

Education and Racial Inequality in Post Apartheid South Africa

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Abstract

The paper looks at the changing patterns of racial inequality in South African labour markets in the Post-Apartheid period. Using nationally representative panel data, it is demonstrated that while opportunities have been significantly equalized, as evidenced by an overall decline in the white-black wage differential, a new form of racial inequality has emerged, operating not directly on income as in the heyday of job reservation and school segregation, but indirectly, through inequality in the rewards to effort, as witnessed by sharply divergent patterns in the returns to education between the races. Differences in the returns to education now account for about 40% of the White-African wage differential, whereas a decade ago this effect was virtually zero. One consequence of this trend is an incentive structure likely to impede or possibly even reverse gains made in the equalization of schooling attainment.

JEL Keywords: Education, Race, Africa, South Africa

1 Introduction

One of the most far-reaching effects of Apartheid was the role it played in generating extreme economic inequality between race groups in South Africa. Not only does South Africa have among the highest levels of income inequality in the world, but this inequality is strongly racial in nature. As Figure 1 shows, the gap between white and black incomes just prior to the first democratic elections was substantial, with average real earnings of whites being more than five times that of blacks. Equally influential was the social engineering via race and language that occurred in the sphere of public education, with the

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introduction of the Bantu Education Act of 1954, which sought to prescribe differential access to education based on race.

The decade since the end of Apartheid, however, has brought about tremendous change in social and political life. South Africa is now heralded as a leading example of a modern liberal society, and many regard its constitutional guarantees as one of the most ambitious statements of individual liberties enshrined in law. Yet these social and political freedoms have not led to tangible change in the economic realm. As figure 2 shows, compared to 1993, when about 36% of the workforce earned less than two US dollars a day, the number of workers falling into this category now stands at 52%. Until recently, little could be said about such a process of change in South Africa, owing both to the lack of nationally representative data, and a coherent framework within which to assess such a change. However, the recent accessibility of large-scale household datasets, capturing information on population groups previously excluded from national statistics, has led to a rapid expansion in studies devoted to understanding the changing structure of labour market opportunities in the post-Apartheid period. Examples are Kingdon and Knight (2003) on the changing structure of unemployment, Moll (1996) on the collapse of the returns to education for Blacks, and most recently Hertz (2003) on measurement error in the returns to Black education.

This paper adds to this new literature by examining changes in the causal structure of racial inequality in South Africa. An analytical framework is derived with testable hypotheses concerning equal opportunity. Using this framework and very recent nationally representative panel data, it is demonstrated that while opportunities have been significantly equalized, as evidenced by an overall decline in the white-black wage differential, a new form of racial inequality has emerged, operating not directly on income as in the heyday of job reservation, influx control and school segregation, but indirectly, through inequality in the rewards to effort, as witnessed by sharply divergent patterns in the returns to education between the races. Differences in the returns to education now account for about 40% of the white-black wage differential, whereas a decade ago this effect was virtually zero. One consequence of this trend is an incentive structure likely to impede or possibly even reverse gains made in the equalization of schooling attainment.

The paper is organised as follows. In section 2, I develop a framework for analysing equality of opportunity in the attainment and reward of human capital. Section 3 presents the empirical approach, while section 4 provides background information on the data used and a discussion of average sample characteristics. The key empirical estimates are presented in section 5, along with a discussion of robustness of the estimates. Section 6 presents concluding remarks.

2 Defining Equality of Opportunity

Two views of equality of opportunity dominate public policy debates. Using the language of Roemer (2000), I refer to these as “non-discrimination” and

“levelling the playing field”. The former is simply the meritocratic principle of the absence of one’s ascriptive characteristics in determining one’s success or failure to acquire some desirable outcome – in this world, such things as “blood, color and sex”, to use the colourful words of Lester Frank Ward (1872), must be made irrelevant to success.

“Levelling the playing field” is a broader conception of both “what” to equalize and “how” to equalize, ranging from exceptionally broad based compensation such as equalization of resources in the Dworkinian sense (where resources are not limited to physical inputs that go into the production of the benefit, but extend also to include things that are beyond the control of any individual such as “talent”) to milder forms of compensation as proposed by Sen (1980)). Both views however, embody a “before” and “after” notion of an intervention to equalize opportunity, the key idea being that once the intervention has taken place, competition should be allowed to play a role. Thus more or less egalitarian views of equal opportunity can be reduced to differences in where to place the gate defining “before” and “after” – the debate often centering on what can reasonably be expected to be outside the scope of one’s control, and what individuals should be held accountable for.¹ Both the liberal view of equal opportunity, as embodied by the meritocratic principle of non-discrimination, as well as modern theories of distributive justice, recognize the importance of rendering the ascriptive aspects of one’s personal characteristics inconsequential through public policy or legislation. Thus it seems uncontroversial that the absence of race as a determining factor in one’s economic success or failure should feature prominently in a definition of a society in which opportunities are equal (Arrow *et al* (2000), Benabou (2000), Bourguignon *et al* (2002), Bowles (1973), Roemer (1996), Dworkin (1981), Sen (1980)).

To derive a framework for analysing equality of opportunity, I begin with the following canonical earnings function² usually attributed to Mincer (1974):

$$\ln y = \ln \alpha + \beta s \tag{1}$$

The constant term accounts for expected earnings in the absence of other factors, where the other factors in this model, relate only to schooling. The schooling coefficient is the marginal private rate of return on education. Since earnings are also likely to be independently influenced by experience, equation 1 is conventionally augmented with a measure of “potential experience” to account for the importance of on-the-job learning. But because this proxy is measured with considerable error in cases where grade repetition is high and spells of unemployment are long (both of which are true for South Africa) convention is not

¹See Roemer (1996) for a survey of the differences in approach taken by Arneson (1989), Dworkin (1981) and Cohen (1989) particularly on the issue of defining “individual responsibility”.

²Note that individual sub-scripts, though applicable, have been suppressed to reduce notational clutter.

followed here. Instead, age is used as opposed to the more standard proxy for “potential experience”. Letting A refer to an individual’s age, we can then write;

$$\ln y = \ln \alpha + \beta s + \psi_1 Age + \psi_2 Age^2 + u \quad (2)$$

where u is an error term, which is assumed to be i.i.d. Equations 1-2 can be derived from a solution to an optimisation problem - i.e., a first order condition that defines the optimal level of schooling that maximises utility. However, I treat equation 2 as a reduced form representation of an income-generating function. This is to account for the fact that in the presence of credit constraints and other institutional impediments that restrict one’s choice set (for example, the set of policies largely grouped under the heading of “Bantu Education”), the structure of choice will be inherently asymmetric across races. The standard microfoundation of the Mincerian Earnings function (see for example Willis (1986)) will be of little value in this context, as it explicitly excludes the possibility that preferences over schooling might be endogenously determined (through constraints imposed by the institutions governing such choices and the like).³

Taking variances of the basic earnings equation gives a measure of the distribution of earnings as a linear function of the distribution of schooling. This is a proxy for earnings inequality attributable to schooling inequality. In a simple two-variable model, the coefficient of determination - the R^2 - measures the fraction of the variation in log earnings explained by education, adjusted for age. Likewise, in a more general model, such as equation 2 the R^2 measure is a useful test of the fraction of earnings inequality explained by the particular combination of explanatory variables in question. However, when earnings is measured with error, unless the mean value of the error (transitory) component is zero, estimates of the intercept terms will be biased, and the R^2 measure is no longer a reliable estimate of so-called “explained inequality”. However, as is well known, if the measurement error is “classical”, then the estimated coefficients remain unbiased, even without correction. For this reason, and because of the likely effect of censoring, which I discuss below, I choose to focus not on the R^2 measure as an indication of the degree of inequality, but on the coefficients themselves.

In the tradition of wage-differential decompositions (Oaxaca (1973), Blinder (1973), Newmark (1988) and Oaxaca and Ransom (1994)), I seek to examine the contribution of personal vs. non-personal characteristics in generating wage differences between groups. In the analysis to follow, I restrict my attention to the Mincerian representation of the earnings function, and therefore consider only the effects of education in generating racial wage differentials.⁴

Now let there be two race groups, Whites and Africans, denoted by w and

³See Card (2001) for one recent attempt at resolving this problem.

⁴I show below that altering the above earnings function to control for other variables which can reasonably be treated as exogenous such as gender, age-education interactions, and within household fixed effects, do not alter the main findings. However, for ease of exposition, I begin with a standard Mincer specification.

a respectively.⁵ Let \bar{A}_j refer to the un-weighted average age for of the two races in the j th year. By making use of the estimated coefficients from equation 2 estimated separately for each group, the total expected earnings differential between Whites and Africans for the j th year can then be represented as:

$$\begin{aligned}
\Delta_j &= \bar{y}_{wj} - \bar{y}_{aj} \\
&= E(y_{wj}|\bar{s}_{wj}, \bar{A}_{wj}, \bar{A}_{wj}^2) - E(y_{aj}|\bar{s}_{aj}, \bar{A}_{aj}, \bar{A}_{aj}^2) \\
&= (\alpha_{wj} - \alpha_{aj}) + (\delta_{wj}\bar{A}_{wj} - \delta_{aj}\bar{A}_{aj}) + (\gamma_{wj}\bar{A}_{wj}^2 - \gamma_{aj}\bar{A}_{aj}^2) \\
&+ (\beta_{wj}\bar{s}_{wj} - \beta_{aj}\bar{s}_{aj})
\end{aligned} \tag{3}$$

This expression can then be decomposed into a “pure-race” effect and a “schooling” effect, where the latter can be further decomposed into a “years” effect, and a “returns” effect. In the Mincerian framework, the pure race effect is sometimes defined as the wage differential that would prevail in the absence of schooling and could be measured as the difference in the intercepts of the two earnings functions (see for example Lam (2001)). However, this is meaningless if either group has a non-zero minimum level of attainment. Since a pure race effect can in principle be calculated for any level of education, one solution is to define the appropriate benchmark at some level of schooling that is at least equal to the larger of the two group-level minima in schooling attainment. I refer to this benchmark as \bar{s}_{a1} . Denoting $\tilde{\Delta}_j$ as the pure race effect (with the same age conditioning as in equation 3), we can then write:

$$\begin{aligned}
\tilde{\Delta}_j &= \bar{y}_{wj} - \bar{y}_{aj} \\
&= E(y_{wj}|\bar{s}_{a1}, \bar{A}_{wj}, \bar{A}_{wj}^2) - E(y_{aj}|\bar{s}_{a1}, \bar{A}_{aj}, \bar{A}_{aj}^2) \\
&= (\alpha_{wj} - \alpha_{aj}) + (\delta_{wj}\bar{A}_{wj} - \delta_{aj}\bar{A}_{aj}) + (\gamma_{wj}\bar{A}_{wj}^2 - \gamma_{aj}\bar{A}_{aj}^2) \\
&+ (\beta_{wj}\bar{s}_{a1} - \beta_{aj}\bar{s}_{a1})
\end{aligned} \tag{4}$$

Intuitively, this expression defines the wage differential that would result, if Whites had the same average schooling attainment as Africans. If the estimation technique is linear in the parameters, equation 4 is always contained in equation 3. Thus the total wage differential given by equation 3 can be expressed as the sum of the pure race effect given by equation 4 and a residual term Ω_j , so that:

$$\begin{aligned}
\Omega_j &= \Delta_j - \tilde{\Delta}_j \\
&= \beta_{wj}\bar{s}_{wj} - \beta_{aj}\bar{s}_{aj} - (\beta_{wj}\bar{s}_{a1} - \beta_{aj}\bar{s}_{a1}) \\
&= \beta_{wj}(\bar{s}_{wj} - \bar{s}_{a1}) - \beta_{aj}(\bar{s}_{aj} - \bar{s}_{a1})
\end{aligned} \tag{5}$$

⁵During Apartheid, there were four racial classifications for the South African population: White, African, Asian, and Coloured. This classification system was the basis for segregating the population in terms of residential areas, schools, and basic economic and political rights. When the term “black” is used in this and other work on South Africa, it is generally meant to refer to the African, Asian, and Coloured populations collectively. In what follows, I will confine my analysis to differences between the “African” and “White” populations. If a statement applies to all non-white populations, the term “Black” will be used. More precisely, “African” refers to individuals who are native to Africa (excluding the Asian and so-called “Coloured” populations), whereas “White” refers to individuals of European decent comprising individuals primarily of British, Dutch, French, Portuguese, and German origin.

We now ask how much of Ω_j can be attributed to differences in means as opposed to differences in rates of return between the two groups? Let the average return to schooling across groups be given as $\bar{\beta}_j = \frac{\beta_{wj} + \beta_{aj}}{2}$. Moreover let the first term of equation 5 equal ω_j and the second term equal τ_j so that

$$\begin{aligned}\omega_j &= \bar{\beta}_j(\bar{s}_{wj} - \bar{s}_{a1}) + (\beta_{wj} - \bar{\beta}_j)(\bar{s}_{wj} - \bar{s}_{a1}) \\ &= \bar{\beta}_j(\bar{s}_{wj} - \bar{s}_{a1}) + (\beta_{wj} - \beta_{aj}) \left(\frac{\bar{s}_{wj} - \bar{s}_{a1}}{2} \right)\end{aligned}\quad (6)$$

$$\begin{aligned}\tau_j &= \bar{\beta}_j(\bar{s}_{aj} - \bar{s}_{a1}) + (\beta_{aj} - \bar{\beta}_j)(\bar{s}_{aj} - \bar{s}_{a1}) \\ &= \bar{\beta}_j(\bar{s}_{aj} - \bar{s}_{a1}) + (\beta_{aj} - \beta_{wj}) \left(\frac{\bar{s}_{aj} - \bar{s}_{a1}}{2} \right)\end{aligned}\quad (7)$$

Substituting equations 6 and 7 into equation 5 yields:

$$\begin{aligned}\Omega_j &= \Delta_j - \tilde{\Delta}_j \\ &= \underbrace{\bar{\beta}_j(\bar{s}_{wj} - \bar{s}_{aj})}_{\text{years}} + \underbrace{(\beta_{wj} - \beta_{aj}) \left(\frac{\bar{s}_{wj} - 2\bar{s}_{a1} + \bar{s}_{aj}}{2} \right)}_{\text{returns}}\end{aligned}\quad (8)$$

Given the above, we can now define precisely what is meant by equality of opportunity. As noted earlier, both the “non-discrimination” and “levelling the playing field” views of equal opportunity take as given that a meritocratic society is one in which race should not be allowed to play a role. Moreover, both views also accept that in a meritocratic society, effort should be rewarded. Taking these two requirements as guiding principles, the following set of conditions can be specified:

Hypothesis 1 *Whites and Africans should have the same expected incomes, contingent on a given level of schooling.*

Hypothesis 2 *Investment in education should reap a positive rate of return.*

Hypothesis 3 *There should be no income differential generated by race - i.e., there should be no “pure race effect”.*

Hypothesis 4 *There should be no income differential generated by differences in mean schooling - i.e., there should be no “years” effect.*

Hypothesis 5 *There should be no income differential generated by differences in rates of return - i.e., there should be no “returns” effect.*

In the above setup, the rejection of conditions 3-5, are necessary and sufficient (in the context of the Mincerian framework) for the rejection of condition

1. Thus, taken together the “race”, “years” and “returns” effects can be thought of as comprising the causal structure of the racial wage differential. Condition 4 might be seen as potentially controversial in a de-racialised schooling system. However, to then extent that present differences in attainment reflect past discrimination, a levelling of the playing field between the races requires special attention be paid to equalising schooling attainment. In this respect, income differentials that arise because of limited progress on this front is no less a form of unequal opportunity than that which is introduced by more direct forms of discrimination.

3 Estimation

A key challenge to estimating standard earnings functions based on household surveys is finding an appropriate method for dealing with zero-earners. This problem is especially acute in the case of South Africa since the rate of unemployment is known to be catastrophically high. Because of this, most now recognise the importance of including the zero earners. However there is little consensus in the econometric literature on precisely how to deal with the zeros once included. While the Tobit estimator has become something of a workhorse in dealing with this problem (see discussion in Green (2003) for example), others such as Deaton (1997) advocate OLS simply on the basis that zero earnings represent valid observations. As is well known, the argument for the latter is especially convincing in the presence of heteroskedastic disturbances (Johnston and Dinardo (1997)). Indeed, Monte Carlo experiments frequently show poor performance of the Tobit estimator relative to OLS under such conditions (Breen (1997)). For this reason, in what follows, I report both OLS and Tobit estimates. Given chronic unemployment in South Africa, arguably, it is instructive to consider the joint process of employment and earnings determination. The Tobit estimator is useful for this purpose as it lends itself to a useful decomposition of each of these effects. To see how this works, consider the following latent condensed form representation of the Mincer model

$$\begin{aligned} y^* &= \mathbf{x}'\beta + u \\ y &= 0 \quad \text{If } y^* \leq 0 \\ y &= y^* \quad \text{If } y^* > 0 \end{aligned} \tag{9}$$

Using theorem 22.4 of Greene (2003), the slope vector of the resulting conditional mean function for left censored data is:

$$\frac{\partial E(y|\mathbf{x})}{\partial \mathbf{x}} = \beta \times Prob(y > 0|\mathbf{x}) \tag{10}$$

An intuitively appealing decomposition of this marginal effect, first proposed by McDonald and Moffit (1980), allows an assessment of the relative earnings and employment effects of a given explanatory variable. The decomposition, obtained by applying the product rule when differentiating the conditional mean

function for left censored data (assuming normality of the errors), is as follows:

$$\begin{aligned} \frac{\partial E(y|\mathbf{x})}{\partial \mathbf{x}} &= pr(y > 0|\mathbf{x}) \frac{\partial E(y|y > 0, \mathbf{x})}{\partial \mathbf{x}} + E(y|\mathbf{x}) \frac{\partial pr(y > 0)}{\partial \mathbf{x}} \\ &= \underbrace{pr(y > 0|\mathbf{x}) \times \beta \left(1 - \sigma \frac{\phi(z)}{\Phi(z)} - \frac{\phi(z)^2}{\Phi(z)^2} \right)}_{\text{earnings}} + \underbrace{E(y|\mathbf{x}) \times \phi(z) \frac{\beta}{\sigma}}_{\text{employment}} \end{aligned} \quad (11)$$

where $\phi(z)$ and $\Phi(z)$ refer to the the probability density and cumulative distributions functions respectively, of z the normalised explanatory variable. The term labelled earnings is a useful way of approximating the rate of return to education when the Tobit estimator is used.

4 Data and Background

Two racially representative data sets covering the decade since 1993 are used in the analysis to follow. The data for 1993 are drawn from the Project for Statistics on Living Standards and Development (PSLSD), the first racially representative national survey of living standards to be conducted in South Africa. Undertaken at a time of great political and economic turmoil, this survey captured for the first time, the conditions under which the majority of South Africans lived during Apartheid. It therefore offers a unique set of information from which to assess changes that have taken place since the advent of democracy. The body of data generated by this study is unique also because it was the first study of its kind to capture outcomes for all population groups, in all parts of the country, on a wide range of issues, and is frequently used as a basis for comparing reforms undertaken since 1994.⁶ Examples are Case and Deaton (1998) on the economic consequences of the de-radicalisation of social pensions, Carter and May (2001) on poverty dynamics, Kingdon and Knight (2003) on unemployment, and Bertrand *et al* (2003) on labour supply.

The second and third sources of data are drawn from the Labour Force Surveys (LFS) of 2001/2002. The LFS is a bi-annual rotating household panel survey, which began in February of 2000 and is designed specifically to track labour market outcomes making it ideal for the topic at hand. Table 1 shows the mean sample characteristics for each of these surveys stratified according to race.

All estimates are for individuals and not households. Earnings are measured as gross monthly pay including overtime and bonuses. Income earners are restricted to full time and casual workers between the ages 15 - 65 for several reasons. First, while full time and casual employment are similarly defined

⁶All previous survey (including population censuses prior to 1991) excluded all homeland areas and the so called TBVC states of Transkei (now part of the Eastern Cape Province), Bophutatswana (now part of the North West Province), Venda (on the Zimbabwe border now part of the Northern Province) and Ciskei (now part of the Eastern Cape). This effectively meant that prior to 1994, the overwhelming mahority of the African population were never included in any of the sampling frames used by governmental statistical agencies.

across the surveys, “self-employment” is poorly defined in both the PSLSD and the LFS thus rendering any meaningful comparisons of this group impossible. Moreover, even in the absence of these problems, including these categories would compound the opposing effects of seasonality (where the measurement error is systematic and therefore predictable) and inter-temporal income fluctuations due to the transitory nature of many types of self-employment in the informal sector.

Years of education are derived from categorical data on level of education attainment. Table 1 indicates that although mean educational attainment rose over the intervening years for both races, Africans still average about four years less schooling attainment than Whites. The distribution of education also shows some interesting asymmetries between the two races, with whites exhibiting a much smaller standard deviation in years of schooling attainment than Africans.

5 Empirical Estimates

5.1 Rates of Return to Education

Table 2 summarises OLS and Tobit estimates of the coefficient on the schooling variable. Those estimates labelled “Ordinary least squares” and “Earnings Effect” are to be interpreted as private rates of return to education (the full set of regression results are reported in table 3). The results are quite striking. At the end of Apartheid the rate of return to education stood at approximately 11% for both races. Although marginally lower than what is typically reported for other parts of Sub-Saharan Africa at roughly the same time period – for example, Psacharopoulos (1994) reports a cross-country average Mincerian return of 13.7% – this estimate is extremely close to what is generally reported for South Africa at the end of the Apartheid period (Erichson and Wakeford (2001), Kingdon and Knight (1999), Hertz (2003)). A decade later however, the return to education for whites stood at a dramatic 43%, whilst that of Africans declined to about 7%. These estimates stand in stark contrast to the most recent estimates for the Sub-Saharan Africa region of 12% reported by Psacharopoulos and Patrinos (2002).

Figure 3 shows a graphical representation of the estimated return functions for the 18-25 year cohort. The picture shows that at the end of Apartheid, both races had virtually identical return functions, whereas a decade later the slopes of the two functions had diverged dramatically.

5.2 Robustness

One potential explanation for the unusually high returns to education for Whites concerns measurement error in schooling. By exploiting the panel structure of the LFS, an inter-temporal schooling correlation of 0.74 was found. By the classical-errors-in variables (CEV) assumption, this would lead to attenuation bias in the estimated rate of return, thus posing no problems for the conclusion

of divergence in the return structures facing the two races. However, there is no obvious reason why one should expect the CEV assumption to hold. Because data on schooling is usually recoded from a categorical structure to one that is discrete, owing to coding errors, the variance of observed schooling in years could in principle be larger than the variance of its unobserved component, thus violating a necessary condition for measurement error to be considered classical. Indeed, Hertz (2003) shows that the rate of return to black education drops to about 5-6% after controlling for mean-reverting measurement error in reported schooling. By averaging reported schooling over successive waves of the LFS, one is able to approximate the true rate of return assuming a non-classical measurement error structure. This method of controlling for measurement error did result in a reduction in the rate of return (to levels roughly on par with what Hertz found to be the case for Africans). However, the reduction for whites was much smaller, with the decline not exceeding 10 %. ⁷

Controlling for measurement error in incomes in a similar fashion (for those who reported incomes in both waves of the survey) however, generally increased rather than decreased the rate of return. Averaging both income and schooling (columns 4 and 7 of table 2) resulted in marginal reductions in the rate of return, but even in this instance, most estimates exceeded 30%.

Neither does censoring account for the result. Table 2 reports the Tobit index coefficient and corresponding marginal effect of the schooling variable (see table 3 panels B and C for the full set of regression results). As is clear from these results, the marginal effect of education was approximately 11% for whites and 9% for Africans at the end of Apartheid, with a similar dramatic rise evident ten years later. Since the Tobit marginal effect captures the combined effect of both earnings and employment, decomposing this coefficient is instructive for the purpose of evaluating the fraction of the total effect of education that can be interpreted as a rate of return *per se*. This is accomplished by applying equation 11 to the Tobit marginal effects, the results of which are summarised in table 2, and shown in more detail in panel C of table 3.

Given the large racial differences in employment, it is no surprise that most of the marginal effect of education for Whites is attributable to changes in earnings, conditional on being employed, whereas for Africans, most of the effect of education operates through marginal changes in the probability of being employed, given average wages. These two effects, separately labelled “Earnings Effect” and “Employment Effect” in table 2, sum to the overall Tobit marginal effect. The rate of return to education for whites estimated in this fashion (i.e., the part of the marginal effect labelled “Earnings Effect” in table 2) follows a similar pattern as that depicted by the OLS estimates. Although this method gives an estimate of the White rate of return that is somewhat lower than the OLS estimate, even after conditioning on the probability of employment, the

⁷In this section, not all results discussed in the text are reported fully (as is the case with the present result) owing to space considerations. Where this occurs, I clearly indicate as much. However, this mostly occurs where the result is either not central to the argument or where it can be excluded because it reappears in a different regression specification without loss of generality.

estimate is still in excess of 34%. Moreover, controlling both for measurement error (as described above) and censoring simultaneously, did not change this finding.⁸

Differences in the age composition between the African and White population could also introduce confounding effects. Recall that table 1 shows slight differences in average age between the two races. Controlling for these effects (results not shown), did little to change the magnitude of the difference in rates of return. Running separate regressions by age cohort revealed that while the youngest cohort of White workers considered (those aged 25-35) exhibited exceptionally high rates of return (in excess of 0.6), all other cohorts considered exhibited rates of return in excess of 0.30. By contrast, Africans exhibited extremely low rates of return, generally ranging between 6-13%.

Finally, the results appear not to be driven by assumptions about functional form. Figure 4 shows semi-parametric estimates of the return functions. The top panel shows a semi-parametric representation of the earnings-education relation, at the end of Apartheid, with the corresponding relation roughly a decade later. For each race group, two sets of curves are shown: the solid curves are for the simple Mincer equation (i.e., equation 2), whereas the underlying regression in the case of the dotted curves includes other regressors that can reasonably be assumed exogenous, such as gender, age-education interactions, and within-household fixed effects. Examining figure 4 tells a remarkably consistent story with that of figure 3, albeit somewhat more nuanced at the top end of the education distribution. We see the same sharp gradient emerging for Whites in the post-apartheid period, and a declining gradient for Africans, for all but those with tertiary qualifications. The emerging convexity of the black return structure evident in 1993 (previously documented by Moll (1996) and others) seems to have become even more pronounced over the decade that followed.

Collectively, these tests suggest that the sharp increase in the gradient of the White return function witnessed over the last decade are robust to potential confounds that could be introduced through measurement error, age-cohort effects, functional form, and model mis-specification.

5.3 The Effect of Omitted School Quality

Reconsider 1 again, but now assume that school quality q is not observed. Under this scenario, q is captured as part of the error term, so

$$y = \ln\alpha + \beta s + \gamma q + v \tag{12}$$

Now if q is correlated with s then u will also be correlated with s . This will tend to bias the estimated return to schooling, the extent of which is given by

$$\begin{aligned} \text{plim } \hat{\beta} &= \beta + \gamma\delta_s \\ &= \beta + \gamma \left(\frac{\text{Cov}(s, q)}{\text{Var}(s)} \right) \end{aligned} \tag{13}$$

⁸These results are not reported here, but can be made available on request.

such that if $\gamma > 0$ & $Cov(s, q) > 0$, then $\hat{\beta}$ is upwardly biased. The same result holds if $\gamma < 0$ & $Cov(s, q) < 0$. A standard solution to this problem is to find a valid proxy for q . However, since q itself is generally measured with error, we employ the multiple indicator method to control for this source of bias, as our data contain more than one measure of school quality. The source from which we derive our measures of school quality is the 1996 School Register of Needs. This is an administrative database maintained by the South African National Department of Education and contains a broad range of questions relating to school inputs. Our primary measure of school quality is the average pupil-teacher ratio of the magisterial district in which an individual resides: $q_j = \sum_i^n (Pupils_i/Teachers_i)/n$ where n refers to the number of schools in a magisterial district and j indexes the district. We then instrument this variable with residential district averages of the number of support staff employed by a school and the number of sheltered areas contained in the school. Our logic is that both of these variables have a direct influence on the pupil-teacher ratio, but should not influence earnings once the pupil-teacher ratio has been controlled for.

Table 4 summarises estimates of the return to schooling accounting for school quality in this manner. The key observation here is that school quality does not appear to have a marked impact on the racial differences in the return structure observed previously.

5.4 Testing for Equality of Opportunity

Given these dramatic changes in the manner in which education is rewarded in the labour market, what can be said about equality of opportunity? Applying the method outlined in equations 3-8, table 5 presents a summary breakdown of the changing structure of the total wage differential over the period 1993-2002 (see table 6 for a more detailed breakdown). At first glance, it appears that the decade since the end of Apartheid has brought about absolute gains in this respect. However, even though there has been an overall reduction in the total wage differential between Whites and Africans (evaluated at mean education of each group, holding age constant as shown in table 6), the structure of the differential has changed. Table 5 indicates that in 1993, almost 90% the White-African wage differential could be said to be driven by pure race effects, with the remainder accounted for by differences in mean schooling attainment between the races. The results of a decade later, however, show a dramatically different causal structure, with more than three quarters of the total differential accounted for by factors relating to schooling. More precisely, by 2002, only 24% of the overall wage differential operated through race directly, whereas about 36% of the difference was explained by differences in average educational attainment. The most striking finding, however, is the large increase in the fraction of the total differential accounted for by changes in the return functions. As table 5 shows, the so-called “returns effect” now explains about 40% of the total wage differential. Calculations based on the estimated tobit coefficients adjusted for measurement error on schooling and earnings from the LFS September 2001 and

February 2002 waves (columns 4 and 7 of table 2), showed that the contribution of the total wage differential explained by differences in the rate of return to education was 32%. Restricting the analysis to the earnings effect reduces this estimate to about 30%.

Finally, the same general pattern holds across both genders though there are large differences in magnitude. In 1993, pure race effects accounted for virtually the entire wage differential between White females and African females. By 2002, however, the race effect is small (about 10%) and negative. The entire wage differential is driven through differences related to schooling with differences in the rate of return now accounting for 58% of the overall wage differential.⁹ The same general pattern holds for the White-male-African-male wage differential (the returns effect increases from 7% in 1993 to 31% in 2002) though pure race effects remain high, at 42%.

Given these changes, what can one say about equal opportunity? Table 7 summarises the main findings concerning each of the five dimensions of equal opportunity outlined earlier. As noted, although gains have been made in the form of moderate reductions in the overall wage differential (including the between-race-within-gender comparison just discussed), on most counts equal opportunity does not exist in present day South Africa, and progress toward levelling the playing field has been slow.

6 Conclusion

While equality in the acquisition of education has seen modest improvements, the translation of educational opportunities into labour market gains has diverged for Africans and Whites over the last decade. Unlike the case during Apartheid, the direct effect of race on earnings is no longer as strong a factor in generating wage differentials between individuals. Rather, race now plays a strong role in determining how educational attainment comes to be valued in the labour market. The potential economic consequences of such a dramatic divergence in opportunities available to Whites and Africans are likely to be far reaching. One implication is that if standard models of human capital accumulation such as Mincer (1974), Becker (1993) and more recently Card (2001), are accurate in their description of the preferences, beliefs and constraints governing these individuals' decisions to acquire more education, then racial differences in the return functions (on the order of magnitude suggested by the evidence presented here) might lead to an incentive structure facing Blacks that is at odds with the further acquisition of schooling. This may impede or possibly even reverse gains made over the past decade in the equalization of schooling attainment. Models of statistical discrimination such as Arrow (1973) and Lundberg and Startz (1998) imply that such a response could lead to a self-fulfilling racial poverty trap in which employers continue to pay members of the disadvantaged group a wage equal to the average marginal product of the disadvantaged group,

⁹The relevant regression results (stratified by gender) that pertain to this decomposition are omitted from the paper for purposes of brevity, but these can be made available on request.

leading to persistence of the unequal reward structure, and hence a continuation of inequalities in educational attainment.

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Table 1: Mean Sample Characteristics

Variable	African	African	White	White
	1993	2002	1993	2002
Age (years)	32.82 (11.34)	34.35 (11.50)	34.67 (11.60)	37.86 (12.81)
Schooling (years)	6.78 (3.89)	8.18 (3.82)	11.19 (3.31)	12.00 (1.94)
Natural Log of Monthly wage (full time and casual)	2.86 (3.20)	3.02 (3.44)	7.15 (2.49)	6.26 (3.74)
Females (percent)	0.48 (0.50)	0.52 (0.49)	0.47 (0.50)	0.47 (0.49)
<i>n</i>	6929	31384	953	1939

Standard deviations are in parentheses. Data for 1993 are from the PSLSD. Data for 2002 are from the LFS. Years of Schooling are derived from categorical data on educational attainment. Monthly earnings are for individuals of working age (15-65) with full time and/or casual employment and includes zero earners classified according to the “broad” definition (i.e., including unemployed individuals not searching for work). See Kingdon and Knight (2000), Nattrass (2000), Wittenberg (2003) and Dinkelman and Pirouz (2002) for more on why the standard ILO definition is considered inappropriate in the context of South Africa

Table 2: Mincerian Returns to Education

Model	White	White	White	African	African	African
	1993	2002	M.Error	1993	2002	M.Error
Ordinary Least Squares	0.11 (4.6)	0.43 (10.3)	0.33 (4.95)	0.11 (10.9)	0.07 (14.1)	0.05 (4.99)
Tobit Index	0.11 (4.4)	0.56 (9.8)	0.33 (4.95)	0.17 (7.9)	0.10 (8.9)	0.05 (4.99)
Tobit Marginal Effect	0.11	0.49	0.33	0.09	0.05	0.04
Earnings Effect: $Pr(y > 0 \mathbf{x}) \times \frac{\partial E(y y>0,\mathbf{x})}{\partial s}$	0.11	0.34	0.32	0.03	0.02	0.02
Employment Effect: $E(y \mathbf{x}) \times \frac{\partial pr(y>0)}{\partial s}$	0.00	0.15	0.01	0.05	0.03	0.02
n	953	1939	531	6929	31384	14966

Table 2 summarises estimates of the coefficient on the schooling variable β from a simple mincer equation. Absolute values of t-ratios shown in parentheses. The dependant variable is the natural log of monthly earnings. OLS and Tobit estimates are presented for comparison. The full set of regression estimates are contained in table 3. The Tobit marginal effects are then decomposed into “earnings” and “employment” effects according to the method proposed by McDonald and Moffit (1980). Those estimates pertaining to earnings can be interpreted as a rate of return applicable only to employed individuals, conditioned on their probability of employment.

Table 3: Returns to Education: Detailed Regression Results

Variable	Whites 1993	Whites 2000	Whites 2002	Africans 1993	Africans 2000	Africans 2002
Panel A						
Ordinary Least Squares Estimates						
Constant	-0.307 (0.43)	-3.167 (5.35)	-4.520 (5.79)	-3.774 (11.88)	-5.755 (34.56)	-5.625 (34.64)
Age	0.335 (8.88)	0.240 (9.35)	0.259 (7.59)	0.272 (15.37)	0.368 (42.89)	0.354 (42.48)
Age ²	-0.004 (8.08)	-0.002 (7.88)	-0.003 (6.36)	-0.002 (10.64)	-0.003 (29.45)	-0.003 (29.31)
Education	0.105 (4.57)	0.418 (13.6)	0.430 (10.3)	0.107 (10.91)	0.106 (21.01)	0.072 (14.14)
R ²	0.11	0.14	0.10	0.10	0.16	0.15
Panel B						
Tobit Index Estimates						
Constant	-0.813 (1.03)	-4.663 (6.73)	-8.422 (7.84)	-13.893 (18.88)	-16.469 (48.24)	-20.063 (49.88)
Age	0.353 (8.49)	0.272 (9.15)	0.340 (7.33)	0.605 (15.45)	0.743 (44.28)	0.864 (44.37)
Age ²	-0.004 (7.72)	-0.003 (7.71)	-0.003 (6.16)	-0.006 (11.22)	-0.007 (32.53)	-0.008 (33.32)
Education	0.113 (4.45)	0.467 (13.15)	0.557 (9.79)	0.166 (7.86)	0.150 (16.10)	0.098 (8.90)
σ	2.581 (40.03)	3.143 (58.29)	4.631 (49.26)	5.824 (68.54)	5.428 (158.09)	6.149 (144.15)
$R^2_{ANNOVA}^\dagger$	0.120	0.171	0.099	0.032	0.034	0.078
Likelihood Ratio [‡]	100.950	426.600	256.670	710.740	673.320	6471.150
Panel C						
Decomposed Tobit Marginal Effects (Earnings and Employment)*						
Constant	-0.811 (1.03)	-4.608 (6.74)	-7.482 (7.90)	-7.163 (19.74)	-10.108 (50.57)	-10.005 (53.30)
Earnings	-0.790 (1.03)	-4.278 (6.76)	-5.191 (7.94)	-2.664 (19.61)	-4.358 (50.06)	-3.628 (52.31)
Employment	-0.021 (1.00)	-0.330 (5.60)	-2.291 (7.42)	-4.498 (19.47)	-5.750 (49.19)	-6.376 (52.43)
Age	0.352 (8.49)	0.269 (9.15)	0.302 (7.34)	0.312 (15.59)	0.456 (44.95)	0.431 (45.42)
Earnings	0.343 (8.49)	0.250 (9.15)	0.209 (7.29)	0.116 (15.31)	0.197 (43.63)	0.156 (43.93)
Employment	0.009 (4.54)	0.019 (7.13)	0.092 (7.12)	0.196 (15.60)	0.259 (44.70)	0.275 (45.40)
Age ²	-0.004 (7.73)	-0.003 (7.72)	-0.003 (6.17)	-0.003 (11.28)	-0.004 (32.82)	-0.004 (33.81)
Earnings	-0.004 (7.72)	-0.003 (7.71)	-0.002 (6.13)	-0.001 (11.17)	-0.002 (32.33)	-0.001 (33.22)
Employment	0.000 (4.41)	0.000 (6.38)	-0.001 (6.04)	-0.002 (11.28)	-0.002 (32.70)	-0.003 (33.78)
Education	0.113 (4.45)	0.462 (13.15)	0.495 (9.81)	0.085 (7.86)	0.092 (16.09)	0.049 (8.90)
Earnings	0.110 (4.45)	0.429 (13.12)	0.343 (9.68)	0.032 (7.80)	0.040 (15.99)	0.018 (8.87)
Employment	0.003 (3.42)	0.033 (8.61)	0.152 (9.32)	0.054 (7.87)	0.052 (16.11)	0.031 (8.90)
<i>n</i>	953	2171	1939	6929	30697	31384

The dependant variable is the natural log of monthly earnings in South African Rand. All explanatory variables are measured in years. Absolute values of t-ratios are in parentheses.

[†] $R^2(ANNOVA)$ is a goodness of fit measure that roughly mimics its OLS counterpart, converging on this measure as the censoring probability goes to zero. It is computed by taking the variance of the predicted mean divided by the variance of the dependant variable (see Veall and Zimmerman (1992) and Greene (2003)).

[‡] The more common goodness of fit measure - the Likelihood ratio - is also reported. It tests the joint hypothesis that all the slope coefficients equal zero and follows a χ^2 distribution with k degrees of freedom. In all cases reported above, the null hypothesis is rejected at the 1% level.

* Panel C of the table shows a decomposition of the marginal effects into two additive components: the independent effects of each explanatory variable on both earnings and employment. The decomposition, first suggested by McDonald and Moffit (1980), is achieved by applying the product rule when differentiating the conditional mean function. Thus: $\partial E(y)/\partial x_k = \Phi(z) \times \beta_k [1 - \sigma(\phi(z)/\Phi(z)) - (\phi(z)/\Phi(z))^2] + [-x' \beta + \sigma(\phi(z)/\Phi(z))] \phi(z) (\beta_k/\sigma)$. See equation 11 for more details.

Table 4: The Effect of School Quality on the Return Structure

Model	White	White	African	African
	OLS	2SLS	OLS	2SLS
<i>Schooling</i>	0.449 (8.71)	0.422 (6.88)	0.071 (6.21)	0.067 (4.48)
<i>q</i>	0.019 (0.59)	-0.053 (0.51)	-0.052 (4.37)	-0.195 (3.35)
<i>n</i>	1905	1256	29065	24678

Absolute values of t-ratios shown in parentheses and are based on standard errors which are robust to district level clustering. 95% confidence intervals for β_w and β_a do not overlap in both OLS and 2SLS specifications.

Table 5: Decomposed White-African Wage Differential

Differential	Pure Race	Years Effect	Returns Effect
	Δ_j	μ_j	ρ_j
1993	0.89	0.11	0.00
2000	0.34	0.34	0.32
2002	0.24	0.36	0.40

The table shows the % breakdown of the total wage differential between Whites and Africans, evaluated at mean educational attainment of each race group, for each time period, holding age constant at the mean of the two means (rounded to the nearest whole number). The “pure race effect” is calculated according to equation 4. The “years effect” and “returns effect” are calculated according to equation 8. The three effects sum to the total differential, given by equation 3.

Table 6: Detailed White-African Wage Differential Decomposition

Component	1993	%	2000	%	2002	%
White Mean Earnings ($E(y_{wj} \bar{s}_{wj}, \bar{A}_j, \bar{A}_j^2)$)	7.648		7.526		6.722	
African Mean Earnings ($E(y_{aj} \bar{s}_{aj}, \bar{A}_j, \bar{A}_j^2)$)	3.309		4.298		3.764	
Total Differential (Δ_j)	4.338		3.228		2.958	
Pure Race Effect ($\tilde{\Delta}_j$)	3.873	(0.89)	1.087	(0.34)	0.711	(0.24)
Schooling Effect ($\Omega_j = \Delta_j - \tilde{\Delta}_j$)	0.465		2.142		2.247	
Years Effect (μ_j)	0.469	(0.11)	1.113	(0.34)	1.066	(0.36)
Returns Effect (ρ_j)	-0.004	(0.00)	1.028	(0.32)	1.181	(0.40)

Table 7: Hypothesis Tests Concerning Equality of Opportunity

Hypothesis	1993	2000	2002	Progress
Hypothesis 1: $\Delta_j = 0$	No	No	No	Yes
Hypothesis 2: $\beta_{kj} > 0$	Yes	Yes	Yes	Yes
Hypothesis 3: $\tilde{\Delta}_j = 0$	No	No	No	Yes
Hypothesis 4: $\mu_j = 0$	No	No	No	No
Hypothesis 5: $\rho_j = 0$	No	No	No	No

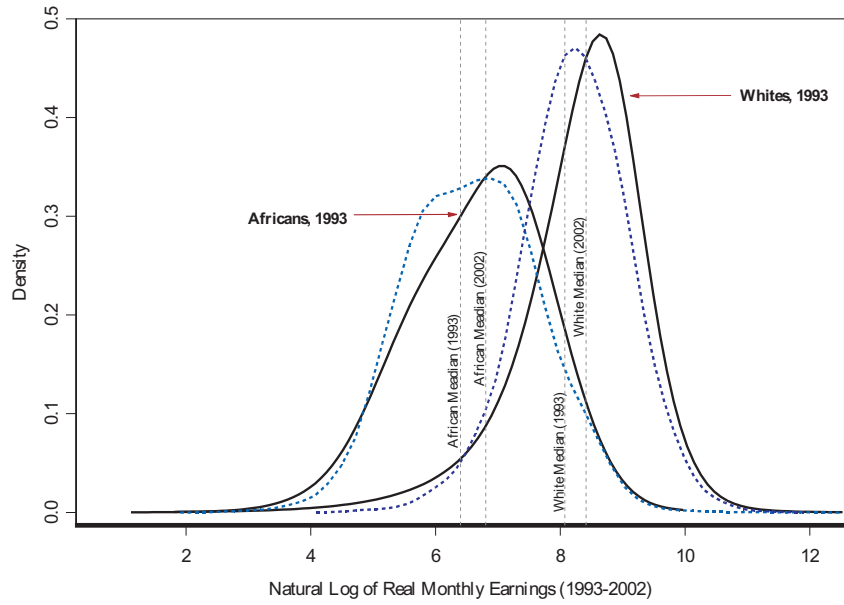


Figure 1: The Racial Distribution of Earnings (1993 & 2000).

The picture shows the evolution of the racial distribution of earnings of full time and casual workers. Data for 1993 are taken from the World Bank living standards measurement survey (also known as the PSLSD). Data for 2002 are drawn from the Labour Force Survey (LFS). Dotted lines represent the 2002 distribution. The income variable is measured in natural logs of monthly earnings in South African Rand. All data are deflated by the average CPI for the months over which the survey work was conducted in the relevant years, with 2000 as the base year.

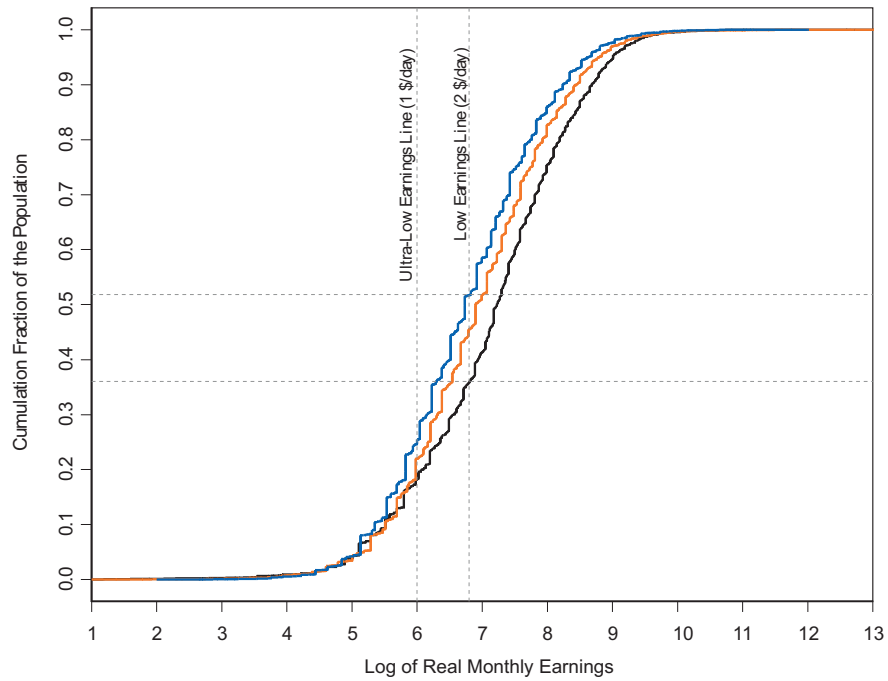


Figure 2: The Evolving Distribution of Real Earnings, 1993 - 2002.

The picture shows the evolution of the distribution of real earnings of full time and casual workers, from right to left, for the years 1993, 2000 and 2002. The “low earnings line” refers to the wage required if a worker’s family were to get to the household poverty line, given the average number of employed and unemployed workers in a household (see Chicello, Fields and Leibbrandt (2001)). The “ultra-low earnings” line is a per capita adult equivalent wage required to take a worker up to the equivalent of the \$1/day poverty line. The Rand amounts used in calculating these reference lines are deflated to account for the new CPI benchmark of 2000.

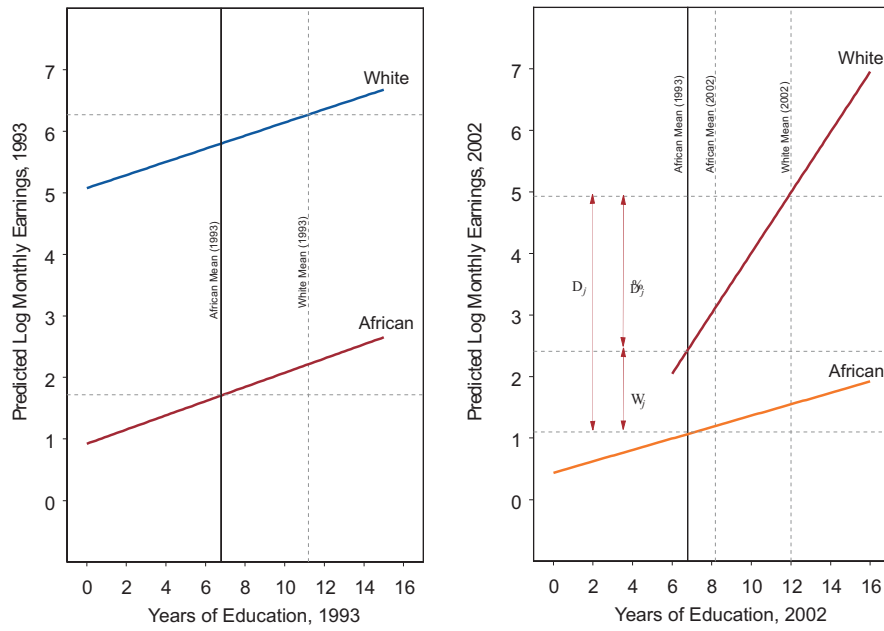


Figure 3: Education and Racial Inequality, 1993 - 2002.

Panel (a) shows predicted earnings of both races in 1993 and 2002. The figures are derived by making use of the OLS predicted values based on the estimates reported in tables 2 and 3, for the 18-25 year cohort only. The vertical reference lines indicate the mean level of attainment for each race group.

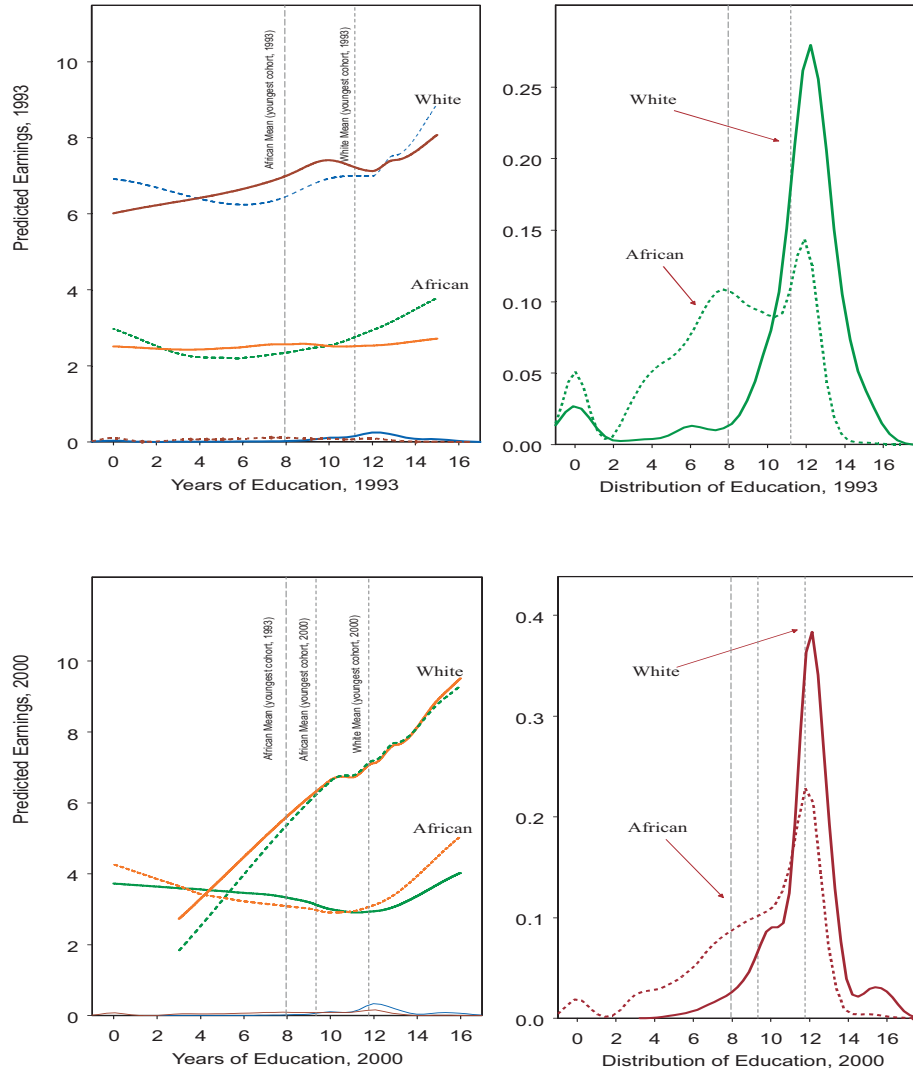


Figure 4: Education and Racial Inequality, 1993 - 2002.

The figures on the left (top and bottom) show non-parametric estimates of the return to education. The figures on the right show the evolution of the distribution of education. The vertical reference lines indicate the mean level of attainment for all 18-25 year old individuals. The solid curves in the figures on the left are non-parametric estimates of the Mincer equation. The dotted curves include other regressors which can reasonably be treated as exogenous such as gender, location, age-education interactions, a remoteness indicator, and household fixed effects. Other regressors such as union membership and occupation are excluded on the grounds that they are endogenous (i.e., they are determined in part by schooling). The fitted values are for 13627 working age individuals in regular and casual wage employment.