HEIGHT AND THE NORMAL DISTRIBUTION: EVIDENCE FROM ITALIAN MILITARY DATA*

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Researchers modeling historical heights have typically relied on the restrictive assumption of a normal distribution, only the mean of which is affected by age, income, nutrition, disease, and similar influences. To avoid these restrictive assumptions, we develop a new semiparametric approach in which covariates are allowed to affect the entire distribution without imposing any parametric shape. We apply our method to a new database of height distributions for Italian provinces, drawn from conscription records, of unprecedented length and geographical disaggregation. Our method allows us to standardize distributions to a single age and calculate moments of the distribution that are comparable through time. Our method also allows us to generate counterfactual distributions for a range of ages, from which we derive age-height profiles. These profiles reveal how the adolescent growth spurt (AGS) distorts the distribution of stature, and they document the earlier and earlier onset of the AGS as living conditions improved over the second half of the nineteenth century. Our new estimates of provincial mean height also reveal a previously unnoticed "regime switch" from regional convergence to divergence in this period.

Height data offer insights into the well-being of populations and historical periods for which other sources are lacking (Fogel 1994). Living conditions during the growing years influence final height through their impact on net nutrition—the balance between the supply of nutrients and the demands of metabolism, physical exertion, and disease (Silventoinen 2003). Though the variation of individual height is dominated by randomly distributed genetic potential, variation in average height over time or across socioeconomic groups is driven by systematic differences in diet, disease environment, workload, and health care. To the extent that these are functions of real income, mean heights provide indirect clues about economic conditions and are a direct measure of health status.

The study of height data has generated new insights into a wide range of issues but has left important questions unresolved. The relative importance of diet and disease continues to be debated, for example. The same is true of the role of income and genetic factors (Deaton 2007). Another set of issues concerns how to model the height distribution and its dependence on covariates such as age.

The basis of statistical models of height has long been the assumption that adult stature approximately follows a normal (Gaussian) distribution in a homogeneous, well-nourished population. This proposition, advanced by Quetelet in 1835 and established more systematically by Galton and Pearson a half-century later (Tanner 1981), is largely accepted among both scholars of human growth and statisticians (see, e.g., Snedecor and Cochran 1989;

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and Tanner 1990), although the log-normal distribution has also been suggested (Limpert, Stahel, and Abbt 2001; Soltow 1992). What has been less well appreciated is the converse: that the distribution of height may depart significantly from normality when net nutrition is inadequate or unequally distributed, or at ages before adult height is reached.

Deprivation can affect not only the mean but also the shape of the height distribution. Poor net nutrition and high childhood mortality can disproportionately cull smaller individuals from the population, negatively skewing the distribution. Further, the nonlinearity of the environment-stature link (genetic potential imposes an upper bound) means that inequality of nutritional status affects the entire distribution. Inequality lowers heights among the disadvantaged more than it raises them among the privileged, which lowers the mean, increases the variance, and can induce skewness and excess kurtosis in the distribution.

Significant departures from normality are also induced by the adolescent growth spurt (AGS). Individual variation in the timing and intensity of the AGS causes the variance of height to increase and its distribution to become negatively skewed during puberty. Among contemporary U.S. boys, for example, annual height increase (velocity) peaks at age 13.5, dispersion peaks at 14, and negative skewness peaks at 15 (data for the year 2000 from the Web site of the National Center for Health Statistics at www.cdc.gov/nchs/about/major/nhanes/growthcharts/datafiles.htm). By age 19, the storm has passed: velocity has fallen to almost zero, dispersion has diminished and stabilized at a lower level, and the distribution looks again bell-shaped. The poor and unequal net nutritional status typical of historical periods, but also of many developing countries today (Deaton 2008; Eveleth and Tanner 1990), delayed and prolonged the AGS beyond the age of 20.

Historical anthropometric studies rely primarily on records drawn from military archives. These data suffer from a variety of problems, such as variation in age at measurement, rounding and heaping, aggregation of individual data into frequency counts, selection effects of recruitment procedures (exemptions, deferments, draft evasion, preferences for robust individuals), and truncation from below due to minimum height requirements (MHR). The last of these problems has attracted the most attention in the literature because it affects virtually all samples drawn from records of enlisted soldiers. It is in solving the truncation problem that the normality assumption has been most frequently invoked.

A widely used technique based on the normality assumption is the quantile bend estimator (QBE) developed by Wachter and Trussell (1982) for composite samples that are not cleanly truncated by a single, strictly enforced MHR but display instead a deficient lower tail. The QBE chooses the parameter values that best match the upper quantiles of a normal distribution to those of the sample, using the point of divergence between the two (the "bend" in the quantile-quantile plot) to identify the point above which the sample is complete. The QBE is not easily adapted to include covariate effects. This is troubling because the failure to control adequately for age effects may lead to incorrect inference during periods over which the pattern of AGS is changing.

Truncated maximum likelihood (TML), also based on the normality assumption, allows estimation of covariate effects under the restrictive assumption that they affect only the mean of the distribution. The restricted TML estimator (A'Hearn 2004) imposes a standard deviation deemed plausible on the basis of nonsample information. Both estimators are bound to be biased upward because they neglect AGS-induced left-skewing. Moreover, their bias will be worse, the worse living conditions become. Thus, to the extent that estimates of a declining trend in mean height in the later 1700s are based on truncated samples, they may have understated the magnitude of the decline.

In this article, we propose a new semiparametric approach to modeling the distribution of height conditional on a set of covariates (e.g., age, calendar time) that dispenses with the normality assumption. Our approach enables us to construct and compare counterfactual distributions of height at different ages and to estimate population statistics of interest by simply evaluating the corresponding statistical functional at these counterfactual distributions. Though developed for tabulated frequency counts from historical military data, the general applicability of our approach makes it equally relevant to the growing literature in economics and human biology on the links between health and wealth in contemporary populations.

Applied to an unusually rich data set of height distributions from nineteenth century Italy, our approach yields new evidence on the effects of the AGS in historical contexts. Our counterfactual height distributions for ages 18 to 22 clearly display the pattern of AGS-induced deviations from normality found at much earlier ages in contemporary data. A general lesson is that historical studies must pay closer attention to such effects or risk distorting estimates of past trends in the distribution of stature. Our age-corrected distributions also provide substantive new insights into trends in the biological standard of living in Italy, particularly its inequality within and between provinces.

DATA

Our data are drawn from a series of annual statistical summaries published by the Italian Ministry of War's Directorate of Conscription and Manpower, which we dub the "Torre reports" after General Federico Torre, who headed the Directorate from its founding in 1861 until 1891. The Torre reports provide tables of height frequencies for each one-centimeter interval from 125 to 199 cm, for each birth cohort from 1855 through 1910, for finely disaggregated administrative regions.

The use of data derived from the Torre reports for anthropometric research is not new. They have been the basis of much subsequent work: the summary statistics calculated by Costanzo (1948), the historical series published by the Italian National Institute of Statistics (Istat 1958), comparisons of height distributions for large regions in selected years (Terrenato and Ulizzi 1983), and more recent analyses of the long-run trend in mean height (Arcaleni 2006; Federico 2003). To construct our data set, we returned to the original volumes of the Torre reports housed at the library of the Italian Parliament in Rome. After entering the data into electronic spreadsheets (thus preserving the original tabulations in electronic format), we subjected them to a series of checks to verify internal consistency and flag likely data errors.

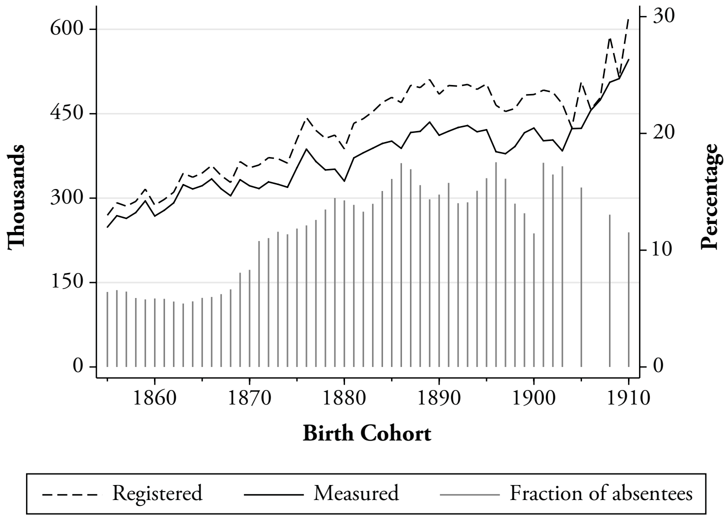
Our data are based on complete enumerations, not samples. The number of young men whose heights were tabulated in the Torre reports rose from about 250,000 individuals for the cohorts born before 1860 to over half a million for those born after 1905 (see Figure 1). Overall, the heights of over 21 million individuals were tabulated in these records. The geographic units of observation changed over time, from subprovincial districts up to the 1892 cohort, to provinces afterward. Using a historical catalog of administrative units to reaggregate the data and dropping territories annexed after World War I, we obtained height distributions for 68 consistently defined provinces for the entire period 1855–1910. At the provincial level, the number of observations for each birth cohort ranges from a minimum of 880 individuals (Livorno, 1855 cohort) to a maximum of 26,866 (Milan, 1910 cohort).

Recruitment Procedures

Statistical modeling of the Torre report data requires an understanding of the recruitment procedures that generated them. These derived from Piedmont's 1854 system of universal male conscription, which was extended to the entire country in the wake of Italian unification (1861).

The regulations specified that young men were to be called to arms in the year they turned 20, and directed municipal authorities to compile a conscription list each January. These conscripts were required to appear before provincial draft councils, which ruled on any claims for exemption or deferment, assigned draft lottery numbers, and conducted physical examinations to ascertain fitness for military service. Conscripts definitively failing the examination were excused from service. Those whose physical condition was

Figure 1. Number of Registered and Measured Individuals and Fraction of Absentees



judged capable of improvement were declared subject to reexamination and deferred to the next draft.

A detailed list of conditions automatically disqualified a conscript from service. These included obvious defects such as blindness or crippled limbs; diseases such as syphilis, tuberculosis, and epilepsy; and insufficient size or strength. Regarding these last attributes, an MHR was initially set at 156 cm; a limit for reexamination deferral, at 154 cm; and a minimum chest circumference, at 80 cm. The MHR was reduced to 155 cm in 1883, to 154 cm in 1913, and to 150 cm in 1917. The reexamination threshold was reduced to 153 cm in 1883 and to 148 cm in 1917.

Data Problems

Various problems in the Torre report data require attention: transitory data errors; heaping and rounding; top- and bottom-coding; interpretation of reported height intervals; variation in the age at measurement; and possible selection effects due to absenteeism, draft deferment, and the naval draft. These data problems may represent important causes of nonnormality in the observed distribution of height. However, because the data are for conscripts rather than recruits and because conscription was universal, they do not suffer the truncation problems typical of historical samples of enlisted soldiers.

Transitory data errors arise from several sources. The first is apparent mistakes in tabulation or printing that affect a single frequency in a single year in a single province. The second is apparent inconsistencies in the way the data from a province were reported to the Ministry, which affect the distribution more broadly, sometimes for more than one year. Rounding to even heights or those ending in 0 or 5 and heaping on the MHR are observed in some times and places. Heaping on the MHR can be attributed to provincial draft councils rounding up measured heights to avoid the burden of reexaminations or to deliver a mandated number of conscripts fit for service. Rounding patterns are inconsistent, and both the extent of heaping and the MHR itself varied over time. Accordingly, we classify these problems, too, as transitory data errors.

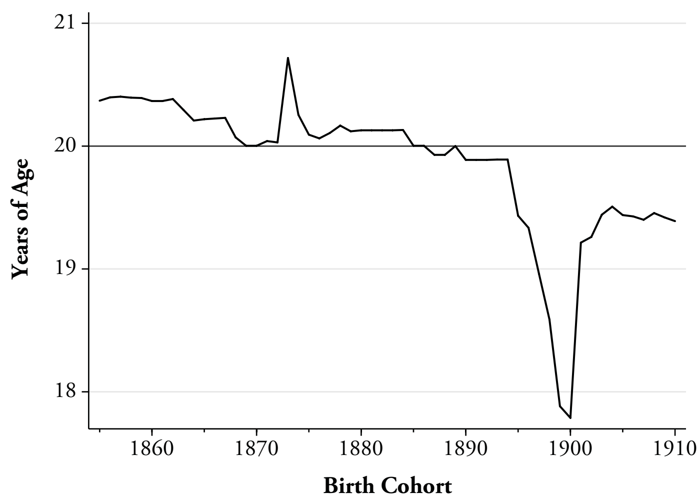
In the Torre reports, heights are top- and bottom-coded, with thresholds that may vary from year to year and across districts. In elaborating our data set, we impose on all years a uniform standard, coding any heights less than 125 cm at 124 cm and any heights over 199 cm as 200 cm. As for the interpretation of the recorded height intervals, Costanzo (1948) interpreted a label such as "stature = 165 cm" as "height in the interval 165.0–165.9," arguing that Ministry of War statisticians dropped (rather than rounded) fractions of centimeters in order to simplify data processing. By contrast, Livi's ([1896] 1905: part I, pp. 23–25) discussion indicates that before they were reported to the Ministry, most heights (three-fourths of them) had already been rounded to the nearest centimeter by military doctors when they measured the conscripts.

Average age at measurement can be estimated from the dates of the provincial draft council operations given in the Torre reports for each cohort, assuming uniform spacing of births over the calendar year. Figure 2 plots these estimates, which display a downward trend from 20.4 years for the class of 1855 to 19.4 for the class of 1910. The full range of variation is about three years between the maximum of 20.7 (birth cohort of 1873) and the minima of 17.9 and 17.8 years, respectively, for the classes of 1899 and 1900, who were called up early to serve in the first World War.

Italian men would still have been growing at these ages, so that intertemporal comparisons of mean height based on the raw distributions would be misleading. More subtly, interprovincial comparisons would also be distorted, since nutritional status affects not only final height but the speed with which it is reached. And in view of the effects of the AGS, it is likely that dispersion, skewness, and possibly kurtosis of height also varied with age. Previous studies have addressed the effects of age on the mean only. The Istat (1958) series is standardized to age 20 using the method of Costanzo (1948): adjusting raw mean height by growth increments derived from an early 1900s longitudinal study of Danish conscripts aged 18–24. Similarly, Terrenato and Ulizzi (1983) asserted that age affects only the mean and made no adjustment of higher moments.

Absenteeism was a significant problem in Italian conscription. The Torre reports allow us to compute the fraction of absentees by province for each cohort. On average over our

Figure 2. Estimated Mean Age at Measurement



period, roughly 12% of the individuals on the conscription rolls failed to appear before the provincial draft council and were not examined. This fraction varies considerably both over time and across provinces. Absenteeism was greatest in areas with known high emigration rates, and its rise among the birth cohorts of the 1870s and decline among those born in the 1900s corresponds well to the rise of emigration in the 1890s and its decline in the 1920s (Hatton and Williamson 1998: chap. 6). Overt draft evasion, which had been a serious problem in the first years of a united Italy (birth cohorts of the 1840s), was no longer a major issue in this period, and special cases due to errors in the conscription lists may be neglected.

The selection effect of emigration on height is ambiguous because emigrants can be selected either positively or negatively depending on travel costs, skill endowments, the relationship between height and skills, and the distribution of income opportunities in origin and destination countries (Borjas 1987). Empirically, there is little direct evidence on the stature of Italian emigrants. Danubio, Amicone, and Vargiu (2005) provided evidence suggesting positive selection on height, but their study relied on a small sample of self-reported data and lacks geographic details.

As for deferment to the next draft, for much of the period under study, the Torre reports offer a single frequency distribution of height covering both the current draft class and those deferred from previous years. Most deferments were for an insufficiently robust constitution, a condition judged by chest circumference and typically associated with above-average height. But other causes of deferment may have been correlated with below-average stature, since poor nutrition renders an individual more susceptible to disease. A small number were automatically deferred because their height fell between the MHR and the reexamination limit. Costanzo's (1948) evidence indicates only a negligible impact of deferred conscripts on the mean and variance of the national distribution of height.

Finally, the army draft was preceded each year by a naval draft, which took only a few thousand individuals nationwide. In principle, naval draftees were removed from the conscription rolls before the army draft. The limited number of individuals involved is unlikely to affect national statistics but may well affect the distribution of heights for a few provinces.

STATISTICAL APPROACH

To deal with the problems discussed in the previous section, we adjust the height distributions in the Torre reports on the basis of four principles. First, we adjust the entire distribution of height, then compute any summary statistics of interest. This can be contrasted with first summarizing the distribution, then attempting ad hoc and independent adjustments of sample moments, as in the Costanzo/Istat standardization of mean height described earlier. Second, we make no assumption about the overall shape of the height distribution. Third, we allow the whole height distribution, not just its mean, to vary with age at measurement, calendar time, and possible selection effects due to absenteeism; for lack of data, however, we cannot also control for the possible selection effects of the naval draft and of deferment to the next draft. Fourth, we adjust at the finest level allowed by the data—namely, the province.

Adjusting for Age and Selection Effects

The unobservable individual measurements underlying the tabulations in the Torre reports may be regarded as draws from a probability distribution that represents the variability of height across the surviving members of a cohort at the time when they are measured. Formally, let the random variable H_t represent measured height for an individual born in year t who is subject to medical examination in year $t + a$ in a given province. Two problems arise when trying to learn about the distribution of H_t from the data.

The first problem is that we have access only to tabulated frequencies by height interval. Hence, all we can hope to identify from the Torre reports is the discrete distribution whose probability mass function coincides with the probability that H_t falls within each available height interval. To recover this discrete distribution, we adopt a semiparametric approach. Let $h_1 < h_2 < \ldots < h_J$ denote the limits of the height intervals observed in the data. The probability associated with each height interval is

$$\pi_{jt} = \begin{cases} \Pr\left(H_t \le h_1\right) & \text{if} \quad j=1 \\ \Pr\left(h_{j-1} < H_t \le h_j\right) & \text{if} \quad j=2,\ldots,J \\ \Pr\left(H_t > h_J\right) & \text{if} \quad j=J+1 \end{cases}. \tag{1}$$

Since the width of each height interval (1 cm) is small relative to the random variability of heights, the discrete approximation to the distribution of H_t provided by the π_{jt} is likely to be sufficiently accurate for most purposes. In principle, these probabilities depend on individual characteristics that were observable at the time of the physical examination (e.g., age, health status, labor force status) but are unavailable in the Torre report data. As explained below, our modeling strategy consists of projecting the π_{jt}, which vary across individuals because of their dependence on unobservables, on a time trend and a few observable cohort characteristics. Our approach is semiparametric because we do not restrict the shape of the distribution of H_t but model parametrically the dependence of the π_{jt} on the observable cohort characteristics under the side constraints that the π_{jt} are strictly between zero and one, and sum to unity.

To impose the adding-up constraint, we exploit the fact that the probability distribution of H_t may equivalently be characterized through the survival function $S_t(h) = \Pr(H_t > h)$. This characterization is particularly convenient for our purposes because of the recursion

$$S_{jt} = (1 - \lambda_{jt})S_{j-1,t}, \quad j = 2, \ldots, J, \tag{2}$$

where $S_{jt} = S_t(h_j) = \Pr\{H_t > h_j\}$ and $\lambda_{jt} = \Pr\{h_{j-1} < H_t \le h_j \mid H_t > h_{j-1}\} = \pi_{jt} / S_{j-1,t}$ is the discrete hazard of H_t. The recursion (2) implies that

$$S_{jt} = S_{1t} \prod_{k=1}^{j} \left(1 - \lambda_{kt}\right), \quad j = 2, \ldots, J. \tag{3}$$

This suggests modeling the first probability $\pi_{1t} = 1 - S_{1t}$ and the discrete hazards λ_{jt}, $j = 2,\ldots, J$, as functions of observable cohort characteristics. This approach is similar to that in Peracchi (2002), except that one does not need to chose a partition of the range of H_t because this partition is already determined by the available data.

The second problem is how to exploit the information contained in a few observable variables: the average age at measurement, calendar time, and the fraction of absentees. These variables are likely to shift the height distribution in ways that may differ across provinces and over time. Because π_{1t} and the discrete hazards λ_{jt} are probabilities and are therefore bounded between zero and one, it is natural to model their log-odds $\eta_{1t} = \log(\pi_{1t} / (1 - \pi_{1t}))$ and $\eta_{jt} = \log(\lambda_{jt} / (1 - \lambda_{jt}))$. A parsimonious but flexible model is the additive specification

$$\eta_{jt} = \alpha_j + b_j(a_t) + f_j(t) + g_j(p_t) + \nu_{jt}, \quad j = 1, \ldots, J, \tag{4}$$

where $b_j(a_t)$ is a function of average age at measurement a_t for the cohort born in year t, $f_j(t)$ is a deterministic time trend, $g_j(p_t)$ is a polynomial in the fraction p_t of absentees in the given province, and ν_{jt} is an approximation error. Smoothing over time using a deterministic trend limits the influence of the transitory errors described in the previous section. It also makes assessment of time trends, such as convergence of provincial mean heights, less dependent on the particular years that are chosen to define the various subperiods.

Given estimates of the unknown parameters in the models for the log-odds, we derive predicted or "adjusted" log-odds:

$$\hat{\eta}_{jt} = \hat{\alpha}_j + \hat{b}_j(\bar{a}) + \hat{f}_j(t) + \hat{g}_j(\bar{p}), \quad j = 1, \ldots, J, \tag{5}$$

where the function $\hat{b}_j(\cdot)$ is evaluated at some reference age $\bar{a}$ and the function $\hat{g}_j(\cdot)$ is evaluated at some reference value $\bar{p}$. Given the adjusted $\hat{\eta}_{jt}$, we compute adjusted $\hat{\pi}_{1t}$ and $\hat{\lambda}_{jt}$ via the inverse logit transformation. Adjusted relative frequencies $\hat{\pi}_{jt}$ are then obtained from the set of recursions

$$\hat{\pi}_{jt} = \hat{\lambda}_{jt}\hat{S}_{j-1,t} \tag{6}$$

$$\hat{S}_{jt} = \hat{S}_{j-1,t} - \hat{\pi}_{jt} \tag{7}$$

for $j = 2, \ldots, J$, starting from $\hat{S}_{1t} = 1 - \hat{\pi}_{1t}$. By construction, these adjusted relative frequencies are bounded between zero and one. By setting $\hat{\pi}_{J+1,t} = \hat{S}_{Jt}$, the adding-up condition is also satisfied by construction. Notice that although age at measurement and calendar time enter additively in the model for the log-odds, this does not prevent age and time from interacting in shaping the distribution function of height. In fact, interaction effects arise both from the use of the inverse logit transformation to obtain probabilities from the estimated log-odds and from the multiplicative nature of the representation (3).

Finally, because the entire height distribution has been adjusted, any population statistics of interest—whether moment- or quantile-based measures—can be obtained by using the adjusted relative frequencies $\hat{\pi}_{jt}$ as weights. For example, if $\hat{F}_t(h) = 1 - \hat{S}_t(h)$ denotes the adjusted distribution function of heights, then the adjusted mean of heights is $\hat{\mu}_t = \int h d\hat{F}_t(h)$.

Implementation Details

After experimenting with a number of different specifications, we take $b_j(\cdot)$ and $g_j(\cdot)$ to be linear in the logarithm of age and the fraction of absentees, respectively, and model the time trend $f_j(\cdot)$ as a cubic polynomial. We also set $h_1 = 129$ cm and $h_J = 190$ cm.

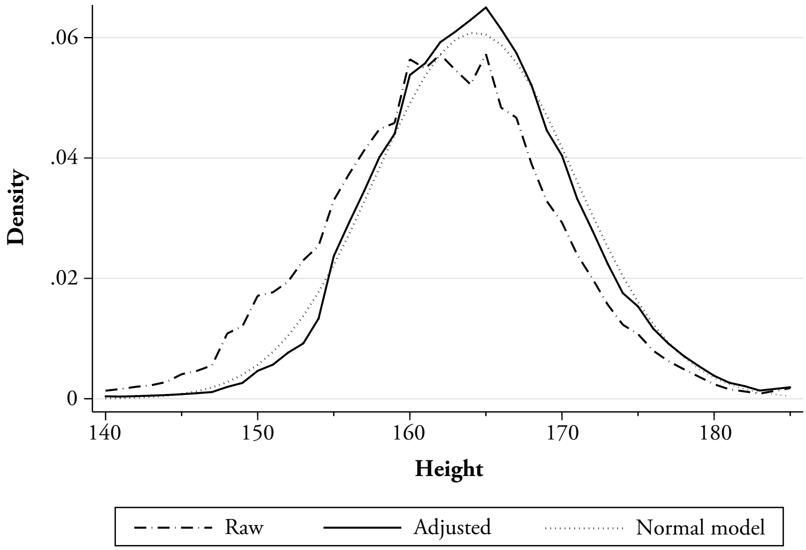
In the case of the cohorts born in 1904, 1906, and 1909, for unclear reasons, most provinces reported only the number examined and not the number enrolled on the conscription lists. This makes it impossible to calculate the fraction of absentees for these three cohorts. For the 1907 birth cohort, the fraction of absentees is nonzero but is implausibly low. We treat the values for these four cohorts as missing and impute them by linear interpolation at the provincial level.

The models for the log-odds have been estimated by weighted least squares regressions of the empirical log-odds on a constant, the logarithm of age at measurement, a cubic time trend, and the fraction of absentees, with weights proportional to the provincial cohort size in each year. We use the estimated model to construct counterfactual distributions of height by varying the reference age between 17 and 22 years, keeping the fraction of absentees constant at $\bar{p} = .05$, which is the median value of the fraction of absentees in the Italian provinces before the emigration-induced increase that begins with the birth cohorts of the early 1870s.

Adjusted absolute frequencies are generated by multiplying adjusted relative frequencies by the size of each provincial cohort (the number of people on the conscription lists minus the reference fraction $\bar{p}$ of absentees). Summing frequencies over provinces yields adjusted frequencies at the national or regional level.

As an illustration, Figure 3 contrasts the raw national density of height for the 1900 birth cohort (dashed line) with the adjusted (age-20) density (solid line) and a normal density with mean and standard deviations equal to those of the adjusted density (dotted line).

Figure 3. Raw and Age-20 Adjusted Densities of Height, 1900 Birth Cohort



Note: The curve labelled "Normal model" is a normal density with the same mean and standard deviation as the adjusted density.

The 1900 cohort was drafted during World War I and measured at an average age of only 17.8 years. Absentees in this cohort were 12.3% of those on the conscription rolls, a fraction roughly equal to the average over the whole period. We take age 20 as the reference for our "counterfactual" height distribution because it is close to the 1855–1910 average of 19.9 and makes our results comparable with Costanzo's (1948) study. Relative to the raw density, the adjusted density is shifted to the right, the mean increasing by 2.5 cm from 161.8 to 164.3 cm. As predicted by our discussion of the AGS, the age adjustment also changes the shape of the distribution: its standard deviation falls by nearly 1 cm from 7.54 to 6.56 cm; its skewness turns from negative (–0.225) to about zero (–0.058); and its excess kurtosis increases slightly from 0.660 to 0.962. The effects of smoothing over time are evident in the disappearance of raw-data peaks at 150 and 148 cm, which were the MHR and reexamination limits for this cohort, and also of the peculiar heaping pattern in the 160–165 cm range.

Relative to a normal distribution with the same mean and standard deviation, our adjusted distribution has more mass in the center and less mass in the tails, a finding similar to that of other studies (Akachi and Canning 2007; Hermanussen, Burmeister, and Burkhardt 1995). The departure from normality of our age-20 smoothed distribution may arise for three main reasons. First, AGS-driven nonnormality may persist even at age 20 under conditions of nutritional stress that were typical of this historical period. Second, clustering of the population into groups, each with a different distribution of height, may lead to a nonnormal mixture of the group-specific distributions. Third, nonnormality may arise from selection effects different from absenteeism. In our case, formal tests of the normality assumption are unlikely to be informative, because even small differences between the observed and the hypothesized distributions will be magnified by the sheer number of observations underlying our data. Although this is the key for consistency of classical testing procedures, it clearly raises the issue of their relevance.

Figure 4 shows, for the 1900 birth cohort, the ratio between the adjusted national densities of height with and without controls for selection due to absenteeism. Overall, the differences between the two densities are confined to the tails of the distribution and tend to be small (the ratio varies between 0.90 and 1.10, except at the extreme tails). In this example, controlling for absenteeism does not change the mean of height and increases the standard deviation and the degree of (negative) skewness only marginally. Its only noticeable effect is to increase excess kurtosis from 0.73 to 0.96.

EMPIRICAL RESULTS

This section describes broad trends in the adjusted distributions of height, both at the national level and the disaggregated provincial level. For comparability with other studies, we focus on means and standard deviations, although at the national level, we also present results for adjusted skewness and excess kurtosis.

National Level

Figure 5 is a time series plot of the raw (unadjusted) national mean of heights, Costanzo's (age-20) adjusted series, and our (age-20) adjusted estimates. Following Livi ([1896] 1905), Costanzo's series has been re-centered by deducting half a centimeter to account for differences in the interpretation of reported height intervals. Standardization to age 20 decreases mean heights of the early cohorts relative to the raw mean and raises those of later cohorts. Smoothing through the use of a polynomial trend removes some of the short-run fluctuations evident in both the raw mean and Costanzo's series. Otherwise, our estimates

Figure 4. Ratio of the Adjusted Densities of Height With and Without Controls for Selection Due to the Absentees, 1900 Birth Cohort

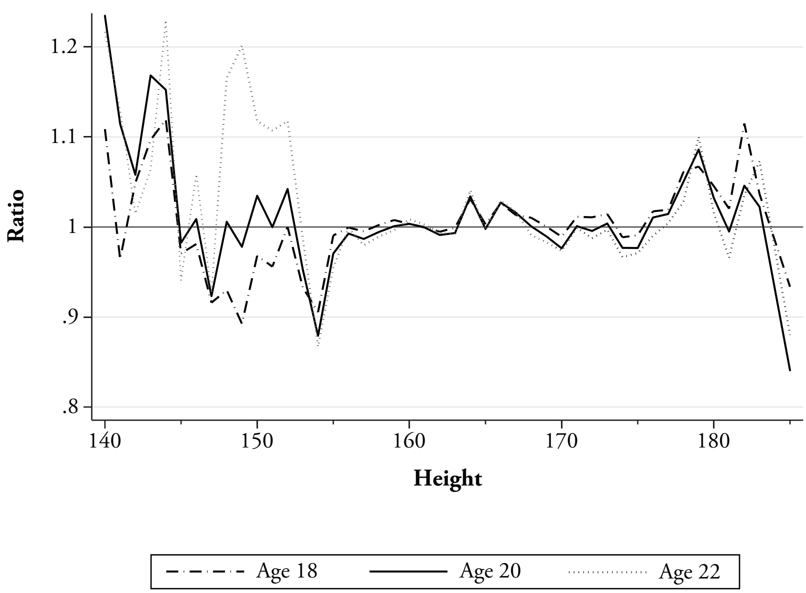
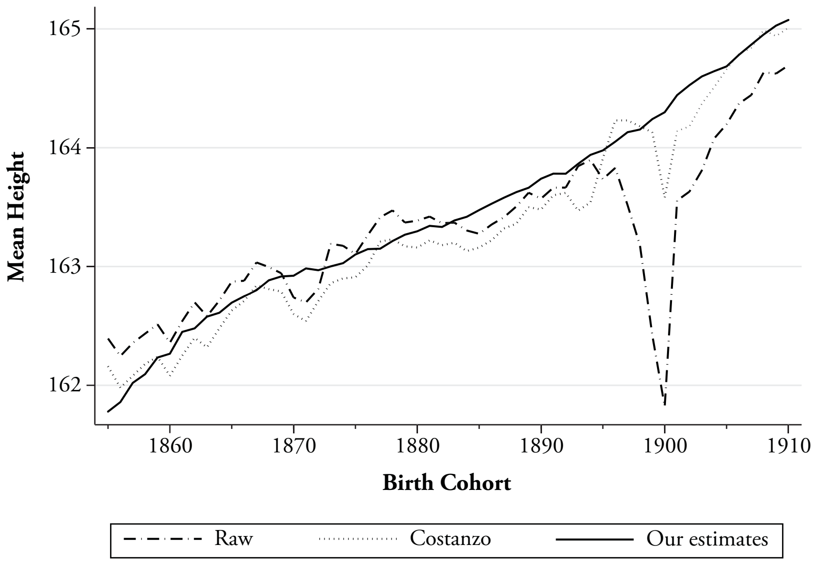


Figure 5. Raw Mean Height by Birth Cohort, Costanzo's (1948) Age-20 Adjusted Series (recentered), and Our Age-20 Adjusted Estimates



and those of Costanzo correspond fairly closely in level and overall trend. Both series indicate a cumulatively substantial increase of about 3.3 cm in mean height.

This trend in mean height corroborates available estimates of individual components of a biological standard of living. Between 1862–1864 and 1910–1912, male mortality rates in the first five years of life fell from over 40% to less than 25%, while male life expectancy at birth increased from 30 to nearly 47 years (Caselli 1987). If the disease environment improved, so too did real income and, with it, nutrition. The latest estimates indicate a 70% increase in real GDP per capita from 1861 to 1910 (Fenoaltea 2006). Per capita availability of foodstuffs, calculated from domestic agricultural production and net imports, is estimated to have risen from approximately 2,500 calories per capita per day to 3,000 over the same period (Federico 2003).

Whether diet or disease was the decisive factor in improving health cannot easily be distinguished in the historical record. Massive investment in water distribution networks beginning in the 1880s is reflected in mortality rates, but the response is subtle (Caselli 1987; Federico 2003), while the upward trend in heights is entirely unperturbed. Neither does the sharp acceleration in real wage growth estimated for this period leave much trace (Fenoaltea 2006). Federico's (2003) regression exercise attributes some 60% of the increase in mean height from 1854 to 1913 to better nutrition, 30% to improvements in sanitary conditions, and the remaining 10% to other variables such as decrease in workload.

Our modeling approach offers a new perspective on improving health. The age-height profiles displayed in Figure 6 are derived from the counterfactual height distributions obtained by varying the reference age over the range from 17 to 22 years, while keeping the fraction of absentees constant. The shifts in the age-height profiles indicate clearly that conscripts born in the first years of the twentieth century were not only taller than their grandfathers; they also reached final height much earlier. This growth anticipation effect is an aspect of the link between net nutrition and height that has been inadequately appreciated. It is ruled out by assumption in the Costanzo/Istat methodology (which adds the same growth increment for a given change in age throughout the period considered) and is revealed here for the first time for Italy.

Our method also allows us to assess trends in the higher moments of the distribution without AGS-induced distortions, a problem entirely ignored in the Costanzo/Istat method. Figure 7 plots the raw and adjusted standard deviations of height at ages 18, 20, and 22, with the standard deviation computed as $\hat{\sigma}_t = \left[\int \left(h - \hat{\mu}_t\right)^2 d\hat{F}_t\left(h\right)\right]^{1/2}$, where $\hat{F}_t$ and $\hat{\mu}_t$, respectively, denote the (raw or adjusted) distribution function and the mean of height for the cohort born at time t. Contrary to the hypothesis of a relatively constant coefficient of variation of height, mean and standard deviation move in opposite directions. The age-20 standard deviation falls by almost one centimeter from the birth cohort of 1855 to those of the early 1890s before rising again slightly. A comparison of the counterfactual age-18 and age-20 series shows that this decline was largely driven by the anticipation of the AGS as net nutrition improved.

Terrenato and Ulizzi (1983) documented a convergence of the Italian height distribution toward a Gaussian shape, with both negative skewness and excess kurtosis diminishing, especially after 1874. However, they made no correction for age at measurement, explicitly claiming that it affects only the mean and not the higher moments of the distribution; nor did they correct for selection effects. Their first conclusion is in line with the evidence in Figures 8 and 9, which shows the raw and adjusted skewness and excess kurtosis of height at ages 18, 20, and 22; here, skewness is computed as $\int \left(h - \hat{\mu}_t\right)^3 d\hat{F}_t\left(h\right) / \hat{\sigma}_t^3$ and excess kurtosis computed as $\left[\int \left(h - \hat{\mu}_t\right)^4 d\hat{F}_t\left(h\right) / \hat{\sigma}_t^4\right] - 3$. Both raw and age-20 skewness and

Figure 6. Estimated Profiles of Age-Adjusted Mean Height for Selected Birth Cohorts

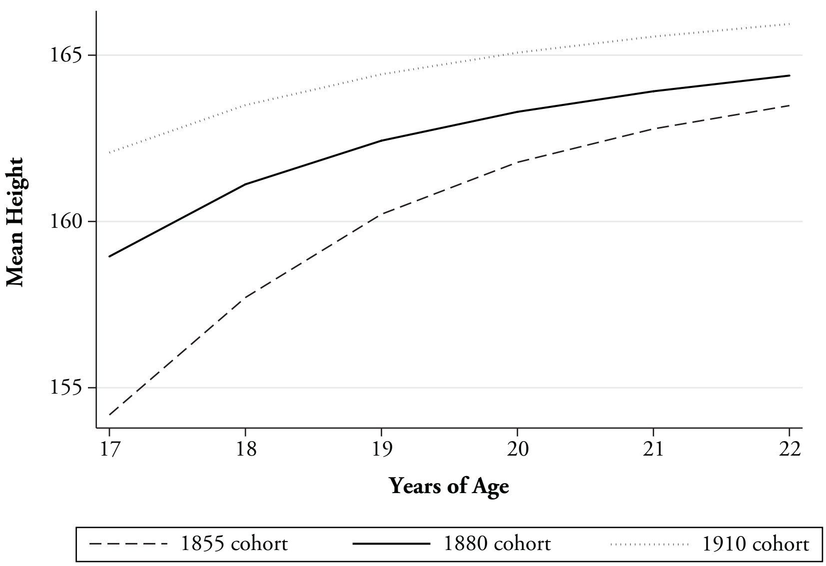


Figure 7. Raw and Age-Adjusted Standard Deviation of Height, by Birth Cohort

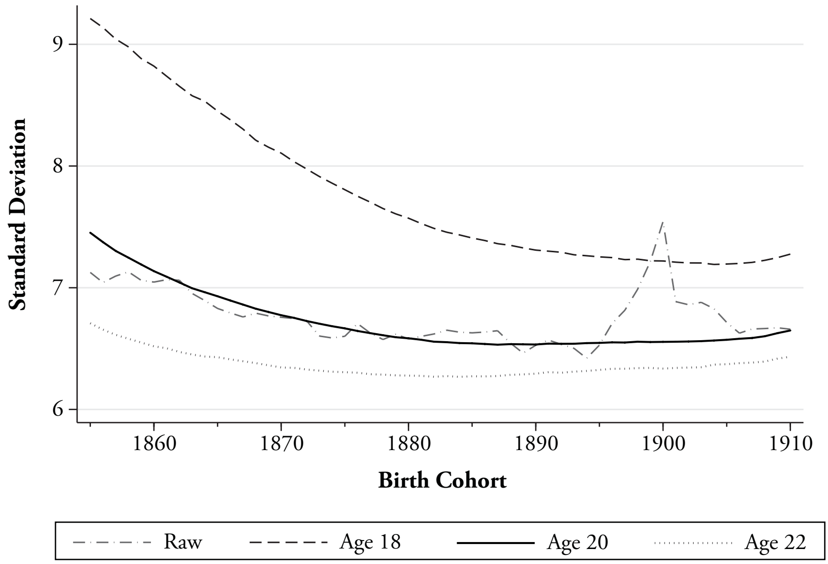


Figure 8. Raw and Age-Adjusted Skewness of Height, by Birth Cohort

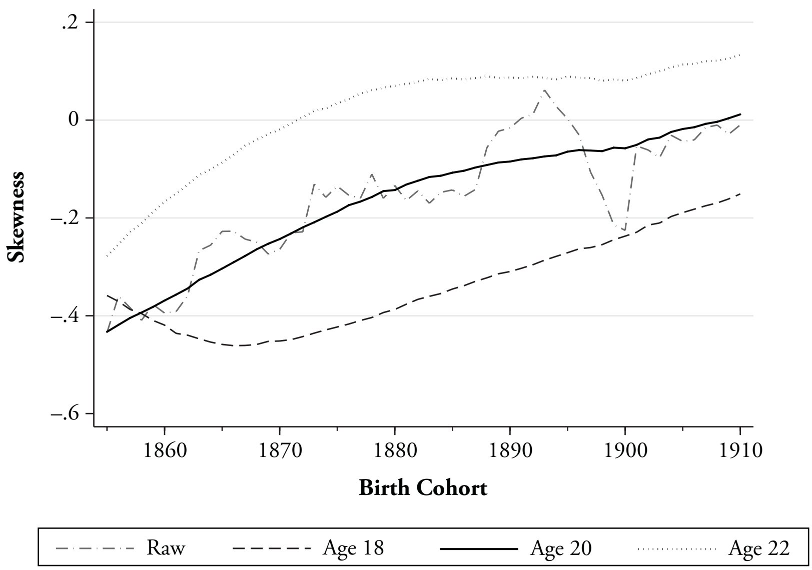
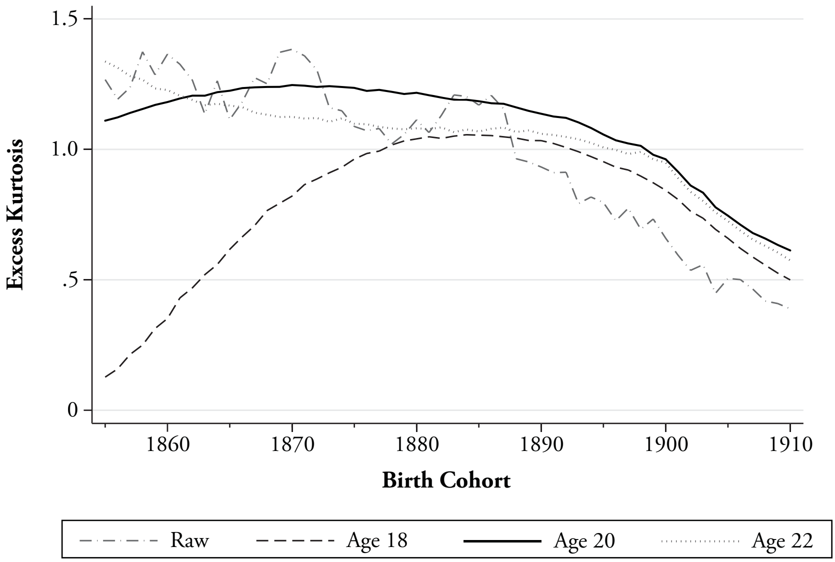


Figure 9. Raw and Age-Adjusted Excess Kurtosis of Height, by Birth Cohort



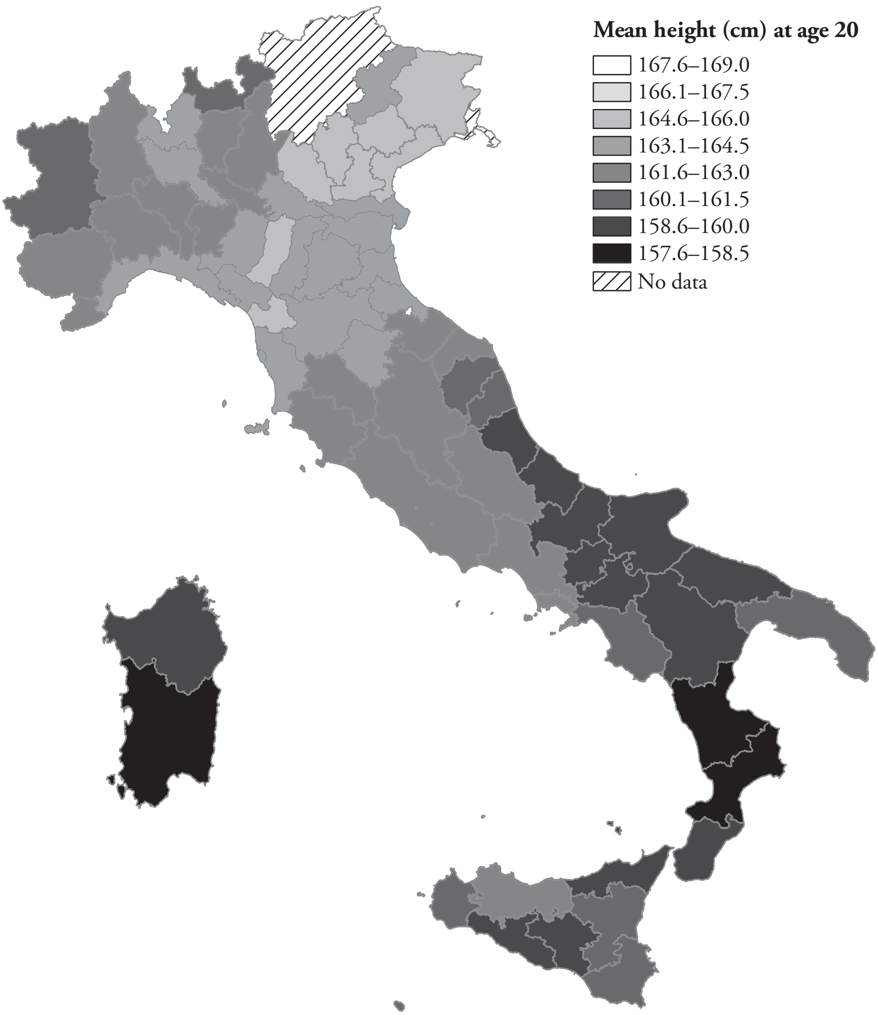
excess kurtosis do converge toward Gaussian values. Terrenato and Ulizzi's second claim, however, is not supported by the evidence; both the levels and the trends of skewness and kurtosis are different at different ages. For the birth cohorts around 1880, for example, skewness is clearly negative at age 18 and age 20 but is slightly positive at age 22. As for the trends, age 18 looks quite different from age 20 and age 22. A simple interpretation of these patterns is the following. Suppose that the peak age for nonnormality induced by the AGS is 19 in 1855, but moves down to 16 by 1910. In this case, higher moments should display a similar trend for ages 20 and 22, which are further and further away from the AGS distortions. For age 18, by contrast, initial improvements in living conditions would at first move the height distribution into the phase of maximum nonnormality. Later, as the onset of the AGS occurs earlier and earlier, 18-year-olds are also past this phase, and their height distribution looks more normal.

Provincial Level

The national distributions of height discussed in the previous section mask considerable diversity at the provincial level—diversity that can be explored in depth on the basis of the new estimates developed here. This section examines the provincial trends in mean height, decomposes the variance of height at the national level into within- and between-province components, and assesses patterns of convergence and divergence at the provincial level.

Figures 10 and 11 map age-20 provincial mean heights for the first (1855) and the last (1910) cohorts in our data. Numerical values are presented in Table 1. For the 1855 cohort, the range of provincial mean height is wide: almost 8 cm, from 157.6 cm (Cagliari, Sardinia) to 165.5 cm (Lucca, Tuscany). While there is considerable variability within each of the three macro-regions of North, Center, and South-Islands, a North-South gradient is evident. For the 1910 cohort, the North-South disparity is still apparent, but age-20 mean

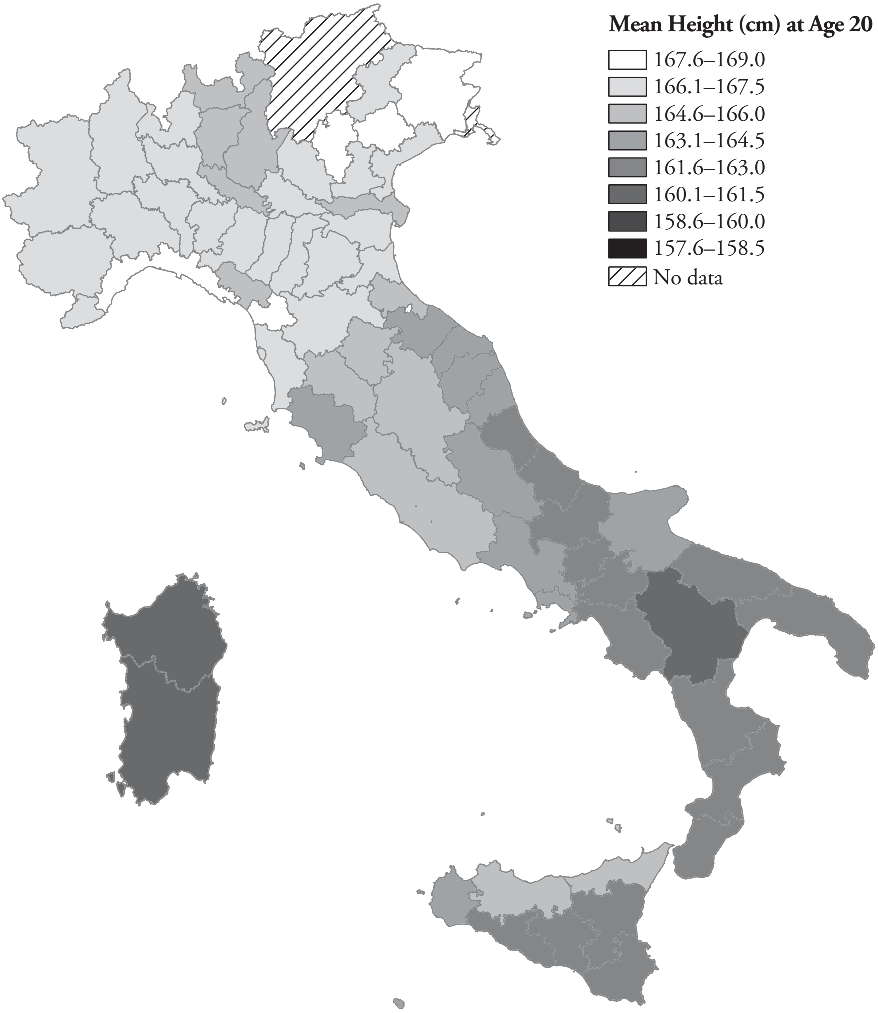
Figure 10. Age-20 Adjusted Provincial Mean Height, 1855 Birth Cohort



height has increased everywhere: in no province by less than 1.3 cm. Clearly, net nutrition improved throughout the country over the 56 years under study here.

Drawing inferences about variation in living standards from variation in mean height is a challenging task, partly because variation in unobserved genetic potential across provinces cannot be ruled out, and partly because separating the effects of higher incomes and better health is difficult. Our age-adjustment procedure provides new evidence that provincial disparities in mean height cannot be attributed entirely to genetic factors. Again it is the counterfactual age-height profile that yields a novel insight, this time at the provincial level. The slope of the age-height profile (or velocity) is systematically greater where mean height is smaller (the provincial-level correlation of velocity and mean height at age 18 in 1855

Figure 11. Age-20 Adjusted Provincial Mean Height, 1910 Birth Cohort



is –.75). Greater velocity at a given age indicates a late-onset AGS. Because the genetic influences controlling AGS onset are largely independent of those controlling final height, this correlation is clear evidence of environmental effects at work (Tanner 1990:122).

Although we cannot exclude a role for genetic factors in cross-province differences, within-province differences over time can be unambiguously attributed to changes in environmental factors because internal migration in this period was limited and short-range. In several provinces, mean height increased by 5 cm or more over the sample period, suggesting that variation in environmental factors can explain significant variation in mean height. A simple way of separating environmental and genetic factors is to represent the mean height of cohort t in province p as $\bar{H}_{pt} = \alpha_p + \delta_t + \mu_{pt}$, where α_p is a time-invariant

Table 1. Provincial Mean Height at Age 20 in 1855 and 1910, Difference in Mean Height Between 1855 and 1910 ($\Delta\mu$), and Difference-in-Difference Coefficient (DiD)

Area	Region	Province	μ_{1855}	μ_{1910}	$\Delta\mu$	DiD
Northwest	Piemonte	Alessandria	162.6	167.1	4.4	1.63
		Cuneo	162.1	166.5	4.4	1.55
		Novara	162.4	167.2	4.8	2.00
		Torino	161.2	167.2	6.1	3.28
	Liguria	Genova	163.6	168.7	5.1	2.28
		Imperia	162.8	166.8	4.0	1.20
	Lombardia	Bergamo	162.9	165.6	2.7	–0.12
		Brescia	162.3	165.4	3.1	0.25
		Como	163.0	166.8	3.8	1.03
		Cremona	162.1	166.0	3.9	1.10
		Mantova	163.4	166.4	3.0	0.20
		Milano	163.9	166.9	3.0	0.23
		Pavia	162.1	167.0	4.9	2.11
		Sondrio	160.4	165.4	5.0	2.17
Northeast	Veneto	Belluno	164.2	167.3	3.1	0.30
		Padova	165.2	167.2	2.1	–0.75
		Rovigo	164.0	165.9	1.9	–0.92
		Treviso	165.0	167.9	2.9	0.09
		Venezia	164.9	167.1	2.1	–0.69
		Verona	164.9	166.8	1.9	–0.89
		Vicenza	165.2	167.5	2.3	–0.54
	Friuli	Udine	165.3	167.7	2.4	–0.37
	Emilia Romagna	Bologna	163.4	167.0	3.6	0.82
		Ferrara	163.1	166.6	3.4	0.63
		Forli	163.6	165.6	2.0	–0.81
		Modena	163.6	166.5	2.8	0.04
		Parma	163.0	166.9	3.9	1.06
		Piacenza	162.6	166.7	4.0	1.22
		Ravenna	163.7	167.2	3.4	0.60
		Reggio Emilia	164.6	166.4	1.8	–1.04
Center	Toscana	Arezzo	163.4	164.7	1.3	–1.48
		Firenze	164.2	167.4	3.2	0.38
		Grosseto	161.8	164.2	2.4	–0.42
		Massa Carrara	163.8	165.7	1.9	–0.91
		Livorno	164.2	167.5	3.3	0.47
		Lucca	165.5	167.9	2.4	–0.43
		Pisa	164.1	167.4	3.3	0.48
		Siena	162.6	165.1	2.5	–0.33

(continued)

(Table 1, continued)

Area	Region	Province	μ_{1855}	μ_{1910}	$\Delta\mu$	DiD
Center (cont.)	Marche	Ancona	162.6	164.6	2.0	–0.82
		Ascoli Piceno	160.0	164.1	4.1	1.29
		Macerata	161.4	163.6	2.2	–0.65
		Pesaro Urbino	162.8	164.1	1.4	–1.45
	Umbria	Perugia	161.9	164.6	2.6	–0.18
	Lazio	Roma	161.8	164.6	2.8	0.00
South	Abruzzi e Molise	Aquila	161.9	164.0	2.1	–0.69
		Campobasso	159.0	162.5	3.5	0.73
		Chieti	159.2	163.0	3.8	0.99
		Teramo	159.3	162.5	3.2	0.38
	Campania	Avellino	158.7	162.4	3.7	0.92
		Benevento	159.5	162.6	3.1	0.31
		Napoli	162.7	164.0	1.4	–1.46
		Salerno	160.1	162.5	2.5	–0.35
	Puglia	Bari	159.3	162.8	3.5	0.65
		Foggia	159.6	163.4	3.9	1.05
		Lecce	160.3	163.1	2.8	–0.02
	Basilicata	Potenza	158.7	160.9	2.2	–0.62
	Calabria	Catanzaro	157.9	162.2	4.4	1.54
		Cosenza	158.0	162.5	4.5	1.73
		Reggio Calabria	159.0	162.0	3.0	0.18
Islands	Sicilia	Caltanissetta	159.1	161.7	2.6	–0.18
		Catania	160.6	163.2	2.7	–0.13
		Agrigento	159.4	162.2	2.8	0.02
		Messina	159.8	165.0	5.2	2.42
		Palermo	161.9	164.6	2.7	–0.12
		Siracusa	159.9	162.6	2.6	–0.18
		Trapani	160.6	163.8	3.2	0.35
	Sardegna	Cagliari	157.6	160.5	2.9	0.09
		Sassari	159.4	160.8	1.4	–1.37

province effect that includes the genetic component, δ_t is a time trend common to all provinces, and μ_{pt} is a time-varying component that captures provincial deviations from the common trend. Under this assumption, which rules out interactions between the genetic component and the time trend, a comparison of provincial mean heights at a given point in time removes the common trend, since

$$\bar{H}_{pt} - \bar{H}_{qt} = \left(\alpha_p - \alpha_q\right) + \left(\mu_{pt} - \mu_{qt}\right), \tag{8}$$

but does not remove the time-invariant province effects. On the other hand, comparing changes over time in provincial mean heights removes both the common trend and the time-invariant province effects, since

$$\left(\bar{H}_{pt} - \bar{H}_{qt}\right) - \left(\bar{H}_{ps} - \bar{H}_{qs}\right) = \left(\mu_{pt} - \mu_{qt}\right) - \left(\mu_{ps} - \mu_{qs}\right). \quad (9)$$

Thus, under the assumed model, the set of differences-in-differences $\left(\bar{H}_{pt} - \bar{H}_{qt}\right) - \left(\bar{H}_{ps} - \bar{H}_{qs}\right)$ provides a useful measure of the differential changes over time in living conditions, being net of genetic components and common trends.

Table 1 and Figure 12 present this set of differences-in-differences, with Rome taken as the reference province. (Rome's mean height for the cohort of 1855 is almost exactly equal to the national average, and the growth in Roman mean height to 1910 is just a few millimeters below average.) Though groups of contiguous provinces tend to share similar growth experiences, it is difficult to detect a systematic pattern in Figure 12 other than the uniformly above-average growth in the Northwest (an area that came to be known as the Industrial Triangle in this period). There is, in particular, no North-South gradient—no tendency for the shorter Southern provinces to experience above-average growth and catch up on the North. There is, in short, no convergence.

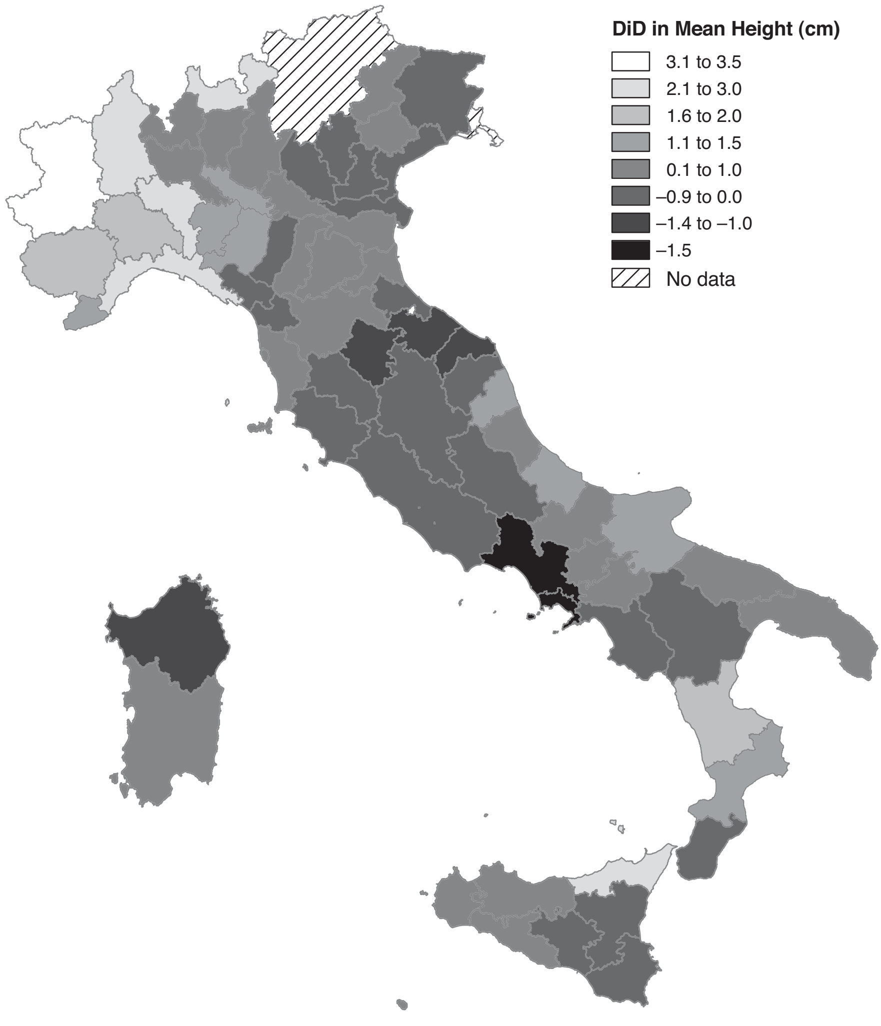
In the economic growth literature, a distinction is typically drawn between beta- and sigma-convergence, the former referring to a negative correlation of the initial level of a variable with its subsequent growth and the latter referring to a decrease in the dispersion of a variable. As documented in the upper panel of Table 2, neither type of convergence is observed in provincial mean heights. The estimated convergence coefficient (β) of –0.13 is so small as to imply only 1 mm slower growth over 56 years for an initial height advantage of 1 cm. Similarly, the change in the standard deviation of provincial mean heights ($\Delta\sigma$) is effectively zero. In the North, there is evidence of significant internal convergence, but elsewhere even within-region convergence is scant.

Yet this overall (lack of) pattern masks a switch from convergence to divergence around 1880. The variance decomposition exercise in Figure 13 shows that within-province variability declines almost monotonically throughout the period considered. It is between-province variability, reaching a minimum in the late 1870s and rising from the late 1880s, that flattens and reverses what would otherwise be a downward trend in Figure 7. The different scales for within- and between-province variance in Figure 13 reveal the predominance of the former component in total height variability. This simply reflects the fact that individual variability due to random genetic factors, which is part of within-province variation, is large compared with systematic differences in mean height.

The switch from convergence to divergence is also evident in Figure 14 and Table 2 (lower panels). The period from 1855 to 1880 displays strong convergence of all sorts: sigma, beta, within-region, and between-region. At the national level, the standard deviation of provincial mean height falls from 2.14 to 1.79, the correlation between initial provincial mean height and its subsequent increase is strong and negative, and the convergence coefficient is nearly three times its full-period value. The average Southern province converges strongly on its taller Northern and Central counterparts, growing by half a centimeter more over 25 years. The negative correlation of initial provincial mean height and its subsequent growth is visually apparent in the left panel of Figure 14, both within and between regions.

With the 1880s begins a very different era. Within regions, convergence largely ceases outside the North. Between regions, a process of outright divergence takes hold. Mean provincial growth rates accelerate by 40% in the North, while slowing by more than 60% in the South. The change is visually apparent in the realignment of the Northern and Southern groups of provinces in the right panel of Figure 14.

Figure 12. Differences-in-Differences in Age-20 Adjusted Provincial Mean Height Between the 1910 and the 1855 Birth Cohort (baseline = Rome province)



Height, Health, and Regional Economic Development

The North-South gradient in height still evident at the end of our sample period is consistent with what is known of regional economic differences at the time. 1911 is the first year for which well-founded estimates of output per capita in Italy's 16 *regioni* are available. (No estimates have been attempted for the 69 provinces.) The most recent estimates (Felice 2006) indicate that the Northwestern regions enjoyed a 24% superiority over the national average at this time. The agricultural South and Islands collectively lagged 16% behind, while the Northeast and Center hovered just at the average. Northwestern Liguria was the

Table 2. Convergence Statistics by Geographic Division and Birth Cohort

Period	Region	μ_0	$\Delta\mu$	ρ	β	$\sigma(\mu_0)$	$\Delta\sigma$
1855–1910	Italy	161.95	3.10	–0.28	–0.13	2.14	–0.04
1855–1910	North	163.38	3.40	–0.80	–0.73	1.23	–0.47
1855–1910	Center	162.86	2.52	–0.20	–0.11	1.43	0.06
1855–1910	South	159.63	3.07	–0.50	–0.37	1.24	–0.13
1855–1880	Italy	161.95	1.41	–0.59	–0.47	2.14	–0.35
1855–1880	North	163.38	1.27	–0.76	–1.02	1.23	–0.41
1855–1880	Center	162.86	1.15	–0.36	–0.38	1.43	–0.08
1855–1880	South	159.63	1.74	–0.62	–0.89	1.24	–0.26
1880–1910	Italy	163.36	1.69	0.26	0.19	1.79	0.31
1880–1910	North	164.65	2.13	–0.47	–0.66	0.81	–0.06
1880–1910	Center	164.01	1.37	–0.07	–0.07	1.34	0.14
1880–1910	South	161.37	1.33	–0.03	–0.03	0.98	0.13

Notes: μ_0 is the initial average of the provincial means; μ_1 is the final average of the provincial means; $\Delta\mu = \mu_1 - \mu_0$; ρ is the correlation coefficient between $\Delta\mu$ and μ_0; β is the slope in the regression of $\Delta\mu$ on μ_0; $\sigma(\mu_0)$ is the standard deviation of the initial provincial means; and $\Delta\sigma$ is the change in the standard deviation of the provincial means. The βs for the 1855–1880 and 1880–1910 subperiods have been standardized to a common 56-year time interval for comparability with the full-period estimates.

Figure 13. Contributions of the Within-Province and the Between-Province Components to the Age-20 Adjusted Variance of Height, by Birth Cohort

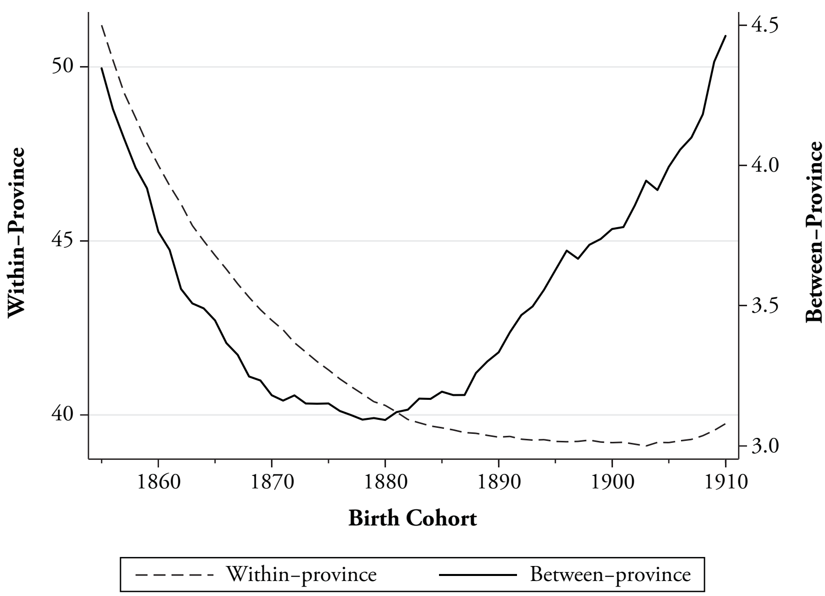
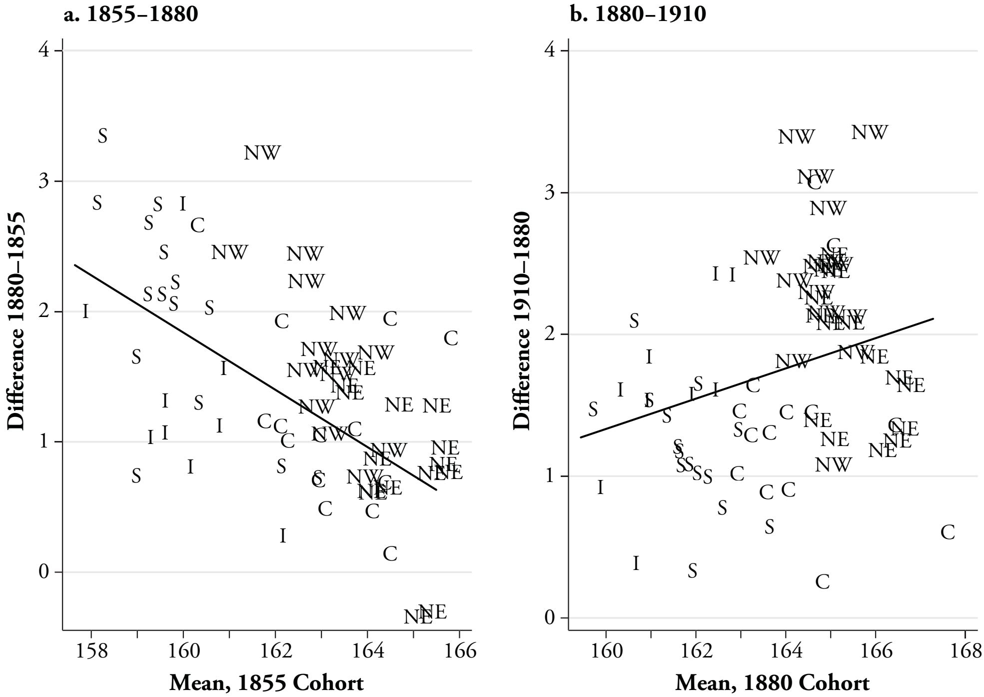


Figure 14. Convergence of Age-20 Adjusted Provincial Mean Height, 1855–1910 Birth Cohorts



highest-income region and also joint tallest in 1911, while Basilicata in the South was second poorest and also second shortest. Though regional rankings by income and height broadly match, a closer look reveals some anomalies. Veneto, in the Northeast, was joint tallest of the regions but also the poorest of the North in 1911, while the island of Sardegna was the shortest but not far below the national average in income and above Veneto's income level.

More industry and, even, greater output per capita do not necessarily generate superior health outcomes. The opposite can be true, for example, when rapid urbanization overwhelms public sanitation systems or causes an increase in the relative price of a nutritious diet. The height evidence suggests that the benefits of higher incomes and the diets, living conditions, and public health measures they could support outweighed such disadvantages in the North circa 1910. Mortality rates offer the only alternative indicator of health outcomes available at a disaggregated level for a wide range of years. Postneonatal infant mortality circa 1910, interpolated from data in Del Panta (1996), broadly matches the pattern of regional heights. Every region of the South and Islands suffered mortality exceeding the national average. Every region of the North and Center had a death rate below the national level, with the exception of Lombardy. Anomalously tall but poor Veneto appears healthy on this measure as well, with a rate of 75 per thousand, almost 20 points below the national average of 94. We focus on infant mortality after the first month because cold weather in the North dramatically increased mortality rates in the first days of life during the winter months, particularly in Veneto (Caselli 1987; Del Panta 1996).

As at the national level, then, the regional height record is broadly consistent with both economic and demographic indicators. This supports the proposition that cross-province differences in mean height are informative regarding living conditions. Our new provincial estimates suggest that there may have been considerable variation in living conditions even within regions. In Sicily, for example, mean heights range from 161.7 cm in the rugged internal province of Caltanissetta to 165.0 cm in the port city of Messina.

There is much less data and considerably less agreement about regional economic trends in earlier periods. In the absence of firm quantitative foundations, interpretations of regional developments after Italian Unification in 1861 have been derived more from theory than evidence. And they have tended to focus on broad North-South comparisons. One school of thought has seen regional disparities as growing from deep institutional or geographical roots and being significant already at the time at the time of Unification. An oft-cited estimate by Eckaus (1961) puts the initial North-South difference in income per capita at 15%–25%. After Unification, disparities gradually widened as the North built on its initial advantages. An alternative view sees North-South divergence as commencing only with Unification, in fact being caused directly by it. In some accounts, fiscal policy drained resources from the South to finance investments that favored Northern industry, which also benefited from tariff protection. Recent studies by Fenoaltea (2006: chap. 6) and Daniele and Malanima (2007) attempted to quantify regional economic disparities prior to 1911. They argued that a North-South gap emerged only 20 years after Unification, beginning in the 1880s with an acceleration of industrialization in the North.

As for demographic indicators, the available data indicate that the North-Center's advantage in postneonatal mortality dates back to at least the 1860s (Del Panta 1996). Veneto in particular had the lowest regional mortality rate among both infants aged 2–12 months and children aged 1–5 years. The advantage for Northern newborns was even more pronounced at this early date, however. Though mortality in the first month of life declined everywhere in the decades following Unification, progress was most dramatic in the North.

The 1855–1880 convergence in provincial mean heights fits none of the accounts of regional economic development. None envisions a period in which Southern regions benefited disproportionately from economic growth, from public infrastructure investment or, for that matter, from exogenous changes in the disease environment or relative price of protein-rich food. The possibility of strong nonlinearity in the relationship between height and net nutrition means that we cannot rule out a generalized, proportionate improvement in living conditions producing convergence. Still, the new evidence on heights suggests that a reassessment of trends in regional economic development in the first decades after Unification may be warranted. The post-1880 divergence of mean heights, by contrast, fits with most of the economic historiography on the period. It is worth reiterating that heights in the South (and throughout Italy) continued to improve after 1880. Though height and health cannot be presumed to track income unerringly, this suggests the economic decline of the South was only relative to a rapidly growing North (Vecchi and Coppola 2006).

CONCLUSIONS

Studies of human stature have frequently relied on the explicit assumption that heights follow a normal distribution and the typically implicit premise that covariates affect only the mean of the distribution. The evidence developed here, on the basis of a modeling technique that eschews such *a priori* restrictions, indicates that both assumptions are problematic.

In particular, the adolescent growth spurt (AGS) has been shown to systematically distort the distribution of heights by increasing dispersion, negative skewness, and excess kurtosis. In the conditions of net nutritional stress typical of the period studied, the effects of the AGS were maximal in the late teens and lingered into the early 20s.

Our semiparametric model of age effects on the height distribution offers more than just a way of controlling for age at measurement; it generates substantive insights into the effects of environmental conditions on human growth. Using our province-specific estimates to construct counterfactual height distributions at different ages, we are able to generate age-height profiles. These profiles indicate that the AGS was systematically later in provinces with shorter mean heights, implying environmental rather than purely genetic influences on provincial mean stature. They also show that improving net nutrition not only increased final height but also led to an earlier onset of puberty and the AGS. Our age-specific estimates of trends in the higher moments of the distribution at the national level confirm this finding. They indicate that the period of maximum AGS-induced distortion of the distribution shifted from above to below 18 years of age.

Our age-adjusted, smoothed, provincial height distributions also offer new insights into regional economic development in nineteenth century Italy. Both a national variance decomposition exercise and a provincial analysis of beta and sigma convergence reveal a "regime switch" from convergence prior to 1880 to divergence thereafter. Post-1880 divergence accords with the perception of widening economic disparities between North and South in this period. The continued increase of heights everywhere even in this period is a useful reminder that regional divergence was a relative rather than an absolute phenomenon. Pre-1880 convergence in height is an entirely novel finding, one indicating that presumptions of constant or widening economic disparities may need to be reassessed.

REFERENCES

A'Hearn, B. 2004. "A Restricted Maximum Likelihood Estimator for Historical Height Samples." *Economics and Human Biology* 2:5–19.

Akachi, Y. and D. Canning. 2007. "The Height of Women in Sub-Saharan Africa: The Role of Health, Nutrition, and Income in Childhood." *Annals of Human Biology* 34:397–410.

Arcaleni, E. 2006. "Secular Trend and Regional Differences in the Stature of Italians, 1854–1980." *Economics and Human Biology* 44:24–38.

Borjas, G.J. 1987. "Self-Selection and the Earnings of Immigrants." *American Economic Review* 77:531–53.

Caselli, G. 1987. *Mortalità e Sopravvivenza in Italia dall'Unità agli Anni '30* [Mortality and survival in Italy from the Unification to the 1930s]. Bologna: CLUEB.

Costanzo, A. 1948. "La Statura degli Italiani Ventenni Nati dal 1854 al 1920" [The stature of 20-year-old Italian males born from 1854 to 1920]. *Annali di Statistica Serie VIII* 2:59–123.

Daniele, V. and P. Malanima. 2007. "Il Prodotto delle Regioni e il Divario Nord-Sud in Italia (1861–2004)" [Regional output and the North-South divide in Italy (1861–1920)]. *Rivista di Politica Economica* 97:267–315.

Danubio, M.E., E. Amicone, and R. Vargiu. 2005. "Height and BMI of Italian Immigrants to the USA, 1908–1970." *Economics and Human Biology* 3:33–43.

Deaton, A. 2007. "Height, Health, and Development." *Proceedings of the National Academies of Science* 104:13232–37.

———. 2008. "Height, Health, and Inequality: The Distribution of Adult Heights in India." *American Economic Review Papers and Proceedings* 98:468–74.

Del Panta, L. 1996. "Dal Settecento ai Giorni Nostri" [From the '700 to the present days]. Pp. 133–212 in *La Popolazione Italiana dal Medioevo ad Oggi* [The Italian population from the middle age to now], edited by L. Del Panta, M. Livi Bacci, G. Pinto, and E. Sonnino. Bari: Laterza.

Eckaus, R. 1961. "The North-South Differential in Italian Economic Development." *Journal of Economic History* 21:285-317.

Eveleth, P. and J.M. Tanner. 1990. *Worldwide Variation in Human Growth,* 2nd ed. Cambridge, UK: Cambridge University Press.

Federico, G. 2003. "Heights, Calories and Welfare: A New Perspective on Italian Industrialization, 1854–1913." *Economics and Human Biology* 1:289–308.

Felice, E. 2006. “Income and Human Development: Measuring Regional Disparities in Income.” Paper presented at the 2nd RTN workshop, “Unifying the European Experience: Historical Lessons of Pan-European Development.” University of Lund, Sweden, October 13–15.
Fenoaltea, S. 2006. *L’Economia Italiana dall’Unità alla Grande Guerra* [The Italian economy from the Unification to World War I]. Rome: Laterza.
Fogel, R. 1994. “Economic Growth, Population Theory, and Physiology: The Bearing of Long-Term Processes on the Making of Economic Policy.” *American Economic Review* 84:369–95.
Hatton, T.J. and J.G. Williamson. 1998. *The Age of Mass Migration: Causes and Economic Impact*. Oxford: Oxford University Press.
Hermanussen, M., J. Burmeister, and V. Burkhardt. 1995. “Stature and Stature Distribution in Recent West German and Historical Samples of Italian and Dutch Conscripts.” *American Journal of Human Biology* 7:505–15.
Istat. 1958. *Sommario di Statistiche Storiche* [Summary of historical statistics]. Rome: Istituto Centrale di Statistica.
Limpert, E., W.A. Stahel, and M. Abbt. 2001. “Log-Normal Distributions Across the Sciences: Keys and Clues.” *Bioscience* 51:341–52.
Livi, R. [1896] 1905. *Antropometria Militare* [Military anthropometry]. Parts I and II. Rome: Tipografia E. Voghera.
Peracchi, F. 2002. “On the Estimation of Conditional Quantile and Conditional Distribution Functions.” *Computational Statistics and Data Analysis* 38:433–47.
Silventoinen, K. 2003. “Determinants of Variation in Adult Body Height.” *Journal of Biosocial Science* 35:263–85.
Snedecor, G.W. and W.G. Cochran. 1989. *Statistical Methods*. Ames, IA: Iowa University.
Soltow, L. 1992. “Inequalities in the Standard of Living in the United States, 1798–1895.” Pp. 121–66 in *American Economic Growth and Standards of Living Before the Civil War,* edited by R.G. Gallman and J.J. Wallis. Chicago: University of Chicago Press.
Tanner, J.M. 1981. *A History of the Study of Human Growth.* Cambridge, England: Cambridge University Press.
———. 1990. *Foetus Into Man: Physical Growth From Conception to Maturity.* Cambridge, MA: Harvard University Press.
Terrenato, L. and L. Ulizzi. 1983. “Genotype-Environment Relationships: An Analysis of Stature Distribution Curves During the Last Century in Italy.” *Annals of Human Biology* 10:335–46.
Vecchi, G. and M. Coppola. 2006. “Nutrition and Growth in Italy, 1861–1911. What Macroeconomic Data Hide.” *Explorations in Economic History* 43:438–64.
Wachter, K. and J. Trussell. 1982. “Estimating Historical Heights.” *Journal of the American Statistical Association* 77:279–303.