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# "The Effect of Adolescent Experience on Labor Market Outcomes: The Case of Height"

by

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# The Effect of Adolescent Experience on Labor Market Outcomes: The Case of Height*

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**Abstract**

Taller workers receive a wage premium, and the disparity in wages is similar in magnitude to the race and gender gaps. We exploit the variation in an individual's height over time to explore the ways in which height affects wages. Specifically, we show that for white males the effect of adult height is essentially eliminated when adolescent height is taken into account. We find that participation in high school sports and clubs, and to a lesser extent schooling, are channels through which teen height affects adult wages. The benefits of being a taller teen seem to accrue equally across income classes and also across broad occupation categories, suggesting that the benefits of teen height do not result from occupational sorting. Because height is heritable and because tall adults tend to have children with each other, the benefits of teen height tend to be perpetuated across generations. Finally, we use our estimates of the teen height premium to perform a simple calculation of the monetary benefits of a newly approved treatment for children that increases height.

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# 1 Introduction

Labor market outcomes are likely to differ depending on a person's outward characteristics. These differences have motivated a large body of research focussed on the disparities across racial and gender groups. Beyond establishing the magnitude of the disparities, a goal of this research is to identify the channels through which the gaps develop. In this paper we take up the same research agenda with respect to height. We start by estimating the magnitude of the height premium, and find it comparable to those associated with race and gender.[1] We then take advantage of a special feature of height relative to race and gender: height varies over time, so that a relatively tall 16-year old may turn out to be a relatively short adult, and vice versa. This time-variation allows us to investigate the stage of development at which having the characteristic (in our case, being short) most strongly determines the wage disparity. We find that being relatively short through the teen years (as opposed to adulthood or early childhood) essentially determines the returns to height. We document that the beneficial effects of teen height are not complementary with any particular vocation path, broadly defined; instead, they manifest themselves in a higher level of achievement in all vocation categories. We point out some social activities that might be important channels for the emergence of the height premium. We use our estimates of the return to teen height to evaluate the monetary incentive to undertake a newly approved treatment that increases teen height, human growth hormone therapy. Finally, we show that teen height is predictably greater for sons of tall parents, meaning that there is an expected wage penalty incurred by the as-yet-unborn children of short parents.

Height is widely believed to be an important ingredient of professional and personal success. Popular books discuss the advantages of being tall.[2] In the past 13 US presidential elections, the taller candidate has won 10 times (the most recent exception being George W. Bush) and, as shown in Figure 1, presidents tend to be distinctly taller than the average population.[3]

[1] While for methodological purposes we will compare our analysis with the literature on racial and gender discrimination, we do not imply that height discrimination, if the term were well-defined, is morally equivalent to racial or gender bias.

[2] See, for example, *The Height of Your Life* by Ralph Keyes: Little, Brown and Co. 1980.

[3] The height of Presidents is taken from http://www.uvm.edu/~tshepard/tall.html. The average height in the population is taken from Steckel (1995) and is the adult height of white males *born* in the US around the year in which the president was in office. Because average height is trending up, in any given year this measure most likely overestimates the average height of *adults* in the US population. The period 1850-1900 over which the height trend flattens is a possible exception.

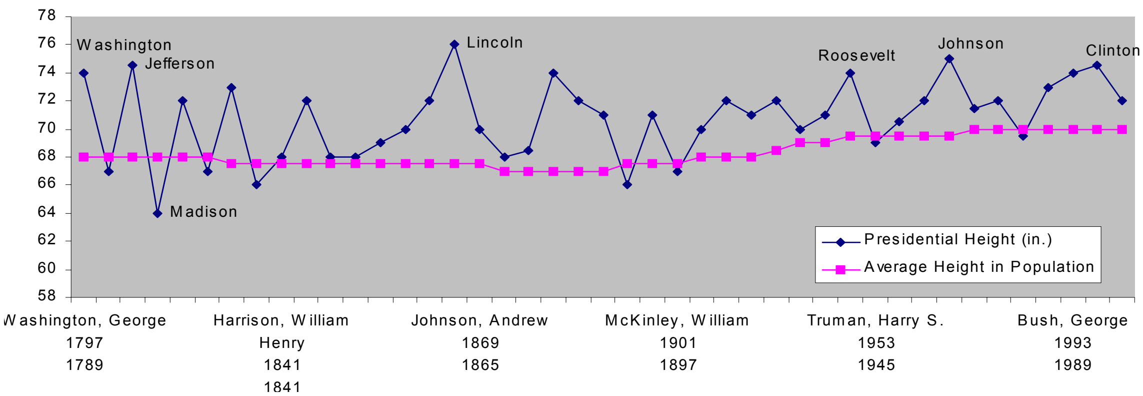


Figure 1: Height of US Presidents

There is an academic literature that investigates the possibility that labor markets reward height separate from ability.[4] Following a standard approach that accounts for differences in productive characteristics and interprets the residual wage differential as a height premium, prior research has estimated that an additional inch of height is associated with a 0.025-5.5 percent increase in predicted wages. Taking into account the potential biases allowed by most of the previous literature,[5] and using data from Britain's National Child Development Survey (NCDS) we find that among white British men every additional inch of adult height is associated with a 2.2 percent increase in wages. In a complementary analysis, drawing on data from the National Longitudinal Survey of Youth (NLSY) we find that among adult white males in the US, every additional inch of height as an adult is associated with a 1.8 percent increase in wages. As the interquartile range of adult heights spans 3.5 to 5 inches in our data (in Britain and the US, respectively), the tallest quarter of the population has a median wage that is more than 13 percent higher than that of the shortest quarter. The impact of this height wage disparity is comparable to those associated with characteristics such as race or gender.[6]

[4] This research is found mainly in sociology and psychology. For a review of the evidence from sociology and psychology see Martel and Biller (1987). A more recent example from this literature is Frieze *et al.* (1992). This evidence is not only drawn from less developed economies where physical size may be an important determinant of productivity (see Steckel (1995) for a review of the connection between height and standard of living). In economics and related literatures see, e.g., Behrman and Rosenzweig (2001), Loh (1993), and Sargent and Blanchflower (1994).

[5] Berhman and Rosenzweig (2001) is an important exception.

[6] Correcting for differences in family background and region of residence, we estimate the black-

Beyond estimating the magnitude of the height premium, our primary focus is to investigate its roots. Several plausible theories have been proposed for why markets might treat shorter people differently. A leading theory in social psychology describes the **interpersonal dominance** derived from height. According to this theory, short people are stigmatized by others, perceived less positively, and thus placed at a disadvantage in negotiating interpersonal dealings.[7] **Evolutionary selection** may also explain the disadvantages of being shorter than competitors. As the human species evolved, to the extent that size provided a direct advantage in the competition for resources, a preference for associating with tall people might have been naturally selected. In addition, greater height may have signaled good health throughout the development process, and therefore a genetic makeup robust to illness and deprivation. To the degree that this signal translated into a preference for taller mates, this may provide an explanation for why, other things equal, shorter people may be viewed as less valuable. These theories are designed to explain a 'taste' for height among employers. A final theory emphasizes **self-esteem** by placing the roots of the height premium in a superior conception of self that is achieved through a comparison with a socially-determined notion of ideal height. A greater self-image leads to higher achievement through a variety of channels, including perseverance and interpersonal skills.

The theories presented above may account for the adverse consequences of a *current* lack of height. Of course, the fact that we observe a height wage premium does not imply that shorter workers are penalized for their current stature. There may be another characteristic, correlated with current height and valuable to employers that is in fact acquired at some pre-market stage. We can think of this characteristic as a form of human capital, a set of skills that is accumulated at earlier stages of development. If this characteristic were unobservable to the researcher, the lower wages of shorter people would be incorrectly ascribed to their lack of height, instead of to their lack of human capital. For example, short children, if stigmatized because of their stature, might find it more difficult to develop interpersonal skills, positive self-conception, or might simply be excluded from participation in groups that foster the development of skills. The mechanisms that generate the disadvantage of short people in acquiring human capital could include any of the channels presented above (interpersonal dominance, self-image, etc.) which may have an impact at any **early stage of development**.

white wage gap to be approximately 15% among full time male workers in the NLSY. Similar analysis indicates that the male-female wage gap is approximately 20% among white full-time workers in both the NCDS and NLSY.

[7] See, Martel and Biller (1987) and Frieze, et al. (1990), (1978).

An alternative theory, **statistical discrimination**, might predict that children who forecast being short adults invest less in human capital because the returns to human capital are smaller for short adults; a rational individual will invest less in assets that provide lower personal return.

Finally, we might entertain a theory that pushes back the source of the height premium to birth: taller people might be endowed to a greater degree with some favorable characteristics. These characteristics could be **family resources** which raise a person's productivity, or other characteristics such as intellectual stamina or work energy, that are **directly productive characteristics** independent of external factors.

Distinguishing among these theories is important for understanding the channels through which outward characteristics affect market outcomes. The magnitude of the height premium alone makes it important to investigate these theories. In addition, understanding the ways in which height affects income may shed light on other labor market disparities such as the race and gender gaps. In the case of the height premium it is possible to make progress on tests of the relative importance of these theories. Alone among common bases of labor market disparities, height is impermanent. Relative stature, as defined below, often changes as an individual grows to his or her full adult height. Participants in the NCDS were measured by physicians at ages 7, 11, 16, and 33, and self-reported height at age 23. Respondents to the NLSY provided self-reported measures of height in 1981, 1982 and 1985. Of particular interest for the present study is that in each of these samples, among those who were relatively short when young, many grew to become average height or even tall adults. These changes in relative stature provide an opportunity to understand better the sources of the height premium.

We show that two adults of the same age and height who were different heights at age 16 are treated differently on the labor market: the person who was taller as a teen earns more. In fact, the preponderance of the disadvantage experienced by shorter adults on the labor market can be explained by the fact that, on average, these adults were also shorter at age 16. This finding suggests that a large fraction of the disparity is not due to a taste for tall adult workers, or to any employer's preference for height when young (which the employer presumably cannot observe); rather, the disparity must reflect a characteristic correlated with height when young. This observation leads our investigation away from a theory of the labor market's taste for height, and toward an analysis of the nature of the unobservable characteristic.

We first show that the teen height premium does not much diminish when we control for variables such as family resources, good health, and native intelligence.

Then, using the fact that NCDS measures height at ages 7 and 11 in addition to 16 and 33 we are able to parse the contribution to the height premium of being tall at different ages. When we regress wages on height measured at ages 7, 11, 16, and 33 we find that only age-16 height has either an economically or statistically significant coefficient. Among the different heights, that is, height at age 16 uniquely influences future wages. The negligible role played by height prior to age 16 (together with other supporting evidence) suggests that the height premium is not simply a premium to early development at any age, and, in fact, our findings place a lower bound on the time at which height appears to play its role.

The wealth of data provided by the NLSY allows an analysis of the channels through which teen height influences later wages. In the US data, those who were relatively short when young were less likely to participate in social activities associated with the accumulation of productive skills and attributes.[8] About half of the wage differential can be accounted for by variation in participation in school-sponsored non-academic activities (such as athletics and clubs), and a smaller fraction of it can be explained by greater levels of schooling. We interpret these findings as evidence of the effects of social factors on the development of human capital and the distribution of economic outcomes. Viewed in this light, our findings suggest that social effects during adolescence, rather than contemporaneous labor market discrimination or correlation with productive attributes may be at the root of the disparity in wages across heights.

In our effort to better understand the channels through which teen height affects wages, we establish that the height premium is not reflected in an observably different choice of occupations, broadly defined. Thus, the height premium is reflected in a higher level of achievement within the same vocation category, broadly defined. Consistent with these findings, we also find that the beneficial effects of teen height are more or less equally manifested across income levels. Overall, the beneficial effects of teen height seem to accrue across the board, as a sort of all-purpose resource.

Finally, we use our estimates of the teen height premium to calculate the monetary return to an investment in teen height, i.e., a treatment with human growth hormone (HGH). From an economist's viewpoint, the HGH treatment represents an indirect investment in human/interpersonal capital of the type that we have suggested. Because

[8] Examples of human capital that might be acquired through such activities include skills of interpersonal negotiation, social adaptablity, and motivation. Productive attributes that are often ascribed to participation in extra-curricular activities include self-esteem and self-discipline. See Heckman (2000) for a complete discussion of the importance of these 'non-cognitive' skills, and their development by social institutions.

the treatment was not available to the individuals in our data set, we have a unique opportunity to perform a cost-benefit analysis of an investment in this type of human capital. Since the return to investment is a percentage of adult wages, taking into account the cost of treatment we find that males with expected average annual earnings exceeding $100,000 have a monetary incentive to undertake the HGH treatment.

Our finding that early factors are important determinants of between-group wage disparities parallels the evidence presented in Neal and Johnson (1996) regarding racial discrimination. Controlling only for 'premarket factors' proxied by a teen test score, Neal and Johnson find that the adult black-white wage gap is significantly reduced for men, and disappears (statistically) for women. Neal and Johnson go on to account for much of the disparity in the test scores with the differences in family and school resources between black and white children. In this dimension, the situation of men who were short when young diverges from Neal and Johnson's analysis of the black-white gap in an important way. Differences in family and school resources explain little of the disparity between tall and short adults. Indeed, conditional on adult height, these pre-market characteristics appear little correlated with height when young. Thus, from a policy perspective, our message with respect to short adolescents is quite different from that of Neal and Johnson. While they place the roots of the racial wage gap early in the development of a child (and possibly even before birth, through the contribution of family circumstances) and make it essentially a problem of resources, we emphasize the teen years and the role of social factors as determinants of the height premium.

## 1.1 Related literature

Our analysis relates most directly to previous studies of the effects of height on market outcomes. Frieze et al. (1990) considers the relationship between height and salary for a sample of M.B.A.'s who graduated during the decade following 1972. Controlling for several demographic and anthropometric characteristics, as well as detailed measures of human capital, Frieze, et al. find that every additional inch of height is associated with an increase in salary of $570. The black-white salary gap in the same sample is estimated to be $1,000. As business school graduates, the sample studied in Frieze, et al. (1990) represents an elite subset of the working population. These findings therefore suggest that the premium received by taller workers should not be attributed solely to their greater physical capacity. Above a low threshold, it is implausible that greater physical capability *per se* contributes to the marginal product of a professional business worker.

Using the variation in height *between* monozygotic female twins, Behrman and Rosenzweig (2001) are able to control thoroughly for both productive genetic and family background characteristics that may be correlated with height.[9] Based on these unique data, Behrman and Rosenzweig estimate that every additional inch of height is associated with a 3.5-5.5 percent increase in women's wages. While the twins data allow Behrman and Rosenzweig to control for genetic and family background characteristics, there is no information on the twins' paths to their adult height, and thus no opportunity to perform our investigation into the sources of height wage differentials.

Sargent and Blanchflower (1994) provide evidence that the US is not unique among industrialized economies in the premia it pays taller workers. Using Britain's NCDS, Sargent and Blanchflower estimate the effect on age-23 wages of height and body-mass index measured at age $x$, for the cohort born in 1958. A separate regression is run for $x = 11, 16$, and 23. Controlling for a number of other characteristics including the respondent's educational qualifications and industry, Sargent and Blanchflower estimate that every additional 10 centimeters in height at age 16 (23) is associated with 2.7 (3.3) percent increase in wages at age 23. As an estimate of the cumulative effect of height on wages, however, Sargent and Blanchflower's study is limited by the fact that their oldest worker is just 23 years old.

Using an early release of the NLSY data, Loh (1993) regresses wages on adult height. He estimates that workers who are shorter (below-average) as adults earn 4%-6% less than above-average workers. These estimates account for differences in adult human capital, occupation, local labor markets, and other demographic characteristics. As in Sargent and Blanchflower (1994), however, Loh's study is limited by the fact that his oldest worker is only 24 years old.

Taken together, the evidence accumulated in this literature suggests an economically important, positive association between height and wages. Loh (1993) provides evidence that this relationship holds across broad sections of the US. population, while Sargent and Blanchflower (1994) demonstrates that the apparent advantage to taller workers is not unique to the US. Behrman and Rosenzweig (2001) shows that the relationship between height and wages cannot be ascribed to different family circumstances or even to differences in genetic endowments. Frieze, et al. (1990) suggest that the estimated height-wage differential is not merely capturing height's directly productive aspects.

[9] There is considerable variation in adult height within a pair of monozygotic twins. See, e.g., Behrman and Rosenzweig (2001).

# 2 Data

Our two main data sources are the NCDS and the 1979 youth cohort of the NLSY. The findings from the NLSY closely parallel those in the NCDS. We draw attention to the aspects in which the results from the US and Britain are substantially different, and where some distinctive features of the each data set provide additional insight.

The NCDS began as a perinatal mortality study of *all* the children born in England, Scotland and Wales during the week beginning March 3, 1958. Seven years later, an attempt (sweep) was made to recontact all of the children who survived infancy. Similar sweeps were made again when the children were ages 11, 16, 23 and 33. At age 33, 11,407 (66%) of the original 17,414 children were recontacted and at least partially surveyed.[10] The NLSY began in 1979 with 12,686 men and women ages 14-21, and has interviewed this cohort every year until 1994, and every other year since then. Respondents to the NLSY were first asked to report their height in 1981, when they were ages 16-23, and most recently in 1985 when they were ages 22-29. We will refer to height measured in 1985 as adult height.

To avoid confounding the effects of race, gender and height discrimination we will carry out our analysis separately by race and gender. We focus our attention primarily on white males. In Britain this implies excluding the small number of native-born non-whites; we also exclude those participants in the NCDS who immigrated to Britain after 1958. In the US we focus on the 2,063 white, non-Hispanic males for whom there exists both adequate height data, and other information.[11]

Table 1 presents summary statistics of the height measures from our primary data sets, along with statistics from an unrelated survey of measured height in the US. We note a few features of the data. First, even when attention is restricted to white men, there is substantial cross-sectional variation in adult height. The data include respondents as short as 60 inches and as tall as 83 inches. The interquartile range spans 3.5 inches (NCDS) and 5 inches (NLSY). Second, while in the NCDS the average change in height between ages 16 and 33 (2.5 inches) represents a substantial fraction of total variation in adult height, in the NLSY the variation in height over time is limited by the fact that respondents were 16 or older when first measured, with more than half

[10] Selection analysis indicates that those from Scotland and the Northwest of England, and those with lower reading test scores at age 7 were less likely to respond to the fifth sweep. Elias and Blanchflower (1988) find similar results with respect to the fourth sweep.

[11] Our interest in adult wages also leads us to exclude the entire NLSY oversample of poor whites who were dropped from the survey after 1990.

being older than 18.[12] On average, the NLSY sample grew just 0.28 inches between 1981 and 1985. Of the NLSY respondents, 618 (30%) reported growth of at least one inch over the period; among those who grew, the average change was 1.68 inches. Nevertheless, as our later analysis shows, this variation in height over time is adequate to provide reasonably precise estimates of the relationship between youth height and adult outcomes, conditional on adult height.

A potentially important limitation of the NLSY data is that height is self-reported to the nearest inch, which raises the issue of measurement error. To illustrate, we note that among our white male subsample, 315 (15.2%) respondents report a height in 1985 that is strictly less than what they reported in 1981, and 75 (3.6%) report a decline in height of more than one inch.[13] By itself, the presence of classical measurement error may strengthen our results, as the error would be expected to bias the coefficients of interest towards zero.[14] Nevertheless, we would like a gauge for the accuracy of self reporting. By one measure the height data recorded in the NLSY appear reasonably accurate. The distribution of the NLSY's self-reported heights is quite similar to that of a national survey of carefully *measured* heights completed in 1980; with the distribution in the NLSY shifted slightly to the right, and having a fatter right tail.[15]

In contrast, in each sweep of the NCDS except for the age-23 sweep, height is measured by a physician. The advantages of the earlier and more accurate height measures in the NCDS are clear. Among the 1,772 white men for whom there exists sufficient data, the average growth between ages 16 and 33 is 2.54 inches, and just 44 (2.5%) report negative growth over this period.[16] We will use age-33 height as our measure of adult height, to eliminate any bias induced by self-reporting. The presence of another measure of adult height—self-reported at age 23—is useful because it allows

[12] It is estimated that in the US adult height is reached at a median age of 21.2 years for males, with a median growth after age 16 of slightly less than 1 inch (Roche 1992, pp. 104-5).

[13] Not all those who report a decrease in height need be in error. Damon, Stoudt, and McFarland (1966) p. 50 reports that adults shrink by an average of 0.95 inches over the course of a day.

[14] In fact, when we omit the extreme tails of the 1981-1985 growth distribution, the magnitude of the estimated effect of height when young increases See table A2.

[15] The National Health and Nutrition Examination Survey conducted by the National Center for Health Statistics between 1976 and 1980 measured standing height against a calibrated bar and used a camera to standardize recording. In this survey, the average measured height for white males ages 18-24 was 69.8 inches, with a standard deviation of 2.8 inches. See National Center for Health Statistics (1987), Table 13. The comparable figures in our subsample of the NLSY are 70.41, and 2.85.

[16] Among the shrinkers, 19 respondents report shrinking by an inch or more during this period. The NCDS has disadvantages as well. For large fractions of those successfully contacted in the later sweeps, data from their earlier sweeps (including height measures) are incomplete.

us to gauge the impact of the bias induced by self-reporting of heights. In the NCDS, at least, this bias is negligible: all the coefficients in this paper are almost identical when age-33 height is replaced by age-23 height.

# 3 Evidence of the Height Premium

Our first task is to examine whether in our data sets, consistent with the literature, there are sizable associations between height on wages. There are some important aspects in which our investigation differs from previous studies. Unlike much of the prior research, we estimate the regression equations separately by race and gender, and focus on the results for white, non-Hispanic males. As noted above, estimating the equations separately avoids confounding the effects of race, gender and height discrimination; moreover this approach allows all of the coefficients to differ by race and gender. In addition, unlike most prior studies, we are able to measure wages at a relatively advanced age (31–38) and thus capture the cumulative effects of differences in height. Finally, our approach to estimating the effect of height takes care to avoid controlling for variables such as education, work experience, and occupation that are endogenous, i.e. choice variables that may be influenced by height. This approach is consistent with the strategy taken by Neal and Johnson (1996) who, along with Heckman (1998), provide detailed arguments against accounting for differences in decision variables when estimating the effect of labor market discrimination.

To begin our assessment of the relationship between height and wages, Table 2 compares summary statistics of the white male subsample by above- and below-median adult height. For the adult outcomes in the NCDS we consider wages at age 33; in the NLSY we consider the data from 1996 when respondents were 31 to 38 years old. The statistics on adult wages summarize only the data for full time workers.[17] Comparing mean log of wages, we find that the average wage of shorter males is 11 percent lower than that of the taller group in the NCDS and 10 percent lower in the NLSY.

Importantly, these shorter and taller males come from family backgrounds that are also different. In particular, Table 2 shows that compared with their taller counter-

[17] In the NCDS, wages are defined as gross pay per reporting period divided by usual hours worked during reporting period. (The reporting period varies depending on how often the respondent is paid.) In the NLSY wages are defined as annual income from wages, salaries and tips, divided by annual hours worked. Full time workers are identified as those who worked more than 1,000 hours in the previous year. The results discussed here and elsewhere in the paper are qualitatively unchanged when other definitions of full time work are used.

parts, shorter males, on average, come from larger families with less educated parents who were less likely to have worked in skilled or professional occupations.[18] Thus, an immediate concern is that the disparities in the average adult outcomes of taller and shorter males reflect these differences in family background rather than any form of height premium. Growing up in families with less human and financial capital, shorter than average men may be placed at a disadvantage in the labor market for reasons that have nothing to do with their lack of height.

To account for the influence of these systematic differences in family background, Table 3 presents OLS estimates of the effect of height on wages holding constant a number of family characteristics.[19] Results from the NCDS are presented along side those from the NLSY. In Columns 1 and 5, the first, simple regression of log wages on height, age, and region of residence[20] indicates that every additional inch of adult height is associated with an increase in wages of 2.7% percent (or an average of approximately £422 ($756) in 1991 full-time equivalent annual earnings) in the NCDS, and 2.5 percent (or an average of $820 in 1996 full-time equivalent annual earnings) in the NLSY. In Columns 2 and 6, after controlling for family characteristics including parents' education and occupation status, and number of siblings, the coefficient on adult height is reduced to 2.2 percent in the NCDS and 1.8 percent in the NLSY.[21]

Neal and Johnson (1996) provides evidence of an important association between the resources of a child's school and the accumulation of valuable pre-market skills. With this evidence in mind, we move on consider whether differences in measures of school quality such as student-teacher ratio, disadvantaged student ratio, dropout rate and teacher turnover rate may explain more of the height wage premium. Because many schools did not respond to the NCDS and NLSY surveys, a substantial portion of our white male subsample (29% and 40%, respectively) is lost when we condition on these

[18] The mothers of shorter men in the NCDS are an exception.

[19] Here, as in our subsequent regression analyses, we implicitly assume that, conditional on other observables, an individual's heights at various ages are exogenously given. This assumption precludes, for example, a model in which, conditional on other observables, height at various ages is determined in part by parents' unobservable investment decisions that also contribute directly to adult productivity and thus to adult wages.

[20] There are small, but statistically significant differences in the distribution of heights across regions. We find, for example, that males from Scotland are on average 0.5 inches shorter than those from the East or the North Midlands of England. In the US, white males in the Northeast are, on average, 0.45 inches shorter than their counterparts in Northcentral states, and 0.59 inches shorter than in the Southern states.

[21] If endogenous variables such as years of completed education and occupation are controlled for, the coefficient remains statistically significant although, it is further reduced.

variables. Restricting attention to the remaining sample with sufficient data, Columns 3-4, and 7-8, present the effects of adding controls for measures school quality in the NCDS and NLSY, respectively. We find that while these measures of school quality are associated with adult wages in the expected way, they are not responsible for the height wage premium. In each data set introducing controls for school quality in the leaves the estimated effect of adult height on adult wages essentially unchanged.

In summary, although the estimated coefficients on height are somewhat reduced after accounting for differences in some external resources, the reduction is minor especially when compared with the analysis of Neal and Johnson (1996), who find that family and school variables may account for a large fraction of the racial wage gap. In our analysis, the fact that taller people tend to come from somewhat more advantaged families does not explain a large part of the height premium.

# 4 It's All in Teen Height

As noted in the introduction, the fact that shorter people are penalized on the labor market does not imply they are penalized for being short. We now argue that much of the wage disadvantage experienced by shorter people can be explained by a characteristic other than adult height, namely height in adolescence. This finding casts substantial doubt on the relevance of a taste for height as an explanation for the observed wage premium.

## 4.1 Irrelevance of Adult Height

We now show that adult height predicts wages only insofar as it is correlated with teen height. The evidence for this claim comes both from the NCDS and the NLSY. (For height when old and young, in the NCDS we use age-33 and age-16 height, in the NLSY height reported in 1985 and 1981.)

To fix the idea of our estimation, consider a random sample of 33 year-old males all of the same height. Among this group, individuals will have been more or less tall at age 16. More specifically, some will not have grown at all in the intervening years, while others may have grown several inches to achieve this adult height. Conditional on adult height, we find a sizable (and statistically significant) difference between the wages of late and early maturers— a 'teen-height premium.' In fact, of the total effect that might be ascribed to adult height discrimination, nearly all can be attributed to the fact that adults who are relatively tall at age 33 tend to be relatively tall at age 16.

A simple comparison of means already suggests that most of the adult-height wage gap is in fact a teen-height wage gap. Table 4 considers only those white male adults of above-median height (69.7 inches or taller in Britain and 71 inches or taller in the US). Among this taller adult group we exclude approximately the tallest 15 percent, and classify the remainder according to their height when young.[22] Approximately 30 (13) percent of these remaining 'tall adults' in Britain (US) were of median height or below when young (67.3 inches or smaller in the British data, 70 in the US data). As adults, 'tall youth-tall adults' are on average less than an inch taller than 'short youth-tall adults'; and yet their adult wages differ considerably. The wages of these 'short youth-tall adults' in Britain are 15 percent lower (10 percent in the US) than those of 'tall youth-tall adults,' a wage gap that meets or exceeds that between short and tall adults (Table 2).

A regression analysis is reported in Table 5. Each specification in Table 5 takes the following form:

$$w_i = \alpha_0 + \alpha_1 H_{i,adult} + \alpha_2 H_{i,youth} + \boldsymbol{\alpha}\mathbf{X}_i + u_i \quad (1)$$

where:

$$\begin{aligned} w_i &= i\text{'s adult wage} \\ H_{i,\text{adult}} &= \text{adult height} \\ H_{i,\text{youth}} &= \text{youth height} \\ \mathbf{X}_i &= \text{a vector of other covariates} \\ u_i &= \text{an error term} \end{aligned}$$

As above, the results from the NCDS and NLSY are presented side-by-side. In the first, basic specification of Table 5 (Columns 1 and 5) we regress adult wages on adult height, youth height, age, and region. In this basic specification we find that, conditional on adult height, every additional inch of height when young, is associated with a 2.6% increase in adult wages in Britain, and a 2.7% increase in the US. Importantly, when we control for youth height, the estimated effect of adult height on

[22] Omitting the tallest of the category of "tall adults" addresses the concern that the results of a simple comparison of above- and below-median may be driven by the fact that those tall adults who were short when young are also likely to be among the shorter 'tall adults,' and for this reason less well paid. Ignoring this fact could lead to an overestimate of the power of youth height to explain adult wage disparities.

wages is nearly zero. The point estimate suggests that conditional on youth height, any additional adult height is associated with a statistically insignificant 0.4% increase in adult wages in the British data. In the US data the estimated coefficient on adult height is a statistically insignificant 0.2%.

As in Section 3 concerning the analysis of adult height alone, we move on to account for a possible relationship between height and aspects of family and school background.[23] Adding controls for family characteristics (Columns 2 and 6) changes the estimates slightly. Accounting for differences in family background we find that, conditional on adult height, every additional inch of age-16 height is associated with a somewhat diminished, but still highly significant 2.1% increase in adult wages in Britain, and a 2.6% increase in the US. Controlling for the effect of youth height, we estimate that adult height is associated with a 0.5% but statistically insignificant increase in adult wages in Britain. In the US, the estimated effect of adult height on wages is -0.4% but not statistically different from zero. Columns 3-4, and 7-8 add controls for measures of school quality from the NCDS and NLSY respectively. As in the analysis of adult height alone (Table 3), we find that adding controls for differences in school quality leaves the estimated effects of youth height and adult height little changed. In the relevant samples, introducing controls for school quality leaves unchanged the estimated effect of an additional inch of youth height in the NCDS, and reduces it from 2.7% to 2.5% in the NLSY. Thus, the finding that pre-adult height, rather than adult height itself, determines the wage premium is robust to the introduction of controls for region, family background, and school quality.

Since in both the NCDS and the NLSY the effect of adult height on wages, conditional on youth height, is nearly zero, this analysis indicates that the adult height-wage disparity is not due to a taste for tall workers. Rather, the different outcomes for taller and shorter workers appear to reflect a characteristic correlated with height when young.

[23]It may be argued that, to capture the gross effect of teen height on adult wages, it is appropriate to condition on the teen's stock of human capital. One argument for such a specification is that, conditional on adult height and family resources, investments in human capital that are positively correlated with later wages are also positively correlated with teen height. Observe, however, that to the extent that these investments are the *result* of greater stature, conditioning on teen human capital would lead to an underestimate of the gross effect of teen height on adult wages. Consistent with this interpretation, our analysis indicates that while pre-teen investments in human capital are unrelated to the teen height premium, post-teen investments may be. See Tables 9 and 11 below.

## 4.2 Irrelevance of Adult Height is Robust to Trimming of Outliers in Height, Growth, and Wage

Given the relatively small degree of intertemporal variation in the NLSY height, it is important to check whether our results for that dataset are sensitive to outliers in the height, growth and wage distributions. (In the NCDS there is less reason for concern since teen-to-adult variation represents a substantial fraction of the total variation in age-33 height.) In Appendix A we experiment with trimming the top and bottom 1 and 5 percent of the distribution of: height, ages 16-33 growth, and wages. Our results in each data set are both qualitatively and quantitatively robust to the exclusion of these outliers.

## 4.3 Irrelevance of Adult Height is Robust to Non-Linear Specifications

Our linear regression analysis points to teen height as the source of the height premium under the maintained assumption that the relationship between these two variables is linear. If, however, we had evidence of non-linearities in the effect of height on wages, then we would have reason to think that the primacy of teen height is driven by the linear specification. The concern is that the linear youth height term is simply correlated with the higher order terms of adult height that actually determine adult wages. The non-parametric comparison of means discussed in Table 4 already casts some doubt on the plausibility of such a correlation driving our results, since that table indicates that most of the adult-height wage gap is in fact a teen-height wage gap. Nevertheless, in this section we summarize the findings of Appendix B, which is devoted to the issue of possible non-linearities. The analysis in Appendix B approaches the issue from three angles.

1. We allow nonlinear specifications for the relationship between teen height and wages. We find that the nonlinear specifications are imprecisely estimated, explain almost no more variation than the linear model, and are less stable than the simple linear equation. More formally, we cannot nearly reject a null hypothesis of a linear specification.[24] The lack of evidence for a non-linear relationship

[24] We do find, however, that a quadratic specification does help slightly with the fit of wages for those with extreme heights, in line with the notion that the a small increase in height should have a large effect on the wages of those who are quite short and a much smaller effect for individuals that are already very tall. The effect is not strong, however. In fact, the improvement in fit all but disappears

between teen height and adult wages suggests that our main finding, namely, the primacy of teen height in explaining the height premium, is not an artifact of the linear specification.

2. We present an analysis in which teen and adult height appear *simultaneously and non-linearly* as explanatory variables. This methodology provides another, more direct test of whether the key finding of section 4.1, namely, the primacy of teen height in explaining the height premium, depends on the linear specification. We find that we cannot nearly reject the hypothesis that adult height is irrelevant.[25]

3. We explore possible non-linearities in the effect of height depending on the level of wages. This question is addressed with quantile regressions.[26] We regress wages on teen height alone as well as coupled with adult height and find limited evidence of a relationship between the size of the height premium and the position in the conditional wage distribution (in the sense that the relationship between the size of the premium and wages does not appear to be monotonic).

## 4.4 Irrelevance of Pre-Teen Height

This section is based solely on data from the NCDS, which afford a unique opportunity to parse further the height premium according to the age at which relatively high stature is attained. Respondents to the NCDS were also measured at ages 7 and 11, allowing us to consider the extent to which height at these ages contributes to the height premium. It is clear that each of these heights, if considered on its own (without conditioning on other heights), will appear to carry a wage premium simply because of the positive correlation between heights at all ages. To determine the extent to which teen height proxies for pre-teen heights, we examine how the estimated contribution of teen height is changed when we introduce earlier heights.

Table 6 considers only those respondents for whom there exists data on age 33, 16, 11 and 7 heights. The basic estimation (Column 1) regresses log of age-33 wages on age-33 height, age-16 height, family background, region. Column 2 adds controls for both pre-teen heights. Comparing the results in Columns 1 and 2 we find first that

when we eliminate the top and bottom 1 percent in height.

[25] An exception is the NLSY when all data points are included (including the 1 percent outliers), where that hypothesis can be rejected (though not strongly). When we exclude 1 percent outliers, this anomaly disappears. This finding reinforces the notion that nonlinearities help fit the wages of height outliers in the NLSY.

[26] We are grateful to a referee for suggesting the use of quantile regressions in this context.

age-11 and 7 heights have no appreciable effect on adult wages. Conditional on all other heights, the estimated effect of an increase in either age-11 or age-7 height is nearly zero. In addition, introducing these controls for earlier height leaves the estimated effects of both age-33 and age-16 height essentially unchanged. Among all recorded heights, only age-16 height is estimated to have an economically large and statistically significant effect on adult wages; no other height makes an appreciable contribution to the height premium.

# 5 Explaining the Teen-Height Premium

## 5.1 Not Due to Employers' 'Taste' for Height

In the previous section we have pushed back the source of the height premium to the pre-adult stages of development. We have shown, in particular, that the wages of two men with the same adult height may reflect the full amount of a 'height gap' depending only on the difference in their heights as teens. We conclude that the height penalty cannot be caused by *current* lack of height. This finding casts doubt on the existence of an employers' taste for current height, or even of a taste for early height (since employers likely cannot differentiate based on height when young).

## 5.2 Not an Effect of Observable External Resources

As shown in Section 4.1, differences in family and school resources explain little of the disparity between tall and short adults. The coefficient on teen height (adult and/or youth) was little changed in both the British and the US data when we controlled for family and school background characteristics, and our finding of a 1.8-2.5% per-inch teen height premium is net of family background characteristics. The conclusion that the height premium is not driven by family background characteristics draws additional support from the analysis of Behrman and Rosenzweig (2001), who find evidence of a sizable wage premium between female twins of varying heights.

One may argue that the robustness of the teen height premium to controlling for resources is due to the necessarily imperfect quality of our measure of resources. According to this argument, teen height is a better indicator than our imperfect measure of resources of some productive unobservable which is directly related to resources. Our measures of resources, however, perform quite well in predicting performance on achievement tests. For example, our measures of family background can explain at least 29 percent of the variation in AFQT scores.

In addition, to the extent that external resources are correlated with height at all ages, if resources were driving the height premium then we should expect heights *at all ages* to be positively associated with wages. Since only teen height seems to matter in the data, the only remaining possibility is that heights at various ages may be proxying for the same unobservable, productive resources but teen height is much better correlated with these unobservable resources. In that case we might obtain the estimates we have, but our interpretation would be misleading.

Plausible candidates for unobservable productive factors may be summarized by the child's home environment. Home factors may contribute to health, cognitive development, and other forms of human capital. One way to assess the influence of such home factors in determining the estimated teen height premium is to test whether teen height is systematically a better predictor than all other heights of observable endowments that are plausibly correlated with productive home factors.[27] If teen height were, in fact, a better predictor of observable home endowments, we would have more reason to be concerned about the influence on our estimates of correlation between teen height and omitted productive home factors. Appendix C presents the results from regressions of various endowments on heights at different ages. Generally, we find that each of the heights is only very weakly associated with these observable endowments, and we find no evidence that teen height is a systematically better predictor of these endowments.

## 5.3 Not an Effect of Unobservable, Intergenerationally Correlated Resources

The previous subsection has ruled out the possibility that the teen height premium reflects a correlation between teen height and certain measurable resources. In addition, the previous subsection casts doubt on the possibility that the teen height premium reflects unobserved resources plausibly correlated with observable endowments. In this section we provide additional evidence that the teen height premium is not merely a reflection of omitted, unobserved resources, such as the wealth of the family, the amount of social connections, etc.

It is clear that unobservable resources such as family wealth and social connections ultimately play an important role in determining adult wages. While controlling for such unobservable variables is impossible in this context, inference may still be drawn about their effects as long as these variables are positively correlated across generations.

[27] We are grateful to Esther Duflo for suggesting this procedure.

Suppose, for example, our concern was that well-connected, high social class individuals are more likely to be tall as teens. Since well-connnectedness is presumably positively correlated across generations, controlling for the father's teen height would diminish the effect of teen height on wages. More generally, if controlling for the father's teen height reduces the size of the son's teen height premium, we can infer that the teen height premium partly reflects the omission of unobserved resources that are correlated across generations.

The NCDS allows us to implement a variant of this empirical strategy because that data set contains a measure of the father's adult height, a relatively good proxy for the father's teen height. Our strategy will be to assess the effect on the estimated teen height premium of introducing father's adult height as a control. The strategy is based on the following argument.

Suppose there is an $X$ factor that determines adult wages, is correlated with teen height, and is, in part, heritable. If we could observe that $X$ factor, then the regression (taking deviations from means)

$$X = \theta_1 H_{teen} + \upsilon$$

would have a positive coefficient $\theta_1$. The coefficient $\theta_1$ captures two relationships. First, conditional on the endowment of $X$ inherited from their fathers, taller teens tend to have more $X$. Second, taller teens tend to have taller fathers, who also tend to have higher $X$ which they pass on to their sons. Therefore, if we were to regress

$$X = \gamma_1 H_{teen} + \gamma_2 H_{dad} + \zeta$$

we would find that

$$\gamma_1 \leq \theta_1, \tag{2}$$

where a strict inequality would indicate that $X$ is at least partly heritable.

Of course we cannot observe this $X$ factor, but we can infer its relationship to teen height from wage regressions. Suppose the true model were:

$$w = \beta_1 H_{teen} + \beta_2 H_{dad} + \beta_3 X + \epsilon \tag{3}$$

The OLS estimator in the regression of $w$ on $H_{teen}$ alone is

$$b_1 = \left(H'_{teen} H_{teen}\right)^{-1} H'_{teen} w.$$

The OLS estimator in the regression of $w$ on both $H_{teen}$ and $H_{dad}$ is

$$\begin{bmatrix} \widetilde{b}_1 \\ \widetilde{b}_2 \end{bmatrix} = \left([H_{teen}, H_{dad}]'[H_{teen}, H_{dad}]\right)^{-1} [H_{teen}, H_{dad}]w.$$

Substituting for $w$ from (3), simplifying, and taking expectations yields

$$E[b_1] = \beta_1 + \beta_2\delta_1 + \beta_3\theta_1$$

$$E\begin{bmatrix}\widetilde{b}_1 \\ \widetilde{b}_2\end{bmatrix} = \begin{bmatrix}\beta_1 \\ \beta_2\end{bmatrix} + \beta_3\begin{bmatrix}\gamma_1 \\ \gamma_2\end{bmatrix},$$

where the coefficient $\delta_1$ denotes the OLS estimator from the regression of the father's adult height on the son's teen height.

When we run these regressions using the NCDS data, we find that $b_1 = \widetilde{b}_1$. (See Table 7). This means that

$$\begin{aligned} \beta_1 + \beta_2\delta_1 + \beta_3\theta_1 &= \beta_1 + \beta_3\gamma_1, \text{ or} \\ \beta_2\delta_1 &= \beta_3[\gamma_1 - \theta_1]. \end{aligned} \tag{4}$$

By assumption, $[\gamma_1 - \theta_1] \leq 0$ and $\beta_3 > 0$. Since height is heritable, $\delta_1 > 0$. Unless we believe that $\beta_2 < 0$, i.e., in the true model having a tall father directly depresses one's wages, then the only way to satisfy equation (4) is to have $\beta_2 = [\gamma_1 - \theta_1] = 0$, which means that (a) in the true model the father's height has no *directly* productive effect on the son's wages, and (b) that the unobserved factor $X$ is not inherited, i.e., we find no evidence of an unobservable resource that can account for the teen height premium and is correlated across generations.

## 5.4 Not a Proxy for Good Health

Another potential explanation for the teen height premium is that teen height proxies for health problems experienced before or during adolescence that inflict lasting damage and depress adult wages. Indirect evidence against this hypothesis is provided by the findings of Section 4.4, which indicate that height before adolescence does not account for the teen height premium. Direct evidence is available, to varying degrees, both in the NCDS and in the NLSY. These data sets allow us to investigate the importance of health in explaining the teen height premium.

In the NCDS, the physical exams that provide our height measures also provide detailed information on the respondents' health status. To the extent that the teen height premium is attributable to better health among the tall, conditioning on this health information would be expected to diminish the coefficient on age-16 height. Table 8 presents results from regressions in which we add to our set of explanatory variables the number of health conditions reported by the physician in the age 7 and age 16 exams.[28] Consistent with the idea that poor child or adolescent health has a lasting impact, these measures of health are have economically important, negative associations with adult wages. However, introducing these measures does nothing to reduce the estimated teen-height premium. Adding the controls leaves the estimated teen-height premium unchanged.

The NLSY lacks detailed health measures until the respondents reach their 40's. In every survey year since 1979, however, respondents were asked whether they have a health condition that limits the kind or amount of work that they can do. We control for these measures in the last four columns of Table 8. The results indicate that the kind of work health limitation has an economically important and statistically significant negative association with adult wages. The amount of work limitation measure has a positive point estimate, but is not statistically distinguishable from zero. In either case, the inclusion of these health measures does not appreciably affect the size of the teen height premium. In sum, controlling directly for health conditions leaves the estimate of the teen height premium essentially unchanged.

A complementary approach is pursued in Appendix C where we regress health measures on heights at different ages (see Table C1). Generally, we find that each of the heights is only very weakly associated with these health measures, and we find no evidence that teen height is a systematically better predictor of the health measures positively associated with wages.

We therefore conclude that, with respect to the height wage premium, teen height is not merely proxying for good health.

## 5.5 Not a Proxy for Weight

Another possibility is that teen height is proxying for weight. If, for example, short teens were more likely to be overweight and being overweight as a teen decreased expected adult wages, then we might incorrectly attribute to height some adverse effects that are in fact due to weight. Weight, however, is a choice variable to a greater degree than, for instance, external resources or even good health, and so we must be especially careful in interpreting any regression result. In Table D1 of the Appendix we regress adult wages on height alone (plus our usual controls), and then we look at how the coefficient on height changes as we introduce weight in the regression. To the extent that weight *is not* a choice variable for the individual, a decrease in the coefficient on

[28] This methodology follows Case, Fertig and Paxson (2003). See notes in Table 8 for a detailed description of the health conditions.

teen height would indicate that teen height is proxying for weight. To the extent that weight *is* a choice variable, a decrease in the coefficient on teen height would suggest that part of the effect of teen height on adult wages is channeled through weight. Table D1 provides estimates from both the UK and US data in which we condition on both adult and teen weight as well as height. We find that adding the controls for weight leaves our estimates of the adult and teen height premia essentially unchanged. We conclude that the teen height premium is largely independent of weight.

## 5.6 Not a Premium to Native Intelligence

Suppose height were proxying for native intelligence; given the pattern we observe whereby age-16 height alone among heights at all ages explains wages, the productive components of native intelligence must be most strongly correlated with age-16 height. Although this hypothesis seems peculiar, we can use the NCDS to investigate it by conditioning on the score of a test of academic achievement taken at age 7. Insofar as academic achievement at age 7 measures native intelligence, conditioning on the test score ought to reduce the coefficient on age-16 height. Table 9 presents the effect of introducing age-7 test scores on the coefficient for age-16 height. Note that all of these estimates account for differences in family backgrounds, so the test scores do not proxy for these characteristics.

Table 9 restricts attention to those respondents to the NCDS with information on height ages 16 and 33, and test scores at age 7. Consistent with the notion that they capture native intelligence, the test scores contribute importantly on their own to explaining adult wages. Each test score is associated with a statistically significant, positive coefficient (1.5% increase in wages per point on the reading test, and 2.4% per point on the math test). However, introducing the scores does not reduce appreciably the estimated teen-height premium. Without controlling for the test scores, the teen-height premium is estimated at 2.1% per inch in this sample. Adding the controls merely reduces the estimated teen-height premium to 1.9% per inch. A complementary approach is pursued in Appendix C where we regress the age 7 test scores on heights at different ages (see Table C1 columns 1-2). Generally, we find that each of the heights is only very weakly associated with these test scores, and we find no evidence that teen height is a systematically better predictor of these test scores.

Comparable measures of native intelligence are not available in the NLSY. The earliest standardized measure of intellectual ability is the AFQT, an achievement test administered in 1981, when the respondents are 16 or older. We will discuss achievement tests in Section 6.1.

## 5.7 Not a Reward to Early Cognitive Development

Later physical maturers might also be later cognitive or emotional maturers. If this were the case, we would expect those maturing later to, for example, get less from the same amount of schooling than their early maturing adult peers and, therefore, complete less school or do worse in the adult labor market. The notion is that at any age, being taller allows one to get more out of education. In this were the case, greater height would be beneficial at all ages and we would expect the coefficient on height at all pre-adult ages to be substantial in Table 6. The fact that we do not see this pattern suggest that there is no advantage to earlier development *per se*.

Alternatively, it might be argued that puberty has a special quality among stages of development. It may be that achieving puberty enables one to start accumulating a special kind of human capital, and those who achieve puberty early (and so are taller as teens) get a head start in the accumulation process. This could be the reason for the pre-eminence of teen height among all heights in explaining adult wages. According to this hypothesis, the teen years are not special because of the environment associated with them; rather, being tall as a teen is merely a symptom of early puberty and thus the precocious achievement of a large fraction of one's ultimate height. This argument can be explored by estimating the extent to which the fraction of one's ultimate height achieved as a teen, rather than height level, matters for adult wages. Table 10 introduces the fraction of final height achieved as a teen along with teen height level. This allows us to distinguish the effect on adult wages of being fully developed as a teen from just being tall on the way to greater heights. If early puberty were the key to larger wages, then the estimated coefficient of percentage height achieved should be large.

Columns 1 and 2 of Table 10 present results from the NCDS. In these specifications we estimate of the effect on wages of teen height alone, and of the effect of teen height conditional on percentage of adult height achieved, respectively. In Column 1 we regress the log of age-33 wages on age-16 height, family background and region, and estimate a 2.4% per inch teen height premium. Column 2 adds a control for the percentage of age-33 height achieved by age 16. In these British data we find no relationship between adult wages and the fraction of adult height achieved by age-16; and adding this control leaves the estimated effect of teen height essentially unchanged. In the US data (Columns 3 and 4) we observe an identical pattern. In Column 3 we estimate a

2.2 percent teen height premium, absent a control for fraction of adult height achieved. As in the British data adding the control for teen development (Column 4) leaves the estimated effect of teen height level basically unchanged. Here, as in the NCDS, this percent of maturity measure does not explain the teen height wage premium.

Another related concern is that, in the NLSY, those who are shorter when measured in 1981 are, after conditioning on year of birth, also on average younger. We observe, however, that controlling for quarter of birth in the regression of adult wages on teen height does not affect the teen height coefficient (see Table 10, column 5). This result suggests that the teen height premium does not merely reflect an age effect. In the NCDS this is not an issue because all subjects are born during the same week.

# 6 Channels

The previous section indicates that the advantage of being a tall teen is not due to some omitted resource variable such as native intelligence, health, etc., and thus we are lead to the conclusion that being tall as an adolescent facilitates the acquisition of some form of human capital. In this section we try to get a better handle on the form of this human capital by exploring some of the channels through which teen height affects wages. Our data sets afford a rich set of alternatives to be explored. We have data on occupation choice, self-esteem measures, high school sports and clubs participation, achievement test scores, and years of completed schooling.

These variables can help illuminate the paths by which taller teens come to have higher wages. If, for example, occupation choice were an important channel, we would expect taller teens to self-select themselves into certain types of occupations (which, incidentally, would imply that certain occupations are richer in the particular form of human capital acquired by tall teens). This phenomenon would indicate that the form of human capital acquired by tall teens is complementary with certain occupations. Conversely, if we found occupation choice to be largely independent of teen height, we would conclude that the form of human capital acquired by tall teens is a "general purpose" asset. Our strategy is to include the variables that represent these channels as controls in the regression of wages on teen height. When a variable substantially reduces the coefficient on teen height we will interpret that variable as an important channel through which teen height affects wages.

**Occupation Choice**

To investigate the correlation between teen height and occupational choice, we rank occupations according to the sample average teen height in the respondent's occupational category. We then use average occupation height as a control in our wage regression. This variable seems to play a limited role in mediating the effect of teen height (see Table 11, Columns 2 and 7). This means that, while tall teens earn more as adults, the premium is largely not a result of sorting across occupations. Consistent with this finding of little sorting, we find that in the US the difference between the average heights of those working in occupations with the 25th and 75th percentile of average height is only 0.48 inches (compare this with the difference between the population's 25th and 75th height percentile, which is 5 inches in the US). Occupational choice appears to be similarly unimportant if we condition directly on occupation codes instead of the average height within occupation (results not shown). Of course, this finding of little sorting takes as given the fact that we divide occupations into 12 categories in the US (and 17 in the UK). We choose this coarse classification to keep the estimates of average height within category relatively precise. This coarse classification presumably obscures more subtle differences in the professional achievement of tall and short teens. Nevertheless, at a first approximation, enjoying the benefits of teen height does not seem to require self-selecting into certain occupations.

**Self-esteem**

Self-esteem, in light of the social-psychological theories described in the introduction, is a natural measure to consider in our search for the channels for the height premium. Our measure of self-esteem is drawn from questions asked in the 1980 wave of the NLSY, when respondents were administered the Rosenberg self-esteem scale.[29] There is no self-esteem measure in the UK data. Self-esteem has a statistically significant and economically important association with wages (see Table 11, Column 8). Thus, individuals who have more self-esteem also earn more, on average. Self-esteem, however, seems to have little to do with the teen height premium. Conditioning on self-esteem leaves the estimated teen height premium essentially unchanged. Thus, tall teens do not appear to earn more because they have greater self-esteem as teens.

**Social Activities**

Having ruled out several possible channels through which teen height influences adult wages, we now provide evidence suggesting that participation in social activities

[29] The Rosenberg Self-Esteem Scale is constructed by adding up the scores (each ranging between 1 and 4) describing the extent to which respondents agreed with 10 statements about themselves. For example, respondents were asked the extent to which they agreed with the statements "I am a person of worth" and "I have very little to be proud of." The average self-esteem score among the white male subsample is 23.93, standard deviation 2.72. We note that self-esteem was measured in 1980, one year prior to the self-reported height measurement. Because of this time lag, our measure may be somewhat less accurate as a correlate of self-esteem in 1981.

in adolescence contributes importantly to the teen-height premium. To this end we restrict attention to the NLSY data set, which contains especially detailed information concerning participation in social activities. Those who were relatively short when young are less likely to participate in social activities that may facilitate the accumulation of productive human capital like social adaptability. We think of athletics, school clubs, and dating as examples of these types of activities. We show that participation in extracurricular activities plays a role in the teen-height premium.

Table 11, Column 9 presents estimates of the effects of height on adult wages, conditional on participation in high school social activities. Retrospective questions about participation in high school activities were asked in 1984, only to those who had finished or were expected to finish high school. Our measure of social activity is the number of non-vocational, non-academic high school clubs in which the respondent participated.[30] Because height is often a criterion for participation in athletics, we separate athletics from these other high school activities.[31] Last, we note that for the younger members of the sample, height in 1981 represents (at least in part) high school height. For those 19 and older in 1981, however, height in 1981 will be a noisier signal of high school height. The analysis is performed both for all white males for whom we have adequate data (Table 11, Column 9), and for those younger than 19 in 1981 alone (Table 12, Column 4).

The results for the all-age sample (Table 11, Column 9) indicate that participation in social activities is associated with a statistically and economically significant wage premium. Controlling for age, height, region and family background, and other club membership, participation in high school athletics is associated with an 11.7% increase in adult wages. Participation in every additional club other than athletics is associated with a 5.1% increase in wages. When we add controls for the levels of participation in high school activities the coefficient on youth height declines by a modest 22%. The estimated effect of adult height is qualitatively unchanged.

The effects of accounting for high school activities are more dramatic, however, when attention is restricted to those who were actually in high school in 1981 when their height was recorded. Table 12, Column 1 presents the basic regression for this younger group. Here we estimate that every inch of youth height is associated with a 2.6% increase in adult wages. Again the effect of adult height, while estimated at

[30] These clubs include youth groups, hobby clubs, student government, newspaper/yearbook, performing arts, and 'other' clubs. This list does not include, in particular, honor societies or vocational clubs. On average, the white male subsample participated in 0.69 clubs, standard deviation 1.00.

[31] Among the white males in our subsample, 51% participated in high school athletics.

-1.6% per inch, is statistically indistinguishable from zero. When, in Column 4, we add controls for the levels of participation in high school activities the coefficient on youth height declines by more than 38% and is no longer statistically significant at the 10% level. Again the coefficients on participation in activities are economically meaningful and statistically significant.

We should emphasize that one must be cautious in interpreting these regressions. Participation in athletics and clubs are choice variables and we have not modelled that choice. Thus, it would be incorrect to conclude that compulsory participation in athletics and/or clubs would raise expected adult wages. Consider sports, for example. If it were success that mattered for the wage premium, requiring shorter boys to participate (and fail) in sports would have no beneficial effect. There are obviously many other plausible models consistent with our finding a relationship between teen participation in athletics and adult wages. Understanding the particular form of that relationship is important, but is beyond the scope of this paper.[32]

## 6.1 Achievement Tests and Years of Schooling

Next we ask what achievement test scores and years of completed schooling reveal about channels for the teen height premium. Controlling for achievement test scores has a considerable effect on the teen height coefficient (see Table 11, Columns 3 and 10, and Table 12, Column 5) suggesting that an important channel through which teen height affects wages is captured by the achievement tests. Depending on the sample, however, the effect on the estimated teen height premium of conditioning on achievement tests depends on whether we control for, among other things, sports and clubs participation. Table 11, Column 12 shows the estimates from the entire US sample, where we condition on all of our potential channels variables. Comparing these results with those from Table 11, Column 10, the estimated AFQT premium is 57% of what it was in the absence of conditioning on the rest of our channels measures, including sports and clubs participation. Comparing columns 9 and 12, we find that conditioning on the achievement tests, and the other channels variables reduces the coefficient on athletics to 29% of what it was unconditional of achievement tests. The estimated coefficient on clubs is now negative, though statistically indistinguishable from zero. Comparing

[32] Barron, Ewing, and Waddell (2000) analyze the link between participation in high school athletics and labor market outcomes. They find that the link can be attributed to lower cost of effort of those who participate in athletics, or to a directly productive role of athletic in training youth for the labor market.

Columns 10 and 12 in Table 11, the degree to which conditioning on the test score explains the teen height premium is unchanged by the inclusion of controls for the other channels. In this sample, therefore, AFQT appears to be a sufficient statistic for the productive components of all of the channels measures.

Table 12 restricts attention to the sample that was actually in high school when height was recorded in 1981. In this sample, participation in high school sports and clubs appears to have an effect on wages and the teen height premium that is independent of what is measured by AFQT. Comparing the results in Columns 5 and 7, we find that, as in the entire sample, the estimated AFQT premium is 57 percent of what it was in the absence of conditioning on sports and clubs participation. Unlike in the entire sample, however, comparing columns 4 and 7 we find that conditioning on the other channels measures reduces the coefficient on athletics symmetrically. The coefficient on athletics conditional on AFQT is 55 percent of what it is unconditional of the achievement test score. Conditioning on the other channels diminishes the estimated coefficient on clubs to 29 percent of its unconditional estimate.

In this younger sample, AFQT is not a sufficient statistic for the productive component of participation in sports and clubs. Comparing Columns 4 and 5 of Table 12, the degree to which conditioning on the test score explains the teen height premium is essentially the same as that explained by conditioning on sports and clubs participation. Comparing Columns 5 and 7 in Table 12, the degree to which conditioning on the test score explains the teen height premium is here improved by the inclusion of controls for sports and clubs participation. Alone, the activities measures and test score each reduce the estimated teen height premium by 38 and 42 percent respectively (Table 12, Columns 4 and 5). Together, along with the other channels controls, these measures reduce the estimated teen height premium by 58 percent (Table 12, Columns 7).

The role played by achievement tests requires a careful interpretation. The AFQT, for example, reflects not only intellectual endowment but also education and other inputs that are the results of past experiences—in this case inputs that are correlated with social activities.[33] The effect on the estimated teen height premium of conditioning

[33] While it is plausible that achievement test scores reflect *some* component of intellectual endowments (these could be intelligence, persistence, social adaptability, etc.) the notion that achievement tests like the AFQT measure *only* pre-adolescent endowments has been discredited. It is known, for example, that the AFQT score increases with age and education. Hansen, Heckman and Mullen (2003), for example, find an important effect of education on AFQT scores (see also Neal and Johnson 1996). Our analysis has added to this evidence; as discussed before, while for individuals in the NLSY who take the AFQT later in life that score is a sufficient statistic (with respect to wages) for

on achievement test scores taken during or after adolescence is consistent, therefore, with a teen height premium that derives from adolescent experiences. The evidence in this paper indicates that the alternative hypothesis, that pre-adolescent intellectual endowments explain the teen height premium, is not very plausible for three reasons. First, we have seen in the NCDS that conditioning on earlier intelligence tests does not diminish the height premium, nor does teen height predict the scores of early intelligence tests (see Section 5.6). Second, since intellectual endowment is partly heritable, the argument developed in Section 5.3 casts doubt on the importance of that endowment as an explanation for the teen height premium. Third, we find no evidence that the teen height premium might reflect an early development premium (see Section 5.7). Overall, the evidence suggests that to the extent that achievement test scores help explain the teen height premium, this is not because they partly reflect pre-adolescent intellectual endowment, but rather because they also reflect schooling and other adolescent experiences.

Further evidence supporting this interpretation is provided by controlling for years of completed schooling. Without controlling for achievement test scores, the estimated teen height premium decreases by 20% - 23% in the UK and US data, respectively, when we control for years of completed schooling (see Table 11 Columns 4 and 11, and Table 12 Column 6). When we control for achievement tests, however, the teen height premium is unaffected by the inclusion of years of completed schooling (results not shown). This means that the ability of the years of schooling to explain the teen height premium is mediated by the achievement tests scores; and at most 40% of the impact of achievement test scores on the teen height premium is mediated by schooling. As our analysis in Section 5.3 suggests, the remaining 60% of the impact cannot be attributed by any pre-adolescent unobservable resource unless that resource is also not inheritable and more closely correlated with teen height than with all other heights. In sum, the evidence suggests that the achievement test scores reflect the same unobservable qualities that tall teens acquire during adolescence and are reflected in greater social interaction and greater schooling.

## 6.2 Summary: The All-Purpose Benefits of Teen Height

The picture emerging from the previous section is one in which the beneficial effects of being tall as an adolescent do not lead adolescents to choose a dramatically different

participation in high school sport and clubs, the same is not true when the AFQT is measured at adolescence (see below).

occupation. Thus, the beneficial effects of teen height do not seem to be especially complementary with any particular vocation path; instead, they are manifest in somewhat higher levels of schooling and achievement within the same vocation category, broadly defined. In this sense, the beneficial effects of teen height seem largely to accrue across the board, as a sort of all-purpose resource. This conclusion is consistent with our findings from the quantile regressions (see part 3 in section 4.3), which show that the beneficial effects of teen height are more or less equally manifested across income levels. We have seen that a natural candidate, self-esteem, does not seem to mediate the teen height premium. Instead, participation in sports and clubs seem to be important channels through which teen height affects adult wages.

# 7 Discussion

## 7.1 Not Statistical Discrimination

We have seen that the teen height premium is mediated by participation in social activities and, perhaps, by greater academic achievement The analysis above does not address the reasons shorter teens have a lower participation rate than taller teens in social activities or score lower on academic achievement tests. The lower rate may be due to obstacles created by the social or learning environment, stigma, etc., that raise the cost for shorter teens, or it may be that shorter teens have the same cost of participation and yet choose to invest less. Shorter teens might invest less in these activities, for example, if short *adults* have a lower return on investments in the activities. If this were the case, since shorter teens forecast themselves as short adults, it is rational of them to invest less relative to taller teens.

In this section we suggest that, to the extent that short adolescents participate less in social or AFQT-enhancing activities, it is not because they anticipate a lower return to these factors when adult. Again restricting attention to the NLSY data, and those who were 19 or younger in 1981, Table 13 presents estimates of the return to participation in social activities and to AFQT. Column 1 considers only those white male workers who were less than median height as adults; Column 2 performs the parallel analysis for those who were at least median height as adults. Conditioning only on age, family background and region, we find that the coefficients on social activities and AFQT do not differ significantly, or systematically between the two regressions. Among those who grew to less than median adult height the estimated coefficient on participation in athletics is 0.092, while for those who grew to at least median adult height the estimated coefficient is larger (0.145), though we can not reject the null hypothesis that the two coefficients are the same. The estimated coefficient on participation in clubs among the shorter adult group is actually larger (0.063) than that for the taller group (0.037), though again we cannot reject the null hypothesis that the coefficients are the same.[34] For the AFQT, the estimated coefficient is also larger for the shorter group (0.007) than for the taller group (0.005). Thus we find little evidence that the returns to investing in social or AFQT-enhancing activities when young are significantly or systematically different depending on whether one forecasts becoming a tall or short adult.

## 7.2 Parents-Children Correlation in Height

It is interesting to explore the correlation between the height of parents and the height of children for at least two reasons. First, if the correlation between parents' teen or adult height and the child's teen height is substantial, then the children of tall parents are advantaged in expectation on the labor market. This advantage could be magnified if parents are found to match assortatively by height (i.e., taller men tend to have children with taller women). Second, if the correlation is high then we might worry that the child's teen height proxies for the parents' height, and thus the estimated coefficient on the child's teen height may reflect some parental endowment that is not completely captured by our measures of resources.

While the NLSY does not report any measure of the subject's parent's height, the NCDS allows us to explore these issues. Unfortunately we do not have a measure of the NCDS parents' teen height, but we have a (self-reported) measure of the parents' adult height. We first compute the correlation coefficient between the parents' adult height and the child's teen height (see Table 14). The coefficient is 0.35 for fathers and 0.40 for mothers. Thus, a son of a tall couple enjoys a relatively large advantage, in expectation, due to his superior expected teen height. This high level of correlation reflects two components, the genetic one (tall parents generate tall children) and the matching one (tall men tend to have children with tall women[35]). In order to parse the relative contribution of the two effects, we compute the partial correlation coefficients between the son's teen height and the father's and mother's adult heights. For fathers, we find that holding constant mother's height, the correlation coefficient is 0.28. For

[34] Similar results hold when we analyze the entire sample of white male workers rather than only this younger subsample.

[35] In our data set, the correlation coefficient between the two parents' height is 0.26 (Table 14).

mothers, the corresponding coefficient is 0.34.

Given the high intergenerational correlation of height, a concern is that the estimated teen height premium may be capturing the benefits of having taller parents. The results in Table 7 indicate, however, that the estimate of the teen height coefficient is unaffected by the inclusion of his father's height (the same is true for mother's height, results omitted). This result suggests that the beneficial effects of teen height do not perpetuate across generations except through height itself.

## 7.3 Part-Time Workers, Blacks, and Females

In both the UK and US data, results that include both full and part-time, white male workers are qualitatively similar to those that restrict attention to full-time workers. In the UK, inclusion of part-time workers generates a somewhat larger point estimate of the coefficient on age-33 height, though it remains statistically indistinguishable from zero. In the US data, the estimates are quantitatively very similar to those that restrict attention to full-time workers only.

With respect to blacks, sample sizes are smaller for this group, and indeed nonexistent for the UK data. The estimates for black men in the US are, however, quite similar to those for white men, though less precisely estimated. The findings appear robust to outliers in the height, and growth and wage distributions (results not shown).

The findings for white women are different with respect to teen height. While we find an economically substantial, and sometimes statistically significant, adult height premium for white women in both the UK and the US, this height premium is not attributable to teen height. Indeed, point estimates of the teen height for women are negative, though not statistically distinguishable from zero. Unlike the male sample, in which the primacy of teen height is a robust finding, we find no prima facie evidence of a teen height premium among women. This finding must be qualified by the recognition that proper estimates of the wage offer functions for women should take account of the important labor market participation decision. The results described here ignore the selection issue. While far from conclusive, we view these results for females as suggesting that the relationship between productive endowments and the timing of physical development is different for males and females. Or, as plausible, in our view, the social-psychological returns to earlier development are different for males and females. Indeed, research in psychology indicates that earlier physical development is a hindrance for girls.[36]

[36] See, e.g., Brooks-Gunn, et al (1985); Ge, et al. (1996), and Graber (1997).

When considered in light of our results for males, the absence of a teen height premium for females reinforces the notion that the estimated teen height premium for males is not merely the result of a correlation between earlier physical development and omitted productive endowments. If the male teen height premium were simply a premium for resources associated with earlier development then we might expect to observe similar benefits to females who developed earlier.

# 8 Human Growth Hormone Replacement Therapy

Since 1985, pathologies resulting in short stature have been treated with injections of synthetic human growth hormone (HGH) replacement in the form of somatropin, the drug's generic name.[37] In addition, somatropin was legally prescribed by physicians "off label" to idiopathic short children, i.e., children who, despite having normal levels of HGH and therefore not being approved by the FDA for treatment, were of modest height relative to their age-peers. In July 2003, Humatrope (the name brand of somatropin from Eli Lilly) was approved by the FDA for use by idiopathic short patients, i.e., patients of short stature who display normal levels of HGH.[38] The FDA's approval goes some way towards endorsing the current practice of prescribing the drug "off label" to short children, and opens the door for the drug to be approved by health insurance companies and HMOs.

FDA approval cited two studies indicating important gains in height for children treated with sufficiently large doses of Humatrope. Table 15, taken from the FDA's Statistical Review and Evaluation of Humatrope (Mele 2003), reproduces the results of these two studies. The first study was a randomized, double-blind experiment with children aged 9-15 years. Subjects received either Humatrope or placebo three times weekly until they achieved their adult heights. After an average treatment of 4.4 years, mean final height of the Humatrope patients exceeded that of the placebo patients by approximately 1.5 inches. In the second study, subjects received one of three increasing doses of Humatrope six times per week. The average length of treatment was 6.5 years. Among the group receiving the highest dose, mean final height exceeded mean height

[37] Before 1985, the only source of HGH replacement was human cadavers. The limited supply allowed just 7,700 US patients to be treated between 1964 and 1985. Sources of synthetic HGH were sought after several recipients of HGH from cadavers were diagnosed with Creutzfeldt-Jakob Disease, a rare, fatal brain disorder closely related to mad cow disease in cattle.

[38] The approval means that the FDA recommends the drug for the treatment of children who are more than 2.25 standard deviations below the mean for age and sex, or the shortest 1.2% of children.

predicted at enrollment in the study by nearly three inches.[39]

Presumably reflecting the perceived efficacy of the treatment, the size of the market for somatropin is already very large. There are five U.S manufacturers of somatropin that have FDA approval, and each manufacturer designates its own brand name.[40] The total sales of the five brands currently on the market reached \$1.5 billion in 2001. For comparison, Prozac generated revenues of about \$2.6 billion in 2000.[41] It is likely that FDA approval will mean an even larger use of somatropin.

From an economists' viewpoint, somatropin is remarkable because it provides a previously unavailable means of control over a variable, teen height, that in our data sets is not under the control of individuals. As such, our data provide a unique opportunity to estimate the monetary returns from the previously unavailable medical option. In other words, we can ask: Is the benefit of treatment worth the cost?

Our analysis, which is exclusively concerned with the monetary benefits of teen height and ignores the psychological benefits, provides a back-of-the-envelope lower bound for the expected benefits of HGH treatment. It is important to keep in mind that the benefit we refer to is an individual benefit, i.e., the benefit that would accrue to an individual who, alone in society, has access to the treatment. If a large section of the population were able to increase their height then presumably the returns to increasing height would change, at least to extent that the benefits from height accrue from being tall relative to one's peers.

**Monetary Cost** Treatment with HGH varies depending on the weight of the patient. For a child who weighs 30 kilograms, the annual treatment at the highest dosage administered in the trials would cost approximately \$25,000 in 2003. The average length of treatment is 6.5 years.[42] A course of somatropin treatment is therefore currently quite expensive. The discounted monetary cost of undergoing the somatropin

[39] Many of the participants in these studies dropped out before achieving their final heights. In assessing the efficacy of Humatrope, Mele (2003) takes care to account for the potential biases introduced by these dropouts.

[40] Genotropin by Pharmacia & Upjohn Company, Humatrope by Eli Lilly, Norditropin by Genentech, Saizen and Serostim by Serono Laboratories.

[41] 2000 is the last year in which Prozac did not compete with generic equivalents. Revenues from Prozac fell precipitously after generic equivalents became available.

[42] In addition, there are non-trivial non-monetary costs to be factored in, as the treatment is administered daily by injection.

treatment for 6 years starting at age 10 equals

$$\sum_{t=1}^{6} \beta^{t} (\$25,000), \tag{5}$$

where $\beta$ denotes the discount factor.

**Monetary Benefit** Let $\rho$ denote the expected percent increase in adult wages from one more inch of teen height, $x$ be the average yearly wages earned over an individual's lifetime, and assume that the individual works from age 22 to 62. We then get an estimate of additional discounted lifetime earnings per inch of teen height computed at age 10 equal to

$$\sum_{t=1}^{40} \beta^{t+12} \rho x, \tag{6}$$

**Cost-Benefit Analysis** We will henceforth fix $\beta$ at 0.97. Regarding $\rho$, in our analysis of the relationship between teen height and wages we find that, after conditioning for background characteristics, the expected return from one more inch of teen height ranges between 1.9 and 2.7 percent of wages. If we substitute $\rho = 0.019$ in expression (6), multiply the expression by 3 (the treatment is expected to increase teen height by around 3 inches), equate to (5) and solve for $x$ we obtain $x = \$149,910$. This is the minimum average yearly wage that would justify investing in the somatropin treatment at age 10, given that an inch in teen height increases wages by 1.9 percent per year. If instead we assume that an inch in teen height increases wages by 2.7 percent per year, we would get an estimate of $x = \$105,500$.

This back-of-the-envelope calculation suggests that the monetary benefits of somatropin treatment are not incommensurate with the cost, and that a sizable fraction of the population might be willing to consider treatment purely on an economic basis. Of course, one can reasonably suppose that there are other benefits associated with greater height that are not pecuniary. These benefits would contribute to the use of the drug. At the same time, the drug must currently be injected, which is quite invasive. The drug companies are currently working on other delivery mechanisms (specifically, oral) which should lower these non-monetary costs.

Of course, our analysis needs to take into account that it is not children who make the decision to invest in somatropin treatment, but the parents, who do not directly enjoy their children's increased future earnings. This would tend to make treatment less prevalent. On the other hand, if insurance companies indeed come to cover somatropin at least for the very short (prior to approval, almost no insurance offered coverage of somatropin treatment for the idiopathic short), we should expect an increase in the diffusion of the treatment. Our analysis is also limited by the fact that it only applies to the return to height *given the current distribution of height* in the population. If somatropin treatment were to become so prevalent as to alter the distribution of heights, it is not clear that our estimates of the returns to height would still apply.

On balance, even taking into account these caveats, we view our back-of-the-envelope calculation as suggesting that somatropin treatments are in the economic self-interest of a sizable fraction of the population. When, moreover, we take into account the likely psychological benefits of height increase, we conclude that we should expect an even greater diffusion of somatropin treatment.

# 9 Conclusions

Labor market outcomes differ depending on a person's physical characteristics, and the resulting disparities have motivated a large body of research focussed on identifying the channels through which these disparities develop. In this paper we took up this agenda with respect to height. First, we found that the magnitude of wage differences associated with height is similar to that associated with racial differences; without correcting for education or occupation choice, we estimate the black-white wage gap to be approximately 15% among males in the NLSY. Depending on our specification, this estimated gap is approximated by the difference in wages for two men whose heights differ by six to eight inches. We then took advantage of a special feature of height relative to other bases of wage disparity, such as race and gender: height changes over time, so that a relatively tall 16-year old may turn out to be a relatively short adult, and vice versa. This time-variation allowed us to investigate the stage of development at which having the characteristic (in our case, being short) most strongly determines the wage disparity. We found that being relatively short through the teen years (as opposed to adulthood or early childhood) essentially determines the returns to height. We documented that the beneficial effects of teen height are not complementary with any particular vocation path; instead, they manifest themselves in a higher level of achievement in all vocation categories, broadly defined. We suggested that social effects might be an important channel for the emergence of the height premium. We found that teen height is predictably greater for sons of tall parents. This suggests that, similar to race-based disparities (but unlike gender-based disparities), there is an expected wage penalty incurred by the as-yet-unborn children of short parents. Finally, we used our estimates of the return to teen height to evaluate the monetary incentive to undertake a relatively new treatment that boosts teen height, human growth hormone therapy.

We emphasized in the introduction that we do not imply that the disparate treatment with respect to height is morally equivalent to racial discrimination—the differences in the history and psychology of these two phenomena are profound. Nevertheless, what is learned from height discrimination may help shed light on racial discrimination. There are, however, very significant differences between black-white wage differences and the wage advantage associated with being taller as an adolescent. For the case of racial differences, Neal and Johnson (1996) provide evidence that some of the disparity in typically unobservable skills can be predicted by differences in the family and school resources of the children. All of these can presumably be affected by public policy. The differences between the background characteristics of blacks and whites is in stark contrast to the differences - or lack thereof - in the background characteristics of youths who were tall or short. When compared to the black-white case, the schools and families of tall and short teens are essentially indistinguishable. As a consequence, it is far more difficult to attribute the effect of short stature as an adolescent on subsequent labor market outcomes to a deficiency in available resources. But if social exclusion and a feeling of 'not fitting in' has long-term, economically important consequences for shorter teens, one must consider the possibility that minority youths might suffer similarly. Obviously, a serious investigation of this is beyond the scope of this paper.

A number of questions pertaining to the height premium are left open by our analysis. Assuming that there are valuable skills that are acquired through participation in clubs and athletics, what precisely is acquired? Likely candidates are the interpersonal skills acquired through social interactions, social adaptability from working in groups, and discipline and motivation that result from participation. We also do not know that it is discrimination within athletics and other extracurricular activities that accounts for shorter teens' lower participation. It may be, for example, that earlier treatment has made these youths more sensitive to slights, and that as a result they withdraw from such interactions. More detailed data on the activities that youths engage in, and the job market consequences would permit a better understanding of the production process of generating social skills.

# A Robustness to Outliers

In this section we explore whether the finding of adult height's irrelevance is robust to the trimming of outliers in the US and UK data.

Tables A1 and A2 present the results of checks for sensitivity to outliers. In each of the specifications we trim both the top and the bottom 1% or 5% of the relevant sample. The point estimates from the UK data are both qualitatively and quantitatively robust to the omission of the 2% and 10% most extreme values of the age-33 height, and age-16 to 33 growth distributions. Standard errors are somewhat larger, but statistical significance at the 5% level is unchanged. The qualitative results from the US data are also robust to the omission of outliers, however the quantitative estimates are less stable than those from the UK data. In particular, omitting the extremes of the height and growth distributions generates larger estimates of the youth height premium and a point estimate of a (statistically insignificant) adult height penalty. In each data set, omitting the extremes of the wage distribution tends to diminish the estimated youth height premium, though it remains statistically and economically significant. The results from the median regressions (quantile 0.5 in Table B6), which are less sensitive to wage outliers than OLS, indicate that our basic results are both qualitatively and quantitatively robust to the exclusion of outliers in the wage distribution.

# B Primacy of Teen Height: Robustness to Non-Linear Specifications

**Nonlinear relationship between teen height and adult wages** Table B1 investigates whether the data suggest a quadratic relationship between teen height and adult wages. The point estimates of the coefficients in Table B1 are imprecisely estimated. When we consider the entire height distribution (including values in the top and bottom 1 percent), an F-test indicates that the non-linear specification improves the fit somewhat, though not nearly enough to obtain statistical significance of the test statistic. This is in line with the notion that a small increase in height should have a large effect on the wages of those who are quite short and a much smaller effect for individuals that are already very tall. When we exclude the top and bottom 1 percent, the F-test indicates that the non-linear specification explains almost the same amount of wage variation as the linear one (see Table B1). Table B1 also shows how the magnitudes of the estimated premia vary considerably when we trim the extremes of the height distribution, and are still imprecisely estimated. We regard the relative sensitivity of the non-linear specification to outliers as additional evidence that the quadratic specification is less reliable than the simple linear equation.[43] We also experimented with additional specifications up to a cubic term, with splines, and with dummy variables for height quantiles. The results (available on request from the authors) indicate that the height coefficients on the non-linear specifications are typically not precisely estimated, and we cannot nearly reject a null hypothesis of a linear specification.

**Non-linearities and the primacy of teen height** We address the concern that the primacy of teen height in explaining the height premium depends on the linear specification. We present an analysis in which teen and adult height appear *simultaneously and non-linearly* as explanatory variables. We allow teen and adult height to enter the wage equation more flexibly, in particular as 3rd or 4th order polynomials. The coefficients on each of the terms of the polynomial are estimated imprecisely, and are often difficult to interpret, but we can nevertheless approximate tests of the following two hypotheses:

$$E(w_i|\mathbf{X}_i, H_{i,\text{teen}}, H_{i,\text{adult}}) = E(w_i|\mathbf{X}_i, H_{i,\text{adult}}) \quad (7)$$

$$E(w_i|\mathbf{X}_i, H_{i,\text{teen}}, H_{i,\text{adult}}) = E(w_i|\mathbf{X}_i, H_{i,\text{teen}}), \quad (8)$$

where $\mathbf{X}_i$ is our usual vector of controls. More precisely, we ask whether in the polynomial specifications we can (1) reject a null hypothesis that all of the coefficients involving teen height terms are zero; and (2) reject a null hypothesis that all of the coefficients involving adult height are zero. Tables B2-B3 and Tables B4-B5 explore these two hypotheses for the UK and US data, respectively.

UK DATA

For the UK data, the first panel of Table B2 presents the results of F-tests for the joint significance of age-16 or age-33 coefficients for different polynomial specifications

[43] To the extent that one takes the coefficients of the non-linear specification seriously, in the UK data a quadratic specification (Table B1) indicates a decreasing age16 height premium, with a 2.9% per inch premium for the short (5'4"), a 2.4% per inch premium for the average (5'7"), and a 1.9% per inch premium for the tall (5'10"). In the US data, the youth height premia from a specification that includes only a linear and quadratic term in 1981 height is very imprecisely estimated (Table B1). Taking the point estimates seriously, they suggest large height premium for the short (3.6% per inch at 5'4") that declines for taller youth (2.1% per inch at 5'10"). Trimming the extremes of the height distribution this specification suggests smaller premia for the short, and very little decline with height.

of the height wage relationship. These results indicate that our basic result is not an artifact of the linear specification, and indeed that the linear specification is capturing nearly all of the wage variation explained by variation in heights. We reject, at least at the 5% level, null hypothesis (1) that the age-16 coefficients equal zero in all but one specification. We are never close to rejecting, even at the 10% level, null hypothesis (2) for any of the specifications. Spline specifications (panel 2) that permit increasing flexibility in the wage-height relationship produce similar results: we reject null hypothesis (1) in each specification, and do not nearly reject null hypothesis (2) in any specification. When the top and bottom 1% of the adult height distribution is trimmed from the sample (Table B3), adult height appears even less relevant as a predictor of adult wages.

US DATA

Tables B4 and B5 present parallel analyses for the US data. Like in the UK data, when we consider the entire sample (Table B4) for each the polynomial specifications we reject null hypothesis (1) at the 5% level. Unlike the UK data, however, we often reject null hypothesis (2) as well. [Only hypothesis (1) is rejected with the spline specifications]. Unlike the linear specifications, then, these results begin to suggest that in the US data both youth and adult height may explain adult wage variation (though, youth height is explaining more of the variation). As Table B5 shows, however, this result is not robust to the exclusion of extreme height outliers. When the top and bottom 1% of the adult height distribution is trimmed from the sample, the importance of adult height for explaining variation in adult wages disappears; we cannot reject null hypothesis (2) for any specification.

**Non-linearities as a function of wage levels** We address possible non-linearities in the effect of height depending on the wage level by means of quantile regressions. Table B6 presents results from quantile regressions for each of the first nine deciles for both the UK and the US data.

UK DATA

In the UK data we find limited evidence of non-linearities in the effect of height depending on the position in the conditional wage distribution. Quantile regressions of wages on teen height alone indicate that the height premium is fairly uniform throughout the wage distribution, varying between 1.5 and 2.5 percent per inch (top panel of Table B6). Quantile regressions on teen and adult height indicate a teen height premium between 0.9 and 2.6 percent per inch (top panel of Table B6). Except at the sixth and seventh deciles, the estimated teen height premium is statistically different from zero at least at the 5% level. None of the estimated adult height premia/penalties is statistically different from zero at the 5% level except at the seventh decile. In both sets of quantile regressions the relationship between the premium and wages does not appear to vary monotonically with the quantile.

US DATA

In the US data we find similar, limited evidence of non-linearities. Quantile regressions of wages on teen height alone indicate that the height premium is fairly uniform throughout the wage distribution, varying between 1.9 and 2.6 percent per inch (see bottom panel of Table B6). Quantile regressions on teen and adult height indicate that youth height premia range from 1.4 percent to 5 percent per inch. Relative to the UK data, the youth height premia are estimated less precisely. Of the nine coefficients only 4 are statistically different from zero at the 5% level. The estimated adult height premia/penalties, range from -3. percent per inch at the ninth decile to 1.1 percent per inch at the third decile, but again there appears to be no monotonic relationship between wages and the size of the height premium. Only the large adult height penalty at the ninth decile is statistically different from zero at the 10% level.

# C Teen Height not a Superior Correlate of Endowments Relative to Other Heights

A concern is that heights at various ages may be proxying for the same unobservable, productive factor or factors. If this were the case, and if teen height were best correlated with these unobservable factors, we might obtain the estimates we have, but our interpretation would be misleading. Table C1 presents the results from regressions of various endowments on heights at different ages. Generally, we find that each of the heights is only very weakly associated with these observable endowments, and we find no evidence that teen height is a systematically better predictor of these endowments.

UK DATA

Age 16 height is not a systematically better predictor of age 7 test scores than any other heights. Conditioning on age 16 and age 33 height alone, age 16 height is a superior partial correlate of age 7 math scores but a weaker partial correlate of age 7 reading schools. Though in each case the estimated marginal effects are very small. When we condition on age 11 height as well, age 16 is no longer a superior predictor of the math score either. There is essentially no relationship between health at age 7 and height at ages 11, 16 or 33; the estimated relationships are again quite small. Among the three heights, age 16 does not seem to be a relatively good predictor of earlier health. Age 16 height is the best predictor of mother's years of schooling, though this relationship is not statistically significant. Results are similar for having a mother working as a professional, which in the UK seems to be correlated with lower economic outcomes. When we condition only on age 33 and age 16 height, we find a significant positive relationship between age 16 height and father's schooling; and the point estimate is larger than that associated with age33 height. This result is not robust, however, to the inclusion of childhood heights. When we also condition on age 11 and age 7 heights, age 16 height is the worst predictor of father's years of schooling. We find similar results for our measure of father's occupation, and for the number of siblings, though none of the heights in either of the father's occupation specifications is a statistically significant predictor of this measure.

US DATA

In the US data, only adult height is a significant predictor of any of our endowment measures. For each endowment measure except siblings, the point estimates indicate adult height is an approximately as good or better predictor of endowment measures as youth height.

# D Teen Height not a Proxy for Weight

It is interesting to control for weight, as the negative wage effects of being short as an adolescent might be due to shorter than average adolescents being overweight. Alternatively, we could imagine that short adolescents are overly thin, and that the teen height premium would simply mask a fitness premium. If these scenarios applied, controlling for weight should either decrease or increase, respectively, the coefficient on teen height. Table D1, however, shows that the coefficients on teen height are not substantially affected by the introduction of teen and adult weight as controls. We conclude that the teen height effect does not reflect a weight effect.

# References


[1] Averett, Susan, and Sanders Korenman (1996). "The Economic Reality of The Beauty Myth." *Journal of Human Resources* **31(2)**:304-330.

[2] Behrman, Jere, and Mark Rosenzweig (2001). "The Returns to Increasing Body Weight." Working Paper. University of Pennsylvania.

[3] Barron, John M., Bradley T. Ewing, and Glen R. Waddell "The Effects of High School Athletic Participation on Education and Labor Market Outcomes." *The Review of Economics and Statistics*, August 2000, 82(3), pp. 409-21.

[4] Blau, Francine and Lawrence Kahn. "Swimming Upstream: Trends in the Gender Wage Differential in the 1980s." *Journal of Labor Economics* 15(1, Part 1) 1997: 1-42.

[5] Cain, Glen (1986). "The Economic Analysis of Labor Market Discrimination: A Survey." In *Handbook of Labor Economics*, vol. 1, Eds. Orley Ashenfelter and Richard Layard. Amsterdam: North-Holland.

[6] Case, Ann, Angela Fertig and Christina Paxson (2003). "From Cradle to Grave? The Lasting Impact of Childhood Health and Circumstance." *Princeton University Center for Health and Wellbeing.* Working Paper.

[7] Damon, Albert, Howard Stoudt, and Ross McFarland *The Human Body in Equipment Design*, Harvard University Press, Cambridge MA 1966.

[8] Cawley, John, Karen Conneely, James Heckman, Edward Vytlacil "Cognitive Ability, Wages, and Meritocracy." In Devlin, Bernie, Stephen Fienberg, Daniel Resnick, Kathryn Roeder, eds. *Intelligence, Genes, and Success: Scientists Respond to the Bell Curve* Springer Verlag, New York, 1997.

[9] Elias, P., and David Blanchflower (1988). "The Occupations , Earnings and Work Histories of Young Adults – Who Gets the Good Jobs?" London England, Department of Employment.

[10] Frieze, Irene H., Olson, Josephine E., and Deborah C. Good (1990). "Perceived and Actual Discrimination in the Salaries of Male and Female Managers. *Journal of Applied Social Psychology* **20(1)**:46-67.

[11] Hamermesh, Daniel S., and Jeffrey E. Biddle (1994). "Beauty and the Labor Market." *American Economic Review* **84(5)**:1174-1194.

[12] Hansen, Karsten, James Heckman and Kathleen Mullen (2003). "The Effect of Schooling and Ability on Achievement Test Scores." Forthcoming, *Journal of Econometrics*.

[13] Heckman, James J. (2000) "Policies to Foster Human Capital." *Research in Economics* **54(1)**:3-56.

[14] _____ (1998). "Detecting Discrimination." *Journal of Economic Perspectives* **12(2)**:101-116.

[15] Loh, Eng Seng (1993). "The Economic Effects of Physical Appearance." *Social Science Quarterly.* **74 (2)**:420-38.

[16] Martel, Leslie F. and Henry B. Biller (1987). *Stature and Stigma: The Biopsychosocial Development of Short Males*. Lexington Books. Lexington, Mass.

[17] Mele, Joy (2003). "Statistical Review and Evaluation: Clinical Studies" NDA #19-640/SE1-033. Food and Drug Administration Center for Drug Evaluation and Research Report (on file with the authors).

[18] National Center for Health Statistics, US. Department of Health and Human Services (1987). "Anthropometric Reference Data and Prevalence of Overweight: United States, 1976-80." DHHS Publication No., (PHS) 87-1688.

[19] Neal, Derek, and William Johnson (1996). "The Role of Premarket Factors in Black-White Wage Differences." *Journal of Political Economy.* 104 (5):869-95.

[20] Pope AW, McHale SM, Craighead WE. *Self-Esteem Enhancement With Children and Adolescents*. New York, NY: Pergamon Press; 1988:2.

[21] Roche, Alex F. *Growth, Maturation, and Body Composition: the Fels Longitudinal Study 1929-1991.* Cambridge University Press, Cambridge UK 1992.

[22] Sargent, James D., and David G. Blanchflower (1994). "Obesity and Stature in Adolescence and Earnings in Young Adulthood: Analysis of a British Birth Cohort." *Archives of Pediatric and Adolescent Medicine* **148:**681-687.

[23] Steckel, Richard H. "Stature and the Standard of Living," *Journal of Economic Literature*, 33(4), Dec. 1995, pp. 1903-40.

| Table 1: Distribution of Height and Change in Height, White Males of the NCDS and NLSY | | | | | |
|---|---|---|---|---|---|
| **Britain -- NCDS**[1] | Height age 16 | Height age 33 | Δ age 16-33 | Height age 11 | Height age 7 |
| Mean | 67.28 | 69.81 | 2.54 | 56.94 | 48.50 |
| Median | 67.32 | 69.69 | 1.97 | 57.01 | 49.02 |
| Standard deviation | 3.01 | 2.62 | 1.99 | 2.63 | 2.16 |
| 25th percentile | 65.35 | 68.11 | 1.18 | 55.00 | 47.01 |
| 75th percentile | 69.29 | 71.65 | 3.54 | 58.74 | 50.00 |
| N | 1772 | 1772 | 1772 | 1684 | 1702 |
| | | | | | |
| **US -- NLSY**[2] | Height 1981 | Height 1985 | Δ 1981-85 | | |
| **Entire subsample** | | | | | |
| Mean | 70.41 | 70.69 | 0.28 | | |
| Median | 70 | 70 | 0 | | |
| Standard deviation | 2.85 | 2.77 | 1.44 | | |
| 25th percentile | 68 | 68 | 0 | | |
| 75th percentile | 73 | 73 | 1 | | |
| N | 2603 | 2603 | 2603 | | |
| **Those with Δ height >0** | | | | | |
| Mean | 69.54 | 71.22 | 1.68 | | |
| Median | 69 | 70 | 1 | | |
| Standard deviation | 3.09 | 2.73 | 1.51 | | |
| 25th percentile | 68 | 69 | 1 | | |
| 75th percentile | 72 | 74 | 3 | | |
| N | 618 | 618 | 618 | | |
| | | | | | |
| ***US Measured*** **heights for white males, ages 18-24, from the National Health and Nutrition Examination Survey 1976-1980**[3] | | | | | |
| Mean | 69.8 | | | | |
| Median | 69.7 | | | | |
| Standard deviation | 2.8 | | | | |
| 25th percentile | 67.9 | | | | |
| 75th percentile | 71.6 | | | | |
| N | 846 | | | | |

[1] The subsample consists only of the full-time, white male workers in the NCDS for whom there is a measure of height at age 33 and 16, and information on wages and family background. The sample is further restricted when we consider those with data on age-11 and age-7 height.
[2] The subsample consists only of the white male respondents to the NLSY for whom there is a measure of height at in 1985 and 1981, and information on family background. The NLSY's oversample of poor whites is excluded.
[3] Source: National Center for Health Statistics (1987), Table 13.

| **Table 2: Summary Statistics, White Males by Adult Height** | | | | |
|---|---|---|---|---|
| | **Britain -- NCDS** | | **US -- NLSY** | |
| | Adult Height Median or Below | Adult Height Above Median | Adult Height Median or Below | Adult Height Above Median |
| **Adult Characteristics:** | Mean (SE) | Mean (SE) | Mean (SE) | Mean (SE) |
| Teen height (inches) | 65.55* (0.08) | 69.12 (0.08) | 68.19* (0.07) | 72.23 (0.06) |
| Adult height (inches) | 67.80* (0.05) | 71.96 (0.06) | 68.25* (0.06) | 72.70 (0.05) |
| Age | 33.0 (…) | 33.0 (…) | 35.34 (0.07) | 35.31 (0.07) |
| Ln(wage per hour) | 1.99* (0.02) | 2.10 (0.02) | 2.58* (0.03) | 2.68 (0.02) |
| Years of completed schooling | 11.04* (0.03) | 11.19 (0.03) | 13.38* (0.09) | 13.79 (0.08) |
| Ever married (%) | 79.89 (1.34) | 80.09 (1.38) | 78.84* (1.42) | 83.15 (1.17) |
| Divorced or separated (%)[1] | 16.74 (1.40) | 14.97 (1.38) | 20.71* (1.59) | 15.90 (1.26) |
| **Family Background:** | | | | |
| Mother's years of schooling | 10.43* (0.05) | 10.57 (0.06) | 11.83* (0.08) | 12.29 (0.07) |
| Mother skilled/ professional (%) | 56.67 (1.64) | 54.66 (1.70) | 7.20* (0.85) | 9.98 (0.89) |
| Father's years of schooling | 11.29* (0.04) | 11.50 (0.06) | 12.14* (0.10) | 12.66 (0.10) |
| Father skilled/ professional (%) | 79.87 (1.32) | 82.87 (1.29) | 12.78 (1.09) | 14.92 (1.06) |
| Number of siblings | 3.14* (0.05) | 2.89 (0.05) | 2.99 (0.06) | 2.91 (0.06) |
| N | 914 | 858 | 931 | 1132 |

* Indicates means are statistically different at the 5% confidence level.
[1] Conditional on having been married.

**Notes:** Teen height is height recorded at age 16 (NCDS), or in 1981 (NLSY), adult height is height recorded at age 33 (NCDS), or 1985 (NLSY). Log wages are in measured in 1991 pounds (NCDS) and 1996 dollars (NLSY), and are for full-time workers only. For the NCDS, years of schooling equal the age at which the respondent (or parent) left school minus five. Parents are identified as skilled (professional) if they work in a professional and or skilled, non-manual (NCDS) or professional/managerial (NLSY) occupation.

| Table 3: OLS Estimates Ln(wage) Equation for Adult, White Male Workers, NCDS and NLSY | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | **Britain -- NCDS** | | | | **US -- NLSY** | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| **Covariates** | | | | | | | | |
| Adult height (inches) | 0.027 (0.0053) | 0.022 (0.0052) | 0.025 (0.0065) | 0.024 (0.0066) | 0.025 (0.0062) | 0.018 (0.0060) | 0.018 (0.0079) | 0.019 (0.0077) |
| Age | … | … | … | … | 0.028 (0.0066) | 0.027 (0.0065) | 0.023 (0.0088) | 0.022 (0.0089) |
| | | | | | | | | |
| Mother's years of schooling | | 0.016 (0.0104) | 0.019 (0.0123) | 0.011 (0.0126) | | 0.025 (0.0092) | 0.030 (0.0117) | 0.030 (0.0118) |
| Mother skilled/ professional (%) | | -0.080 (0.0357) | -0.057 (0.0424) | -0.048 (0.0436) | | 0.019 (0.0608) | 0.092 (0.0769) | 0.087 (0.0766) |
| Father's years of schooling | | 0.008 (0.0086) | 0.010 (0.0099) | 0.004 (0.0097) | | 0.030 (0.0065) | 0.026 (0.0083) | 0.025 (0.0084) |
| Father skilled/ professional (%) | | 0.135 (0.0467) | 0.112 (0.0557) | 0.095 (0.0576) | | 0.050 (0.0459) | 0.100 (0.0570) | 0.101 (0.0597) |
| Number of siblings | | -0.033 (0.0084) | -0.025 (0.0101) | -0.021 (0.0102) | | -0.023 (0.0077) | -0.023 (0.0094) | -0.023 (0.0095) |
| | | | | | | | | |
| Student/teacher ratio | | | | -0.0005 (0.0073) | | | | -0.002 (0.0057) |
| Disadvantaged student ratio | | | | -0.0016 (0.0008) | | | | -0.001 (0.0011) |
| Dropout rate | | | | -0.0015 (0.0080) | | | | -0.001 (0.0010) |
| Teacher turnover rate | | | | -0.0004 (0.0022) | | | | -0.006 (0.0029) |
| | | | | | | | | |
| N | 1772 | 1772 | 1257 | 1257 | 1577 | 1577 | 943 | 943 |
| Adjusted $R^2$ | 0.032 | 0.047 | 0.045 | 0.051 | 0.031 | 0.092 | 0.104 | 0.108 |
| F-Statistic (K,N-K-1) | 9.99 | 10.25 | 8.00 | 7.45 | 9.86 | 15.52 | 10.34 | 9.04 |

Robust standard errors are in parentheses.
**Notes:** See notes for Table 2. Sample consists only of white male, full-time workers. Each specification includes controls for region and a constant term, results omitted. In the NCDS the disadvantaged student ratio is defined as the percentage of the school's population with a father in a non-skilled, manual occupation. Starting in the academic year beginning in the fall of 1972, the age of mandatory schooling was raised to 16 years in Britain. The dropout rate used here is the percentage of students who, in the 1971-72 academic year, left the respondent's school at or before their earliest legal opportunity.

| Table 4: Summary Statistics, Somewhat Taller White Males, By Relative Teen Height | | | | |
|---|---|---|---|---|
| | Britain -- NCDS | | US -- NLSY | |
| | Adult Height 69.7-72.8 inches | | Adult Height 71-74 inches | |
| | Teen Height Median or Below | Teen Height Above Median | Teen Height Median or Below | Teen Height Above Median |
| **Adult Characteristics:** | Mean (SE) | Mean (SE) | Mean (SE) | Mean (SE) |
| Teen height (inches) | 65.92** (0.11) | 69.52 (0.06) | 69.13** (0.17) | 72.24 (0.04) |
| Adult height (inches) | 70.88** (0.06) | 71.36 (0.04) | 71.50** (0.07) | 72.37 (0.03) |
| Age | 33.0 (…) | 33.0 (…) | 34.05** (0.18) | 35.48 (0.07) |
| Ln(wage per hour) | 1.98** (0.03) | 2.13 (0.03) | 2.57* (0.06) | 2.67 (0.02) |
| Years of completed schooling | 11.07 (0.07) | 11.20 (0.05) | 11.79 (0.50) | 12.02 (0.20) |
| Ever married (%) | 78.64 (2.96) | 81.08 (1.86) | 71.67** (4.11) | 84.66 (1.29) |
| Divorced or separated (%)[2] | 14.57 (2.87) | 16.11 (1.94) | 23.26** (4.56) | 14.76 (1.38) |
| **Family Background:** | | | | |
| Mother's years of schooling | 10.47 (0.07) | 10.49 (0.11) | 12.38 (0.22) | 12.26 (0.07) |
| Mother skilled/ professional (%) | 56.12 (3.54) | 55.00 (1.81) | 12.21 (2.86) | 9.23 (0.99) |
| Father's years of schooling | 11.30 (0.086) | 11.48 (0.087) | 12.40 (0.32) | 12.63 (0.11) |
| Father skilled/ professional (%) | 78.06* (2.96) | 83.48 (1.73) | 10.69 (2.70) | 15.77 (1.25) |
| Number of siblings | 3.22** (0.12) | 2.82 (0.06) | 2.76 (0.17) | 2.94 (0.06) |
| N | 196 | 460 | 131 | 856 |

** Indicates means are statistically different at the 5% confidence level.
* Indicates means are statistically different at the 10% confidence level.
[1]Conditional on having been married. **Notes:** See Table 2.

| Table 5: OLS Estimates Ln(wage) Equation for Adult, White Male Workers, NCDS and NLSY | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Britain -- NCDS | | | | US -- NLSY | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| **Covariates** | | | | | | | | |
| Adult height (inches) | 0.004 (0.0074) | 0.005 (0.0073) | 0.005 (0.0087) | 0.004 (0.0086) | 0.002 (0.0096) | -0.004 (0.0091) | -0.006 (0.0113) | -0.003 (0.0114) |
| Youth height (inches) | 0.026 (0.0066) | 0.021 (0.0066) | 0.023 (0.0077) | 0.023 (0.0077) | 0.027 (0.0095) | 0.026 (0.0090) | 0.027 (0.0108) | 0.025 (0.0111) |
| Age | … | … | … | … | 0.024 (0.0067) | 0.023 (0.0065) | 0.020 (0.0088) | 0.018 (0.0089) |
| | | | | | | | | |
| Mother's years of schooling | | 0.016 (0.0104) | 0.020 (0.0123) | 0.012 (0.0126) | | 0.023 (0.0092) | 0.028 (0.0116) | 0.028 (0.0116) |
| Mother skilled/ professional (%) | | -0.074 (0.0356) | -0.048 (0.0420) | -0.040 (0.0431) | | 0.024 (0.0606) | 0.101 (0.0767) | 0.095 (0.0765) |
| Father's years of schooling | | 0.007 (0.0087) | 0.008 (0.0100) | 0.003 (0.0097) | | 0.030 (0.0065) | 0.026 (0.0083) | 0.025 (0.0083) |
| Father skilled/ professional (%) | | 0.130 (0.0465) | 0.106 (0.0553) | 0.089 (0.0571) | | 0.052 (0.0458) | 0.103 (0.0593) | 0.103 (0.0595) |
| Number of siblings | | -0.029 (0.0084) | -0.022 (0.0104) | -0.017 (0.0102) | | -0.023 (0.0077) | -0.023 (0.0094) | -0.023 (0.0094) |
| | | | | | | | | |
| Student/teacher ratio | | | | -0.0013 (0.0073) | | | | -0.001 (0.0057) |
| Disadvantaged student ratio | | | | -0.0016 (0.0008) | | | | -0.001 (0.0012) |
| Dropout rate | | | | -0.0014 (0.0008) | | | | -0.001 (0.0010) |
| Teacher turnover rate | | | | -0.0005 (0.0022) | | | | -0.006 (0.0029) |
| | | | | | | | | |
| N | 1772 | 1772 | 1257 | 1257 | 1577 | 1577 | 943 | 943 |
| Adjusted $R^2$ | 0.037 | 0.049 | 0.043 | 0.054 | 0.034 | 0.094 | 0.107 | 0.110 |
| F-Statistic (K,N-K-1) | 11.47 | 10.97 | 8.54 | 7.91 | 8.82 | 14.31 | 9.63 | 8.82 |

Robust standard errors are in parentheses. Equations include controls for region and a constant term, results omitted
**Notes:** See notes for Tables 2 and 3.

**Table 6: OLS Estimates Ln(wage) Equation for White Male Workers of Britain's NCDS, at Age 33, Controlling for Prior Physical Development**

| | (1) | (2) |
|---|---|---|
| **Covariates** | | |
| Height age 33 (inches) | 0.003 (0.0076) | 0.002 (0.0079) |
| Height age 16 (inches) | 0.021 (0.0069) | 0.019 (0.0084) |
| Height age 11 (inches) | | 0.003 (0.0107) |
| Height age 7 (inches) | | 0.003 (0.0109) |
| | | |
| Mother's years of schooling | 0.016 (0.0109) | 0.016 (0.0109) |
| Mother skilled worker | -0.095 (0.0368) | -0.095 (0.0370) |
| Father's years of schooling | 0.007 (0.0089) | 0.007 (0.0090) |
| Father skilled worker | 0.132 (0.0491) | 0.132 (0.0491) |
| Number of siblings | -0.031 (0.0090) | -0.030 (-0.0090) |
| | | |
| N | 1617 | 1617 |
| Adjusted $R^2$ | 0.056 | 0.055 |
| F-Statistic (K,N-K-1) | 10.02 | 8.92 |

Robust standard errors are in parentheses.

Sample consists only of full time workers. Equations include controls for region and a constant term, results omitted

**Notes:** See notes for Tables 2 and 3.

| Table 7: OLS Estimates Ln(wage) Equation for White Male Workers of Britain's NCDS, at Age 33, Controlling for Father's Adult Height | | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| **Covariates** | | | | |
| Height age 33 (inches) | 0.003 (0.0075) | 0.002 (0.0081) | 0.003 (0.0075) | 0.003 (0.0079) |
| Height age 16 (inches) | 0.025 (0.0065) | 0.025 (0.0065) | 0.020 (0.0065) | 0.020 (0.0065) |
| Father's Adult Height (inches) | | 0.003 (0.0055) | | -0.001 (0.0056) |
| Other controls for family background | no | no | yes | yes |
| N | 1713 | 1713 | 1713 | 1713 |
| Adjusted $R^2$ | 0.039 | 0.038 | 0.051 | 0.051 |
| F-Statistic (K,N-K-1) | 11.26 | 10.36 | 10.77 | 10.14 |

Robust standard errors are in parentheses.
Sample consists only of full time workers. Equations include controls for region and a constant term, results omitted. Other controls for family background in columns (3) and (4) include education and occupation of parents and number of siblings.
**Notes:** See notes for Tables 2 and 3.

**Table 8: OLS Estimates Ln(wage) Equation for Adult, White Male Workers, NCDS and NLSY Controlling for Measures of Past Health**

| | Britain -- NCDS | | | | | US -- NLSY | | | |
|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | | (5) | (6) | (7) | (8) |
| **Covariates** | | | | | **Covariates** | | | | |
| Height age 33 (inches) | 0.007 (0.0079) | 0.006 (0.078) | 0.007 (0.0079) | 0.006 (0.0078) | Height in 1985 (inches) | -0.006 (0.0092) | -0.006 (0.0092) | -0.006 (0.0092) | -0.006 (0.0092) |
| Height age 16 (inches) | 0.020 (0.0073) | 0.020 (0.0073) | 0.020 (0.0073) | 0.020 (0.0073) | Height in 1981 (inches) | 0.028 (0.0092) | 0.027 (0.0093) | 0.028 (0.0092) | 0.027 (0.0092) |
| Age | ** | ** | ** | ** | Age | 0.023 (0.0066) | 0.023 (0.0066) | 0.023 (0.0066) | 0.024 (0.0066) |
| # health conditions age 16 | | -0.052 (0.0262) | | -0.049 (0.0276) | Health limits kind of work 79 | | -0.156 (0.0788) | | -0.216 (0.100) |
| # health conditions age 7 | | | -0.046 (0.0406) | -0.029 (0.0434) | Health limits amount of work 79 | | | 0.026 (0.086) | 0.158 (0.120) |
| | | | | | | | | | |
| Mother's yrs of Schooling | 0.012 (0.0114) | 0.013 (0.0114) | 0.013 (0.0114) | 0.013 (0.0114) | Mother's yrs of Schooling | 0.024 (0.0093) | 0.024 (0.0092) | 0.024 (0.0092) | 0.024 (0.0092) |
| Mother skilled worker | -0.076 (0.0391) | -0.073 (0.0391) | -0.076 (0.0391) | -0.073 (0.0391) | Mother skilled worker | 0.022 (0.0609) | 0.023 (0.0606) | 0.022 (0.0610) | 0.024 (0.0606) |
| Father's yrs of Schooling | 0.006 (0.0091) | 0.006 (0.0091) | 0.006 (0.0091) | 0.006 (0.0091) | Father's yrs of schooling | 0.030 (0.0066) | 0.030 (0.0066) | 0.030 (0.0066) | 0.030 (0.0066) |
| Father skilled worker | 0.136 (0.0513) | 0.132 (0.0513) | 0.136 (0.0513) | 0.132 (0.0514) | Father skilled worker | 0.052 (0.0046) | 0.055 (0.0046) | 0.052 (0.0046) | 0.055 (0.0046) |
| Number of siblings | -0.035 (0.0083) | -0.035 (0.0084) | -0.035 (0.0083) | -0.035 (0.0084) | Number of siblings | -0.023 (0.0078) | -0.022 (0.0078) | -0.023 (0.0078) | -0.022 (0.0078) |
| | | | | | | | | | |
| N | 1546 | 1546 | 1546 | 1546 | N | 1558 | 1558 | 1558 | 1558 |
| Adjusted $R^2$ | 0.049 | 0.049 | 0.093 | 0.049 | Adjusted $R^2$ | 0.094 | 0.095 | 0.093 | 0.096 |
| F-Stat. (K,N-K-1) | 10.26 | 9.79 | 9.66 | 9.26 | F-Stat. (K,N-K-1) | 15.68 | 13.39 | 12.86 | 12.50 |

Robust standard errors are in parentheses.
**Notes:** See notes for Tables 2 and 3. In the NCDS, at age 7 the possible health conditions include slight, moderate, or severe: general motor handicap, disfiguring condition, mental retardation, emotional maladjustment, head or neck abnormality, upper limb abnormality, lower limb abnormality, spine abnormality, respiratory system problem, alimentary system problem, urogenital system problem, heart condition, blood abnormality, skin condition, epilepsy, other central nervous system condition, diabetes, and any other condition. The possible conditions at age 16 are the same except: disfiguring condition is replaced by general physical abnormality, emotional maladjustment is replaced by emotional/behavioral problem, and eye, hearing and speech defects are also included. In the NLSY, respondents were asked in 1979 a) whether their health limited the kind of work they could do and b) whether their health limited the amount of work they could do.

| **Table 9: OLS Estimates Ln(wage) Equation for White Male Workers of Britain's NCDS, at Age 33, Controlling for Age-7 Academic Test Scores** | | |
|---|---|---|
| | (1) | (2) |
| **Covariates** | | |
| Height age 33 (inches) | 0.004<br>(0.0074) | -0.000<br>(0.0073) |
| Height age 16 (inches) | 0.021<br>(0.0066) | 0.019<br>(0.0065) |
| Age 7 reading test score (0-30) | | 0.015<br>(0.0025) |
| Age 7 math test score (0-10) | | 0.024<br>(0.0080) |
| | | |
| Mother's years of schooling | 0.015<br>(0.0106) | 0.000<br>(0.0108) |
| Mother skilled worker | -0.078<br>(0.0362) | -0.072<br>(0.0354) |
| Father's years of schooling | 0.006<br>(0.0089) | 0.004<br>(0.0086) |
| Father skilled worker | 0.139<br>(0.0473) | 0.102<br>(0.0461) |
| Number of siblings | -0.029<br>(0.0086) | -0.018<br>(0.0086) |
| | | |
| N | 1726 | 1726 |
| Adjusted $R^2$ | 0.049 | 0.093 |
| F-Statistic (K,N-K-1) | 10.98 | 17.56 |

Robust standard errors are in parentheses.
Sample consists only of full time workers. Equations include controls for region and a constant term, results omitted
**Notes:** See notes for Tables 2 and 3.

**Table 10: OLS Estimates Ln(wage) Equation for White Male Workers of both the NLSY and Britain's NCDS, Youth Height Level vs. Percentage of Adult Height**

| | Britain -- NCDS | | US -- NLSY | | |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| **Covariates** | | | | | |
| Youth height (inches) | 0.024 (0.0047) | 0.026 (0.0055) | 0.022 (0.0059) | 0.022 (0.0063) | 0.022 (0.0063) |
| 100*(Youth height)/ (adult height) | | -0.004 (0.0053) | | 0.002 (0.0062) | 0.002 (0.0062) |
| Born in 1st quarter | | | | | 0.049 (0.0430) |
| Born in 2nd quarter | | | | | 0.008 (0.0450) |
| Born in 3rd quarter | | | | | 0.0007 (0.0426) |
| | | | | | |
| Mother's years of schooling | 0.016 (0.0104) | 0.016 (0.0104) | 0.024 (0.0092) | 0.023 (0.0092) | 0.023 (0.0093) |
| Mother skilled worker | -0.074 (0.0356) | -0.074 (0.0356) | 0.023 (0.0604) | 0.024 (0.0606) | 0.023 (0.0607) |
| Father's years of schooling | 0.007 (0.0087) | 0.007 (0.0087) | 0.030 (0.0064) | 0.030 (0.0065) | 0.031 (0.0065) |
| Father skilled worker | 0.130 (0.0465) | 0.130 (0.0465) | 0.052 (0.0458) | 0.052 (0.0458) | 0.053 (0.0457) |
| Number of siblings | -0.029 (0.0084) | -0.029 (0.0084) | -0.023 (0.0077) | -0.023 (0.0077) | -0.022 (0.0077) |
| | | | | | |
| N | 1772 | 1772 | 1577 | 1577 | 1577 |
| Adjusted $R^2$ | 0.050 | 0.049 | 0.095 | 0.094 | 0.094 |
| F-Statistic (K,N-K-1) | 11.72 | 10.97 | 15.71 | 14.29 | 15.92 |

Robust standard errors are in parentheses.
Sample consists only of full time workers. Equations include controls for region and a constant term, results omitted
**Notes:** See notes for Tables 2 and 3.

**Table 11: OLS Estimates Ln(wage) Equation for Adult, White Male Workers, NCDS and NLSY Controlling for Other Outcome Measures**

| | Britain -- NCDS | | | | | | US -- NLSY | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| **Covariates** | | | | | | | | | | | | | |
| Height age 33 (inches) | 0.006<br>(0.0084) | 0.005<br>(0.0082) | 0.005<br>(0.0081) | 0.005<br>(0.0082) | 0.005<br>(0.0080) | Height in 1985 (inches) | -0.004<br>(0.0098) | -0.004<br>(0.0097) | -0.003<br>(0.0097) | -0.001<br>(0.0097) | 0.003<br>(0.0093) | -0.002<br>(0.0095) | 0.002<br>(0.0093) |
| Height age 16 (inches) | 0.020<br>(0.0071) | 0.018<br>(0.0068) | 0.010<br>(0.0068) | 0.016<br>(0.0068) | 0.011<br>(0.0065) | Height in 1981 (inches) | 0.023<br>(0.0095) | 0.022<br>(0.0094) | 0.022<br>(0.0095) | 0.018<br>(0.0095) | 0.011<br>(0.0092) | 0.018<br>(0.0093) | 0.011<br>(0.0092) |
| Average height of occupation age 33 | | 0.352<br>(0.0320) | | | 0.243<br>(0.0328) | Average height of occupation in 1996 | | 0.070<br>(0.0486) | | | | | 0.020<br>(0.0449) |
| Reading test score age 16 | | | 0.018<br>(0.0029) | | 0.013<br>(0.0029) | Self-esteem in 1980 | | | 0.028<br>(0.0061) | | | | 0.014<br>(0.0059) |
| Math test score age 16 | | | 0.014<br>(0.0025) | | 0.005<br>(0.0027) | Participation in HS athletics | | | | 0.117<br>(0.0325) | | | 0.034<br>(0.0315) |
| Yrs of completed schooling age 33 | | | | 0.179<br>(0.0191) | 0.084<br>(0.0241) | No. of HS clubs participated in | | | | 0.051<br>(0.0179) | | | -0.011<br>(0.0179) |
| | | | | | | AFQT percentile in 1980 | | | | | 0.007<br>(0.0007) | | 0.004<br>(0.0008) |
| | | | | | | Yrs of completed schooling in 1996 | | | | | | 0.078<br>(0.0076) | 0.049<br>(0.0087) |
| | | | | | | | | | | | | | |
| N | 1376 | 1376 | 1376 | 1376 | 1376 | N | 1485 | 1485 | 1485 | 1485 | 1485 | 1485 | 1485 |
| Adjusted $R^2$ | 0.045 | 0.125 | 0.123 | 0.109 | 0.172 | Adjusted $R^2$ | 0.094 | 0.096 | 0.106 | 0.109 | 0.162 | 0.164 | 0.185 |
| F-Stat. (K,N-K-1) | 8.56 | 18.96 | 13.46 | 13.76 | 22.64 | F-Stat. (K,N-K-1) | 13.85 | 12.86 | 14.80 | 13.31 | 22.30 | 23.42 | 18.49 |

Robust standard errors are in parentheses.

**Notes:** Sample consists only of full time workers. Equations include controls for region and a constant term, and in the US age, results omitted. Average height of occupation is defined as the sample average teen height of all the workers in the respondent's occupational category, excluding the respondent himself. Occupations are divided into 17 categories in the UK and 12 in the US. In the UK, years of completed schooling are defined as the age at which the respondent left school minus five. See notes for Tables 2 and 3.

| **Table 12: OLS Estimates Ln(wage) Equation for Adult, White Male Workers Less than 19 Years old in 1981, NLSY Controlling for Other Outcome Measures** | | | | | | | |
|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
| **Covariates** | | | | | | | |
| Height in 1985 (inches) | -0.016<br>(0.0127) | -0.016<br>(0.00125) | -0.015<br>(0.00126) | -0.009<br>(0.0124) | -0.009<br>(0.0117) | -0.010<br>(0.0122) | -0.006<br>(0.0117) |
| Height in 1981 (inches) | 0.026<br>(0.0118) | 0.025<br>(0.0113) | 0.024<br>(0.0118) | 0.016<br>(0.0120) | 0.015<br>(0.0111) | 0.020<br>(0.0115) | 0.011<br>(0.0112) |
| Average height of occupation in 1996 | | 0.199<br>(0.0662) | | | | | 0.127<br>(0.0604) |
| Self-esteem in 1980 | | | 0.023<br>(0.0087) | | | | 0.009<br>(0.0085) |
| Participation in HS athletics | | | | 0.165<br>(0.0461) | | | 0.090<br>(0.0445) |
| No of HS clubs participated in | | | | 0.097<br>(0.0231) | | | 0.028<br>(0.0234) |
| AFQT percentile in 1980 | | | | | 0.007<br>(0.0010) | | 0.004<br>(0.0010) |
| Yrs of completed schooling in 1996 | | | | | | 0.078<br>(0.0011) | 0.045<br>(0.0122) |
| | | | | | | | |
| N | 627 | 627 | 627 | 627 | 627 | 627 | 627 |
| Adjusted $R^2$ | 0.114 | 0.136 | 0.138 | 0.153 | 0.192 | 0.194 | 0.229 |
| F-Stat. (K,N-K-1) | 7.12 | 7.62 | 6.86 | 8.94 | 11.52 | 11.12 | 9.43 |

Robust standard errors are in parentheses. Sample consists only of white, male full time workers. Equation includes controls for family background, region and a constant term, results omitted.

| **Table 13: OLS Estimates of the Returns to Social Activities and AFQT, by Adult Height, for White Male Workers of the NLSY, Younger Sample Only** | | |
|---|---|---|
| | | |
| **Covariates** | Adult Height Below Median | Adult Height Median or Above |
| Age | -0.006<br>(0.0386) | -0.014<br>(0.0330) |
| Participation in HS athletics | 0.092<br>(0.0696) | 0.145<br>(0.0605) |
| No. of HS clubs participated in | 0.063<br>(0.0348) | 0.037<br>(0.0326) |
| AFQT percentile in 1980 | 0.007<br>(0.0015) | 0.005<br>(0.0012) |
| | | |
| N | 284 | 343 |
| Adjusted $R^2$ | 0.237 | 0.151 |
| F-Statistic (K,N-K-1) | 6.78 | 6.76 |

Robust standard errors are in parentheses.
Sample consists only of full time workers, ages 19 or younger in 1981. Equation includes controls for family background, region and a constant term, results omitted.

**Table 14: Correlations Between Parents' and Son's Heights and Between Parents' Heights: for White Male Workers in the NCDS**

| | Father's Adult Height | | Mother's Adult Height | |
|---|---|---|---|---|
| | Correlation | Partial Correlation | Correlation | Partial Correlation |
| Son's height age 16 | 0.353 | 0.280 | 0.397 | 0.337 |
| Son's height age 33 | 0.406 | 0.332 | 0.470 | 0.411 |
| Mother's adult height | 0.264 | ** | 1.000 | ** |
| | | | | |
| N | 1705 | | 1705 | |

Robust standard errors are in parentheses.
Sample consists only of white male full time workers. Partial correlation coefficients present the correlation with relevant parent's height holding constant the other parent's height.

| **Table 15. Final Height Means Adjusting for Baseline Predicted and Target Height for Participants in Two Randomized Trials of Human Growth Hormone Treatment for Children with Idiopathic Short Stature** | | | | | |
|---|---|---|---|---|---|
| | Study 1 | | Study 2 | | |
| | Placebo | Humatrope 0.22 | Humatrope 0.24 | Humatrope 0.24/0.37 | Humatrope 0.37 |
| FH – Baseline PH | | | | | |
| Standard devs | -0.18 | +0.33 | +0.83 | +1.10 | +1.29 |
| cm | -0.9 | +2.3 | +5.4 | +6.5 | +7.3 |
| (95% CI for cm) | (-3.3, 1.5) | (0.6, 3.9) | (2.9, 8.0) | (3.9, 9.1) | (4.7, 9.8) |
| FH – Target Height | | | | | |
| Standard devs | -0.96 | -0.68 | -0.46 | -0.64 | -0.26 |
| cm | -7.0 | -4.8 | -3.3 | -4.8 | -1.9 |
| (95% CI for cm) | (11.3, -2.6) | (-7.6, -2.0) | (-7.7, 1.0) | (-9.2, -0.4) | (-6.3, 2.4) |

**Source Mele (2003)**

**Table A1: Tests of Sensitivity to Height, Growth, and Wage Outliers for the Wages of White Males in the NCDS**

| | whole sample | | omit top and bottom 1% of height distribution | | omit top and bottom 5% of height distribution | | omit top and bottom 1% of Δ 16-33 distribution | | omit top and bottom 5% of Δ 16-33 distribution | | omit top and bottom 1% of wage distribution | | omit top and bottom 5% of wage distribution | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Height 33 | 0.005 | 0.002 | 0.003 | 0.001 | 0.009 | 0.008 | 0.003 | 0.000 | 0.004 | 0.004 | 0.004 | 0.000 | 0.004 | 0.001 |
| | (0.0073) | (0.0079) | (0.0082) | (0.0089) | (0.0093) | (0.0098) | (0.0085) | (0.0092) | (0.0099) | (0.0107) | (0.0061) | (0.0069) | (0.0048) | (0.0052) |
| Height 16 | 0.021 | 0.019 | 0.021 | 0.018 | 0.021 | 0.018 | 0.022 | 0.020 | 0.023 | 0.019 | 0.019 | 0.015 | 0.014 | 0.009 |
| | (0.0066) | (0.0084) | (0.0068) | (0.0087) | (0.0071) | (0.0091) | (0.0078) | (0.0096) | (0.0099) | (0.0110) | (0.0053) | (0.0066) | (0.0043) | (0.0055) |
| Height 11 | | 0.003 | | 0.003 | | 0.004 | | 0.001 | | 0.006 | | 0.008 | | 0.008 |
| | | (0.0107) | | (0.0109) | | (0.0113) | | (0.0108) | | (0.0108) | | (0.0087) | | (0.0069) |
| Height 7 | | 0.003 | | 0.005 | | 0.003 | | 0.005 | | -0.000 | | -0.001 | | -0.000 |
| | | (0.0109) | | (0.0118) | | (0.0120) | | (0.0110) | | (0.0115) | | (0.0095) | | (0.0073) |
| N | 1772 | 1617 | 1726 | 1574 | 1619 | 1480 | 1735 | 1584 | 1599 | 1449 | 1737 | 1585 | 1595 | 1459 |

Dependent variable is ln(wages) at age 33. Sample is restricted to full-time white male workers. Specification includes controls for family background, region, and a constant term, results not presented. Robust standard errors are in parentheses.

**Table A2: Tests of Sensitivity to Height, Growth, and Wage Outliers for the Wages of White Males in the NLSY**

| | whole sample | omit top and bottom 1% of height distribution | omit top and bottom 5% of height distribution | omit top and bottom 1% of Δ 16-33 distribution | omit top and bottom 5% of Δ 16-33 distribution | omit top and bottom 1% of wage distribution | omit top and bottom 5% of wage distribution |
|---|---|---|---|---|---|---|---|
| 1985 height | -0.004 | -0.009 | -0.014 | -0.013 | -0.040 | -0.001 | -0.007 |
| | (0.0091) | (0.0113) | (0.0122) | (0.0158) | (0.0217) | (0.0081) | (0.0071) |
| 1981 height | 0.026 | 0.029 | 0.032 | 0.035 | 0.058 | 0.021 | 0.018 |
| | (0.0090) | (0.0108) | (0.0113) | (0.0162) | (0.0218) | (0.0079) | (0.0068) |
| N | 1577 | 1548 | 1359 | 1550 | 1421 | 1545 | 1417 |

Dependent variable is ln(wages) in 1996. Sample restricted to full-time white male workers. Includes controls for age, family background, region and a constant term, results not presented. Robust standard errors in parentheses.

**Table B1: Quadratic Specifications of the Relationship Between Teen Height and Adult Wages**
**White Male, Full-time workers**

## Britain -- NCDS

| | whole sample | | trim top and bottom 1% of adult height | |
|---|---|---|---|---|
| height16 | 0.024 | 0.1344 | 0.023 | 0.0969 |
| | (0.0047) | (0.1024) | (0.0050) | (0.1234) |
| $(\text{height16})^2$ | | -0.0008 | | -0.0006 |
| | | (0.0008) | | (0.0009) |
| N | 1772 | 1772 | 1726 | 1726 |
| root MSE | 0.5914 | 0.5915 | 0.5964 | 0.5965 |
| RSS | 614.2030 | 613.9717 | 608.2143 | 608.1318 |
| Adj. R2 | 0.0511 | 0.0509 | 0.0466 | 0.0462 |

F-statistic for null of linear model

| whole sample | trimmed sample |
|---|---|
| 0.661 | 0.232 |

| height at 16 | predicted marginal effects: whole sample | predicted marginal effects: trimmed sample |
|---|---|---|
| 62 | 0.0318 | 0.0285 |
| 63 | 0.0302 | 0.0274 |
| 64 | 0.0285 | 0.0263 |
| 65 | 0.0269 | 0.0252 |
| 66 | 0.0252 | 0.0241 |
| 67 | 0.0236 | 0.0230 |
| 68 | 0.0219 | 0.0219 |
| 69 | 0.0202 | 0.0208 |
| 70 | 0.0186 | 0.0197 |
| 71 | 0.0169 | 0.0186 |
| 72 | 0.0153 | 0.0175 |
| 73 | 0.0136 | 0.0164 |
| 74 | 0.0120 | 0.0153 |
| 75 | 0.0103 | 0.0142 |
| 76 | 0.0087 | 0.0131 |

## US -- NLSY

| | whole sample | | trim top and bottom 1% of adult height | |
|---|---|---|---|---|
| height81 | 0.0220 | 0.1947 | 0.0219 | 0.0481 |
| | (0.0059) | (0.1995) | (0.0057) | (0.2126) |
| $(\text{height81})^2$ | | -0.0012 | | -0.0002 |
| | | (0.0014) | | (0.0015) |
| N | 1577 | 1577 | 1548 | 1548 |
| root MSE | 0.5968 | 0.5967 | 0.5916 | 0.5918 |
| RSS | 557.6885 | 557.2788 | 537.9655 | 537.9591 |
| Adj. R2 | 0.0949 | 0.0950 | 0.0974 | 0.0968 |

F-statistic for null of linear model

| whole sample | trimmed sample |
|---|---|
| 1.151 | 0.018 |

| height in 1981 | predicted marginal effects: whole sample | predicted marginal effects: trimmed |
|---|---|---|
| 62 | 0.0411 | 0.0248 |
| 63 | 0.0386 | 0.0245 |
| 64 | 0.0361 | 0.0241 |
| 65 | 0.0336 | 0.0237 |
| 66 | 0.0312 | 0.0233 |
| 67 | 0.0287 | 0.0230 |
| 68 | 0.0262 | 0.0226 |
| 69 | 0.0237 | 0.0222 |
| 70 | 0.0213 | 0.0218 |
| 71 | 0.0188 | 0.0215 |
| 72 | 0.0163 | 0.0211 |
| 73 | 0.0138 | 0.0207 |
| 74 | 0.0114 | 0.0203 |
| 75 | 0.0089 | 0.0200 |
| 76 | 0.0064 | 0.0196 |

Dependent variable is ln(wages) at age 33, in 1996, in the UK and US data, respectively. Sample is restricted to full-time white male workers. Specification includes controls for family background, region and a constant term, results not presented. Robust standard errors are in parentheses.

| Table B2: F-tests of Non-linear Specifications of the Relationship Between Height and Age-33 Wages, NCDS White Male, Full-Time Workers | | | | | | | |
|---|---|---|---|---|---|---|---|
| Specification | Unrestricted RSS | Restricted RSS h16 coefs=0 | Restricted RSS h33 coefs=0 | Fstat for null: h16 coefs=0 | Fstat for null: h33 coefs=0 | Approximate 5% critical value | Approximate 10% critical value |
| $y_i=\beta_1 h33 + \beta_2 h16 + \gamma X_i + \varepsilon_i$ | 614.086 | 616.883 | 614.203 | 7.99 | 0.33 | 3.84 | 2.71 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h16^2 + \gamma X_i + \varepsilon_i$ | 613.447 | 616.276 | 613.972 | 4.04 | 0.75 | 3.00 | 2.30 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \gamma X_i + \varepsilon_i$ | 613.477 | 616.276 | | 8.00 | | 3.84 | 2.71 |
| | 613.477 | | 614.203 | | 1.04 | 3.00 | 2.30 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta 3 h16^2 + \gamma X_i + \varepsilon_i$ | 613.737 | 616.883 | | 4.50 | | 3.00 | 2.30 |
| | 613.737 | | 613.972 | | 0.67 | 3.84 | 2.71 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h16^2 + \beta_5 h16*h33 + \gamma X_i + \varepsilon_i$ | 613.033 | 616.276 | 613.972 | 3.09 | 0.89 | 2.60 | 2.08 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h16^2 + \beta_5 h33^3 + \beta_6 h16^3 + \gamma X_i + \varepsilon_i$ | 613.287 | 616.100 | 613.774 | 2.68 | 0.46 | 2.60 | 2.08 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h33^3 + \gamma X_i + \varepsilon_i$ | 613.333 | 616.100 | | 7.91 | | 3.84 | 2.71 |
| | 613.333 | | 614.203 | | 0.83 | 2.60 | 2.08 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h16^2 + \beta_5 h33^3 + \beta_6 h33^4 + \gamma X_i + \varepsilon_i$ | 613.256 | 616.077 | | 4.03 | | 3.00 | 2.30 |
| | 613.256 | | 613.972 | | 0.51 | 2.37 | 1.94 |
| two piece spline | 613.626 | 616.442 | 614.153 | 4.02 | 0.75 | 3.00 | 2.30 |
| three piece spline | 613.266 | 616.415 | 613.632 | 3.00 | 0.35 | 2.60 | 2.08 |
| four piece spline | 612.838 | 616.308 | 613.244 | 2.48 | 0.29 | 2.37 | 1.94 |
| five piece spline | 609.540 | 615.239 | 611.170 | 3.27 | 0.93 | 2.21 | 1.85 |

Dependent variable is ln(wages) at age 33. Sample is restricted to full-time white male workers. Specification includes controls for family background and region.

**Table B3: F-tests of Non-linear Specifications of the Relationship Between Height and Age-33 Wages, NCDS White Male, Full-time Workers Omitting Top and Bottom 1% of Age 33 Height Distribution**

| Specification | Unrestricted RSS | Restricted RSS h16 coefs=0 | Restricted RSS h33 coefs=0 | Fstat for null: h16 coefs=0 | Fstat for null: h33 coefs=0 | Approximate 5% critical value | Approximate 10% critical value |
|---|---|---|---|---|---|---|---|
| $y_i=\beta_1 h33 + \beta_2 h16 + \gamma X_i + \varepsilon_i$ | 608.160 | 610.938 | 608.214 | 7.81 | 0.15 | 3.84 | 2.71 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h16^2+\gamma X_i + \varepsilon_i$ | 607.955 | 610.749 | 608.132 | 3.92 | 0.25 | 3.00 | 2.30 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \gamma X_i + \varepsilon_i$ | 607.980 | 610.749 | | 7.78 | | 3.84 | 2.71 |
| | 607.980 | | 608.214 | | 0.33 | 3.00 | 2.30 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta 3h16^2+\gamma X_i + \varepsilon_i$ | 608.029 | 610.938 | | 4.09 | | 3.00 | 2.30 |
| | 608.029 | | 608.132 | | 0.29 | 3.84 | 2.71 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h16^2+ \beta_5 h16*h33+ \gamma X_i + \varepsilon_i$ | 607.571 | 610.749 | 608.132 | 2.97 | 0.53 | 2.60 | 2.08 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h16^2+ \beta_5 h33^3 + \beta_6 h16^3+ \gamma X_i + \varepsilon_i$ | 607.054 | 610.084 | 607.915 | 2.84 | 0.81 | 2.60 | 2.08 |
| $y_i=\beta_1 h33 + \beta_2 h16 + \beta_3 ht33^2 + \beta_4 h33^3 + \gamma X_i + \varepsilon_i$ | 607.101 | 610.084 | | 8.39 | | 3.84 | 2.71 |
| | 607.101 | | 608.214 | | 1.04 | 2.60 | 2.08 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2+ \beta_5 h85^3 + \beta_6 h85^4+ \gamma X_i + \varepsilon_i$ cannot be estimated | | | | | | | |
| two piece spline | 608.002 | 610.794 | 608.209 | 3.92 | 0.29 | 3.00 | 2.30 |
| three piece spline | 607.617 | 610.781 | 607.701 | 2.96 | 0.08 | 2.60 | 2.08 |
| four piece spline | 607.151 | 610.619 | 607.378 | 2.43 | 0.16 | 2.37 | 1.94 |
| five piece spline | 603.540 | 609.262 | 605.331 | 3.23 | 1.01 | 2.21 | 1.85 |

Dependent variable is ln(wages) at age 33. Sample is restricted to full-time white male workers. Specification includes controls for family background and region.

**Table B4: F-tests of Non-linear Specifications of the Relationship Between Height and 1996 Wages, NLSY White Male, Full-time workers**

| Specification | Unrestricted RSS | Restricted RSS h81 coefs=0 | Restricted RSS h85 coefs=0 | Fstat for null: h81 coefs=0 | Fstat for null: h85 coefs=0 | Approximate 5% critical value | Approximate 10% critical value |
|---|---|---|---|---|---|---|---|
| $y_i=\beta_1 h85 + \beta_2 h81 + \gamma X_i + \varepsilon_i$ | 557.633 | 559.616 | 557.689 | 5.57 | 0.16 | 3.84 | 2.71 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2 + \gamma X_i + \varepsilon_i$ | 554.772 | 557.727 | 557.279 | 4.16 | 3.53 | 3.00 | 2.30 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \gamma X_i + \varepsilon_i$ | 555.373 | 557.727 | | 6.63 | | 3.84 | 2.71 |
| | 555.373 | | 557.689 | | 3.26 | 3.00 | 2.30 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta 3h81^2 + \gamma X_i + \varepsilon_i$ | 557.272 | 559.616 | | 3.29 | | 3.00 | 2.30 |
| | 557.272 | | 557.279 | | 0.02 | 3.84 | 2.71 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2 + \beta_5 h81*h85 + \gamma X_i + \varepsilon_i$ | 554.771 | 557.727 | 557.279 | 2.77 | 2.35 | 2.60 | 2.08 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2 + \beta_5 h85^3 + \beta_6 h81^3 + \gamma X_i + \varepsilon_i$ | 554.771 | 557.682 | 556.882 | 2.73 | 1.98 | 2.60 | 2.08 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h85^3 + \gamma X_i + \varepsilon_i$ | 555.372 | 557.682 | | 6.50 | | 3.84 | 2.71 |
| | 555.372 | | 557.689 | | 2.17 | 2.60 | 2.08 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81*h85 + \beta_5 h85^3 + \gamma X_i + \varepsilon_i$ | 555.166 | 557.682 | | 3.54 | | 3.00 | 2.30 |
| | 555.166 | | 557.689 | | 1.77 | 2.37 | 1.94 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2 + \beta_5 h85^3 + \beta_6 h85^4 + \gamma X_i + \varepsilon_i$ | 554.634 | 557.498 | | 4.03 | | 3.00 | 2.30 |
| | 554.634 | | 557.279 | | 1.86 | 2.37 | 1.94 |
| two piece spline | 555.698 | 558.027 | 556.796 | 3.27 | 1.54 | 3.00 | 2.30 |
| three piece spline | 555.765 | 558.287 | 557.147 | 2.36 | 1.29 | 2.60 | 2.08 |
| four piece spline | 554.940 | 557.896 | 556.678 | 2.08 | 1.22 | 2.37 | 1.94 |
| five piece spline | 554.190 | 557.110 | 556.989 | 1.64 | 1.57 | 2.21 | 1.85 |

Dependent variable is ln(wages) in 1996. Sample is restricted to full-time white male workers. Specification includes controls for age, family background ,region and a constant term.

**Table B5: F-tests of Non-linear Specifications of the Relationship Between Height and 1996 Wages, NLSY White Male, Full-time Workers Omitting Top and Bottom 1% of 1985 Height Distribution**

| Specification | Unrestricted RSS | Restricted RSS h81 coefs=0 | Restricted RSS h85 coefs=0 | Fstat for null: h81 coefs=0 | Fstat for null: h85 coefs=0 | Approximate 5% critical value | Approximate 10% critical value |
|---|---|---|---|---|---|---|---|
| $y_i=\beta_1 h85 + \beta_2 h81 + \gamma X_i + \varepsilon_i$ | 537.802 | 539.791 | 537.965 | 5.68 | 0.47 | 3.84 | 2.71 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2+\gamma X_i + \varepsilon_i$ | 537.618 | 539.737 | 537.959 | 3.02 | 0.49 | 3.00 | 2.30 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \gamma X_i + \varepsilon_i$ | 537.743 | 539.737 | | 5.69 | | 3.84 | 2.71 |
| | 537.743 | | 537.965 | | 0.32 | 3.00 | 2.30 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta 3h81^2+\gamma X_i + \varepsilon_i$ | 537.799 | 539.791 | | 2.84 | | 3.00 | 2.30 |
| | 537.799 | | 537.959 | | 0.46 | 3.84 | 2.71 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2+ \beta_5 h81*h85+ \gamma X_i + \varepsilon_i$ | 537.572 | 539.737 | 537.959 | 2.06 | 0.37 | 2.60 | 2.08 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2+ \beta_5 h85^3 + \beta_6 h81^3+ \gamma X_i + \varepsilon_i$ | 537.139 | 539.724 | 537.864 | 2.46 | 0.69 | 2.60 | 2.08 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h85^3 + \gamma X_i + \varepsilon_i$ | 537.704 | 539.724 | | 5.76 | | 3.84 | 2.71 |
| | 537.704 | | 537.965 | | 0.25 | 2.60 | 2.08 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81*h85+ \beta_5 h85^3 + \gamma X_i + \varepsilon_i$ | 537.570 | 539.724 | | 3.07 | | 3.00 | 2.30 |
| | 537.570 | | 537.965 | | 0.28 | 2.37 | 1.94 |
| $y_i=\beta_1 h85 + \beta_2 h81 + \beta_3 ht85^2 + \beta_4 h81^2+ \beta_5 h85^3 + \beta_6 h85^4+ \gamma X_i + \varepsilon_i$ cannot be estimated | | | | | | | |
| two piece spline | 537.568 | 539.577 | 537.814 | 2.87 | 0.35 | 3.00 | 2.30 |
| three piece spline | 537.592 | 539.632 | 537.936 | 1.94 | 0.33 | 2.60 | 2.08 |
| four piece spline | 536.921 | 539.269 | 537.353 | 1.67 | 0.31 | 2.37 | 1.94 |
| five piece spline | 536.085 | 538.813 | 537.786 | 1.56 | 0.97 | 2.21 | 1.85 |

Dependent variable is ln(wages) in 1996. Sample is restricted to full-time white male workers. Specification includes controls for age, family background, region and a constant term.

**Table B6: Quantile Ln(Wage) Regression Coefficients for White, Male, Full-time Workers in the NCDS and NLSY**

**Britain -- NCDS**

quantile

| | 0.1 | | 0.2 | | 0.3 | | 0.4 | | 0.5 | | 0.6 | | 0.7 | | 0.8 | | 0.9 | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| height33 | | -0.003 | | -0.005 | | -0.001 | | 0.005 | | 0.008 | | 0.013 | | 0.016 | | 0.009 | | 0.008 |
| | | (0.0083) | | (0.0092) | | (0.0081) | | (0.0082) | | (0.0077) | | (0.0083) | | (0.0074) | | (0.0077) | | (0.0107) |
| height16 | 0.015 | 0.017 | 0.020 | 0.024 | 0.025 | 0.026 | 0.022 | 0.018 | 0.022 | 0.017 | 0.020 | 0.011 | 0.019 | 0.009 | 0.021 | 0.018 | 0.022 | 0.023 |
| | (0.0046) | (0.0073) | (0.0052) | (0.0082) | (0.0043) | (0.0071) | (0.0051) | (0.0072) | (0.0048) | (0.0068) | (0.0041) | (0.0074) | (0.0038) | (0.0065) | (0.0047) | (0.0070) | (0.0054) | (0.0097) |
| N | 1772 | | 1772 | | 1772 | | 1772 | | 1772 | | 1772 | | 1772 | | 1772 | | 1772 | |

**US -- NLSY**

quantile

| | 0.1 | | 0.2 | | 0.3 | | 0.4 | | 0.5 | | 0.6 | | 0.7 | | 0.8 | | 0.9 | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1985 | | -0.001 | | 0.002 | | 0.011 | | 0.008 | | -0.003 | | -0.005 | | -0.013 | | -0.006 | | -0.033 |
| height | | (0.0189) | | (0.0125) | | (0.0121) | | (0.0127) | | (0.0125) | | (0.0131) | | (0.0112) | | (0.0134) | | (0.0153) |
| 1981 | 0.020 | 0.021 | 0.026 | 0.025 | 0.024 | 0.014 | 0.024 | 0.017 | 0.022 | 0.024 | 0.022 | 0.027 | 0.022 | 0.034 | 0.019 | 0.025 | 0.025 | 0.050 |
| height | (0.0109) | (0.0171) | (0.0073) | (0.0117) | (0.0064) | (0.0117) | (0.0066) | (0.0124) | (0.0057) | (0.0123) | (0.0066) | (0.0130) | (0.0057) | (0.0114) | (0.0074) | (0.0140) | (0.0102) | (0.0152) |
| N | 1577 | | 1577 | | 1577 | | 1577 | | 1577 | | 1577 | | 1577 | | 1577 | | 1577 | |

Dependent variable is ln(wages) at age 33, in 1996, in the UK and US data, respectively. Sample is restricted to full-time white male workers. Specification includes controls for family background, region, and a constant term, results not presented. Robust standard errors are in parentheses.

**Table C1: Conditional Relationships Between Heights at Various Ages and Endowments, White, Male, Full-time Workers**

**Britain -- NCDS**

| | Math test score age 7 | | Reading test score age 7 | | Number of health conditions age 7 | | Mother's years of completed schooling | | Mother in a skilled profession | | Father's years of completed schooling | | Father in a skilled profession | | Number of silblings | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| height33 | 0.055 | 0.043 | 0.212 | 0.182 | 0.001 | 0.002 | 0.015 | 0.007 | 0.004 | 0.000 | 0.027 | 0.019 | 0.001 | 0.000 | 0.023 | 0.046 |
| | (0.0335) | (0.0357) | (0.0961) | (0.1007) | (0.0051) | (0.0055) | (0.0223) | (0.0254) | (0.0068) | (0.0075) | (0.0205) | (0.0245) | (0.0051) | (0.0056) | (0.0215) | (0.0236) |
| height16 | 0.067 | 0.041 | 0.099 | 0.026 | -0.002 | 0.001 | 0.031 | 0.041 | -0.010 | -0.016 | 0.038 | 0.007 | 0.008 | 0.003 | -0.099 | -0.024 |
| | (0.0292) | (0.0383) | (0.0851) | (0.1103) | (0.0049) | (0.0057) | (0.0174) | (0.0217) | (0.0059) | (0.0079) | (0.0157) | (0.0231) | (0.0047) | (0.0062) | (0.0199) | (0.0259) |
| height11 | | 0.046 | | 0.128 | | -0.006 | | -0.001 | | 0.019 | | 0.054 | | 0.006 | | -0.055 |
| | | (0.0417) | | (0.1117) | | (0.0057) | | (0.0287) | | (0.0099) | | (0.0344) | | (0.0077) | | (0.0282) |
| height07 | | ** | | ** | | ** | | -0.001 | | -0.008 | | 0.005 | | 0.004 | | -0.087 |
| | | | | | | | | (0.0307) | | (0.0108) | | (0.0313) | | (0.0088) | | (0.0324) |
| | | | | | | | | | | | | | | | | |
| N | 1729 | 1645 | 1743 | 1658 | 1679 | 1597 | 1772 | 1617 | 1772 | 1617 | 1772 | 1617 | 1772 | 1617 | 1772 | 1617 |
| R-squared | 0.0178 | 0.0183 | 0.0154 | 0.0160 | 0.0003 | 0.0012 | 0.0069 | 0.0078 | 0.0022 | 0.0041 | 0.0115 | 0.0148 | 0.0042 | 0.0063 | 0.0289 | 0.0406 |

**US -- NLSY**

| | Health limits kind of work '79 | Health limits amount of work '79 | Mother's years of completed schooling | Mother in a skilled profession | Father's years of completed schooling | Father in a skilled profession | Number of silblings |
|---|---|---|---|---|---|---|---|
| 1985 height | 0.002 | 0.004 | 0.054 | 0.010 | 0.130 | 0.007 | -0.023 |
| | (0.0034) | (0.0030) | (0.0455) | (0.0048) | (0.0565) | (0.0051) | (0.0353) |
| 1981 height | -0.004 | -0.004 | 0.056 | -0.004 | -0.001 | -0.004 | 0.011 |
| | (0.0037) | (0.0036) | (0.0440) | (0.0051) | (0.0538) | (0.0050) | (0.0352) |
| | | | | | | | |
| N | 1558 | 1558 | 1577 | 1577 | 1577 | 1577 | 1577 |
| R-squared | 0.001 | 0.002 | 0.0172 | 0.0042 | 0.0125 | 0.0009 | 0.0005 |

Robust standard errors are in parentheses. In addition, each specification controls only for a constant term, results omitted.

**Table D1: Estimates of Height Premia Conditional on Weight**
**White, Male, Full-Time Workers**

| | **Britain -- NCDS** | | **US -- NLSY** | |
|---|---|---|---|---|
| Adult height | 0.006 | 0.006 | -0.004 | -0.006 |
| (inches) | (0.0074) | (0.0076) | (0.0091) | (0.0091) |
| Adult weight | | -0.0006 | | 0.004 |
| (kg) | | (0.0003) | | (0.0023) |
| Teen height | 0.020 | 0.022 | 0.026 | 0.028 |
| (inches) | (0.0066) | (0.0088) | (0.0090) | (0.0096) |
| Teen weight | | -0.000 | | -0.004 |
| (kg) | | (0.0020) | | (0.0025) |
| N | 1753 | 1753 | 1577 | 1577 |

In the UK data the dependent variable is ln(wages) at age 33, in the US data the dependent variable is ln(wages) in 1996. Each sample is restricted to full-time white male workers. In the UK data, one extreme value of the age 16 weight distribution, a respondent weighing 252lbs was omitted from the sample. Specifications include controls for family background, region, and a constant term results not presented. Robust standard errors are in parentheses.