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The Persistence of Poverty in Britain : Evidence from Patterns of Intergenerational Mobility

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Abstract

In this paper we provide estimates of poverty persistence in the U.K by examining the degree of intergenerational mobility. Our findings support the view that poverty is a culture which is transmitted across generations. Furthermore we find that the correlation between fathers' and sons' earnings remains significantly positive after controlling for inherited endowments. This suggests that government programs aimed at providing equal opportunities, such as equal access to education, have the potential to substantially reduce inequality.

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1. Introduction

The transmission of poverty across generations and the channels through which such transfers take place should play an important role in any attempt to reduce inequality. Early work on U.S data concluded that individuals lived in open societies where practically all advantages or disadvantages of ancestors were eliminated after only three generations (Becker and Tomes 1986). Solon (1992) and Zimmerman (1992) challenge this view, suggesting that the earlier estimates were biased downwards as a result of measurement error and homogenous samples. Using nationally representative samples and correcting for measurement error, these studies report a correlation between fathers' and sons' earnings of approximately .45, implying substantially less mobility than the earlier results.

Less is known about the persistence of poverty across generations in the U.K. In the late 1970's Atkinson, Maynard, and Trinder, henceforth AMT, conducted a survey to obtain information on earnings and income for families originally surveyed by Rowntree in York in 1950. Using information on approximately 300 father-son pairs they found that the correlation between fathers' and sons' earnings was .44 (AMT (1983)). However as noted by AMT the original Rowntree sample was restricted to 'working-class households', defined as households with total earnings of less than £550 [2]. Furthermore the sample was restricted to the city of York, which the authors acknowledge may not have been typical of the average U.K city at that time. As a result the extent to which the AMT findings generalise to the population as a whole is uncertain.[3]

[2] AMT estimate that this reduced the population in the Rowntree study by approximately 40%.

[3] The restriction to working-class households may be particularly problematic in that an implication of the intergenerational transfer model which we describe in the next section is that regression to the mean in earnings will be faster among richer families. As a result the large correlation coefficient obtained by AMT may simply reflect the nature of their sample.

In this paper we extend the work of AMT by using the National Child Development Survey (NCDS) to examine the extent of intergenerational mobility in the U.K. The NCDS is a nationally representative sample which contains information on individuals' income and family background as well as detailed information on work histories. We use the NCDS to examine several dimensions of intergenerational mobility including earnings mobility and the relationship between fathers' and sons' unemployment histories. We also use information on ability tests administered to the children at the ages of 7, 11 and 16 in an attempt to control for inherited endowments. We find that a 1% increase in fathers' earnings increases sons' earnings by .6% after correcting for measurement error. This highlights a tendency for poverty to be passed from generation to generation. For example, a coefficient of .6 on the log of fathers' earnings implies that the earnings of a son of a father at the top decile of the earnings distribution will be twice that of the son of a father at the bottom decile.[4] The relative importance of this can be gauged by comparing it to recent estimates of wage gains associated with obtaining a college education. For instance Murphy and Welch estimated that in 1986 individuals in the U.S with a college degree earned about 1.7 times that of individuals with only a high-school degree (Murphy and Welch 1989).

The persistence of poverty is not restricted to earnings. We find that sons of fathers who have been unemployed are also more likely to be unemployed. Finally, we find that a substantial correlation remains after controlling for inherited endowments. In the context of the model specified in the paper this suggests that inequality in opportunity may have an important role to play in transmitting poverty across generations and that programs aimed at providing equal access to health and education facilities can substantially reduce inequality.

2. Theory

[4] This calculation is based on a ratio of top to bottom decile of 3.16 in 1991 (taken from the New Earnings Survey of weekly earnings for full-time males whose pay was not affected by absences).

In this section we outline a model of intergenerational transfers which we use to interpret the empirical results presented later in the paper. The model developed originally by Becker and Tomes (1986) is based on an altruistic family that is unable to borrow on capital markets to finance investments in its childrens' education.[5] The model consists of two time periods, childhood and adulthood. Human capital accumulation (H_t) occurs in the first period and is transferred into earnings and consumption in the second period according to the following equation :

$$Y_t^* = H_t + e_t \tag{1}$$

where the subscript t refers to the generation to which the individual belongs and e_t is a measure of market luck. H_t is in turn determined by endowments (E_t) and parental expenditure on the child's development (τ_{t-1}). It is assumed that endowments are transmitted across generations via a linear Markov model, with parameter h.

Since human capital is poor collateral for a loan it is assumed that parents must finance investments in children either by selling assets or reducing their own consumption. Furthermore, it is assumed that parents are altruistic in that children's consumption/income enters directly into the parents' utility function. We denote the degree of altruism by δ. Solving for optimal transfers (τ^o_{t-1}) we find that parental expenditure on children depends directly on market luck, the child's endowments, parental income, the return on parental expenditure, individual endowments and parental generosity. We can see this more clearly by considering a specific form for the utility function such as a log-linear specification.[6] The problem facing the parent

[5] For a model of intergenerational transfers without altruism see Cigno (1993). In his model transfers from middle-aged individuals to the young generation are a means of providing for the middle-age generation upon retiring.

[6] In this paper we only consider solutions where $\tau_{t-1} >= 0$. This rules out cases where children make transfers to parents. We also ignore the fertility decisions of parents. For a model examining the implication of fertility on intergenerational mobility see Becker and Barro (1988).

can be written as :

$$\underset{\tau_{t-1}\, C_{t-1}}{\text{Max}} \quad \delta \ln(C_{t-1}) + (1-\delta) \ln(Y^*_t) \tag{2}$$
$$\text{s.t.} \quad Y^*_t = H_t + e_t$$
$$H_t = \alpha_o \tau_{t-1} + \alpha_1 E_t$$
$$C_{t-1} + \tau_{t-1} = Y^*_{t-1}$$

Solving this maximisation problem results in an optimal level of τ_{t-1} equal to:

$$\tau^o_{t-1} = (1-\delta)\, Y^*_{t-1} - \frac{\alpha_1}{\alpha_o}\, \delta\, E_t - \frac{\delta}{\alpha_o}\, e_t \tag{3}$$

The effect of parental income on transfers is positive but less than one provided that consumption is a normal good ($\delta > 0$). This implies regression to the mean in transfers in that the children of rich parents receive proportionately less from their parents than the children of poor parents.

Substituting (3) into the earnings equation (1) we get the following expression for sons' earnings :

$$Y^*_t = \rho\, Y^*_{t-1} + \theta\, E_t + e_t + v_t \tag{4}$$

where $\rho = \alpha_o(1-\delta)$; $\theta=\alpha_1 (1-\delta)$ and $\upsilon_t= -\delta e_t$.

Equation (4) forms the basis for the empirical work presented in section 4. The parameter ρ measures parents propensity to invest in a child's human capital resulting from credit constraints in the capital market. However childrens' earnings also depend indirectly on parental earnings through the transmission of endowments. In developing policy proposals it is important to distinguish between these two channels.

3. Data

The data used in this paper are taken from the NCDS, a longitudinal data set following the lives of all those living in Great Britain who were born between the 3rd and 9th of March 1958. To date there have been 5 follow-up surveys of the individuals. These took place in 1965 (when they were aged 7), 1969 (aged 11), 1974 (aged 16), 1981 (aged 23) and 1991 (aged 33).

The information in sweeps 1-3 was provided mostly by the parents of the children and contains data on parents' education, social class, earnings, income and work history. The sweep 5 survey contains detailed self-reported information on the child's labour market status until the age of 33, including wages and past work history. The NCDS also includes several measures of the child's ability, including the scores from reading and math tests administered when the child was aged 7. In this paper we use these measures to proxy for the endowments (E_t) of the child.

In this paper we examine only father-son pairs in order to focus on individuals with a strong attachment to the labour force. We recognise that the female labour supply decision plays an integral part in the poverty process. An analysis of this issue requires modelling the joint female participation decisions of mothers and daugthers, which we aim to examine in future research. In estimating earnings mobility we restrict our attention to father-son pairs in which the father was working in 1974 and the son was working in 1991. We relax this restriction when we look at unemployment histories. When analysing the earnings data we further restrict the sample to fathers and sons who earned more than £40 a week and less than £1200 a week (measured in 1991 prices)[7]. After imposing these restrictions we had 1330 father-son pairs with valid data.[8,9] A description of the variables used in this study along with summary statistics used are given in the data appendix.

[7] The average weekly earnings for all males in 1991 was £318.

[8] In total 11,400 individuals were interviewed at sweep 5. Of these 7906 reported information on a regular father figure, of which 6385 reported valid earnings data for their father. Of this sample 3021 were males with valid own wage data. The further reduction of our sample to the 1330 used in the paper reflects the requirements that the individuals report information both on the ability tests and the variables used as instruments for father's earnings.

[9] Of this sample 96% of the pairings represent natural father-son pairings. We have also carried out the analysis excluding the 4% for which this was not the case, however this had little impact on the results.

4. Results

a. Intergenerational Earnings Mobility

In this section we concentrate on estimating equation (4) of section 2 :

$$Y_t^* = \rho\, Y_{t-1}^* + \theta\, E_t + w_t \qquad (5)$$

Our primary interest in this paper is in estimating ρ, the extent to which differences in father's earnings *directly cause* differences in earnings among children. Two potential sources of bias are likely to occur when one attempts to estimate ρ. The first is an upward bias in ρ as a result of omitting information on endowments. We return to this issue later in the paper.

The second source of bias results from the fact that rather than observing the lifetime earnings of fathers, Y^*_{t-1}, we observe a point in time measure of earnings Y_{t-1}. The relationship between the two may be expressed as :

$$Y_{t-1} = Y_{t-1}^* + \varepsilon_{t-1}$$

where ε_{t-1} is a random error term. It is well known that under these circumstances the OLS estimate of the correlation coefficient underestimates the true parameter. To correct for this bias we estimate equation (5) using instrumental variables.

The results of our estimation are given in table 1[10]. The first column gives the standard OLS estimate of ρ when measures of endowments are omitted from the model. The estimated coefficient is significant and equal to .29. This is similar to the OLS estimate of .295 obtained by Zimmerman and the estimate of .2 reported by Behrman and Taubman (1985). This estimate implies significant regression to the mean in earnings and substantial mobility. Using the ratio of top to bottom decile earnings of 3.16 taken from the 1991 New Earnings Survey our estimate of .29 implies that ancestral advantages will be almost entirely eliminated after 4 generations.

[10] In the results presented here we do not adjust fathers earnings for life-cycle differences. Jenkins (1987) discusses life-cycle biases in studies of intergenerational mobility. In particular he shows that the direction of the bias is unclear and argues that in certain cases traditional approaches to the correcting for the bias may be unsatisfactory. To examine the robustness of our findings, we also carried out the analysis using a life-cycle adjusted measure of fathers' earnings and obtained results very similar to those presented in the paper.

To examine the impact of measurement error in fathers' earnings we use the instrumental variables (IV) approach replacing fathers' reported earnings with a measure which is correlated with Y^*_{t-1} but uncorrelated with either the transitory error term ε or the equation error w. Among the variables used to instrument for parental earnings are father's education, father's social class and a set of variables relating to the accommodation occupied by the father. These latter variables include a dummy for whether or not the father owned the house they lived in, proxies for the size of the house and for the availability of both indoor toilet facilities and hot water. The housing variables are likely to serve as good instruments in that they are correlated with lifetime earnings but uncorrelated with the transitory component of observed earnings (it's unlikely that an individual will buy a house as a result of a temporary increase in earnings).[11,12] All of our instruments have the expected significant signs and a likelihood ratio test that the coefficients on the instruments are jointly zero is easily rejected. [13] The second stage estimate of ρ is given in column two of table 1. As expected correcting for measurement error increases the estimated ρ to .62. This is of similar magnitude to Solon's (1992) IV estimate of .53 and suggests that our earlier estimate of .29 overstates the extent of earnings mobility. Rather than characterizing a society in which mobility is prevalent our corrected estimate suggests that the U.K is best characterized by an economy in which ancestral advantages or disadvantages persist

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[11] As pointed out by both Solon (1992) and Zimmerman (1992) the IV approach may lead to an upward bias in the estimate of ρ if the instruments also appear directly in the structural model. We test for instrument validity using a Sargan test, the results of which we report in Table 1. This test fails to reject the hypothesis that our instruments are uncorrelated with the error in the sons equation. Solon and Zimmerman also use average father's earnings over a number of years as an instrument. However we do not have multiple measures of parents' earnings available in the NCDS. We have multiple measures for sons' earnings. However under standard assumptions about measurement error in a left hand side variable using these multiple measures will improve the efficiency of our estimates but will not affect the estimate of ρ.

[12] Since father's earnings are reported in intervals we estimate the first stage equation by maximum likelihood.

[13] See Bound and Jaeger (1994) for a discussion of the finite sample biases which may arise using IV if the correlation between the instruments and the instrumented variable is weak.

for several generations. For example our estimate of .62 implies that a son born to a man earning twice the mean earnings will earn 2.36 times that of a son born to a man earning half the average earnings and that differences between the top and bottom deciles of the earnings distribution persist for over 9 generations.

To examine the extent to which this estimate of ρ reflects the impact of credit constraints we control for inherited endowments by including the results of two ability tests carried out when the children were aged 7. The availability of test scores when the children were young is important in that the tests precede much of the child's formal schooling. The IV results when these measures are included are given in column 3 of table 1. Looking at the effect of ability on earnings we see that in both cases higher ability, even at age 7, leads to higher earnings at age 33. While the coefficient on father's earnings falls from .64 to .55, it is remains large and significant.

There are three possible problems associated with identifying ρ using ability tests taken when the child is aged 7. Firstly, the results from a battery of tests administered on the same day may be a poor reflection of the child's ability if the child's performance on that day was not typical of the child's true capabilities. Secondly, this ability measure may reflect human capital acquired through formal or informal schooling and finally test scores may be poor indicators of endowments coefficients on all variables are affected which could cause our estimate of ρ to be biased.

To correct for test/retest measurement error we instrument our endowment variable using the results from similar tests administered when the child was aged 11 and aged 16. The results are presented in the final column of Table 1. Our estimate of ρ is still significant and equal to .42 which is similar to the raw correlation reported by AMT. In the context of the model specified in section 2 these results suggest that a substantial proportion of the correlation in earnings across generations comes from parental financing of the child's development rather than directly inherited endowments. To the extent that this is the case programs that equalise the marginal cost of funds to finance human capital will lead to a more efficient outcome with lower inequality.

If our ability measure reflects not only inherited ability but also skills acquired through formal schooling or from interaction with parents, controlling for E_t also purges the initial coefficient of some skills which may have been acquired through investments . Having ability measures at age 7, reduces the likelihood of this but it is possible that important skills are acquired at early levels of schooling or through informal learning at home, preceding formal schooling. To the extent that these skills reflect the ability of rich parents to place their children in better nurseries or the fact that richer parents may be better able to afford to take time off work to spend with children then we would want these effects to be included in ρ. However this argument would suggest that our estimate of .42 is biased downwards and only strengthens the case for programs which equalise opportunities.

If our measure of endowments capture test taking ability rather than productive abilities, as would be the case if children simply learned how to score well on tests, then our use of test scores to control for endowments would be invalid. While we recognise that this may be a problem we would argue that it is less likely to be true of tests administered at age 7.

b. Relationship between father's and son's unemployment experiences

The previous section examined the intergenerational links in earnings. However as noted by Altonji and Dunn (1991), few studies have attempted to examine family links in the main components that influence earnings and income. In this section we carry out such an analysis by examining intergenerational links in unemployment experience. The model outlined in section 1 assumes that differences in transfers from parents account for differences in individuals' human capital and thus earnings while working. However differences in human capital may also affect the probability of working and therefore provide another channel through which poverty is transmitted across generations. To measure the unemployment history of the children we use the diary information provided in sweeps 4 and 5 of the NCDS.[14] This provides us with a complete month by month record of individuals' work histories over the 17 year period from January 1974 to January 1991. For fathers we use information from sweeps 2 and 3, at which stage the fathers were on average 41 and 46 years of age respectively. In both these surveys fathers were asked to report the number of weeks off work in the previous year through unemployment. Using unemployment records rather than earnings relaxes the assumption that the father and son be working and provides an alternative measure of economic status.

[14] The work histories used in constructing childrens' unemployment histories are derived by merging the diary information in sweeps 4 and 5 and using subsidiary questions such as age when finished schooling to fill in missing data. For a more detailed discussion of these work histories see Dolton and O'Neill (1995).

To examine the relationship between fathers' and sons' unemployment histories we estimate a generalised Poisson regression model (Hausman, Hall and Griliches (1984)). A familiar problem with the traditional poisson approach is that count data tend to exhibit overdispersion in that the mean is less than and not equal to the variance as implied by the Poisson process. Extra poisson variation may arise for instance if some of the causes of the poisson process are unobserved. The generalised Poisson approach controls for this by introducing an unobserved random variable z into the traditional poisson regression. Conditional on z the dependent variable y_i is assumed to reflect random drawings from a Poisson distribution with parameter $z\lambda_i$:

$$f(y_i \mid z) = \frac{e^{-z\lambda_i} (z\lambda_i)^{y_i}}{y_i\,!} \qquad y_i = 0, 1, 2, ..$$

In our specification λ_i is related to the X's by the following equation :

$$\ln \lambda_i = X_i \beta$$

The unconditional likelihood is obtained by integrating over z. Assuming that z has a gamma distribution[15] it can be shown that the unconditional distribution of y_i is a negative binomial. The introduction of the unobserved random variable z thus breaks the mean-variance equality imposed by the standard Poisson model.

In our analysis we use two measures of sons unemployment histories. The first measures the number of months the son was unemployed between 1974 and 1991, while the second measures the number of spells of unemployment the son experienced between 1974 and 1991. The advantage of the first measure is that it captures the duration of unemployment spells. The disadvantage of this measure is that it is unlikely to satisfy the Poisson assumption that events occurring in disjoint time intervals be independent. This is due to the tendency for months of unemployment to cluster. Using spells of unemployment compresses information on

[15] The gamma assumption is common in this literature and results in a closed form expression for the unconditional likelihood.

unemployment duration but is more likely to satisfy the independent increments assumption.

As a regressor we use the number of weeks spent unemployed by the father in 1969 and 1974. Approximately 10% of fathers in our sample experienced a spell of unemployment in either 1969 or 1974. Having unemployment histories 5 years apart during the fathers peak earnings period may provide a better indicator of permanent income than earnings at a point in time, as well as allowing us to examine another dimension of intergenerational poverty transmission.

The results are presented in Table 2[16]. Rather than report the coefficient estimates β we report the relative incidence rate ratios. The incidence rate ratio for a Δx_i change in x_i, holding all other x's constant is $e^{\beta i \Delta xi}$. The final row of the table reports the Likelihood ratio statistic testing that the data follow a Poisson process. In each case the Poisson assumption is strongly rejected justifying the use of the generalised Poisson approach.

The estimates in columns 1 and 2 relate to months of unemployment and spells of unemployment respectively. Both these results show a significant positive relationship between fathers' and sons' unemployment histories, with sons of fathers who had been unemployed being significantly more likely to experience a period of unemployment. For instance the coefficient of weeks unemployed in column 2 implies that the son of a father who has 6 months unemployment is 1.5 times more likely to experience a spell of unemployment than the son of a father with no unemployment.

To examine the channels through which poor unemployment prospects may be passed on from generation to generation the results in column 3 and 4 include a measure of son's region of residence. As expected sons living in the Greater London area have lower unemployment rates than those living outside. However, controlling for region has only a small impact on the

[16] The small sample sizes in this table reflect non-response of the father to the unemployment question in 1974. However we have also estimated the same specifications using only 1969 unemployment history for the father, for which the sample sizes were over 1500 and obtained even stronger persistence effects.

intergenerational coefficient. In the months equation the coefficient falls from 1.021 to 1.019, while in the spells equation it falls from 1.039 to 1.037. The p-values associated with testing that whether these coefficients are equal to 1 (the null-hypothesis of no effect) are .055 and .0047 for the months and spells equation respectively. These results suggest that immobility of workers has only a small role to play in the intergenerational transfer process.

5. Conclusion

In this paper we have used a nationally representative sample of father-son pairs to examine the persistence of low earnings and inequality across generations in the U.K. Our findings are consistent with the recent literature in the U.S which finds limited earnings mobility when allowances are made for measurement error in the earnings data. Our estimate of the correlation between fathers and sons earnings is approximately .6, which implies that the son of a person at the top decile of the earnings distribution earns twice as much as the son of a person at the bottom decile. This advantage is larger than estimates of the increase in earnings associated with obtaining a college degree over a high-school degree and imply a static society in which family circumstances at birth play an important role in determining a child's future.

We also examine broader aspects of economic success, namely the propensity to become unemployed. This allows us to examine more general aspects of poverty and may also provide a more reliable indicator of an individual's lifetime economic worth than earnings or income observed at a point in time. Using this measure we again find that poverty persists across generations, with sons of fathers who have experienced unemployment having unemployment rates which were significantly higher than sons of fathers who were not unemployed.

Finally we use measures of individual ability obtained from tests administered to the children at the age of seven to identify the role of credit constraints in transmitting inequality across generations. We find that ability is an important determinant of economic success, however the persistence of inequality remains after controlling for individual endowments. In the context of the model specified in the paper this implies that part of the transmission of poverty may work through constraints on borrowing and that programs which equalise the marginal cost of funds for financing human capital accumulation will lead to a more efficient outcome with lower inequality.

Table 1
Estimates of Intergenerational Mobility Parameters
(Standard Errors in parentheses)

Variable	OLS	IV1	IV2[a]	IV3[b]
Log Father's Earnings	.29* (.031)	.62* (.056)	.53* (.056)	.43* (.058)
Log Father's Earnings Squared				
Middle Math			.04* (.027)	.11* (.035)
Higher Math			.11* (.029)	.18* (.034)
Middle Reading			.07* (.025)	.05 (.029)
Higher Reading			.11* (.028)	.11* (.035)
Sample Size	1330	1330	1330	1330
Sargan Test (p-value)		.072	.144	.062

[a] The test scores are divided into deciles, with 'Higher' referring to the top 30 percent and 'Middle' referring to the middle forty percent of the total distribution. This allows for nonlinear effects of endowments on earnings.
[b] In this specification both the fathers' earnings and the endowment measures are instrumented.

Table 2
Estimates of Intergenerational Links
Between Father and Sons Unemployment Histories
(Standard Errors in parentheses)

Variable	(1) Months	(2) Spells	(3) Months	(4) Spells
Father's weeks unemployed	1.02* (.01)	1.016* (.005)	1.0199 (.0105)	1.015* (.005)
Son's Region			.68977* (.1259)	.779* (.089)
Sample Size	935	935	935	935
Log Likelihood	-2204.5	-1217.5	-2202.5	-1215.2
LR Test χ^2 (1)	20525.9	268.4	20326.0	261.6

Data Appendix
Variable Description
Sample Means are given in parenthesis

Son's Weekly Wage in 1991 (246.23)

Father's Weekly Wage in 1974 (37.08)

Middle Math: dummy variable=1 if individual was in the middle 40% of the entire arithmetic score (administered at age 7) distribution (.41)

Higher Math: dummy variable =1 if individual was in the top 30% of the entire arithmetic score (administered at age 7) distribution (.38)

Middle Reading: dummy variable=1 if individual was in the middle 40% of the entire reading score (administered at age 7) distribution (.41)

Higher Reading: dummy variable=1 if individual was in the top 30% of the entire reading score (administered at age 7) distribution. (.34)

Son's Unemployment : months unemployed between Jan. 1974 and Jan. 1991. (8.18)

Father's Weeks Unemployed : total weeks unemployed in 1969 and 1974 (1.87)

Son's Region: dummy variable=1 if son lived in London (.31).

<u>Instruments for Father's Income</u> (Sample size 7211)

Fathers Education constructed from age father left school (yrs) (10.45)
Professional: Dummy variable =1 if father's social class was professional (.05)
Intermediate: Dummy variable =1 if father's social class was Intermediate (.29)
Skilled Manual: Dummy variable =1 if father's social class was skilled manual. (.61)
House Owner: Dummy variable =1 if father lives in owner occupied dwelling. (.51)
House Size. Dummy variable =1 if number of rooms in house are greater than 6. (.89)
Kitchen Size: Dummy variable =1 if kitchen > 6ft wide and not used as a living room. (.65)
Toilet Facilities Dummy variable =1 if have indoor toilet facilities (.96)
Hot Water Supply Dummy variable =1 if have hot water supply (.99)

<u>Instruments for Test scores at age 7.</u> (Sample size in bold)

Reading Comprehension Score at age 11	**9486**	(16.33)
Mathematics Test Score at age 11	**9404**	(17.37)
Reading Comprehension at age 16	**9486**	(25.63)
Mathematics Comprehension at age 16	**9404**	(12.98)

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