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Older or Wealthier? The Impact of Age Adjustment on Wealth Inequality*

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Abstract

Differences in individual wealth holdings are widely viewed as a driving force of economic inequality. However, as this finding relies on cross-sectional data, a concern is that older is confused with wealthier. We propose a new method to adjust for age effects in cross-sections, which eliminates wealth inequality due to age, yet preserves inequality arising from other factors. Using a new cross-country comparable database, we examine the impact of age adjustments on wealth inequality across countries and over time. We find that the most widely used method yields a substantially different picture of age-adjusted wealth inequality than our method.

Keywords: Wealth inequality; life cycle; age adjustments; Gini coefficient

JEL classification: *D*31; *D*63; *D*91; *E*21

I. Introduction

The distribution of wealth is an important determinant of overall economic inequality as well as a marker for the types of activities that are rewarded in an economy. Wealth inequality is also a matter of considerable interest in the bodies of literature on economic growth, institutions and development, occupational choice and entrepreneurship, as well as in asset pricing.[1] New sources of cross-country comparable microdata suggest that the wealth

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[1] See, for example, Gollier (2001) for a study of wealth inequality and asset pricing, Mookherjee and Ray (2002) for a review of the literature on wealth inequality, economic growth, institutions, and development, and Meh (2005) for a recent analysis of wealth inequality in relation to occupational choice and entrepreneurship.

holdings of individuals vary substantially within and across countries. In most countries, the Gini coefficient for wealth is reported to be twice that of income. Moreover, the world distribution of wealth is found to be much more concentrated than the world distribution of income.[2]

Because of data availability, however, this recent evidence on wealth inequality is based on cross-sectional data. This is potentially problematic as both theoretical models and empirical results suggest a strong age–wealth relationship (see, for example, Davies and Shorrocks, 2000). The age–wealth profile is firmly established as increasing during the working lifespan and usually declining somewhat after retirement. Hence, a snapshot of wealth inequality within a country runs the risk of confusing older with wealthier, and thereby providing a misleading picture of the differences in the lifetime wealth of its citizens.

For this reason, it has long been argued that age adjustments of cross-sectional measures of inequality are necessary (see, for example, Atkinson, 1971). Age adjustment allows us to utilize the cross-sectional data at our disposal, while avoiding some of the potential pitfalls associated with using these data. In particular, age-adjusted inequality measures can be used to evaluate whether changes in wealth inequality over time occur because of changes in the age structure or whether there are other forces at play. Moreover, age adjustments can be useful when comparing wealth inequality across countries, by controlling for differences resulting from cross-country variation in age–wealth profiles and age structure.

In this paper, we investigate whether cross-sectional wealth inequality measures are sensitive to differences in wealth holdings over the life cycle. We consider how age adjustments might influence the wealth-inequality ranking of countries as well as the time trend in wealth inequality in a country. In some respects, our approach recalls the pioneering paper of Paglin (1975), which first raised the question of the effect of age on inequality, and its trend. While the validity of the Paglin–Gini (PG) has been questioned from several perspectives, which we address in our analysis, the issue of age adjustment of inequality measures remains an important research question.[3] In fact, given the aging of the large baby boom cohorts born after World War II, the issue can be viewed as potentially more important now than in the earlier period (1947–1972) considered by Paglin.

Our first contribution is that we propose a new method to adjust for age effects. Unlike existing methods, this addresses the fact that individuals

[2] See, for example, Wolff (1996), Davies and Shorrocks (2000), Davies *et al.* (2006), and Sierminska *et al.* (2006).

[3] The Paglin approach to age adjustment was subject to three rounds of comments and replies in the *American Economic Review* (Paglin, 1977, 1979, 1989); it has numerous citations, and it continues to be subject to controversy.

differ both with respect to age and with respect to other wealth-generating factors. For example, an individual's education level is not only an important determinant of his wealth, but it is also correlated with his age. Existing methods assume that the unconditional distribution of mean wealth by age represents the age effects and will, therefore, not only eliminate wealth inequality attributable to age but also differences because of factors correlated with age, such as education. In contrast, the method proposed in this paper eliminates inequality because of age, yet preserves inequality arising from all other factors. To this end, a multivariate regression model is employed, allowing us to isolate the net age effects while holding other determinants of wealth constant. Next, we derive a new, age-adjusted Gini coefficient, where perfect equality requires that each individual receives a share of total wealth equal to the proportion of wealth he would hold if all wealth-generating factors except age were the same for everyone in the society. Our method can be viewed as a generalization of the approach to age adjustments proposed by Wertz (1979) and it is important in situations where omitted-variables bias is a major concern.[4]

Our second contribution is that we provide a theoretical foundation to assess the properties of age-adjusted inequality measures. In particular, we propose a set of conditions that are similar to those underlying the classical Gini coefficient in all respects but one: the equalizing wealth is not given by the mean wealth in the society as a whole, but depends on the ages of the individuals. In the spirit of Paglin (1975), an inequality in a society that is not age-adjusted requires that all individuals have equal lifetime wealth, but not that individuals at all ages must have equal wealth holdings in any given year. Furthermore, we explore the relationship between our age-adjusted Gini coefficient, the classical Gini coefficient, and alternative age-adjusted inequality measures.

Our final contribution is that we examine empirically the impact of age adjustments on the wealth-inequality ranking of countries as well as on the time trend for wealth inequality in Italy and the US. To this end, we use data from Canada, Finland, Germany, Italy, Sweden, the UK, and the US, collected from the new, cross-country comparable Luxembourg Wealth Study (LWS) database. We find that the ranking of wealth distributions is quite sensitive to the method used to make age adjustments. In particular, the much-used PG is shown to yield a substantially different picture of wealth inequality from our method. Interestingly, our new age-adjusted Gini coefficient provides a wealth-inequality ranking of countries that comes close

[4] Even though Danziger *et al.* (1977), Minarik (1977), and Kurien (1977), in early comments to Paglin (1975), point out that adjusting appropriately for age effects requires a well-specified multivariate model, we are not aware of any study that adjusts for age effects while controlling for other determinants of individual income or wealth holdings.

to the ranking based on the classical Gini coefficient, which disregards age effects. A possible interpretation is that age adjustments might be less important than previous studies have suggested, albeit this conclusion might not necessarily hold true for other applications.

This is the first study to examine the impact of age adjustments on the wealth-inequality ranking of countries. However, several studies have investigated the effect of adjusting for age effects on wealth and income inequality in a given country. Paglin (1975) studied the effect of age adjustment on the distribution of income and wealth in the US. He concluded that the classical Gini coefficient overstates wealth and income inequality and, moreover, that age adjustments convert a flat time trend in income inequality into a declining time profile. Formby *et al.* (1989) extend this work by analyzing the time period 1980–1986. They found that inequality has risen faster according to the PG than the classical Gini coefficient over this period.[5] Mookherjee and Shorrocks (1982) have studied income inequality in the UK. They find that the adjustment for age converts an apparent upward trend in overall income inequality into a declining time profile, according to the PG, which is horizontal when using strictly decomposable inequality measures to make age adjustments. By contrast, Pudney (1993) suggests that only a small part of observed income and wealth inequality in China can be explained by age effects. None of the above studies uses methods that adjust for age effects while controlling for other income or wealth-generating factors.

In Section II, we set out the proposed method to identify and adjust for age effects, and we explore its relationship to the classical Gini coefficient as well as to existing age-adjusted inequality measures. In Section III, we describe the data and clarify definitional issues. In Section IV, we discuss the results using the different age-adjusted wealth-inequality measures, before concluding in Section V.

II. Age Adjustment of Inequality

The proposed method for age adjustment of inequality can be described as a three-step procedure. First, a new age-adjusted Gini coefficient (*AG*) is derived. Second, a multivariate regression model is employed, allowing us to isolate the net age effects on wealth while holding other determinants of wealth constant. Third, the wealth distribution that characterizes perfect equality in age-adjusted wealth is determined.

[5] Other studies that have attempted to adjust for age effects on income inequality estimates for the US include Danziger *et al.* (1977), Minarik (1977), Nelson (1977), and Friesen and Miller (1983). For a review, see Formby *et al.* (1989).

Below, we describe the three steps of our method, before examining the relationship between AG, the classical Gini coefficient (G), and alternative age-adjusted inequality measures.

A New Age-Adjusted Gini Coefficient

Consider a society consisting of n individuals where every individual i is characterized by the pair $(w_i, \widetilde{w}_i)$, where w_i denotes the actual wealth level and $\widetilde{w}_i$ is the equalizing wealth level in a given year. If actual and equalizing wealth are the same for all individuals and they live equally long, there is perfect equality of lifetime wealth in this society. As is clear when we define the equalizing wealth level formally (in the third subsection of Section II), the equalizing wealth is the same for all individuals belonging to the same age group in this society; it is a function of individual i's age, but not of any other individual characteristics. If no other wealth-generating factor is correlated with age, the equalizing wealth is simply the mean wealth of each age group. Furthermore, if there are no age effects on wealth, the equalizing wealth will be equal to the mean wealth in the society as a whole.

The joint cross-sectional distribution Y of actual and equalizing wealth is given by

$$Y = [(w_1, \widetilde{w}_1), (w_2, \widetilde{w}_2), ..., (w_n, \widetilde{w}_n)].$$

Let Ξ denote the set of all possible joint distributions of actual and equalizing wealth, such that the sum of actual wealth equals the sum of equalizing wealth. Suppose that the social planner imposes the following modified versions of the standard conditions on an inequality partial ordering defined on the alternatives in Ξ, where $A \preceq B$ represents that there is at least as much age-adjusted inequality in B as in A.[6] Let μ denote the mean wealth of the population as a whole, and let Δ_i represent the difference between individual i's actual wealth w_i and equalizing wealth $\widetilde{w}_i$. Let the distributions of such differences for the two distributions $[\Delta_i(A) = w_i(A) - \widetilde{w}_i(A)$ and $\Delta_i(B) = w_i(B) - \widetilde{w}_i(B)]$ be sorted in ascending order.

Condition 1. ***Scale invariance***: *for any* $a > 0$ *and* $A, B \in \Xi$, *if* $A = aB$, *then* $A \sim B$.

Condition 2. ***Anonymity***: *for any permutation function* ρ: $n \to n$ *and for* $A, B \in \Xi$, *if* $[w_i(A), \widetilde{w}_i(A)] = [w_{\rho(i)}(B), \widetilde{w}_{\rho(i)}(B)]$ *for all* $i \in n$ *then* $A \sim B$.

[6] See Almås *et al.* (forthcoming) for analogous conditions imposed to study the equality of opportunity.

Condition 3. ***Unequalism***: *for any* A, $B \in \Xi$ *such that* $\mu(A) = \mu(B)$, *if* $\Delta_i(A) = \Delta_i(B)$ *for every* $i \in n$, *then* $A \sim B$.

Condition 4. ***Generalized Pigou–Dalton criterion***: *for any* A, $B \in \Xi$, *if there exist two individuals* s *and* k *such that* $\Delta_s(A) < \Delta_s(B) \leq \Delta_k(B) < \Delta_k(A)$, $\Delta_i(A) = \Delta_i(B)$ *for all* $i \neq s, k$, *and* $\Delta_s(B) - \Delta_s(A) = \Delta_k(A) - \Delta_k(B)$, *then* $A \succ B$.

Scale invariance states that, if all actual and equalizing wealth levels are rescaled by the same factor, then the level of age-adjusted inequality remains the same. Anonymity implies that the ranking of alternatives should be unaffected by a permutation of the identity of individuals. Unequalism entails that the social planner is only concerned with how unequally each individual is treated, defined as the difference between his actual and equalizing wealth.[7] Finally, the generalized version of the Pigou–Dalton criterion states that any fixed transfer of wealth from an individual i to an individual j, where $\Delta_i > \Delta_j$, reduces age-adjusted inequality.

The AG is based on a comparison of the absolute values of the differences in actual and equalizing wealth between all pairs of individuals, and is defined as

$$AG = \frac{\sum_j \sum_i |(w_i - \widetilde{w}_i) - (w_j - \widetilde{w}_j)|}{2\mu n^2}. \quad (1)$$

It is straightforward to see that AG satisfies Conditions 1–4. Note that these conditions are similar to those underlying G in all respects but one: the equalizing wealth is not given by the mean wealth in the society as a whole, but depends on the age of the individuals.

Because it is straightforward to construct age-adjusted Lorenz curves based on the distribution of differences between actual and equalizing wealth, it is by no means necessary to focus on the Gini coefficient: other inequality indices that are based on the Lorenz curve, such as the Bonferroni index, can also form the basis for age adjustments.

Identifying the Net Age Effects

Suppose that the wealth level of individual i at a given point in time depends on the age group a that he belongs to as well as his lifetime resources given as a function h of a vector X of individual characteristics:

$$w_i = f(a_i)h(X_i). \quad (2)$$

[7] This condition can therefore be viewed as analogous to the Focus axiom in poverty analysis, stating that a poverty index should focus entirely on the incomes of the poor (see, for example, Foster and Shorrocks, 1991).

The functional form of f depends on the underlying model of wealth accumulation. In the simplest life-cycle model, there is no uncertainty, individuals earn a constant income until retirement, and the interest rate, as well as the rate of time preference, is zero. In this model, the wealth of an individual increases up to retirement and declines afterwards. If the earnings profile is upward sloping, the model predicts borrowing in the early part of the life cycle. The fact that this is not always observed could be explained by credit-market imperfections. The introduction of lifetime uncertainty and non-insurable health hazards induces the elderly to hold assets for precautionary purposes, which reduces the rate at which wealth declines during retirement. If the sole purpose of saving is to leave a bequest to one's children, individuals behave as if their horizons were infinite and wealth does not decline with age.

Given the theoretical ambiguity of f, we specify a flexible functional form, yielding the wealth-generating function

$$\ln w_i = \ln f(a_i) + \ln h(X_i) = \delta_i + X_i' B, \tag{3}$$

where δ_i gives the percentage wealth difference of being in the age group of individual i relative to some reference age group, holding all other variables constant. Because of the right skewness combined with the sparse tail of the wealth distribution, our log-linear specification is preferable to a linear specification. As net wealth might be negative, we therefore add to each wealth observation a constant equal to the absolute value of the minimum wealth observation when estimating the log-linear specification. This is simply a matter of adjusting the location of the distribution.[8] Equation (3) is estimated by ordinary least-squares (OLS) separately for each country. The key assumption underlying this estimation is that there are no omitted factors correlated with age that determine individual wealth holdings. In that case, we obtain consistent estimates of the net age effects on wealth.

It is important to emphasize that the objective of the estimation of equation (3) is not to explain as much variation as possible in wealth holdings, but simply to obtain an empirically sound estimate of the effects of age on wealth. Drawing on the findings of Jappelli (1999) and Hendricks (2007) of individual characteristics correlated with wealth, X includes educational attainment in our baseline specification. When performing robustness analysis, we extend the set of controls to include sex, number of children, industry and occupation of household head, region of residence, marital status, immigration status, and spouse's characteristics. The reasons for not including these variables in the baseline specification are twofold. First,

[8] In this regard, it should be noted that the properties of inequality measures based on the Gini coefficient are preserved when applied to distributions with zero and negative values (see, for example, Amiel *et al.*, 1996).

we do not have data on all the variables for every country under study. In addition, some of the variables are potentially endogenous to individual wealth holdings. In any case, we show that our results are robust to the inclusion of the additional controls.

Existing age-adjusted inequality measures, discussed in detail in the final subsection of Section II, implicitly assume a stationary economy, implying no cohort effects. Consequently, they risk confounding age effects with cohort effects, as these factors are perfectly collinear in a cross-section. A novelty of this paper is that we make an effort to separate age effects from cohort effects. As pointed out by Heckman and Robb (1985), it is necessary to impose some structure on the cohort effects in order to address this identification problem. Jappelli (1999) and Kapteyn *et al.* (2005) explore reasons why different cohorts accumulate different amounts of wealth. They have found that productivity growth is the primary determinant of differences in wealth across cohorts; productivity growth generates differences in permanent incomes across cohorts, which feeds into the wealth accumulation of individuals belonging to different generations. Following Masson (1986), we assume that the age cross-sections and the cohort profiles of wealth (in constant prices) coincide, except for a constant state of real growth. If wealth grows at the rate g, then the typical profile for any given cohort is $(1+g)$ times larger than that for the one-year-older cohort. When estimating equation (3), we therefore inflate each individual's wealth value by the factor $(1+g)^{\text{age}}$. Mirer (1979) shows that under commonly accepted assumptions in life-cycle theory, the growth rate of wealth is equal to the growth rate of income between successive cohorts. To adjust the observed wealth levels for economic growth across cohorts, we use an annual growth rate of 2.5 percent. Our results are robust to other choices of growth rates.

The assumption of a stationary economy also implies no intra-cohort mobility in individual wealth holdings, which has been criticized by, for example, Johnson (1977) and Friesen and Miller (1983). By conditioning on individual characteristics, the assumption of parallel age–wealth profiles might be more reasonable for *AG* than for existing age-adjusted inequality measures. However, just as any other study measuring inequality using cross-sectional data, this paper admittedly comes short of fully accounting for the effects of intra-cohort mobility. Yet, it is reassuring that several studies suggest that accounting for mobility has little impact on country ranking by income inequality (see, for example, Burkhauser and Poupore, 1997; Aaberge *et al.*, 2002).

Defining Equalizing Wealth

Identifying the net age effect is only part of the job; we also need to find a consistent way of adjusting for age effects when there are other wealth-generating factors. There is a considerable body of literature concerning the problem of how to adjust for some, but not all, income-generating factors when the income function is not additively separable (see, for example, Bossert and Fleurbaey, 1996; Kolm, 1996). The problem of adjusting for age effects on wealth is analogous. To eliminate wealth differences attributable to age but preserving inequality arising from all other factors, we employ the so-called general proportionality principle proposed by Bossert (1995) and Konow (1996), and further studied in Cappelen and Tungodden (2007). Then, the absence of age-adjusted inequality requires that any two individuals belonging to a given age group have the same wealth level. Moreover, in any situation where everyone has the same wealth-generating factors except age, there should be no lifetime wealth inequality.

More formally, the equalizing wealth level of individual i depends on his age as well as every other wealth-generating factor of all individuals in the society. It is formally defined as

$$\widetilde{w}_i = \frac{\mu n \sum_j f(a_i)h(X_j)}{\sum_k \sum_j f(a_k)h(X_j)} = \frac{\mu n e^{\delta_i}}{\sum_k e^{\delta_k}}, \tag{4}$$

where e^{δ_k} gives the net age effect of belonging to the age group of individual k after integrating out the effects of other wealth-generating factors correlated with age. No age-adjusted inequality corresponds to every individual i receiving $\widetilde{w}_i$, which is the share of total wealth equal to the proportion of wealth an individual from his age group would hold if all wealth-generating factors except age were the same for everyone in the population. If there is no age effect on wealth, the equalizing wealth level is equal to the mean wealth level in the society.

Relationship to the Classical Gini Coefficient

The classical Gini coefficient is defined in equation (5):

$$G(Y) = \frac{\sum_j \sum_i |(w_i - \mu) - (w_j - \mu)|}{2\mu n^2}. \tag{5}$$

By comparing this expression to equation (1), we can see that there is a very close link between G and AG. Both measures are based on a comparison of the absolute values of the differences in actual and equalizing wealth levels between all pairs of individuals. The distinguishing feature is how equalizing wealth is defined. For G, the equalizing wealth level is assumed to be μ. Perfect equality requires not only equal lifetime wealth, but additionally that individuals of all ages must have the same wealth holding in any given year, which can be realized only if there is a flat age–wealth profile. However, a flat age–wealth profile runs counter to both consumption needs over the life cycle and productivity variation depending on human capital investment and experience. Indeed, the relationship between wealth and age can produce wealth inequality at a given point in time even if everyone is completely equal in all respects but age. As transitory wealth differences even out over time, a snapshot of inequality produced by G runs the risk of producing a misleading picture of actual variation in lifetime wealth. In comparison, AG abandons the assumption of a flat age–wealth profile and allows equalizing wealth to depend on the age of the individuals. In doing so, AG purges the cross-sectional measure of inequality of its inter-age or life-cycle component. If $\widetilde{w}_i = \mu$ for all individuals in every age group, the age–wealth profile is flat and AG coincides with G.

To obtain further intuition on the similarities and differences between G and AG, it is helpful to see the correspondence between the standard representation of the Lorenz curve and a Lorenz curve expressed in differences between actual wealth and mean wealth in the society as a whole. Figure 1 displays standard and difference-based Lorenz curves for the same wealth distribution. The area between the standard Lorenz curve and the diagonal of the upper diagram (the line of equality) is identical to the area between the difference-based Lorenz curve and the horizontal axis (the line of equality) in the lower diagram. In both cases, G is equal to twice the area ‘A’, between the Lorenz curve and the line of equality.

In a similar way, we can draw the age-adjusted Lorenz curve underlying AG, expressing the differences between actual wealth and the equalizing wealth in the population. Just as for G, AG is equal to twice the area between this difference-based Lorenz curve and the horizontal axis (line of equality). When drawing age-adjusted Lorenz curves, however, individuals are ordered not by their wealth per se, as in Figure 1, but according to the difference between actual and equalizing wealth.

Both G and AG reach their minimum value of 0, if everyone receives their equalizing wealth. Moreover, both measures take their maximum when the difference between actual and equalizing wealth is at its highest possible level. Specifically, G reaches its maximum value of 1, if one individual holds all wealth. In comparison, AG takes its maximum of 2 in the hypothetical situation where the equalizing wealth of the individual who has all the wealth is zero, and the equalizing wealth of one of the individuals with no wealth is equal to the aggregate wealth in the economy. The fact that AG and G range over different intervals is therefore a direct result of their different views of perfect equality; age-adjusted inequality is not only a result of differences in individuals' actual wealth holdings, but also a result of differences in equalizing wealth between individuals at different points in the life cycle.

By the same token, AG will be smaller (greater) than G whenever the differences in individuals' wealth holdings because of age are positively

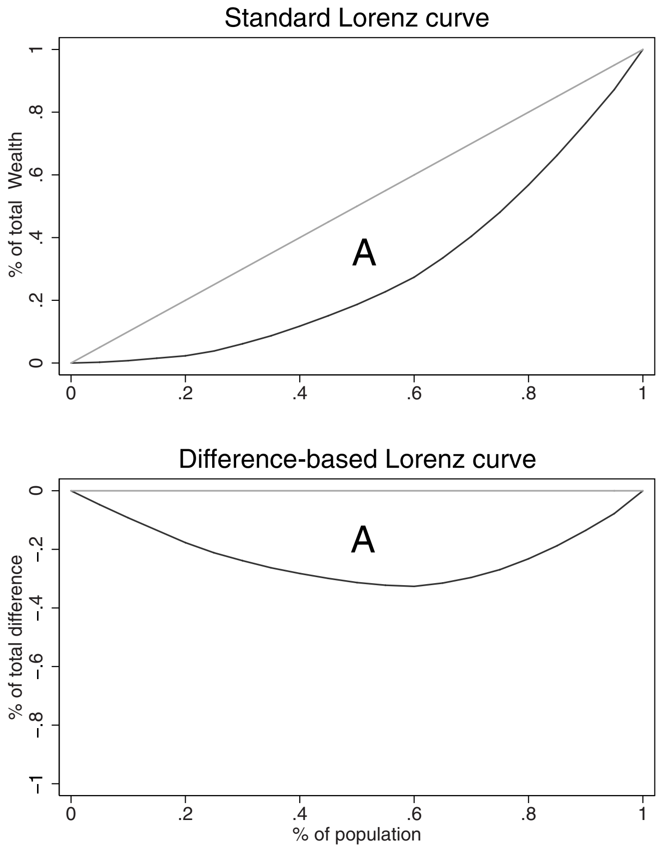

Fig. 1. Standard and difference-based representations of the classical Lorenz curve

(negatively) correlated with differences in individuals' wealth attributable to other wealth-generating factors.[9] For example, an individual with zero wealth will contribute less to inequality in AG than in G whenever his equalizing wealth level is lower than the mean wealth in the society.

Relationship to Existing Age-Adjusted Inequality Measures

There are two distinguishing aspects of age-adjusted inequality measures. First, they hold different views on how equalizing wealth should be

[9] To see this, let $\epsilon_i = w_i - \widetilde{w}_i$ for any individual i, and note that AG and G have the same denominator. While the numerator of AG aggregates $|\epsilon_i - \epsilon_j|$ over all pairs of individuals, the numerator of G aggregates $|(\widetilde{w}_i + \epsilon_i) - (\widetilde{w}_i + \epsilon_j)|$ of all pairs of individuals. Hence, $G > AG$ whenever $\mathrm{cov}(\widetilde{w}, \epsilon > 0)$.

measured. Second, they differ in the way they aggregate up the differences between actual and equalizing wealth. In this paper, we consider two alternative age-adjusted inequality measures: *PG* and Wertz–Gini (*WG*). These both have the same objective as *AG*, namely to purge *G* applied to snapshots of wealth inequality of its inter-age or life-cycle component. In particular, the condition of a flat age–wealth profile is relaxed. Below, we use Conditions 1–4 to assess the properties of *PG* and *WG*, and to characterize their relationships to *AG*.

Because of its close relationship to *AG*, it is convenient to first consider *WG*. This was proposed by Wertz (1979), who claims that *PG* fails to adjust properly for age effects. However, his comment has been largely ignored, perhaps because he does not put up conditions that allow a formal assessment of the properties of *PG* and *WG*. Let *WG* be defined by

$$WG(Y) = \frac{\sum_j \sum_i |(w_i - \mu_i) - (w_j - \mu_j)|}{2\mu n^2}, \tag{6}$$

where μ_i and μ_j denote the mean wealth levels of all individuals belonging to the age group of individual i and j, respectively. Like *AG*, *WG* is based on a comparison of the absolute values of the differences in actual and equalizing wealth levels between all pairs of individuals and ranges over the interval [0, 2]. It is also straightforward to see that it satisfies Conditions 1–4.

The distinguishing feature between *AG* and *WG* is that the latter measure defines the equalizing wealth of an individual i as the unconditional mean wealth levels in his age group, μ_i, whereas the former measure defines his equalizing wealth as the net age effect of belonging to his age group after integrating out the effects of other wealth-generating factors correlated with age, $\widetilde{w}_i$. Any differences between *AG* and *WG* are therefore a result of omitted-variables bias in using μ_i to measure equalizing wealth. As is well known, the omitted-variables bias in μ_i is given by the effect of the omitted variables on wealth times the regression of omitted variables on age (see, for example, Angrist and Pischke, 2009). For example, cohort is perfectly collinear with age and will therefore bias the age effect in so far as it is correlated with wealth. Another example is education, which is correlated with both age and wealth. The omitted-variables bias formula tells us that *WG* will be equal to *AG* whenever age is uncorrelated with omitted wealth-generating factors. Hence, *AG* can be viewed as a generalization of *WG*, and it is important in situations where omitted-variables bias is a major concern.

Next, consider the much-used *PG*, which can be expressed as

$$PG(Y) = \frac{\sum_j \sum_i (|w_i - w_j| - |\mu_i - \mu_j|)}{2\mu n^2}, \tag{7}$$

where μ_i and μ_j denote the mean wealth levels of all individuals belonging to the age group of individuals i and j, respectively. Applying the standard Gini decomposition, PG can be rewritten as

$$PG = G - G_{\mathrm{b}} = \sum_i \theta_i G_i + R. \tag{8}$$

Here, G_{b} represents the Gini coefficient that would be obtained if the wealth of each individual in every age group were replaced by the relevant age group mean μ_i, G_i represents the Gini coefficient of wealth within the age group of individual i, θ_i is the weight given by the product of this group's wealth share $(n_i\mu_i)/\mu n$ and population share n_i/n (where n_i is the number of individuals in the age group of individual i), and R captures the degree of overlap in the wealth distributions across age groups (see, for example, Lambert and Aronson, 1993).[10]

Both WG and PG define the equalizing wealth of an individual as the mean wealth level of the age group he belongs to, disregarding the fact that other wealth-generating factors are correlated with age. Unlike AG, these might eliminate not only inequality because of age but also inequality because of these other factors.

In addition, PG stands out in the way it aggregates up the differences in actual and equalizing wealth. Specifically, PG is based on a comparison of differences in the absolute values of actual and equalizing wealth levels between all pairs of individuals, $|(w_i - w_j)| - |(\mu_i - \mu_j)|$. This runs counter to the unequalism condition, because $|(w_i - w_j)| - |(\mu_i - \mu_j)| = 0$ does not necessarily imply that $|(w_i - \mu_i) - (w_j - \mu_j)| = 0$. The following numerical example shows that PG violates this condition. Consider two countries A and B with two age groups, each consisting of two individuals. Suppose that country A's distribution of actual and equalizing wealth, $[w_i(A), \mu_i(A)]$, is given by

$$A = [(20, 60), (100, 60), (60, 80), (100, 80)],$$

whereas country B's distribution of $[w_i(B), \mu_i(B)]$ is given by

$$B = [(0, 40), (80, 40), (80, 100), (120, 100)].$$

In both countries, the distribution of differences between the actual and equalizing wealth, $w_i - \mu_i$, is given by $[\{-40, 40\}, \{-20, 20\}]$. According to the unequalism condition, age-adjusted inequality measures should be the same when the distributions of differences between actual and

[10] Overlap implies that the wealth holdings of the richest person in an age group with a relatively low mean wealth level exceeds the wealth holdings of the poorest person in an age group with a higher mean wealth level (i.e., $w_i < w_j$ and $\mu_i > \mu_j$ for at least one pair of individuals i and j).

equalizing wealth are the same. While WG satisfies this condition, PG violates it.[11]

Arguably, the unequalism condition is an intuitively appealing condition as it ensures that age-adjusted inequality measures follow G in measuring inequality according to the differences in actual and equalizing wealth, between all pairs of individuals, rather than the aggregated differences in actual wealth minus the aggregated differences in equalizing wealth.[12]

As $|(w_i - w_j) - (\mu_i - \mu_j)|$ provides an upper bound for $|(w_i - w_j)| - |(\mu_i - \mu_j)|$, it follows that $WG \geq PG$. This raises the questions, under which conditions will WG be equal to PG and, subsequently, can we be sure that the two measures produce the same inequality ranking? As stated in Proposition 1, PG will differ from WG if there is any age effect on wealth, provided that there is some within-age-group wealth variation.

Proposition 1. *For any distribution Y, $WG(Y) \geq PG(Y)$, with strong inequality whenever $\mu_i \neq \mu_j$ for at least one pair of individuals and $w_i \neq \mu_i$ or $w_j \neq \mu_j$ for at least one of these individuals.*

(The Proof is provided in Appendix A.)

As shown in the Proof of Proposition 1, overlap in the wealth distributions across age groups (i.e., $R > 0$) is a sufficient, but not a necessary, condition for $WG > PG$. A corollary to Proposition 1 is therefore that PG is likely to yield a different ranking from WG in situations where countries differ substantially in the degree of overlap. This result relates to a major controversy surrounding PG, namely whether or not R should be treated as an inter-age or a within-age-group component.[13] Until recently, the issue was unsettled simply because little was known about the overlap term; Shorrocks and Wan (2005), for example, refer to R as a 'poorly specified' element of the Gini decomposition. However, Lambert and Decoster (2005) provide a novel characterization of the properties of R, showing first that R unambiguously falls as a result of a within-group progressive transfer, and second that R increases when the wealth holding in the poorer group

[11] Specifically, $WG(A) = WG(B) = 0.25$, whereas $PG(A) = 0.179 \neq PG(B) = 0.107$.

[12] Our numerical example illustrates the difference. Consider distribution A and the contribution to age-adjusted inequality from the comparison of the richest individuals in the two age groups, for which $[w_i(A), \mu_i(A)]$ is given by (100, 60) and (100, 80). Paglin advocates that perfect equality corresponds to everyone receiving the mean wealth of their age group. A wealth comparison of this pair of individuals should thus contribute with 20 to age-adjusted inequality, which is captured by the numerator of WG. By contrast, the numerator of PG records a -20 contribution to age-adjusted inequality – the rationale for which is hard to grasp.

[13] Nelson (1977) and others argue that R is part of inter-age inequality and should thus be netted out when constructing age-adjusted inequality measures. Paglin (1977), however, maintains that R is capturing within-group inequality and that PG is accurately defined.

is scaled up, and reaches a maximum when means coincide. This makes Lambert and Decoster (2005, p. 378) conclude that "the overlap term in R is at once a between-groups and a within-groups effect: it measures a between-groups phenomenon, overlapping, that is generated by inequality within groups". Therefore, $R=0$ is necessary (although not sufficient) for PG to net out the inter-age component, and nothing but the inter-age component, from cross-sectional inequality measures.

III. Data and Definitions

Recently, the availability and quality of data on household wealth have improved. Household surveys of assets and debt have previously suffered from non-sampling errors because of high non-response and misreporting rates. In addition, comparative studies of wealth distributions have been beset by comparability problems because of methodological and data issues ranging from the basic problem of index numbers to differences in the methods and definitions used in the various countries. Today, the data problems are mitigated by oversampling of wealthy people in surveys as well as by utilizing supplementary information, such as administrative data from tax and estate registers. The LWS – an international project to collect and harmonize existing microdata on household wealth into a coherent database – has reduced the comparability problems. We use the LWS database, and select the following seven countries because of data availability: Canada, Finland, Germany, Italy, Sweden, the UK, and the US.[14]

It should be noted that we follow previous studies of wealth distributions using the LWS database in excluding Austria, Norway, and Cyprus from the analysis (see, for example, Sierminska *et al.*, 2006). We drop Norway because of the inconsistency stemming from valuing real estate on a taxable basis and debt at market prices,[15] Cyprus because over 60 percent of the observations lack information on net wealth, and Austria because it lacks data on net wealth. Finland's 1994 survey is also excluded because this dataset lacks information on education.

We follow common practice and focus on the distribution of household net wealth, which refers to material assets that can be sold in the marketplace less any debts, thereby excluding pension rights as well as human capital. Net wealth consists of financial assets and non-financial assets net of total debt. Total debts refer to all outstanding loans. Financial assets include deposit accounts, stocks, and mutual funds, whilst non-financial

[14] See Sierminska *et al.* (2006) and the LWS homepage http://www.lisproject.org/lwstechdoc.htm for a detailed description of the LWS database.

[15] Statistics Norway estimates that in the 1990s the taxable value of houses was, on average, less than a third of their market values (see Harding *et al.*, 2004). The majority of Norwegians are therefore registered with negative net wealth.

assets consist of the principal residence and other real-estate investments.[16] Business assets are not included.

In this paper, we use the household as the economic unit. This is in part because assets are recorded at the household level but also to conform to previous studies of wealth distributions. Households with missing values for wealth, education, or age of household head are dropped. To compare the wealth holdings of singles and couples, we assign each married/cohabiting spouse a wealth level equal to his or her net household wealth divided by the square root of 2. Robustness analysis demonstrates that our results are unaffected by the choice of equivalence scale.

To define age groups, we follow common practice and rely on information about the age of the household head. To be specific, we define seven age groups: 24 years and younger, 25–34 years, 35–44 years, 45–54 years, 55–64 years, 65–74 years, and 75 years and older.[17] There are no household heads older than 75 years in the Swedish data. In all countries, we categorize the education variable into four educational groups. The four groups correspond as close as possible to the following categories: 'high-school dropout', 'high-school graduate', 'non-university post-secondary certificate', and 'university degree or certificate'.

In the robustness analysis, we run a battery of specification checks, adding further controls to equation (3), including number of children, marital status, region of residence, immigrant status, as well as sex, occupation, and industry of household head. Marital status is divided into five categories: 'single without children', 'single with children', 'couple without children', 'couple with children', and 'others'. Industry and occupation are included using the country's own categories.

IV. Empirical Analysis

Descriptive Statistics

Overall, the descriptive statistics are consistent with previous evidence in showing substantial variation among OECD countries in the age structure

[16] The self-assessed current value of the principal residence and other real-estate investments is reported for all countries except for Sweden, where the tax value is reported. However, Statistics Sweden calculates the ratios of purchase prices to tax values for different types of houses and geographical locations, and uses these to inflate the tax values. For comparability issues, it is also comforting that the principal residence represents almost the same share of total assets in Sweden as in neighboring country Finland (61 versus 64 percent).

[17] Formby *et al.* (1989) and Paglin (1989) discuss the theoretical effects of the choice of the widths of the age groups on age adjustments of inequality. The results of Formby *et al.* (1989) suggest, however, that age-adjusted inequality estimates are not substantially different for age groups of one-, five-, and ten-year intervals.

(Burkhauser *et al.*, 1997; Banks *et al.*, 2003) as well as savings patterns (Borsch-Supan, 2003).[18]

Table 1 demonstrates that there is considerable variation in the demographic structure of the seven OECD countries examined in this study. First and foremost, the age structure differs substantially across the countries. For instance, Italy has, on average, older household heads, which might be because Italians move out from their parents' house later in life than is typical in most OECD countries (see, for example, Manacorda and Moretti, 2006). By contrast, Canada as well as the Nordic countries, Sweden and Finland, have relatively young household heads. The fact that the age structure differs means that the inequality ranking of countries might be affected by age adjustments, even if countries have the same age–wealth profile.

Table 1 also reveals a considerable change over time in the age structure in the US. As a result of the large, but temporary, increase in the population growth rate following World War II, the population shares of middle-aged and older household heads have increased significantly from 2000 to 2006. Because the middle-aged and elderly have, on average, accumulated more wealth than the young, changes in age composition might potentially affect the trend in inequality.

Furthermore, Table 1 demonstrates significant cross-country differences in educational attainment. In particular, the educational level is substantially lower in Italy compared with the US and Germany. The US also stands out with the highest mean wealth, whereas Canada and the Nordic countries have the lowest. This might not be a surprise given the differences in the scope of the public savings programs between these countries (see, for example, Klevmarken *et al.*, 2003).

Figure 2 reveals that there is not only a considerable variation in the age structure across the countries, but also a substantial variation in the age–wealth relationship. This also indicates that cross-country comparisons of inequality could potentially be affected by age effects. In particular, the US has a markedly more hump-shaped age–wealth profile than the rest of the countries. In contrast, there seems to be relatively little life-cycle savings in Sweden, which corresponds to what is found by Klevmarken (2006).

Estimation Results

Equation (3) is estimated separately by OLS for each country, and separately for each cross-section for the US and Italy, for which we have data for more than one year.

[18] See Sierminska *et al.* (2006) for a detailed discussion of the descriptive statistics of the LWS database.

Table 1. *Descriptive statistics by country*

	Canada 1999	Germany 2001	Italy 2002	Italy 2004	Sweden 2002	UK 2000	US 2000	US 2003	US 2006	Finland 1998
Mean age	48.6	50.6	55.7	55.9	47.2	49.1	50.4	50.9	51.8	47.2
Age comp. (percent)										
24 and younger	4.32	2.44	0.52	0.66	4.10	5.84	3.56	3.25	3.08	4.30
25–34 years	16.99	13.44	7.17	7.14	16.57	17.16	13.21	12.25	12.02	15.47
35–44 years	24.27	23.43	18.60	18.47	21.86	21.81	21.43	19.59	18.39	25.10
45–54 years	20.51	21.41	21.83	21.45	24.43	18.47	24.71	24.60	23.47	27.34
55–64 years	14.01	18.78	21.37	20.60	22.84	14.90	17.15	21.16	21.89	16.08
65–74 years	11.52	13.57	18.53	18.62	10.19	12.13	11.50	11.70	12.85	7.30
75 years and older	8.37	6.94	11.98	13.06	NA	9.69	8.44	7.46	8.29	4.41
Education (percent)										
High-school dropout	27.75	1.48	35.48	33.00	21.48	27.65	6.70	5.96	5.31	30.87
High-school graduate	21.83	12.89	28.07	28.41	44.17	24.34	28.81	27.72	28.89	36.49
Post-secondary	27.12	50.05	28.28	29.80	6.24	45.37	16.00	15.50	15.13	22.16
University degree	23.30	35.58	8.17	8.80	27.31	2.62	48.49	50.83	50.67	9.44
Female head (percent)	NA	31.47	30.59	32.95	29.29	39.69	13.27	13.17	12.34	28.87

Notes: Data sources are national household wealth surveys included and harmonized in the LWS database. Household weights are used to ensure nationally representative results. Wealth levels are expressed in 1000 US dollars (PPP adjusted using Penn World Table 6.3).

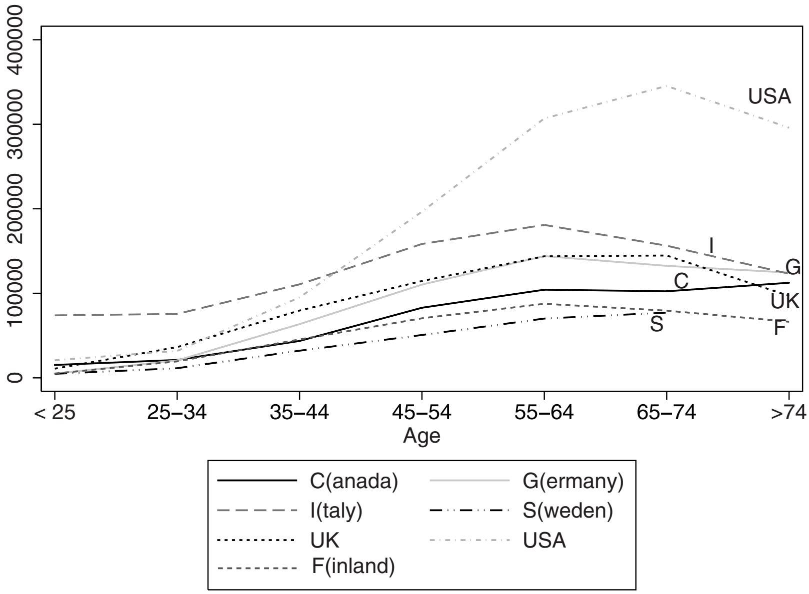

Fig. 2. Age–wealth profile by country
Notes: Data sources are national household wealth surveys included and harmonized in the LWS database. The studies for Italy (2002) and the US (2000) are included. Household weights are used to ensure nationally representative results. Wealth levels are expressed in international dollars from Penn World Table 6.3.

The fairly precise estimation results presented in Table 2 reveal a standard hump-shaped age–wealth profile where wealth increases during the working lifespan and declines somewhat after retirement in most countries. Wealth generally increases with education; the increase is, however, larger in Canada, the UK, and Italy, than in the other countries.

It is also evident from Table 2 that the explanatory power differs substantially across the countries. Canada and Italy have higher R^2 than the other countries, whereas Sweden has by far the lowest R^2, mirroring the cross-country differences in life-cycle saving. As emphasized in the second subsection of Section II, the main purpose of these regressions is not to explain as much of the wealth-generating process as possible, but rather to back out an empirically sound estimate of the net age effect. Hence, variation in goodness-of-fit measures across countries is a concern in so far as it reflects cross-country differences in omitted-variables bias, rather than differences in unobservables unrelated to age. Below, we report the results from a battery of robustness checks addressing the concern for omitted-variables bias, none of which changes the results of the analysis.

Table 2. *Estimation results of the log-linear wealth regression: baseline specification*

	Canada 1999	Germany 2001	Italy 2002	Italy 2004	Sweden 2002	UK 2000	US 2000	US 2003	US 2006	Finland 1998
25–34 years	0.008	−0.003	−0.005	−0.029	0.000	0.027	−0.002	−0.003	−0.001	0.003
	(0.010)	(0.000)	(0.037)	(0.041)	(0.000)	(0.005)	(0.000)	(0.000)	(0.001)	(0.005)
35–44 years	0.155	0.004	0.147	0.138	0.004	0.097	−0.000	0.003	0.007	0.054
	(0.010)	(0.001)	(0.036)	(0.040)	(0.000)	(0.006)	(0.000)	(0.000)	(0.001)	(0.005)
45–54 years	0.365	0.019	0.401	0.373	0.010	0.195	0.006	0.010	0.025	0.103
	(0.012)	(0.001)	(0.037)	(0.041)	(0.000)	(0.007)	(0.000)	(0.000)	(0.001)	(0.015)
55–64 years	0.585	0.037	0.641	0.637	0.017	0.257	0.016	0.028	0.046	0.200
	(0.014)	(0.001)	(0.037)	(0.041)	(0.001)	(0.033)	(0.000)	(0.001)	(0.001)	(0.008)
65–74 years	0.751	0.044	0.750	0.833	0.028	0.377	0.025	0.036	0.075	0.225
	(0.016)	(0.001)	(0.038)	(0.044)	(0.001)	(0.014)	(0.001)	(0.001)	(0.002)	(0.013)
75 years and older	0.879	0.055	0.778	0.841	NA	0.353	0.029	0.047	0.092	0.248
	(0.020)	(0.002)	(0.042)	(0.046)	(NA)	(0.013)	(0.001)	(0.001)	(0.002)	(0.019)
High-school graduate	0.125	0.010	0.189	0.192	0.005	0.089	0.007	0.008	0.013	0.039
	(0.009)	(0.002)	(0.016)	(0.017)	(0.000)	(0.009)	(0.000)	(0.001)	(0.001)	(0.008)
Post-secondary	0.123	0.016	0.404	0.405	0.010	0.135	0.010	0.014	0.019	0.061
	(0.009)	(0.002)	(0.019)	(0.020)	(0.001)	(0.015)	(0.000)	(0.001)	(0.001)	(0.020)
University degree	0.276	0.032	0.590	0.631	0.010	0.261	0.023	0.032	0.063	0.174
	(0.010)	(0.002)	(0.026)	(0.026)	(0.001)	(0.028)	(0.000)	(0.001)	(0.001)	(0.018)
Constant	12.729	16.418	13.038	13.111	16.526	14.250	17.960	17.594	17.052	14.240
	(0.010)	(0.002)	(0.037)	(0.042)	(0.000)	(0.012)	(0.000)	(0.001)	(0.001)	(0.007)
R^2	0.309	0.056	0.227	0.243	0.010	0.083	0.090	0.098	0.134	0.065
Number of observations	26,035	24,731	13,386	13,240	24,640	6,953	36,935	37,355	36,840	6,737

Notes: The table reports the results from the OLS estimation of equation (3). Data sources are national household wealth surveys included and harmonized in the LWS database. Household weights are used to ensure nationally representative results. Reference categories: 24 years and younger and high-school dropout. Heteroskedasticity-robust standard errors in parentheses. There are no households in the sample for Sweden with the age of household head 75 years and older.

Table 3. *Wealth-inequality ranking of countries according to different measures*

	Canada 1999	Germany 2001	Italy 2002	Sweden 2002	UK 2000	US 2000	Finland 1998
G	0.752(4/5)	0.752(4/5)	0.576(1)	0.880(6)	0.694(3)	0.914(7)	0.584(2)
PG	0.446(3)	0.500(5)	0.476(4)	0.612(7)	0.428(2)	0.528(6)	0.380(1)
WG	0.760(5)	0.754(4)	0.572(2)	0.862(6)	0.678(3)	1.080(7)	0.548(1)
$AG^{\text{no controls}}$	0.728(4)	0.749(5)	0.576(2)	0.878(6)	0.681(3)	0.912(7)	0.572(1)
AG	0.730(4)	0.750(5)	0.587(2)	0.878(6)	0.680(3)	0.912(7)	0.572(1)

Notes: Data sources are national household wealth surveys included and harmonized in the LWS database. Household weights are used to ensure nationally representative results. Country ranking is given in parentheses.

Age-Adjusted Estimates of Wealth Inequality

In this section, we investigate how age adjustments might influence the wealth-inequality rankings of countries as well as the time trend in wealth inequality within a country. First, it should be noted that the age-adjusted inequality measures, such as G, are ordinal in nature, and any monotonic transformation of such a measure will preserve its ranking of distributions. This means that the numerical values of these inequality measures are primarily of interest as a way of comparing and ordering the distributions. The fact that the measures range over different intervals is therefore beside the point.[19]

The first row of Table 3 reports wealth-inequality results using G for the seven countries under study. We can see that the reported G for wealth is substantially larger than that for income.[20] It is also evident that Italy has the least unequal wealth distribution followed by Finland, whereas the US and Sweden have the strongest concentrations of wealth among their citizens. Figure 3 shows the time trend in wealth inequality for Italy and the US. We can see that G suggests a slight decrease in inequality in both countries.

The low wealth inequality in Finland corresponds well to its low income inequality. In comparison, the high wealth inequality in Sweden contrasts with its low income inequality, but conforms to findings from other studies (see, for example, Domeij and Klein, 2002; Sierminska *et al.*, 2006). This is, to a large extent, driven by the large fraction of households with zero or negative net wealth in Sweden compared with other countries. Domeij and

[19] As shown in the fourth subsection of Section II, G can range from 0 to 1, PG from 0 to G, and AG and WG from 0 to 2. Normalizing these measures so that they range over the same interval is possible, but it will not affect the ranking of the wealth distributions for any of the measures.

[20] G for income for the seven countries under study is reported to be as follows: Canada, 0.33 (2000); Finland, 0.27 (2000); Germany, 0.28 (2000); Italy, 0.36 (2000); Sweden, 0.25 (2000); the UK, 0.36 (1999); the US, 0.41 (2001) (WDI, 2010).

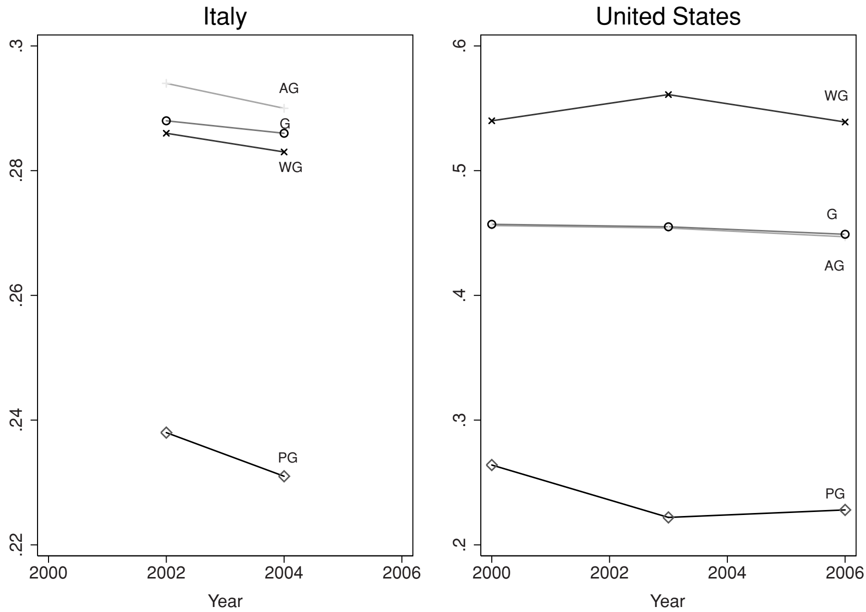

Fig. 3. Time trend in wealth inequality for the US and Italy
Notes: Data sources are national household wealth surveys included and harmonized in the LWS database. Household weights are used to ensure nationally representative results.

Klein (2002) suggest that Sweden's redistributive public pension scheme can account for much of the difference between the degree of inequality in its income and its wealth distribution. However, it is not clear that the explanation of the public pension scheme is consistent with the evidence that Finland and Sweden have quite similar income inequality but widely different wealth inequality. An alternative explanation for the high wealth concentration in Sweden is that it was not affected by the main economic and geopolitical shocks that have been identified as major causes of decreased top wealth shares in other developed countries: Sweden did not participate in either of the world wars and was not affected by the Great Depression (Roine and Waldenström, 2009).

Rows 2–5 of Table 3 report age-adjusted inequality measures for the seven countries under study, while Figure 3 shows the time trends in age-adjusted wealth inequality for Italy and the US. The insights from these results can be summarized in three conclusions. First, the country rankings of wealth distributions and the time trends in wealth inequality are quite sensitive to the method used to make age adjustments. In particular, the much-used *PG* is shown to yield a substantially different picture of wealth inequality than *AG* and *WG*. Second, *AG* produces a wealth-inequality ranking of countries that comes close to the ranking based on *WG*, albeit the wealth-inequality time trend for the US differs substantially when using *AG* compared with *WG*. Nevertheless, the way age-adjusted inequality measures aggregate up the difference between actual and equalizing wealth seems to play a larger role than omitted-variables bias, as the inequality rankings differ more between *WG* and *PG* than between *AG* and *WG*. Third, *AG* produces a very similar ranking as *G*. Although this might be reassuring for statistical offices and government agencies, which regularly rely on *G* to evaluate cross-sectional wealth distributions, this conclusion might not necessarily hold true for other applications.[21]

Turning to a more detailed investigation of the empirical results using the different measures, let us first consider the results using *PG*, reported in the second row of Table 3. We can see that *PG* yields a very different picture of wealth inequality than *G*. For example, according to *PG*, the wealth inequality in Sweden is higher than that in the US, a result that runs counter to findings from other age-adjusted inequality measures as well as *G*. Moreover, *PG* alters the ranking of Italy from having clearly the most equal wealth distribution to being more unequal than Finland, the UK, and Canada. It is also evident that Canada and Germany change order in the country rankings when using *PG*. In addition, Figure 3 reveals that *PG* produces a different time trend of wealth inequality in the US, compared with *G*. Overall, our findings for *PG* conform well to the Paglin (1975) study of income and wealth inequality in the US over the period 1947–1972, in suggesting that age adjustments change the picture of inequality.

As shown in equation (8), *PG* might yield a different wealth-inequality ranking than *G* in so far as there is significant cross-country variation in between-group inequality, G_b. Because G_b is a population weighted average of the different age-group means, it increases as a result of larger disparity in mean wealth across age groups. For example, Figure 3 shows that the US has a much stronger age–wealth relationship than Sweden, which explains why *PG* alters the rankings of the two countries. Furthermore, G_b increases with the number of people in the age groups with relatively low and relatively high mean wealth levels. For instance, Italy has a relatively compressed age distribution compared with Canada, which explains the change in the country rankings when measuring inequality using *PG* instead of *G*. In comparison, Canada and the UK experience a similar

[21] For example, Almås *et al.* (forthcoming) use the method proposed in this paper to study the time trend in earnings inequality in Norway over the last few decades. They find that *G* and *AG* yield substantially different time trends in earnings in Norway. Furthermore, the time trends in *AG* and *WG* differ substantially. A possible explanation is that the correlation between education and earnings is, in fact, much stronger than the correlation between education and wealth.

decrease from G and PG because their age–wealth profiles and age structure are quite similar.[22]

Next, consider the results using WG, reported in the third row of Table 3 and Figure 3. We can see immediately that the country rankings of wealth distributions and the time trends in wealth inequality are quite sensitive to the way age-adjusted inequality measures aggregate up the differences between individuals' actual and equalizing wealth. In line with Proposition 1, WG is greater than PG for all countries. This is, in part, because of considerable overlap in the wealth distributions across age groups. Indeed, the overlap term R, defined in equation (7), ranges between 0.196 (Finland, 1998) and 0.355 (US, 2000) in the countries under study. This cross-country variation in the degree of overlap also contributes to explaining the large change in the wealth-inequality ranking of countries. For example, WG evaluates Germany as more equal than Canada, whereas PG evaluates Canada as more equal than Germany. At the same time, R is considerably larger in Canada than in Germany (0.232 for Germany and 0.260 for Canada).

The two last rows report the inequality rankings based on AG. Specifically, the last row uses the estimated age effects reported in Table 2 to compute the equalizing wealth levels defined by equation (4) and the associated AG given by equation (1). In comparison, the fourth row drops the controls for education in equation (3), so that the only distinguishing feature from WG is the adjustment for economic growth across cohorts in the identification of age-group mean wealth levels. Any difference between WG and AG without controls is therefore attributable to omitted-variables bias in the former measure because of cohort effects, whereas the difference between AG without controls and AG with controls is a result of omitted-variables bias in the former measure because of education.

We can see that the country rankings according to AG without controls are quite similar to those of WG. The exceptions are the rankings of Canada and Germany. In addition, Figure 3 reveals that WG suggests a rise in wealth inequality in the US from 2000 to 2002, whereas AG without controls indicates a small decline. When comparing the last two rows, it is clear that AG with and without controls produces the same picture of wealth inequality. In fact, the point estimates are very similar. This implies that education is not an important source of omitted-variables bias in age-adjusted inequality. To understand why, recall that the omitted-variables bias depends on the effect of the omitted variables on wealth times the

[22] Specifically, the values of G_b for the different countries are as follows: 0.306, Canada (1999); 0.204, Finland (1998); 0.251, Germany (2001); 0.102, Italy (2002); 0.110, Italy (2004); 0.269, Sweden (2002); 0.268, UK (2000); 0.386, US (2000); 0.466, US (2003); 0.443, US (2006).

regression of omitted variables on age. Table 2 shows that wealth generally increases with education. Furthermore, when regressing age on education we find that younger cohorts have a higher level of education than older cohorts.[23] However, the magnitude of these effects is too small to change the wealth-inequality ranking. The relatively small omitted-variables bias is mirrored in Table 4, showing that the estimated age effects on wealth without controls for education are quite similar to those with controls for education, reported in Table 2. We also see that the omitted-variables bias is strongest in Italy, because of the relatively strong effects of education on wealth and age on education in that country.

Finally, it should be noted that the country ranking according to *AG* is quite similar to that produced by *G*. The exceptions are that the age adjustment makes Finland more equal than Italy and Canada more equal than Germany. As discussed in Section II, *AG* will be smaller (greater) than *G* whenever the differences in individuals' wealth holdings because of age are positively (negatively) correlated with differences in individuals' wealth holdings attributable to other wealth-generating factors. The fact that the estimates of *G* and *AG* are generally quite similar therefore implies that the correlation is fairly small. This suggests that individuals who have relatively high equalizing wealth because of the age group they belong to do not have systematically different wealth holdings because of other wealth-generating factors.

Robustness Analysis of the Age-Adjusted Inequality Measure

We run a battery of robustness checks to examine whether the results from our age-adjusted inequality measure are sensitive to the inclusion of additional controls, choice of growth adjustment, and use of equivalence scale. In some cases, the robustness analysis is performed only for a subset of the countries because of data availability. As summarized in Table 5, the main picture is that the country ranking by wealth inequality is robust to the alternative specifications.[24]

To be specific, the country ranking is unaffected by adding number of children and marital status to the set of controls for all countries (*AG*(1)). Moreover, extending the set of controls to include occupation, and industry, and sex of household head (*AG*(2)) does not alter the picture of inequality. The same holds true when we also control for immigration status and region (*AG*(3)), and when using the subsample of couple households to control for age and education of the spouse (*AG*(4)). Acknowledging the

[23] The regression results of age on education are reported in Almås and Mogstad (2010).
[24] The robustness analysis undertaken is described in more detail in Almås and Mogstad (2010).

Table 4. *Estimation results of the log-linear wealth regression: no controls*

	Canada 1999	Germany 2001	Italy 2002	Italy 2004	Sweden 2002	UK 2000	USA 2000	USA 2003	USA 2006	Finland 1998
25–34 years	0.048	0.003	0.051	0.010	0.001	0.031	0.001	0.001	0.003	0.019
	(0.009)	(0.000)	(0.044)	(0.041)	(0.000)	(0.004)	(0.000)	(0.000)	(0.000)	(0.002)
35–44 years	0.183	0.010	0.196	0.156	0.005	0.094	0.003	0.007	0.013	0.065
	(0.010)	(0.001)	(0.043)	(0.039)	(0.000)	(0.005)	(0.000)	(0.000)	(0.000)	(0.003)
45–54 years	0.395	0.025	0.418	0.359	0.011	0.176	0.009	0.015	0.031	0.107
	(0.012)	(0.001)	(0.044)	(0.040)	(0.000)	(0.007)	(0.000)	(0.000)	(0.001)	(0.017)
55–64 years	0.586	0.042	0.582	0.549	0.017	0.225	0.018	0.033	0.054	0.190
	(0.014)	(0.001)	(0.045)	(0.040)	(0.001)	(0.030)	(0.000)	(0.001)	(0.001)	(0.007)
65–74 years	0.724	0.048	0.602	0.651	0.027	0.319	0.026	0.037	0.076	0.205
	(0.015)	(0.001)	(0.045)	(0.042)	(0.001)	(0.012)	(0.001)	(0.001)	(0.002)	(0.013)
75 years and older	0.836	0.058	0.597	0.623	NA	0.289	0.029	0.047	0.088	0.222
	(0.020)	(0.002)	(0.048)	(0.043)	(NA)	(0.012)	(0.001)	(0.001)	(0.002)	(0.017)
Constant	12.838	16.433	13.298	13.417	16.531	14.365	17.970	17.607	17.078	14.278
	(0.009)	(0.000)	(0.042)	(0.038)	(0.000)	(0.003)	(0.000)	(0.000)	(0.000)	(0.001)
R^2	0.276	0.047	0.128	0.148	0.009	0.062	0.055	0.064	0.079	0.050
Number of observations	26,035	24,731	13,386	13,240	24,640	6,953	36,935	37,355	36,840	6,737

Notes: The table reports the results from the OLS estimation of equation (3). Data sources are national household wealth surveys included and harmonized in the LWS database. Household weights are used to ensure nationally representative results. Reference category: 24 years and younger. Heteroskedasticity-robust standard errors in parentheses.

Table 5. *Wealth-inequality ranking of countries according to AG by specification*

	Canada 1999	Germany 2001	Italy 2002	Sweden 2002	UK 2000	US 2000	Finland 1998
AG	0.730(4)	0.750(5)	0.587(2)	0.878(6)	0.680(3)	0.912(7)	0.572(1)
AG(1)	0.731(4)	0.749(5)	0.589(2)	0.878(6)	0.679(3)	0.912(7)	0.572(1)
AG(2)	NA	0.749(4)	0.608(2)	0.880(5)	0.680(3)	0.912(6)	0.572(1)
AG(3)	NA	0.749(3)	0.606(1)	0.878(4)	0.680(2)	NA	NA
AG(4)	NA	0.713(4)	0.568(2)	0.841(5)	0.645(3)	0.921(6)	0.548(1)
AG(5)	0.831(4)	0.880(5)	0.687(1)	1.076(7)	0.768(3)	0.983(6)	0.698(2)
AG(6)	0.727(4)	0.751(5)	0.592(2)	0.879(6)	0.683(3)	0.913(7)	0.576(1)
AG(7)	0.728(4)	0.749(5)	0.582(2)	0.878(6)	0.680(3)	0.912(7)	0.572(1)
AG(8)	0.733(4)	0.749(5)	0.593(2)	0.878(6)	0.680(3)	0.912(7)	0.571(1)
AG(9)	0.729(4)	0.749(5)	0.589(2)	0.878(6)	0.678(3)	0.912(7)	0.572(1)

Notes: Data sources are national household wealth surveys included and harmonized in the LWS database. Household weights are used to ensure nationally representative results. Country ranking is given in parentheses. *AG* is the baseline specification controlling for education.
AG(1): Estimation adding number of children and marital status as a control variables to baseline specification.
AG(2): Estimation adding sex of household head, number of children, occupation, industry, and marital status as control variables to baseline specification.
AG(3): Estimation adding number of children, occupation, industry, marital status, region, and immigration status as control variables to baseline specification.
AG(4): Estimation adding spouse's education and age as control variables to baseline specification.
AG(5): Estimation based on the subsample of single households using baseline specification.
AG(6): Estimation based on the EU equivalence scaling using baseline specification.
AG(7): Estimation based on a growth rate of 2 percent using baseline specification.
AG(8): Estimation based on a growth rate of 3 percent using baseline specification.
AG(9): Estimation based on polynomials of continuous age variables using baseline specification.

inherent arbitrariness in the choice of equivalence scale, we use an alternative equivalence scale (*AG*(6)) and find that the ranking is unchanged. On top of this, we make sure that the choice of economic growth rate does not affect our results by applying alternative growth rates (*AG*(7)–*AG*(8)). Finally, we check that using polynomials of continuous age variables instead of age-group dummies does not change the country ranking (*AG*(9)).

However, when we restrict our sample to singles (*AG*(5)), the ranking changes somewhat. A motivation for this specification check is that the common practice of using equivalence scales to capture pooling of wealth and economics of scale within the household might be too crude. However, as being single is potentially endogenous to individuals' wealth holdings, we need to be cautious in interpreting these results. With this caveat in mind, we can see that restricting the sample to singles alters the ranking of Finland from having the most equal distribution to being more unequal than Italy. It is also evident that Sweden and the US change order in the country ranking when looking only at singles. Furthermore, our results demonstrate that *AG* is generally higher within the sample of singles compared with the population as a whole. There are several possible explanations. On the one hand, negative marital sorting on wealth could contribute to lower inequality for the full sample compared with the subsample of singles. On the other hand, the high inequality within the sample of singles could simply reflect that this a very heterogeneous group of people, making the comparison difficult both across and within age groups.

In line with the other results above, the time trends in inequality in Italy and the US are robust to the alternative specifications. However, the results change somewhat when we examine the time trend in the subsamples (see Almås and Mogstad, 2010).

V. Concluding Remarks

A strong relationship between age and wealth implies that inequality of wealth at a given point in time is likely to exist even in a society where everyone is completely equal in all respects other than age. It has therefore been argued that age adjustments of inequality measures based on cross-sectional data are necessary.

In this paper, we have proposed a method to adjust for age effects in cross-sectional data, which eliminates transitory inequality but preserves inequality arising from other factors. Applying a cross-country comparable wealth database, we have found smaller effects of age adjustment than existing approaches. Interestingly, our new age-adjusted Gini coefficient provides a wealth-inequality ranking of countries that comes quite close to the ranking based on the classical Gini coefficient, which disregards age effects. A possible interpretation is that age adjustments are less important than previous studies have suggested, albeit this conclusion might not necessarily hold true for other applications.

There are a number of other applications where life-cycle effects matter. For example, theoretical models and empirical results suggest a strong relationship between age and earnings. This raises several interesting questions. Is the substantial increase in earnings inequality in developed countries over the last decades an artifact of the baby boomers growing older? Can reported divergence in global income inequality be explained by increased differences in the age structure of rich and poor countries? Our age-adjusted inequality measure can be used to investigate these questions.

Appendix: Proof of Proposition 1

Proof. The triangle inequality theorem states that $|x-y| \geq |x| - |y|$, and the inequality holds if and only if one of the following conditions is satisfied:

(i) $x > 0$ and $y < 0$;
(ii) $x < 0$ and $y > 0$;
(iii) $x > y$ and $y < 0$;
(iv) $x < y$ and $y > 0$.

Let $x = (w_i - w_j)$ and $y = (\mu_i - \mu_j)$. It follows that $WG > PG$ if and only if one of the above conditions holds for at least one pair of individuals i and j.

Without loss of generality (because of the symmetry of the conditions), let the age groups be sorted by mean wealth such that $\mu_i \geq \mu_j$. Let $\min(w_i)$ denote minimum wealth in the age group of individual i, and let $\max(w_j)$ denote the maximum wealth in the age group of individual j.

Assume that $\mu_i > \mu_j$, implying $y > 0$.

No overlap in age-group distributions: assume that $\min(w_i) \geq \max(w_j)$. Then, $\min(w_i) - \mu_i < \max(w_j) - \mu_j$ whenever $w_i \neq \mu_i$ for at least one individual in the age group of individual i or $w_j \neq \mu_j$ for at least one individual in the age group of individual j. In that case, $x < y$ and condition (iv) holds.

Overlap in age-group distributions: assume that $\min(w_i) < \max(w_j)$. Then, $x > 0$ and condition (ii) holds.

Hence, $\mu_i \neq \mu_j$ for at least one pair of individuals and $w_i \neq \mu_i$ or $w_j \neq \mu_j$ for at least one of these individuals are sufficient conditions for $WG > PG$. □

References

Aaberge, R., Bjørklund, A., Jäntti, M., Palme, M., Pedersen, P., Smith, N., and Wennemo, T. (2002), Income Inequality and Income Mobility in the Scandinavian Countries Compared to the United States, *Review of Income and Wealth 48*, 443–469.

Almås, I. and Mogstad, M. (2010), Older or Wealthier? The Impact of Age Adjustment on Cross-Sectional Inequality Measures, NHH, Department of Economics working paper, no. 9.

Almås, I., Cappelen, A., Lind, J., Sørensen, E., and Tungodden, B. (forthcoming), Measuring Unfair Inequality. Evidence from Norwegian Data, NHH, *Journal of Public Economics*.

Almås, I., Havnes, T., and Mogstad, M. (2010), Baby Booming Inequality? Demographic Change and Inequality in Norway, NHH, Department of Economics working paper, no. 6.

Amiel, Y., Cowell, F., and Polovin, A. (1996), Inequality Among the Kibbutzim, *Economica 63*, 563–585.

Angrist, J. and Pischke, J. (2009), *Mostly Harmless Econometrics: An Empiricist's Companion*, Princeton University Press, NJ.

Atkinson, A. (1971), The Distribution of Wealth and the Individual Life Cycle, *Oxford Economic Papers 23*, 239–254.

Banks, J., Blundell, R., and Smith, J. (2003), Understanding Differences in Household Financial Wealth between the United States and Great Britain, *Journal of Human Resources 38*, 241–279.

Borsch-Supan, A. (2003), *Life Cycle Savings and Public Policy: A Cross-National Study of Six Countries*, Academic Press, San Diego.

Bossert, W. (1995), Redistribution Mechanisms Based on Individual Characteristics, *Mathematical Social Sciences 29*, 1–17.

Bossert, W. and Fleurbaey, M. (1996), Redistribution and Compensation, *Social Choice and Welfare 13*, 343–355.

Burkhauser, R. and Poupore, J. (1997), A Cross-National Comparison of Permanent Inequality in the United States and Germany, *Review of Economics and Statistics 79*, 10–17.

Burkhauser, R., Frick, J., and Schwarze, J. (1997), A Comparison of Alternative Measures of Economic Well-Being for Germany and the United States, *Review of Income and Wealth 43*, 153–171.

Cappelen, A. and Tungodden, B. (2007), Fairness and the Proportionality Principle, NHH, Department of Economics working paper, no. 31.

Danziger, S., Haveman, R., and Smolensky, E. (1977), The Measurement and Trend of Inequality: Comment, *American Economic Review 67*, 505–512.

Davies, J. and Shorrocks, A. (2000), The Distribution of Wealth, in A. Atkinson and F. Bourguignon (eds.), *Handbook of Income Distribution*, Elsevier, Amsterdam, 605–675.

Davies, J., Sandstrom, S., Shorrocks, A., and Wolff, E. (2006), The World Distribution of Wealth, Working Paper, WIDER.

Domeij, D. and Klein, P. (2002), Private Pensions: To What Extent Do They Account for Swedish Wealth Inequality? *Review of Economic Dynamics, Elsevier for the Society for Economic Dynamics 5*(3), 503–534.

Formby, J., Seaks, T., and Smith, J. (1989), On the Measurement and Trend of Inequality: A Reconsideration, *American Economic Review 79*, 256–264.

Foster, J. and Shorrocks, A. (1991), Subgroup Consistent Poverty Indices, *Econometrica 59*, 687–709.

Friesen, P. and Miller, D. (1983), Annual Inequality and Lifetime Inequality, *Quarterly Journal of Economics 98*, 139–155.

Gollier, C. (2001), Wealth Inequality and Asset Pricing, *Review of Economic Studies 68*, 181–203.

Harding, T., Solheim, H., and Benedictow, A. (2004), House Ownership and Taxes, Discussion Paper 395, Statistics Norway, Oslo.

Heckman, J. and Robb, R. (1985), Using Longitudinal Data to Estimate Age, Period, and Cohort Effects in Earnings Equation, in W. Mason and S. Fienberg (eds.), *Cohort Analysis in Social Research: Beyond the Identification Problem*, Springer, New York.

Hendricks, L. (2007), Retirement Wealth and Lifetime Earnings, *International Economic Review 48*, 421–456.

Jappelli, T. (1999), The Age–Wealth Profile and the Life Cycle Hypothesis: A Cohort Analysis with a Time Series of Cross-Sections of Italian Households, *Review of Income and Wealth 45*, 57–75.

Johnson, W. (1977), The Measurement and Trend of Inequality: Comment, *American Economic Review 67*, 502–504.

Kapteyn, A., Alessie, R., and Lusardi, A. (2005), Explaining the Wealth Holdings of Different Cohorts: Productivity Growth and Social Security, *European Economic Review 49*, 1361–1391.

Klevmarken, A. (2006), On Household Wealth Trends in Sweden over the 1990s, *Journal of Human Resources 38*, 322–353.

Klevmarken, A., Lupton, J., and Stafford, F. (2003), Wealth Dynamics in the 1980s and 1990s, in E. N. Wolff (ed.), *International Perspectives on Household Wealth*, Edward Elgar, UK.

Kolm, S. (1996), *Modern Theories of Justice*, MIT Press, Cambridge, MA.

Konow, J. (1996), A Positive Theory of Economic Fairness, *Journal of Economic Behavior and Organization 46*, 13–35.

Kurien, J. (1977), The Measurement and Trend of Inequality: Comment, *American Economic Review 67*, 517–519.

Lambert, P. and Aronson, J. (1993), Inequality Decomposition Analysis and the Gini Coefficient Revisited, *Economic Journal 103*, 1221–1227.

Lambert, P. and Decoster, A. (2005), The Gini Coefficient Reveals More, *Metron - International Journal of Statistics 63*, 373–400.

Manacorda, M. and Moretti, E. (2006), Why Do Most Italian Youths Live With Their Parents? Intergenerational Transfers and Household Structure, *Journal of the European Economic Association 4*, 800–829.

Masson, A. (1986), A Cohort Analysis of Wealth–Age Profiles Generated by a Simulation Model in France (1949–1975), *Economic Journal 96*, 173–190.
Meh, C. (2005), Entrepreneurship, Wealth Inequality, and Taxation, *Review of Economic Dynamics 8*, 688–719.
Minarik, J. (1977), The Measurement and Trend of Inequality: Comment, *American Economic Review 67*, 513–516.
Mirer, T. (1979), The Wealth–Age Relation Among the Aged, *American Economic Review 69*, 435–443.
Mookherjee, D. and Ray, D. (2002), Contractual Structure and Wealth Accumulation, *American Economic Review 92*, 818–849.
Mookherjee, D. and Shorrocks, A. (1982), A Decomposition Analysis of the Trend in UK Income Inequality, *Economic Journal 92*, 886–902.
Nelson, E. (1977), The Measurement and Trend of Inequality: Comment, *American Economic Review 67*, 497–501.
Paglin, M. (1975), The Measurement and Trend of Inequality: A Basic Revision, *American Economic Review 65*, 598–609.
Paglin, M. (1977), The Measurement of Inequality: Reply, *American Economic Review 67*, 520–531.
Paglin, M. (1979), The Measurement of Inequality: Reply, *American Economic Review 69*, 673–677.
Paglin, M. (1989), On the Measurement and Trend of Inequality: Reply, *American Economic Review 79*, 265–266.
Pudney, S. (1993), Income and Wealth Inequality and the Life Cycle: A Non-Parametric Analysis for China, *Journal of Applied Econometrics 8*, 249–276.
Roine, J. and Waldenströ m, D. (2009), Wealth Concentration over the Path of Development: Sweden, 1873–2006, *Scandinavian Journal of Economics 111*, 151–187.
Shorrocks, A. and Wan, G. (2005), Spatial Decomposition of Inequality, *Journal of Economic Geography 5*, 59–81.
Sierminska, E., Brandolini, A., and Smeeding, T. (2006), The Luxembourg Wealth Study – A Cross-Country Comparable Database for Household Wealth Research, *Journal of Economic Inequality 4*, 375–383.
WDI (2010), World Development Indicators, http://www.worldbank.org (February 18, 2010).
Wertz, K. (1979), The Measurement of Inequality: Comment, *American Economic Review 69*, 670–672.
Wolff, E. (1996), International Comparisons of Wealth Inequality, *Review of Income and Wealth 42*, 433–451.

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