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Author:	Aaberge, Rolf Mogstad, Magne Peragine, Vito
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Measuring Long-term Inequality of Opportunity

Rolf Aaberge, Magne Mogstad and Vito Peragine

Abstract

In this paper, we introduce a new family of rank-dependent measures of inequality and social welfare consistent with the Equality of Opportunity (EOp) principle. The proposed framework can be used to measure long-term as well as short-term EOp, depending on whether we let permanent income or snapshots of income form the basis of the analysis. Further, it allows for both an ex-ante and an ex-post approach to EOp. There is long-term ex-post inequality of opportunity if individuals who exert the same effort have different permanent incomes. In comparison, the ex-ante approach focuses on differences in the expected permanent income between groups of individuals with identical circumstances. To demonstrate the empirical relevance of a long-run perspective on EOp, we exploit a unique panel data from Norway on individuals' incomes over their working lifespan. This allows us to examine how well analysis of opportunity inequality based on snapshots of income approximate the results based on permanent income.

1. Introduction

Over the last decades, increasing discontent has been expressed with distributional analysis based on observations of income for a single year. The reason is twofold: (i) transitory income components, and (ii) life-cycle variation in income. On the one hand, inequality in annual income is expected to overestimate the extent of long-term income inequality, since idiosyncratic shocks to income average out over time. On the other hand, measuring income early (late) in individuals' working lifespan is expected to understate (overstate) long-term income inequality, as individuals with high permanent income tend to be those with high income growth. This has led to a spur of research on inequality and social welfare in long-term or permanent income according to the Equality of Outcome (EO) principle.[1]

At the same time, a large and growing literature has addressed the question: what exactly should be equalized? Egalitarian theories of justice, especially after Sen's (1980, 1985, 1992) influential work, emphasize equality of opportunity (EOp) rather than equality of income. The distinction between functionings and capabilities made by Sen is closely related to the EOp concept, although the notion of individual responsibility plays a different role in Sen (1980, 1985, 1992) than in the theories of Dworkin (1981,a,b), Arneson (1989), Cohen (1989) and Roemer (1993).[2] Following Roemer (1993, 1998) and Fleurbaey (1995), economists have explored different ways of evaluating income distributions according to the EOp principle. However, snapshots of income still form the basis of EOp analyses.[3] The aim of this paper is to bridge these two strands of the literature by introducing and applying a general framework for evaluating long-term income distributions according to the EOp principle.

We follow the basic idea of the EOp literature closely in assuming that individuals' outcomes arise from two different types of variables: variables which they should not be held responsible for (circumstances), and variables which belong to the sphere of individuals' responsibility (effort). Once this basic partition has been made, the concept of EOp can be decomposed into two distinct ethical principles: *the Compensation Principle*, which states that differences in outcomes due to circumstances are ethically unacceptable and should be compensated, and *the Reward Principle*, which states that differences due to effort are to be considered ethically acceptable and do not justify any redistribution.

The existing literature on the measurement of opportunity inequality has explored two main approaches to measure opportunity inequality,[4] namely the ex-ante and the ex-post approach.[5] The two approaches rely on different definitions of EOp and embody the above ethical principles in

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[1] See e.g. Shorrocks (1978), Chakravarty et al. (1985), Maasoumi and Trede (2001), and Aaberge and Mogstad (2009).

[2] See Fleurbaey (2009) for a discussion.

[3] A notable exception is Bourguignon et al. (2007), which will be discussed in detail below.

[4] See, among others, Peragine (2002, 2004a,b), Bourguignon et al. (2003), Dardanoni et al. (2006), Lefranc et al. (2006), Ferreira and Guignoux (2008), Peragine and Serlenga (2008) and Checchi and Peragine (2009).

[5] Notice that the ex ante and the ex post approaches can be incompatible (see Fleurbaey and Peragine, 2010).

different ways. According to the ex-ante approach, there is equality of opportunity if the set of opportunities is the same for all individuals, regardless of their circumstances. Thus, inequality of opportunity is reduced if inequality between individual opportunity sets decreases. This approach partitions the population into different types, where each type is formed by individuals endowed with the same set of circumstances. The type-specific outcome distribution is interpreted as the opportunity set of individuals with the same circumstances. Accordingly, it focuses on inequality between types, and is neutral with respect to inequality within types.[6] In comparison, the ex-post approach states that there is equality of opportunity if all those who exert the same effort have the same outcome. Inequality of opportunity decreases if outcome inequality decreases among individuals with the same degree of effort. The ex post approach emphasizes inequalities within groups of individuals at the same effort levels, so-called tranches. Differences between tranches are interpreted as due to individual effort, and are considered as ethical acceptable.[7]

The approaches discussed above have been formulated in a static context, where current incomes (or some other snapshot of individual advantage) form the basis of the analysis. In this paper, we propose a framework that can be used to measure short-term or long-term EOp, depending on whether we let snapshots of income or permanent income form the basis. Our framework allows for both an ex-ante and an ex-post approach to EOp. There is long-term ex-post inequality of opportunity if individuals who exert the same effort have different permanent incomes. In comparison, the ex-ante approach focuses on the expected permanent income for groups of individuals with identical circumstances. Hence, the ex-ante approach pays attention to inequalities in expected permanent income between different types of individuals.

Measuring long-term EOp requires aggregation in two steps. The first step consists of aggregating the income stream of each individual into an interpersonal comparable measure of permanent income. To this end, we follow Aaberge and Mogstad (2009) in using a measure of permanent income which incorporates the costs of and constraints on making inter-period income transfers. The second step consists of aggregating individuals' permanent incomes into EOp measures of social welfare and inequality. Specifically, we employ an axiomatic approach to justify the introduction of a generalized family of rank-dependent measures of ex post as well as ex ante opportunity inequality and social welfare. When measuring short term EOp, the measures of opportunity and social welfare are simply applied to the distributions of annual income.

The gain in studying long-run EOp is twofold. First of all, by focusing on the distribution of permanent income, we eliminate life-cycle bias in opportunity inequality. In particular, we overcome the problem of separating out how much of the life-cycle variation in annual income that is due to circumstances, and how much that is due to effort. Secondly, by letting permanent income form the

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[6]This is the approach proposed by van de Gaer (1993) and Peragine (2004a) and used by Ferreira and Guignoux (2008) and Lefranc et al. (2006).

[7] This is the approach proposed by Roemer (1993, 1998), and used by Roemer et al. (2003) and Aaberge and Colombino (2010). Peragine (2002, 2004b) and Checchi and Peragine (2010) explore both the ex ante and the ex post approaches.

basis for the EOp analysis, we eliminate opportunity inequality due to idiosyncratic shocks to income. This can be important since such transitory income components average out over time, and therefore do not call for compensation according to the Compensation Principle. The long-term perspective on EOp can also be important from a policy perspective, by highlighting that inequality opportunities may accumulate over time, if (dis)advantages beget (dis)advantages. Moreover, it allows for the possibility that groups viewed as (dis)advantaged based on snapshots of income may change positions over time. In such cases, policies that minimize snapshots of opportunity inequality may deviate from those optimizing equality of opportunity in long-term income.

To demonstrate the empirical relevance of our EOp framework, we exploit a unique panel data set from Norway on individuals' incomes over their working lifespan. This allows us to examine how well analyses of inequality and social welfare based on snapshots of income approximate the results based on long-term income. Further, we investigate how sensitive distributional comparisons between the EOp and the EO principle of justice are to the choice of accounting period of income. Finally, we assess to what extent the ex-ante and the ex-post approaches yield different pictures of opportunity inequality and social welfare.

This paper proceeds as follows. Section 2 deals with the problem of aggregating (permanent or snapshots of) income across individuals into measures of (long-term or short-term) income inequality and social welfare. Section 3 introduces a new family of rank-dependent measures of ex-ante and ex-post opportunity inequality and social welfare, after which we clarify the differences between our framework and the method for measuring long-term EOp proposed by Bourguignon et al. (2007). Section 4 reports the empirical results, and Section 5 concludes.

2 Rank-dependent measures of inequality and social welfare

Empirical analyses of inequality in income distributions are conventionally based on the Lorenz curve and associated summary measures of inequality, which means that concern is directed towards income shares without taking account of differences in mean incomes. To summarize the information content of the Lorenz curve and to achieve rankings of intersecting Lorenz curves, the standard approach is to employ the Gini coefficient in combination with one or two inequality measures from the Atkinson family or the Theil family. However, since they have distinct theoretical foundations, it is difficult to evaluate their capacity as complementary measures of inequality. An alternative approach, which will be chosen in this paper, is to use measures of inequality which can be shown to have a common theoretical justification as criteria for ranking Lorenz curves.

The Lorenz curve L for a cumulative income distribution F with mean μ is defined by

(2.1) $$L(v) = \frac{1}{\mu} \int_0^v F^{-1}(t)\,dt, \quad 0 \le v \le 1,$$

where $F^{-1}(t) = \inf\{x : F(x) \geq t\}$ is the left inverse of F. Thus, the Lorenz curve $L(v)$ shows the share of total income received by the poorest $100\,v$ per cent of the population. Note that F can either be a discrete or a continuous distribution function. Although the former is what we actually observe, the latter often allows simpler derivation of theoretical results and is a valid large sample approximation. Below F will be assumed to be a continuous distribution function.

Under the restriction of equal mean incomes, the problem of ranking Lorenz curves formally corresponds to the problem of choosing between uncertain prospects. This relationship has been utilized by, for example, Atkinson (1970) to characterize the criterion of non-intersecting Lorenz curves in the case of distributions with equal mean incomes. This was motivated by the fact that in cases of equal mean incomes, the criterion of non-intersecting Lorenz curves is equivalent to second-degree stochastic dominance, which means that the criterion of non-intersecting Lorenz curves obeys the Pigou-Dalton principle of transfers. To perform inequality comparisons with Lorenz curves we can either limit our attention to distributions with equal means, or alternatively simply abandon the assumption of equal means and consider distributions of relative incomes.[8] The latter approach normally forms the basis of empirical studies of income inequality and will be used in this paper.

Let $\boldsymbol{L}$ denote the family of Lorenz curves, and let a social planner's ranking of members of $\boldsymbol{L}$ be represented by a preference ordering $\succeq$, which will be assumed to satisfy the following basic axioms:

Axiom 1 (Order). $\succeq$ *is a transitive and complete ordering on $\boldsymbol{L}$.*

Axiom 2 (Dominance). *Let $L_1, L_2 \in \boldsymbol{L}$. If $L_1(v) \geq L_2(v)$ for all $v \in [0,1]$ then $L_1 \succeq L_2$.*

Axiom 3 (Continuity). *For each $L \in \boldsymbol{L}$, the sets $\{L^* \in \boldsymbol{L} : L \succeq L^*\}$ and $\{L^* \in \boldsymbol{L} : L^* \succeq L\}$ are closed (w.r.t. L_1-norm).*

Given the above continuity and dominance assumptions for the ordering $\succeq$, Aaberge (2001) demonstrated that the following axiom,

Axiom 4 (Independence). *Let L_1, L_2 and L_3 be members of $\boldsymbol{L}$ and let $\alpha \in [0,1]$. Then $L_1 \succeq L_2$ implies*

$$\alpha L_1 + (1-\alpha) L_3 \succeq \alpha L_2 + (1-\alpha) L_3,$$

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[8] See e.g. Smith's (1979) discussion of necessities and Sen's (1992) discussion of relative deprivation.

characterizes the rank-dependent family of inequality measures $J_{EO,p}$ defined by[9]

$$(2.2) \qquad J_{EO,p}(L) = 1 + \int_0^1 L(v)dp(v) = 1 - \frac{1}{\mu}\int_0^1 p(v)F^{-1}(v)dv,$$

where subscript *EO* refers to inequality of outcomes, and subscript *p* refers to a positive and non-increasing weight-function defined on the unit interval such that $\int p(v)dv = 1$ and $p(1) = 0$. The latter condition ensures that $J_{EO,p}$ has the unit interval as its range. Note that *p* can be interpreted as a preference function of a social planner who assigns weights to the incomes of the individuals in accordance with their rank in the income distribution. Therefore, the functional form of *p* reveals the attitude towards inequality of a social planner who employs J_p to judge between Lorenz curves.

As was recognized by Ebert (1987), the justification of the social welfare function associated with $J_{EO,p}$ can be made in terms of value judgement of the trade-off between the mean and (in)equality in the distribution of income

$$(2.3) \qquad W_{EO,p} = \mu(1 - J_{EO,p}) = \int_0^1 p(v)F^{-1}(v)dv,$$

where $W_{EO,p}$ can be interpreted as the equally distributed (equivalent) level of income. Yaari (1988) provides an alternative axiomatic justification for using the latter term of (2.3) as a criterion for ranking distribution functions *F*, rather than as a trade-off between mean and equality. Note that the normative justification for the family of rank-dependent measures of inequality J_p is analogous to the justification for Atkinson's expected utility type of inequality measures. The essential differences between these two approaches for measuring inequality and social welfare arise from the independence axioms. Whilst the expected utility independence axiom requires that the ordering of distributions of individual welfare is invariant with respect to identical mixing of the distributions being compared, the rank-dependent independence axiom requires that the ordering is invariant with respect to identical mixing of Lorenz curves (or identical mixing of the inverses of distributions) being compared. For further discussion, see Yaari (1988) and Aaberge (2001).

When specifying $p(v) = 2(1-v)$, it follows that $J_{EO,p}$ becomes equal to the Gini coefficient. To complement the information provided by the Gini coefficient, it appears attractive to use members of the rank-dependent family $C_k = J_{EO,p_k}, k = 1,2,...,$ where

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[9] Mehran (1976) introduced an alternative version of (2.2) based on descriptive arguments. For alternative motivations of the J_P-family and various subfamilies we refer to Donaldson and Weymark (1980, 1983), Weymark (1981), Ben Porath and Gilboa (1994), and Aaberge (2000, 2007).

$$(2.4) \qquad p_k(v) = \begin{cases} -\log v, & k = 1 \\ \dfrac{k}{k-1}\left(1 - v^{k-1}\right), & k = 2,3,.... \end{cases}$$

As indicated by Aaberge (2000, 2007), the family $\{C_k : k = 1, 2, ...\}$ uniquely determines the Lorenz curve L. Thus, by restricting to this family of inequality measures no information is lost. However, in applied work it is for practical reasons convenient to employ a few measures of inequality. To this end, Aaberge (2007) draws on standard statistical practice to justify the use of C_1 (the Bonferroni coefficient), C_2 (the Gini coefficient) and C_3 as a basis for summarizing the inequality information in an income distribution, and the associated social welfare functions W_1, W_2 and W_3 to assess the trade-off between efficiency and (in)equality. Moreover, these three measures of inequality also prove to complement each other with regard to sensitivity to transfers at the lower, the central and the upper part of the income distribution.

3 Generalized families of rank-dependent measures of opportunity inequality and social welfare

This section introduces the generalized family of rank-dependent measures of ex post as well as ex ante opportunity inequality and social welfare. The proposed framework is valid for measuring short-term as well as long-term EOp, depending on whether we let snapshots of income or permanent income form the basis of the analysis. But first, we will discuss the setup.

3.1 The setup

We consider a continuum of individuals. Individuals are described by a list of factors, which can be partitioned into two different classes: factors beyond the individual control, represented by a person's vector of circumstances c belonging to an abstract and finite set $\Omega = \{c^1, \ldots, c^m\}$; and factors for which the individual is fully responsible, represented by a scalar variable[10] e (effort), $e \in \Theta \subseteq \Re_+$. We observe c but not e. Assume that the circumstances of an individual consist of factors that are constant over time, such as family background.[11]

By partitioning the population into m subpopulations, each representing a class identified by variable c, we call "type" i the set of individuals whose set of circumstances is c_i. Within type i, there will be a cumulative distribution of (permanent or period-specific) income F_i with mean μ_i and a population share denoted by q_i such that $F(x) = \sum_{i=1}^{m} q_i F_i(x)$, where $\sum_{i=1}^{m} q_i = 1$. A second partition is

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[10] In the analysis it could equally well be represented as a vector.

[11] The same assumption is made by Bourguignon et al. (2007).

based on the responsibility e. For all degrees of responsibility $e \in \Theta$ we call *tranche e* the set of individuals whose responsibility level is equal to e.

Let the permanent income be defined as $y = g(c,e)$, where g is a "production function" that describes the relationship between an individual's permanent income and her effort and circumstances. We do not know the form of the function g; we know, however, that it is increasing in e and the same for all individuals. Roemer (1993, 1998) suggests taking the quantile of the effort distribution of the type an individual belongs to as an inter-type comparable measure of effort, which due to the monotonicity of the income function, will correspond to the quantile in the income distribution of the type. We adopt this solution and hence we say that all individuals at the s quantile of their respective type income distributions have the same effort. In the analysis of long-term EOp, we define the tranche s as the subset of individuals whose permanent incomes $F_1^{-1}(s), F_2^{-1}(s), ..., F_m^{-1}(s)$ are at the sth quantile of their respective type permanent income distributions.

As discussed above, using current income to measure income inequality raises two problems: (i) transitory income components, and (ii) life-cycle variation in income. In particular, these problems might lead to a distorted assessment of opportunity inequality. To see this, suppose we define *tranche s* as the subset of individuals whose period-specific incomes $F_1^{-1}(s), F_2^{-1}(s), ..., F_m^{-1}(s)$ are at the sth quantile of their respective type period-specific income distributions. Let the individual income of period t be defined as

(3.1) $$y_t = f_t(c, e_t, u_t)$$

where f_t is a time-specific "production function" that relates an individual's income of period t to circumstances c, (accumulated) effort e_t, and idiosyncratic shocks to income u_t. We observe c but not e_t or u_t. Furthermore, we assume that the shape of the function f_t is unknown; we only know that it is increasing in e_t and the same for all individuals.

Note that (3.1) departs from the standard framework used in static analysis of opportunity inequality in two ways. First, it distinguishes between two types of factors that are outside the individual responsibility; circumstances c that are constant over time, and idiosyncratic factors u_t that are time-varying but assumed to average out over time. Given the circumstances c at birth, it is impossible to disentangle the effect of e_t and the effect of u_t. Second, (3.1) acknowledges that the time-specific production function f_t may differ from the production function of lifetime income g, reflecting that the effects of circumstances and effort may vary over the life-cycle. For both these reasons, the quantiles of period-specific income distributions of each type can in general not be considered inter-type comparable measures of effort. In the empirical analysis, we will examine how

well analysis of opportunity inequality based on period-specific income approximates the results based on permanent income.

3.2 Ex post approach

Let $\tilde{F}_1^{-1}(s) \leq \tilde{F}_2^{-1}(s) \leq ... \leq \tilde{F}_m^{-1}(s)$ be the ordering of incomes $F_1^{-1}(s), F_2^{-1}(s),...,F_m^{-1}(s)$ across types of tranche *s*. Since the type-specific income distributions might intersect, note that the type ordering by income might change across quantiles; i.e. $\tilde{F}_k^{-1}(s)$ and $\tilde{F}_k^{-1}(v)$ might represent different types. Accordingly, the proportion of people associated with the lowest income, the second lowest income, etc. might change across quantiles. Thus, let $q_i(s)$ be the population share associated with the individual having rank *i* at tranche *s*. Moreover, let $a_j(s) = \sum_{i=1}^{j} q_i(s)$ and $b_j(s) = 1 - a_j(s)$.

The quantile-specific mean $\mu(s)$ and *EOp* Lorenz curve $L_{EOp}(\cdot; s)$ are defined by

$$(3.2) \qquad \mu(s) = \sum_{i=1}^{m} q_i(s)\tilde{F}_i^{-1}(s) = \sum_{i=1}^{m} q_i F_i^{-1}(s)$$

and

$$(3.3) \qquad L_{EOp}(a_j(s); s) = \frac{\sum_{i=1}^{j} q_i(s)\tilde{F}_i^{-1}(s)}{\mu(s)}, \quad j = 1, 2,...,m, \quad 0 \leq s \leq 1.$$

Inserting (3.2) and (3.3) in (2.2) yields the following family of *quantile-specific rank-dependent measures of inequality of opportunity*

$$(3.4) \qquad J_{EOp,p}(s) = 1 - \frac{1}{\mu(s)} \sum_{j=1}^{m} q_j(s) p(a_j(s)) \tilde{F}_j^{-1}(s),$$

where *p* satisfies the normalization condition $\sum_{j=1}^{m} q_j(s) p(a_j(s)) = \sum_{j=1}^{m} q_j p(a_j) = 1$. Thus, $J_{EOp,p}(s)$ measures income inequality due to different circumstances for those who belong to tranche *s* and thus have exerted the same degree of effort. According to the *Compensation Principle*, $J_{EOp,p}(s)$ can be interpreted as the inequality of opportunity in tranche *s*.

In order to define an overall measure of opportunity inequality, a method for aggregating the tranche-specific inequality measures is required. Treating inequality between tranches as ethically

acceptable, we can aggregate across the tranche-specific inequality measures to obtain the following family of rank-dependent measures of overall inequality of opportunity

$$(3.5) \qquad J_{EOp,p} = \int_0^1 J_{EOp,p}(s)\,ds \, .$$

Thus, inequality of opportunity decreases if outcome inequality decreases among the individuals with the same degree of effort. The social welfare function associated with $J_{EOp,p}$ is given by

$$(3.6) \qquad W_{EOp,p} = \mu(1 - J_{EOp,p})$$

where $\mu_i = \int_0^1 F_i^{-1}(t)\,dt$ and $\int_0^1 \mu(s)\,ds = \sum_{i=1}^m q_i \mu_i = \mu \, .$

Note that p_e defined by

$$(3.7) \qquad p_e(a_j(s)) = \begin{cases} \dfrac{1}{q_1(s)}, & j = 1 \\ 0, & j = 1,2,...,m \end{cases}$$

represents the upper limit of inequality aversion exhibited by the family of non-increasing weight functions p. By inserting (3.7) in (3.4), (3.5) and (3.6) we get

$$(3.8) \qquad J_{EOp,e}(s) = 1 - \frac{\tilde{F}_1^{-1}(s)}{\mu(s)} = 1 - \frac{\min\limits_{1 \le i \le m} F_i^{-1}(s)}{\mu(s)} \, ,$$

$$(3.9) \qquad J_{EOp,e} = 1 - \int_0^1 \frac{\min\limits_{1 \le i \le m} F_i^{-1}(s)}{\mu(s)}\,ds$$

and

$$(3.10) \qquad W_{EOp,e} = \mu \int_0^1 \frac{\min\limits_{1 \le i \le m} F_i^{-1}(s)}{\mu(s)}\,ds \, .$$

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By inserting the specification $p(a_j(s)) = \dfrac{(1 - a_j(s))}{\sum\limits_{i=1}^{m} q_i b_i}$ where $b_j = 1 - a_j$ in (3.4) and (3.5), we get the

Gini version of $J_{EOp,p}(s)$ defined by

$$(3.11) \qquad G_{EOp}(s) = 1 - \frac{\sum\limits_{j=1}^{m} q_j(s) b_j(s) \tilde{F}_i^{-1}(s)}{\mu(s) \sum\limits_{j=1}^{m} q_j b_j} ,$$

and

$$(3.12) \qquad G_{EOp} = \int_0^1 G_{EOp}(s)\,ds = 1 - \frac{\int_0^1 \left[\sum\limits_{j=1}^{m} q_j(s) b_j(s) \dfrac{\tilde{F}_i^{-1}(s)}{\mu(s)} \right] ds}{\sum\limits_{j=1}^{m} q_j b_j}$$

3.3 Relationship between outcome inequality and ex post opportunity inequality

An appropriate decomposition method for inequality of outcome with respect to inequality of opportunity and inequality of effort should satisfy the following conditions:

(i)   If there is complete EOp, i.e. $J_{EOp,p} = 0$, then $J_{EO,p}$ measures inequality of effort.

(ii)   If every individual have exerted the same degree of effort, i.e. $\mu(s) = \mu$ *for all s*, then

$J_{EO,p}$ measures inequality of opportunity, i.e., $J_{EO,p} = J_{EOp,p}$.

Inserting for the type-specific decomposition of the outcome distribution F; i.e. $F(x) = \sum\limits_{i=1}^{m} q_i F_i(x)$, in

(2.2) yields

$$(3.13) \qquad J_{EO,p} = 1 - \frac{1}{\mu} \int_0^1 \left[\sum_{j=1}^{m} q_j\, p(F(F_j^{-1}(s)))\, F_j^{-1}(s) \right] ds .$$

Next, by ordering types by increasing income for each tranche and by observing that

$a_j(s) = F(\tilde{F}_j^{-1}(s))$, we get the following alternative expression for (3.13)

$$(3.14) \qquad J_{EO,p} = 1 - \frac{1}{\mu} \int_0^1 \left[\sum_{j=1}^{m} q_j(s)\, p(a_j(s))\, \tilde{F}_j^{-1}(s) \right] ds .$$

Inserting for (3.4) in (3.14) yields

$$(3.15) \qquad J_{EO,p} = 1 - \frac{1}{\mu} \int_0^1 \mu(s)(1 - J_{EOp,p}(s))ds = J_{EOp,p} + \int_0^1 (1 - \frac{\mu(s)}{\mu})(1 - J_{EOp,p}(s))ds .$$

It follows immediately from (3.15) that $\mu(s) = \mu$ for all s implies that $J_{EO,p} = J_{EOp,p}$, which means that condition (ii) is fulfilled. When there is complete equality of opportunity, i.e. $J_{EOp,p}(s) = 0$ for all s, then $J_{EOp,p}$ defined by (3.15) becomes a measure of inequality of effort as required by condition (i). Accordingly, the decomposition (3.15) satisfies the requirements (i) and (ii). Note that the latter term of the decomposition (3.15) can be considered as a measure of inequality of effort, where the contribution from each tranche depends on the tranche-specific inequality of opportunity.

The decomposition (3.15) characterizes the relationship between EO and ex-post EOP, and demonstrates that it is meaningful to treat $J_{EOp,p}$ as a share of $J_{EO,p}$. Section 4 provides empirical estimates of the share of outcome inequality that is opportunity inequality, based on permanent as well as current incomes.

3.4 Ex ante approach

The ex ante approach focuses attention on the differences in the outcome prospects for individuals with identical circumstances. Accordingly, it focuses on inequality between types, and is instead neutral with respect to inequality within types. Hence, the ex ante approach is more focused on *inequalities between social groups*. Referring to the Principle of Utilitarian Reward introduced by Fleurbaey (2008), which states the social irrelevance of inequalities due to differences in effort, this implies that any equalizing transfer with identical circumstances should not change income inequality or social welfare, whatever the effort level of the persons involved. This means that an ex ante approach for measuring inequality of opportunity will solely depend on the type-specific mean incomes.

Assume that the type-specific means can be ordered in the following way $\mu_1 \leq \mu_2 \leq ... \leq \mu_m$ and that q_i is associated with μ_i. Now consider a hypothetical distribution where every individual of type i is assumed to receive the same income equal to the type-specific mean μ_i. In this case, the type-specific distributions are defined by

$$(3.16) \qquad F_i^*(x) = \begin{cases} 0, x < \mu_i \\ 1, x \geq \mu_i \end{cases}, i = 1,2,...,m,$$

whereas the distribution function of the society is defined by $F^*(x) = \sum q_i F_i^*(x)$, and the associated Lorenz curve[12] is defined by

$$(3.17) \qquad L^*(a_j) = \frac{\sum_{i=1}^{j} q_i \mu_i}{\sum_{i=1}^{m} q_i \mu_i}, \; j = 1, 2, \ldots, m.$$

Moreover, the associated rank-dependent family of inequality measures is defined by

$$(3.18) \qquad J_{EO_p,p}^* = 1 - \frac{\sum_{j=1}^{m} q_j p(a_j) \mu_j}{\mu \sum_{j=1}^{m} q_j p(a_j)}\; .$$

The $J_{EO_p,p}^*$ measure captures the inequality between types, which can be interpreted as ex ante inequality of opportunity. When p represents the Gini weight function, then J_p^* is given by

$$(3.19) \qquad G_{EO_p}^* = 1 - \frac{\sum_{j=1}^{m} q_j b_j \mu_j}{\mu \sum_{j=1}^{m} q_j b_j}\; .$$

The welfare function associated with $J_{EO_p,p}^*$ is defined by

$$(3.20) \qquad W_{EO_p,p}^* = \mu(1 - J_{EO_p,p}^*) = \frac{\sum_{j=1}^{m} q_j p(a_j) \mu_j}{\sum_{j=1}^{m} q_j p(a_j)}$$

By inserting (3.7) in (3.18) and (3.20) we get

$$(3.20) \qquad J_{EO_p,e}^* = 1 - \frac{\min\limits_{1 \le i \le m} \mu_i}{\mu}$$

and

$$(3.21) \qquad W_{EO_p,e}^* = \min_{1 \le i \le m} \mu_i\; .$$

Note that $W_{EO_p,e}^*$ corresponds to the EOp welfare function introduced by Roemer (1993, 1998), when applied to snapshots of income. When μ_i represents the mean permanent income of type i, then $W_{EO_p,e}^*$ coincides with the EOp welfare function introduced by Bourguignon et al. (2007), if and only if

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[12] The partial ordering based on the Lorenz curve below was characterized in Peragine (2002).

μ_i can be expressed as weighted average of the period-specific type mean incomes. As shown by Aaberge and Mogstad (2009), μ_i can be given such a representation only under the assumption of a perfect capital market, that is, when there are no liquidity constraints and the real interests on borrowing and savings are equal.

3.5 Relationship with previous approaches

In a pioneering paper, Bourguignon et al. (2007) propose an alternative approach for measuring long-term EOp, which is in line with an ex ante perspective. First, individuals are divided into types who are homogeneous with respect to exogenous circumstances at a given point in time, where the income distributions represent the opportunity set of individuals belonging to specific type in a given period. Second, for each type the opportunity sets are aggregated across time periods to obtain the long-tern opportunity sets. Finally, the distribution of long-term opportunity sets across types is evaluated. Bourguignon et al. (2007) discuss two different methods to do so.

As a first methodology, Bourguignon et al. (2007, p. 243) define EOp as follow: there is EOp if, in each period, there is not dominance between types. Hence, in each period of time, there should be equality or, as a weaker definition, non-dominance of opportunity sets. This can be seen as the natural extension of the approach used by Lefranc et al (2006) and Peragine and Serlenga (2008) to a long-term perspective. As second methodology, Bourguignon et al. (2007) propose a long-term extension of the utilitarian version of EOp, according to which there is EOp if the long-term expected value of all types are equal. They first use a utilitarian evaluation function which evaluates the opportunity set of each type at a given period t by its mean μ_{it}. Next, they introduce a *time aggregator* $\phi_t(\mu_{it})$, in order to obtain an expression of the long-term value of the opportunity set. Finally, to evaluate the distribution of long-term opportunity sets, they use the following social welfare function:

$$W(F) = \min_i \int_T \phi_t \mu_{it} dt .$$

As demonstrated above, we take a different approach to long-term EOp. In our methodology, the aggregation over time is performed for each individual. This allows us to be specific with respect to the underlying model of intertemporal choice that underlies the construction of the distribution of long-term income. Our approach is also more general in the sense that it allows both an ex ante and an ex post approach to long-term EOp. Moreover, both in the ex ante and in the ex post approach, our framework allows for different attitudes towards inequality aversion.

4 Empirical analysis

This section implements the proposed framework for measuring opportunity inequality and social welfare, using panel data from Norway on individual income over the working lifespan.

Data. Our empirical analysis utilizes several register databases provided by Statistics Norway. The data include a rich longitudinal dataset containing records for every Norwegian from 1967 to 2006. The variables captured in this dataset include individual demographic information (sex, year and municipality of birth, number of siblings) and socio-economic data (income and education). Importantly, the dataset includes personal identifiers, allowing us to link children to their parents. Our income measure is defined in close agreements with international recommendations (see Expert Group on Household Income Statistics, 2001), and incorporates annual wages, capital income, and all public cash transfers. We use the consumer price index to make earnings and incomes from different periods comparable. Throughout this paper we focus on individual income, because our family data does not allow us to link spouses living in cohabitation. Nor can we identify married spouses before the mid 1970s. Therefore, we cannot construct a measure of family or household income.

Our empirical analysis focuses on the 1942-1944 cohorts in order to ensure availability of data on income for more or less the entire working lifespan. In particular, these cohorts are between 23 and 25 years old in 1967 and between 62 and 64 in 2006.[13] Our analytical sample is restricted to males, given their role of breadwinner and primary wage-earner for these cohorts. Also, we exclude individuals whose information on annual income is missing. Finally, we drop observations where information on parents is missing. The final sample used in the analysis consists of 26 090 individuals.

Permanent income. As a measure of permanent income, we follow Aaberge and Mogstad (2009) in defining permanent income as the minimum annual expenditure an individual would need in order to be as well off as he could be by undertaking inter-period income transfers. To derive an interpersonal comparable measure of permanent income, they follow standard practice in assuming that inter-period income transfers are carried out in accordance with an intertemporal utility function that is common to all individuals. The common utility function, which is to be determined by the social planner based on his ethical value judgement, contains within it interpersonal comparability of both welfare levels and welfare differences, and can be viewed as a normative standard where individuals are treated symmetrically.[14] Specifically, the social planner is assumed to employ the conventional discounted utility model with perfect foresight, where preferences are intertemporal separable and additive.

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[13] Note that people are eligible for early retirement when they have become 62 years old. See Aaberge and Mogstad (2010) and Almås et al. (2010) for analyses of the time trends in inequality of outcome in Norway.

[14] The use of a common utility function is well-established in the public economic literature and has e.g. been proposed by Deaton and Muellbauer (1980) and Hammond (1991). It also forms the basis for the definition and measurement of a money-metric measure of utility in for example King (1983) and Aaberge et al. (2004).

The measure of permanent income proposed by Aaberge and Mogstad (2009) is compatible with a rather general structure on intertemporal preference and the credit market. Moreover, it can be considered as a money-metric measure of the utility level associated with the income stream for a given individual. In addition, it encompasses much used measures of permanent income, such as the average income and the annuity value of an income stream,[15] depending on the assumptions made about the intertemporal preferences and the credit market. In contrast to studies using average income or the annuity income as the permanent income, we allow for individual-specific interest rates on borrowing and savings in individuals' budget constraint. Thus, our measure of permanent income incorporates the cost of making inter-period income transfers, and hence account for the utility loss that may be associated with income fluctuations. The annual real interest rates on borrowing and savings are computed from Norwegian official statistics on interest rates on loans and deposits in commercial banks over the period 1967-2006.

We further assume that the instantaneous utility term of the intertemporal utility function belongs to the Bergson family, which is a much used specification of the instantaneous utility function in intertemporal choice theory (Davies and Shorrocks, 2000). In the results displayed below, we have set the rate of time preferences equal to 2 percent, which is approximately the same as the average real interest rates on borrowing and savings during the period of study. Finally, the intertemporal elasticity of substitution is assumed to be equal to 0.5, which is in the mid range of the estimates reported in previous research. In any case, the estimated inequality and social welfare are quite robust to substantial changes in the preference parameters.[16]

Circumstances and types. How to define the exact set of individual circumstances is a controversial question. Besides, in empirical work, observing the entire set of circumstances is clearly out of reach. In this paper, we focus on the dependence of individual opportunity on a restricted set of circumstances relating to family background. The role of families in determining socio-economic outcomes has been extensively discussed in economics as well as from other scientific perspectives. Parents influence their children via several channels: investment in their children's education, transmission of cultural values and social skills, and genetic endowments. As argued by Roemer (1998) and others, equality of opportunity requires compensating the influence of family background on individual outcomes, as it lies beyond individual responsibility and choice.

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[15] In Shorrocks (1978) as well as in most subsequent empirical studies of long-term inequality and income mobility according to the EOp principle, the average income is used as an approximation for permanent income.

[16] There is no consensus on the exact value of the intertemporal elasiticity of substitution. On the one hand, consistency of calibrated dynamic macroeconomic models with aggregate data requires relatively large value of this parameter. For example, Kydland and Prescott (1982) calibrated this value to be around 0.66. On the other hand, direct estimates of this parameter seem to suggest considerably lower values (see e.g. Hall, 1988). As a robustness check, we have used values of the intertemporal elasiticity ranging from .1 to 1.0, and values of the rate of time preferences ranging from 0 to .1. It is reassuring to find that our estimates of the measures of inequality and social welfare are quite robust to the changes in the preference parameters. The results are available on request.

A large body of empirical research shows a strong degree of intergenerational transmission, as parental characteristics like education and income are highly correlated with the outcomes of children along the same dimensions (see e.g. Solon 1999). As a starting point, we therefore split the sample by birth cohort and further divide it into three groups based on the highest educational attainment of the parents: compulsory school (grades 1-7), middle school (grades 8-10), and higher education (11 or more years of education). Next, we partition the sample according to an indicator for urbanity, equal to one if the individual was born in one of the five bigger cities of Norway (Oslo, Bergen, Trondheim, Stavanger, Drammen). The idea is that urbanity may proxy for neighbourhood effects, which may arise from varying local resource bases, such as availability of institutions like schools and childcare, as well as through social interaction in peer-groups, like attitudes and preference formation as well as the existence and enforcement of social norms.[17] Finally, we partition the sample according to an indicator for being raised in a large family, equal to one if the individual has two or more siblings. Drawing on the quantity-quality model of Becker and Lewis (1973), a large body of empirical research has shown a fairly strong association between number of siblings and children's outcome. In particular, it appears to be substantial negative effects of large sibship size (see Mogstad and Wiswall, 2010). Thus, to measure EOp, the sample is divided into 36 mutually exclusive and collectively exhaustive types based on circumstances. Within each type, we will allow for 100 tranches, that is, quantiles in the type-specific income distributions.

Ideally, we would have liked to use more information on family background to define the set of individual circumstances. However, the Norwegian register data is quite limited when it comes to family background of individuals born almost sixty years ago. Furthermore, our sample size restricts us for having more types, as we need a sufficient number of observations to identify the tranche of each type. Yet as a robustness check, we have replaced the grouping based on the highest educational attainment of the parents with a grouping based on fathers' earnings in 1960. Specifically, the individuals are divided into three groups of equal size based on the distribution of fathers' earnings: 33 percentile and below, 33-66 percentile, and 66 percentile and above. The data on fathers' earnings is collected from the 1960 census. We find that the estimated measures of inequality and social welfare change very little when using fathers' income instead of parental education to define the set of individual circumstances.[18]

Descriptive statistics. Table 1 displays the composition of the sample according to the variables used to define circumstances. We see that about 20 percent of the sample was born in urban areas, and about 45 percent come from a large family. Moreover, it is evident that about 53 percent has parents with only compulsory schooling, almost 30 percent has at least one parent who attended middle

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[17] Raaum et al. (2006) show significant effects of childhood neighbourhood on children's long-run outcomes.
[18] The results are available on request.

school, and less than 18 percent has at least one parent with higher education. It is further evident that the composition is fairly similar across cohorts.

Table 1. Descriptive statistics

| | Birth cohort | | | |
	42	43	44	42-44
Income (NOK)				
Period-specific income:				
1967-1968	157,338	150,824	141,850	149,191
1975-1976	240,324	239,477	233,859	237,538
1985-1986	244,191	244,085	239,967	242,518
1995-1996	277,150	283,057	279,302	279,926
2005-2006	311,261	329,562	328,496	323,958
Mean income (1967-2006)	251,824	258,005	249,505	252,955
Permanent income (1967-2006)	234,976	237,163	230,058	233,787
Circumstances (populaton shares)				
Urban	18.18	20.09	21.06	19.92
Large family	46.02	44.51	43.69	44.62
Parental Education				
Comp. school only	53.48	52.59	51.39	52.38
Middle school	29.00	29.39	30.67	29.77
Higher education	17.52	18.02	17.95	17.85
Sample size	7,399	8,572	10,119	26,090

Note: Period-specific income is defined as the average income over the two-year period.

Figure 1. Marginal distribution of permanent income for entire sample and by parental education
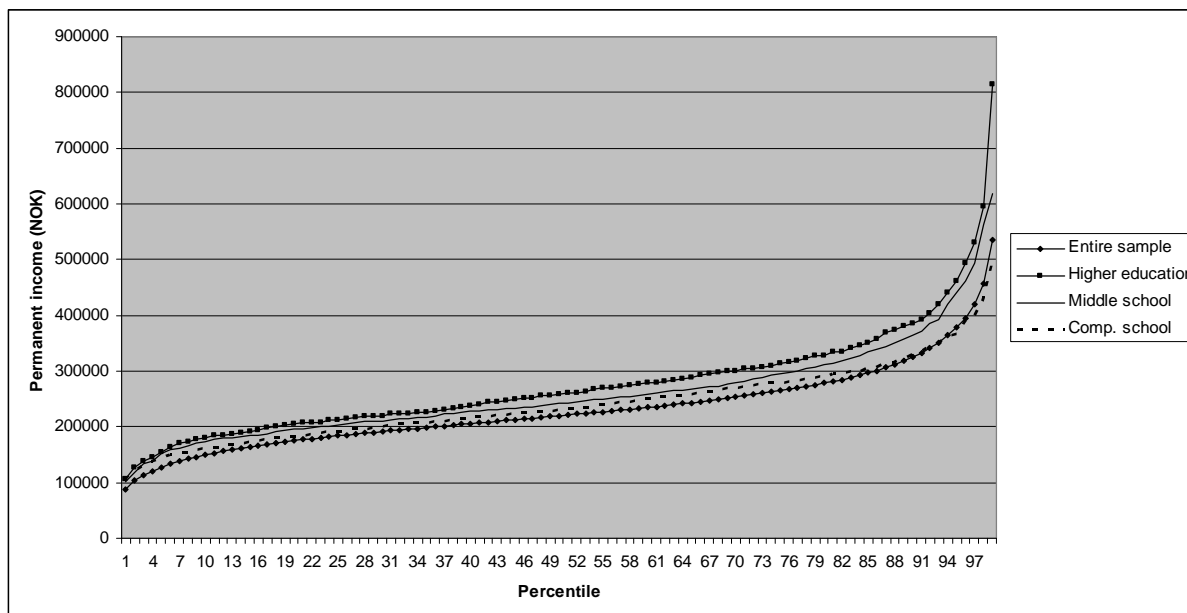
Notes: The distribution denoted entire sample includes the incomes of all individuals. The distributions denoted Higher education, Middle school and Compulsory school are graphed separately by parental education for the sub-sample of individuals from the 1944 cohort who were born in an urban area and living in a small family.

Table 1 also shows that the income of these cohorts increase over time. In particular, the period-specific incomes are, on average, lower than the mean income over the period 1967-2006 up until the early 1990s. It is also clear that the long-term mean income exceeds the permanent income measure, as only the latter takes into the costs of inter-period income transfers. As expected, the income of the

older cohort are relatively high during the first years as they have more labor market experience, but are surpassed by the younger cohorts when they grow older. The mean and the permanent income are rather similar across cohorts. Figure 1 graphs the cumulative distribution of permanent income for the sample as a whole, and for different types according to parental education. As expected, permanent income is increasing in parent's education.

Equality of outcome.

Figure 2 graphs the time-trend in inequality according to the EO principle based on period-specific (two years) income (cf. Period-specific, EO), and compare it to the EO measure of inequality based on permanent income (cf. Permanent, EO.) We see that inequality is fairly stable during the 1970s when the individuals are their mid 30s, with a Gini-coefficient ranging from .18 to .20. After this point in time, inequality is increasing steadily, almost doubling as the individuals reach the peak of the age-earnings profile.[19] In contrast, the Gini coefficient in permanent income is as low as .17, which suggests substantial income mobility. Moreover, our results indicate that snapshots of inequality based on income early in the working lifespan may provide a reasonable approximation of inequality in permanent income when the EO principle is employed. The intuition is that individuals with high permanent income tend to be those with high income growth.

**Figure 2. Gini-inequality according to EO and ex-post EOp based on period-specific income and permanent income**
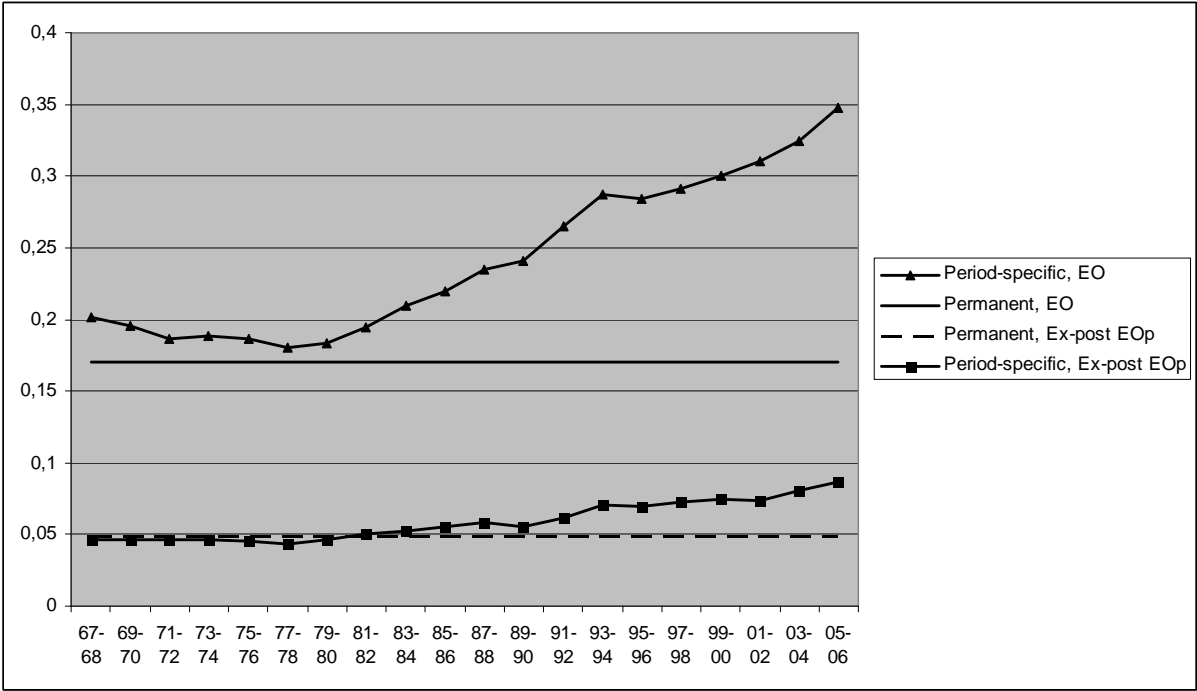
Notes: The ex-post EOp Gini measure of inequality is defined by (3.12).

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[19] A possible explanation is that the age-earnings profile differs systematically across individuals, with relatively high earnings growth among individuals with high education. Another explanation is that skill-biased technical change has led to increased earnings inequality. Investigating these two explanations is beyond the scope of this paper.

Figure 3 reports the time-trend in social welfare according to the EO principle based on period-specific income (cf. Period-specific, EO), and compares it to EO measure of inequality in permanent income (cf. Permanent, EO). We see social welfare is increasing substantially during the 1970s, as mean incomes rise and inequality is stable. Over the next decade, social welfare declines as the welfare loss from increased inequality dominates the increase in mean incomes. During the economic boom in the 1990s, social welfare increased once again owing to a large increase in mean incomes, despite substantially higher income inequality. We also see that social welfare in permanent income mirrors social welfare in period-specific inequality fairly well when the individuals are between 30 and 50 years of age, in contrast to the inequality results.

Figure 3. Social welfare according to EO and ex-post EOp based on period-specific income and permanent income
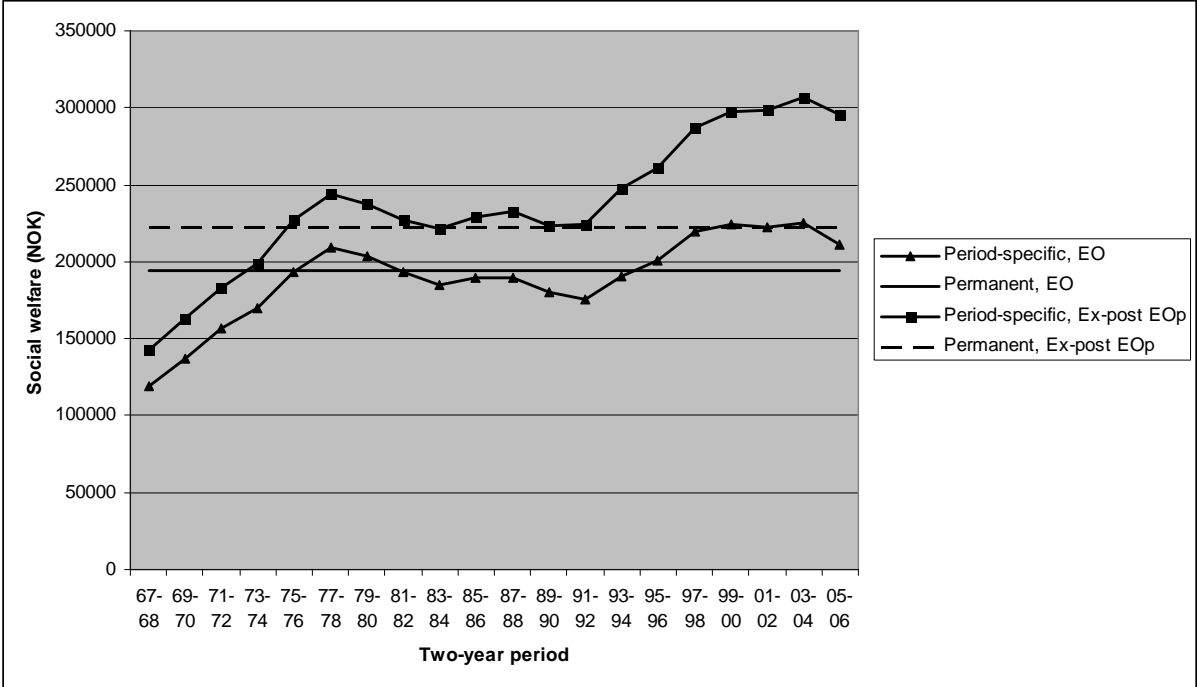

Note: The ex-post EOp social welfare measure is defined by (3.6), with the ex-post EOp Gini measure as the inequality index.

Equality of opportunity.

Turning attention to our ex-post measures of EOp, Figure 2 shows the trend in inequality based on period-specific income (cf. period-specific, ex-post EOp) and compares it to inequality in permanent income (cf. permanent, ex-post EOp). We see that opportunity inequality is fairly stable until individuals reach their early 40s, with ex-post EOp Gini less than .05. After this point in time, the ex-post EOp Gini increases steadily, almost doubling as individuals reach the end of the working lifespan. In contrast, the ex-post EOp Gini in permanent income is slightly lower than .05, which suggests that groups viewed as (dis)advantaged based on snapshots of income regularly change position over time. Figure 4 breaks down the ex-post EOp Gini into the different percentile-specific Gini coefficients. We see that inequality is remarkably stable across the percentiles. In particular, the percentile-specific Gini

coefficients vary between .04 and .06 for quantiles located between the 3rd and 85th percentile, and somewhat higher at the tails of the distribution. Another interesting finding in Figure 2 is that snapshots of inequality based on income early in the working lifespan may provide a reasonable approximation of inequality in permanent income, also when the EOp principle is employed. This implies that outcome inequality increases over the life cycle among individuals who have exerted the same degree of effort.

**Figure 4. Percentile-specific Gini-inequality according to ex-post EOp based on permanent income**
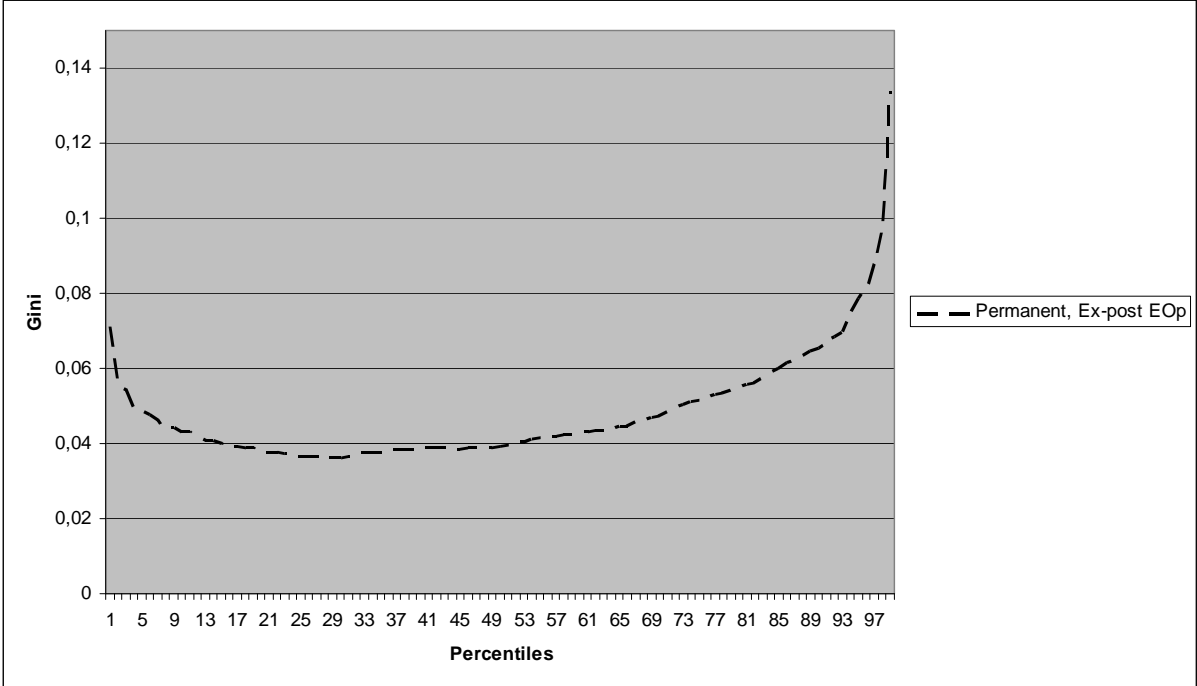
Note: The percentile-specific ex-post EOp measure is given by (3.4), with the ex-post EOp Gini measure as the inequality index.

Figure 2 also demonstrates that the ex-post EOp measures of inequality are much smaller than the corresponding EO measures. As shown in equation (3.15), we can interpret the difference between the EO measure of inequality and the ex-post EOp measure of inequality as inequality due to effort. Specifically, our findings suggest that opportunity inequality makes up about 28 percent of inequality in outcome when the analysis is based on permanent income and between 23 and 26 percent when the period-specific incomes form the basis. However, the time trend in inequality according to the ex-post EOp Gini is quite similar to that produced by the EO Gini. This implies that the life-cycle pattern in outcome inequality among individuals who have exerted the same degree of effort, is much like the life-cycle pattern in outcome inequality in the population as a whole.

Figure 3 reports social welfare measures according to the ex-post EOp principle based on period-specific income (cf. period-specific, EOp) and permanent income (cf. permanent, EOp). Compared to the EO results, social welfare is considerably higher when the ex-post EOp principle forms the basis for the analysis, owing to lower inequality. Specifically, the ex-post EOp measure of social welfare in permanent income is almost 15 percent higher than the EO measure. Yet the time

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trend in social welfare based on period-specific incomes is quite similar across the two measures. In particular, we see that the ex-post EOp measure of social welfare based on snapshots of income during the period the individuals are between 30 and 50 years of age are fairly similar to the corresponding measures based on permanent income.

Figure 5. Gini-inequality according to ex-ante and ex-post EOp based on period-specific income and permanent income
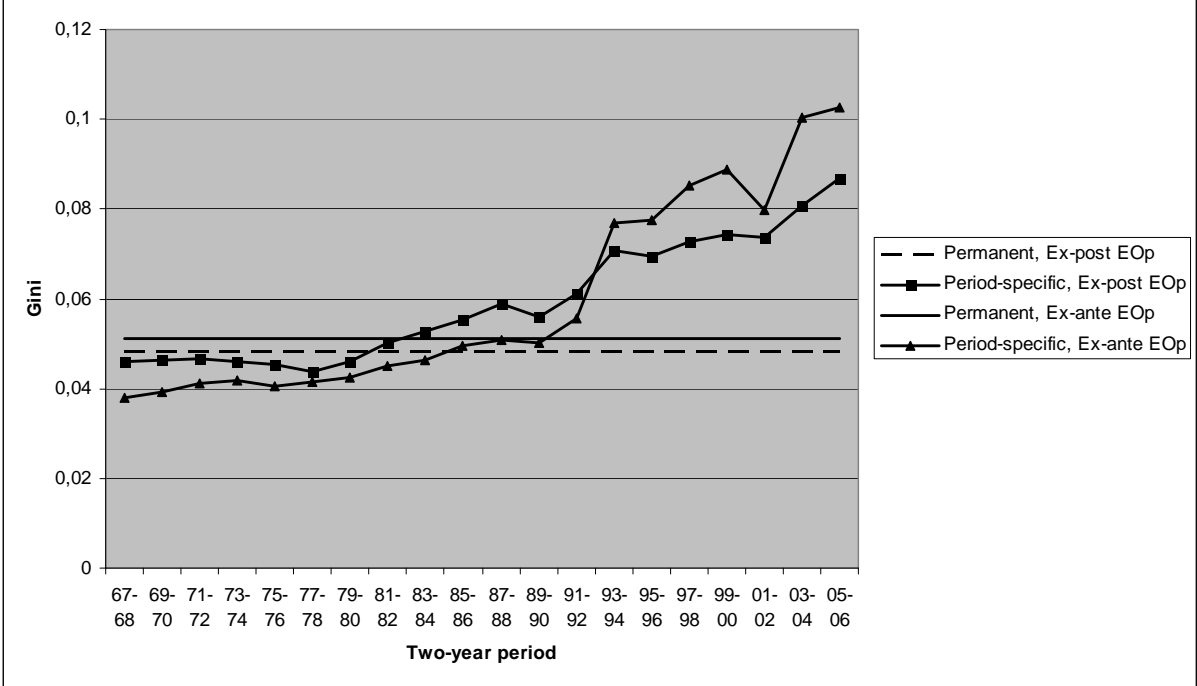
Notes: The ex-post and ex ante EOp Gini are defined by (3.12) and (3.18).

Figure 6. Difference in social welfare according to ex-ante and ex-post EOp based on period-specific income and permanent income
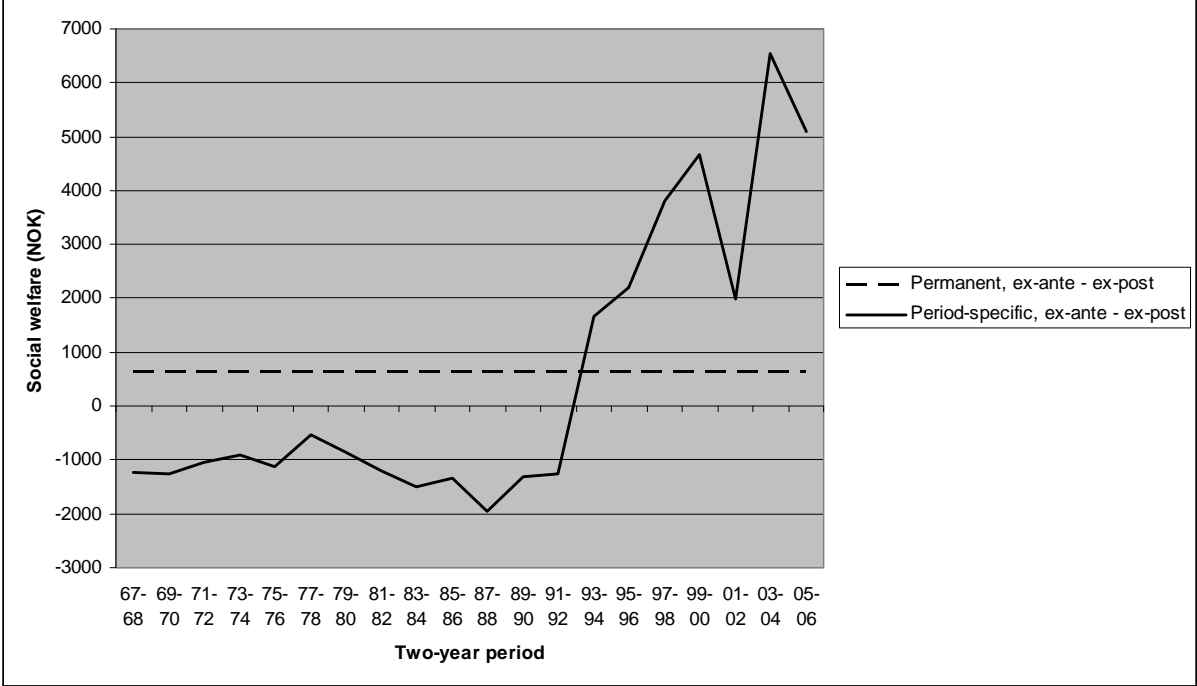
Notes: The ex-post and ex ante EOp social welfare measures are defined by (3.6) and (3.20), using the Gini measure as the inequality index.

Figure 5 compares the ex-ante and the ex-post EOp measures of inequality. The ex ante long term measure is slightly higher than the ex post long term measure. We also see that the period-specific measures show a common trend. In particular, for the first 25 years ex post inequality of opportunity dominates the ex-ante measures, while the situation is reversed for the latest 15 years of the working lifespan. While the ex ante measures are focused on the inequality between social types, the ex post measures are able to capture in a finer way the individual income gaps due to circumstances. Therefore, these results indicate that the distance between social groups, as defined by the circumstances, become increasingly important, relatively to the individual differences, in the last part of the working lifespan. Figure 6 graphs the difference in social welfare according to the ex-ante and ex-post EOp measures. These results mirror well the similarity in inequality across the two types of EOp measures.

Life-cycle variation in the association between current and lifetime income

We wrap up this empirical analysis with a comparison between our results and those produced by the literature on life-cycle variation in the association between current and lifetime income. In the pioneering studies by Haider and Solon (2006) and and Bohlmark and Lindquist (2006), they analyze life-cycle variation in the association between current and permanent or lifetime income, to evaluate the appropriateness of the textbook errors-in-variables model. They find that it does not accurately characterize current income as a proxy for lifetime income. In particular, the slope coefficient in a regression of log current income on log lifetime income is not generally equal to one, as the textbook-errors-in-variablese model assumes, varying instead systematically over the life cycle. An important implication is that using current income as a proxy for lifetime income can generate substantial life-cycle bias in regression analysis.

Along the same lines, we have estimated the slope coefficient in a regression of current income in year t, Y_{it}, on permanent income, Z_i, which is given by:

$$(4.1) \qquad \lambda_t = \frac{cov(Y_{it}, Z_i)}{var(Z_i)} .$$

As in Haider and Solon (2006) and Bohlmark and Lindquist (2006), λ_t starts out at a value less than one at the outset of the career, then increases over the life cycle, and exceeds one later in the career. At the age when λ_t equals 1, one may use current income as a proxy for permanent or lifetime income without generating life cycle bias in regression analysis. In our data set, λ_t is approximately equal to 1 around age 50.[20]

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[20] In comparison, Haider and Solon (2006) find that λ_t comes close to 1 when individuals are in their early 40s (and mid 30s), whereas Bohlmark and Lindquist (2006) report that λ_t is approximately 1 when individuals are aged 46-53 (and around age

However, this finding does not justify using current income at age 50 as a proxy for lifetime income in the measurement of inequality. To see this, note the life-cycle variation in the association between the Gini coefficient in current and permanent income can be expressed as:

$$(4.2) \qquad \delta_t = G(Y_{it}) - G(Z_i) = \frac{2cov[Y_{it}, F(Y_{it})]}{E(Y_{it})} - \frac{2cov[Z_i, F(Z_i)]}{E(Z_i)} \,.$$

In general, $\lambda_t = 1$ does not imply that $\delta_t = 1$ (or vice versa). The reason is that the δ_t depends on the covariance between individuals' incomes and their ranks in the distributions of current and permanent income, whereas λ_t simply depends on the covariance between individuals current and permanent income. As shown in Figure 2, δ_t is minimized when individuals are in their mid 30s, and later (earlier) in individuals' working lifespan, current income inequality overstates (understates) the extent of permanent income inequality. This illustrates that that life-cycle variation in the association between inequality in the distributions of current and lifetime income is a separate issue from life-cycle variation in the association between levels of current and lifetime income.

5 Summary and conclusion

More than half a century ago, Friedman (1957, p38) argued that "*the identification of low measured income with 'poor' and high measured income with 'rich' is justified only if measured income can be regarded as an estimate of expected income over a lifetime or a large fraction thereof"*. His concern was that studies based on fluctuating annual incomes may provide a misleading picture of the consumption possibilities of individuals, and consequently, also the extent of inequality and social welfare in a society. Instead, methods used for distributional analysis should reflect that income can be, and regularly is, used for consumption not only within the year that it is obtained, but also in other years.

Since Friedman's work, a large and growing literature has addressed the question: what exactly should be equalized? Egalitarian theories of justice, especially after Sen's (1980, 1985, 1992) pioneering work, emphasize individuals' responsibility in determining their own well-being. Following this literature, economists have explored different ways of evaluating income distributions according to the EOp principle. However, snapshots of income still form the basis of EOp analyses.

The aim of this paper is to introduce and apply a general framework for evaluating long-term income distributions according to the EOp principle. The proposed framework is valid for measuring short-term or long-term EOp, depending on whether we let snapshots of income or permanent income

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33). It should be noted, however, that these two studies measure income in logs rather than levels. When measuring income in logs, our estimate of λ_t is equal to 1 around age 40.

form the basis. Further, it allows for both an ex-ante and an ex-post approach to EOp. There is long-term ex-post inequality of opportunity if individuals who exert the same effort have different permanent incomes. In comparison, the ex-ante approach focuses on the expected permanent income for groups of individuals with identical circumstances.

To demonstrate the empirical relevance of a long-run perspective on EOp, we have used a unique panel data from Norway on individuals' incomes over their working lifespan. The insights from our empirical results may be summarized with three conclusions. First, EOp measures of inequality are less than one third of the corresponding EO measures, suggesting that a large fraction of inequality of outcomes is attributable to initial circumstances. This is true both when permanent income and snapshots of income form the basis for the analysis. Hence, social welfare is much higher according to the EOp principle than the EO principle. Second, snapshots of income overstate inequality compared to analysis based on permanent income, suggesting substantial income mobility. However, our results indicate that snapshots of inequality based on income early in the working lifespan may provide a reasonable approximation of inequality in permanent income; in comparison, social welfare measures based on snapshots of income during the period the individuals are between 30 and 50 are fairly similar to the corresponding measures based on permanent income. This is true both when the EO and the EOp principle form the basis for the analysis. Third, we find some similarity between ex-ante and ex-post EOp measures of inequality and social welfare. This holds both for permanent income and the time-trend in period-specific income, except for the latest 15 years of the working lifespan.

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