Infant mortality, income and adult stature in Spain

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Abstract

This paper presents new evidence on the relationship between infant mortality at the year of birth and adult stature using regional data for five cohorts in Spain, born between 1969 and 1986, a period of significant economic and social transformation. Consistent with previous studies, we find that there is a strong negative correlation between infant mortality and adult height, even after controlling for: secular changes affecting both infant mortality and adult height, constant differences across regions, and economic conditions at birth. Interestingly, we do not find a role for either GDP per capita or income inequality in the year of birth in explaining average cohort height after accounting for infant mortality in the year of birth. Disease, not income, appears to have been the constraining factor in Spain, at least after 1969, suggesting that the burden of disease in childhood can have long-lasting effects on health, reflected in differences in adult stature. Our results resonate on recent empirical findings for developed and developing countries, and suggest that the epidemiological transition in the 20 years leading to Spain's entry into European Union led to subsequent improvements in adult height.

Keywords: Mortality; Height; Income; Spain.
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1. Introduction

There is a growing interest by economists on the determinants of height (Steckel, 2009), a well-known biological indicator of the populations (Fogel, 2004; Komlos and Baten, 1998). The attention that this topic generates can be explained by the existing empirical evidence documenting a systematic relationship between adult height, income and life expectancy (Waaler, 1984; Leon et al., 1995; Jousilahti et al., 2000; Battya et al., 2009). Taller individuals tend to earn more and this "height premium" appears to be linked to early (cognitive) childhood outcomes (Case and Paxson, 2008) or to the positive association between height and physical capacity (Lundborg et al., 2009). Moreover, taller individuals are also more likely to report positive emotions such as enjoyment and happiness (Deaton and Arora, 2009), mainly because they have higher incomes (and more education).

An emerging body of evidence links the determinants of adult height with early conditions in life. Recently, Bozzoli, Deaton and Quintana-Domeque (2009) unveil that across a range of European countries and the United States there is a strong inverse relationship between postneonatal (one month to one year) mortality and the mean height of those children as adults. Hatton (2009) using town-level panel data on the heights of school children in Britain also finds a negative effect of infant mortality on children's height.

This paper investigates the relationship between disease environment in early life (or childhood disease, proxied by infant mortality in the year of birth) and adult height using a panel of 5 birth-cohorts (1969-72, 1975-77, 1978-80, 1981-83 and 1984-86) and 17 Spanish regions (CCAA, Comunidades Autónomas). These cohorts experienced a gradual increase in adult height in European countries, and Spain was the country (among Southern European countries) that experienced the greatest gains in stature with respect to Northern Europe countries (Garcia and Quintana-Domeque, 2007). The nearly two decades between 1969 and 1986 constitutes a period of pronounced economic and social transition from the dictatorial regime of Franco to Spain's entry in the European Union: Spanish GDP per capita grew at an average compounded rate of 3.0% per year, about 1 percentage point higher that growth rates experienced by both Western European countries and the United States during this period. Even more salient is the sharp drop in infant mortality rates: in 1969 the probability of a Spanish child dying in the first year of life was about 2.5 times higher than that of a child born in England and Wales (31 vs. 13 per 1,000 live births); in contrast, by 1987 Spain had lowered its infant mortality rate below the levels achieved by England and Wales (8.8 vs. 9.4 deaths per 1,000 births). Adult height increased from 168.5 cm from those born in 1969 to 171.1 cm for those born in 1986 (own calculations based on the Encuesta Nacional de Salud 2006). These rapid socioeconomic changes make Spain the ideal case study to analyze the role of economic and health conditions in childhood in predicting adult height outcomes, which may also be relevant for countries experiencing similar transitions.

We find that child mortality is indeed a strong predictor of future height. A reduction of infant mortality rate of 20 units per 1,000 (similar to that experienced by Spain between 1969 and 1986) explains an increase in height of about 1.6 cm, about 60 percent of the gain in adult stature during this period. Assuming that infant mortality is a good proxy for the disease environment in the first year of life, our results suggest that conditions in early life substantially affect lifetime health. Interestingly, this relationship survives when we include GDP per capita in the year of birth. In fact, GDP per capita results to be non-statistically significant in explaining adult height, indicating that all the explanatory power of income is already captured by infant mortality dynamics.

This paper complements the existing previous research in at least three different ways: first, exploiting regional rather than cross-country variation; second, providing a simple procedure to adjust average regional height when this includes individuals born in different regions; and third, exploring the association between income inequality and average height.

The rest of the paper is organized as follows. Next section presents a brief literature review on infant mortality and adult stature. Section 3 presents the sources of data. Section 4 discusses patterns in infant mortality and adult height and presents the main results. Section 5 presents some robustness checks. Section 6 concludes.

2. Infant mortality and adult stature

Epidemiologists and human biologists have studied the importance of adult height and its determinants (e.g., Eveleth and Tanner, 1990; Bogin, 2001; Silventoinen, 2003). This large body of research concludes that the average stature of an (adult) population reflects the cumulative impact of net nutrition, which depends on the intake of nutrients, the nutritional requirements, and the ability of the body to absorb them. Of course, these in turn depend on physical exercise, environmental conditions and the health status of the individual. Further, inherited and genetic factors are also relevant in explaining individuals’ height (e.g., Estrada et al., 2009). However, environmental circumstances (e.g., disease environment) and socioeconomic conditions (e.g., income) at birth have been shown to be very important determinants of the stature of a population. In particular, after a review of studies covering populations in Europe, New Guinea, and México, Malcolm (1974) concludes that differences in average height between populations are almost entirely the product of the environment.

It is important to emphasize that the fact that a population has foodstuff resources or a high level of income to buy them (gross nutrition) does not necessarily imply that net nutrition (and their height) will be higher. The incidence of disease may prevent richer populations from fully achieving their biological potential stature: the prevalence of gastrointestinal diseases during childhood (and/or adolescence), for instance, may prevent richer populations to absorb nutrients efficiently during the important stages of human growth, which will be translated in shorter individuals.

Usually, infant mortality rate (or post-neonatal mortality when the data are rich enough) is considered a proxy for exposure to infections among survivors (e.g., Forsdahl, 1977), and it has been found to be a strong predictor of the average (adult) height of the survivors (e.g., Sobral, 1990). Other studies have used different components of infant mortality rate (most prominently postneonatal mortality rate, PNM, the ratio of deaths among infants 28-364 days of age over live births) to measure the negative effect of the burden of disease in infancy on adult height (e.g., Schmidt , Jørgensen, and Michaelsen, 1995; Bozzoli, Deaton and Quintana-Domeque, 2009).

Previous work using Spanish data for the twentieth century has found a correlation between health (infant mortality and life expectancy), wealth, income and height. Most of this research has been conducted by Martínez Carrión (1994, 2002 and 2005) using height data from the Spanish military records. Martínez Carrión (2002) highlights the divergence between economic wellbeing and physical welfare in the beginning of the modern economic growth in Spain: income by itself cannot explain the evolution of height. He also finds that the higher exposure of children to infectious disease could leave them suffering from the after-effects, reflected in the bodies of those who survived, something referred as scarring or debilitation effect in Bozzoli, Deaton and Quintana-Domeque (2009).

3. Data

We compute average cohort heights using data from the Encuesta Nacional de Salud 2006 (ENS, Spanish National Health Survey 2006). The ENS is a national representative survey that contains information regarding several demographic and health characteristics, collected between June 2006 and June 2007. For the purposes of our work, the ENS has a question about individual height, country of birth, region of current residence, age and gender.

Specifically, the ENS contains the following question about individual height: "*¿Y cuánto mide, aproximadamente, sin zapatos*?" (Approximately, what is your height without shoes?). Since we are interested in exploring the relationship between average attained height by cohort and infant mortality at the year of birth, we focus on individuals aged 20 and above in 2006-2007, e.g., those who have already attained their adult stature by the time the survey was carried out. Hence, we do not use data for the period after 1986. This gives rise to a panel of 5 birth-cohorts (1969-72, 1975-77, 1978-80, 1981-83 and 1984-86) for the 17 Spanish regions (CCAA, Comunidades Autónomas).

Individuals self-report their heights, and hence we must be aware of potential measurement error. Thomas and Frankenberg (2002) and Ezzati et al. (2006) show that, in the United States, self-reported heights exaggerate actual heights on average, but the difference is close to constant for ages 20 to 50. Our cohorts are born between 1969 and 1986, so they are 20-37 years old when reporting their heights. If we assume that patterns of height self-reporting in Spain are similar to those in the United States, the (average) difference between actual and self-reported height will be captured by the constant term in the regression.

Given the information in the ENS, we can exclude individuals who are born outside of Spain from our sample. However, we have no information on the individual's region of birth within Spain. This is an important point, already noted by Martínez-Carrión (2002), who alerts of the possibility that individuals who migrate are, on average, taller, but we postpone its discussion to the robustness checks section.

Adult heights for each cohort and region (CCAA) are first computed for women and men separately and then averaged out. This reduces the chances of imprecision in the case of small cohorts where the ratio of females to males can be quite different from its population counterpart.

Infant mortality rates at the regional level (CCAA) come from mortality tables published in Instituto Nacional de Estadística (1988) (covering the periods 1969-72 and 1974-77), and from the online TEMPUS time series database maintained by the Instituto Nacional de Estadística (INE, National Statistical Institute of Spain, http://www.ine.es, in annual frequency for the period 1975-88). For the cohort 1969-72 we use data from Instituto Nacional de Estadística (1988), and for the cohorts starting in 1975 we compute averages from the TEMPUS time series database. Although these two sources do not apply exactly the same methodology, the differences between both series are unsystematic. Indeed, in a preliminary analysis not reported here, we could not reject that the average difference in mortality rates between these two sources is just classical measurement error. More information for the mortality tables can be found in Goerlich and Pinilla (2006).

It would be interesting to perform the analysis decomposing infant mortality rate (IMR) on neonatal mortality (NNM) and postneonatal mortality (PNM), however we have not been able to find data on either NNM or PNM disaggregated at the regional level, although Gutiérrez and Regidor (1993) document trends in NNM and PNM in Spain (at the country level) for the period 1975-88.

GDP regional data for the year of birth come from de la Fuente (2009), and uses regional price levels to deflate nominal values.

4. Results

Table 1 presents the data for IMRs (infant mortality rates, per 1,000 live births) and average adult heights (in 2006-07) for five birth cohorts in 17 Spanish regions (CCAA). A couple of facts merit attention. First, the average infant mortality declines substantially during the period 1969-86 from 29 to 10 per 1,000 births, while adult height increases for the 5 cohorts analyzed from 169 to 171 centimeters. Second, there is substantial cross sectional variation in both measures: for the period (cohort) 1969-72, Catalunya has the lowest IMR, 21 per 1,000, while Extermadura has the highest one, 37 per 1,000. This cross sectional variation decreases over time: in the period (cohort) 1984-86, the highest IMRs are found in Galicia, Navarra and La Rioja, 12 per 1,000, while Catalunya and Cantabria have the lowest ones, 7 per 1,000. There is also substantial cross-sectional variation in average height: the average height for the 5 cohorts in Canarias, Madrid, Murcia and Navarra is 171 cm, while in Asturias it is 168 cm, a difference of 3 cm.

[Insert Table 1 about here]

We now explore the relationship between the main variables in the paper. Figure 1 plots the mean height (cm), the infant mortality rate, and the log GDP per capita in our data. It shows a negative relationship between mean height and IMR. The figure also allows us to perform an eyeball test to infer the signs of the relationships between log GDP per capita and both average height and IMR: the first one is positive; the second one negative.

[Insert Figure 1 about here]

Further, Table 2 shows a correlation matrix that displays the pairwise correlation coefficients for the variables average adult height, IMR and log GDP. The signs of the pairwise correlation coefficients are the expected ones after looking at the 3 graphs in Figure 1. The correlation coefficient between IMR at the year of birth and average adult height is –0.572 and it is statistically significant at the 1% level, suggesting a negative relationship between IMR and height. The correlation coefficient between log GDP at the year of birth and adult height is 0.344 and it is also statistically significant at the 1% level. Finally, the correlation between contemporaneous IMR and log GDP is –0.542 (also statistically significant at the 1% level).

[Insert Table 2 about here]

Table 3 presents the main results of the paper. It contains a series of regressions in which mean population height is the dependent variable. Column (1) shows that in the 85 observations (17 regions × 5 cohorts) pooled time-series cross-section, variation in infant mortality explains almost one half of the variation in average height. The parameter estimate is –0.10, so that a reduction in infant mortality by 20 per 1,000, which is the one observed for Spain in the period, is associated with an increase in average height by 2 centimeters, which is more or less the actual increase shown in Table 1. Column (2) includes a time trend (mean cohort year centered around 1978), to capture secular changes affecting both infant mortality and adult height. This last variable is not statistically significant, but reduces the estimate of IMR to –0.07. Note that the standard error of this new estimate is 2.5 times larger than before. Column (3) replaces the time trend with cohort dummies. The parameter estimate on IMR is reduced to –0.05, but both this estimate and the cohort effects appear to be imprecisely estimated, so none of them result to be statistically significant (the F-test for the joint significance of the cohort dummies is 1.32, so we cannot reject the null hypothesis of absence of cohort effects). Moreover the adjusted-R^2 in column (3) is smaller than that displayed in column (2). The fact that we do not find a statistically significant effect of IMR once we exploit within-cohort variation is not surprising. First, we lose most of our variation in average height. Second, the precision of our estimate decreases substantially. Finally, the attenuation bias of white noise measurement error is exacerbated. Column (4) differs from column (1) in that the former includes region (CCAA) dummies. Interestingly, while the (adjusted) explanatory power of our regression (adjusted-R^2) increases from 46% to 58%, there is no significant change in the coefficient on IMR. Adding a time trend to specification (4) has no significant change on the coefficient attached to IMR either, as shown in column (5).

[Insert Table 3 about here]

Columns (6) to (8) show the same specifications as in columns (2), (4) and (5) with the inclusion of log GDP per capita at the year of birth. Three important observations deserve special attention. First, the qualitative results for infant mortality remain unchanged, with the point estimates ranging from –0.055 in column (6) to –0.074 in columns (7) and (8). Second, controlling for log GDP at the year of birth does neither add much to the explanatory power of our regressions –the adjusted-R^2s are lower (or the same) than those excluding log GDP displayed columns (2), (4) and (5). Third, we do not find a role for GDP per capita in the year of birth in explaining average cohort height after accounting for infant mortality in the year of birth. Hence, although log GDP and average height are positively correlated (Table 2), once we control for IMR the correlation disappears. Note that although not statistically significant, we could expect a positive relationship between average height and log GDP if, first, regional income (or resources) is well proxied by log GDP per capita and, second, higher regional income (resources) leads to better nourished populations or populations with more resources to fight against diseases or both, and hence to taller populations.

Our results suggest that it is the decrease in disease (IMR), and not the increase in income (GDP), what explains the increase in adult stature in Spain between 1969 and 1986. This finding is entirely consistent with both previous research on cohorts born between 1950 and 1980 for Europe and the United States (Bozzoli, Deaton and Quintana-Domeque, 2009) and historical evidence in Spain (Martínez Carrión, 2002).

Martínez Carrión (2002) already acknowledges that in Spain the increase in stature has been much more strongly linked to the drop in (infant) mortality than to the modern economic growth (in terms of income or GDP per capita). The emphasis in the relationship between average income and average stature at the population level comes from the exploratory essay by Steckel (1983), who shows a strong correlation between average height and income per capita in both developed and developing countries during the mid-twentieth century. However, recent studies show that there are divergences between height and conventional economic indicators (e.g., Geissler, 1993). As pointed out by Martínez Carrión (1994), the case of UK clearly shows that there is no clear and direct connection between average income and average stature at the population level: the drop in stature of British conscripts in the mid nineteenth century does not correlate well with the increase in the national income (or GDP per capita).

Quantitatively, and bearing in mind that we are using IMR rather than PNM, our results are similar to those obtained in Bozzoli, Deaton and Quintana-Domeque (2009). The authors report that in the 316 pooled time-series cross-section observations for 10 continental European countries, the United Kingdom and the United States over 31 years of birth (1950-1980), variation in postneonatal mortality by itself explains 62% of the variation in average height. Their parameter estimate is –0.16, so that a reduction in PNM by 20 per thousand is associated with an increase in average height by 3.2 cm.

Although our findings appear to be very sensible, both qualitatively and quantitatively, there are at least three important caveats that one should take into account when interpreting them. First, we do not have data on region of birth but region of current residence. Hence, we need to be aware of potential selective migration. Second, there could be omitted variables that are related to IMR, GDP and average height, such as income inequality. Finally, there could be nonmonotonicites (or nonlinearities) in the mortality-height relationship. We investigate the extent to what our results may be biased by these factors in the next section.

5. Robustness checks

5.1. Selective migration

One important feature of our data is that we can measure adult height of those currently living in a region but we cannot identify the individual region of birth. If taller people born in poor regions systematically migrated to better regions (i.e., regions with lower IMR and/or higher GDP in the year of birth), we may find a negative (positive) relationship between average regional height (which includes the heights of those born in other regions) and regional IMR (log GDP) in the year of birth.

There is some evidence of positive health selectivity in the demography and epidemiology literatures: migrants tend to be in better health than non-migrants (e.g., Marmot, Adelstein and Bulusu, 1984; Jasso et al., 2004; Swallen, 2002). Hence, we cannot discard the possibility that individuals in better health (and taller) tend to migrate to regions with low infant mortality rates (and high gross domestic products). Indeed, Quiroga (2001) realizes that in Spain conscripts (military) who migrated from underdeveloped regions to developed ones were (on average) taller than those who remained in their region of birth. Nevertheless, Bentolila (1997) discusses the puzzle of comparatively low regional migration rates coexisting with high unemployment, finding that migration from poorer to richer regions has dwindled over time, and even being compensated by "affluent" migration from richer to poorer regions.

Suppose that we have data on $c = 1, \dots, C$ cohorts and $r = 1, \dots, R$ regions. Assume a simple model stating a (linear) relationship between the average height of a cohort c born in a region r, $\bar{h}^c_{r,r}$, and the infant mortality in that region in the cohort-year of birth, IMR^c_r:

$$\bar{h}^c_{r,r} = \alpha + \beta IMR^c_r + \varepsilon^c_r \tag{1}$$

Provided that the model is well specified, OLS provides a consistent estimate of β:

$$plim\hat{\beta}_{OLS} = \frac{cov(\bar{h}^c_{r,r}, IMR^c_r)}{var(IMR^c_r)} = \beta \tag{2}$$

Unfortunately, we do not observe (and we cannot compute) the average height of a cohort born in a region, $\bar{h}^c_{r,r}$, with the ENS data, but the average height of a cohort currently residing in a region, $\bar{h}^c_r$. Nevertheless, we know the relationship between these averages. Indeed, the average height of a cohort c in a region of residence r can be expressed as a weighted average of the mean height of a cohort c who was born in the same region as the one of current residence, $\bar{h}^c_{r,r}$, and the average mean height of a cohort c who was born in a different region than the one of current residence, $\bar{h}^c_{-r,r}$:

$$\bar{h}^c_r = \lambda^c_{r,r}\bar{h}^c_{r,r} + (1 - \lambda^c_{r,r})\bar{h}^c_{-r,r} \tag{3}$$

where $\lambda^c_{r,r}$ is the fraction of individuals of cohort c living in region r who were born in region r. Notice that $\left(1 - \lambda^c_{r,r}\right) = \sum_{\substack{j=1 \\ j\neq r}}^{R} \lambda^c_{j,r}$, where $\lambda^c_{j,r}$ is the fraction of individuals of cohort c living in region r who were born in region j. Hence, $\bar{h}^c_{-r,r} = \sum_{\substack{j=1 \\ j\neq r}}^{R} \left(\frac{\lambda^c_{j,r}}{\sum_{\substack{j=1 \\ j\neq r}}^{R} \lambda^c_{j,r}} \bar{h}^c_{j,r} \right)$.

We can adjust $\bar{h}^c_r$ dividing it by $\lambda^c_{r,r}$, as long as $0 < \lambda^c_{r,r} \leq 1$:

$$\frac{\bar{h}^c_r}{\lambda^c_{r,r}} = \bar{h}^c_{r,r} + \left(\frac{1 - \lambda^c_{r,r}}{\lambda^c_{r,r}}\right)\bar{h}^c_{-r,r} \tag{4}$$

Hence, we can estimate:

$$\frac{\bar{h}^c_r}{\lambda^c_{r,r}} = \gamma + \delta IMR^c_r + u^c_r \tag{5}$$

It is straightforward to show that:

$$plim\hat{\delta}_{OLS} = \beta + \frac{cov\left(\left(\frac{1 - \lambda^c_{r,r}}{\lambda^c_{r,r}}\right)\bar{h}^c_{-r,r}, IMR^c_r\right)}{var(IMR^c_r)} \tag{6}$$

Equation (6) illustrates the bias due to selective migration: inter-regional migration (residential mobility across regions) is selective when both the *fraction* of a cohort of individuals living in region r who were born in other regions and their *average height* are correlated to the *IMR* in the year of birth of the region where they actually live.

Modeling selective migration is not straightforward in our setting, with both multiple regions (seventeen) and multiple potential determinants of migration (height, mortality and income). This is beyond the scope of our paper. However, we can use a reduce-form approach (see Halliday and Kimmitt (2008) for micro-foundations of this reduced-form model).

Suppose that given a (behavioral) model of selective migration, we can obtain the following (statistical) model of average height for immigrants to region *r* coming from $-r$ (all regions other than *r*):

$$\bar{h}^{c}_{-r,r} = \pi + \Psi(IMR^{c}_{r}) + v^{c}_{r} \tag{7}$$

where Ψ is a continuous differentiable function of IMR^{c}_{r}. If $\Psi' = 0$, average height of immigrants to region *r* is unrelated to infant mortality in the region of destiny. Let's assume that $\Psi = \theta IMR^{c}_{r}$, so

$$\bar{h}^{c}_{-r,r} = \pi + \theta IMR^{c}_{r} + v^{c}_{r} \tag{8}$$

where θ reflects the effect of selective migration on average height of immigrants.

Plugging equations (1) and (8) into equation (4), we obtain the following equation:

$$\frac{\bar{h}^{c}_{r}}{\lambda^{c}_{r,r}} = \alpha + \beta IMR^{c}_{r} + \pi\left(\frac{1-\lambda^{c}_{r,r}}{\lambda^{c}_{r,r}}\right) + \theta\widetilde{IMR}^{c}_{r} + \varepsilon^{c}_{r} + \left(\frac{1-\lambda^{c}_{r,r}}{\lambda^{c}_{r,r}}\right)v^{c}_{r} \tag{9}$$

where $\widetilde{IMR}^{c}_{r} = \left(\frac{1-\lambda^{c}_{r,r}}{\lambda^{c}_{r,r}}\right)IMR^{c}_{r}$.

Of course, we cannot discard the possibility that selective migration is based not only on IMR but also on GDP in the region of destiny at the year of birth. In particular, a more general (statistical) model of environmental conditions at birth and adult stature accounting for selective migration can be described by:

$$\bar{h}^{c}_{r,r} = \alpha + \beta IMR^{c}_{r} + \rho log\,(GDP^{c}_{r}) + \varepsilon^{c}_{r} \tag{10}$$

$$\bar{h}^{c}_{-r,r} = \pi + \theta IMR^{c}_{r} + \mu\log\,(GDP^{c}_{r}) + v^{c}_{r} \tag{11}$$

Plugging equations (10) and (11) into equation (4), we obtain the following equation:

$$\frac{\bar{h}^{c}_{r}}{\lambda^{c}_{r,r}} = \alpha + \beta IMR^{c}_{r} + \rho\, log(GDP^{c}_{r}) + \pi\left(\frac{1-\lambda^{c}_{r,r}}{\lambda^{c}_{r,r}}\right) + \theta\widetilde{IMR}^{c}_{r} + \mu\log(\widetilde{GDP}^{c}_{r}) + \varepsilon^{c}_{r} + \left(\frac{1-\lambda^{c}_{r,r}}{\lambda^{c}_{r,r}}\right)v^{c}_{r} \tag{12}$$

where $\widetilde{GDP}^{c}_{r} = \left(\frac{1-\lambda^{c}_{r,r}}{\lambda^{c}_{r,r}}\right)GDP^{c}_{r}$. Under no selection, $\theta = \mu = 0$.

Table 4 shows the percentages of Spaniards in 2006 by region and cohort who were born in the region. Andalucía, Asturias, Catalunya, Galicia and País Vasco have the highest percentages of individuals who were born in the region of current residence for the cohorts under analysis, while Balears, Castilla - La Mancha and La Rioja have the lowest ones.

[Insert Table 4 about here]

Table 5 contains a series of regressions in which adjusted mean population height is the dependent variable. As shown above, the regression coefficients for IMR (and GDP) can be compared with those of Table 3 without scaling. We realize that, in general, the standard errors of the estimates of the effect of IMR are much higher now. However, the results are qualitatively similar. In fact, based on equations (10) to (12) a test for non-selection of migrants ($\theta = \mu = 0$) never results in rejecting the null hypothesis, suggesting that interregional migration had no impact on the link between adult heights and conditions at birth. In light of this, results in Table 3 are preferred to those displayed in Table 5 (in which redundant variables are included).

[Insert Table 5 about here]

5.2. Income Inequality

In our models we account for economic conditions during childhood by including regional GDP per capita in the year of birth. However, this indicator although informative of the average economic conditions in a given year for a given region, it omits the effects of income distribution. Income inequality may affect average population height, if poor people are more likely to be exposed to a worse disease environment or are more likely to suffer more than rich people when exposed to the same disease environment (Steckel, 1995). Having less income means having fewer resources to fight against disease and to buy foodstuff. It also may imply to live in a poor environment, not only in economic terms, but also in terms of sanitation and hygienic measures. Hence, income inequality may affect the average height of the population by reducing the potential adult height reached by poor people while leaving unaffected the potential adult stature reached by rich people. Thus, we should expect a negative effect of income inequality on average height. Recently, Deaton (2008) has explored the relationship between income inequality and the distribution of heights in India. Although the effect of inequality on height is statistically significant in some specifications, its sign is opposite to the expected one.

Individual inequality has fluctuated in the last decades in Spain, albeit with a downward trend. During the 70s and 80s, inequality levels (proxied by the Gini index) were similar to those observed contemporaneously in the United States. The Gini index declined between 1973 and 1991, and grew between 1991 and 2001 (Goerlich and Mas, 2004). Richer regions (in terms of income per capita) have, on average, a more equal distribution of income (in terms of the Gini inequality index).

To account for economic inequality, we use the Gini inequality measure obtained from Ayala Cañón, Málaga and Chaparro (2005). Unfortunately, the available data allows us to identify intra-region inequality only for two cohorts: 1973-74 and 1980-81, which reduces the sample size from 85 to 34 observations. As expected, the correlation coefficient between IMR and income inequality is 0.49, and it is statistically significant at the 1% level. The correlations between income inequality and both average height and log GDP are negative (–0.20 and –0.24, respectively) but not statistically significant.

Table 6 investigates the role of income inequality in determining adult height. Not surprisingly, given the small sample size, the results do not indicate that income inequality at time of birth is a predictor of adult height. Column (1) shows that the significant relationship between height and IMR holds in this smaller sample. This relation is weakened by the addition of a time trend (as displayed in column (2)), which nonetheless results not statistically significant. Columns (3) to (6) report different specifications including a time trend, regional fixed effects and log GDP. In all but one of these specifications, the coefficient of IMR remains significant. None of the other regressors included in these specifications is individually statistically significant, and in terms of adjusted-R^2, column (1) still provides the best fit. Column (7) reports the results of a specification which includes log GDP and inequality together with IMR. Inequality has a statistically insignificant positive coefficient, and IMR at time of birth remains a significant predictor of adult height. Column (8) reports results adding region fixed effects, which are similar in qualitative terms to those shown in column (1). It is important to emphasize that these results are based on regressions with small sample sizes, 34 observations. The combination of a reduced sample size and loosely specified equations results in imprecise estimates, specifically in columns (5), (6) and (8) where standard errors of IMR are about 3 times larger than those displayed in the tightly specified model in column (1). Thus, results displayed in the table must be taken with a grain of salt. Nevertheless, it is reassuring that the qualitative results for IMR hold across most of these specifications.

[Insert Table 6 about here]

5.3. Non-monotonic and/or non-linear relationships

5.3.1. Selection versus debilitation

Adult height and its early life determinants may involve a non-monotonic relationship. That would be the case if, for example, child mortality operated on adult height (of survivors) through opposing scarring and selection mechanisms, with selection increasing the adult height of survivors at high child mortality levels (Deaton, 2007). The relationship between adult height and child mortality is positive (negative) for high (low) levels of mortality when the so-called “selection” (survival into adulthood) effect dominates (is dominated by) the “debilitation” (scarring on survivors) effect. Bozzoli, Deaton and Quintana-Domeque (2009) find evidence of the debilitation effect in developed countries (Europe and the United States), while they find evidence of the selection effect in the poorest and highest mortality countries of the world.

Looking at our data, Figure 2 shows the relationship between average height and IMR based on raw quadratic prediction with 95% confidence intervals, which is close to linear. Thus, we find no evidence of either non-monotonicities or nonlinearities in this relationship for Spain in the period studied.

[Insert Figure 2 about here]

5.3.2. Individual versus population relations

Another source of non-linearities is discussed in Deaton (2003): income inequality may also affect average height at the *population level* if there is a concave relationship between income and adult height at the *individual level*. Let individual height be related to individual income as described by

$$h_{i,r} = \alpha + \beta y_{i,r} + \gamma y_{i,r}^2 + \varepsilon_{i,r} \qquad (13)$$

where $h_{i,r}$ is the adult height of individual i living in region r and $y_{i,r}$ is the log of individual income. If there is a concave relationship, it is naturally to expect $\beta > 0$ and $\gamma < 0$, that is, income affects height positively but its marginal effect decreases as income increases. Taking expectations over the individuals in each region and adding and subtracting the term $\gamma \bar{y}_r^2$ equation (13) becomes:

$$\bar{h}_r = \alpha + \beta \bar{y}_r + \gamma var(y_{i,r}) + \gamma \bar{y}_r^2 + \bar{\varepsilon}_r \qquad (14)$$

Figure 3 shows traces of non-linear relationship between height and log GDP, based on raw quadratic prediction, suggesting the existence of a concave relationship between height and log GDP: the higher is log GDP, the higher is average height, but the effect seems to dwindle as log GDP increases.

[Insert Figure 3 about here]

Given the concerns of omitted variables due a potential non-linear relationship between IMR and log GDP, we run some specifications in Table 7 including quadratic terms for IMR and log GDP.

[Insert Table 7 about here]

Column (1) in Table 7 allows for a non-monotonic effect of IMR on average height. Estimates suggest that, as expected from Figure 2, the relationship is close to linear: the quadratic term is statistically non-significant. The F-statistic for the coefficients on IMR and IMR^2 equal to zero is 4.18 suggests that IMR and its squared value are jointly a strong predictor of adult height, but the effect of each of them cannot be estimated with precision. Column (2) adds the log GDP. Again, we reject the existence of a non-monotonic relationship between IMR and average height, and the relationship between IMR (and its squared value) and height becomes insignificant in the joint significance test. Column (3) shows estimates of a specification where both IMR and log GDP have its squared terms included, together with a time trend and regional dummies. In this case it is impossible to distinguish any individual effect, although a joint test (not shown) for the joint significance of IMR, IMR^2, log GDP and log GDP^2 rejects the null ($p = 0.0563$). It is possible that the sample size and the inclusion of a time trend and region fixed effects put a limit to disentangling the individual contributions of income and IMR to adult height. Finally, column (4) shows estimates of a specification with the inequality measure, together with region fixed effects and a nonlinear effect from IMR and log GDP. Although the sample size decreases dramatically, IMR and IMR^2 are jointly significant, with other variables exerting no effect on height. However, and in accordance with results in previous columns, there is no evidence of nonlinearities in this relationship. We hypothesize that even though IMR decreased markedly during the 70s and 80s, the range of variation may not be enough to detect potential nonlinearities.

There is evidence from a non-linear relationship between height and mortality, but it comes from different periods and countries. Bozzoli, Deaton and Quintana-Domeque (2009) report nonlinearities in the relationship between mortality and height only when pooling results from both developed and developing countries (some of them having infant mortality rates an order of magnitude higher than those of Spain in 1970). Schmidt , Jørgensen, and Michaelsen (1995; Figure 4) provide graphical evidence suggesting a non-linear relationship between conscript height and neonatal mortality in some European countries, but they do not report the result from nonlinear specifications. It may be possible thus, that such nonlinearities prevailed in Europe, but it does not be the case in Spain, at least not for our sampling period.

6. Conclusions

We have used data on 5 birth-cohorts from 17 Spanish communities to describe the relationship between infant mortality and adult height. Our study uses both regional and time variation, and hence provides complementary evidence for studies carried out with pooled cross-sections of countries. Our results are in line with those of others found in the literature. Infant mortality results to be a good predictor of adult height even after controlling for economic conditions. A reduction in IMR of the magnitude seen the period studied (1969-86) explains about 60% of the increment in adult height reported, or about 1.6 cm. Using information about income inequality at a regional level, we find no evidence of a relationship between this indicator and adult height, although the evidence comes from using a subset of our original data due to the lack of observations for income inequality for most of our birth-cohorts of interest. Since our results may be affected by inter-regional migration, we present a simple procedure to adjust for such possibility. We find that migration does not affect the strength or the sign of the relationship between adult height and infant mortality.

Taken altogether, these results resonate on recent empirical findings for developed and developing countries. They suggest that the large drop in infant mortality that Spain experienced in the 20 years prior to its entry in the European Union may be an important factor in explaining the improvements in current standard of living (proxied by adult height).

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Table 1: Mean heights (2006-07) and infant mortality for selected birth cohorts

(IMR: Infant mortality rate per 1,000 births, Height: mean heights of men and women, cm.)

	1969-72		1975-77		1978-80		1981-83		1984-86		Average	
	IMR	Height	IMR	Height	IMR	Height	IMR	Height	IMR	Height	IMR	Height
Andalucia	30	169	19	169	15	171	12	170	10	171	17	170
Aragón	24	169	16	170	14	171	12	170	9	171	15	170
Asturias (Principado de)	25	168	20	168	17	170	13	169	11	168	17	168
Balears (Illes)	28	169	15	169	13	172	10	170	10	170	15	170
Canarias	26	171	18	172	15	171	12	171	8	171	16	171
Cantabria	31	169	19	170	13	172	12	170	7	171	17	170
Castilla y León	33	170	22	169	16	170	11	169	10	172	19	170
Castilla - La Mancha	31	167	18	171	14	171	11	170	9	172	17	170
Catalunya	21	169	14	170	11	170	10	170	7	171	13	170
Comunidad Valenciana	24	168	16	170	13	169	11	171	8	171	14	170
Extremadura	37	168	20	169	17	171	11	170	11	169	19	169
Galicia	35	168	22	171	17	170	13	172	12	171	20	170
Madrid (Comunidad de)	26	169	14	171	11	171	12	172	8	170	14	171
Murcia (Región de)	30	169	20	170	13	171	10	171	9	172	16	171
Navarra (Comunidad Foral de)	25	169	17	172	15	170	14	173	12	172	16	171
Pais Vasco	25	168	17	170	16	169	12	172	11	169	16	170
La Rioja	33	169	17	170	16	171	14	169	12	170	19	170
Total	29	169	18	170	15	171	12	171	10	171	17	170

Source: Authors’ calculations from the ENS 2006.

Figure 1: Pairwise relation between Height, IMR and log GDP

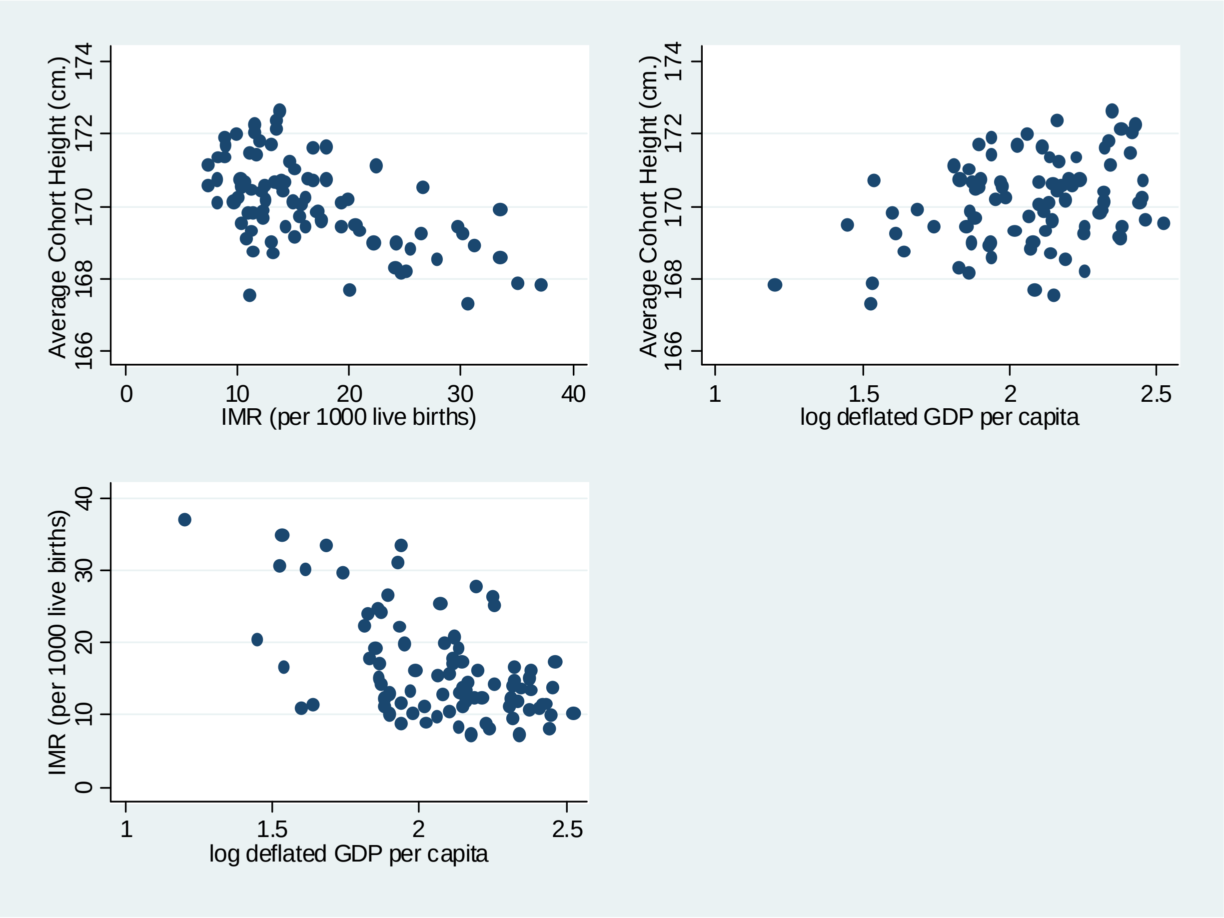

Table 2: Correlation matrix

	height	IMR	log GDP
height	--	−0.5719*** (0.0000)	0.3443*** (0.0013)
IMR	−0.5719*** (0.0000)	--	−0.5416*** (0.0000)
log GDP	0.3443*** (0.0013)	−0.5416*** (0.0000)	--

Source: Authors' calculations. p-values in parentheses. *** p-value < 0.01, ** p-value < 0.05, * p-value < 0.1

Table 3: Regressions of height on infant mortality and other variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
IMR	−0.103***	−0.068**	−0.048	−0.113***	−0.111***	−0.055*	−0.074**	−0.074
	(0.011)	(0.027)	(0.031)	(0.012)	(0.039)	(0.033)	(0.035)	(0.046)
ln(GDP)	--	--	--	--	--	0.274	2.31	2.31
						(0.459)	(2.00)	(2.03)
Time trend	--	0.061	--	--	0.003	0.072*	--	0.001
		(0.039)			(0.054)	(0.043)		(0.053)
Cohort dummies?	NO	NO	YES	NO	NO	NO	NO	NO
Region dummies?	NO	NO	NO	YES	YES	NO	YES	YES
F-test cohort dummies = 0	--	--	1.32	--	--	--	--	--
F-test region dummies = 0	--	--	--	3.78***	3.38***	--	3.87***	3.35***
R^2	0.47	0.48	0.50	0.66	0.66	0.48	0.67	0.67
Adjusted-R^2	0.46	0.47	0.46	0.58	0.57	0.46	0.58	0.57
N	85	85	85	85	85	85	85	85

Notes: Heteroskedasticity robust standard errors are reported in parentheses. Time trend is defined as the difference between mean cohort year and 1978. Observations have been weighed using the number of individual observations that gave rise to the cohort-region average.

*** p-value < 0.01, ** p-value < 0.05, * p-value < 0.1

Table 4: % of Spaniards residing in each region (CCAA) who were born in that region, by age-group (cohort), 2006

	20-24 (1982-86)	25-29 (1977-81)	30-34 (1972-76)	35-39 (1967-71)
Andalucía	94	91	90	91
Aragón	89	86	83	81
Asturias (Principado de)	94	92	91	89
Balears (Illes)	78	69	68	65
Canarias	90	85	85	87
Cantabria	91	86	81	80
Castilla y León	90	86	84	85
Castilla - La Mancha	82	75	72	74
Catalunya	95	93	90	84
Comunidad Valenciana	91	88	84	80
Extremadura	89	84	83	86
Galicia	96	94	93	93
Madrid (Comunidad de)	90	86	80	74
Murcia (Región de)	92	89	86	85
Navarra (Comunidad Foral de)	89	85	81	78
País Vasco	95	94	92	89
Rioja (La)	84	79	73	70

Source: INE.

Table 5: Regressions of adjusted height on infant mortality and other variables

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
IMR	−0.088***	−0.048	−0.106***	−0.102	−0.101***	−0.042	−0.052	−0.052
	(0.030)	(0.049)	(0.033)	(0.064)	(0.035)	(0.062)	(0.062)	(0.074)
ln(GDP)	--	--	--	--	−0.601	−0.051	3.12	3.12
					(1.193)	(1.30)	(3.44)	(3.55)
Time trend	--	0.067	--	0.006	--	0.077	--	0.000
		(0.048)		(0.068)		(0.056)		(0.069)
$\left(\frac{1-\lambda}{\lambda}\right)$	173.29***	173.50***	173.40***	173.45***	165.02***	168.94***	180.53***	180.52***
	(3.41)	(3.59)	(4.10)	(4.22)	(13.22)	(13.56)	(42.67)	(43.38)
$\widetilde{\text{IMR}}$	−0.193	−0.190	−0.148	−0.148	−0.121	−0.154	−0.212	−0.212
	(0.145)	(0.159)	(0.155)	(0.157)	(0.149)	(0.158)	(0.312)	(0.314)
$\ln(\widetilde{\text{GDP}})$	--	--	--	--	3.22	1.69	−2.95	−2.95
					(5.61)	(5.79)	(16.97)	(17.31)
Region dummies?	NO	NO	YES	YES	NO	NO	YES	YES
F-test $\theta = \mu = 0$	1.77	1.44	0.91	0.88	0.83	0.72	0.59	0.59
N	85	85	85	85	85	85	85	85

Notes: Adjusted height = $\frac{\bar{h}_{c,r}}{\lambda_{c,r}}$. Unadjusted height = $\bar{h}_{c,r}$. Heteroskedasticity robust standard errors are reported in parentheses. Time trend is defined as the difference between mean cohort year and 1978. Observations have been weighed using the number of individual observations that gave rise to the cohort-region average. *** p-value < 0.01, ** p-value < 0.05, * p-value < 0.1

Table 6: Regression of height on infant mortality and other variables: testing the role of inequality

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
IMR	−0.117***	−0.050	−0.135***	-0.108**	−0.032	−0.225**	−0.108***	−0.255**
	(0.033)	(0.063)	(0.044)	(0.041)	(0.112)	(0.089)	(0.039)	(0.093)
ln(GDP)	--	--	--	0.503	1.49	−10.24	0.499	−11.28
				(1.16)	(1.45)	(9.29)	(1.180)	(9.36)
Time trend	--	0.949	--	--	1.72	--	--	--
		(0.889)			(1.28)			
Inequality	--	--	--	--	--	--	0.292	10.84
							(7.76)	(16.07)
Region dummies?	NO	NO	YES	NO	NO	YES	NO	YES
R^2	0.24	0.26	0.58	0.25	0.29	0.61	0.25	0.62
Adjusted- R^2	0.22	0.21	0.14	0.20	0.22	0.15	0.17	0.11
N	34	34	34	34	34	34	34	34

Notes: Heteroskedasticity robust standard errors are reported in parentheses. IMR for the first cohort is the simple average of infant mortality rates for cohorts born in 1969-72 and 1974-77. Other variables calculated as stated in the main text. Observations have been weighed using the number of individual observations that gave rise to the cohort-region average.

*** p-value < 0.01, ** p-value < 0.05, * p-value < 0.1

Figure 2: Average Height and IMR

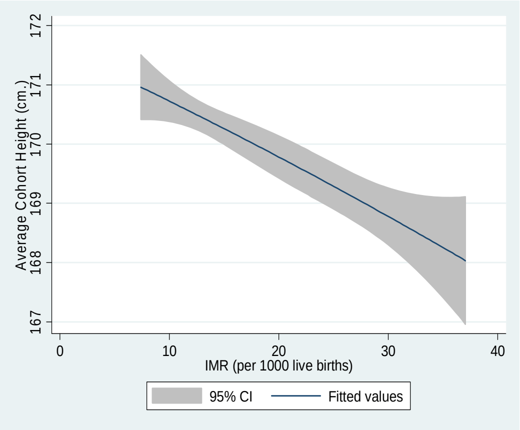

Figure 3: Average Height and log GDP

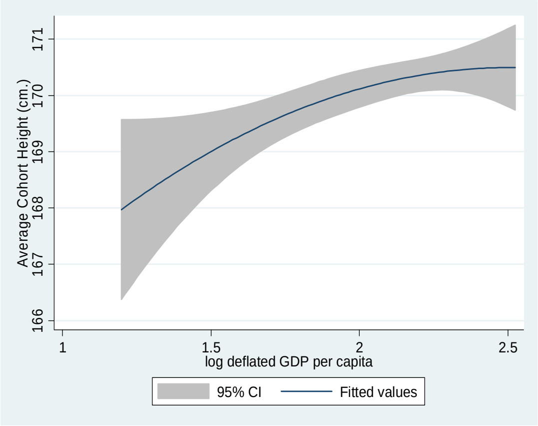

Table 7: Regressions of height on infant mortality and other variables: controlling for non-montonicities and non-linearities

	(1)	(2)	(3)	(4)
IMR	0.028	−0.004	0.026	−0.665
	(0.158)	(0.161)	(0.181)	(0.578)
IMR^2	−0.002	0.001	−0.002	0.010
	(0.003)	(0.003)	(0.003)	(0.015)
ln(GDP)	--	1.84	−1.62	−44.15
		(2.12)	(7.44)	(58.38)
$\ln(GDP)^2$	--	--	0.805	8.48
			(1.59)	(14.23)
Time trend	0.065	0.036	0.041	--
	(0.083)	(0.090)	(0.093)	
Inequality	--	--	--	27.02
				(20.19)
Region dummies?	YES	YES	YES	YES
F-test IMR and IMR^2	4.18**	1.35	1.50	5.25**
F-test ln(GDP) and $\ln(GDP)^2$	--	--	0.57	0.95
N	85	85	85	34

Notes: F-test NL in IMR = value of the F-statistic for the test of all the coefficients on the IMR variables equal to zero. F-test NL in ln(GDP) is similarly defined. Heteroskedasticity robust standard errors are reported in parentheses. Time trend is defined as the difference between mean cohort year and 1978. Inequality is measured by the Gini coefficient. Observations have been weighed using the number of individual observations that gave rise to the cohort-region average.

*** p-value < 0.01, ** p-value < 0.05, * p-value < 0.1