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Growth pro-poorness from an Intertemporal Perspective with an Application to Indonesia, 1997–2007

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Abstract

The impact of growth on the distribution of income or consumption is regularly debated at both the scientific and policy levels. Within the micro-oriented literature dedicated to growth pro-poorness evaluation issues, the focus is specifically on the poverty impacts of growth. Considering a cross-sectional perspective for poverty measurement, early contributions have logically assessed these distributional effects in an anonymous fashion. But this means ignoring both the income dynamics and mobility impacts of growth. The paper extends the growth pro-poorness framework in two important ways. First, a longitudinal perspective is adopted which accounts independently for anonymous and mobility growth effects. Second, the paper's treatment of mobility encompasses both the gain of "mobility as equalizer" and the variability cost of poverty transiency. Several decompositions are introduced to evaluate the relative contribution of each of these effects on the pro-poorness of distributional changes. An empirical illustration is performed using Indonesian data for the period 1997–2007.

Keywords: growth pro-poorness, income mobility, intertemporal poverty, Indonesia.

JEL codes: D31, D63, I32

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1 Introduction

The dynamic relationship between economic growth and distribution changes is a long-lasting subject of investigations from both the micro- and macroeconomic perspectives. In particular, a specific and micro-oriented branch of the literature, known as "pro-poor growth," is generating sustained scrutiny from both the scientific and policy spheres, with the prime objective of assessing how growth is associated with poverty changes. This literature resulted in the development of numerous analytical tools for that purpose (see notably, Ravallion and Chen 2003; Son 2004; Essama-Nssah 2005; Essama-Nssah and Lambert 2009; Duclos 2009; Bérenger and Bresson 2012).

In line with the traditional focus on cross-sectional poverty, a crucial role is played in these tools by the "anonymity" assumption that the identity of the growth beneficiaries shall not be regarded as relevant in the analysis. This is an often uncontroversial hypothesis, in particular, if the aim is to identify the purely cross-sectional impact of growth. However, postulating anonymity means that income dynamics are then disregarded, namely that mobility observed during the growth process is not of measurement and normative interest. To illustrate that point, consider the following two separate income transformations A and B undergone by a four-person distribution of income from period t to $t+1$:

$$(40, 60, 90, 90) \underset{A}{\rightarrow} (90, 90, 40, 60), \tag{1}$$

$$(40, 60, 90, 90) \underset{B}{\rightarrow} (40, 60, 90, 90) \tag{2}$$

Let's assume that, in both periods, the poverty line is equal to 70. In both cases, traditional indexes used to assess the pro-poorness of such growth processes like the Rate of Pro-Poor Growth (RPPG) (Ravallion and Chen 2003) would return zero values as the final marginal distribution of income is strictly identical to the initial marginal distribution.[1] Yet, the two income dynamics are quite different: considerable mobility is implied by A whereas B leaves everyone's income unchanged. We may therefore wish a pro-poorness index to behave differently when considering the two growth patterns.

To circumvent these limitations, it is argued that a "non-anonymous" perspective should be endorsed for growth pro-poorness assessments (see notably Grimm 2007; Jenkins and Van Kerm 2011; Bourguignon 2011; Palmisano and Peragine 2015; Palmisano and Van de gaer 2016). Proponents of this position emphasize the crucial role of mobility in the distributional effects associated with growth. While measurement aspects of growth pro-poorness and of mobility are both quite developed, the analysis of the impact of mobility on growth pro-poorness is a promising field that has yet to be developed to our knowledge.[2]

[1] See also Kakwani and Pernia (2000), Kakwani and Son (2003) and Kakwani and Son (2008) for alternative grow pro-poorness indexes.

[2] See for instance the reviews on mobility measurement in Fields and Ok (1999); Fields (2008), or

Bringing together these two issues means considering the individual poverty trajectories over time, hence considering an intertemporal evaluation of poverty. Mobility will then have converse effects on intertemporal poverty. On the one hand, consistent with Friedman (1962), mobility generally implies some equalization of permanent incomes across individuals. On the other hand, mobility induces variability costs, since risk-averse individuals may experience welfare losses with time variability. In the present study, the pro-poor or anti-poor nature of growth is determined by comparing observed intertemporal poverty with a counterfactual situation consisting of the absence of any kind of distributional change.

Various pro-poorness features of growth are also explored in this paper through a set of additive decompositions. The first one disentangles the measurement of anonymous growth from that of its non-anonymous component. The second decomposition isolates the snapshot effects of income changes from multitemporal ones. The third decomposition separates the contribution of reranking, inequality changes, and pure growth in explaining growth pro-poorness. Finally, a fourth decomposition makes it possible to estimate the contribution of each subperiod to intertemporal poverty changes.

The approach suggested in the present paper differs both methodologically and conceptually from past contributions on this topic. For instance, the Individual RPPG introduced by Grimm (2007), defined as the average income growth of the initially poor individuals, specifically focuses on the impact of growth on the initially poor and does not take into account the negative income effects of those who experience deprivation after growth. Foster and Rothbaum (2012) propose using cutoff-based mobility measures to identify variations of poverty over time, but, their method restricts poverty measurement to two specific snapshot poverty indexes, namely the headcount index and the mean poverty gap whose limitations are widely aknowledged (Sen 1976).

This paper's contribution to the literature is twofold. The first contribution is to account for the impact of a growth process on intertemporal poverty, hence making it possible to disentangle the anonymous impact of growth from its mobility impact (the non-anonymous growth). The second contribution is an extension of the "mobility as equalizer" framework to take into account the effect of horizontal mobility on poverty, corrected for poverty transiency costs as well as for social welfare losses due to inequality in the distribution of intertemporal poverty among the population.

The rest of the paper is organized as follows. Section 2 introduces a family of intertemporal indexes that can be interpreted as a representative income shortfall, that is the welfare loss, expressed as a share of the poverty line, due to the existence of poverty over the whole period. Section 3 describes our conceptual framework for assessing intertemporal growth pro-poorness and describes its properties when used with the suggested intertemporal poverty indexes. Section 4 suggests various decomposi-

Jäntti and Jenkins (2015).

tions of the proposed indexes that help to understand the pro-poor or anti-poor nature of observed growth processes. An empirical illustration of this framework is contained in Section 5 considering Indonesia during the period 1997–2007. It is notably shown that, unless variability aversion is large relatively to inter-individual inequality aversion, growth can be deemed intertemporally pro-poor in Indonesia during this period. Section 6 concludes.

2 Intertemporal poverty assessment

The analysis is focuses on the dynamics of a distribution of living standards (incomes, without loss of generality) for a population of n persons, with individuals denoted $i = 1, ..., n$ over $T > 1$ time periods (annual or monthly for instance) of their life. Each generic period is denoted by $t = 1, ..., T$ and the duration T is supposed to be the same for the whole population, *viz*, we are comparing people's living conditions over the same spell.

Periodic income $y_{i,t}$ is supposed to be non-negative. Let $\boldsymbol{y}_{(i)} \equiv (y_{i,1}, \ldots, y_{i,t}, ..., y_{i,T})$ then be the vector of individual i's incomes across the T periods and $\boldsymbol{y}_t$ be a cross-sectional vector of incomes at time t. The income profile $\boldsymbol{y}_i$ is the ith row of the $n \times T$ matrix $\boldsymbol{Y}$. For the sake of simplicity, we normalize incomes at time t by the corresponding poverty line $z_t > 0$. Poverty lines can either be absolute (constant in real terms) or relative (to income norms that are likely to vary across time). Censoring incomes at the corresponding poverty line yields $\tilde{y}_{i,t} \equiv \min(y_{i,t}, 1)$. Then poverty can be measured over an individual's lifetime by $p\big(\boldsymbol{y}_{(i)}\big)$ with $p\big(\boldsymbol{y}_{(i)}\big) \geq 0$ whenever $\exists t \in \{1, \ldots, T\}$ such that $y_{i,t} < 1$ and $p\big(\boldsymbol{y}_{(i)}\big) = 0$ otherwise. Intertemporal poverty at the population level is measured by the index $P(\boldsymbol{Y})$.

2.1 Individual illfare

Let the (normalized) poverty gap for person i at period t be defined by $g_{i,t} \equiv 1 - \tilde{y}_{i,t}$. Then vector $\boldsymbol{g}_{(i)} \equiv (g_{i,1}, \ldots, g_{i,t}, ..., g_{i,T})$ describes the sequence of poverty gaps for this person i across T periods, and $\boldsymbol{G}$ is the $n \times T$ matrix of normalized poverty gaps for the whole population. Finally, the vector $\boldsymbol{g}_t \equiv (g_{1,t}, \ldots, g_{n,t})$ gives the cross-sectional distribution of gaps at time t. In the literature, the income gap $g_{i,t} \in [0, 1]$ is a standard measure of individual poverty for both snapshot and intertemporal poverty measurement. For instance, the widely used FGT class (Foster, Greer, and Thorbecke 1984) of additive poverty indexes relies on the aggregation of simple transformations of poverty gaps.[3] Using an FGT-like formulation, the poverty of each individual i over

[3]It also serves as a basis for the intertemporal generalizations of FGT indexes proposed in Foster (2009); Canto, Gradín, and del Rio (2012), or Busetta and Mendola (2012), not to mention specific members of the family of indexes introduced by Hoy and Zheng (2011); Bossert, Chakravarty, and d'Ambrosio (2012), and Dutta, Roope, and Zank (2013)

the T periods can be measured by:

$$p_\gamma\left(\boldsymbol{y}_{(i)}\right) \equiv \sum_{t=1}^{T} \omega_t g_{i,t}^{\gamma}, \quad \text{with } \gamma \geq 0, \tag{3}$$

where the $\omega_t > 0$, $t \in \{1, \ldots T\}$ and $\sum_{t=1}^{T} \omega_t = 1$, define a weighing scheme that indicates the sensitivity of poverty to the sequence of experienced deprivations. With decreasing weights, priority is given to eradicating poverty experienced earlier in life, for instance in childhood; with weights increasing through time, more importance is on the contrary given later deprivations.[4]

The parameter γ measures the social evaluator aversion to inequality and variability in a person's poverty gaps. A larger value for γ means higher weight is given to income losses for severe deprivations when compared with light deprivations. For $\gamma = 1$, the index (3) is the simple weighted average of i's poverty gaps across time. For $\gamma > 1$, a sequence of income increments and decrements that leaves the weighted mean of income gaps unchanged but shrinks intertemporal variability reduces $p_\gamma\left(\boldsymbol{y}_{(i)}\right)$. It is worth stressing that the index relies on a "union" definition of the poverty domain since individuals are regarded as poor, from an intertemporal perspective, whenever they experience at least one deprivation during the whole period.[5]

So as to account explicitly for the cost of time variability, we suggest using the poverty counterpart of the "equally distributed equivalent income" introduced by Atkinson (1970) for the assessment of inequality and social welfare. This equally distributed equivalent (EDE) poverty gap for person i, $\pi_\gamma\left(\boldsymbol{g}_{(i)}\right)$, is defined by:

$$\pi_\gamma\left(\boldsymbol{g}_{(i)}\right) \equiv p_\gamma^{-1}\left(p_\gamma\left(\boldsymbol{y}_{(i)}\right)\right) = \left(\sum_{t=1}^{T} \omega_t g_{i,t}^{\gamma}\right)^{\frac{1}{\gamma}}. \tag{4}$$

The EDE gap $\pi_\gamma\left(\boldsymbol{g}_{(i)}\right)$ is the gap level that, if experienced at each period of i's lifetime, would result in the same level of poverty for i over time as that generated by its observed sequence of relative deprivations. For $\gamma = 1$, $\pi_\gamma\left(\boldsymbol{g}_{(i)}\right)$ then corresponds to the simple weighted average gap over time, i.e. $\pi_1\left(\boldsymbol{g}_{(i)}\right) = \sum_{t=1}^{T} \omega_t g_{i,t}$. For $\gamma > 1$, $\pi_\gamma\left(\boldsymbol{g}_{(i)}\right)$ is never lower than $\pi_1\left(\boldsymbol{g}_{(i)}\right)$ because variability is regarded as a social bad. The difference between these two values can be interpreted as the cost of individual i's deprivation variability:

$$c_\gamma(\boldsymbol{g}_{(i)}) \equiv \pi_\gamma\left(\boldsymbol{g}_{(i)}\right) - \pi_1\left(\boldsymbol{g}_{(i)}\right). \tag{5}$$

[4] The index (3) is a specific version of the lifetime individual poverty measure introduced by Hoy and Zheng (2011). See also Bresson and Duclos (2015).

[5] A generalization with other definitions of the poverty domain using a counting approach *à la* Alkire and Foster (2011) can easily be performed by censoring vectors $\boldsymbol{g}_{(i)}$ whose (weighted) number of deprivations is less than a given threshold $\in]1, T]$.

Hence, intertemporal poverty for i can be expressed as:

$$\pi_\gamma\left(\boldsymbol{g}_{(i)}\right) = \pi_1\left(\boldsymbol{g}_{(i)}\right) + c_\gamma(\boldsymbol{g}_{(i)}). \tag{6}$$

Consequently, $\pi_\gamma\left(\boldsymbol{g}_{(i)}\right)$ is the sum of the (weighted) average intertemporal income gap and of the intertemporal cost of mobility.

2.2 Social illfare

Here, we consider the aggregation of these individual EDE gaps so as to obtain a comparable value for the whole population. As in the case of traditional snapshot poverty, many functional forms can be proposed to perform this social aggregation. Here, we also make use of the FGT formulation for aggregation:[6]

$$P_{\alpha,\gamma}\left(\boldsymbol{Y}\right) \equiv \frac{1}{n}\sum_{i=1}^{n}\left(\pi_\gamma\left(\boldsymbol{g}_{(i)}\right)\right)^{\alpha}, \tag{7}$$

where parameter $\alpha \geq 1$ measures aversion to poverty inequality across individuals. A socially representative EDE gap for the population, $\Pi_{\alpha,\gamma}\left(\boldsymbol{G}\right)$, is then given by:

$$\Pi_{\alpha,\gamma}\left(\boldsymbol{G}\right) \equiv \left(\frac{1}{n}\sum_{i=1}^{n}\left(\pi_\gamma\left(\boldsymbol{g}_{(i)}\right)\right)^{\alpha}\right)^{\frac{1}{\alpha}}. \tag{8}$$

In general, individual dynamics are taken into account with this intertemporal index, but an anonymous evaluation of intertemporal poverty can be performed using $\Pi_\alpha \equiv \Pi_{\alpha,\alpha}$. Switching two poor persons' income at any t will then not impact the social evaluation of intertemporal poverty, whatever the income streams of the two individuals in the other periods.[7]

Indices $P_{\alpha,\gamma}$ and $\Pi_{\alpha,\gamma}$ are ordinally equivalent and so can be used equally for comparing any pair of distributions. However, $\Pi_{\alpha,\gamma}\left(\boldsymbol{G}\right)$ can be usefully interpreted as the relative gap level which, if assigned uniformly to all individuals at every time period, would yield the same poverty level as that observed with the intertemporal distribution $\boldsymbol{G}$. It is thus a representative gap that indicates the social cost, expressed as a fraction of the poverty line, of observed poverty.

[6] The resulting index P_α^θ is the one proposed by Bourguignon and Chakravarty (2003) in the context of multidimensional poverty measurement. It also generalizes Duclos, Araar, and Giles (2010), where $\alpha = \gamma$ and $\omega_t = \frac{1}{T}\ \forall t \in \{1,\dots,T\}$.

[7] This can be more easily seen if we express $\Pi_\alpha\left(\boldsymbol{G}\right)$ as:

$$\Pi_\alpha\left(\boldsymbol{G}\right) = \left(\sum_{t=1}^{T}\omega_t\frac{1}{n}\sum_{i=1}^{n}g_{i,t}^{\alpha}\right)^{\frac{1}{\alpha}} = \left(\sum_{t=1}^{T}\omega_t P_\alpha(\boldsymbol{g}_t)\right)^{\frac{1}{\alpha}}. \tag{9}$$

The poverty ranking of two distributions showing the same marginal income distributions but different joint distributions will depend on the preferences of the social evaluator with respect to poverty variability and poverty inequality. Note that, in that case, the cross-sectional distributions of poverty gaps are the same under the two processes. With aversion toward inequality and variability is the same (i.e. $\alpha = \gamma$), the two distributions will then be judged tantamount in terms of poverty. Let $\tilde{\boldsymbol{G}}$ be a permutation of $\boldsymbol{G}$ so that individual ranks are kept unchanged during the whole growth process. Distribution $\tilde{\boldsymbol{G}}$ is regarded as no worse than $\boldsymbol{G}$ with indifference toward variability ($\gamma = 1$), while insensitivity toward inequality ($\alpha = 1$) makes distribution $\boldsymbol{G}$ no worse than $\tilde{\boldsymbol{G}}$. Hence, whether poverty is more severe in $\boldsymbol{G}$ or $\tilde{\boldsymbol{G}}$ will crucially depend on the chosen values for α and γ.

As with individual illfare, useful decompositions can be performed for the poverty index Π_α. Let

$$c_{\alpha,\gamma}(\boldsymbol{G}) \equiv \Pi_{\alpha,\gamma}(\boldsymbol{G}) - \Pi_{1,\gamma}(\boldsymbol{G}) \tag{10}$$

be the cost of inequality of intertemporal poverty across individuals. It shall not be confused with:

$$\frac{1}{n}\sum_{i=1}^{n} c_\gamma(\boldsymbol{g}_{(i)}) = \Pi_{1,\gamma}(\boldsymbol{G}) - \Pi_{1,1}(\boldsymbol{G}), \tag{11}$$

that is the average cost of deprivation variability at the aggregate level. Associating (11) with (10) and solving for $\Pi_{\alpha,\gamma}(\boldsymbol{G})$ we obtain:

$$\Pi_{\alpha,\gamma}(\boldsymbol{G}) = \Pi_{1,1}(\boldsymbol{G}) + \frac{1}{n}\sum_{i=1}^{n} c_\gamma(\boldsymbol{g}_{(i)}) + c_{\alpha,\gamma}(\boldsymbol{G}). \tag{12}$$

Equation (12) additively decomposes aggregate intertemporal poverty into three components: the average individual intertemporal poverty gap, the average cost of deprivation variability, and the cost of inequality in intertemporal poverty.

3 Measurement of pro-poorness in an intertemporal setting

3.1 General framework

Usually, that is in the context of cross-sectional analyses of poverty, assessing the pro-poor nature of a given growth process implies comparing the observed poverty level at the end of the period with the level that would have been observed under some given benchmark. This benchmark could be either a targeted poverty level or a counterfactual one. Let $\hat{\boldsymbol{Y}}$ denote that reference distribution.

The suggested measurement of pro-poor growth is anchored to an intertemporal pro-poorness evaluation function $IPP\big(P(\hat{\boldsymbol{Y}}), P(\boldsymbol{Y})\big)$ that takes the simple linear form in the present paper for expositional simplicity:

$$IPP(P(\hat{\boldsymbol{Y}}), P(\boldsymbol{Y})) \equiv P(\hat{\boldsymbol{Y}}) - P(\boldsymbol{Y}), \tag{13}$$

that satisfies standard appealing properties. For instance, $IPP(P(\hat{\boldsymbol{Y}}), P(\boldsymbol{Y})) = 0$ if observed poverty is identical to benchmark poverty. Moreover, the measure will be deemed pro-poor (anti-poor) if estimated intertemporal poverty is lower (larger) than the chosen counterfactual poverty level. Finally, values of the index can be compared, a larger (lower) value for one given growth spell being qualified as more pro-poor (anti-poor).[8]

The definition of the counterfactual situation is crucial as different benchmark distributions will naturally result in different evaluations of growth pro-poorness. A crucial element is whether an absolute or a relative definition of growth pro-poorness is chosen—the former view considers that growth is pro-poor when poverty decreases absolutely speaking while the latter states that growth is pro-poor when the incomes of the poor rise faster than some norm (often proportional to mean income). For the sake of simplicity, this paper follows an absolute approach. However, it is worth pointing out that generalizing to a relative approach simply means dividing incomes by the chosen norm.

Similarly, "mobility means different things to different people," in the words of Fields (2008, p. 1), and some agreement is necessary with respect to that concept. In this paper, mobility is interpreted as any temporal change in individual income. A natural candidate for the counterfactual scenario is then the status quo, namely the absence of distributional changes. The benchmark $\boldsymbol{Y}_1$ is then a counterfactual distribution in which every person would receive exactly the same income as the one he got initially.[9] The IPP index is consequently the difference between poverty in a counterfactual situation in which the first period deprivation is extended over the T-period growth spell and observed intertemporal poverty.[10]

Of course, as known in the growth pro-poorness literature (Duclos 2009), rival versions can be proposed for the counterfactual distribution. For instance, the counterfactual distribution could only refer to the absence of exchange mobility, hence resulting in a counterfactual distribution showing the same marginal distributions as the observed distribution but without reranking from year to year. Another possibility is to take a relative view on, that is to consider a "neutral" growth process (in terms of

[8] Fields (2010) uses similar properties for the measurement of mobility.

[9] A similar approach is used by Chakravarty, Dutta, and Weymark (1985) and Fields (2010), although the benchmark in the former study is based on relative immobility, that is the *share* of each person in total income is assumed to be constant across time.

[10] This property relates to the normalization axiom proposed by Hoy and Zheng (2011) that requires a person's lifetime poverty to be represented by snapshot poverty if this person gets every period the same income level.

snapshot inequality) over the studied period.[11] However it is worth stressing that some of the decompositions proposed in section 4 make it possible to obtain quite easily the corresponding values of the IPP as components or sum of components of our preferred version of the IPP.

3.2 Intertemporal pro-poorness indexes

Using the benchmark deprivation matrix $\boldsymbol{G}_1$ referring to $\boldsymbol{Y}_1$, we have $\Pi_{\alpha,\gamma}(\boldsymbol{G}_1) = \Pi_\alpha(\boldsymbol{g}_1)$, that is:

$$\Pi_\alpha(\boldsymbol{g}_1) = \left(\frac{1}{n}\sum_{i=1}^{n} g_{i,1}^{\alpha}\right)^{\frac{1}{\alpha}}, \tag{14}$$

This is then the EDE income gap corresponding to the value of the FGT index at year $t = 1$. Using the family of poverty indexes introduced in the previous section, we obtain an operational expression for (13):

$$IPP_{\alpha,\gamma} = \Pi_\alpha(\boldsymbol{g}_1) - \Pi_{\alpha,\gamma}(\boldsymbol{G}). \tag{15}$$

The index equals 0 when everyone's deprivation level is left unchanged during the whole growth spell. It takes a positive value if intertemporal poverty is less severe than initial cross-sectional poverty, and negative in the opposite case. If growth is associated with the eradication of poverty at the subsequent periods, then $IPP_{\alpha,\gamma}$ will be equal to $(1-\omega_1)\Pi_\alpha(\boldsymbol{g}_1) > 0$. This is an upper bound for the IPP index and is equal to the amount of intertemporal poverty that is eliminated through growth, which corresponds to discounted value of poverty experienced in the first period.

The cost of individual variability as well as the benefits of a potential reduction of intertemporal inequalities, both resulting from mobility, are incorporated in the $IPP_{\alpha,\gamma}$ index.[12] $IPP_{\alpha,\gamma}$ satisfies the usual properties of anonymity (in the identity of individual gap vectors), scale invariance, continuity, population invariance, and subgroup consistency required for social evaluations. $IPP_{\alpha,\gamma}$ increases with initial poverty and decreases with intertemporal poverty. Nevertheless, changes in first-period

[11] As stressed by an anonymous referee, a possible issue is that year $t = 1$ was an abnormal year during the period of interest, hence resulting in large values of the IPP, in particular if T is relatively large. We acknowledge this possible issue but note that the same problem is likely to hold with usual growth "pro-poorness" tools. A possible solution to fix that issue could be to test the sensitivity of the results by considering a contiguous year as the reference or averaging individual incomes for the very first years of the growth spell. However, one can simply argue that no interpretation of the IPP should be given without any ex ante description of the studied growth spell.

[12] It is worth underlining that the family of indexes proposed in eq. (18) are normative in nature. Such normatively grounded indexes are derived from explicit social illfare functions and are measures of the change in intertemporal social illfare resulting from mobility. Such measures contrast with indexes of mobility that aim at describing some aspects of mobility. Hence our framework is not meant to provide statistical measures of income changes but to assess the impact of such changes on intertemporal illfare. By using a welfare function to perform this comparison, our pro-poorness indexes allow us to determine whether the observed changes were desirable in terms of poverty or social illfare reduction.

gaps have ambiguous effects since both poverty levels are affected.

To illustrate the behavior of the index, consider the example (1) used in the introduction. As the average income gap is left unchanged during this growth process, the sign of $IPP_{\alpha,\gamma}$ will uniquely depend on the chosen values for the aversion to poverty variability and aversion to intertemporal poverty parameters—assigned values for the weighing scheme do not determine the sign of the index here. In particular, for $\gamma > \alpha$ variability aversion dominates aversion to poverty inequality and the other way around for $\gamma < \alpha$. Let us consider the case of $\omega_1 = \omega_2$. With more emphasis given to variability aversion, for instance $\alpha = 3$ and $\gamma = 4$, the index becomes negative (e.g. $IPP_{3,4} = -0.016$). Because of the cost of temporal variability the growth process is not regarded as pro-poor. With $\alpha = 3$ and $\gamma = 2$ the index takes a positive value (e.g. $IPP_{3,2} = 0.029$) and the transformation can be deemed pro-poor because of the poverty equalization effect of mobility.

4 Decompositions

In this section, we suggest four decompositions of the $IPP_{\alpha,\gamma}$ index that show the respective contributions of mean income growth, mobility, inequality, and subperiod changes. For the sake of simplicity, we set $T = 2$ for the first three decompositions.[13]

The first decomposition disentangles the anonymous and the mobility components of growth:

$$IPP_{\alpha,\gamma} = \underbrace{\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_1) - \Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_2)}_{AG} + \underbrace{\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_2) - \Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_2)}_{M}. \tag{16}$$

The index $\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_2)$ returns an anonymous evaluation of intertemporal poverty. Consequently, it does not account for the social evaluation of the benefits and costs of mobility: AG accordingly assesses the poverty effect of an anonymous growth process, while M captures the non-anonymous effects of observed mobility during the growth spell. The component AG is positive if we observe both a decrease in the mean poverty gap and a contraction in the periodic distribution of poverty gaps. The component M is positive if inter-individual inequality aversion is stronger than temporal variability aversion ($\alpha > \gamma$), zero for $\gamma = \alpha$, and otherwise negative. The sign of the two effects is not determined by the weights ω_t. With example (1), we obtain $AG = 0$ and $M = 0.029$ with $\alpha = 3$ and $\gamma = 2$. As the anonymous growth impact is nil, the beneficial impact of the whole growth process on intertemporal poverty can exclusively be attributed to a (pro-poor) effect of observed mobility.

The distinction between standard anonymous pro-poorness and our intertemporal approach is further highlighted with the second decomposition. For that purpose, it is

[13] A generalization to larger values of T is provided in the appendix.

worth noting that the poverty cost of inter-person inequality in $\Pi_\alpha(\boldsymbol{g}_1)$ is the poverty cost of initial inequality, that is, $c_\alpha(\boldsymbol{g}_1)$. Using (10), a decomposition of the benchmark poverty level is:

$$\Pi_\alpha(\boldsymbol{g}_1) = \Pi_1(\boldsymbol{g}_1) + c_\alpha(\boldsymbol{g}_1), \tag{17}$$

that is the sum of the average poverty gap in the first period and the cost of inequality in the initial distribution of individual poverty gaps. In a two-period setting, equation (12) can then be rewritten as:

$$\Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_2) = \omega_1 P_1(\boldsymbol{g}_1) + \omega_2 P_1(\boldsymbol{g}_2) + \frac{1}{n}\sum_{i=1}^{n} c_\gamma(\boldsymbol{g}_{(i)}) + c_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_2). \tag{18}$$

The following decomposition of the IPP index can then be proposed:

$$\begin{aligned} IPP_{\alpha,\gamma} = & \underbrace{\omega_2 \left[P_1(\boldsymbol{g}_1) - P_1(\boldsymbol{g}_2)\right]}_{\Delta P^c} + \underbrace{\omega_2 \left[c_\alpha(\boldsymbol{g}_1) - c_\alpha(\boldsymbol{g}_2)\right]}_{\Delta c^c} \\ & + \underbrace{\left[\omega_1 c_\alpha(\boldsymbol{g}_1) + \omega_2 c_\alpha(\boldsymbol{g}_2)\right] - c_{\alpha,\gamma}(\boldsymbol{g})}_{M^c} \underbrace{- \frac{1}{n}\sum_{i=1}^{n} c_\gamma(\boldsymbol{g}_{(i)})}_{CV}. \end{aligned} \tag{19}$$

The interpretation for those four components is the following:

- ΔP^c captures changes in the average cross-sectional gaps, $P_1(\boldsymbol{g}_1)$ and $P_1(\boldsymbol{g}_2)$, and so does not depend on variability and intertemporal inequalities.

- Δc^c is, up to a multiplicative term, the difference between the cost of inequality in the initial and in the final periods. Δc^c can be both positive or negative, depending on whether inequality in cross-sectional poverty has fallen or has increased between the two periods.

- M^c, is the difference between the weighted sum of the cost of cross-sectional inequalities and the cost of intertemporal inequality, which is mobility's ability to decrease inequality between individuals, taking the cost of variability into account.

- CV reflects the cost of the longitudinal variability induced by mobility. CV is always negative when $\gamma > 1$ since variability aversion then systematically assigns a social cost to the variability associated with mobility.

Disregarding the weighing term ω_2, the first two components ΔP^c and Δc^c capture the usual components of anonymous pro-poor growth in the spirit of Ravallion and Chen (2003).[14] Conversely, the two components M^c and CV reflect the social evaluator's trade-off between the costs and benefits of mobility, that is the intertemporal pro-poorness effects. It can be noted that $\Delta c^c = 0$ and $M^c = 0$ with $\alpha = 1$, while $CV = 0$ when $\gamma = 1$. In the specific case of $\alpha = \gamma = 1$, $\Delta c^c = M^c_{\alpha,\gamma} = CV = 0$, and consequently $IPP_{\alpha,\gamma} = \Delta P^c$, the difference in the average poverty gap.

Turning back again to example (1), the first two components, ΔP^c and Δc^c, are nil as the (anonymous) cross-sectional income distribution is the same in both periods. For $\alpha = 3, \gamma = 2$, $M^c = 0.089$ shows a positive value indicating that growth has shrunk deprivation inequalities from an intertemporal point of view. Income inequalities are the same in both periods, but considering a larger two period time-horizon, they have decreased in comparison with the benchmark case. Lastly, $CV = -0.059$.

With the third decomposition, the emphasis is put on the reranking effect of growth. It is obtained by making use of two counterfactual distributions $\boldsymbol{g}_1^I$ and $\boldsymbol{g}_1^{IR}$. The counterfactual distribution $\boldsymbol{g}_1^I$ is obtained starting from the distribution of individuals' poverty gaps at the final period but scaling them to obtain the average poverty gap of the first period and ordering them on the base of their rank in the first period, that is $\boldsymbol{g}_1^I \equiv \tilde{\boldsymbol{g}}_2 \frac{\Pi_1(\mathbf{g}_1)}{\Pi_1(\mathbf{g}_2)}$ with $\tilde{\boldsymbol{g}}_2 \equiv r(\boldsymbol{g}_2, \boldsymbol{g}_1)$ where $r(\boldsymbol{a}, \boldsymbol{b})$ orders elements from $\boldsymbol{a}$ according to observed ranks in $\boldsymbol{b}$.[15] It is clear that the only feature that differs between $\boldsymbol{g}_1$ and $\boldsymbol{g}_1^I$ is inequality. The counterfactual distribution $\boldsymbol{g}_1^{IR}$ is obtained starting from the previous counterfactual distribution $\boldsymbol{g}_1^I$, but ordering individuals on the base of their rank at the end of the growth spell, that is $\boldsymbol{g}_1^{IR} \equiv \boldsymbol{g}_2 \frac{\Pi_1(\boldsymbol{g}_1)}{\Pi_1(\boldsymbol{g}_2)}$.[16] So, the unique difference between $\boldsymbol{g}_1^I$ and $\boldsymbol{g}_1^{IR}$ is reranking. As a consequence, $\boldsymbol{g}_1^{IR}$ and $\boldsymbol{g}_2$ only differ with respect to their average poverty gap. Note that the counterfactual distributions $\mathbf{g}_1^{IR}$ and $\mathbf{g}_1^I$ are computed by considering the inequality structure and the ranks of the poverty gaps distribution and not of the income distribution. Although this choice may seem debatable, it is in line with considering an index showing sensitivity to deprivations variability across time (through γ) and to inequalities of intertemporal poverty across persons (through α).

[14] It deserves to be noted that AG and the sum $\Delta P^c + \Delta c^c$ generally differ, since we have:

$$AG = \left(P_\alpha(\boldsymbol{g_1})\right)^{\frac{1}{\alpha}} - \left(\omega_1 P_\alpha(\boldsymbol{g_1}) + \omega_2 P_\alpha(\boldsymbol{g_2})\right)^{\frac{1}{\alpha}}, \tag{20}$$

$$\Delta P^c + \Delta c^c = \omega_2 \left(\left(P_\alpha(\boldsymbol{g_1})\right)^{\frac{1}{\alpha}} - \left(P_\alpha(\boldsymbol{g_2})\right)^{\frac{1}{\alpha}} \right). \tag{21}$$

Note that $AG = \Delta P^c + \Delta c^c$ when $\alpha = 1$; when $\alpha > 1$, we have instead $AG \leq \Delta P^c + \Delta c^c$.

[15] Consider a situation in which the distribution of income is the initial distribution from example (1) and $(10, 6, 5, 8)$ at the $T = 2$. Given the poverty line $z = 7$, $\boldsymbol{g_2}$ is then $(0.29, 0.14, 0, 0)$. Since $\tilde{\boldsymbol{g}}_2 = (0, 0.14, 0.29, 0)$, $\Pi_1(\boldsymbol{g}_1) = 0.143$ and $\Pi_1(\boldsymbol{g}_2) = 0.108$, we consequently have $\boldsymbol{g}_1^I = (0, 0.14, 0.29, 0) \times \frac{0.143}{0.108}$.

[16] Considering the example proposed in footnote 15, $\boldsymbol{g}_1^{IR}$ will then be $(0.29, 0.14, 0, 0) \times \frac{0.143}{0.108}$.

Noting that $\Pi_\alpha(\boldsymbol{g}_1) = \Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_1)$, the third decomposition is then:[17]

$$IPP_{\alpha,\gamma} = \underbrace{\Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_1) - \Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_1^I)}_{I} + \underbrace{\Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_1^I) - \Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_1^{IR})}_{R}$$
$$+ \underbrace{\Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_1^{IR}) - \Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_2)}_{PG}. \quad (22)$$

The interpretation of each component is the following:

- I captures the intertemporal effects of inequality and variability in poverty ($\boldsymbol{g}_1$ and $\boldsymbol{g}_1^I$ share the same arithmetic mean and they rank individuals in the same manner). More specifically, I assesses the effects of inequality across time and individuals when initial ranks are preserved. Increasing inequalities will systematically result in a negative value for I, no matter the chosen values for the aversion parameters α and γ. With $\alpha = \gamma = 1$, I will be null as the index becomes neutral with respect to intertemporal variability and inequality in poverty.

- R, measures the effect of reranking on intertemporal poverty ($\boldsymbol{g}_1^{IR}$ and $\boldsymbol{g}_1^I$ show the same mean and the same degree of cross-sectional inequality, but differ with respect to the way individuals are ranked). Naturally, if reranking is observed during the growth spell, then $R = 0$. When individual ranks change, the values of the aversion parameters determines the sign of the R component. In the case $\alpha < \gamma$, R is strictly negative because reranking induces deprivation variability at the individual level and the costs of variability are deemed larger than the benefits of inequality associated with reranking. Alternatively, in the case $\alpha > \gamma$, R is strictly positive since reranking has an equalizing effect on poverty over time and this beneficial effect is more valued than the costs of variability. Finally, $\alpha = \gamma = 1$ implies $R = 0$.

- PG assesses a "pure" growth effect on intertemporal poverty ($\boldsymbol{g}_2$ and $\boldsymbol{g}_1^{IR}$ only differ with respect to their average value). This component is positive (negative) if "pure" growth is associated with a reduction in individuals' intertemporal poverty. Its sign is not determined by the values of α and γ, though the higher is γ with respect to α, the higher tends to be the absolute value of the effect.

When $\alpha = \gamma = 1$, $IPP_{\alpha,\gamma} = PG$: the pro-poor nature of any growth process is then solely determined by the "pure" growth effect. It can be noted that, contrary to PG, the component AG from the first decomposition is not purged from the inequality and reranking effects.[18]

[17] Ruiz-Castillo (2004) proposes a similar decomposition of the ethical index of mobility introduced by Chakravarty et al. (1985).

[18] As indicated earlier, this decomposition is characterized by path dependency. The value of the components would differ with alternative "paths" for the decomposition. For instance, we could have

With the example in (1), $I = 0$ given that inequality is identical in both periods; $R = -0.016$ for $\alpha = 3, \gamma = 4$, since there is a reshuffling of individuals in the distributions (the two initially poor individuals become the two richest), but the variability costs are higher than the benefits. Finally, $PG = 0$ given that the average gap is unchanged.

Finally, the studied growth spell is likely to last over a relatively long period and it may be desirable to isolate the contribution of a specific subperiod, provided the available data make it possible to perform a multi-period analysis ($T > 2$).

Let the intertemporal poverty measure $\Pi_{\alpha,\gamma}(\boldsymbol{G})$ be denoted by $\Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \ldots \boldsymbol{g}_T)$ and benchmark poverty, $\Pi_\alpha(\boldsymbol{g}_1)$, by $\Pi_\alpha(\boldsymbol{g}_1, \ldots \boldsymbol{g}_1)$. Assuming $T = 3$ and noting C^t the contribution of growth to $IPP_{\alpha,\gamma}$ from t to $t+1$, we then have the following decomposition of $IPP_{\alpha,\gamma}$:

$$IPP_{\alpha,\gamma} = \underbrace{\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_1, \boldsymbol{g}_1) - \Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_2, \boldsymbol{g}_2\big)}_{C^1} + \underbrace{\Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_2, \boldsymbol{g}_2\big) - \Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_2, \boldsymbol{g}_3\big)}_{C^2}. \quad (23)$$

As the result of the decomposition is likely to be path dependent, it may be worth considering a Shapley decomposition (Shorrocks 2013).[19]

considered to capture first the growth effect, then the impact of reranking, and lastly, the inequality one. No sequence can be regarded as necessarily more appropriate than another (see e.g. DiNardo, Fortin, and Lemieux 1996). A possible way of dealing with that issue is to apply a Shapley-Shorrocks decomposition, consisting of computing the Shapley value of each effect across all possible sequences (see Shorrocks 2013).

[19]The two components can then be computed as:

$$C^1 = \frac{1}{2}\Bigg(\Big(\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_1, \boldsymbol{g}_1) - \Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, f_{1,2}(\boldsymbol{g}_1), f_{1,2}(\boldsymbol{g}_1)\big)\Big)$$
$$+ \Big(\Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_1, f_{2,3}(\boldsymbol{g}_1)\big) - \Pi_{\alpha,\gamma}\Big(\boldsymbol{g}_1, f_{1,2}(\boldsymbol{g}_1), f_{2,3}\big(f_{1,2}(\boldsymbol{g}_1)\big)\Big)\Big)\Bigg), \quad (24)$$
$$= \frac{1}{2}\Bigg(\Big(\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_1, \boldsymbol{g}_1) - \Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_2, \boldsymbol{g}_2\big)\Big)$$
$$+ \Big(\Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_1, f_{2,3}(\boldsymbol{g}_1)\big) - \Pi_{\alpha,\gamma}(\boldsymbol{g}_1, \boldsymbol{g}_2, \boldsymbol{g}_3)\Big)\Bigg), \quad (25)$$
$$C^2 = \frac{1}{2}\Bigg(\Big(\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_1, \boldsymbol{g}_1) - \Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_1, f_{2,3}(\boldsymbol{g}_1)\big)\Big)$$
$$+ \Big(\Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, f_{1,2}(\boldsymbol{g}_1), f_{1,2}(\boldsymbol{g}_1)\big) - \Pi_{\alpha,\gamma}\Big(\boldsymbol{g}_1, f_{1,2}(\boldsymbol{g}_1), f_{2,3}\big(f_{1,2}(\boldsymbol{g}_1)\big)\Big)\Big)\Bigg), \quad (26)$$
$$= \frac{1}{2}\Bigg(\Big(\Pi_\alpha(\boldsymbol{g}_1, \boldsymbol{g}_1, \boldsymbol{g}_1) - \Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_1, f_{2,3}(\boldsymbol{g}_1)\big)\Big)$$
$$+ \Big(\Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_2, \boldsymbol{g}_2\big) - \Pi_{\alpha,\gamma}\big(\boldsymbol{g}_1, \boldsymbol{g}_2, \boldsymbol{g}_3\big)\Big)\Bigg). \quad (27)$$

where $f_{t,t+1}(\boldsymbol{g}_k) \equiv 1 - \boldsymbol{\delta}_{t,t+1}(1 - \boldsymbol{g}_k)$ with $\boldsymbol{\delta}_{t,t+1} \equiv \Big(\frac{\tilde{y}_{1,t}}{\tilde{y}_{1,t+1}}, \ldots \frac{\tilde{y}_{n,t}}{\tilde{y}_{n,t+1}}\Big)$.

5 Empirical illustration

Data are from the second, third and fourth rounds of the Indonesian Family Life Survey (IFLS) conducted by RAND, UCLA and the Demographic Institute of the University of Indonesia. The IFLS is an ongoing longitudinal socioeconomic and health survey, that contains over 30,000 individuals representing 83% of the Indonesian population living in 13 (out of 26) provinces, mostly on Sumatra and Java. Data are collected on individual respondents, their families, their households, the communities in which they live, and the health and education facilities they use (Strauss, Witoelar, Sikoki, and Wattie 2009). For the present study, we rely on expenditure estimates provided for the years 1997, 2000, and 2007. More specifically, our estimates are computed using per capita expenditures adjusted for inflation, using official CPI, and for regional price level differences, using regional poverty lines provided with the IFLS. Using Jakarta in 2007 as a reference for price levels, the poverty line is set at Rp264,383. It is worth noting that, though the time span of the growth spell is relatively large, we only have three observations for each household over the period. As a consequence, our results will mostly emphasize long-term dynamics. Short-terms dynamics are then not taken into account, hence resulting in an underestimation of the social cost or benefits (depending on $\alpha \gtrless \gamma$ of income variability at the individual level.

For the present study, each period is given the same weight for the estimation of the IPP index and its components.

Table 1: Cross-sectional and intertemporal EDE gaps for Indonesia, 1997–2007.

	Snapshot poverty			**Intertemporal poverty**		
	1997	2000	2007	$\beta = 1$	$\beta = 2$	$\beta = 3$
$\alpha = 1$	0.0419	0.0403	0.0162	0.0328	0.0476	0.0548
	(0.00201)	(0.00175)	(0.00107)	(0.00111)	(0.00155)	(0.0017)
$\alpha = 2$	0.13	0.123	0.073	0.0822	0.112	0.127
	(0.00373)	(0.00333)	(0.00311)	(0.00204)	(0.00235)	(0.00275)
$\alpha = 3$	0.202	0.191	0.13	0.123	0.159	0.18
	(0.00503)	(0.005)	(0.0048)	(0.0031)	(0.00288)	(0.00302)

Note: Boostraped standard errors in parentheses (200 replications).

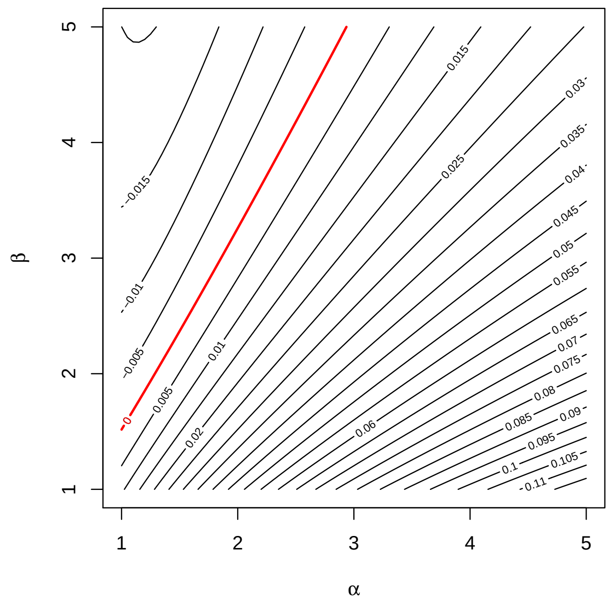
Table 1 shows both snapshot and intertemporal poverty estimates for values of α and γ within the set $\{1, 2, 3\}$. First, it can be seen that during the whole period, cross-sectional poverty has decreased. More specifically, poverty has not significantly changed between 1997 and 2000, but decreased substantially during the later subperiod whatever the value for α. These results are robust, i.e. do not depend on the specific value for the poverty line or the chosen poverty index within the set of monotone subgroup-consistent indexes.[20] The Asian crisis explains the deceiving results for the

[20]Cdf curves (not reported here but available upon request) are crossing and look very close for the

earliest subperiod, the per capita income representing in 2000 only 85% of its level in 1997.[21] The recovery and the sustained growth (about 4% per year between 2000 and 2007) have later been associated with poverty alleviation. It is worth noting that the pace of poverty alleviation over the period shrinks with the chosen value for α. This means that the growth process was less successful in lowering extreme poverty than moderate poverty.

The decreasing values for cross-sectional EDE gaps can be directly compared with the reported values for the intertemporal EDE gap as the same metric is used in both cases. Disregarding the welfare costs of income variability ($\gamma = 1$) the value for the intertemporal EDE gap is a simple average of snapshot EDE gaps. Raising the value of the income variability sensitivity parameter γ increases the EDE gap and thus offsets the observed improvement in cross-sectional poverty. When inequality aversion dominates variability aversion, the compensation is partial and the intertemporal EDE gap is lower than the corresponding value for 1997. But in the opposite situation, the social cost of income variability is regarded as so important that it fully cancels the observed improvement after 2000.

Figure 1: Sensitivity of $IPP_{\alpha,\gamma}$ with respect to α and γ, Indonesia 1997–2007.



The values of $IPP_{\alpha,\gamma}$ reported in Table 2 reflect these opposite effects, but nevertheless show that, unless variability aversion is large relatively to inequality aversion (see Figure 1), growth can be deemed intertemporally pro-poor in Indonesia during the period 1997–2007. The beneficial effect may even be regarded as substantial for some values of the parameters α and γ. For instance, with $\alpha = 3$ and $\gamma = 1$, we observe

years 1997 and 2000. The curve for the year 2007 is always lower for all income values.

[21] The same figures are reported for GDP per capita in 2011 PPP by the World Bank.

Table 2: Values of the IPP index for Indonesia, 1997–2007.

	$\beta = 1$	$\beta = 2$	$\beta = 3$	Max
$\alpha = 1$	0.00911	-0.00563	-0.0129	0.0279
	(0.00119)	(0.00119)	(0.0012)	(0.0008)
$\alpha = 2$	0.0479	0.0184	0.00286	0.0867
	(0.00282)	(0.00234)	(0.00245)	(0.0017)
$\alpha = 3$	0.0792	0.043	0.0225	0.134
	(0.00396)	(0.00326)	(0.0035)	(0.0022)

Note: Boostraped standard errors in parentheses (200 replications).

that the overall well-being shift and the mobility-as-equalizer effect have contributed to decrease by 7.9 percentage points the initial corresponding EDE gap. Compared with the maximum theoretical values of the $IPP_{\alpha,\gamma}$, our results underline significant progress with respect to poverty alleviation in Indonesia if we only consider the equalizing effects of mobility.

Table 3: Decomposition into anonymous (AG) and non-anonymous (M) for Indonesia, 1997–2007.

	AG	M		
		$\beta = 1$	$\beta = 2$	$\beta = 3$
$\alpha = 1$	0.00911	0	-0.0147	-0.022
	(0.00122)	..	(0.000464)	(0.000686)
$\alpha = 2$	0.0184	0.0295	0	-0.0155
	(0.00243)	(0.000842)	..	(0.00037)
$\alpha = 3$	0.0225	0.0567	0.0206	0
	(0.00307)	(0.00181)	(0.000563)	..

Note: Boostraped standard errors in parentheses (200 replications).

Table 3 shows that, taking an anonymous perspective, growth in Indonesia was unambiguously pro-poor during the whole period since the anonymous growth component AG is significantly positive. However, that anonymous effect is rather small over a 10-year period. Once the effects of mobility on intertemporal poverty are taken into account (i.e. $\gamma \neq \alpha$), mobility plays a decisive role in determining the sign of our intertemporal pro-poorness index. Indeed, the magnitude of the mobility sensitivity effect is relatively large in comparison with the anonymous growth component. This can be explained by the relatively low correlation between individual incomes in 1997 and 2000—Pearson's correlation coefficient is only 0.12 and becomes not significantly different from zero considering only those identified as poor from an intertemporal point of view— hence showing that mobility was high in the aftermath of the Asian crisis.[22]

Table 4: Subperiod contributions to the IPP index for Indonesia, 1997–2007.

	$\beta=1$		$\beta=2$		$\beta=3$	
	C^1	C^2	C^1	C^2	C^1	C^2
$\alpha=1$	0.00489	0.00423	-0.00573	0.000099	-0.0108	-0.00202
	(0.00127)	(0.00044)	(0.00125)	(0.00044)	(0.00118)	(0.00052)
$\alpha=2$	0.0325	0.0154	0.0109	0.00746	-0.000263	0.00312
	(0.0031)5	(0.00091)	(0.00267)	(0.00075)	(0.00255)	(0.00083)
$\alpha=3$	0.0564	0.0228	0.0297	0.0134	0.0146	0.00794
	(0.00445)	(0.00142)	(0.00404)	(0.00107)	(0.00362)	(0.00094)

Note: Boostraped standard errors in parentheses (200 replications).

This relatively high mobility associated with the first subperiod growth process explains why its contribution is relatively large (Table 4) though the cross-sectional income distributions are almost identical. With a marked aversion for extreme poverty ($\alpha = 3$, for instance), the mobility-as-equalizer effect during the period 1997–2000 was a large contributor to observed intertemporal growth pro-poorness between 1997 and 2007. Regarding the pattern of growth during the subperiod 2000–2007, the contribution has generally been positive, but it can be stressed that the magnitude of the contribution was relatively low compared with the first subperiod growth pattern.

The results of a further decomposition into anonymous changes and mobility effects are presented in Table 5. It can be seen that changes in the average poverty gap have contributed little to growth pro-poorness. Consequently, the anonymous component of the IPP index is mostly explained by changes in coss-sectional gap inequalities between the poor. The relative size of the mobility-as-equalizer effect M^c with respect to the cost of individual income variability CV depends primarily on the chosen values for the parameters α and γ.

Rank mobility was effective during the considered period in Indonesia and our results (Table 6) show its significant influence on intertemporal pro-poorness when income variability sensitivity is low ($\gamma = 1$). Our estimates finally show that changes in the cross-sectional relative distributions of gaps, net of the reranking effect, were significantly anti-poor from an intertemporal perspective and have been offset by the pro-poor effect of pure growth.

[22]Considering the subperiod 2000–2007, the two values for this correlation coefficient were respectively 0.43 and 0.04, both significantly different from zero.

Table 5: Decomposition into average poverty gap (ΔP^c), cross-sectional inequality (Δc^c), difference between intertemporal and unitemporal inequality (M^c), and variability (CV) for Indonesia, 1997–2007.

	$\beta = 1$				$\beta = 2$				$\beta = 3$			
	ΔP^c	Δc^c	M^c	CV	ΔP^c	Δc^c	M^c	CV	ΔP^c	Δc^c	M^c	CV
$\alpha = 1$	0.00911	0	0	0	0.00911	0	0	-0.0147	0.00911	0	0	-0.022
	(0.00128)	..	..	..	(0.00123)	..	..	(0.00048)	(0.00115)	..	..	(0.00071)
$\alpha = 2$	0.00911	0.0122	0.0266	0	0.00911	0.0122	0.0118	-0.0147	0.00911	0.0122	0.00353	-0.022
	(0.00116)	(0.00161)	(0.00073)	..	(0.00118)	(0.00151)	(0.00051)	(0.00044)	(0.00115)	(0.00158)	(0.00053)	(0.00067)
$\alpha = 3$	0.00911	0.0187	0.0513	0	0.00911	0.0187	0.0299	-0.0147	0.00911	0.0187	0.0166	-0.022
	(0.00128)	(0.00332)	(0.00163)	..	(0.00123)	(0.00304)	(0.00111)	(0.00051)	(0.00123)	(0.00297)	(0.00094)	(0.00077)

Note: Boostraped standard errors in parentheses (200 replications).

Table 6: Decomposition into inequality change (I), reranking (R), and pure growth (PG) for Indonesia, 1997–2007.

	$\beta = 1$			$\beta = 2$			$\beta = 3$		
	I	R	PG	I	R	PG	I	R	PG
$\alpha = 1$	0	0	0.0092	-0.00261	-0.0164	0.0135	-0.0042	-0.0243	0.0157
	..	..	(0.000481)	(0.000135)	(0.000289)	(0.000752)	(0.000209)	(0.00042)	(0.000821)
$\alpha = 2$	-0.0156	0.0349	0.0287	-0.0218	0	0.0403	-0.0273	-0.0168	0.0471
	(0.000975)	(0.000601)	(0.00138)	(0.00148)	..	(0.0023)	(0.00165)	(0.000283)	(0.00231)
$\alpha = 3$	-0.0393	0.0679	0.0507	-0.0502	0.0226	0.0708	-0.0608	0	0.0835
	(0.00242)	(0.00136)	(0.00252)	(0.0033)	(0.000403)	(0.0039)	(0.0037)	..	(0.00415)

Note: Boostraped standard errors in parentheses (200 replications).

6 Conclusion

Many studies have challenged the issue of testing the pro-poor nature of growth, but focusing on snapshot evaluations of poverty. In the present paper we argue that a comprehensive assessment of the pro-poor nature of a growth spell may require a shift from the traditional cross-sectional perspective to a longitudinal one, so as to account fully for the dynamics of individual deprivations over time.

For that purpose, a family of aggregate indexes of intertemporal pro-poorness is introduced. While previous studies are essentially based on the comparison of the initial and final income distributions, we suggest here performing an evaluation of growth pro-poorness using the joint distribution of income, hence considering more information than usually provided by marginal or conditional income distributions. The proposed family of intertemporal pro-poorness indexes aggregates "equally distributed equivalent" measures of the sequence of poverty gaps experienced by each individual in the population. An appealing feature of these indexes is their ability to capture both the cost of deprivation variability and the benefit of intertemporal equalization associated with mobility. Different decomposition procedures are also introduced to disentangle the different contributions of pure growth, cross-sectional and intertemporal inequalities, exchange mobility, and temporal variability in explaining the intertemporal pro-poorness of any growth process.

This measurement framework is illustrated using panel data for Indonesia between 1997 and 2007. Although the Indonesian population was severely hit by the Asian crisis in the late 1990s, we show that growth could be deemed pro-poor from an intertemporal perspective unless we assumed marked aversion with respect to individual income variability. Changes in cross-sectional poverty have positively contributed to these beneficial changes, but mobility was also substantial during the period of analysis and had noticeable effects on intertemporal poverty.

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Appendix

Generalization to T periods

As mentioned in the main text, the decompositions provided in this paper can be generalized to time horizons of $T > 2$ periods.

The first decomposition is obtained by adding and subtracting in (15) the EDE of periodic individual poverty as follows:

$$\underbrace{\Pi_\alpha(\boldsymbol{g}_1) - \Pi_\alpha(\boldsymbol{g})}_{AG} + \underbrace{\Pi_\alpha(\boldsymbol{g}) - \Pi_{\alpha,\gamma}(\boldsymbol{g})}_{M}$$

To generalize the second decomposition, observe that (12) can be rewritten as:

$$\Pi_{\alpha,\gamma}(\boldsymbol{g}) = \omega_1 P_1(\boldsymbol{g}_1) + \omega_2 P_2(\boldsymbol{g}_2) + ... + \omega_T P_T(\boldsymbol{g}_T) + c_{\alpha,\gamma}(\boldsymbol{g}) + \frac{1}{n}\sum_{i=1}^{n} c_\gamma((g))$$

$IPP_{\alpha,\gamma}$ can then be decomposed as:

$$\underbrace{\omega_2\left[P_1(\boldsymbol{g}_1) - P_1(\boldsymbol{g}_2)\right] + \omega_3\left[P_1(\boldsymbol{g}_1) - P_1(\boldsymbol{g}_3)\right] + ... + \omega_T\left[P_1(\boldsymbol{g}_1) - P_1(\boldsymbol{g}_T)\right]}_{\Delta P^c} +$$

$$+\underbrace{\omega_2\left[c_\alpha(\boldsymbol{g}_1) - c_\alpha(\boldsymbol{g}_2)\right] + \omega_3\left[c_\alpha(\boldsymbol{g}_1) - c_\alpha(\boldsymbol{g}_3)\right] + ... + \omega_T\left[c_\alpha(\boldsymbol{g}_1) - c_\alpha(\boldsymbol{g}_T)\right]}_{\Delta c^c}$$

$$+\underbrace{\left[\omega_1 c_\alpha(\boldsymbol{g}_1) + \omega_2 c_\alpha(\boldsymbol{g}_2) + \omega_3 c_\alpha(\boldsymbol{g}_3) + ... + \omega_T c_\alpha(\boldsymbol{g}_T)\right] - c_{\alpha,\gamma}(\boldsymbol{g})}_{M^c} +$$

$$+\underbrace{\frac{1}{n}\sum_{i=1}^{n} c_\gamma(\boldsymbol{g}_{(i)})}_{CV}$$

Lastly, when $T > 2$, the third decomposition can be obtained as:

$$\underbrace{\left[\Pi_{\alpha,\gamma}(\boldsymbol{g}_1) - \Pi_{\alpha,\gamma}(\boldsymbol{g}_1^I)\right]}_{I} + \underbrace{\left[\Pi_{\alpha,\gamma}(\boldsymbol{g}_1^I) - \Pi_{\alpha,\gamma}(\boldsymbol{g}_1^{IR})\right]}_{R} +$$

$$+\underbrace{\left[\Pi_{\alpha,\gamma}(\boldsymbol{g}_1^{IR}) - \Pi_{\alpha,\gamma}(\boldsymbol{g})\right]}_{PG}$$

Here, $\boldsymbol{g}^I = (\boldsymbol{g}_1, ..., \boldsymbol{g}_t^I, ..., \boldsymbol{g}_T^I)$, where $\boldsymbol{g}_t^I$ denotes the counterfactual distribution of poverty gaps at time t obtained by preserving the same average poverty gaps and ranks as observed in the first period distribution. Similarly, $\boldsymbol{g}^{IR} = (\boldsymbol{g}_1, ..., \boldsymbol{g}_t^{IR}, ..., \boldsymbol{g}_T^{IR})$, where

$\boldsymbol{g}_t^{IR}$ denotes the counterfactual time-specific distribution of poverty gaps obtained by keeping the same average poverty gap as that of the first period distribution.