of close-income groups for policy purpose. The UL approach has been recently applied to investigate the *RE* of taxation in the USA (Kim and Lambert, 2009). However, to our knowledge, there has been no previous attempt to apply such methodological improvement in the specific area of health care financing.

Another limitation of previous work on inequality measurement in health care finance, which may also fuel unnecessary misinterpretations, is related to the fact that most of the previous studies have rarely assessed the statistical significance of inequality measures. The very few studies (Klavus, 2001; Cissé et al., 2007) that have attempted to do so have used the classical asymptotic method, which has its own limitations (Mills and Zandvakili, 1997). However, statistical inference based on bootstrap methods were shown to lead to more subtle treatment for statistical problems associated with the measures of inequality (Andres and Calonge, 2005; Moran, 2006). The bootstrap method takes into account the specific bounds of inequality measures while no underlying function of distribution is being imposed. Besides comparing standard errors and probability intervals, an obvious advantage of the bootstrap method is that it allows for testing the relationship between two interdependent curves according to the dominance criterion, and therefore, reduces the risk of biased interpretations due to sample structure (Davidson and Flachaire, 2007; Abu-Zaineh et al., 2008).

The purpose of this paper is to apply the above methodological advances initially developed for inequality measurement of taxation to the specific domain of health care financing, and to illustrate how these developments can significantly help clarifying debates about health care policies in the context of developing countries, using the particular case of the Occupied Palestinian Territory (OPT). The financing structure of health care in the OPT is expected to be associated with a major risk of exacerbation of inequalities: the country lacks a universal system of health care financing and a substantial share of health care expenditures is funded through households' direct out-of-pocket payments PCBS, 2004. The main public financing scheme is the one known as the Governmental Health Insurance (GHI). Initially, this was on a compulsory basis for public sector employees. However, the scheme has been opened up to others – in the private and informal sectors – on a voluntary basis. By 2002, over 60% of the Palestinian households were covered by public scheme, a little less than half of these being covered on a voluntary basis (MoH-PHIC, 2006). On the other hand, the ongoing political crisis in the region, which has brutally increased the proportion of the population living under the poverty line (PCBS, 2006b), tends to aggravate existing inequities in access to health care (Mataria et al., 2006). Consequently, the Ministry of Health has started to provide an almost free of charge coverage to the mostly affected classes of the population. This insurance scheme was lately known as "Al-Aqsa Intifada Insurance". However, the indigent performance of the local economy, the freeze of financial support and tax transfers to the Palestinian Authority (PA) threaten to negatively affect the current health care delivery system. Furthermore, due to the lack of effective and sufficient redistributive polices (e.g., proper system of tax transfer), which may help reduce inequality and ensure more equitable distribution of income, the redistributive effect of the current health care financing structure are expected to aggravate the global existing inequality in the prevalent income distribution in the country.

The empirical analysis of this paper is based on data from a recent household health expenditure survey, originally designed for initiating a system of National Health Accounts for the OPT. The survey, which was carried out in 2004 by the Palestinian Central Bureau of Statistics PCBS, 2004, covered a national representative sample of Palestinian households residing in the West Bank (WB) and Gaza Strip (GS). Collected data provide detailed information on households' incomes and expenditures, individuals' health care seeking behaviours, health care expenditures, and government and private insurance premiums. Consequently, the survey offers a unique opportunity to assess some equity features of health care financing, and to discuss the extent to which recent developments in inequality measurement, and its statistical inference, may add to our understanding of equity issues in the Palestinian context. Some of the methodological developments that we try to transfer to the field of equity measurement in the case of health care financing in the OPT may also be worthwhile for other contexts in developing countries.

The remainder of the paper is organized as follows. Section 2 discuses some methodological issues in measuring and decomposing the total *RE* into *VE*, *HE*, and *RR*, using the measurement approach proposed by Urban and Lambert (2008); this is followed by describing the estimation procedures and statistical inferences for inequality measures. Section 3 describes our data and variables definitions used for the empirical analysis. Results are reported in Section 4. The last two sections contain our discussion and conclusions.

2. Methodology

2.1. Measurement model

Total *RE* of health care payments can be defined as the (dis)equalizing effect associated with a move from pre- to post-payment income distributions (Reynolds and Smolensky, 1977). Thus, if G_x and G_y are pre- and post-payment Gini coefficients, respectively, then *RE* can be assessed as:

$$RE = G_x - G_y \tag{1}$$

A positive (negative) value of *RE* indicates that health payments tend to reduce (exacerbate) income inequality; and thus, the payment scheme is qualified as "pro-poor" ("pro-rich"). Segregating *RE* into *VE*, *HE*, and *RR* – as suggested by Urban and Lambert (2008) – necessitates the estimation of a set of concentration coefficients for post-payment income distribution, C_y, each corresponds to a particular ordering of income units as per groups of close-income equals, Ω_k^w (k taking the values form 1 to K, with K being the number of groups as determined by the income bandwidth, w, and the maximum income in the sample). Consequently, *RE* can be decomposed as follows:

$$RE = VE - HE - RR \tag{2}$$

The *VE* component, which measures the counterfactual inequality change that would arise should horizontal equity in payments prevail, can then be assessed as follows:

$$VE = G_x C_y^1 \tag{3}$$

where C_y^1 is the post-payment income concentration coefficient that would be obtained if each $x \in \Omega_k^w$ is reduced by $\mu_{g,k}$—where $\mu_{g,k}$ is the mean payment of health care of the group k.

HE, which measures the *pure* horizontal inequity, is assessed by comparing person-by-person departures of the actual post-payment incomes from those generated by a reference schedule constructed, counterfactually, to be horizontal inequality-free within groups of close-equals,

$$HE = C_y^2 - C_y^1 \tag{4}$$

where C_y^2 is the post-payment income concentration coefficient that is obtained by ranking households by their *pre-payment incomes*. Finally, *RR*, which measures the reranking that occurs in the transition from pre-payment to post-payment income period, is conceptualized to incorporate three forms of rank reversals: R_{WG}, R_{EG}, and R_{BG}. These are related as follows:

$$\begin{aligned} RR &= R_{WG} + R_{BG} + R_{EG} = [C_y^3 - C_y^2] + [C_y^4 - C_y^3] \\ &\quad + [G_y - C_y^4] = G_y - C_y^2 \end{aligned} \tag{5}$$

where C_y^3 is obtained by ranking households within groups by post-payment incomes, and the groups by their pre-payment group means, $\mu_{x,k}$; whereas C_y^4 is obtained by ranking households within groups by post-payment incomes and the groups by post-payment group means, $\mu_{y,k}$.

It is important to note that each form of reranking corresponds to a different change in the order of income distribution and indicates different possible causes. The presence of "differential treatment of equals" *per se* can induce reranking (Wagstaff and Doorslaer, 1997). For instance, risk-rating premiums can cause both horizontal inequality and reranking. Such form of reranking can be detected by the term R_{WG}, which measures rank reversals within groups. Reranking can also arise in the absence of horizontal inequality. For instance, a flat-rate premium induces no horizontal inequality, though such (highly regressive) payment schedule can induce changes in ranking-order of the whole groups, hence the term R_{EG}. Lastly, reranking can occur when marginal payment rates exceed unity in some parts of income distribution. For instance, this can be due to catastrophic health payments, which lead to reranking between-groups (R_{BG}). This form of reranking shows how households with different incomes are affected by health financing.

2.2. Statistical inference of inequality measures

Statistical significance of observed variations in the computed values of each of the above measures was tested using the bootstrap method (BTS). BTS is a re-sampling method used to simulate the empirical distribution of a statistic (Effron and Tibshirani, 1993). The same method is used to evaluate the null hypothesis of no change between the two periods (e.g., pre-payment and post-payment periods). The applications of BTS for inequality measures are described and documented in Duclos and Araar (2006), and Andres and Calonge (2005). However, to illustrate the use of the BTS methods for the measures in the present paper, we briefly describe the implemented procedures for the case of Gini coefficients denoted as G.

Consider a statistic $\hat{G}$ based on a sample size N; hence, instead of assuming the shape of the distribution of $\hat{G}$ statistic, the entire sampling distribution of $\hat{G}$ is approximated through investigating its variation over a large number of pseudo-samples obtained by randomly selecting, with replacement, a large number (R) of sub-samples of size n ($n < N$), out of the same dataset—the BTS re-samples. The statistic $\hat{G}$ is then computed for each BTS re-sample, yielding $\hat{G}^*$—the so-called BTS replication of the statistic $\hat{G}$. The sampling variation of $\hat{G}$ can be estimated by applying the expression of standard errors to the R-length vector of BTS replications. Regarding the estimation of the probability confidence intervals $[\hat{G}_I, \hat{G}_S]$ at a confidence level of $(1-\alpha)$%, BTS provides us with different possible methods to construct tail probabilities for the statistic $\hat{G}$; e.g., the percentile method (Mills and Zandvakili, 1997). The latter procedure involves an estimation of an empirical function $f(_{\hat{R}})$ of the statistic of interest (G) from R of BTS re-samples. The empirical percentiles ($\alpha/2$) and ($1-\alpha/2$) can, then, be computed for a significance level (α).

Using the underlying relationship between confidence intervals and hypothesis testing, tail probability values for hypothesis testing can be computed from the BTS distribution. Since we are interested in comparing different values of an inequality measures such as the difference between pre-payment and post-payment Gini coefficients, we adopt the following tests: let V_1 and V_2 be vectors of BTS values of measure (G) in the periods 1 and 2, for instance, pre-payment income and post-payment income measured over the same households in a particular year, then $D=V_2-V_1$. The test rejects the null hypothesis that the coefficients for the two periods are the same; i.e., the equality between coefficients, if the estimated confidence interval for the difference of the Gini coefficients does not include zero. This can be formulated as follows: H_0: $G_1=G_2$ against H_A: $G_1 \neq G_2$ which can be re-defined as H_0: $D=0$ against H_A: $D \neq 0$. Previous authors have suggested that R should be at least 200 in order to obtain estimates significant at the 1% level (Duclos and Araar, 2006). In this paper we simulate 200 re-samples with replacements to estimate the BTS standard errors and the tail probability intervals at $\alpha=5\%$. Of course, one can allow the sample size of BTS to vary and then to examine the effects of such variation on our measures of dispersion. However, this exercise is beyond the objective of this paper.

2.3. Estimation procedures

The computation of inequality measures – defined above – can be conducted using either the convenient covariance methods (Jenkins, 1988) or the integration methods available in software programs such as GAUSS, Stata and DAD (van Doorslaer et al., 1999; Duclos and Araar, 2006). However, in this paper, we have chosen to use purpose-built procedures in MATLAB software program. In case where household is the unit of observation, one has to adjust for differences in the demographic composition of the household in order to generate individual-level estimates. Among the several ways to do so (Lambert, 2003; Muellbauer and Ven, 2003), we used the WHO/FAO equivalence scale, suggested for the case of developing countries (Cissé et al., 2007). This assigns a value of 1 for an adult man, 0.8 for an adult woman, and 0.5 for children younger than 15 years old (Aho et al., 1997). Household level of resources is then divided by the equivalence factor, z_i, to generate an adjusted income per equivalent-adult.

The different inequality coefficients are estimated using the following formula (Lerman and Yitzhaki, 1989; Kim and Lambert, 2009):

$$G=\frac{2\mathrm{cov}\left(W_h\left(\sum_{i=1}^{h}w_iz_i/N\right)\right)}{\mu}=\frac{(2/N)\sum_{h=1}^{n}w_hz_h(W_h-\mu)\left(\left(\sum_{i=1}^{h}w_iz_i/N\right)-(1/N^2)\sum_{h=1}^{n}\sum_{i=1}^{h}w_iz_i\right)}{\mu} \quad (6)$$

where W_h is household h income per equivalent adult (h taking the values form 1 to n), n is the total number of sample observations, N is the total number of equivalent adults, $N=\sum_{h=1}^{n}w_hz_h$, and μ is the mean income overall.

The computations of RE and RR measures are independent of the choice of income bandwidth (w). As for VE, HE, and the various components of reranking (R_{BG}, R_{WG}, R_{EG}), an income bandwidth must be defined in order to construct groups of close-income equals. Since there are different methods to specify income bandwidth of close-equals (e.g., Ven et al., 2001; Wagstaff, 2002; Duclos et al., 2003), and since the relative importance of decomposition components are shown to be sensitive to different definitions of income bandwidth (Aronson et al., 1994; Urban and Lambert, 2008), we used an annual income interval ranging from 250 New Israeli Shekel (*NIS*) to 15,000 *NIS*[2] to assess the sensitivity of the results to alternative groupings of income bandwidth. A bandwidth of 500 *NIS* – being around 12.5% of the mode value of pre-payment income vector (x) in the study sample – is then, picked up for the initial results.[3] Having selected a bandwidth for close-income equals, the decomposition components are computed directly as a sample statistic—each according to the measurement system specified above.

3. Data, variables definitions

The Palestinian Household Health Expenditure Survey PCBS, 2004 is a two-stage stratified cluster-random sample of 4496 Palestinian households. For the purpose of this paper, the household is taken as the unit of analysis. After the process of deleting the cases with missing relevant information, the WB sub-sample contained 2504 households and GS sub-sample contained 1322 households. The data have been weighted to compensate for non-response and to recover the population profile as per the OPT population Census of 1997. The survey questionnaire included information about household's recent health experiences – utilization and expenditures – during the last month preceding the survey. The questionnaire starts by gathering information about the nature of the health problem(s) – if any – including: acute, chronic, injury/accident, dental and/or psychological. Respondents were then asked about the number of visits made to the health care provider, spectrum of service(s) received during the visit, and about all direct and indirect health expenditures resulting from this episode, including: transportation and any other related expenses, and cost sharing by other stakeholders. Socioeconomic characteristics of households were also collected, including: number of household's members, household mean monthly income and expenditures, household expenditures on the various health care services and on insurance premiums; as well as individual respondents' age, sex, education, marital status, occupation and employment status, type of housing (Urban, Rural, and Camp) and region (West Bank, Gaza Strip) of residence.

The direct measures of living standards available from the survey are household mean monthly income and household mean monthly expenditures. As the later was judged by PCBS (2004) to be more accurate and definitely a better measure of households' incomes we used this variable as a measure of households' living standards. This choice is coherent with current experience of data collection in developing countries (Deaton and Grosh, 2000). Therefore, household pre-payment income variable (x) is apprehended through the total household expenditure, gross of payments for health care, while household post-payment income is pre-payment income, so defined, net of payment for health care. The total household income – as reflected by the total household expenditure – is then annualized and divided by the equivalence factor to generate an adjusted income per equivalent-adult.

[2] At the time of the study 1 *NIS* was equivalent to 0.23 US$.

[3] This is in line with the optimal bandwidth suggested by van de Ven et al. (2001) in which they advocate a bandwidth that maximizes the vertical contribution of the RE.

Payments for health care were computed for three sources of health care financing: direct out-of-pocket payments, GHI, and private health insurance (PHI) schemes.[4] As for out-of-pocket payments, they represent total actual spending, directly and indirectly, paid by the patient in exchange of services offered by governmental, non-governmental, UNRWA and private health care providers. Data on health care spending was gathered by questioning each household about its monthly expenditures on 25 separate health care items, including: consultation fees, laboratory tests, medications (counting for auto-medication), expenditures on traditional healers, transportation costs, and any other related health expenditures. On the other hand, GHI and PHI premiums have been recorded, and were included, in the estimation of household total health expenditures. As for income, total health expenditures were annualized and equivalised using the aforementioned equivalence scale.

4. Results

4.1. Principal decomposition results

Table 1 presents the UL decomposition of the redistributive effects for each, and all source(s) of health care financing in the WB and GS, along with the corresponding values of BTS standard errors and 95% BTS confidence intervals. It can first be noted that among the principal financing sources out-of-pocket payments cause a statistically significantly negative *RE* (−0.0370 and −0.0247 in the WB and GS, respectively). Direct financing is thus "pro-rich" in the sense that payments increase inequality in the post-payment income distributions compared to pre-payment ones. By contrast, GHI and PHI schemes appear to be "pro-poor" in their redistributive effects, as reflected by the positive values of *RE* in both regions. Table 1 shows, however, that the magnitudes of *RE* associated with the two schemes are quite marginal (0.0007 and 0.0001 for GHI and PHI, respectively), and statistically insignificant (at $\alpha = 0.05$). As a result, health care financing remains overall significantly "pro-rich" with *RE* of −0.0379 and −0.0254 in the WB and GS, respectively. Such results exhibit a significant increase in inequality induced by the current financing system in the two Palestinian regions.

The contributions of *VE*, *HE*, and *RR* to income inequality can be better reflected by expressing them as a percentage of the total *RE* (Wagstaff and Doorslaer, 1997). In the case of out-of-pocket payments the values of *VE* (−0.0158) and (−0.0084) would account for 42.7% and 34.9% of the total "pro-rich" *RE* in the WB and GS, respectively. This indicates that the majority of the income inequality generated by such financing modality is due to both *HE* and *RR*. As for GHI, the estimated values of *VE*, which emerge to be significantly positive in the two regions, would account for 172.1% and 148.4% of the total "pro-poor" *RE* in the WB and GS, respectively; indicating that GHI would have been 72.1% and 48.4% more "pro-poor" redistributive if there had been no *HE* and *RR*. Turning to PHI, the value of *VE* would account for 128.5% and 113.6% of the total "pro-poor" *RE* in the WB and GS, respectively; indicating that the PHI scheme would have been more redistributive by 28.5% and 13.6% in the absence of both *HE* and *RR*. However, the positive values of *VE* of private insurance prove to be rather indeterminate and statistically insignificant (at $\alpha = 0.05$).

[4] Since the health expenditure survey did not provide any information on taxes, it was impossible to estimate the amount that would have been paid in taxes for health care by a given household. Therefore, the analysis was confined to three sources of financing which are considered to be the major sources of health care financing in the context of the Palestinian health care system.

Table 1
Decomposition of redistributive effects of principal health care financing schemes in the OPT.[a,b,c]

West Bank	*RE*	*VE*	*HE*	*RR*	R_{WG}	R_{EG}	R_{BG}	*HE*/*RE*	*HE*/*RE*	*RR*/*RE*	R_{WG}/*RE*	R_{EG}/*RE*	R_{BG}/*RE*
Out-of-pocket payments	−3.7004 (0.260) [−4.170, −3.130]	−1.5801 (0.261) [−2.092, −1.051]	0.0601 (0.011) [0.041, 0.071]	2.0602 (0.150) [1.765, 2.354]	0.2402 (0.020) [0.260, 0.354]	0.0270 (0.023) [0.020, 0.063]	1.7930 (0.142) [1.500, 2.030]	42.7%	−1.6%	−55.7%	−6.5%	−0.7%	−48.5%
Governmental Health Insurance	0.0698 (0.040) [−0.010, 0.160]	0.1201 (0.0401) [0.0498, 0.1887]	0.0038 (0.001) [0.001, 0.007]	0.0465 (0.010) [0.030, 0.063]	0.0169 (0.0011) [0.0120, 0.0204]	0.0001 (0.0001) [0.0000, 0.0006]	0.0295 (0.007) [0.019, 0.043]	172.1%	5.5%	66.6%	24.2%	0.1%	42.3%
Private health insurance	0.0140 (0.012) [−0.140, 0.040]	0.0180 (0.015) [−0.010, 0.048]	0.0002 (0.002) [0.0001, 0.0005]	0.0038 (0.001) [0.0002, 0.0101]	0.0022 (0.0010) [0.0012, 0.0033]	0.0000 (0.0000) [0.0000, 0.0000]	0.0016 (0.0008) [0.0006, 0.0033]	128.5%	1.4%	27.1%	15.7%	0.0%	11.4%
Total payments	−3.7908 (0.282) [−4.301, −3.271]	−1.4701 (0.280) [−1.960, −0.943]	0.0602 (0.009) [0.040, 0.080]	2.2605 (0.172) [1.922, 2.591]	0.2603 (0.014) [0.230, 0.299]	0.0200 (0.020) [0.013, 0.131]	1.9802 (0.150) [1.590, 2.250]	38.8%	−1.6%	−59.6%	−6.9%	−0.5%	−52.2%
Gaza Strip	*RE*	*HE*	*HE*	*RR*	R_{WG}	R_{EG}	R_{BG}	*HE*/*RE*	*HE*/*RE*	*RR*/*RE*	R_{WG}/*RE*	R_{EG}/*RE*	R_{BG}/*RE*
Out-of-pocket payments	−2.4773 (0.385) [−3.211, −1.780]	−0.8401 (0.405) [−1.511, −0.080]	0.0610 (0.020) [0.030, 0.092]	1.5762 (0.181) [1.302, 1.955]	0.3002 (0.020) [0.260, 0.354]	0.0210 (0.020) [0.007, 0.093]	1.2550 (0.160) [1.010, 1.610]	33.9%	−2.5%	−63.6%	−12.1%	−0.8%	−50.7%
Governmental Health Insurance	0.0741 (0.041) [−0.010, 0.163]	0.1100 (0.041) [0.030, 0.195]	0.0083 (0.003) [0.003, 0.025]	0.0276 (0.005) [0.020, 0.045]	0.0150 (0.0010) [0.012, 0.018]	0.0000 (0.000) [0.000, 0.000]	0.0126 (0.004) [0.011, 0.022]	148.4%	11.2%	37.2%	20.2%	0.0%	17.0%
Private health insurance	0.0125 (0.011) [−0.004, 0.030]	0.0142 (0.008) [−0.002, 0.032]	0.0004 (0.0003) [0.0003, 0.0014]	0.0013 (0.0006) [0.0003, 0.0260]	0.0010 (0.0004) [0.0003, 0.0017]	0.0000 (0.000) [0.000, 0.000]	0.0003 (0.0002) [0.0002, 0.0010]	113.6%	3.2%	10.4%	8.0%	0.0%	2.4%
Total payments	−2.5406 (0.371) [−3.232, −1.820]	−0.7302 (0.364) [−1.430, −0.043]	0.0703 (0.020) [0.020, 0.110]	1.7401 (0.192) [1.370, 2.111]	0.3100 (0.030) [0.260, 0.370]	0.0201 (0.010) [0.011, 0.095]	1.4100 (0.170) [1.070, 1.710]	28.7%	−2.8%	−68.5%	−12.2%	−0.8%	−55.5%

[a] Income bandwidth set at 500 *NIS*.
[b] All estimates have been multiplied by 100 for readability.
[c] Bootstrap standard errors are in parentheses and bootstrap confidence intervals at 95% are in brackets.

These results suggest that despite the importance of vertical differences (progressivity or regressivity) in the income redistribution induced by the financing system, there is a fairly high degree of *HE* and *RR* associated with each source of financing. Table 1 shows, however, that there are considerable variations in the relative importance of these two effects in generating more income inequality. Unexpectedly, the value of *HE* of (0.0006) in the case of out-of-pocket payments appears to account for only less than −2.5% of the total "pro-rich" *RE* in the two Palestinian regions.[5] This indicates that horizontal differences are responsible for only a quite small amount of the total "pro-rich" *RE* associated with out-of-pocket payment, and therefore, the total *RE* induced by direct payment for health care would have been marginally less "pro-rich" redistributive in the absence of horizontal inequity. In the case of GHI, the values of *HE* appear to be more important and clearly reduce the redistributive "pro-poor" effect of this scheme compared to what it would have been in the absence of horizontal inequity by about 5.5% and 11.2% in the WB and GS, respectively. By contrast, the value of *HE* for PHI is rather small and accounts for only 1.4% in the WB and 3.2% GS of the total *RE*. By considering, lastly, the overall health care payment, the contributions of *HE* to the total "pro-rich" *RE*, remains fairly small, being about −1.6% and −2.8% in the WB and GS, respectively.

These results indicate that the majority of the "additional" increase in the income inequality that is not due to "pure" vertical and horizontal inequities is due to reranking (*RR*). Indeed, in the two cases: the regressive out-of-pocket payments and the progressive GHI, *RR* appears to be responsible for the largest amount of additional variation in the *RE*. As shown in Table 1, our methodological approach further decomposes the total reranking (*RR*) into three components: R_{WG}, R_{EG} and R_{BG}. This decomposition is useful to ascertain the sources of reranking and the contribution of each of them in the total *RE*. By so doing, the inter-group reranking (R_{BG}) emerges to be responsible for the majority of the total reranking induced by out-of-pocket payments and by the total health care payment, and would alone account for about half of the increase in the income inequality in the two regions. On the other hand, intra-group reranking (R_{WG}) would account for −6.5% and about −12.1% of the additional increase in *RE* in the WB and GS, respectively. The contribution of entire-group reranking (R_{EG}) is quite small and would slightly increase the regressive *RE* of out-of-pocket payments (by less than 1.0%) in both regions. In the case of GHI, the relative importance of the R_{BG} and R_{WG} are significantly different between the two regions: in the WB the R_{BG} appears to be responsible for the majority of the decrease in the "pro-poor" *RE* (42.3%) compared to (24.2%) attributed to R_{WG}, whereas in GS the majority of reranking-induced decrease is attributed to R_{WG} (20.2%) compared to (17.0%) due to R_{BG}. Regarding PHI scheme, the R_{WG} constitutes the largest share of reranking-induced decrease in *RE*, whilst R_{BG} comes in the second place. Lastly, in the two regions and for the two schemes, the R_{EG} remains zero.

4.2. Sensitivity of decomposition results to income bandwidth

The results presented in the above section highlight the relative importance of the vertical and horizontal inequities, as well as that of the different forms of reranking, in the total variation of income inequality (*RE*) for an income bandwidth (w) of 500 *NIS*. The concentration coefficients C_y – used to decompose the total *RE* – are bivariate measures of inequality—i.e., they measure *inequality* in one variable; e.g., post-payment income (y), in relation to ranking of another; e.g., pre-payment income (x) (Koolman and Doorslaer, 2004). Since the ranking is variant with respect to the definition of w, changing the size of w – used to group x – is likely to affect the relative magnitudes of the *VE*, *HE*, as well as the relative contributions of each of the three reranking components, R_{BG}, R_{WG}, R_{EG} to the *RE*. This section explores the sensitivity of the decomposition components to different choices of income bandwidths (w).

Results are presented in Figs. 1–6 where the relevant values of decomposition components are plotted against a large range of w – where w taking the values from 250 to 15,000 *NIS* – for both

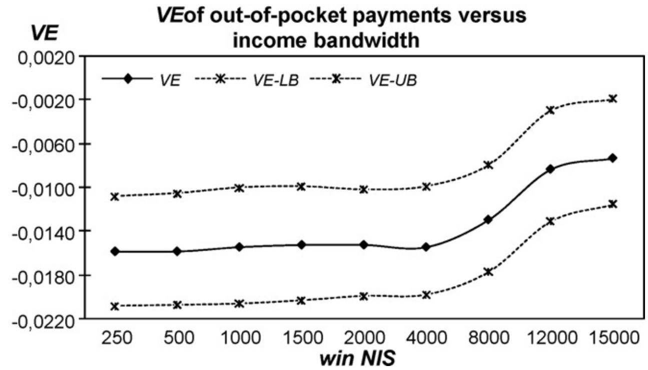

Fig. 1. *VE* of out-of-pocket payments versus income bandwidth.

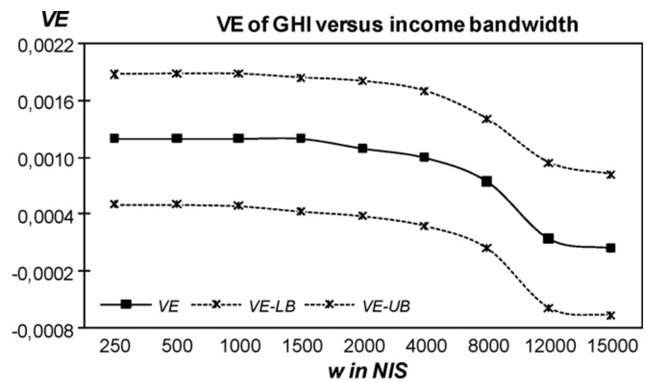

Fig. 2. *VE* of GHI versus income bandwidth.

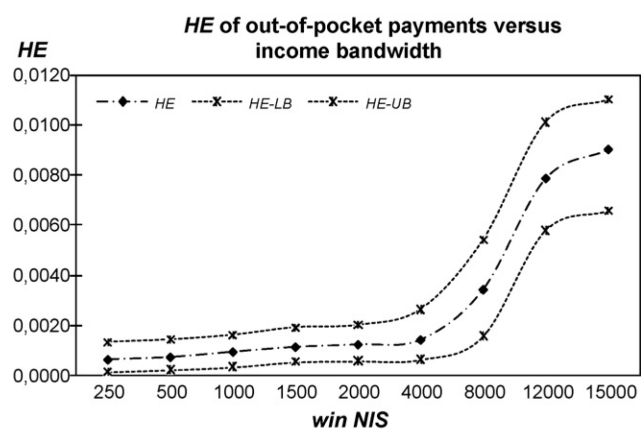

Fig. 3. *HE* of out-of-pocket payments versus income bandwidth.

[5] Note that *HE* appears to be negative when expressed as a percentage of *RE* since the out-of-pocket is a regressive source of financing, and therefore *HE* increases the *RE* in absolute term.

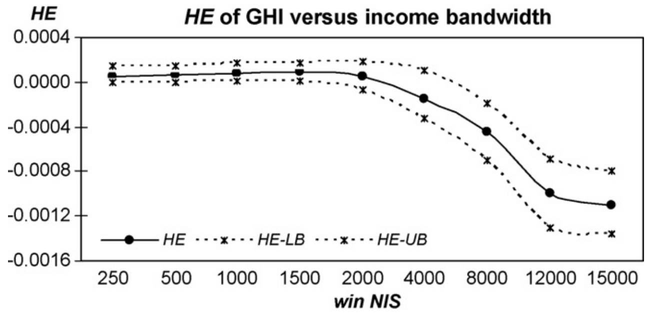

Fig. 4. *HE* of GHI versus income bandwidth.

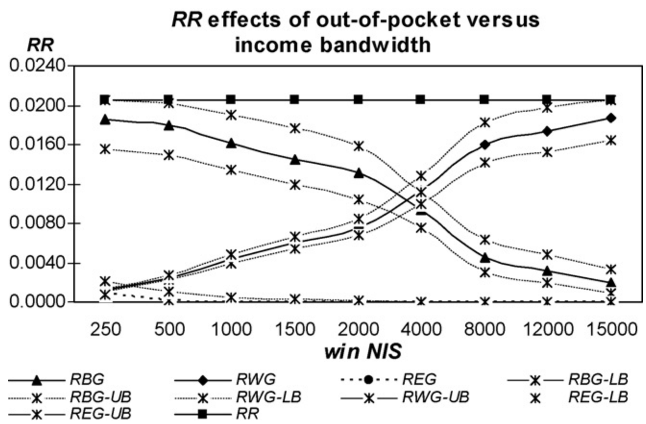

Fig. 5. *RR* effects of out-of-pocket versus income bandwidth.

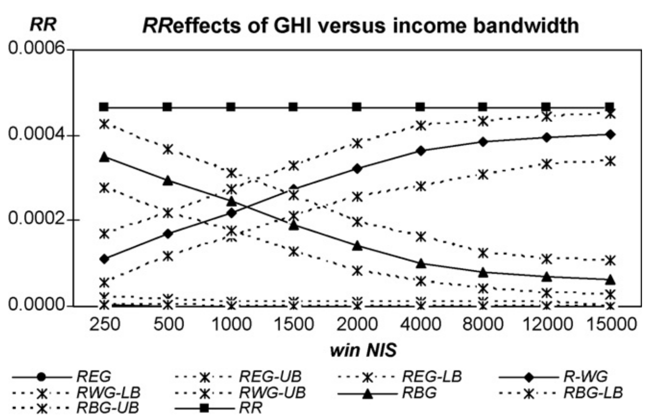

Fig. 6. *RR* effects of GHI versus income bandwidth.

out-of-pocket payments and GHI in the case of WB.[6] The corresponding BTS confidence intervals at a significance level of 95% are presented as dashed lines along w. Figs. 1 and 2 indicate that the *VE* in the two cases: the regressive out-of-pocket payments and the progressive GHI tend to *fall* as w increases; this clearly implies lower contributions of *VE* in the total *RE*. As shown in Figs. 1 and 2, the lower and upper bounds of the confidence intervals of *VE* associated with out-of-pocket payments are both negative along all income bandwidths, while in the case of GHI they are positive along income bandwidths except for the very large bandwidth where the lower bound turns to be negative. This indicates that the *VE* is always significantly negative ($VE < 0\ \forall\, w$ at $\alpha = 0.05$) in case of out-of-pocket payments, and significantly positive ($VE > 0\ \forall\, w < 8000$ *NIS* at $\alpha = 0.05$) in the case of GHI.

Results concerning *HE*, as represented in Figs. 2 and 3, show that in the case of out-of-pocket payments the values of *HE* tend to rise as w increases, suggesting higher contributions of *HE* in the total *RE*. The lower and upper bounds of the confidence intervals in this case confirm that *HE* is always significantly positive ($HE > 0\ \forall\, w$ at $\alpha = 0.05$). In the case of GHI, the estimated values of *HE* tend to rise and are found to be significantly positive only in the small income bandwidths ($HE > 0\ \forall\, w < 2000$ *NIS* at $\alpha = 0.05$). Fig. 3 shows, however, that *HE* turns to be significantly large and negative in the large bandwidths ($HE < 0\ \forall\, w > 2000$ *NIS* at $\alpha = 0.05$). Regarding, the relative importance of different forms of reranking, Figs. 5 and 6 suggest that the values of R_{BG} and R_{EG} get inferior and even approach zero, while the values of R_{WG} get larger as the bandwidth increases. To sum up, the general trend of each decomposition component appears as follows: the larger the bandwidth (w) that is used to construct close-income equals, the lesser the contribution of vertical effect (*VE*), the higher the contribution of horizontal effect (*HE*) and within-group reranking (R_{WG}), and lastly, the lesser the contribution of both inter-groups and entire-groups reranking (R_{BG} and R_{EG}) to *RE*.

5. Discussion

This paper has attempted to transfer recent methodological development in inequality measurement of taxation to the specific domain of health care financing. The analysis of the redistributive impact of health care financing on the overall income inequality has been extended beyond the partial analysis of progressivity and the commonly used AJL approach. A modified decomposition approach that disentangles the total *RE* of health care payment into vertical, horizontal and reranking effects has been implemented, using the measurement model proposed by Urban and Lambert (2008). Such approach provides appropriate measures of inequality that can be normatively distinct in the context where households' incomes are regrouped into close – rather than – exact-income equals, and where the actual payments made by households belonging to these groups may further affect their intra- and entire-group ranking-order. In addition to simultaneously estimate each measure of inequality as sample statistics, the analysis presented in this paper has attempted to examine the statistical inference of each particular measure of inequality using the bootstrap method. Such method provides a basis for assessing the extent of sampling error associated with estimated inequality measures and allows testing statistical significance of each of them within the dominance framework. The analysis was conducted for the three main health care financing schemes proper to the Palestinian context, which has recently experienced sudden and severe impoverishment effects imposed by the chronic political crises.

The decomposition analysis clearly confirms that the "differential treatments" – as reflected by both "unequal treatment of equals" and "improper treatment of unequals" ($HE + RR$) – are together fairly more important in determining the degree of the overall income inequality induced by health care financing than the progressivity (regressivity) contribution that had previously attracted the most attention in the literature. Indeed, the effect of reranking appears to be even more important than the pure horizontal inequity and represent the major factor behind the adverse effect on income inequality. The factors underlying the "improper treatment of unequals" are therefore, of considerable interest. The decomposition approach was able to identify sources of reranking

[6] Similar analysis was also conducted for GS; however, since the analysis demonstrates similar magnitudes of the sensitivity of decomposition components to the choice of bandwidth, we have chosen to include in this paper the results for the WB case only and for two financing sources: out-of-pocket payments and GHI premiums.

in terms of inter-, intra- and entire-groups reranking. The latter two forms of reranking (R_{WG} and R_{EG}), that were not explicitly envisaged in previous research proved be prevalent sources of reranking induced by health care payments. The analysis also reveals a significant effect of the pure horizontal inequity – as identified by person-by-person comparisons of actual and counterfactual post-payment income distributions – in the overall variation of income inequality.

The detailed analysis of the impact of different health care financing schemes on income inequality reveals even more interesting information. It was shown that the impact of out-of-pocket payments on income inequality not only derives from their regressivity but also from differential treatment due to horizontal inequity and reranking. Indeed, horizontal inequity and total reranking combined were responsible for more than half of the “pro-rich” income redistribution associated with out-of-pocket payments. These figures are far greater than those reported for the OECD countries by van Doorslaer et al. (1999) – *HE* and *RR* were estimated, using somewhat different definitions, to be between 3 and 30% of the total *RE* – but remain close to others estimated for Vietnam—61.5% and 70.8% of the total *RE* in 1993 and 1998, respectively (Wagstaff, 2002). Quite interestingly, the decomposition approach shows that the overall reranking effects are more important in terms of their redistributive effect than the vertical and horizontal differences. This is again consistent with previous results found for out-of-pocket payments in Vietnam (Wagstaff, 2002), Nigeria (Ichoku, 2005), and Netherlands (Wagstaff and van Doorslaer, 1997) even if the extent of *RR* in the two Palestinian regions appears to be higher than in these other countries. This indicates that in the context of the predominantly “market-driven” health care financing in the OPT, out-of-pocket health care payments tend to force households not only to buy health care disproportionately to their income but also to affect their income status and, therefore, to exacerbate poverty.

Further examination of the sources of reranking reveals that the so-specified inter-groups reranking (R_{BG}) is responsible for the most part of reranking-induced variation in income inequality associated with out-of-pocket payments (with $R_{BG}\% = -48.5\%$ and -50.7% of the total *RE* in the WB and GS, respectively). This may be due to the stochastic nature of most illnesses, which affects households with different income status; resulting for instance, in richer sick individuals to be overtaken by poorer but healthy ones. On the other hand, the relatively small value of intra-group reranking (R_{WG}) compared with inter-groups reranking (R_{BG}) may, however, reflect quite small disparities in the actual payments within the specified group of equals. The same is true when considering the disparities between the actual and counterfactual income distributions. The latter resulted in relatively small values of pure horizontal inequality, which account for 1.6% and 2.5% of the total pro-rich *RE* in the WB and GS, respectively. These figures appear to be quite small when compared to those previously obtained for developed countries; e.g., 11.3% for Netherlands (Wagstaff and van Doorslaer, 1997), and developing countries; e.g., 25% for Vietnam (Wagstaff, 2002). It should be noted, however, that some of these differences may reflect, in part, the different methodologies used to account for pure horizontal inequity; in previous studies, horizontal inequity was measured using the AJL approach as a residual term (Wagstaff and van Doorslaer, 1997; van Doorslaer et al., 1999; Wagstaff, 2002). It is, as noted earlier, where the AJL approach is applied to close-income equals setting, the *HE* component tends to be overestimated by de facto incorporating the R_{WG} term.

By pointing out that the “improper treatment of unequals” may be a more serious problem associated with out-of-pocket payments than the “differential treatment of equals”, our results raise an important question from a policy perspective, regarding the potential causes of such *RR*. In the context of developed countries a number of factors, in addition to the randomness incidence of illness, were identified to be responsible for the presence of reranking; e.g., variations in private insurance coverage against public sector co-payments, variations in health services utilization and institutional arrangements of public insurance systems (van Doorslaer et al., 1999). In the context of OPT – where no universal insurance coverage exists, and where private insurance is so far limited, the randomness of illness and the size of payments involved seem to be the most likely factor behind such reranking. Yet, another potential explanatory factor may be the variability in practical difficulties to access health care facilities according to different locations and political realities (e.g., refugee camps, effects of Israeli military occupation, etc.).

In addition, it must be noted that the current structure of out-of-pocket payments in the OPT is a rigid one, with generally no price-discrimination policies that may take into account inter-households' contributive capacities and no exemptions or reduction in the amount of payments that may account for non-income criteria, such as age, pregnancy and disablement. This is especially the case of the private health care delivery sector, which plays a non-negligible role in health care provision and finance: about 21.4% of total health care visits take place at private health care institutions and result in 40.5% of total national health care expenditures (PCBS, 2006a). In this context, the relatively small observed variations in payments at each income level, as reflected by both the pure horizontal inequity and the intra-group reranking terms in our estimations, may be due to such rigid payment structure of out-of-pocket payments in the OPT, while the randomness of illness remains the most probable source of this inequality. It seems, therefore, evident to consider the “improper treatment of unequals” as a more serious policy concern than the “differential treatment of equals”.

Results on the overall inequality variation associated with health insurance schemes – both the governmental and private schemes – appear to be less conclusive: the two schemes appeared to have marginal and statistically insignificant equalizing effects on the income distribution. Results such as these seem to reflect, in part, the relatively low shares of the public and private insurance schemes in the overall financing mix in the two Palestinian regions. Some, though, is undoubtedly due to the fact that converge is not universal, and hence, the *RE* associated with the insurance schemes are likely to be driven by variations in coverage across different income groups. On the other hand, the substantial values of *HE* and R_{WG} observed in the case of GHI may be related to the way premiums are established at each income level, resulting from households of similar (equivalent) incomes making dissimilar (equivalent) contributions. In addition, the large observed values of inter-group reranking (R_{BG}) may also reflect the great diversity of GHI institutional arrangements and variations in coverage at each income level in relation to the fact that GHI is compulsory for public sector employees only, whereas it is of a voluntary nature for others.

Indeed, the governmental insurance scheme in the OPT involve four different types of enrolments: the public sector employees are compulsory enrolled and pay a fixed percentage (5%) of their basic income up to a ceiling of 75 *NIS*; self-employed individuals and wage-earners can be voluntary enrolled by paying respectively a monthly premium of 75 *NIS* and of 50 *NIS*, i.e. lump sum payment regardless of individual incomes; the last category concerns the exempted households, e.g., hardship cases, who are covered by the ministry of social affairs with minimum premiums of 45 *NIS* being paid on their behalf (Schoenbaum et al., 2005). On the other hand, a recent extension of GHI coverage has been achieved through the so-called “Al-Aqsa Intifada” insurance scheme, which was set up by the

Palestinian ministry of health following the current crises in 2000. The Ministry has offered an almost "free of charge" coverage to the mostly affected classes of population, and a very low-premium insurance was later introduced to offer coverage for a high percentage of uninsured households in the WB and GS.

It is well-known that introducing exemptions can enhance the progressivity of a financing scheme to the extent that lower income deciles are concerned (Wagstaff, 2002). Although the extension of GHI coverage under "Al-Aqsa Intifada insurance" was frequently decided on some income-related criteria, like the loss of jobs due to the strict closure, as well as non-income-related criteria, like disablement or injuries during the Second Palestinian Intifada, the latest extension in coverage has randomly opened the enrolment in GHI regardless of these criteria. Moreover, although the aim of such extension was to promote equity in the provision of publicly financed health care, the recent increase in the number of households entitled for public services through the "Al-Aqsa Intifada insurance" has not been associated with a parallel improvement in the capacity of health services delivery. This has led to a further deterioration in the quality of care provided (MAS, 2000), a significant decrease in voluntary enrolment, and consequently in GHI total revenues (PMoH-MHIS, 2002; PCBS, 2004). Our results strongly suggest that these unplanned evolutions have undoubtedly affected the magnitude of progressivity of GHI contributions and resulted in households on different incomes making disproportionate contributions and in households with relatively high incomes to contract out of this public insurance scheme. These evolutions have clearly limited the potential positive effects of such insurance scheme in protecting poor people from the impoverishing effects of catastrophic health care payments and in reducing the adverse impact of health care payment on the overall generalized income inequality in the OPT.

Although the analysis undertaken in this study has tried to benefit from the latest methodological developments in the field of inequality measurement, some practical limitations that might have affected the study results are worth mentioning. Firstly, the analysis was conducted using data from the recent national Household Health Expenditure Survey (PCBS, 2004). Although, household expenditure surveys are generally considered to be an adequate source of data for inequality analysis (Ravallion and Chen, 1997; Cowell, 2000), our survey lacks information on health care expenditures paid through direct and indirect taxation. This is a clear limitation knowing that the survey was originally designed to help build a system of National Health Accounts for the OPT, including such data would have offered the opportunity to assess the equity implications of overall health care financing in the OPT. Secondly, data on "third-party financial assistance" – other than private or government insurance – are also missing or unreliable which is also a limitation since patients, especially those who are members of poor households, may have received financial aide from third-party (e.g., NGOs, charitable institutions, etc.) to help cover part of their health care costs.

Another serious limitation is the short recall period for which a payment source was recorded. This is the case for out-of-pocket payments, where data are based upon a one-month recall period and may be subjected to eventual biases due to stochastic and seasonal nature of illness. Annualising out-of-pocket payments in the presence of such seasonalities, by multiplying with some scaling factor, may lead to over- or under-estimation of total health expenditures. Although such practice is considered to be adequate for distribution analysis of aggregates (e.g., deciles in the case of progressivity), it might not be reasonably adequate for an accurate analysis in the case of decomposition (van Doorslaer et al., 1999). This might be avoided in future studies should health expenditure information be collected over a longer period of time.

However, in spite of this fact, a short recall period may help reduce the potential recall bias about health expenditures. Indeed, in the present study, recall bias might not have been a major problem. This is not only due to the relatively short reference period (one month) but also due to the detailed methodology of the survey, where respondents were asked about all different forms of health expenditures one by one. Compared to a scenario where the households are asked about the total sum of their health expenditures over a longer period of time, this detailed data collection method may have reduced the possibility that some items be neglected by the households. Finally, it is worth noting that this direct measure of health care expenditure ignores some important aspects, such as waiting time, loss of productivity, quality of care and variations in tariffs charged by different health care providers (e.g., public, private for profit, and private not for profit health care), all of which may have a significant impact on total expenditures (Juster and Stafford, 1991; Preker and langenbrunner, 2005).

6. Conclusion

In spite of their limitations, the results presented in this paper provide a useful and detailed picture of the overall inequality variation associated with the current Palestinian health care financing structure. Such results should help shape policy toward building an equitable and efficient health care financing system for the OPT. Firstly, given the finding that out-of-pocket payments are associated with pronounced adverse effects on the already unbalanced income distribution, an urgent need is there to identify innovative financing mechanisms capable to reduce the financial burden of health care expenditure and to limit the existing strong regressivity in the system. Among the potential policy measures is a reduction in the real cost of health care (e.g., medications and health professionals' tariffs). Indeed, in the current context, the cost of medications and health professionals' tariffs – especially those charged by specialists – absorb the biggest share of health care expenditures (PCBS, 2004). Secondly, although the above mentioned policy shall enhance vertical equity in the system, it would not per se be able to significantly reduce the considerable amount of differential treatment (horizontal inequity and reranking), which was found to be the most important factor behind the adverse effect of health care payments on households' incomes. It is well-established that the bulk of this differential treatment is largely driven by the stochastic nature of illness and the size of direct payment involved (Wagstaff, 2002). Therefore, a far bigger reduction would only be possible through a shifting from ex-post-payments to ex-ante mode of financing. This might be accomplished by introducing properly designed community-based insurance arrangements.

Although proven to be promising, the current structure of the Governmental Health Insurance scheme needs to be reconsidered to further enhance its positive intrinsic capacities. Indeed, this public insurance scheme, which includes various enrolment arrangements, appears to have a considerable potential vertical effect when horizontal inequity and reranking are eliminated or reduced. Given that the sizable amount of such differential treatment is associated with the current insurance arrangements, it seems vital to reconsider the structure of insurance contributions that are associated with various enrolment arrangements. This also requires reconsidering the unplanned extension of insurance coverage through the so-called "Al-Aqsa Intifada insurance". In addition, despite the fact that private health insurance schemes are far limited in the OPT and cater for a very small proportion of the population, it was found that such financing modality could still play a positive role in protecting households against the adverse effects of ex-post-payments, should enrolment's premiums be suitably linked to households' various abilities-to-pay (Pauly et al., 2006).

Finally, although this paper attempted to shed the light on the sources of inequality associated with the current health care financing arrangements in the OPT, a number of issues still call for further research. Among these are the determinants of health care seeking behaviour in the OPT, not only the classical socioeconomic factors but also the political realities (i.e. locality types and regions), which might affect the horizontal inequity and reranking. This is germane to the Palestinian situation where access to health care is highly influenced by Israeli measures of security (e.g., Separation Wall, checkpoints, etc.). Another interesting area of research would be to examine the value-added of extending health insurance coverage and the possibility of making insurance available for poor people without increasing inequalities.

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