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MEASURING MULTIDIMENSIONAL INEQUALITY AND WELL-BEING: METHODS AND AN EMPIRICAL APPLICATION TO ITALIAN REGIONS

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1. INTRODUCTION

In recent years there has been a growing consensus in favour of including other dimensions, beyond monetary indicators, in analysing well-being and a broad theoretical literature on the subject of multidimensional inequality and well-being, mainly based on the conceptualisation of Sen's (1985, 1987) "capability approach", has emerged (Maasoumi, 1986; Tsui, 1995, 1999; Bourguignon, 1999; Bourguignon and Chakravarty, 2003; Weymark, 2006; Decancq et al., 2009).

This shift to a multidimensional view of welfare analysis has not been confined to academic research and has extended to policy-oriented analysis. As a confirmation of this, since 1990 the United Nations Development Programme has brought into question the primacy of GDP per capita as a welfare measure by proposing the Human Development Index (HDI), which combines income with life expectancy and educational achievements and is now one of the most influential measures of well-being. The report of the Stiglitz-Sen-Fitoussi Commission (Stiglitz et al., 2009) is one of the most recent attempts of measuring well-being moving beyond GDP and following a multidimensional view of human development.

The main differences between the alternative strategies proposed essentially relate to whether each dimension is singularly or jointly investigated, and whether or not multidimensionality is collapsed into a summary indicator (Brandolini, 2008). If it is assumed that the different dimensions cannot be directly compared, other indicators of well-being can be considered in conjunction with information on income, but carrying out a one-by-one analysis of each component. According to this *item-by-item* approach, followed among others by Sen (1985) and Fahey et al. (2005), no attempt to reduce complexity and to aggregate the attributes is made. The advantage of this strategy rests on its simplicity; however, as the indications on inequality of the different dimensions diverge, its lack of synthesis makes it impossible to draw any general conclusion on overall inequality.

Conversely, in all aggregation procedures the basic problem is how to select and summarize the multiple dimensions of individuals' well-being, the so called *indexing problem* (Rawls, 1971). In this context, a first approach consists in specifying a composite index of well-being (thus making the problem unidimensional) and then, in a second stage, computing a standard univariate inequality index (Maasoumi, 1986, 1999). A second approach proposed in the literature is to derive multivariate indices of inequality that satisfy some desirable properties and can be directly applied to the vectors of attributes (Tsui, 1995, 1999; Bourguignon, 1999).

The aggregate measures of well-being and multidimensional inequality require in any case decisions to be made regarding the functional form of the social welfare function, the degree of substitutability between the different attributes of well-being, their weights in the composite indicator, the transfer sensitivity (aversion to inequality) of well-being between individuals. Many empirical analyses demonstrate that multidimensional inequality and well-being measures are very sensitive to all these choices. Rather then choosing an arbitrary combination of such parameters, calculating a set of measures across a range of them is, therefore, a more appropriate strategy in multidimensional inequality and well-being comparisons across time and space. As outlined by Brandolini (2008), far from being a weakness of multidimensional approach, the investigation of the effects of alternative normative assumptions and the lack of uniqueness in the results enrich the informative value of the multidimensional perspective.

In this paper, focusing mainly on normative multidimensional inequality measures, these two approaches are presented and discussed (Section 2). The more commonly used inequality adjusted multidimensional well-being indicators are also reviewed. We show, in particular, that such indicators can be seen as special cases of the well-being measures implicit in the multidimensional inequality indices (Section 3). Using Italian data on individual income, education and health status from the 2005 and 2008 Italian Survey on Income and Living Conditions (IT-SILC), an empirical analysis of multidimensional inequality and inequality adjusted well-being levels in Italian regions is then performed. To the best of our knowledge, this is one of the first analyses of multidimensional inequality and well-being at the regional level and exploits the fact that the IT-SILC sample assures estimates' consistency at the NUTS 2 level. The analysis mainly aims at illustrating the impact of different normative choices concerning the degree of substitutability between dimensions and the degree of inequality aversion on regional indicators of multidimensional inequality and well-being (Section 4). Given the variability that characterizes multidimensional indicators, depending on the uncertainty connected to the survey nature of the data as well as on the alternative parameter combinations chosen, the regional well-being scores are presented together with the corresponding confidence intervals, based on bootstrapped standard errors. Some remarks conclude the paper (Section 5).

2. THE NORMATIVE APPROACH TO THE MEASURE OF MULTIDIMENSIONAL INEQUALITY IN WELL-BEING

2.1 *Inequality and well-being in one dimension*

Before introducing multidimensionality, it is useful recalling the Atkinson-Kolm-Sen approach in measuring inequality and well-being in one dimension, say income.

Let $y = \{y_1, ..., y_n\}$ be a vector of incomes of *n* individuals (or households) and $W(y) = \sum_i U(y_i)$ an additively separable social welfare function (SWF), where $U(y_i)$ is an increasing and strictly concave utility function used by the social evaluator to assess the well-being of a generic unit and need not to coincide with individual's utility function. The inequality measure is based on the concept of *equally distributed equivalent* (*ede*) *income* y_e, defined as the level of income that, if obtained by each individual, produces the same social welfare as the observed distribution:

$$W(y_1, ..., y_n) = W(y_e, ..., y_e) \Rightarrow \sum_i U(y_i) = nU(y_e). \tag{1}$$

The corresponding Kolm-Atkinson relative inequality index is:

$$I(y) = 1 - \frac{y_e}{\mu}, \text{ where } \mu = \frac{1}{n}\sum_i y_i, \tag{2}$$

and the relationship between inequality and well-being is:

$$W(y) = y_e = \mu \cdot (1 - I(y)). \tag{3}$$

In particular, the Atkinson (1970) index is related to the following class of utility functions

$$\begin{cases} U(y_i) = \dfrac{1}{1-\varepsilon} y_i^{1-\varepsilon} & \text{for } \varepsilon \geq 0,\ \varepsilon \neq 1 \\ U(y_i) = \ln y_i & \text{for } \varepsilon = 1 \end{cases}, \tag{4}$$

where the parameter ε reflects the social aversion to inequality. When $\varepsilon > 0$ there is aversion to inequality and as ε rises society attaches more weight to transfers at the lower end of the distribution, coherently with the Pigou-Dalton transfer principle.

From the SWF associated to welfare function $U(y_i)$ the following expression for *ede* is obtained:

$$\begin{cases} y_e = \left(\dfrac{1}{n}\sum_i y_i^{1-\varepsilon}\right)^{\frac{1}{1-\varepsilon}} = \mu_{1-\varepsilon} & \text{for } \varepsilon > 0,\ \varepsilon \neq 1 \\ y_e = \prod_i (y_i)^{\frac{1}{n}} = \mu_0 & \text{for } \varepsilon = 1 \end{cases} \tag{5}$$

where $\mu_{1-\varepsilon}$ is the generalized mean of order $1-\varepsilon$ (μ_0 is the geometric mean), while the inequality index can be expressed as:

$$I_\varepsilon(y) = 1 - \frac{\mu_{1-\varepsilon}}{\mu}, \tag{6}$$

and the relationship between well-being and inequality is $\mu_{1-\varepsilon} = \mu \cdot (1 - I_\varepsilon(y))$.

2.2 *The multidimensional setting*

2.2.1 *Notation, axioms and properties*

We assume that the domains of well-being have been identified and that the achievements in all the dimensions are interpersonally comparable. Consider that there is a fixed set of individuals (households, regions, countries, etc.) $N = \{1,...,n\}$ (with $n \geq 2$) and the set of dimensions of well-being (attributes) is $K = \{1,...,k\}$. A distribution of attributes among the population is an $n \times k$ real-valued matrix X, with the ij-th element x_{ij} representing individual i's quantity of the j-th attribute (with $x_{ij} \in \mathbb{R}_+$). The i-th row of X is denoted $\underline{x}_i = \{x_{i1},...,x_{ik}\}$ and can be interpreted as a well-being vector for individual i, since it summarizes the achievement of the individual on all the dimensions considered. Therefore, the *indexing problem* here consists in the search for an appropriate well-being index, S, that maps the well-being bundles $\underline{x}_i$ on the set of real numbers ($S(X): X \mapsto \mathbb{R}$), so that they can be naturally ordered and used to assess the position of any two individuals and the distance between them (Decancq and Lugo, 2010).

As in the single-dimensional case, each multi-attribute inequality index satisfies, either implicitly or explicitly, a set of properties which define the specific functional form of the index. Following Weymark (2006), there is a number of basic properties that a multidimensional inequality index should satisfy. These can be grouped in two sets of axioms. The first consists in those properties that are not concerned with the distributional sensitivity of the inequality measure. These *non-distributional axioms*, which are straightforward generalization of their unidimensional counterparts, are: Continuity, Anonymity, Normalization, Replication Invariance, Scale Invariance, Decomposability and Additive Separability (by population subgroups and by dimensions)[1].

The second set includes distributional properties (or majorization criteria), which provides partial orders that ranks distribution matrices in terms of their degree of inequality. Several authors have tried to provide multivariate generalizations of the Pigou-Dalton principle of transfer by directly imposing conditions in the space of the distribution matrices X. In particular, Kolm (1977) proposed the *Uniform Majorization Principle* (UM), defining the condition that premultiplication of a distribution matrix by a bistochastic matrix (i.e. a non negative square matrix with row and column sums equal to one) should lead to a socially preferred state. The UM principle imposes that a mean-preserving averaging performed uniformly on all dimensions leads to an increase in social welfare. Another multi-

[1] For a detailed discussion of each non-distributional axiom see Weymark (2006) and Lugo (2007).

attribute Pigou-Dalton transfer principle is the Uniform Pigou-Dalton Majorization Principle (UPD) which defines the condition that pre-multiplying a distribution matrix for the product of a finite number of $n \times n$ Pigou-Dalton transfers leads to a reduction in multidimensional inequality. The UPD condition is less restrictive than the UM principle.

The UM criterion allows to measure inequality with respect to the dispersion of the multidimensional distribution of the attributes, but fails in addressing the second dimension of multivariate inequality. In fact, Atkinson and Bourguignon (1982) argued that a multidimensional inequality index should also account for the dependence between dimensions and developed a dominance criterion, later formalized by Tsui (1999) as the *Correlation Increasing Majorization* (CIM). This criterion is based on the idea that, given two distribution matrices with the same marginal distributions for the dimensions but with different degree of correlation between dimensions, the one with less correlation is socially preferred. Atkinson and Bourguignon show that this condition is satisfied for any increasing well-being index with negative cross-derivatives.

2.2.2 *The two-step approach*

Pioneered by Maasoumi (1986), the two-step approach has the advantage of making the aggregation procedure explicit, arriving firstly to a single composite well-being measure for each individual and then applying some univariate inequality index. In the first step, Maasoumi use information theory and in particular a multivariate generalization of the general entropy measures, obtaining the following class of optimal aggregation functions showing constant elasticity of substitution (CES):

$$S_\beta(\underline{x}_i) = \left[\sum_{j=1}^{k} w_j x_{ij}^{\beta} \right]^{1/\beta} . \tag{7}$$

The individual well-being index $S_\beta(\underline{x}_i)$ is a generalized (weighted) mean of order β of the achievements in each well-being dimension. The dimension-weights $w_1, \ldots, w_k$ are equal across individuals and assumed to sum up to one. Weights determine how functionings contribute to well-being, and alternative weighting structures reflect different views on the relative importance of the attributes.

The parameter β is related to the degree of substitutability between attributes σ (with $\sigma = 1/(1-\beta)$) and determines the shape of the contours for all pair of attributes. The smaller is β, the smaller is the substitutability between dimensions, that is the more one has to give up of one attribute to get an extra unit of a second attribute while keeping the level of well-being constant[2]. Generally, for $\beta \leq 1$ (i.e. non-negative elasticity of substitution) the well-being index is a weakly

[2] The CES formulation is criticized by Bourguignon and Chakravarty (2003) on the ground that in case of more than two attributes it implies the same elasticity of substitution between all dimensions.

concave function, which reflects a preference for well-being bundles that are more equally distributed. In the limit, as $\beta \to -\infty$ and $\sigma \to 0$, dimensions are treated as perfect complements and the well-being function is of Leontief type, thus favouring individuals with more balanced achievements among all the dimensions.

Significant special cases are obtained for $\beta = 0$ and $\beta = 1$. When β is equal to 0 the composite indicator of well-being $S_\beta(\underline{x}_i)$ is a Cobb-Douglas function, with unitary elasticity of substitution:

$$S_0(\underline{x}_i) = \prod_{j=1}^{K} x_{ij}^{w_j} \ . \tag{8}$$

When $\beta = 1$ the well-being indicator is a linear function of the k attributes and $\sigma \to \infty$, that is attributes are perfect substitutes so that low levels on one of them can be perfectly compensated by high levels on another.

Returning to the expression of $S_\beta(\underline{x}_i)$, the original values of the indicators in X are often firstly transformed, for three main reasons: to capture the diminishing returns in the conversion into well-being of some attributes of the composite indicator, especially income; to reduce the effect of extreme values and outliers when the original distribution is markedly skewed; to rescale well-being attributes before they can be sensibly aggregated, as they are generally measured in different units of measurement. If we define $f_j(\cdot)$ $(j = 1,...,k)$ to be the dimension-specific transformation functions, it is possible to obtain the transformed distribution matrix Z and reformulate equation (7) to account for the transformation introduced:

$$S_\beta(\underline{z}_i) = \left[\sum_{j=1}^{k} w_j [f(x_{ij})]^\beta\right]^{1/\beta} = \left[\sum_{j=1}^{k} w_j z_{ij}^\beta\right]^{1/\beta} . \tag{9}$$

Obviously, different choices for β, for the weighting structure and for the functions $f_j(\cdot)$ will lead to different composite indices. As an illustrative example, within the general framework outlined in equation (9) it is possible to obtain the Human Development Index (HDI), which is a composite index of three well-being indicators at the country level: standard of living (the logarithm of GDP per capita), health (life expectancy at birth), and education (measured by a composite index of adult literacy rate and school enrolment rate). These indicators are normalized such that they reflect the achievements in terms of percentage from the minimum to the maximum values and are aggregated by a simple weighted mean of order one (with weights $w_j = 1/3$). This implies that the parameter β in equation (9) is set equal to 1, i.e., the dimensions are assumed as perfect substitutes.

Once a composite index of well-being has been defined, it is possible to calculate overall inequality by applying a unidimensional inequality measure in a second step. The pioneering inequality measure proposed by Maasoumi (1986) is obtained by calculating a Generalized Entropy index on the vector of $S_\beta(\underline{z}_i)$.

The two-step inequality indices can be also obtained within the normative approach (see Decancq et al (2009) by specification of an additively separable SWF defined over $S_\beta(\underline{z}_i)$:

$$W(Z) = \sum_{i=1}^{n} U[S_\beta(\underline{z}_i)] = \frac{1}{1-\varepsilon}\sum_{i=1}^{n}[S_\beta(\underline{z}_i)]^{1-\varepsilon}\,, \tag{10}$$

with $U(\cdot)$ defined as in (4) and the parameter ε (with $\varepsilon \geq 0$) reflecting the social aversion to inequality in the composite indicator of well-being.

The unidimensional Atkinson-Kolm-Sen inequality $I_\beta^U(Z)$ can be then obtained as the scalar that solves:

$$W[(1 - I_\beta^U(Z)) \cdot \mu(S_\beta)] = W(Z)\,, \tag{11}$$

where $\mu(S_\beta)$ is the mean composite well-being index across the individuals. $I_\beta^U(Z)$ will then measure the overall well-being that could be given up if well-being is equalized among individuals, while holding the overall social welfare unaltered. From SWF (10), the unidimensional Atkinson measure of inequality in $S_\beta(\underline{z}_i)$ can be written as:

$$I_\beta^U(Z) = 1 - \left[\frac{1}{n}\sum_{i=1}^{n}\left(\frac{S_\beta(\underline{z}_i)}{\mu(S_\beta)}\right)^{1-\varepsilon}\right]^{1/(1-\varepsilon)}. \tag{12}$$

In terms of generalized means, $I_\beta^U(Z)$ can also be rewritten as:

$$I_\beta^U(Z) = 1 - \frac{\mu_{1-\varepsilon}(\mu_\beta(\underline{z}_i))}{\mu(\mu_\beta(\underline{z}_i))}\,, \tag{13}$$

where $\mu_\beta(\underline{z}_i)$ is the weighted mean of order β of the achievements in each well-being dimension and $\mu_{1-\varepsilon}$ is the mean of order $1-\varepsilon$ calculated on the individual values of the composite well-being index $S_\beta(\underline{z}_i)$.

Within this framework, the choice of the transformation functions in equation (9) must be carefully considered, as $I_\beta^U(Z)$ is invariant for proportional changes in well-being levels (i.e. is *scale invariant*), but other transformations will generally lead to changes in the inequality measure (Decancq *et al.*, 2009).

2.2.3 *The multivariate approach*

Even if the definition of a composite well-being indicator and the application of a univariate inequality index may seem a natural approach, it does not allow to fully capturing the multidimensional nature of well-being. For this reason, a second approach proposed in the literature is to derive multivariate indices of inequality that satisfy some desirable properties and can be directly applied to the vectors of attributes.

Following Weymark (2006), a relative multidimensional inequality measure $I_{\beta}^{M}(Z)$ can be derived starting from a continuous, strictly increasing, anonymous, strictly quasi-concave, separable and scale invariant multidimensional SWF $W(Z)$, as:

$$W[(1-I_{\beta}^{M}(Z))\cdot Z_{\mu}]=W(Z), \tag{14}$$

where Z_{μ} is a distribution matrix, where every observation is replaced by its column mean. $I_{\beta}^{M}(Z)$ is a multidimensional generalization of the standard unidimensional Atkinson-Kolm-Sen index, and represents the fraction of the aggregate amount of each attribute that can be destroyed if every dimension is equalized and the resulting distribution is socially indifferent to the original one.

A pioneering example of multidimensional generalization of the Atkinson-Kolm-Sen approach is proposed by Tsui (1995) who derives the following inequality index, starting from a multidimensional SWF which satisfies the above mentioned properties:

$$I^{T}(Z)=1-\left[\frac{1}{n}\sum_{i=1}^{n}\left[\prod_{j=1}^{k}\left(\frac{z_{ij}}{\mu_{j}}\right)^{r_j}\right]\right]^{1/\sum_j r_j}, \tag{15}$$

where μ_j is the mean of attribute z_j.

By assuming $w_j(1-\varepsilon)=r_j$, such that $(1-\varepsilon)=\sum_j r_j$ (see Lugo, 2007; Brandolini, 2008), the Tsui index (15) can be rewritten as:

$$I^{T}(Z)=1-\left[\frac{1}{n}\sum_{i=1}^{n}\left(\frac{\prod_{j=1}^{K}z_{ij}^{w_j}}{\prod_{j=1}^{K}\mu_{j}^{w_j}}\right)^{1-\varepsilon}\right]^{1/(1-\varepsilon)} = 1-\left[\frac{1}{n}\sum_{i=1}^{n}\left(\frac{S_0(\underline{z}_i)}{S_0(\underline{\mu})}\right)^{1-\varepsilon}\right]^{1/(1-\varepsilon)}, \tag{16}$$

where $S_0(\underline{z}_i)$ is a special case of the Maasoumi individual composite well-being index $S_\beta(\underline{z}_i)$, with $\beta = 0$ (Cobb-Douglas function) and $S_0(\mu_j)$ is the corresponding composite well-being of the mean individual, that is the individual endowed with the mean value of each attribute. It is worth noting that this normalization by the "representative" well-being, instead of the mean value of the aggregator $S_\beta(\underline{z}_i)$ as in (12), is the one suggested by Bourguignon (1999) in his comment to Maasoumi (1999).

Weymark (2006) points out that the Cobb-Douglas aggregation function of the Tsui index is quite restrictive, mainly because of the *strong ratio-scale invariance* assumption. As shown by Decancq and Lugo (2009), relaxing this assumption in favour of *weak ratio-scale invariance* is the condition to obtain, in presence of other properties (monotonicity, normalization and separability), the more flexible CES functional form as a suitable aggregation function between attributes. In other words, there is a trade-off between flexibility of the aggregation function and the rescaling procedure of the different dimensions.

The strong ratio-scale invariance assumption was originally questioned by Bourguignon (1999), who proposed, in fact, a CES functional form for the composite well-being indicator, which also includes the inequality aversion parameter (see also Lugo, 2007):

$$S_{\beta,(1-\varepsilon)}(\underline{z}_i) = \left[\left(\sum_{j=1}^{k} w_j z_{ij}^{\beta}\right)^{1/\beta}\right]^{1-\varepsilon}, \tag{17}$$

with $\beta \leq 1$ and $\varepsilon > 0$.

Solving $W(Z) = \sum_{i=1}^{n} S_{\beta,(1-\varepsilon)}(z_i) = W[(1-I^B)\cdot\mu]$, where $\mu = \left[\left(\sum_{j=1}^{K} w_j \mu_j^{\beta}\right)^{1/\beta}\right]^{1-\varepsilon}$,

the corresponding multidimensional inequality index is then obtained:

$$I^B = 1 - \frac{1}{n}\frac{\sum_{i=1}^{n}\left[\sum_{j=1}^{K}(w_j z_{ij}^{\beta})^{1/\beta}\right]^{1-\varepsilon}}{\left[\sum_{j=1}^{K}(w_j \mu_j^{\beta})^{1/\beta}\right]^{1-\varepsilon}}. \tag{18}$$

A similar index has been proposed by Decancq et al. (2009), based again on a CES functional form (9), from which they obtain the following generalization of the Tsui (1995) index:

$$I_{\beta}^{M}(Z)=1-\left[\frac{1}{n}\sum_{i=1}^{n}\left[\frac{\left[\sum_{j=1}^{k}w_j z_{ij}^{\beta}\right]^{1/\beta}}{\left[\sum_{j=1}^{k}w_j \mu_j^{\beta}\right]^{1/\beta}}\right]^{1-\varepsilon}\right]^{1/(1-\varepsilon)}=1-\left[\frac{1}{n}\sum_{i=1}^{n}\left(\frac{S_{\beta}(\underline{z}_i)}{S_{\beta}(\underline{\mu})}\right)^{1-\varepsilon}\right]^{1/(1-\varepsilon)} \quad (19)$$

It is worth noting that index (19) is linked to the Bourguignon index (18) by the following relationship: $I^{B}=1-(1-I_{\beta}^{M})^{1-\varepsilon}$.

As for $I_{\beta}^{U}(Z)$, also $I_{\beta}^{M}(Z)$ and $I^{T}(Z)$ can be rewritten in terms of generalized means as:

$$I_{\beta}^{M}(Z)=1-\frac{\mu_{1-\varepsilon}(\mu_{\beta}(\underline{z}_i))}{\mu_{\beta}(\underline{\mu})}, \quad (20)$$

and

$$I^{T}=I_{0}^{M}(Z)=1-\frac{\mu_{1-\varepsilon}(\mu_{0}(\underline{z}_i))}{\mu_{0}(\underline{\mu})}. \quad (21)$$

where $\mu_{\beta}(\underline{\mu})$ is the mean of order β of the (arithmetic) mean achievements in each well-being dimension[3].

With respect to the framework defined in (9), and for the corresponding class of multidimensional inequality indices, the CIM principle is satisfied provided that $\varepsilon+\beta>1$. So, for the Tsui index CIM is valid only if at least a relatively high aversion to inequality is assumed ($\varepsilon>1$), while for its generalization (19), as for the Bourguignon index (18), the same principle is satisfied only if the substitutability parameter β is over some level which decrease with the increase of the aversion to inequality parameter ε (e.g. for $\varepsilon=0$ $\beta>1$, for $\varepsilon=1$ $\beta>0$, and so on).

As already noted for the Tsui index, comparing multidimensional index (19) and the corresponding unidimensional measure (12), it is possible to notice that they only differ for their denominator. This leads to only small differences between the results of the empirical applications; however, from an axiomatic point of view, the two measures significantly differ: the two-step approach does not necessarily satisfy neither the UM or the CIM criteria, while the multidimensional inequality index does not always satisfy the Pigou-Dalton principle in the space of the individual well-being indices (Decancq et al., 2009).

[3] A special case of (20) is the inequality index obtained by List (1999).

3. INEQUALITY-ADJUSTED WELL-BEING MEASURES

The composite indicators of well-being recently proposed to compare individuals, countries, regions, etc. with respect to their achievements in different dimensions, can be often seen as special cases of the well-being measures implicit in the multidimensional inequality indices reviewed in the previous Section 2. For instance, within the general framework defined by equation (9) it is possible to obtain, as already noted, the Human Development Index, whose three well-being dimensions are aggregated by a weighted arithmetic mean, so that the parameter β is set equal to 1 and dimensions are assumed as perfect substitutes.

The assumption that the dimensions of well-being are perfect substitutes is one of the criticisms that can be raised against HDI. The main one of such conceptual criticism is, however, the absence of any distributional concern in the evaluation of well-being. Differences in individual achievements are ignored, so that two countries (or regions) having the same mean achievements will have the same HDI values even if they have very different distributions of achievements among individuals.

In response to the latter criticism, following the suggestion of Anand and Sen (1997), Hicks (1997) proposes to deflate the mean value of each of the *k* (three) dimensions of the HDI by the factor $(1-G_j)$, where G_j is the Gini coefficient for the dimension z_j, yielding the Sen welfare indicator $S(z_j)=\mu_j\cdot(1-G_j)$, and then aggregating this deflated mean values by an arithmetic mean. The following inequality-adjusted multidimensional index of well-being is so defined:

$$W_G(Z)=\mu[S(z_1),...,S(z_k)]\,. \qquad (22)$$

However, as noted by Foster *et al.* (2005), $W_G(Z)$ does not satisfy some of the desirable properties for a well-being index, because: i) deflating the mean values by a Gini index violates subgroup consistency (or aggregation consistency, see Shorrocks, 1984), so that changes in well-being in a subgroup of the population (e.g. in a region) could be not associated with a corresponding change in the population as a whole; ii) aggregating by an arithmetic mean, as in the HDI, is not sensitive to inequality between dimensions, so that countries (or regions) will have the same HDI values even if they have very different mean achievements among dimensions.

Foster *et al.* (2005) propose a new family of development indices compatible also with the properties of subgroup consistency and sensitiveness to inequality among dimensions. Based on the concept of generalized mean, this family of indices is obtained: i) by deflating the mean values of each dimension by the factor $(1-I_\varepsilon(z_j))$, where $I_\varepsilon(z_j)$ is the Atkinson inequality measure with parameter ε for dimension z_j, that is substituting in the composite indicator of well-being μ_j with $\mu_{1-\varepsilon}(z_j)$, the corresponding mean of order $1-\varepsilon$; ii) by aggregating across dimensions again with a mean of the same order $1-\varepsilon$ instead of a simple arithmetic mean, as in $W_G(Z)$, in order to take account for inequality across dimensions of well-being and to treat them as not completely substitutes, with the degree of complementarity rising as ε rises. So the family of inequality-adjusted development indices proposed by Foster et al. (2005) is the following:

$$W_{\varepsilon}(Z)=\mu_{1-\varepsilon}[\mu_{1-\varepsilon}(z_1),...,\mu_{1-\varepsilon}(z_j)] \quad (\varepsilon \geq 0), \tag{23}$$

where $1-\varepsilon$ can be interpreted as both an inequality aversion parameter and a parameter measuring the degree of substitutability between dimensions. It is worth noting that (23) is a special case of the class of well-being indices implicit in (20) (or in (19)) when $\beta=1-\varepsilon$, the parameters combination that corresponds to the limiting case between more or less inequality being produced by increasing correlation among well-being dimensions (Bourguignon, 1999).

Moreover, setting the aversion parameter $\varepsilon=1$, so that the expression (23) became a geometric mean of geometric means, we obtain the new Inequality-adjusted Human Development Index (IHDI) recently introduced by the United Nations Development Programme (UNDP, 2010) to adjust the traditional HDI for inequality both in distribution of each dimension and across dimensions:

$$W_1(Z)=IHDI=\mu_0[\mu_0(z_1),...,\mu_0(z_K)]. \tag{24}$$

The lack of sensitiveness to correlation among dimensions of the class of indices (24) has been pointed out also by Seth (2009a, 2009b) on the ground that the same parameter ($1-\varepsilon$) is used for both the degree of substitution and inequality aversion, so that the indices are *path independent* and can be obtained also following the "column-first" two step aggregation procedure (Pattanaik et al., 2008), that excludes any compliance with correlation sensitivity[4].

In order to ensure the sensitivity to association among dimensions, Seth (2009a, 2009b) propose a "row-first" strategy of aggregation, first applying a generalized mean of order β, then a generalized mean of different order, say $1-\varepsilon$, to aggregate across persons. The Seth class of well-being indices so assumes the same expression of well-being implicit in (20):

$$W_{1-\varepsilon}(Z)=\mu_{1-\varepsilon}(\mu_{\beta}(\underline{z}_i)). \tag{25}$$

Coherently with the CIM condition ($\beta+\varepsilon>1$) for the inequality index (20), Seth (2009b) shows that if $1-\varepsilon<\beta$ $W_{1-\varepsilon}(Z)$ decreases by increase in association between dimensions.

[4] A similar "column first" two-step approach has been recently proposed also by Herrero et al. (2010).

4. EMPIRICAL APPLICATION TO ITALIAN REGIONS

In this Section we introduce the dataset used in the empirical analysis (4.1) and describe the well-being attributes considered to build up the multidimensional well-being measure, with specific attention devoted to data treatment and transformation procedures necessary to make dimensions comparable. In Section 4.2 we present the measures of inequality in multidimensional well-being by firstly discussing the sensitivity of inequality measurement to alternative normative choices at the country level results and then focusing on outcomes at the regional level. Finally (Section 4.3), we analyse differences in inequality-adjusted multidimensional well-being across Italian regions, and their variation over time.

4.1. *Data and Well-Being Dimensions*

The empirical application is based on the Italian survey on "Statistics on Income and Living Conditions" (IT-SILC), conducted by Istat on yearly basis since 2004, which is part of the "European Statistics on Income and Living Conditions" (EU-SILC) project. For the aims of the present analysis, the use of the IT-SILC dataset allows obtaining consistent regional estimates of multidimensional inequality and well-being indicators, as the sample is representative of the Italian population at the regional (NUTS 2) level.

We use data taken from the 2005 and 2008 cross-sectional waves, which refers to the income levels and living conditions at the end of 2004 and 2007, respectively. We focus on three indicators for the respective dimensions of human well-being: equivalised disposable income, an indicator of health status and years of schooling attained.

Equivalised disposable income (variable HX090 of the IT-SILC dataset) is obtained by dividing total disposable household income, adjusted with the within-household non-response inflation factor, by the equivalised household size. As inequality indicators are known to be potentially sensitive to the presence of extreme incomes in the tails of the distribution (see van Kerm, 2007), we decide not to consider negative and zero incomes and adopt a winsorizing procedure for the lowest 0.25% and the highest 0.1% income observations, replacing those extreme values with the values of trimming thresholds[5].

We use two different indicators as proxies of health status. Firstly, we consider a categorical variable measuring self-assessed health status (variable PH010 of the IT-SILC)[6]. This variable, as every subjective indicator, has the advantage of providing a global assessment of health that is informative for all the population, but its subjectivity may significantly limits interpersonal comparability. Moreover, as pointed out by van Doorslaer and Jones (2002), categorical measures may cause problems for the measurement of health inequalities. For these reasons, we define

[5] This procedure implies modifying 162 and 153 observations in the two surveys.

[6] The original scale of the variable has been inverted so that its level rises as the self-assessed heath status goes to "very bad" (1) to "very good" (5).

a composite cardinal indicator of health status based on the predicted values of an ordinal logit regression of self-assessed health status on three other health-related variables (namely, indicators of chronic illness, physical limitations and unmet health treatments) and socio-demographic characteristics. Complete estimation results for year 2007 are presented in Table A.1 in the Appendix[7]. This procedure imposes cardinality and the constructed measure is as close as possible to the self-reported health status, while ensuring that individuals with the same observed characteristics obtain the same health measure (Decancq and Lugo, 2009).

The third dimension relates to years of schooling and is constructed by combining information on the highest education level attained and the number of years in post-secondary education. As our sample include individuals who are still in education, for these individuals it is necessary to increase the number of years corresponding to the highest level of education completed. For individuals over 25 years old, we assign a value equal to the years of schooling corresponding to the course currently attended. For students under 25 years old, we increase the number of years of schooling by the difference between their age and the years of education corresponding to the highest level completed. The whole distribution is then translated by one unit to avoid the presence of zeros which may cause problems in the computation of inequality measures.

To compute multidimensional inequality indices, four assumptions have to be made, involving: the weighting scheme; the normalization of attributes; the degree of substitutability; and the inequality aversion parameter. With respect to the weighting scheme, we use equal weights for simplicity. We then consider a dimension-specific rescaling and divide the values of each attribute by its respective mean in 2007. It is worth remarking that, while rescaling by the mean values does not modify the distribution of the attributes with respect to the untransformed values, the min-max linear transformation commonly adopted in empirical studies (see Decancq and Lugo, 2010), implying a scaling and a translation operation, modify inequality assessment when translation sensitive indexes are used. Finally, we compare alternative values of the degree of substitutability and of the inequality aversion parameter. We allow the parameter β to vary between 1 (perfect substitutability) and -5 (relatively high complementarity), focusing attention on $\beta = 0$ for which we obtain the Tsui index (16). The parameter ε varies between 0.3 and 3, a range which is commonly adopted in empirical analyses.

Table 1 provides some basic descriptive analysis and shows the evolution of inequality between 2004 and 2007 for the attributes considered[8].

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[7] The predicted values are then linearly transformed so that they ranges from slightly more than 0 (most unhealthy individual) to 1 (the healthiest individual).

[8] Individual sample weights (IT-SILC variable PB040) are used in all the computations presented.

TABLE 1

Descriptive statistics and univariate inequality measures

	2004 (N = 47079)			2007 (N = 43692)		
	Income	Health Status	Education	Income	Health Status	Education
Mean	16952.01	0.6788	11.1886	18102.13	0.6700	11.5165
Std. Dev.	11757.44	0.1773	5.1594	11170.28	0.1868	5.1889
Min	686.67	0.0262	1	666.67	0.0204	1
Max	141883.3	1	23	122900.0	1	23
Gini	0.3168	0.1402	0.2578	0.3009	0.1506	0.2522
	(0.0019)	(0.0009)	(0.001)	(0.0016)	(0.0009)	(0.001)
ATK0.5	0.0854	0.0223	0.0642	0.0760	0.0256	0.0619
	(0.0011)	(0.0003)	(0.0006)	(0.0009)	(0.0003)	(0.0006)
ATK1	0.1657	0.0502	0.1450	0.1504	0.0577	0.1405
	(0.0020)	(0.0007)	(0.0016)	(0.0016)	(0.0008)	(0.0016)
ATK2	0.3456	0.1331	0.3866	0.3240	0.1514	0.3796
	(0.0056)	(0.0030)	(0.0040)	(0.0046)	(0.0024)	(0.0043)
ATK3	0.5865	0.3301	0.6371	0.5732	0.3018	0.6349
	(0.0108)	(0.0456)	(0.0033)	(0.0094)	(0.0089)	(0.0036)

Note: bootstrapped (500 replications) standard errors in parentheses.
Source: authors' elaboration on 2005 and 2008 IT-SILC data.

Analysing the distribution of the variables, it is possible to note that income and to a less extent education attainments are characterized by decreasing inequality over the period of analysis for all the inequality measures. On the other hand, health status shows higher inequality in 2007 than in 2004, with a reversal of the ranking of the two distributions when the Atkinson index with higher inequality aversion is considered. Comparing the distribution of each dimension in the two years, inequality of health status is significantly lower than that of income and education, irrespective of the index considered. Moreover, focusing on Atkinson measures, as the weight attached to the lowest part of the distribution increases (i.e. for $\varepsilon = 2$ and $\varepsilon = 3$) inequality in education attainments becomes higher than that of income both in 2004 and 2007: this is due to the presence of a still significant fraction of individuals without any educational attainment. This evidence is further confirmed when the analysis of univariate inequality is carried out at the regional level. As it is shown in Table A.2 in the Appendix, inequality in education levels becomes noticeably higher than that of income as ε rises, especially in the Southern regions.

These diverging patterns suggest the difficulty of drawing a univocal picture on overall inequality by means of an item-by-item approach, and the opportunity of summarizing the multiple dimensions of individuals' well-being into a synthetic indicator.

In Table 2 we present the correlation structure between the dimensions of well-being. As it can be noted, correlation coefficients between dimensions are all significant at the 1% level. Correlation is particularly high between health and education, while it is relatively low between income and the other two dimensions (equal to 0.1515 and 0.297 in 2007), thus justifying the inclusion of non-monetary dimensions, together with disposable income, in the assessment of individual well-being. All the correlation coefficients show an increasing pattern over the period considered that will lead, *ceteris paribus*, to an increase in inequality measured by means of multidimensional indexes satisfying the correlation increasing majorization principle.

TABLE 2

Correlation structure: pair-wise correlation coefficients

2004			
Variable	Income	Health Status	Education
Income	1		
Health Status	0.1454***	1	
Education	0.2815***	0.5489***	1
2007			
Variable	Income	Health Status	Education
Income	1		
Health Status	0.1515***	1	
Education	0.2970***	0.5601***	1

Note: *** denotes significance at the 1% level.
Source: authors' elaboration on 2005 and 2008 IT-SILC data.

4.2. *Multidimensional well-being and inequality in Italian regions*

In order to measure multidimensional well-being inequality and to analyse its pattern over time, we start by computing the inequality index (19) at the national level. In Figure 1, we present the values of the inequality index as a function of β and ε parameters. The two panels of the Figure show the surface and the contour plots of I^M for the distribution of the three attributes considered in 2007, with β and ε values ranging from -5 to 1 and from 0 and 3, respectively. Inequality increases as the parameter of inequality aversion rises and dimensions are treated as complements: obviously, inequality is equal to zero when the dimensions are treated as perfect substitutes and there is no social aversion to inequality ($\beta = 1$ and $\varepsilon = 0$) and reaches its maximum (equal to 0.65) when the parameters combination $\beta = -5$ and $\varepsilon = 3$ is considered (i.e. assuming complementarity between dimensions and high aversion to inequality). Moreover, the contour plot in panel b) shows that I^M becomes particularly sensitive to changes in the correlation structure between attributes when high aversion to well-being inequality is assumed. In particular, when $\varepsilon > 2$ and we consider parameter combinations for which the CIM axiom is satisfied ($\beta + \varepsilon > 1$), it is possible to note that as β goes from 0 to -2 inequality tends to move to higher contours: in order to remain on the same iso-inequality curve, a strong decrease in inequality aversion is required to compensate the effect of a small increase in the degree of complementarity between attributes.

The above mentioned inequality pattern for different β and ε combinations emerges also from Table 3, panel a), where it is worth remarking that all the measures are statistically significant at the 1% level, based on bootstrapped standard errors. This further confirms the opportunity of carrying out a detailed sensitivity analysis of the multidimensional index considered, as different normative choices lead to a significantly different picture of overall inequality in well-being.

More interestingly, as it can be noticed from panel b) of the Table, multidimensional inequality shows an overall tendency to decrease over the period of analysis, due to the already discussed reduction of inequality in income and education levels, which dominates the increase in health inequality. However, differences are statistically significant only for the parameters combinations that do not *a) Surface plot* *c) Contour plot*

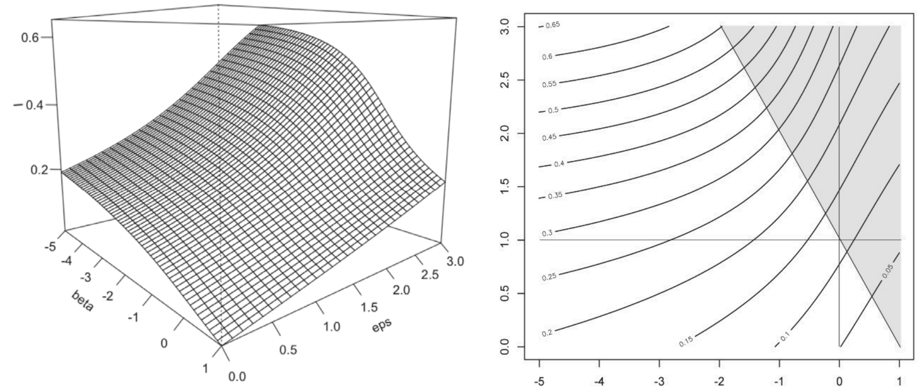

Figure 1 – Multidimensional inequality as a function of β and ε parameters
Note: the parameter combinations for which the multidimensional inequality index in (19) satisfies the CIM axiom are highlighted in grey.
Source: authors' elaboration on 2008 IT-SILC data.

TABLE 3

Multidimensional inequality measures: different β and ε parameters combinations

a) 2007 values

β	ε 0.3	1	1.5	2	2.5	3
-5	0.2193	0.2973	0.3715	0.4636	0.5625	0.6504
	(0.0019)	(0.0023)	(0.0029)	(0.0036)	(0.0039)	(0.0039)
-2	0.1569	0.2270	0.2924	0.3733	0.4645	0.5534
	(0.0019)	(0.0024)	(0.0028)	(0.0035)	(0.0041)	(0.0044)
-1	0.1201	0.1805	0.2339	0.2973	0.3693	0.4443
	(0.0018)	(0.0022)	(0.0025)	(0.0031)	(0.0037)	(0.0041)
0	0.0701	0.1176	0.1546	0.1958	0.2409	0.2893
	(0.0018)	(0.0019)	(0.0021)	(0.0023)	(0.0028)	(0.0034)
0.5	0.0427	0.0855	0.1185	0.1537	0.1916	0.2321
	(0.0017)	(0.0018)	(0.0019)	(0.0022)	(0.0025)	(0.0031)
1	0.0170	0.0579	0.0884	0.1205	0.1545	0.1906
	(0.0017)	(0.0019)	(0.0019)	(0.0021)	(0.0024)	(0.0029)

b) Differences with respect to 2004

β	ε 0.3	1	1.5	2	2.5	3
-5	-0.0122‡	-0.0133‡	-0.0148‡	-0.0144‡	-0.0130†	-0.0122
	(0.0028)	(0.0034)	(0.0042)	(0.0053)	(0.0063)	(0.007)
-2	-0.0106‡	-0.0119‡	-0.0128‡	-0.0129‡	-0.0126	-0.0130
	(0.0028)	(0.0034)	(0.0041)	(0.0053)	(0.0066)	(0.0079)
-1	-0.0084‡	-0.0097‡	-0.0095‡	-0.0094†	-0.0091	-0.0098
	(0.0027)	(0.0031)	(0.0037)	(0.0045)	(0.0056)	(0.007)
0	-0.0045	-0.0042	-0.0041	-0.0031	-0.0016	0.0003
	(0.0025)	(0.0027)	(0.0029)	(0.0033)	(0.0038)	(0.0045)
0.5	-0.0023	-0.0033	-0.0017	-0.0006	0.0010	0.0032
	(0.0025)	(0.0026)	(0.0027)	(0.0029)	(0.0034)	(0.0041)
1	-0.0004	-0.0011	0.0000	0.0011	0.0025	0.0046
	(0.0025)	(0.0026)	(0.0026)	(0.0028)	(0.0032)	(0.0038)

Note: bootstrapped (500 replications) standard errors in parentheses.
In panel b) of the Table, daggers ‡ and † denote significance at the 1% and 5% levels, respectively.
The parameter combinations for which the multidimensional inequality index satisfies the CIM axiom are highlighted in grey.
Source: authors' elaboration on 2005 and 2008 IT-SILC data.

satisfy the CIM axiom (i.e. $\beta + \varepsilon > 1$), characterized by low inequality aversion and moderately high complementarity between attributes. In this region of the parameters space, in which less inequality is produced by more correlation as pointed out by Bourguignon (1999), the increasing correlation between dimensions from 2004 to 2007 further enhances the diminishing effect of the overall drop in unidimensional inequalities. The statistical significance of changes in inequality tends to disappear as we move to the region satisfying the CIM condition and assume substitutability and high ε values. The difference in inequality levels between 2004 and 2007 becomes even positive, but not statistically significant when $\beta + \varepsilon$ is much larger than 1: in these cases the increase in correlation between attributes offsets the effect of the decreasing inequality in each dimension.

In Table 4 we move to the sub-national analysis, focusing on the analysis of the inequality adjusted well-being levels defined in equation (25) at the NUTS 2 level, for all the twenty Italian regions, with Trentino Alto Adige split into the two autonomous provinces of Bolzano and Trento, in year 2007. The regional well-being indices have been normalized by the corresponding index at the national level and then multiplied by 100, so that $W_{1-\varepsilon}^{Italy}(Z) = 100$.

TABLE 4

Regional multidimensional well-being measures (year 2007)

	Tsui index $\beta = 0$				ε=2				
	$\varepsilon = 1$	$\varepsilon = 1.5$	$\varepsilon = 2$	$\varepsilon = 2.5$	β=−2	β=−1	β=1	ε=2.5, β=−1	ε=1, β=1
Region	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Piemonte	104.7	105.5	106.7	108.1	114.3	110.8	104.1	113.3	103.1
Valle d'Aosta	106.3	107.5	109.1	110.8	118.0	113.9	106.3	116.6	104.5
Lombardia	108.3	109.4	110.9	112.7	119.8	115.8	108.0	119.6	106.6
Bozen-Bolzano	115.6	117.3	119.5	122.3	133.1	126.9	115.4	132.9	113.3
Trento	110.8	112.9	115.8	119.4	134.6	125.9	110.2	134.6	107.5
Veneto	104.7	105.7	106.9	108.5	116.3	112.1	104.1	115.8	102.8
Friuli-Venezia Giulia	106.2	107.4	109.1	111.1	118.2	114.1	106.0	117.8	104.3
Liguria	104.7	105.6	106.7	107.7	113.0	110.3	104.4	112.3	103.4
Emilia-Romagna	105.5	105.8	106.4	107.3	107.6	107.1	105.6	107.8	105.1
Toscana	106.0	106.7	107.6	108.7	113.0	110.6	105.7	112.5	104.9
Umbria	101.3	101.4	101.6	101.9	101.7	101.6	101.7	101.7	101.2
Marche	101.6	102.0	102.6	103.3	105.2	104.0	101.7	105.1	101.1
Lazio	104.6	104.5	104.2	103.7	106.8	105.7	103.7	105.7	104.3
Abruzzo	94.6	93.6	92.6	91.7	87.9	89.8	94.7	88.5	95.5
Molise	90.3	89.7	89.2	88.8	88.4	88.9	89.7	89.4	90.7
Campania	91.5	90.7	89.9	89.1	81.9	85.1	93.0	83.5	93.7
Puglia	90.5	90.0	89.6	89.3	81.7	85.0	92.2	84.0	92.2
Basilicata	87.2	85.5	83.8	82.2	75.9	79.1	87.1	77.2	89.3
Calabria	84.0	82.3	80.4	78.4	71.3	74.8	85.0	72.6	87.1
Sicilia	87.6	87.0	86.6	86.2	81.1	83.4	89.1	83.3	89.5
Sardegna	93.8	93.6	93.5	93.5	90.5	91.7	94.8	91.7	94.7
Italia	100	100	100	100	100	100	100	100	100

Source: authors' elaboration on 2008 IT-SILC data.

The alternative combinations of β and ε parameters considered leave the rankings of the regions with respect to the multidimensional well-being indices almost unchanged. As it can be noticed, the regions of the Centre and North of Italy are characterized by significantly higher well-being levels than the Southern regions.

In particular, the autonomous provinces of Bolzano and Trento, followed by Lombardia, remain at the top of the rankings, irrespective of the parameters choices. These evidences not only depend on the higher average levels of the well-being dimensions considered, but also on the sensibly lower levels of multidimensional well-being inequality in the Central and Northern regions, as it can be noted in Table A.3 in the Appendix, that are related to the more equal unidimensional distribution of each well-being component (as reported in Table A.2 in the Appendix). On the other hand, Abruzzo and Sardegna always perform better than the other regions of the South, due to the lower multidimensional inequality levels (see Table A.3) mainly connected to the significantly lower inequality in educational attainments (see Table A.2) of these two regions.

Deepening the sensitivity analysis of multidimensional well-being with respect to different parameters combinations, it is possible to pick out some regularities in the patterns of the regional gaps. Firstly, due to the more homogeneous distribution of well-being both across individuals and dimensions, the positive divide of the Central and Northern regions with respect to the national level generally tends to widen as inequality aversion increases and as dimensions are treated as more and more complements, with the exceptions of Emilia-Romagna and Umbria which remain substantially stable. The opposite happens in the Southern regions, whose well-being indicators worsen as ε rises and β becomes negative. These regions, in fact, not only suffer from higher inequality in the interpersonal distribution of the attributes, but also from a lack of homogeneity in the average attainments of each dimension, which significantly reduces multidimensional well-being when more weight is given to the bottom of the distribution and when the worse-off dimension receives more weight in the SWF.

In Figure 2 we provide a graphical representation of the regional well-being scores, together with the corresponding 95% bootstrapped confidence interval, focusing attention on the well-being measures implied by the Tsui index (21), assuming low ($\varepsilon = 1$) and relatively high ($\varepsilon = 2.5$) inequality aversion.

The Figure confirms the existence of a clear-cut distance between Central-Northern and Southern regions in terms of multidimensional well-being levels, irrespective of the different preferences for redistribution assumed, even when the sample variability of the estimates is taken into account. Again, Bolzano and Trento show the highest well-being, despite the indices are characterized by a significant variability and wide confidence intervals due to the small size of the samples in the two provinces. Among the regions of the South, Calabria displays the worst well-being performances, especially when we assume a high degree of inequality aversion. From the analysis of the Figure it is also possible to note that the sample variability of the estimates tends to increase with the ε parameter. One of the key messages emerging form Figure 2 is that confidence intervals can be a valuable instrument for assessing robustness of well-being rankings, despite they are not commonly used in empirical works, as pointed out by Ravallion (2010). Høyland et al. (2011) show that welfare rankings can be highly sensitive when uncertainty is taken into account, though less so at the extremes. Our results confirm this evidence: including sampling variability we are still able to unambigu-

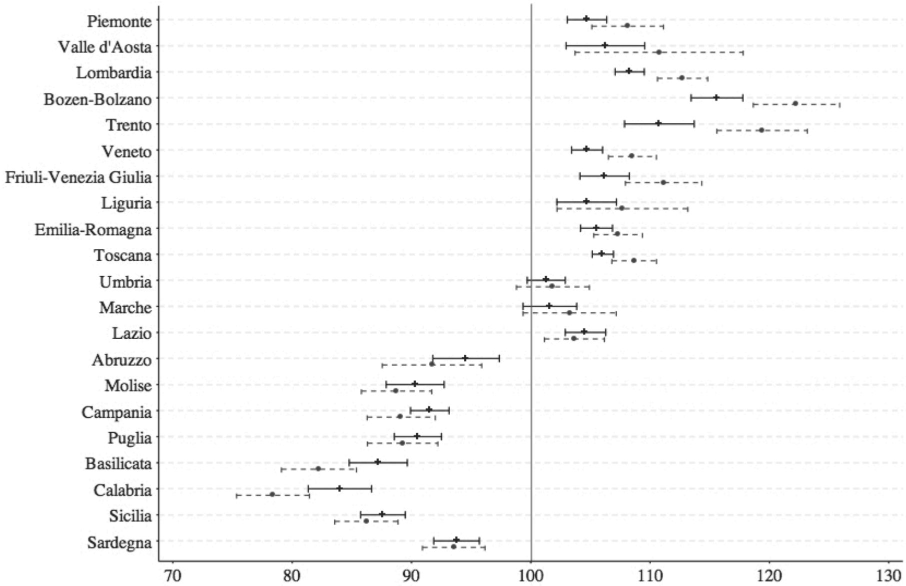

Figure 2 – Inequality adjusted multidimensional well-being comparisons
Note: each bar indicates the 95% confidence interval for the welfare measures (normalized so that Italy=100) implied by the Tsui inequality indices with $\varepsilon=1$ (continuous line) and $\varepsilon=2.5$ (dotted line).
Source: authors' elaboration on 2008 IT-SILC data.

ously distinguish the best regions from the worst and to highlight a clear dualism between Northern and Southern regions. However, it is particularly difficult to univocally rank all the regions, as the confidence intervals of the well-being indices often overlap[9].

In Table 5 we present the percentage changes of regional well-being indices between 2004 and 2007, together with indications on their statistical significance.

From the analysis of the Table, it clearly emerges an overall tendency of well-being to increase over the period considered, both at the aggregate and at the regional levels. The regions of the Centre and North of Italy are characterized by the highest (and statistically significant) growth rates, showing an overall tendency towards regional divergence in multidimensional well-being. Among Southern regions, Molise and Sicilia represent an exception to the general trend, showing a significant improvement over the period, with increasing growth rates as ε rises and as β becomes more and more negative. This evidence characterizes almost all the regions with the best growth performances (together with Molise and Sicilia, Marche, Emilia Romagna and, to a less extent, Trento, Veneto, Toscana and Liguria), for which it is possible to observe that well-being increases at higher rates

[9] This point is clearly explained by Ravallion (2010), who asserts that the use of composite development indices, such as the HDI, would be more appropriate "to try to identify a few reasonably robust country groupings than these seemingly precise but actually rather uncertain country rankings".

TABLE 5

Regional multidimensional well-being measures: percentage changes 2004-2007

	Tsui index $\beta = 0$				ε=2				
	$\varepsilon = 1$	$\varepsilon = 1.5$	$\varepsilon = 2$	$\varepsilon = 2.5$	β=–2	β=–1	β=1	ε=2.5, β=–1	ε=1, β=1
Region	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Piemonte	1.28	1.07	0.85	0.66	-0.83	-0.12	1.06	-1.08	1.32
Valle d'Aosta	2.82	2.72	2.38	1.68	2.52	2.21	2.38	-0.04	2.29
Lombardia	2.30‡	2.36‡	2.37‡	2.31	2.88	2.59	2.31‡	2.56	2.17‡
Bozen-Bolzano	6.36‡	6.16‡	5.74‡	5.04‡	5.05	5.38†	6.06‡	4.04	6.14‡
Trento	4.21†	4.25†	4.32†	4.44	5.73	5.11	4.13†	6.26	4.14‡
Veneto	5.77‡	5.73‡	5.65‡	5.54‡	7.60‡	6.65‡	5.09‡	6.63‡	5.13‡
Friuli-Venezia Giulia	0.71	0.31	-0.25	-0.98	-2.44	-1.50	0.33	-3.39	0.84
Liguria	4.61‡	4.60‡	4.30	3.48	7.10	5.84	3.59	5.44	3.89‡
Emilia-Romagna	3.83‡	4.17‡	4.61‡	5.13‡	5.35†	4.99‡	4.09‡	5.52†	3.64‡
Toscana	3.69‡	3.68‡	3.73‡	3.87‡	5.04†	4.44‡	3.35‡	4.86†	3.64‡
Umbria	2.50	2.34	2.07	1.60	0.21	0.85	2.66	-0.57	2.78†
Marche	5.90‡	6.41‡	7.00‡	7.61‡	14.10‡	11.00‡	5.34‡	13.40‡	5.13‡
Lazio	1.01	0.83	0.45	-0.28	2.06	1.34	0.36	1.33	0.84
Abruzzo	1.45	1.28	1.21	1.28	0.16	0.56	1.41	0.28	1.32
Molise	4.45†	4.75†	5.19†	5.79†	13.75‡	10.27‡	2.19	11.92‡	2.29
Campania	4.38	3.88	3.15	2.20	4.25	3.74	2.32	2.03	3.27‡
Puglia	3.69	3.45	3.19	2.94	4.81	4.23	2.04	3.86	2.41
Basilicata	3.57	2.61	1.53	0.42	1.67	1.46	1.36	-0.14	2.90
Calabria	1.27	0.65	-0.04	-0.58	3.07	1.75	0.05	3.70	0.96
Sicilia	6.70‡	6.98‡	7.26‡	7.53‡	16.08‡	12.55‡	4.75‡	14.12‡	4.38‡
Sardegna	1.49	1.71	2.02	2.40	1.13	1.42	1.88	1.40	1.44
Italia	3.35‡	3.36‡	3.26‡	3.08‡	5.15‡	4.31‡	2.68‡	4.43‡	2.85‡

Note: Daggers ‡ and † denote significance at the 1% and 5% significance levels, respectively (based on bootstrapped standard errors).
Source: authors' elaboration on 2005 and 2008 IT-SILC data.

when we assume high inequality aversion and dimensions are assumed as complements. In these regions well-being growth appears to be related to significant improvements in the equity of the interpersonal welfare distribution and, especially, in the homogeneity of the attainments in each dimension, with a recovery in the worse-off dimension of well-being. On the other hand, in the regions with the worst growth performances the opposite pattern can be picked out: the growth rates further shrink as ε increase and as β becomes negative. In particular, the low growth rates of these regions are related either to a worsening in the homogeneity of the distribution of well-being both across individuals and dimensions (this is the case of Piemonte, Friuli-Venezia Giulia and Umbria where, for some parameters combinations, growth rates become even negative) or to an overall stability in multidimensional inequality with respect to alternative values of ε and β (as in Lombardia, Abruzzo and Sardegna).

5. CONCLUDING REMARKS

In this paper, focusing mainly on the normative approach, the more commonly used multidimensional inequality measures and inequality adjusted multidimensional well-being indicators are reviewed. In particular, we show that such well-being indicators can be seen as special cases of the well-being measures implicit in multidimensional inequality indices.

Using Italian data on individual income, education and health status from the 2005 and 2008 Italian Survey on Income and Living Conditions (IT-SILC), an empirical analysis of multidimensional inequality and inequality adjusted well-being levels in Italian regions has been performed. Given the variability that characterizes inequality and well-being indicators, depending on the uncertainty connected to the survey nature of the data as well as on the alternative parameter combinations chosen, the regional well-being scores are presented together with the corresponding confidence intervals, based on bootstrapped standard errors. To the best of our knowledge, this is one of the first analyses of multidimensional inequality and well-being performed at the regional level and exploits the fact the IT-SILC sample assures estimates' consistency at the NUTS 2 level.

Two main results can be outlined. Firstly, including sampling variability we are still able to highlight a clear dualism between Central-Northern and Southern regions, as well as to assess the highest well-being levels of some Northern regions (Bolzano, Trento, Lombardia) and the worst performances of some Southern regions (Calabria). However, it is not possible to univocally rank all the regions, as the confidence intervals of the well-being indices often overlap. Secondly, distributional concerns across individuals and across dimensions significantly affect both regional discrepancies and growth rate. More specifically, increasing inequality aversion and decreasing substitutability between attributes progressively widen the regional well-being differences, revealing a more homogeneous distribution of well-being both across individuals and dimensions in the Central-Northern regions with respect to the Southern ones. The same pattern emerges from the analysis of the evolution of regional well-being over time, revealing that high growth rates are related with improvements in the equity of interpersonal well-being distribution and, especially, with a recovery in its worse-off dimensions.

These evidences confirm that the investigation of a multiplicity of normative assumptions, regarding the degree of substitutability between dimensions and inequality aversion, enriches the multidimensional analysis of well-being. The lack of uniqueness in the results, which is a typical feature of multidimensional analyses of inequality and well-being, far from being a weak aspect of the approach, deepens the informative value allowing to reveal how the two distributional concerns affect both regional gaps and growth patterns of well-being.

Comparative analysis of well-being across regions needs further research on data quality. We mainly refer to two sources of inaccuracy in the measurement of well-being dimensions based on survey data. The first one is that, due to underreporting, household surveys generally provide underestimated figures on disposable income. Besides the possible different rates of underestimation among regions, underreporting affects comparative inequality-adjusted well-being levels also through the inequality component of the indicator, as there is no reason to assume that such bias is distribution neutral (Ravallion, 2000). Anchoring the disposable income to national (regional) accounts and, most of all, taking into account the distributional effects of underreporting is a main topic for further research. The second source of inaccuracy concerns the quality aspects of education, as years of schooling is becoming an increasingly poor measure of educational attainments (Cipollone, 2010). It would make more sense to use outcome indicators of the education systems (Kovacevic, 2010), or at least to correct the years of schooling with some measure of outcome quality, such as those provided by the international PISA survey. This is another topic for future research to improve the accuracy of regional multidimensional inequality and well-being analyses.

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SUMMARY

Measuring multidimensional inequality and well-being: methods and an empirical application to Italian regions

In this paper, focusing on the normative approach, we review and discuss the main multidimensional inequality measures and inequality-adjusted multidimensional well-being indicators.

Using Italian data on individual income, education and health status from the 2005 and 2008 Italian Survey on Income and Living Conditions (IT-SILC), an empirical analysis of multidimensional inequality and inequality-adjusted well-being levels in Italian regions has been performed. Given the variability that characterizes inequality and well-being indicators, depending on the uncertainty connected to the survey nature of the data as well as on the alternative parameter combinations chosen, the regional indices are presented together with the corresponding confidence intervals, as an instrument for assessing the robustness of well-being rankings.

A significant result of the analysis is that distributional concerns, both across individuals and between dimensions, remarkably affects discrepancies in regional well-being. More specifically, increasing inequality aversion and decreasing substitutability between attributes progressively widen regional well-being differences. The same pattern emerges from the analysis of the evolution of regional well-being over time, revealing that high growth rates are related with improvements in the equity of interpersonal well-being distribution and, especially, with a recovery in its worse-off dimensions.

APPENDIX

TABLE A1

Ordered logit estimation of self-assessed health status (year 2007)

Dependent variable: Self- assessed health status		Coefficient	Standard Error
Eq. Income (in logs)		0.1551***	(0.0180)
Student		-0.8360***	(0.0854)
Retired		-0.0602*	(0.0325)
Years of education		0.1444***	(0.0069)
Widowed		-0.0206	(0.0392)
Divorced		-0.1551**	(0.0754)
Single		0.1466***	(0.0291)
Male		0.1766***	(0.0205)
Housewife		0.5635***	(0.0444)
Chronic illness		-1.4081***	(0.0303)
Physical limitations		-1.2093***	(0.0175)
Unmet health treatment		-0.7114***	(0.0392)
Age 35-44		-0.5294***	(0.0358)
Age 45-54		-1.0804***	(0.0388)
Age 55-64		-1.5545***	(0.0433)
Age 65-69		-1.8264***	(0.0547)
Age 70-74		-2.0526***	(0.0573)
Age 75-79		-2.1551***	(0.0618)
Age over 80		-2.1102***	(0.0628)
Region 1		-0.0323	(0.0879)
Region 3		0.0801	(0.0837)
Region 4		0.9545***	(0.1046)
Region 5		0.2865***	(0.1072)
Region 6		0.1217	(0.0859)
Region 7		0.1746*	(0.0919)
Region 8		0.4432***	(0.0922)
Region 9		0.2274***	(0.0861)
Region 10		0.2632***	(0.0863)
Region 11		0.2714***	(0.0905)
Region 12		0.0752	(0.0892)
Region 13		0.0316	(0.0862)
Region 14		-0.0451	(0.1001)
Region 15		0.0276	(0.1040)
Region 16		0.5754***	(0.0872)
Region 17		0.2722***	(0.0892)
Region 18		0.1665*	(0.1002)
Region 19		-0.3000***	(0.0957)
Region 20		0.2759***	(0.0890)
Region 21		-0.0463	(0.0973)
N	43692		
Pseudo-R^2	0.2924		

Note: ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.

TABLE A.2

Univariate inequality measures by region (year 2007)

Region	N. Obs.	Income			Health Status			Education		
		Gini	ATK1	ATK2	Gini	ATK1	ATK2	Gini	ATK1	ATK2
Piemonte	2597	0.288	0.135	0.291	0.146	0.053	0.140	0.238	0.109	0.280
Valle d'Aosta	674	0.264	0.116	0.248	0.134	0.048	0.132	0.234	0.107	0.277
Lombardia	4604	0.284	0.134	0.281	0.135	0.046	0.119	0.232	0.106	0.275
Bozen-Bolzano	908	0.275	0.124	0.251	0.120	0.036	0.087	0.220	0.092	0.236
Trento	767	0.245	0.095	0.192	0.129	0.039	0.096	0.221	0.084	0.190
Veneto	3269	0.261	0.113	0.241	0.147	0.054	0.137	0.236	0.110	0.286
Friuli-VeneziaGiulia	1801	0.259	0.113	0.257	0.148	0.053	0.138	0.230	0.102	0.260
Liguria	1748	0.283	0.133	0.296	0.140	0.047	0.120	0.235	0.111	0.290
Emilia-Romagna	3166	0.285	0.135	0.292	0.154	0.056	0.138	0.255	0.140	0.374
Toscana	3107	0.274	0.125	0.275	0.146	0.053	0.134	0.247	0.126	0.327
Umbria	2015	0.275	0.129	0.287	0.163	0.065	0.162	0.254	0.146	0.393
Marche	2259	0.276	0.128	0.287	0.159	0.063	0.161	0.251	0.132	0.348
Lazio	3233	0.304	0.151	0.318	0.154	0.062	0.162	0.233	0.129	0.362
Abruzzo	1057	0.278	0.126	0.260	0.161	0.067	0.180	0.282	0.196	0.502
Molise	902	0.294	0.138	0.271	0.167	0.066	0.166	0.276	0.173	0.448
Campania	3017	0.316	0.168	0.371	0.145	0.057	0.148	0.270	0.171	0.453
Puglia	2330	0.287	0.145	0.328	0.145	0.055	0.145	0.281	0.186	0.477
Basilicata	1072	0.310	0.159	0.347	0.174	0.074	0.190	0.293	0.211	0.527
Calabria	1401	0.307	0.157	0.341	0.187	0.097	0.280	0.290	0.213	0.534
Sicilia	2471	0.308	0.153	0.315	0.158	0.065	0.173	0.276	0.178	0.461
Sardegna	1294	0.292	0.141	0.309	0.169	0.073	0.186	0.258	0.156	0.421
Italy	43692	0.301	0.150	0.324	0.151	0.058	0.151	0.252	0.141	0.380

Source: authors' elaboration on 2008 IT-SILC data.

TABLE A3

Regional multidimensional inequality measures (year 2007)

	Tsui index $\beta = 0$				ε=2				
	$\varepsilon = 1$	$\varepsilon = 1.5$	$\varepsilon = 2$	$\varepsilon = 2.5$	β=−2	β=−1	β=1	ε=2.5, β=−1	ε=1, β=1
Region	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Piemonte	0.100	0.131	0.164	0.200	0.300	0.240	0.109	0.303	0.054
	(0.007)	(0.008)	(0.009)	(0.011)	(0.015)	(0.012)	(0.008)	(0.017)	(0.007)
Valle d'Aosta	0.091	0.119	0.150	0.185	0.280	0.223	0.095	0.286	0.047
	(0.014)	(0.017)	(0.020)	(0.026)	(0.032)	(0.028)	(0.016)	(0.038)	(0.013)
Lombardia	0.096	0.125	0.157	0.191	0.287	0.229	0.103	0.285	0.051
	(0.005)	(0.006)	(0.006)	(0.008)	(0.011)	(0.009)	(0.006)	(0.011)	(0.005)
Bozen-Bolzano	0.085	0.110	0.138	0.167	0.248	0.197	0.092	0.246	0.045
	(0.009)	(0.009)	(0.011)	(0.013)	(0.017)	(0.014)	(0.010)	(0.019)	(0.010)
Trento	0.073	0.095	0.117	0.141	0.199	0.161	0.082	0.194	0.040
	(0.012)	(0.013)	(0.013)	(0.014)	(0.018)	(0.015)	(0.013)	(0.019)	(0.013)
Veneto	0.093	0.123	0.155	0.191	0.284	0.226	0.102	0.283	0.049
	(0.006)	(0.006)	(0.007)	(0.008)	(0.008)	(0.007)	(0.006)	(0.009)	(0.005)
Friuli-Venezia Giulia	0.090	0.118	0.148	0.181	0.280	0.221	0.095	0.278	0.047
	(0.009)	(0.010)	(0.011)	(0.012)	(0.016)	(0.014)	(0.009)	(0.018)	(0.009)
Liguria	0.098	0.128	0.162	0.202	0.308	0.243	0.104	0.308	0.049
	(0.011)	(0.012)	(0.015)	(0.021)	(0.018)	(0.016)	(0.015)	(0.022)	(0.012)
Emilia-Romagna	0.111	0.146	0.183	0.222	0.353	0.280	0.116	0.349	0.057
	(0.006)	(0.006)	(0.007)	(0.008)	(0.013)	(0.011)	(0.005)	(0.013)	(0.005)
Toscana	0.102	0.134	0.170	0.208	0.319	0.253	0.109	0.318	0.053
	(0.004)	(0.004)	(0.006)	(0.007)	(0.011)	(0.009)	(0.004)	(0.012)	(0.004)
Umbria	0.114	0.150	0.190	0.234	0.368	0.292	0.114	0.364	0.055
	(0.007)	(0.008)	(0.009)	(0.012)	(0.013)	(0.011)	(0.009)	(0.013)	(0.007)
Marche	0.108	0.142	0.180	0.220	0.343	0.273	0.111	0.340	0.053
	(0.010)	(0.012)	(0.013)	(0.015)	(0.020)	(0.017)	(0.011)	(0.020)	(0.009)
Lazio	0.115	0.153	0.196	0.245	0.357	0.287	0.126	0.360	0.059
	(0.007)	(0.008)	(0.008)	(0.009)	(0.012)	(0.010)	(0.008)	(0.012)	(0.007)
Abruzzo	0.131	0.176	0.224	0.275	0.426	0.343	0.133	0.419	0.064
	(0.013)	(0.014)	(0.016)	(0.017)	(0.021)	(0.019)	(0.013)	(0.020)	(0.011)
Molise	0.127	0.169	0.214	0.262	0.388	0.313	0.138	0.380	0.067
	(0.012)	(0.012)	(0.012)	(0.013)	(0.020)	(0.017)	(0.011)	(0.018)	(0.011)
Campania	0.134	0.177	0.224	0.274	0.439	0.353	0.131	0.430	0.062
	(0.008)	(0.009)	(0.010)	(0.012)	(0.014)	(0.014)	(0.009)	(0.015)	(0.007)
Puglia	0.130	0.171	0.216	0.262	0.436	0.346	0.122	0.420	0.059
	(0.010)	(0.011)	(0.012)	(0.012)	(0.014)	(0.013)	(0.010)	(0.014)	(0.009)
Basilicata	0.150	0.201	0.255	0.310	0.470	0.383	0.158	0.460	0.075
	(0.012)	(0.012)	(0.013)	(0.013)	(0.016)	(0.015)	(0.011)	(0.015)	(0.011)
Calabria	0.157	0.209	0.265	0.323	0.486	0.399	0.154	0.476	0.072
	(0.014)	(0.014)	(0.014)	(0.013)	(0.014)	(0.014)	(0.013)	(0.013)	(0.013)
Sicilia	0.133	0.175	0.220	0.266	0.421	0.337	0.129	0.406	0.062
	(0.009)	(0.010)	(0.011)	(0.012)	(0.013)	(0.012)	(0.010)	(0.013)	(0.009)
Sardegna	0.124	0.162	0.204	0.248	0.399	0.318	0.118	0.388	0.056
	(0.009)	(0.010)	(0.010)	(0.011)	(0.020)	(0.016)	(0.008)	(0.018)	(0.007)
Italia	0.118	0.155	0.196	0.241	0.373	0.297	0.121	0.369	0.058
	(0.002)	(0.002)	(0.002)	(0.003)	(0.004)	(0.003)	(0.002)	(0.004)	(0.002)

Note: bootstrapped (500 replications) standard errors in parentheses.
Source: authors' elaboration on 2008 IT-SILC data.