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**WELL-BEING DURING THE GREAT RECESSION:
NEW EVIDENCE FROM A MEASURE OF MULTI-DIMENSIONAL LIVING
STANDARDS WITH HETEROGENEOUS PREFERENCES**

Romina Boarini, Marc Fleurbaey, Fabrice Murtin and Paul Schreyer¹

ABSTRACT

We compute a distribution-adjusted welfare measure that aggregates income, unemployment and longevity for 28 OECD countries since the late-2000s, while using individual weights that reflect people's heterogeneous preferences. The estimated shadow prices of one percentage point of unemployment and one year of longevity average 3.1% and 5.7% of income respectively, while varying across countries and individuals. The study reveals that the rate of economic growth poorly reflects the social cost of the Great Recession, especially in hard-hit countries and for poorer households. On average across OECD countries, GDP per capita stagnated with a 0.0% annual change while the living standards of poor households fell by 5.3% annually between 2008 and 2013.

Keywords : Welfare, well-being, living standards, beyond GDP, shadow prices, value of statistical life, measurement, indicator.

JEL codes: I31, I32, J17, J18, I38, I18.

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1. Introduction

Stiglitz, Sen and Fitoussi (2009) make a strong case to complement well-established measures of economic well-being such as income or consumption with dimensions that capture the quality of life, for example health, education or social relations. This can be achieved by elaborating dashboards of indicators as in OECD (2011, 2013, 2015): this approach has the obvious advantage of covering a broad number of dimensions but also the drawback that no overall comparison can be undertaken across countries or in time absent a single welfare metric. Nordhaus and Tobin (1973) were first to devise a monetary summary measure of well-being that incorporated some aspects of quality of life. Today, there is a myriad of composite indexes, most prominently represented by the Human Development Index. While these indexes provide a single overall metric, they typically suffer from aggregation weights that are chosen ad hoc and do not allow quantifying trade-offs between dimensions in a credible way. For instance, Ravallion (2012) focuses on the puzzling implicit valuations of life in the Human Development Index. By attempting to give stronger foundations to a composite index, this paper addresses Ravallion's concerns. Similar, most of these indexes fail to take into account distributional elements, looking at average achievements only.

While there is a long-standing literature on social welfare that addresses the aggregation of income or consumption across individuals (Atkinson, 1970, Jorgenson and Slesnick, 1983), recent theoretical and empirical advances in welfare economics have also provided foundations for aggregating across various dimensions of outcomes and so create a single-valued welfare metric. Leading examples include Fleurbaey and Gaulier (2009), Jones and Klenow (2016), Fleurbaey and Blanchet (2013), Schokkaert and Decancq (2016) and Boarini et al. (2015). The contributions to the literature differ somewhat in their theoretical foundations and in their empirical strategy.²

This paper relies on an equivalent income approach to address the issue of aggregation across several well-being outcomes, an approach that has the virtue of providing consistent welfare

² Another direction taken consists of multivariate dominance tests (Atkinson and Bourguignon, 1982). This approach avoids forming a composite index and asks instead how far one can go in ranking countries using a welfare function defined on income and longevity for which only a set of theoretical assumptions are required.

evaluations even when individuals do not hold the same preferences over the bundle of items being evaluated as illustrated in Decancq *et al.* (2015).³

We estimate people's preferences as follows. We first derive a monetary measure of the marginal welfare loss of unemployment by estimating the relative elasticity of unemployment to income in a life satisfaction regression. This econometric approach permits capturing the effects of unemployment above and beyond the loss of income. These effects relate to both the impact of unemployment on the unemployed and the impact of the unemployment rate on everyone, as an 'environmental' variable. We then calibrate a life-time utility model where utility is defined over two quality of life dimensions (longevity and jobs) as well as consumption to derive the value of a statistical life. Importantly, we identify some interaction effects between individual characteristics and income, which allow for heterogeneous preferences across individuals.

The paper contributes to the literature on welfare measurement in several ways. First, it provides a unified theoretical framework that connects the empirical literature on subjective well-being with that devoted to measuring the value of a statistical life. The model used for evaluating welfare losses due to unemployment and reduced longevity is consistent with standard decision theory (i.e. the maximization of an instantaneous utility function over the overall life time) and derives a distribution-adjusted welfare measure that we label multi-dimensional living standards (MDLS). Unlike other theoretical approaches in the field (e.g. Becker et al., 2005, Jones and Klenow, 2016), the framework does not assume an identical utility function for all individuals in the society and all countries considered. The adjustment for changes in the welfare distribution introduces a normative element into the computation and our MDLS depends on the degree of aversion to inequality chosen for a social welfare function in the tradition of Atkinson (1970). Second, it estimates the distributions of shadow prices for unemployment and longevity that is sufficiently robust and coherent with other evaluation methods in the literature. Finally, it calculates MDLS for 28 countries during the Great Recession and shows how this welfare measure has been impacted by the economic crisis. Results distinguish between different segments of the population (i.e., different degrees of aversion to inequality).

³ The use of equivalent income functions has been advocated by many economists, among them Deaton and Muellbauer (1980), Diewert (1985) and King (1983).

The main findings of the paper are as follows. First, the shadow price of one point of unemployment expressed as a percentage of household income is on average equal to 3.1% over the period, while the average shadow price of one year of longevity amounts to 5.7%. The latter empirical results constitutes a key finding of the paper and a significant extension of Jones and Klenow (2016)'s framework. There is moderate heterogeneity in shadow prices both within and between countries, except for the group of unemployed people as their shadow price of unemployment is at least twice the average shadow price observed at country level. Finally, we find that the living standards of the median households have decreased significantly during the Great Recession, while economic growth has remained slightly positive. On average across OECD countries, poor households have recorded a stark fall in living standards of 5.3% annually between 2008 and 2013. Overall, the rate of economic growth falls short of adequately reflecting the social cost of the great Recession, especially in countries most hardly hit by the crisis and among poor households.

The paper is organized as follows. The first section describes the theoretical framework, including the identification strategy and the derivation of shadow prices of unemployment and longevity. The second section presents the estimates of the shadow prices of unemployment. The third section focuses on the shadow prices of longevity, discussing how they have been calculated building on well-established approaches in the literature that are then adapted to our theoretical framework. The following section discusses the MDLS for median and poorer households, looking at the levels and the evolution of our welfare metric. The final section concludes.

2. *The setup*

2.1. Defining equivalent income and multi-dimensional living standards

The approach adopted here borrows the concept of equivalent income (or money-metric utility) that was introduced by Samuelson (1974) and later extended to non-market dimensions by Willig (1981), King (1983) and Hammond (1994). Deaton and Muellbauer (1980) used the approach as a convenient measure of people's access to resources. More recently, the equivalent income approach has been given a new foundation through the theory of fair social orderings (Fleurbaey and Maniquet 2011, Fleurbaey and Blanchet 2013).

Consider an individual i with disposable lifetime income Y_i who enjoys non-market quality of life q_i , all represented by an indirect utility function $V(Y_i, q_i)$. The equivalent income is computed as the level of income Y_i^* that would give this person the same indirect utility if she enjoyed a reference quality of life q^* :

$$V(Y_i, q_i) = V(Y_i^*, q^*).$$

Lifetime income (and, later, annual income and annual consumption) is measured with regards to a single reference price vector. We approximate this price vector by national accounts private consumption deflators and purchasing power parities so that all monetary variables are expressed in 2005 international dollars⁴.

We can now rewrite the definition of the equivalent income as⁵

$$V(Y_i, q_i) = V(Y_i^*, q^*) = V(Y_i \cdot (1 - \delta_i^q), q^*),$$

where δ_i^q is the share of real income compensating individual i for the change in quality of life from q_i to q^* . Equivalent income $Y_i^* = Y_i \cdot (1 - \delta_i^q)$ thus corresponds to real disposable lifetime income with a quality of life adjustment. When all individuals' incomes are made comparable, in that they reflect the same reference quality of life q^* and the same set of market reference prices, these equivalent incomes lend themselves to welfare comparisons between individuals.

One advantage of the equivalent income approach is its capacity to preserve consistency in welfare evaluations even if preferences among individuals are heterogeneous. Consistency is ensured if there is a single vector of reference outcomes q^* against which the outcomes for individual i , q_i , are compared. Similarly, real income is valued in terms of a single vector of reference prices⁶. Consistency in preference orderings is an advantage over measures based on

⁴ In a market context, and for welfare comparisons in time, Jorgenson and Slesnick (1983, 2014) define a welfare quantity index as $\frac{m(u^1, p^0)}{m(u^0, p^0)}$ where u^1 and u^0 are utility levels in the comparison and in base period, respectively and $m(u^i, p^0)$ is minimum expenditure on consumption products in period $i=0,1$ given prices p^0 . The corresponding quantity index implied by our current approach is $\frac{m(u^1, p^*)}{m(u^0, p^*)}$, and thus similar in spirit. See Allen and Diewert (1983) and Diewert (2014) for a discussion of price index calculation and its implication for total factor productivity growth.

⁵ With an abuse of notation, we keep V as the label of the indirect utility function with deflated income.

⁶ For a more extensive discussion of this point see Fleurbaey and Blanchet (2013) and Fleurbaey (2015).

compensating variations such as Becker et al. (2005), for which the references are moving or Jones and Klenow (2016), where the main measure (equivalent variation) takes the US values as the fixed reference, whereas the second measure (compensating variation) takes each country's values as reference. Jorgenson and Slesnick (2014), in their market-based welfare comparisons for the United States, use a chain welfare index where the reference price evolves (although every binary comparison is for a given set of prices, in line with standard practice in national accounts and price indices). Moving references create the risk of cycles and inconsistent chains, even though they have the advantage to underpin outcomes (of quality of life) or quantities (of market consumption) that are closer to the geographical location or time period of the shadow prices or market prices that the individual faces. For the purpose at hand – valuation of quality-of-life components – we have a preference for single reference vectors of market prices and quality-of-life.

In practice, two dimensions of quality of life will be selected, namely longevity and job prospects. The proxies used for job prospects are the unemployment rate and the individual labour status. Regarding the choice of reference values for the components of q^* , a “natural” reference value is the best possible outcome for longevity as well as for unemployment. The status of “not being unemployed” is thus our reference for the latter. For the former, we consider the maximum value of longevity in the countries of our sample⁷.

This choice is the only one guaranteeing that among individuals who share identical objective situations, and fail to achieve the best level for those goods, the individuals who care more about these goods are indeed considered worse off. As soon as one selects another reference level for

⁷ A detailed discussion of the choice of references q^*, p^* can be found in Fleurbaey and Blanchet (2013, appendix). The key reasoning here is as follows: “Good health is not always an unambiguous notion [...], but by and large it is well defined, and, more importantly, it is something that is considered a ‘normal’ state. In this state, people are generally compared with respect to other aspects of life, and it would be strange to argue that of two healthy individuals, the one who is less worried about keeping a good health is pro tanto less well-off than the one who cares more about health. In such a case, picking the normal value as the single reference makes good sense. Then the equivalent income systematically measures the loss for those who do not reach the normal state. [...] Unemployment is like sickness, and provides another example in which everyone aspires to avoiding it, which suggests a natural reference.” (p. 221)

these goods, one can find paradoxical situations in which the individuals who suffer more from failing to reach the best level are considered better off.⁸

Once equivalent incomes are calculated at the individual level, society's multi-dimensional living standards (MDLS) can be defined by inserting equivalent incomes into a social welfare function. Our choice falls on Atkinson's (1970) formulation, a generalised mean over individual welfare⁹ with a coefficient of inequality aversion τ :

$$MDLS = \left(\frac{1}{n} \sum_i (Y_i^*)^{1-\tau} \right)^{\frac{1}{1-\tau}} = \left(\frac{1}{n} \sum_i Y_i^* \right) (1 - I), \quad (1)$$

where $I \equiv 1 - (\sum_i (Y_i^*)^{1-\tau} / n)^{\frac{1}{1-\tau}} / (\sum_i Y_i^* / n)$ is the Atkinson-Kolm-Sen index of inequalities in equivalent incomes. As will be elaborated further below, expected longevity is taken to be the same across individuals within a country, but job prospects differ by labour status, and the marginal utility of income also differs across individuals. Then, individuals of a given country have a different valuation (as a percentage share of income) of expected longevity and job prospects. One implication for equation (1) is that inequality in equivalent income at the country level is driven by the country's income inequality, plus differences in labour status and differences in preferences for income. When MDLS is computed across all individuals and countries, the resulting inequality measure reflects both differences in individual characteristics (income and labour status) as well as differences in preferences.

⁸ Consider an individual A who cares more about health than individual B (i.e. allocates a price $p^A > p^B$). The monetary equivalents of the health level T given the reference level T^* are $p^A \cdot (T - T^*)$ and $p^B \cdot (T - T^*)$. If T^* is chosen as the maximum outcome of T , it is always the case that $p^A \cdot (T - T^*) < p^B \cdot (T - T^*)$, logically implying that A has lower utility than B because she suffers more from not achieving the best outcome. However, if T^* is not taken as the maximum outcome, it is always possible to find a paradoxical situation where A has higher utility than B (when $-T^* > 0$), although A still suffers more than B from failing to reach the best outcome. The same paradox can arise when health is given a non-zero price and one looks at the "equivalent budgets" (i.e., budget tangent to the indifference curves): It is then possible for an individual who cares more about health to have a greater equivalent budget if the price of health is not zero.

⁹ For a discussion of Atkinson's measure see Diewert (1985), Deaton and Muellbauer (1980) and Blackorby and Donaldson (1978). Another well-established social welfare specification is Jorgenson and Slesnick (1983, 2014). Jorgenson and Schreyer (2015) compare the Atkinson formula with the Jorgenson-Slesnick formula in a national accounts context. They also discuss a number of practical issues such as the transition from survey-based to national accounts data or the use of equivalence scales, issues that are not further addressed in the paper at hand.

2.2. Specification of instantaneous utility

The next task is selecting a specific form for the utility function at hand. One of the innovations in the present paper is that we allow for the inclusion of several “environmental” variables that generate additional utility or disutility for people (i.e., positive or negative externalities), and also bring about preferences’ heterogeneity across groups of people.

Unemployment is the first set of “environmental” variables that generate disutility for people. Social adverse effects of experienced unemployment such as the loss of social connections and purpose in life have been found to be detrimental to well-being in empirical studies (Blanchflower and Oswald 2004). Evidently, unemployment also has a direct negative effect on income that is taken into account (as this paper does) when considering individual income that incorporates the income shock arising from the unemployed status. In addition to direct status effects for the unemployed, there are negative externalities such as increased insecurity and lower social cohesion for others. We proxy these effects by the aggregate unemployment rate measured at country level, which affects everyone negatively. Typically, economic studies have investigated the welfare impact of unemployment only via monetary effects (Krusell and Smith 1999), and often failed to find any significant impact of unemployment on discounted utility due to the capacity of people to save and cushion income shocks. We go beyond this approach by adding a negative externality from the country’s unemployment rate.

Average disposable income in the country is the second type of environmental factor that this paper considers. Several studies have considered national income as a determinant of life satisfaction in addition to individual income, which reflects the individuals’ utility gain from living in a relatively affluent country (see Di Tella et al., 2003, and Clarck et al., 2008, for a review). The relationship between life satisfaction and income is significant both across countries and within countries in recent data (Stevenson and Wolfers, 2008), which suggests controlling for both individual income (as a classical determinant of individual utility) and average income (as an environmental factor) in order to disentangle the two effects. In doing so, however, we follow Decancq et al. (2019) and account for people’s heterogeneity in preferences by introducing some interaction between personal characteristics and the average income variable (see the discussion of the econometric model below).

In terms of the functional form, we use a Constant Relative Risk aversion (CRRA) utility function that depends positively on current consumption and is also influenced by the set of environmental factors. The CRRA is common in the literature. Murphy and Topel (2005), Hall and Jones (2007), Jones and Klenow (2016) and Murtin et al. (2015) calibrate the rate of risk aversion¹⁰ to 1.25, while Becker et al. (2005) and Gaulier and Fleurbaey (2009) retain a value of 0.8. Following De Serres and Murtin (2016) and empirical evidence provided by Layard et al. (2008)¹¹, we select a logarithmic utility function, defined over the current consumption bundle with an inter-temporal elasticity of substitution equal to unity, and an index X_i of several environmental factors that are both individual and country-specific:

$$u(c_i, X_i) = \alpha \log \left(\frac{c_i}{\omega} \right) + X_i \cdot \Gamma_i; \quad (2)$$

where X is a matrix of observables, Γ a vector of coefficients and ω parameterizes the consumption level just above survival for which instantaneous utility is equal to zero. In previous studies this parameter is calibrated on the Value of a Statistical Life to reflect the willingness-to-pay for a longer life.

2.3. *Specification of indirect utility*

To assess discounted lifetime income and utility, we place instantaneous utility (2) into an inter-temporal context, using the same valuation framework as in Becker et al. (2005) or Jones and Klenow (2016). As stressed by Rosen (1988), the willingness-to-pay for longer life is determined by the expected discounted present value of lifetime utility. Let Y_i and S_π be respectively lifetime income and the survival function that depends on the crude death rate π . The indirect lifetime utility $V(Y_i, \pi, X_i)$ derived from consumption $c_i(t)$ and affected by the set of externalities $X_i(t)$ is

$$V(Y_i, \pi, X_i) = \max \int e^{-\rho t} S_\pi(t) u(c_i(t), X_i(t)) dt$$

Assuming that utility is separable between periods, the rate of risk aversion equals the intertemporal elasticity of substitution.

¹¹

Deriving conclusions on cardinal utility from subjective well-being surveys is obviously controversial, since respondents may not pick satisfaction levels that reflect their risk aversion or their elasticity of intertemporal substitution. But the fact that the logarithmic form appears salient in these data provides a useful anchor.

$$\text{subject to } Y_i = \int e^{-rt} S_\pi(t) y_i(t) dt = \int e^{-rt} S_\pi(t) c_i(t) dt$$

where ρ is the rate of time preference, $u(\cdot)$ is instantaneous utility, r the interest rate and $y_i(t)$ annual income in period t . The budget constraint stipulates that the lifetime expected discounted value of future consumption equals expected lifetime income, assuming perfect annuity markets. In line with Becker et al. (2005), we consider a hypothetical life-cycle individual who receives the same income per capita $y_i(t) = y_i$ and faces the same set of environmental factors $X_i(t) = X_i$ throughout her lifetime, and who is subject throughout life to the survival function of the country at a given date. Under the assumption that the interest rate r equals the rate of time preference ρ , the solution of the maximisation programme specifies that consumption remains constant throughout the life cycle, and that for all t , $c_i(t) = y_i$, so that indirect utility is simply given by

$$V(y_i, \pi, X_i) = u(y_i, X_i) A(\pi) \quad (3)$$

where $A(\pi) = \int e^{-rt} S_\pi(t) dt$ is the value of an annuity discounted by the survival function $S_\pi(t)$. A standard assumption in this context is a constant mortality rate π over the life cycle, that is, $-\frac{S'_\pi(t)}{S_\pi(t)} = \pi$ and $S_\pi(t) = e^{-\pi t}$. This yields a simple expression for life expectancy T – calculated as the integral of the survival function – as $T = 1/\pi$ and for the value of one annuity: $A(\pi) = 1/(r + \pi)$. Indirect utility now reads:

$$V(y_i, T = 1/\pi, X_i) = \frac{u(y_i, X_i)}{r + \pi(T)} = \frac{u(y_i, X_i)}{r + 1/T} \quad (4)$$

A constant mortality rate throughout the life cycle is not realistic from an empirical perspective, but importantly, the latter assumption is innocuous for the valuation of longevity. To see this, consider the ‘true’ observed distribution of mortality rates by age $-\frac{S'_\pi(t)}{S_\pi(t)} = \pi(t)$. To relate it to the Value of a Statistical Life (see below), it is convenient to consider an average Crude Death Rate equal to the average mortality rate calculated over the stationary distribution of the population where each age group t has a population share $S_\pi(t)/\int_0^\infty S_\pi(t) dt = S_\pi(t)/T$. This yields an average mortality rate equal to $\bar{\pi} = \int_0^\infty \frac{-S'_\pi(t)}{S_\pi(t)} \frac{S_\pi(t)}{T} dt = \frac{1}{T}$, which is the same relationship between mortality and longevity as the one derived above under the simplifying assumption of a constant mortality rate. In other words, the most natural Crude Death Rate to be considered in this

simple valuation framework is always equal to the inverse of longevity, irrespective of the distribution of mortality rates by age.

We shall now explicitly introduce the spatial, cross-country dimension to our setup when combining equations (2) and (3) so that the indirect utility function of an individual i who lives in country j with life expectancy T_j and environmental factors $X_{i,j}$ and who enjoys an annual disposable income $y_{i,j}$ is given by:

$$V(y_{i,j}, T_j, X_{i,j}) = \frac{\alpha \log\left(\frac{y_{i,j}}{\omega}\right) + X_{i,j} \cdot \Gamma_i}{r + 1/T_j} \quad (5)$$

Indirect utility thus depends on fundamental parameters $(\alpha, \Gamma_i, \omega, r)$. Parameters (α, Γ_i) determine the willingness-to-pay to change external factors $X_{i,j}$, while parameters (ω, r) determine the value of life or the willingness-to-pay for extended life expectancy.

2.4. Identification strategy

To identify the required parameters we rely on life satisfaction regressions for parameters (α, Γ_i) and on a calibration approach for the longevity parameter ω that is based on the Value of a Statistical Life (VSL). The value of the discount rate r is fixed at 3% as in Becker et al. (2005).

Identification of parameters relies on the separability of indirect utility into longevity on the one hand and other factors on the other hand. Said differently, instantaneous utility is shaped by contemporary effects of income and environmental factors (including unemployment), but not by progress in longevity. The valuation of environmental factors $X_{i,j}$ with respect to target $X_{i,j}^*$ is derived from the following equation:

$$V(y_{i,j}, T_j, X_{i,j}) = V(y_{i,j}^*, T_j, X_{i,j}^*)$$

which can be simplified due to separability as

$$u(y_{i,j}, X_{i,j}) = u(y_{i,j}^*, X_{i,j}^*) \quad (6)$$

Our key identifying assumption is that life satisfaction can be viewed as a linear transformation of instantaneous utility, implying from (2) and (6) that

$$\begin{aligned} LS(y_{i,j}, X_{i,j}) &= \mu \cdot u(y_{i,j}, X_{i,j}) + \tau + \theta_{i,j,t} \\ &= \mu \cdot \alpha \log(y_{i,j}) + \mu \cdot X_{i,j} \cdot \Gamma_i + \tau - \mu \cdot \alpha \cdot \log(\omega) + \theta_{i,j,t} \end{aligned} \quad (7)$$

where $\theta_{i,j,t}$ represents an error term with a structure discussed below, and (μ, τ) two unknown parameters. Equation (7) shows that life satisfaction regressions allows for the identification of the ratio of interest Γ_i/α , but not for the identification of parameter ω as τ is unknown. Without loss of generality, one can select $\mu = 1$.

Our identifying assumption is supported by the empirical finding that most respondents to life satisfaction surveys are not forward looking and have a relatively short time horizon in mind as shown by Benjamin et al. (2016). Hence, life satisfaction is deemed to reflect contemporary and large changes in socio-economic conditions rather than long-term transformations with small contemporary changes such as progress in longevity. Building on previous work in this area (Murtin et al., 2015) we nonetheless test whether longevity has a significant and robust effect on life satisfaction and our evidence is mixed at best (see below). Although some specifications yield plausible results that are fully in line with this paper's main results, they seem insufficiently robust. This is corroborated by other research, in particular Deaton (2008) who also finds no robust evidence that changes in longevity and in life satisfaction are correlated.

A proper econometric specification of life satisfaction will then yield estimates of parameters $(\alpha, \Gamma_{i,j})$, while equations (6) and (7) will allow for the valuation of environmental factors $X_{i,j}$ that verify:

$$LS(y_{i,j}, X_{i,j}) = LS(y_{i,j}^*, X_{i,j}^*) \quad (8)$$

In turn, the estimate of parameter ω and the valuation of longevity are obtained from a calibration of the instantaneous utility function on the VSL. The vast majority of studies devoted to the subject derive the value of longevity from a measure of the VSL. As explained in De Serres and Murtin (2015) there is a tight connection between the VSL and the notion of compensating income presented in Section 2.1. Consider a population of N workers who pay each year the compensating income $\delta.y$ to observe a marginal decrease $d\pi$ in the crude death rate π . Each year, the statistical number of saved life is $-N d\pi$ and the amount invested for the mortality risk reduction is $N \delta.y$, which corresponds to a VSL

$$VSL = -\frac{\delta.y}{d\pi}$$

For marginal changes in the crude death rate π , the compensating income δ also satisfies $V(y - \delta.y, \pi + d\pi, X) = V(y, \pi, X)$ with $d\pi < 0, \delta > 0$, which yields after a first-order expansion and the use of (3):

$$VSL = -\frac{\delta}{d\pi} = -\frac{\frac{\partial V}{\partial \pi}}{\frac{\partial V}{\partial y}} = -\frac{A'(\pi)}{A(\pi)} \frac{u(y,X)}{u'(y,X)} \quad (9)$$

Equation 9 provides an additional moment condition, which will be used to estimate parameter ω , conditionally on other estimates $(\hat{\alpha}, \widehat{\Gamma_{l,j}})$ and the value of the VSL in a given country and time.

In summary, we use a hybrid approach where the valuation of environmental factors (including unemployment) is derived from life satisfaction regressions while the valuation of longevity is based on values of statistical life (VSL), itself based on willingness-to-pay for reducing mortality risk (as typically derived from wage differentials).

2.5. *Econometric framework*

The vector $\theta_{i,j,t}$ comprises country fixed-effects that reflect systemic cross-country cultural (or other time-invariant) differences in the relationship between life satisfaction and individual utility, a period-specific component allowing momentary shifts in the latter relationship, individual characteristics such as age and gender reflecting systemic differences in life satisfaction across population groups enjoying the same level of instantaneous utility, as well as an error term:

$$\theta_{i,j,t} = a_j + b_t + \pi \cdot Z_i + \varepsilon_{i,j,t} \quad (10)$$

As explained above, the set of environmental factors include : i) an individual unemployment dummy U_i ; ii) country's unemployment rate $U_{j,t}$; iii) the log of country's average disposable household income, interacted with individual characteristics W_i including age, gender, unemployment's status and income group (i.e. high, medium or low income in the national income distribution), which yields:

$$X_{i,j,t} \cdot \Gamma_i = \beta \cdot U_i + \theta \cdot U_{j,t} + (\gamma + W_i \cdot \Lambda) \cdot \log(y_{j,t}) \quad (11).$$

As a result, the estimated life satisfaction regression combines equations (7), (10) and (11) and writes (while ignoring constant terms):

$$LS_{i,j,t} = a_j + b_t + \pi \cdot Z_i + \alpha \cdot \log(y_{i,j,t}) + (\gamma + W_i \cdot \Lambda) \cdot \log(y_{j,t}) + \beta \cdot U_i + \theta \cdot U_{j,t} + \varepsilon_{i,j,t} \quad (12)$$

where $\varepsilon_{i,j,t}$ is the residual. Importantly, in all regressions we calculate robust standard errors clustered at the country level in order to avoid the underestimation of standard errors pertaining to

country-level variables (i.e. $\log(y_{j,t})$ and $U_{j,t}$). Other econometric specifications have been tested, including interactions between individual income and individual characteristics. In most cases, they were not statistically significant and did not explain much of the variance of life satisfaction.¹²

Our attempt of estimating the welfare effect of unemployment from life satisfaction regressions faces several econometric challenges, some of which are described in Ravallion and Lokshin (2001) and Ravallion et al. (2016). First, life satisfaction regressions are capable of overestimating the welfare losses from unemployment given that unemployment is likely to be correlated with key personality traits that also negatively influence personal happiness and satisfaction with life. To reduce the influence of unobserved heterogeneity, individual fixed effects could be introduced inside the regressions if a longitudinal panel was available, which is not the case as the available data are repeated cross-sections. Meanwhile, one introduces country fixed-effects that reduce the risk of biased estimates due to country-level unobserved heterogeneity (i.e., omitted time-invariant country effects that are correlated to the regressors). In the sample of countries under study, there appears to be very low correlation (i.e., below 0.25) between changes in log income, longevity and unemployment, and hence little risk of encountering multicollinearity problems.

Second, individual income in particular is affected by large measurement errors in the Gallup World Poll, which may bias its estimated coefficient towards zero. In this regard, Murin et al. (2015) show that the use of individual income drawn from the Gallup survey yields a large attenuation bias on the income variable and hence a crude overestimation of shadow prices. Several other studies have raised the same econometric issue. For instance, Powdthavee (2010) highlights the role of large measurement errors on income, which yield a sizeable attenuation bias. Similarly, Fujiwara (2013) argues that the “wellbeing valuation” provides biased estimates of the value of non-market goods unless the income variable is instrumented. Strikingly, he finds in his case-study that the coefficient on log income jumps from 0.16 (non-instrumented income variable) to 1.10 (instrumented variable). However, we are less subject to this risk as we also control for average disposable income.

¹²

It is also possible to introduce some interaction between individual characteristics and the unemployment rate. For the sake of model’s parsimony and better readability of the results, we included a single source of interaction.

2.6. Willingness-to-pay to eliminate unemployment and increase longevity

From equation (12), the willingness to eliminate unemployment considers the hypothetical situation with zero unemployment for both individuals and society ($U_i = U_{j,t} = 0$) in exchange for a reduction in income for both individuals and society by a factor $\delta_{i,j}^U$ that verifies¹³

$$\alpha \cdot \log(y_{i,j}) + (\gamma + W_i \cdot \Lambda) \cdot \log(y_j) + \beta \cdot U_i + \theta \cdot U_j = \alpha \cdot \log((1 - \delta_{i,j}^U) \cdot y_{i,j}) + (\gamma + W_i \cdot \Lambda) \cdot \log(1 - \delta_{i,j}^U) \cdot y_j \quad (13)$$

which yields the individual shadow price

$$\delta_{i,j}^U = 1 - \exp\left(\frac{\beta \cdot U_i + \theta \cdot U_j}{\alpha + \gamma + W_i \cdot \Lambda}\right) \quad (14)$$

Conveniently, the econometric results will report the average ‘social cost of unemployment’ δ^U that is calculated after averaging the unemployment dummy and other group dummies in equation (13), which writes:

$$\delta^U = 1 - \exp\left(\frac{\beta \cdot \bar{U}_i + \theta \cdot U_j}{\alpha + \gamma + \bar{W}_i \cdot \Lambda}\right) \quad (15)$$

Likewise, the willingness to increase longevity from T_j to T^* implies a reduction in individual and aggregate incomes by a factor $\delta_{i,j}^T$ that verifies $V(y_{i,j}, T_j, X_{i,j}) = V(y_{i,j}^*, T^*, X_{i,j}^*)$, which yields after using (5) and (11) and applying a few algebraic manipulations: :

$$\delta_{i,j}^T = 1 - \exp\left(-\frac{\alpha \cdot \log\left(\frac{y_{i,j}}{\omega}\right) + X_{i,j} \cdot \Gamma_i}{\alpha + \gamma + W_i \cdot \Lambda} \cdot \left(1 - \frac{r+1/T^*}{r+1/T_j}\right)\right) \quad (16)$$

As before, an average shadow price of longevity can be calculated by averaging dummies in the sample.¹⁴

¹³ In this calculation, one assumes a certain degree of myopia as agent i applies her own willingness-to-pay on income's reduction of all agents in society. Thus, she is ignorant of others' preferences. In formula (13), time subscripts are dropped by convenience.

Finally, the willingness-to-pay to eliminate unemployment and simultaneously raise longevity from T to T^* is solution of $A(T_j).u(y_{i,j}, X_{i,j}) = A(T^*).u(y_{i,j}^*, X_{i,j}^*)$. The corresponding shadow price $\delta_{i,j}^{T,U}$ is used to renormalize $\delta_{i,j}^U$ and $\delta_{i,j}^T$ by the ratio $(\delta_{i,j}^U + \delta_{i,j}^T)/\delta_{i,j}^{T,U}$ for the sake of decomposability in willingness-to-pay.¹⁵

3. *The value of work*

This section describes the data used to construct the MDLS index as well as the results from subjective well-being regressions.

3.1. *The Data*

An original database gathering various measures of income, longevity, unemployment as well as income distribution (i.e. average income across 10 deciles) has been constructed for 28 OECD countries. Time coverage varies but the final index of MDLS is available from the early 2000s and the latest year of the database is 2013.

Our measure of average income is national accounts-based household net disposable income. This is the sum of primary household incomes from various sources (compensation of employees, self-employed income, and property income), minus taxes paid on income and wealth, minus social security contributions paid, plus government transfers received and minus depreciation of household assets (essentially real estate). All variables are expressed at constant prices and constant PPPs with 2005 as the reference year. Along with the general unemployment rate (number of unemployed as a percentage of the labour force) and life expectancy at birth, variables were extracted from the various sections of the online OECD database (OECD, 2015a). This includes also information from the OECD Income Distribution database (OECD, 2015b) from which we draw information on the distribution of disposable household income. It is of note that

¹⁴ Parameter ω is equal to $\overline{y_{i,j}} \cdot \exp \left(-\frac{\frac{VSL}{y_j} \cdot \left(r + \frac{1}{T_j} \right) \cdot (\alpha + \gamma + W_i \cdot \Lambda) \cdot -X_{i,j} \cdot \Gamma_i}{\alpha} \right)$ where $\overline{y_{i,j}}$ is the average individual income in the reference country and year (the United States in 2007).

¹⁵ The exact formula is $\delta_{i,j}^{T,U} = 1 - \exp \left(\frac{A(T_j) - A(T^*)}{A(T^*)} \cdot \frac{\alpha \cdot \log \left(\frac{y_{i,j}}{\omega} \right) + (\gamma + W_i \cdot \Lambda) \cdot \log(y_j)}{\alpha + \gamma + W_i \cdot \Lambda} + \frac{A(T_j)}{A(T^*)} \cdot \frac{\beta \cdot U_i + \theta \cdot U_j}{\alpha + \gamma + W_i \cdot \Lambda} \right)$

the income definition underlying the distributional data differs somewhat from the income definition of the aggregate national accounts data as the distributional information is directly based on household surveys and administrative registers. Survey measures of household income refers to cash income – excluding imputed components such as home production and imputed rents – that is disposable after transfers and taxes. It includes earnings (broken down into those of the household head, of the spouse and of other household members), self-employment income, capital income (rents, dividends and interest), public transfers and household taxes. Incomes are equivalised within the household using the square root of household size as an equivalence scale¹⁶. Data on life satisfaction come from the Gallup World Poll (2004-2013). Table 1 shows values of the main indicators of interest for 2013.

Table 1. Descriptive statistics - 2013

	Household disposable income <i>2005 USD PPP</i>	Life expectancy at birth	Unemployment rate <i>in per cent</i>	Income ratio of top and bottom quintiles	Life satisfaction (0-10 scale)
AUS	24920	82.2	5.8	5.4	7.4
AUT	22178	80.9	5.0	4.4	7.6
BEL	19201	80.3	8.5	4.0	7.3
CAN	23534	82.1	7.2	5.2	7.6
CHE	26075	82.8	4.5	4.4	7.7
CHL	10157	78.9	6.2	13.0	6.7
CZE	11919	78.4	7.1	3.7	6.5
DEU	22363	81.2	5.4	4.4	7.0
ESP	15590	82.4	26.2	6.7	6.2
EST	9713	76.6	8.8	5.4	5.2
FIN	17330	80.8	8.4	3.7	7.5
FRA	19939	82.0	10.0	4.7	6.7
GBR	21632	81.0	7.8	5.6	7.0
GRC	12250	80.6	27.5	6.3	4.7
HUN	9153	75.4	10.3	4.5	4.9
IRL	17568	81.2	14.1	4.7	6.8
ITA	17064	82.3	12.4	5.6	6.1
JPN	18590	83.7	4.3	6.2	6.0
KOR	15019	81.6	3.3	5.5	5.9
MEX	9774	74.6	5.2	13.7	7.5
NLD	18395	81.1	6.7	4.1	7.5

¹⁶

As part of an ongoing work programme at the OECD, distributional data is being developed that is fully consistent with the national accounts. First results have been provided by Fesseau et al. (2013a, 2013b) but information for longer time periods and broader country coverage are needed before this work can be incorporated into the present calculations.

NZL	16141	81.8	6.4	5.2	7.4
POL	9952	76.9	10.5	4.8	5.7
PRT	13648	80.4	17.0	5.8	5.2
SVK	10935	76.3	14.3	3.9	5.9
SVN	14044	80.3	10.3	3.6	6.1
SWE	18965	81.7	8.2	4.1	7.5
USA	34808	78.9	7.5	8.2	7.2

Note: Data for household income, life expectancy, unemployment and income distribution are extracted from OECD (2015a, 2015b). Data on life satisfaction is from Gallup World Poll (2013).

3.3. *Main results*

Table 2 reports the estimation of simplified versions of Equation (12) that ignore interaction effects as a first step (i.e. $\Lambda = 0$). Column I presents a ‘macro-economic model’ that excludes variables observed at the individual level (i.e. $\alpha = \beta = 0$), Column II uses a ‘micro-economic’ model that excludes variables observed at the country level (i.e. $\gamma = \theta = 0$), while Column III combines both models. In all cases, the average social cost of unemployment depicted by equation (15) is reported. It is divided by the average unemployment rate observed across countries and time, so that it represents the average percentage of income lost per point of unemployment. In Columns II and III, it is also possible to calculate the social cost of unemployment for the unemployed, by using $U_i = 1$ in equation 15.

Table 2. The shadow price of unemployment – first estimates

	Macro regression I	Micro regression II	Combined regression III
Dependent variable is life satisfaction			
Log of average household disposable income	1.263** (0.467)		1.053** (0.466)
Log of equivalised disposable income		0.461*** (0.046)	0.459*** (0.050)
Country's unemployment rate	-0.048*** (0.006)		-0.037*** (0.006)
Individual unemployment dummy		-0.797*** (0.037)	-0.779*** (0.041)
Social cost of unemployment for the unemployed (%) <i>(% income lost by the unemployed per point of the unemployment rate)</i>		9.8	6.1
Social cost of unemployment per capita (%) <i>(% income lost per point of the unemployment rate)</i>	3.3	0.7	2.4
Controls for gender and age dummies	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes
Country dummies	Yes	Yes	Yes
N	2.2e+05	2.6e+05	2.2e+05
R ²	0.132	0.181	0.174

Note: robust standard errors clustered at the country level; for the calculation of unemployment's average shadow price, one retains a sample's unemployment rate $U=0.034$ and a country's unemployment rate $UN=0.084$.

Table 2 reveals several interesting features. First, the coefficient on country's average disposable income is always strongly significant even after controlling for individual income (Column III), and its coefficient is larger than that on individual income. This finding is consistent with our previous claims of potential measurement errors affecting the individual income variable. Likewise, country's unemployment rate is always strongly significant even after controlling for the individual unemployment dummy. This supports the view that changes in aggregate unemployment are associated with changes in life satisfaction among the whole population, and not just among the unemployed. More importantly, country's unemployment rate explains the bulk of the social cost of unemployment for society, which falls from 3.25% (Column I) to 0.68% (Column II) when the aggregate unemployment rate is excluded from the regression. This is explained by the fact that the aggregate unemployment rate affects everyone in the country, while the individual unemployment status concerns a much smaller population. On the other hand,

Table 2 also reveals that the cost of unemployment is much larger for the unemployed (6.1% on Column III) as for the rest of the population (2.4%), which is explained by effect carried out by the unemployed individual dummy. Finally, the estimate of 2.4% of income lost per point of unemployment (Column III) is broadly in line with previous results in Murtin et al. (2016).

Table 3 presents the estimation of the full model laid out in equation (15) with interaction effects. All explanatory variables are statistically significant, often at 5% or 1% confidence level. Across all specifications, the average social cost of unemployment varies between 2.4% and 3% per point of the aggregate unemployment rate. The elasticity of national income is between 2 and 5 times larger than that of individual income. Table 3 reveals interesting differences in people's preferences for higher national income. It is stronger for females (Column 1), highly educated people (Column 2), older people (Column 3), middle and high-income people (Column 4), and lower among unemployed people (Column 5). Column 6 includes all interaction terms and serves as a base model to value unemployment from equation (14), and calculate equivalent incomes across societies and time while accounting for preferences' heterogeneity.

Figure 1 depicts the resulting distribution of unemployment's shadow price. The top panel reflects only the role of interaction effects in the estimated model as it imposes a similar aggregate unemployment rate across countries (equal to 8.4%). It shows a bimodal distribution that mainly derives from the gender difference in income preferences. The group of unemployed people is too small among the sample to be visible on that graph. The second panel allows for within and between country heterogeneity. It also depicts a bimodal distribution as implied by gender differences, but the dispersion of shadow prices is also larger as a result of the cross-country variation in the unemployment rate. Across countries and time, the mean and median shadow prices of unemployment are equal to 3.1% with a 5th percentile at 2.6% and a 95th percentile at 3.7%.¹⁷

¹⁷ The average of individual shadow prices (3.1%) is slightly different from the shadow price of the average individual in the sample as reported in Table 3 (3.0%).

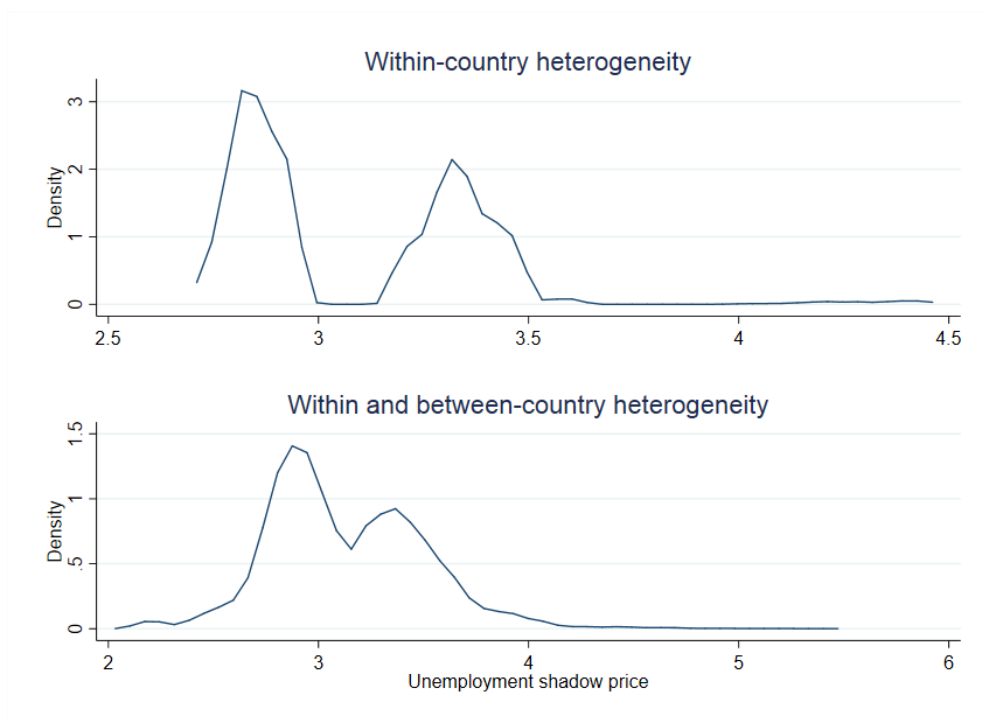
Table 3. The shadow price of the employment ratio - baseline estimates

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable is life satisfaction						
Log of average household disposable income (y)	1.165** (0.467)	0.916* (0.479)	1.064** (0.465)	1.139** (0.461)	1.065** (0.466)	1.147** (0.469)
Log y x male	-0.224*** (0.059)					-0.235*** (0.055)
Log y x higher education		0.039*** (0.004)				0.037*** (0.004)
Log y x young			-0.023*** (0.008)			-0.024*** (0.008)
Log y x middle-age			-0.010* (0.005)			-0.011* (0.006)
Log y x high income				0.062*** (0.007)		0.060*** (0.007)
Log y x middle income				0.034*** (0.004)		0.034*** (0.003)
Log y x unemployed					-0.270** (0.110)	-0.280** (0.102)
Log of equivalised disposable income	0.460*** (0.050)	0.436*** (0.048)	0.459*** (0.050)	0.209*** (0.055)	0.459*** (0.050)	0.197*** (0.052)
Country's unemployment rate	-0.037*** (0.006)	-0.039*** (0.007)	-0.037*** (0.006)	-0.041*** (0.006)	-0.038*** (0.006)	-0.042*** (0.007)
Individual unemployment dummy	-0.778*** (0.041)	-0.763*** (0.040)	-0.778*** (0.041)	-0.751*** (0.038)	-0.720*** (0.050)	-0.673*** (0.046)
Social cost of unemployment - All (% income lost per point of the unemployment rate)	2.4	2.7	2.4	2.8	2.4	3.0
Males	2.5					3.4
High-educated		2.7				2.8
Young			2.4			2.9
Middle-age			2.4			2.9
High-income				2.7		2.8
Middle-income				2.8		2.8
Unemployed					6.7	7.4
Controls for gender and age dummies	Yes	Yes	Yes	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes
Country dummies	Yes	Yes	Yes	Yes	Yes	Yes
N	2.2e+05	2.2e+05	2.2e+05	2.2e+05	2.2e+05	2.2e+05
R ²	0.175	0.178	0.174	0.179	0.174	0.184

Note: robust standard errors clustered at the country level. For the calculation of the social cost of unemployment, one retains a share of unemployed people in the sample $U=0.034$ and an average country's unemployment rate $UN=8.4\%$. * $p<0.10$; ** $p<0.05$; *** $p<0.01$.

These numbers and Figure 1, however, do not reveal the large unemployment's shadow prices recorded among the unemployed, which constitute only 3% of the sample's population and hence are not visible on Figure 1. Yet, the average shadow price among unemployed workers amounts to 7.9% per point of unemployment, and it varies broadly between 3.3% (5th percentile) and 13.3% (95th percentile) across countries and time. On average, the unemployment's shadow price of the unemployed is more than the double of the average shadow price in the population.

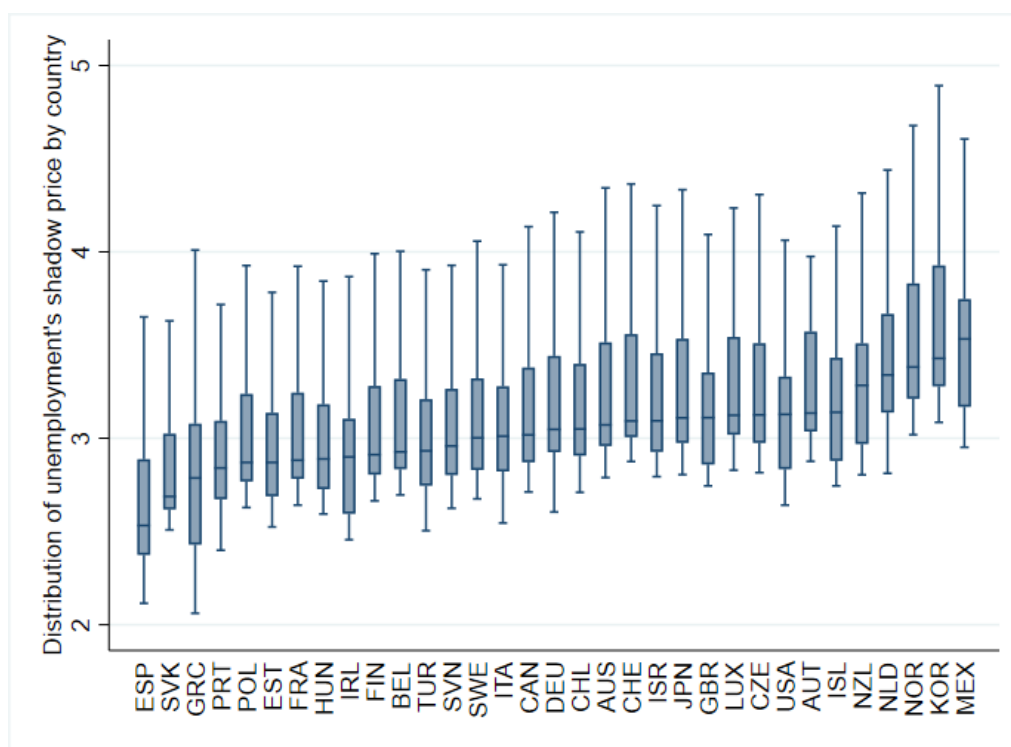
Figure 1. The distribution of unemployment's shadow price with within and between country heterogeneity - 2013



Note: The shadow price of unemployment is calculated as the monetary value of eliminating unemployment divided by the unemployment rate and expressed as a share of household average income. Distributions are smoothed using Epanechnikov kernel.

Figure 2 depicts the distribution of unemployment's shadow prices by country over the 2007-2013 period. Boxes depict the 75th percentile (upper hinge), the median and the 25th percentile (lower hinge) of national distributions of shadow prices. Median shadow prices range from 2.5% in Spain to 3.5% in Mexico. The large majority of OECD countries record a median shadow price comprised between 2.8% and 3.2%.

Figure 2. The distribution of unemployment's shadow price by country



Note: The shadow price of unemployment is calculated as the monetary value of eliminating unemployment divided by the unemployment rate and expressed as a share of household average income.

4. *The value of life*

This section provides details on the calculation of the willingness-to-pay for longer life and it highlights the cross-country differences in this valuation.

4.1. *Choice of the VSL*

A large number of studies have been devoted to the valuation of life using the approach of the VSL. While Murphy and Topel (2006) have provided the theoretical foundations of the valuation framework, Viscusi (1993) and Viscusi and Aldy (2003) have surveyed the empirical studies that have sought to assess the VSL. Typically, these studies assess the willingness-to-pay of workers for reductions in risks of accidental death by using wage differences across jobs with different occupational mortality risks. The marginal willingness-to-pay of workers is then aggregated on a statistical population and expressed in the form of the VSL, which is defined as the aggregate monetary value that a given population would be collectively willing to pay to save one

(statistical) life by reducing mortality risks. In practice, estimates of the VSL range between 4 and 9 million in 2004 USD, and the US Environmental Protection Agency uses a default value of 6.3 million in 2004 USD. As shown by Murphy and Topel (2005), the latter amount corresponds to the value of a life-year of about USD 200,000 at the beginning of active life, which peaks at USD 350,000 around age 50 and flattens around USD 100,000 after age 90.

For the purpose at hand, the parameter ω is calibrated on the average VSL value of 6.3 million 2004 USD, the US Environmental Protection Agency benchmark.¹⁸ This is a conservative choice as Viscusi-Aldy (2003), in their meta-analysis, report an average VSL of 6.7 million USD. After reviewing over 100 empirical studies, the authors conclude that “the mean predicted VSL vary from 5.5 to 7.6 million USD for the U.S. sample” (Viscusi and Aldy, 2003, p.42). Finally, we follow Becker et al. (2005) and select an annual interest rate of $r = 0.03$. The VSL of 6.3 million in 2004 is converted to a value of 6.6 million in 2007, the first of observation in the United States’ Gallup survey.

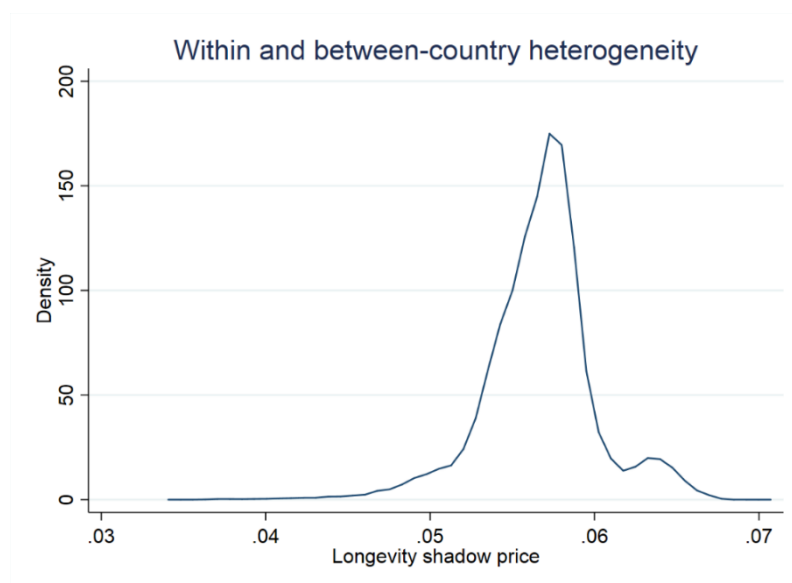
As argued by Murtin et al. (2015), VSL studies pertain to adult workers and ignore the unrevealed value of life-years during childhood, but the surveys by Hammitt and Haninger (2010) suggest a significantly larger VSL for a child. To ensure consistency between the empirical VSL and the theoretical framework, one considers an adult aged 25 years and we correct life expectancy at birth by subtracting 25 years. As life expectancy at age 25 is not available for all countries and years, this adjustment is convenient from an empirical point of view. Moreover, child mortality is not large in our sample of OECD countries and this adjustment is very close to reality. For instance, Murtin et al. (2017) find an average life expectancy at age 25 years equal to 54.5 years among a sample of 15 OECD countries over the period 2009-2013, while the corresponding life expectancy at birth equals 78.9 years, entailing a gap of 24.4 years.

¹⁸ See Dockins et al. (p.4): “EPA’s Guidelines advise analysts to use a central VSL estimate of \$4.8 million in 1990 dollars. Based on the gross domestic product (GDP) deflator this converts to approximately \$6.2 million in 2002 dollars. This value is derived from 26 estimates assembled for EPA’s first retrospective analysis of the Clean Air Act (USEPA, 1997). Each estimate is from a different study, with 21 of the estimates from hedonic wage studies and the remaining five derived from contingent valuation (CV) studies. Until 2003, the estimate from EPA’s Guidelines was uniformly applied to mortality risk reductions across program offices. EPA recently used an estimate of \$5.5 million (1999 dollars) in its analysis of reduced mortality from air regulations.”

As a result, the parameter ω yielding a VSL of 6.6 million in the United States in 2007 is equal to 130.96. Considering the infra-marginal increase of one year in longevity in the United States in 2007, equation (16) yields an average value of 6.6% of household disposable income, or 2,230 USD. This figure is almost fully consistent with the VSL and equation (9) that pertains to marginal changes. Indeed, after adjustment for life expectancy at age 25, the latter increase implies a decrease in average mortality equal to $1/52.4 - 1/53.4 = 3.57 \times 10^{-4}$ or 3.57 deaths per 10,000 people, implying a VSL of $2,230 / (3.57 \times 10^{-4}) = 6.3$ million USD. The difference with the benchmark VSL of 6.6 million USD only arises from the difference in formulas for marginal and infra-marginal changes.

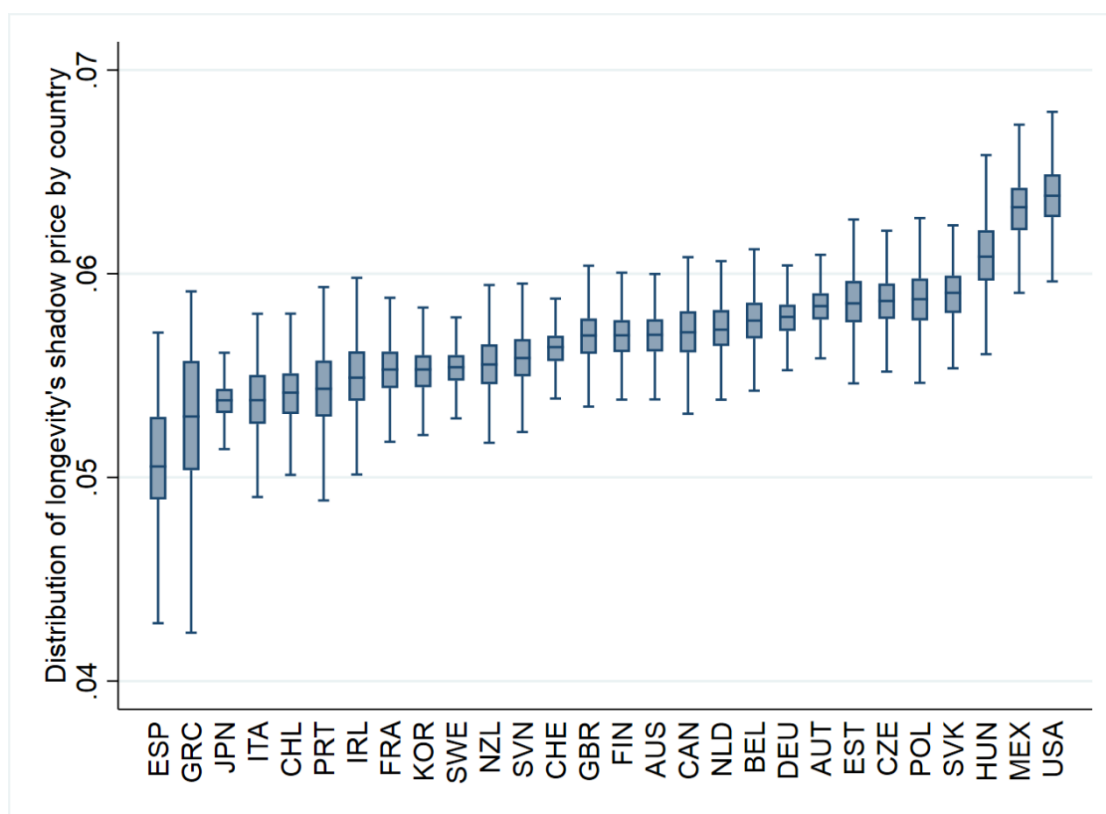
As for unemployment, equation (16) makes clear that the shadow price of longevity varies across individuals, countries and time. Figure 3 plots its distribution among the sample. Over the period, the shadow price of one additional year of life expectancy is on average equal to 5.7%, with a 5th percentile equal to 5.1% and a 95th percentile equal to 6.3%. Figure 4 further describes the distribution of longevity's shadow prices by country in 2013. Median shadow prices are lowest in Spain (at 5.1%) and highest in the United States (at 6.3%). There is more cross-country differences than within-country in longevity's shadow price.

Figure 1. The distribution of longevity's shadow price across OECD countries



Note: The shadow price of longevity is expressed as a share of household income. The distribution is smoothed using Epanechnikov kernel.

Figure 2. The shadow price of longevity across countries in 2013



Note: The shadow price of longevity is calculated as the monetary value of raising longevity by one year expressed as a share of household average income.

5. *Well-being during the Great Recession*

This section calculates multi-dimensional living standards based on heterogeneous preferences and depicts its variation during the Great Recession.

5.1. *Computation of multi-dimensional living standards*

Once the willingness to pay to reduce unemployment and increase longevity are calculated, multi-dimensional living standards can be computed as specified in equation 1:

$$MDLS = \left(\frac{1}{n} \sum_i Y_i^* \right) \cdot (1 - I) = \left(\frac{1}{n} \sum_i y_i \cdot (1 - \delta_i^T - \delta_i^U) \right) \cdot (1 - I) \quad (17)$$

In this setting, the calculation of MDLS relies on individual-specific corrections for unemployment and longevity. The Atkinson-Kolm-Sen Index of inequalities I is calculated for two levels of aversion to inequality τ , which is calibrated so that the average MDLS across countries matches the median equivalent income across countries ($\tau=0.89$) or the average 20th percentile of equivalent income across countries ($\tau=3.36$). We will thus present two indices that pertain either to the median household or to relatively poor households in terms of multi-dimensional living standards.

While relying on micro-data was essential to derive shadow prices at the individual level (as it accounts for individual characteristics), the shape of the income distribution is poorly reflected in the Gallup survey and the use of individual income is not warranted to derive a credible country-level MDLS. Hence in this section, we impute individual income by matching each decile from the micro-data (i.e. the Gallup survey) with the corresponding decile's average income from the OECD Income Distribution Database. The calculation of MDLS is nonetheless completed at the individual level.

Formula (17) provides an exact calculation of MDLS but does not allow decomposing the various contributions of income, unemployment and longevity due to the interaction between individual-specific components. For the purpose of providing some accounting of living standards, one

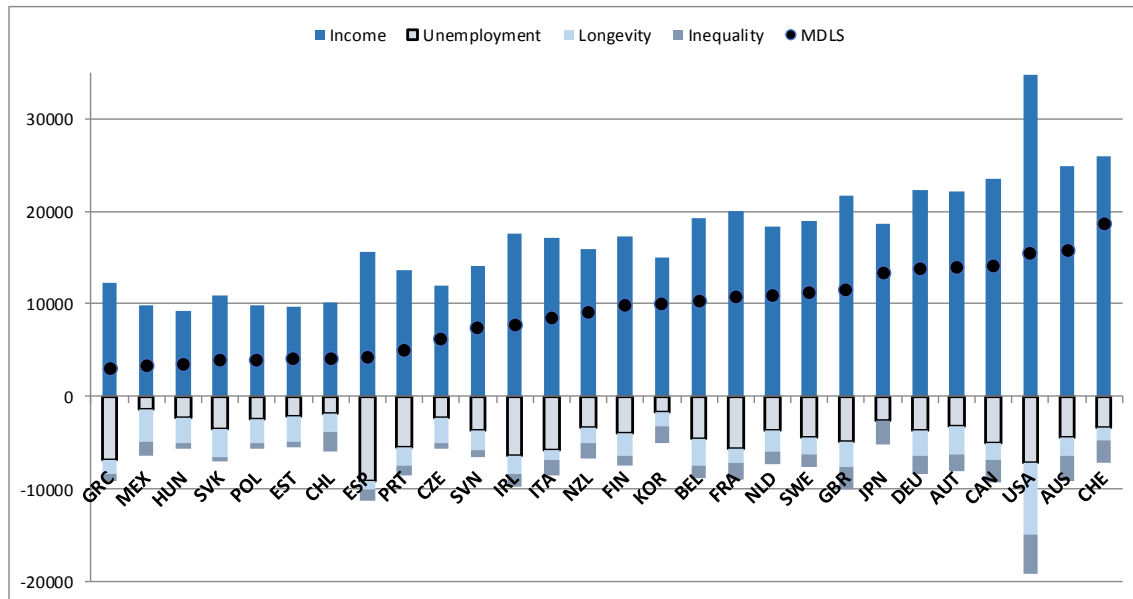
calculates an approximation of MDLS in country j by replacing individual shadow prices by their national average, which writes:

$$\widehat{MDLS}_j = y_j \cdot (1 - \delta_j^T - \delta_j^U)(1 - I) = y_j - y_j \cdot \delta_j^T - y_j \cdot \delta_j^U - I \cdot y_j \cdot (1 - \delta_j^T - \delta_j^U) \quad (18)$$

where \widehat{MDLS}_j is the proxied MDLS, y_j is average national income and (δ_j^T, δ_j^U) are national averages of individual-specific shadow prices. Numerically speaking, there is very little difference between the actual MDLS and its approximation \widehat{MDLS}_j . The relative error between the two variables records a mean of 0.54%, a 5th percentile of -0.15% and a 95th percentile of 1.82%. Thus, one can conveniently present results pertaining to the proxy \widehat{MDLS}_j .

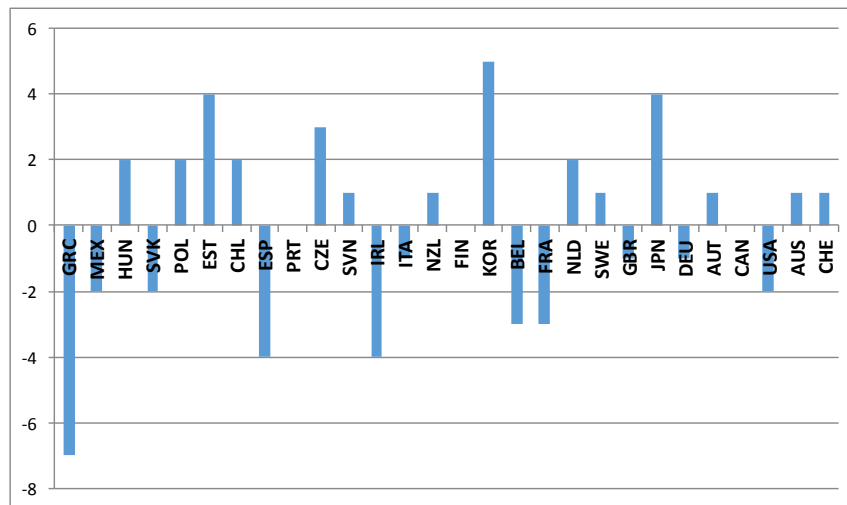
Formula (18) decomposes \widehat{MDLS}_j into four terms, namely average income and the corrections for unemployment, the longevity gap and inequality in living standards. Figure 5 depicts the MDLS and its four components for the median household and the latest available year. MDLS vary in a large interval, from 18,737 USD for Switzerland to 3,008 USD for Greece. Across the sample, average living standards (9,100 USD) represent about half of average disposable income (17,150 USD). The average correction for unemployment is the largest (-4,150 USD) followed by longevity (-2,300 USD) and the inequality penalty related to the median household (-1,600 USD). The cross-country correlation between average income and MDLS is large and equal to 0.90. Although income is the main component of MDLS for all countries under study, moving to MDLS as a broader welfare measure changes countries' positions in the large majority of cases (Figure 6). The largest changes in rank are observed for Greece and Korea due to high and low unemployment respectively, followed by Spain, Ireland, Japan and Estonia, where both unemployment and longevity play a role.

Figure 3. Living standards of the median household in 2013



Note: Living standards are expressed in 2005 USD.

Figure 4. Changes in country rankings when moving from income to MDLS

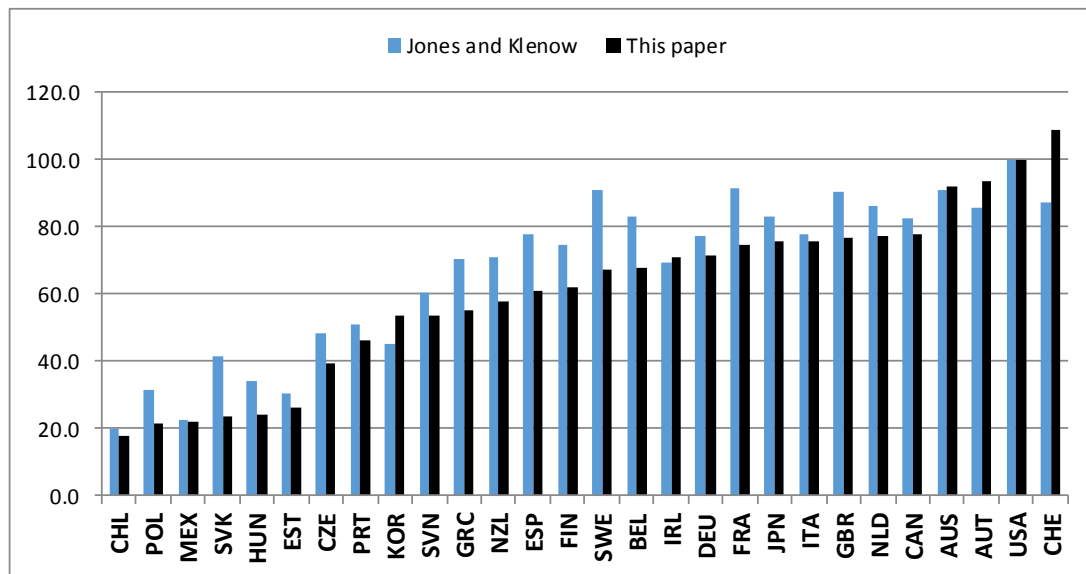


Note: The change in country ranking is calculated as the rank in income minus the rank in living standards.

Figure 7 compares our welfare measure to the results by Jones and Klenow (2016) for the year 2007, the year with the largest number of common country points. The two welfare measures are highly correlated (0.926), though results differ from some countries such as Switzerland, Sweden, France, Spain or Greece. These differences may stem from differences in the theoretical

frameworks (e.g. Jones and Klenow use leisure on top of income, mortality and inequality) and the different sources of data. Overall, however, the two welfare measures are rather consistent. Also, one advantage of our approach is to provide a set of heterogeneous shadow prices for unemployment and longevity, which could be used for policy evaluation within countries. This analysis is left aside for future research.

Figure 5. Comparison with Jones and Klenow's measure of welfare 2007



Note: Living standards and Jones and Klenow's measure of welfare are expressed in relative terms with respect to the US in 2007.

5.2. Well-being during the Great Recession

Finally, we calculate the growth rates of the proxy \widetilde{MDLS}_j and of its four components between 2008 and 2013 for the median household and the households at the 20th and 80th percentiles of the equivalent income distribution. Table 4 provides a comparison of GDP growth with MDLS growth over the period. Focusing first on the median household ($\tau = 0.89$), in 20 OECD countries out of 28, MDLS growth has been lower than GDP growth (the OECD averages being respectively -2.1% and 0.0% per year). Longevity has been the main positive driver of MDLS growth with an average contribution of 1.5% per year. Income has had a slightly negative contribution (-0.2%), while the rise in inequality has decreased MDLS by 0.4% per year. The rise

in unemployment has had a major negative impact with an annual decrease of 3% in MDLS. In countries most hardly hit by the crisis, soaring unemployment has created a wedge between GDP growth and MDLS growth, which were respectively equal to -1.8% and -14.7% in Spain, -6% and -26.5% in Greece, -1.3% and -8% in Ireland, -1.3% and -7.9% in Portugal. A major finding of this paper is that GDP growth has failed to reflect the social cost of the Great Recession and the magnitude of the loss in living standards for the median household, especially in hardly hit countries.

Turning to the living standards of households situated at the 20th percentile, Table 4 shows a huge decrease in living standards by 5.3% annually, which reflects the strong negative contribution of the inequality component (-3.6% per year). Rising unemployment and deteriorating economic conditions have primarily hit households at the bottom of the living standards distribution, entailing a gap of 3.6% per year in the growth of average living standards and the growth of living standards of poor households. In Greece and Spain, living standards have been slashed by 25% and 15% respectively every year for 5 years. In Spain, living standards of households at the 20th percentile have fallen from 5,200 USD in 2008 to 1,360 USD in 2013, and in Greece they have fallen from 5,000 USD to 940 USD between 2009 and 2013. Those figures reflect the huge increase in unemployment over the period (+15 percentage points in Spain and +19 percentage points in Greece) as well as the strong valuation of unemployment for the unemployed, whose shadow price is at least twice larger than the average shadow price as mentioned before.

Finally, relatively rich households situated at the 80th percentile have suffered relatively less as living standards have decreased by ‘only’ 1.4% per year. The contribution of the inequality component has been positive (0.29%), implying that the gap between the average and 80th percentile’s equivalent incomes has narrowed. There are only 3 countries in which poor households have fared better than rich ones in terms of living standards’ growth, namely Korea, Slovak Republic and Switzerland. In Greece and Spain, the gap in the annual growth rates between the two groups was about 15%, meaning that inequality in living standards has exploded in those 2 countries.

Table 4. Annual growth of living standards of households during the economic crisis

	Period		Economic growth Annual rate	Growth of living standards								
	Start	End		Annual rate			Contributions					
				Median ($\tau=0.89$)	Percentile 20 ($\tau=3.36$)	Percentile 80 ($\tau=-1.9$)	Average household income	Unemployment	Longevity	Median ($\tau=0.89$)	Inequality Percentile 20 ($\tau=3.36$)	Percentile 80 ($\tau=-1.9$)
Australia	2008	2013	0.95	0.29	-0.69	-0.07	0.72	-1.37	0.91	0.02	-0.96	-0.33
Austria	2008	2013	0.05	-1.63	-4.52	-0.54	-0.70	-1.01	0.58	-0.50	-3.39	0.59
Belgium	2008	2013	-0.37	-1.82	-4.12	-1.08	-0.59	-1.65	0.78	-0.35	-2.65	0.38
Canada	2008	2013	0.36	1.62	-1.21	1.74	1.16	-1.20	1.84	-0.18	-3.01	-0.06
Chile	2008	2013	2.97	7.45	6.61	7.80	5.75	0.81	0.97	-0.08	-0.92	0.27
Czech Rep.	2010	2013	0.16	-0.15	-0.33	-0.05	-1.71	0.10	1.50	-0.04	-0.22	0.05
Estonia	2008	2013	0.39	-0.48	-3.56	0.31	-0.70	-3.16	3.81	-0.43	-3.51	0.36
Finland	2008	2013	-1.65	-0.35	-2.23	-0.23	0.40	-1.88	1.25	-0.11	-1.99	0.00
France	2008	2013	-0.16	-2.11	-5.44	-0.54	-0.01	-2.45	1.01	-0.65	-3.98	0.92
Germany	2008	2013	0.45	2.56	0.01	3.25	0.67	1.02	1.20	-0.34	-2.89	0.35
Greece	2009	2013	-5.98	-26.47	-39.84	-24.26	-9.23	-17.14	1.63	-1.72	-15.09	0.49
Hungary	2009	2013	0.98	-0.16	-6.19	1.31	-0.03	-1.14	1.84	-0.83	-6.85	0.65
Ireland	2008	2013	-1.32	-7.96	-13.75	-6.41	-1.91	-6.99	1.79	-0.84	-6.63	0.70
Italy	2008	2013	-1.99	-6.59	-9.83	-5.63	-2.47	-4.73	1.11	-0.50	-3.74	0.46
Japan	2008	2013	0.46	1.87	0.34	2.37	1.21	-0.33	1.23	-0.24	-1.76	0.27
Korea	2010	2013	2.40	4.83	6.10	4.27	1.92	0.77	1.89	0.25	1.53	-0.31
Mexico	2008	2012	0.70	-1.82	-4.53	-0.29	0.07	-2.37	1.21	-0.73	-3.44	0.79
Netherlands	2008	2013	-1.03	-3.30	-5.69	-3.27	-1.05	-3.01	0.95	-0.18	-2.57	-0.16
New Zealand	2008	2012	0.98	0.06	-3.74	0.16	1.94	-3.36	1.81	-0.33	-4.13	-0.23
Poland	2009	2012	3.30	-0.53	-5.93	0.50	0.88	-3.68	2.87	-0.60	-6.00	0.43
Portugal	2008	2013	-1.35	-7.86	-13.00	-7.48	-1.58	-7.63	1.80	-0.44	-5.59	-0.06
Slovak Republic	2010	2013	2.00	1.14	1.67	0.69	-0.59	-0.02	1.65	0.10	0.63	-0.35
Slovenia	2009	2013	-0.68	-3.97	-6.26	-3.84	-1.68	-4.10	1.94	-0.14	-2.42	-0.01
Spain	2008	2013	-1.78	-14.68	-24.93	-10.68	-1.93	-12.80	2.07	-2.03	-12.28	1.97
Sweden	2008	2013	0.01	0.31	-3.02	1.41	1.94	-1.69	0.53	-0.48	-3.81	0.63
Switzerland	2009	2012	1.11	2.00	2.92	1.05	0.73	-0.06	1.04	0.29	1.21	-0.66
United Kingdom	2008	2013	-0.44	-1.17	-2.74	-0.24	-0.15	-2.42	1.71	-0.30	-1.87	0.63
United States	2008	2013	0.38	-0.36	-3.71	0.62	0.47	-1.63	1.30	-0.50	-3.85	0.48
Average			0.03	-2.12	-5.27	-1.40	-0.23	-2.97	1.51	-0.42	-3.58	0.29

6. Conclusion

This paper has proposed a methodology to measure MDLS incorporating the effect of income, unemployment and longevity while accounting for heterogeneous preferences. It confirms the result of other studies showing that this new measure is correlated with GDP, but is nevertheless quite different. MDLS incorporate population preferences over unemployment estimated from SWB surveys as well as a value of longevity calibrated from a model with intertemporal utility. The combination of survey-based and model-based valuations appears useful in order to take account of the fact that SWB surveys provide information about flow utility rather than lifetime utility. The study reveals that the rate of economic growth poorly reflects the severity of the Great Recession in living standards terms, especially in hardly hit countries and for poorer households. On average across OECD countries, GDP per capita stagnated while living standards of poor households fell by 5.3% annually between 2008 and 2013.

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