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Tue Gørgens
Xin Meng
Rhema Vaithianathan

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Forschungsinstitut
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Tue Gørgens
Australian National University

Xin Meng
Australian National University and IZA

Rhema Vaithianathan
University of Auckland

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IZA

P.O. Box 7240
53072 Bonn
Germany

Phone: +49-228-3894-0
Fax: +49-228-3894-180
E-mail: iza@iza.org

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ABSTRACT

Stunting and Selection Effects of Famine: A Case Study of the Great Chinese Famine*

The Great Chinese Famine of 1959-1961 is puzzling, since despite the high death rates, there is no discernable diminution in height amongst the majority of cohorts who were exposed to the famine in crucial growth years. An explanation is that shorter children experienced greater mortality and that this selection offset stunting. We disentangle stunting and selection effects of the Chinese famine, using the height of the children of the famine cohort. We find significant stunting of about 2cm for rural females and slightly less for rural males who experienced the famine in the first five years of life. Our results suggest that mortality bias implies that raw height is not always a good measure of economic conditions during childhood.

JEL Classification: C33, I12, N950, O15

Keywords: famine, height, China, panel data, GMM

Corresponding author:

Xin Meng
Division of Economics
Research School of Pacific and Asian Studies
Australian National University
Canberra 0200
Australia
E-mail: Xin.Meng@anu.edu.au

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1 Introduction

China's "Great Leap Forward" famine of 1959–1961 was one of the worst human catastrophes of the 20th Century, yet it is only recently that researchers have started to piece together its long-term consequences. One of the more puzzling aspects of the famine discussed by Schultz (e.g. Schultz, 2001) is that while there were an estimated 30 million excess deaths the adult height of the cohort who experienced famine during their early childhood appears no shorter than children born during non-famine years. This is puzzling because people exposed to such a drastic famine during a crucial growth phase would have been expected to be appreciably stunted.

There are two competing hypothesis as to why there was no apparent effect of the Chinese famine on height. One is the hypothesis of complete catch-up, which postulates that once the period of famine was over, food-adequacy was restored to such an extent that children were able to catch-up and achieve their full adult height. Incomplete catch-up is commonly referred to as stunting (e.g. Tanner, 1981). The hypothesis of complete catch-up paints the Chinese famine as a hiatus in an otherwise plentiful time, with no lasting consequences on the height of survivors.

The alternative explanation offered by Yan (1999) and Schultz (2001) is that the famine caused selection, creating a survivor bias in the height of the current Chinese population. If shorter children are less likely to survive a famine (e.g. Fawzi et al., 1997; Smedman et al., 1987), then this selection effect has to be controlled for before we may conclude that the famine did not cause stunting. The surviving populations may be taller or shorter depending on the relative size of these two effects. Schultz (2001) points out the importance of disentangling the stunting and selection effects of the Chinese famine but argues that "there is insufficient time-series evidence on mortality and health series indicators to know under what conditions one empirical force (i.e. stunting or selection by mortality) would dominate".

The objective of this paper is to estimate the stunting effects of famine, allowing for possible selection effects, in order to assess the merit of the two hypotheses. Our empirical strategy relies on modern data and uses the children of cohorts who were exposed to famine during early childhood and the children of cohorts who were less affected by famine to control for the selection effect. Children inherit their parents' genotype (selection) and not their phenotype (stunting). If famine survivors have greater potential height due to selection, then on average their children will inherit this potential and be taller than children of the control group. We can therefore use child-height to control for the effects of selection.

Using the height of children born to famine survivors and a control group, we are able to disentangle stunting from selection. We find that the average adult height of rural people who were exposed to the famine in the first 5 years of life is between 1 and 2cm shorter than our control group.

Our study echoes the finding of other researchers concerned about the possibility that height might be affected by selection bias due to the correlation between mortality and height (e.g. Vaupel et al., 1979; Waaler, 1984; Bairagi and Chowdhury, 1994). This is an increasingly important issue given the widely accepted practice in economics of using height to proxy for economic conditions, when income figures are unavailable or unreliable (e.g. Fogel et al., 1982; Fogel, 1994; Steckel, 1995; Micklewright and Ismail, 2001). Our findings sound a caution, that anthropometric measures may not be completely reliable.

Famine may also have important consequences for productivity. It is well established that in both developed and developing countries, height and wages are positively correlated.[1] However, it is not clear that genetic factors and environmental factors have the same effect (e.g. Schultz, 2002). If productivity is related to a person's realized rather than potential height, and famine causes stunting, then famine retards the person's productivity. Average productivity will be unaffected if the selection effects of famine leaves average height unchanged. Conversely, if productivity is related to a person's potential height, famine leaves individual productivity unaffected, but the selection effects of famine may increase average productivity.

This paper is set out as follows. Section 2 provides an overview of the Chinese famine. Section 3 discusses the genetic and environmental factors which determine the height of

[1] See Strauss and Thomas (1998) for a survey of some of the empirical evidence of this correlation.

a person and the relationship between the height of family members. Section 4 describes the data. In Section 5, we conduct some preliminary analysis, while Section 6 presents the econometric model and our estimation strategies. The results are discussed in Section 7, and Section 8 concludes the paper.

2 The Great Chinese Famine

The Great Chinese Famine started in 1959 and ended in 1961. There is still some controversy over the exact cause of the famine, although it was certainly associated with a reduction in grain output resulting from disruption in production attending the Great Leap Forward campaign and the collectivization of agriculture (e.g. Yao, 1999). This caused a drastic fall in grain production in 1959. However, it is generally accepted that the decline in food availability alone did not cause the estimated 20 to 30 million excessive deaths between 1958 and 1961.

It is widely held that overzealous officials, keen to make a good impression about the success of collectivization, exaggerated grain production. The central planners therefore, mistakenly believing there to be adequate grain supplies, exported rice, continued the wasteful practice of free grain and consumption in communal dining halls (e.g. Yang and Su, 1998) and acquired large amounts of grain for urban populations (e.g. Johnson, 1998; Lin and Yang, 2000). Widespread famine in the rural areas quickly followed.

Why the famine ended is still not certain. Johnson (1998) argues that it was associated with a wide-array of policy changes including the abolition of communal kitchens, importation of grain, and a reduction in the urban appropriation of grain. Land was returned to peasant control, and collectivization scaled back (e.g. Yang and Su, 1998).

Because of the lack of contemporaneous evidence, researchers have relied on mortality figures from the China statistical yearbook of 1983 to piece together what happened during this crisis (e.g. Coale and Banister, 1994; Lin and Yang, 2000; Wei and Yang, 2005). Riskin (1998) however points out that the reliability of the mortality statistics published in the Chinese Yearbook for the famine period cannot be corroborated since there is little information about how mortality was calculated.

Although the famine lasted only a short time, even based on official figures the annual mortality rate peaked at 28 per 1,000 in the rural areas, more than doubling the rate recorded in the pre-famine years (e.g. Lin and Yang, 1998). Between 1957 and 1960, death rates increased from 10.08 to 25.43 per 1,000 and the birth rate during the same period fell from 34 to 21 per 1,000. From the perspective of excessive deaths, the Great Chinese Famine outstrips any other recorded famine (e.g. Smil, 1999).[2]

During the 1950s, China was mainly a rural society, with 85% of the total population classified as rural dwellers. As Lin and Yang (2000) point out, even though farmers produced grain products, the centralized distribution and the urban-biased development strategy implied that when food was limited the rural population had to sacrifice their consumption. While both urban and rural populations experienced an increase in their mortality rate during the famine years, the urban death rate in 1960 was 1.6 times the pre-famine rate, while the rural rate over the same time period rose by a factor of 2.6.

There is also evidence that females suffered more than males. Coale and Banister (1994) use data from four censuses that were held between 1953 to 1990 as well as retrospective fertility surveys conducted in 1982 and 1988 to study the cohort-specific mortality rates. They find that although the gap between male and female mortality rates declined over the course of the 20th century, the decline was interrupted for cohorts who were children during the time of the famine. For these cohorts, girls were around 7% more likely to die than boys. They attribute this to a general neglect of female health during the famine, reflecting a cultural bias towards boys. They suggest that girls bore the brunt of the excess deaths caused by the famine.

There are to date a few papers that have found long-term consequences of the Chinese famine. St Clair et al. (2005) find that famine cohorts have an elevated risk of schizophrenia. Yan (1999) looks for long-term stunting by plotting the average height of females and males. She finds a reduction in average height for males born in the famine years, whereas for females she observes a peculiar spike in height.[3]

[2] However, the actual death rates during the Irish famine of 1845–1849 and the Bengali famine of 1943 were higher (e.g. O'Rourke, 1994).

[3] Chen and Zhou (2002) also look for stunting among famine survivors. Using data from the 1991 CHNS survey, they regress the height of rural adults on a constant, the excess death rate in the person's

3 Determination of Height

The objective of this paper is to estimate the stunting effects of famine in the presence of selection. Our estimation strategy is based on a comparison of the height of two generations. It is convenient to denote the older generation the "parents" and the younger the "children", although it should be kept in mind that many of the "parents" were children during the famine years and that many of the "children" are adult and parents at the time of data collection.

To disentangle stunting from selection, we compare the intergenerational height relationship between families where the parents experienced famine during their childhood and families where the parents did not. To motivate the strategy, this section discusses factors which determine a person's height and the relationship between the heights of family members. Each family consists of a mother, a father, and one or more children. Accordingly, we index the family members by $j = m, f, 1, \ldots, J$.

Our estimating strategy relies on comparing famine cohorts with a control group. To define the famine cohorts, one needs to understand the effect of famine on different age groups and to select those age groups which were most severely affected. While we have no information on the age profile of those who died during the Chinese famine, Salama et al. (2001) follow a sample of Ethiopians through a short famine period (December 1999 to July 2000) and find that 80% of those who died were children less than 14 years of age. There is also evidence that nutritional deficiencies in early childhood is more important for determining adult height than later childhood (e.g. Micklewright and Ismail, 2001; Glewwe and King, 2000; Hoddinott and Kinsey, 2001). To allow for this, we define two famine cohorts. The "old" famine cohort consists of those born between 1948 and 1956; they were aged between 5 and 13 in 1961. The "young" famine cohort consists of those born between 1957 and 1961, who were aged under 5 during the famine.

The control group is defined as those who were born up to 10 years immediately

home province in 1960, the birth year and the birth year interacted with the excess death rate. Their approach yields mixed and conflicting estimates. While they find some evidence of stunting, the pattern is neither consistent between females and males nor among different birth cohorts. Their estimates imply stunting of certain cohorts who were born *after* the famine.

before (1938 to 1947) and immediately after the famine cohorts (1962 to 1971).[4] The control group is chosen so as to extract a reasonably sized sample, while at the same time ensuring that it is close to the famine cohorts in birth years in order to minimize the possible impact of economic growth on height. Since the Great Famine affected all of China, it is impossible to find families who were not affected by famine in some generation. The genetic pool of any group of people, who were alive at some point in time after the famine, is therefore affected by selection. However, famine deaths are mostly amongst the very young and as the duration of the famine was short, the control group's genetic pool was subject to much less selection.[5]

A person's height at time t is determined by three major influences (e.g. Schultz, 2002): genetic factors including hormonal and biochemical factors, environmental factors which influence nutrition and health conditions during childhood, and his/her age at the time of measurement. Let h_{ijt} denote the height of the jth individual in the ith family in period t. Then

$$h_{ijt} = f(\text{age}_{ijt}, \text{sex}_{ij}) + G_{ij} + E_{ij} + U_{ijt}, \qquad j = m, f, 1, \ldots, J, \tag{1}$$

where f is some function of age and sex, G_{ij} represents the effect of genetic factors, E_{ij} the effect of environmental factors, and U_{ijt} is measurement error. The heights of the family members are related through both genetic factors and common environmental factors.

Medical research suggests that up to 60% of the height variation in a population can be attributed to genetic factors, but the exact inheritance process is not well understood (e.g. Ginsburg et al., 1998). A simple model of heritability (e.g. Goldberger, 1978) postulates that

$$G_{ij} = \tau_m G_{im} + \tau_f G_{if} + \tilde{G}_{ij}, \qquad j = 1, \ldots, J, \tag{2}$$

where G_{im} and G_{if} are the genotypes of the mother and father, τ_m and τ_f are weights

[4] Where our sample has three generations in a household, and all three generations are born after 1938, we discard the family unit where the parent is part of the control group. If there is no such choice, then we discard the younger family.

[5] It is possible that the 1938–1947 cohort, which were aged 14 to 23 in 1961, might also have been stunted. We discuss this possibility in Section 7.2, and check our results against this assumption.

with $\tau_m + \tau_f = 1$, and $\tilde{G}_{ij}$ is an individual-specific component. The latter is assumed to have mean 0 and be uncorrelated with G_{im} and G_{if}. There is no evidence that genes on the X or the Y chromosomes have any major effects (e.g. Carter and Marshall, 1978), whence it may be assumed that τ_m and τ_f equal 1/2.[6]

The determination of height is a result of a complex interaction of genetic and environmental factors which are not well understood (e.g. Tanner, 1981). Environmental factors such as restrictions on diet, exposure to diseases and physical activity can retard height. These environmental factors are affected by family and community characteristics. Parental income and education (e.g. Hoddinott and Kinsey, 2001) and birth order (e.g. Horton, 1986) have all been found to be relevant in explaining stature. The supply of public health services and clean drinking water are also important as chronic diarrhoea is a major cause of stunting in poor communities (e.g. Moore et al., 2001). In this paper, we focus on famine as a key environmental factor that determines height.

A model of the effect of environmental factors on height must accommodate unobserved as well as observed factors. It is particularly important to allow for unobserved factors which are common to all members of a family, because the characteristics of the local environment, socioeconomic status and lifestyle are strongly correlated between generations, which means that parents' nutritional intake, health and treatment in case of illness when young may be similar to that of their children.

An error-components model is highly flexible and well suited for our purposes. Singling out exposure to famine as an important determinant, we decompose the effect of environmental factors on height as follows:

$$E_{ij} = F'_{ij}\alpha_j + E^o_{ij} + E^c_i + \tilde{E}_{ij}, \qquad j = m, f, \tag{3}$$

$$E_{ij} = E^o_{ij} + E^c_i + \tilde{E}_{ij}, \qquad j = 1, \ldots, J, \tag{4}$$

where $F_{ij} = (F^o_{ij}, F^y_{ij})'$ is a vector whose two components indicate if the individual was born between 1948 and 1956 ($F^o_{ij} = 1$) or between 1957 and 1961 ($F^y_{ij} = 1$), the parameter $\alpha_j = (\alpha^o_j, \alpha^y_j)'$ is the amount of famine-related stunting for the old and the young famine

[6] We test the sensitivity of this assumption in Section 7.2.

cohorts, E_{ij}^{o} represents the effect of other observed factors (see Section 5), E_{i}^{c} represents the effect of unobserved factors which are common to all members of family i, and $\tilde{E}_{ij}$ the effect of unobserved factors which are specific to individual j. The latter is assumed to be uncorrelated with the observed and the unobserved common factors.

There are no famine dummies in the children's equations (Equation (4)), because no one in the second generation grew up during the Great Famine. Moreover, we exclude the parents' famine dummies from the children's equations, on the assumption that whether or not a parent experienced famine during his/her own childhood has no *direct* effect on their children's height. Indirect effects through E_{ij}^{o} and E_{i}^{c} are allowed. For example, the assumption does not rule out that on average parents in the famine cohorts feed their children better than parents in the control group. However, the difference must have arisen because parents who feed their children better were more likely to survive the famine (a selection effect through E_{i}^{c}), not because the famine caused survivors to change feeding patterns. Finally, we exclude each parent's famine dummy from the spouse's equation, because the fact that a person suffered famine during childhood cannot *directly* affect the partner's adult height. The assumptions that F_{im} and F_{if} do not appear in the spouse's nor in the children's equations are crucial for identifying the stunting effects, as we show in Section 6.

Combining equations (1), (2) and (3) yields the following model. The heights of each member of family i in time period t are given by

$$
\begin{aligned}
h_{imt} &= f(\mathrm{age}_{imt}, \mathrm{sex}_{im}) + F_{im}'\alpha_m + E_{im}^{o} + G_{im} + E_{i}^{c} + \tilde{E}_{im} + U_{imt}, \\
h_{ift} &= f(\mathrm{age}_{ift}, \mathrm{sex}_{if}) + F_{if}'\alpha_f + E_{if}^{o} + G_{if} + E_{i}^{c} + \tilde{E}_{if} + U_{ift}, \\
h_{i1t} &= f(\mathrm{age}_{i1t}, \mathrm{sex}_{i1}) + E_{i1}^{o} + \tau_m G_{im} + \tau_f G_{if} + E_{i}^{c} + \tilde{G}_{i1} + \tilde{E}_{i1} + U_{i1t}, \\
&\vdots \\
h_{iJt} &= f(\mathrm{age}_{iJt}, \mathrm{sex}_{iJ}) + E_{iJ}^{o} + \tau_m G_{im} + \tau_f G_{if} + E_{i}^{c} + \tilde{G}_{iJ} + \tilde{E}_{iJ} + U_{iJt},
\end{aligned}
\tag{5}
$$

where the unobserved specific variables $\tilde{G}_{ij}$, $\tilde{E}_{ij}$ and U_{ijt} are assumed to be uncorrelated with the observed variables as well as with the unobserved common variables, G_{im}, G_{if}

and E_i^c.

As mentioned in the Introduction, there are two competing hypotheses as to why there is no apparent effect of the famine on the average height of the current adult Chinese population. The first hypothesis is that after the famine was over children were able to catch-up and achieve their full adult height. In other words, the famine has no lasting effect on any individual. Model (5) is consistent with this hypothesis if α_m and α_f are both 0.

The second hypothesis claims that parents who experience famine during childhood are stunted. However, those who survived the famine are a select group who would, in the absence of a famine shock, have been taller than those who died (e.g. Yan, 1999; Schultz, 2001). The stunting and selection effects operate in opposite directions and may offset each other. The final height of the famine survivors may be no different to a control group of adults who did not suffer the famine. Model (5) is consistent with the stunting-selection hypothesis if α_m and α_f are positive.

There is no independent evidence that those who died during the Chinese famine were indeed shorter. However, Fawzi et al. (1997) studied Sudanese children between 6 months and 6 years of age during a famine and finds that after adjusting for a number of factors including age, sex, socio-economic status, and vitamin A levels, children in the shortest height-for-age category had a significantly higher mortality rate than taller children. Smedman et al. (1987) found similar results for children in Guinea-Bassau.

While the theoretical mechanism through which such a selection process might work is speculative, evidence that there is differential mortality by height has been established by Waaler (1984). Using a large sample of Norwegian individuals, Waaler found a clear reduction in mortality with increased body height. Waaler's findings have been corroborated on Swedish data by Peck and Vagero (1989) and British data by Leon et al. (1995). Kemkes-Grottenthaler (2005) investigated age of death and height of skeletons from various time periods. She found that taller individuals had a considerably heightened life-expectancy.

Friedman (1982), using data on slave mortality, observed that shorter slaves experienced higher mortality rates and concluded that "it is necessary to standardize for mortality differences before comparing the mean height of groups with substantially different mortality experiences".

Our model accommodates such selection effects by allowing G_{im}, G_{if} and E_i^c to be correlated with F_{im} and F_{if}. That is, famine survivors may have larger or smaller values of G_{im}, G_{if} and E_i^c than the control group who did not experience famine during their childhood.[7] These differential values will be passed on to their children, who as a consequence will be taller or shorter than children whose parents are in the control group.

Note that our model does not impose selection, nor does it require that selection be in favor of short people. Our model is therefore not inconsistent with Deaton's (2005) speculation that shorter people may be more efficient at using food than larger people and therefore more able to survive. If Deaton's hypothesis is true, then we would expect any selection effects to exacerbate the stunting effects, and for the children of famine survivors to be shorter.

The evidence indicates that during the famine, birth rates also fell dramatically (e.g. Coale and Banister, 1994). Therefore, members of the famine cohorts born between 1959 and 1961 are also censored by the falling birth rates. From the point of view of estimating the stunting of famine survivors, people who were never born do not pose a separate problem from people who were born but did not survive. In terms of our model, both kinds of selection causes the distribution of G_{im}, G_{if} and E_i^c to be different in the famine cohorts and the control group.

In closing this section, we note that the biological literature has identified assortative mating as a major confounding factor in the analysis of the inheritability of height (e.g. Carter and Marshall, 1978; Ginsburg et al., 1998). It is well established that people tend to marry people of similar characteristics, be it education, socio-economic status, or height. Thus, while there can be no direct effect of the father's stunting and selection effects on the mother's height and vice versa, there may be an indirect effect because of assortative mating: a man who is stunted is more likely to marry a short woman, and therefore more likely to marry a woman who is short for genetic reasons. Our model

[7] Survival probabilities may also be related to factors unrelated to height, but that is not a concern here.

accommodates assortative mating behavior by allowing the mother's and the father's observed and unobserved variables to be (positively) correlated (see Section 6).

4 Data, Famine Cohorts and Control Group

The data used in this study are from the China Health and Nutrition Survey (CHNS) conducted by the Carolina Population Center at University of North Carolina at Chapel Hill. We use the first four waves of the panel. The CHNS contains rich information including individual and household demographic and economic characteristics, health and nutrition status, living environment, and community characteristics. Most of this information refer to the time of the interview; historical information is limited. Importantly for our purposes, the survey included a physical examination of all members of each household by medical specialists with regard to height, weight, blood pressure, etc.[8]

The survey population is drawn from the provinces of Guangxi, Guizhou, Henan, Hubei, Hunan, Jiangsu, Liaoning, Shandong and Heilongjiang. Guangxi and Guizhou are located in the south-west, Hunan and Hubei in the inland, Jiangsu in the southeast, and Henan, Liaoning, Shandong and Heilongjiang are located in northern China. Average height varies significantly across provinces. People from the northern provinces tend to be taller than people in the south. This has been noted in research which compares the height of mainland Chinese with Hong-Kong Chinese and finds that despite the better economic conditions in Hong-Kong, northern mainland Chinese children are taller. While our sample is restrictive in terms of only covering eight provinces, these eight provinces are a reasonable representation in terms of size and the severity of famine. Using Lin and Yang's calculations (Lin and Yang, 2000, Table 3), we note that three of our eight provinces had higher death rates in 1960 than the national average. In the Lin and Yang data, 9 out of the 28 had rates higher than the national average. Therefore, our provinces slightly over-represent famine prone regions, although our data does not include information on Anhui, which was the most severely affected province.

[8]Further details on the CHNS can be found on the Carolina Population Center web site at http://www.cpc.unc.edu/china.

One of the unique features of this dataset is that it is a three-dimensional panel, varying across individuals, households, and time periods. The panel is unbalanced. First, in each year some households and survey sites are dropped and new households and survey sites added. Heilongjiang was not included until 1997, in which year Liaoning was dropped. Second, the number of individuals in each household changed over the eight-year period because of births, deaths, marriages etc.

It is well known that the death rate in rural areas was much higher than that in urban areas during the famine (e.g. Lin and Yang, 2000) and we therefore carry out our analysis separately for rural and urban areas. However, people living in an urban area at the time of the survey may have been in a rural area during the famine (and vice versa).

Between the late 1950s and mid 1980s the household registration system restricted labor mobility and largely confined people to their birthplaces. However, centrally controlled population movement did occur immediately before and after the famine period. During the Great Leap Forward (1957–1958), some people from rural areas were sent to cities to work. After the famine (1961–1962) these people were sent back (e.g. Zhao, 1999). Therefore some of those in our sample who are classified as rural passed the famine years in urban centers. Given that rural areas were more severely affected by famine, our estimates on the selection and stunting effects in the rural areas would be biased towards zero. Based on Zhao's (1999) estimation, between 1961 and 1962 around 20 million people were sent to the countryside. This amounts to only 3.5% of the 1962 rural population so the bias should not be significant.

In the post-famine period, the main concern is contamination of the urban data by migration from rural areas. Between 1964–1985, the population in the urban areas grew by 2.43% per annum due to internal migration.[9] In a 2002 survey of urban households, 18% had changed their status from rural to urban after 1959.[10] This migration may result in an overestimation of the famine effect on urban population. In conclusion, the potential

[9]Calculated by the authors using data on migration inflows from Zhao (1997) and total urban population from Comprehensive Statistical Data and Materials on 50 Years of New China, Beijing, National Bureau of Statistics.

[10]Calculated by the authors using data from Question 124 of the 2002 Urban Household Income Distribution Survey.

effect of internal migration on our results is to underestimate the effect of famine on the rural population and to overestimate the effect on the urban sample.

From each household in the CHNS we select a family unit which consists of a mother, a father and at least one child living with his/her parents.[11] Our final dataset, after excluding observations with missing information, consists of 2,115 families in the rural sample and 1,080 families in the urban sample. As previously mentioned, not all families were interviewed in each wave of the survey, and the number of family members may change from one wave to the next. Table 1 provides a cross-tabulation of the mother's and father's birth year for the rural sample (top panel) and urban sample (bottom panel).

For the rural mothers, 36% of the sample is in the old famine cohort and 17% in the young famine cohort while 16% are born before the famine and 31% after the famine. For the rural fathers, 35% are in the old famine cohort and 16% in the young famine cohort, while 22% are born before the famine, and a further 27% are born after the famine. The numbers for urban sample are approximately the same.

To check how representative our sample is, we compared our sample with the 2000 Chinese Census (0.1% sample). Our sample appears to contain a slightly smaller proportion of individuals in the famine cohorts and a larger proportion in the control group; most of the latter were born after the famine. Presumably the reason for this skewness is that individuals born after the famine are more likely to have children living at home relative to the other two groups.

Summary statistics of the data are provided in Table 2. The average heights of the rural and urban mothers are 155.2 and 156.0cm, respectively. For fathers, the rural-urban height difference is 1cm. The urban sample would be expected to be taller because of their relatively better economic conditions. The average ages of rural and urban mothers and fathers are 37, 37, 38 and 39, respectively.

The children are 11 years of age on average, ranging from 0 to 33 (not shown in the table). While there are 13% of boys aged 20 and above, only 9% of girls are in this age range. Older male children are more likely than their female counterparts to live with their

[11]The CHNS collects information about every individual living in each selected household at the time of the survey. No information is collected for family members living outside the household.

parents; this may explain large proportion of male children in the sample. In the rural areas, young mothers are 1.0cm taller and young fathers are 1.1cm taller than the control groups. The old mothers and fathers are 7mm shorter and 4mm taller, respectively. In the urban areas, both young mothers and young fathers are 6mm shorter, while there is virtually no difference in average height between the old famine cohorts and the control groups. The age and education differences are negligible for older famine fathers, while young famine fathers are slightly younger and better educated. Old famine mothers are slightly older and less educated than the control group while young famine mothers are younger and more educated.

5 Preliminary Analysis

As a preliminary step, we start by graphing height by birth cohort of the rural mothers and fathers (Figure 1). Recall that the Chinese famine was short and sharp and that by 1962 the crisis had all but passed (Lin and Yang, 2000). We would therefore expect there to be a jump in height for people born after 1962. Not only is this not apparent, there is also a peculiar jump in height for the 1960 female cohort, similar to that which puzzled Yan (1999).

However, this might be due to systematic differences in birth cohorts such as age, education and provinces. We therefore estimate the following model by OLS using data for all families and all years,

$$h_{ijt} = F'_{ij}\alpha_j + x'_{ijt}\beta_j + u_{ijt}, \qquad j = m, f, \tag{6}$$

where F_{ij} is the vector of famine dummies defined previously, x_{ijt} is a vector of other explanatory variables, and u_{ijt} is a residual. For reasons explained in Section 3, x_{ijt} consists of age, years of education (a proxy for the permanent income, socioeconomic status, health and nutrition during childhood), province (a measure of race), birth year (to capture the trend in economic development), and survey year dummies (to capture variations in measurement error between survey waves).[12]

The estimate of α_j is a measure of the average height difference between the famine cohorts and the control group, controlling for age etc. If there is no selection effect (and the correlation between the unobserved and the observed variables is negligible), then this would be an estimate of the famine-related stunting of the old and young famine cohorts.

Selected estimates are reported for mothers and fathers separately in Table 3.[13] We find that the young famine fathers show stunting of 0.62cm (t-ratio –1.68) in the rural sample and 0.98cm (t-ratio –1.87) in the urban sample which are significant at below the 10% level. However, the coefficients for young famine mothers is positive for the rural sample of 0.43cm and negative for the urban sample of 0.73cm but are both insignificant (t-ratio of 1.35 and –1.60 respectively). This suggests that either young famine mothers experienced full catch-up or that the stunting and selection effects cancel each other out.

For older famine fathers, positive height differentials of 0.50cm in the rural and 0.93cm in the urban sample are observed and both are significant at the 10% level (t-ratio of 1.80 and 2.31 respectively). For mothers, the difference is insignificant in both the rural and urban samples (–0.07cm with t-ratio of –0.27 and 0.64cm with t-ratio of 1.64 respectively).

Our data therefore confirms earlier findings by Yan (1999) that there is no apparent consistent pattern of stunting amongst famine cohorts, and a peculiar positive differential amongst some cohorts who passed through the famine at an older age. Further analysis is therefore needed to establish whether these results are due to the offsetting effects of stunting and selection.

We now turn to a simple test for selection. The idea is to exploit the fact that children inherit the parents' genotype, not their actual height (phenotype). Everything else being equal, if famine survivors were destined to be unusually tall but were stunted by the

[12]We calculate the age of each respondent at the time of the interview using his/her exact birth date and the date of the interview. Since the survey is carried out over several months, this means that birth year, age and survey year are not perfectly collinear in our data, and we include all three variables in our analysis. However, as this kind of identification is fragile and we do not want to interpret the effect of these variables separately. The estimated stunting effects are virtually unaffected whether we include all three variables or just (any) two.

[13]The t-ratios reported here and elsewhere are robust to heteroskedasticity and correlation across individuals and across time within a family and to heteroskedasticity across families. It is assumed that observations are independent across families.

famine, we would expect their children to be taller than children whose parents are in a suitably chosen control group. Conversely, if there is no selection bias in the height of the famine cohorts, we would expect no height difference between children of the famine cohorts and children of the control group.

To compare the height of the children of the famine cohorts and the control group, we estimate the following model by OLS,

$$h_{ijt} = F'_{im}\alpha^*_m + F'_{if}\alpha^*_f + x'_{ijt}\beta^*_c + u^*_{ijt}, \qquad j = 1, \ldots, J, \tag{7}$$

where F_{im} and F_{if} are vectors of the parents' famine dummies, x_{ijt} is a vector of other explanatory variables, and u^*_{ijt} is a residual.

The most important explanatory variable is the child's age. We show in Appendix A.1 that the height-age relationship for children is very well captured by cubic splines. Thus, for a child, x_{ijt} includes a cubic spline in age, sex, the spline interacted with sex, the mother's and father's years of schooling (proxies for family income during childhood and parents' knowledge about health and nutrition), the total number of children observed in the family and that number squared (to capture family resources per child), the birth order, the child's birth year and birth year squared, the mother's birth year, province dummies and survey year dummies.[14]

With a caveat on assortative mating explained below, the coefficients on the parents' famine dummies indicate the selection effect of famine. Separate identification of the coefficients of the mother's and the father's famine dummies requires there to be a sufficient number of families where one parent belongs to a famine group and the other to the control group. That is, if there were complete sorting and both parents belonged either to a famine cohort or to the control group, then F_m and F_f would be perfectly collinear and estimation would fail. While the children may be taller, it is impossible to tell how much is coming from the mother and how much from the father. It would be possible,

[14]The order of the children is defined according to the birth order of those children who live with their parents in one or more of the survey years. Since some children may not live with their parents (e.g. adult children), the order is not necessarily the birth order within the total number of children in the family. The maximum number of children observed in a family is six.

however, to estimate a joint effect on the children's height of having both parents in a famine cohort.

Table 1 shows that the proportion of marriages across famine cohorts and the control group in our data are 17% in the rural sample and 19% in the urban sample. While not zero, these are low figures. With such a high level of collinearity, it is difficult to estimate separate effects of the mother and the father. Insignificant t-tests should therefore be interpreted with caution, as the insignificance may be due to the difficulty in separating the effect of the mother from that of the father, rather than to there being no effect at all. Where relevant, we therefore supplement t-tests with Wald tests to examine joint significance.

In the later 1970s and early 1980s, the Chinese government introduced a "one-child" policy, which was more strictly enforced in the urban areas. It is possible that urban families with more than one child are a selected group. To check the robustness of our conclusion we re-estimate the relationship on a restricted sample using only the first child in each family. The results are similar. Table 4 presents selected OLS results.

For the rural sample, the estimates of α_m^* and α_f^* for all famine cohorts are positive and the Wald test suggests that for both the young and old cohorts, the famine dummies for mother and father are jointly significantly different from 0 (p-value of 0.00).

For the urban sample, the estimates are insignificant for both the young and the old famine cohort (p-value 0.42 and 0.58 respectively). The insignificance of parental cohort on child height is consistent with other evidence given in the literature that the famine had a more severe impact on the rural population than on the urban population (e.g. Lin and Yang, 2000).

It is possible that the estimates of α_m^* and α_f^* are affected by assortative mating. As discussed in Section 3, many studies have found significant correlation between the heights of a husband and wife. Assortative mating implies that a person who is stunted by famine is more likely to marry a short person, and if that person is short for any inheritable reason their children will be shorter. Consequently, the estimates of α_m^* and α_f^* may underestimate the selection effect of famine.

Let us summarize the story so far. For the rural sample, our preliminary results show that only the young famine fathers demonstrate visible evidence of stunting of a statistically significant nature. Yet, the children of all famine cohorts are taller than the control group, suggesting that there may have been some selection amongst all these groups. For the urban sample, the results are less clear. From the parent's estimates, we find some visible stunting amongst both young famine cohorts while the old famine cohorts demonstrate a positive height differential. The child height equation shows no evidence of selection.

6 Disentangling Stunting From Selection

While the results in the previous section were suggestive, they are not conclusive. Simple estimation methods such as OLS used in the preliminary analysis are inconsistent when unobserved variables are correlated with the explanatory variables. In this section, we present our econometric model of the relationship between the height of parents and their children. We escribe how to obtain consistent estimates of the stunting effects while controlling for selection by utilizing the information provided by children's height about the genotype of their parents.

Econometric Model

To simplify the discussion, let g_{ij} represent unobserved terms which are common to all family members and may be correlated with the explanatory variables, and let ϵ_{ijt} represent terms which are specific to individual j and uncorrelated with the explanatory variables. Specifically, for the mother and father, define $g_{ij} = G_{ij} + E_i^c$ and $\epsilon_{ijt} = \tilde{E}_{ij} + U_{ijt}$. For a child, define $g_{ij} = \tau_m G_{im} + \tau_f G_{if} + E_i^c$ and $\epsilon_{ijt} = \tilde{G}_{ij} + \tilde{E}_{ij} + U_{ijt}$, and note that $g_{ij} = \tau_m g_{im} + \tau_f g_{if}$ since $\tau_m + \tau_f = 1$.

The explanatory variables were discussed in Section 5. The vector x_{ijt} includes variables which represent age (and sex in case of children), observed environmental factors, and variation in measurement error between survey waves.

We model the effect of the explanatory variables on the height of various family members linearly, $x'_{ijt}\beta_j$. We assume the coefficients are the same for all children. To capture differential treatment, we include birth order, the total number of children observed in the family and the number of children squared as explanatory variables.[15] With this additional structure, model (5) implies that the heights of the members of family i in period t satisfy the equations

$$\begin{aligned} h_{imt} &= F'_{im}\alpha_m + x'_{imt}\beta_m + g_{im} + \epsilon_{imt}, \\ h_{ift} &= F'_{if}\alpha_f + x'_{ift}\beta_f + g_{if} + \epsilon_{ift}, \\ h_{i1t} &= x'_{i1t}\beta_c + \tau_m g_{im} + \tau_f g_{if} + \epsilon_{i1t}, \\ &\vdots \\ h_{iJt} &= x'_{iJt}\beta_c + \tau_m g_{im} + \tau_f g_{if} + \epsilon_{iJt}. \end{aligned} \tag{8}$$

The assumptions already imposed imply that the ϵ_{ijt}s are uncorrelated with all other right-hand side variables. However, g_{im} and g_{if} may be correlated with the explanatory variables (F_{im}, F_{if} , x_{imt}, x_{ift}, x_{i1t}, $\ldots$, x_{iJt}). Note that this does not rule out correlation and heteroskedasticity in ϵ_{ijt} across persons and across time.

The model is somewhat more complicated than a standard panel data model. First of all, we have a three dimensional panel (family, individual, time) rather than the usual two-dimensional panel (group, time).[16] Second, there are two unobserved group effects (g_{im}, g_{if}) instead of one. Third, the parameters in (8) are time-invariant but vary across individuals within a family, whereas in a standard panel data model the parameters are the same for all observations within a group. The remainder of this section outlines our two estimation methods, the within-group estimator and the GMM estimator.

For the purposes of estimation, we assume that observations are independent and identically distributed (iid) across families.[17] We also assume that τ_m and τ_f are known. This assumption greatly simplifies the estimation problem, because the model is linear

[15] Due to the possible endogeneity of the fertility decision, we also estimated a model without the variables "number of children" and "birth order". Omitting these variables does not alter our results.

[16] A single cross-section is sufficient for identification in our model. We use four time periods in order to reduce the influence of measurement errors and to increase the efficiency of the estimators.

[17] The iid assumption concerns the sampling method and is satisfied for our data with the usual caveat for survey non-response and attrition from the panel.

in the remaining parameters when τ_m and τ_f are fixed. In most of the analysis we take $\tau_m = \tau_f = 1/2$. We investigate the sensitivity of the estimates to this assumption in Section 7.2.

Within-Group Estimator

The principle behind the within-group estimator is to transform the equation system (8) in order to eliminate unobserved variables g_{im} and g_{if} which may be correlated with the explanatory variables. The parameters of the transformed equations can then be consistently estimated by OLS.

In the standard panel data model, the relevant unobserved variables appear in the same form in all equations within the group, and they may therefore be eliminated by subtracting from each variable its group mean (e.g. Hsiao, 1986, chapter 3). In our model, g_{im} and g_{if} do not appear in the same form in all equations. However, it is possible to estimate the stunting effects, α_m and α_f, by applying the within-group estimator after first combining the mother's and father's equations into an equation for the average parental height. Specifically, given fixed values of τ_f and τ_m, define $h_{ipt} = \tau_m h_{imt} + \tau_f h_{ift}$ and $\epsilon_{ipt} = \tau_m \epsilon_{imt} + \tau_f \epsilon_{ift}$.

Model (8) then implies that

$$\begin{aligned}
h_{ipt} &= \tau_m F'_{im}\alpha_m + \tau_m x'_{imt}\beta_m + \tau_f F'_{if}\alpha_f + \tau_f x'_{ift}\beta_f + \tau_m g_{im} + \tau_f g_{if} + \epsilon_{ipt}, \\
h_{i1t} &= x'_{i1t}\beta_c + \tau_f g_{if} + \tau_m g_{im} + \epsilon_{i1t}, \\
&\vdots \\
h_{iJt} &= x'_{iJt}\beta_c + q'_i\gamma_c + \tau_f g_{if} + \tau_m g_{im} + \epsilon_{iJt}.
\end{aligned} \tag{9}$$

Since the unobserved genetic heights enter each equation in (9) in the same form, namely $\tau_f g_{if} + \tau_m g_{im}$, they will be eliminated by subtracting group means from all variables as in the standard model.

As is usual in panel data models with "fixed effects", coefficients of variables which are constant within the family (e.g. province dummies) are not identified, because they are indistinguishable from the unobserved common variables. Fortunately, these parameters are not of particular concern in this paper.[18]

GMM Estimator

It is unlikely that the within-group estimator is efficient, because it puts equal weight on all equations.[19] In many applications this is not a serious problem. However, the stunting and selection effects of famine are likely to be small relative to the overall variation in height, and efficiency may therefore be an issue here. This leads us to our second estimation strategy, Generalized Method of Moments (GMM).

Let z_i be a linearly independent subset of the observed variables F_{im}, F_{if}, and x_{ijt} for all j and t. The unobserved variables for each individual in each period consists of g_{im}, g_{if} and ϵ_{ijt}. By assumption z_i and ϵ_{ijt} are uncorrelated. Following Chamberlain (1982) we capture the correlation between z_i and (g_{im}, g_{if}) using "nuisance" parameters.[20] Let $\phi_j = (\rho_{jm}, \rho_{jf}, \zeta_j')'$ be defined as the vector of coefficients obtained from projecting g_{ij} linearly on z_i. That is, define ϕ_j such that

$$\begin{aligned} g_{im} &= z_i'\phi_m + \eta_{im} = F_{im}'\rho_{mm} + F_{if}'\rho_{mf} + \tilde{x}_i'\zeta_j + \eta_{im}, \\ g_{if} &= z_i'\phi_f + \eta_{if} = F_{im}'\rho_{fm} + F_{if}'\rho_{ff} + \tilde{x}_i'\zeta_j + \eta_{if}, \end{aligned} \tag{10}$$

where $\tilde{x}_i$ is the vector such that $z_i = (F_{im}', F_{if}', \tilde{x}_i')'$. By definition, η_{im} and η_{if} are (un-

[18] As a consequence combining the parents' equations, certain parameters are no longer separately identified for the mother and father, but this does not affect the estimation of α_m and α_f.

[19] See for example Wooldridge (2002, chapter 11.3) for a discussion. Another advantage of GMM over the within-group estimator is that it is less sensitive to measurement errors in the explanatory variables (e.g. Hsiao, 1986, chapter 3.9; Ashenfelter and Krueger, 1994).

[20] The model can be thought of as an extension of Chamberlain's (1982) model. Chamberlain's assumptions are virtually identical to those made in this paper, but he considered a simple two-dimensional panel with a single common unobserved effect, no time-invariant explanatory variables, and no time-varying parameters of interest. It can also be shown that his minimum distance estimation procedure is equivalent to the GMM procedure used here.

conditionally) uncorrelated with z_i. Substituting (10) into (8) yields

$$\begin{aligned} h_{imt} &= F'_{im}(\alpha_m + \rho_{mm}) + F'_{if}\rho_{mf} + x'_{imt}\beta_m + \tilde{x}'_i\zeta_m + \tilde{\epsilon}_{imt}, \\ h_{ift} &= F'_{im}\rho_{fm} + F'_{if}(\alpha_f + \rho_{ff}) + x'_{ift}\beta_f + \tilde{x}'_i\zeta_f + \tilde{\epsilon}_{ift}, \\ h_{i1t} &= F'_{im}(\tau_m\rho_{mm} + \tau_f\rho_{fm}) + F'_{if}(\tau_f\rho_{ff} + \tau_m\rho_{fm}) + x'_{i1t}\beta_c + \tilde{x}'_i(\tau_m\zeta_m + \tau_f\zeta_f) + \tilde{\epsilon}_{i1t}, \\ &\vdots \\ h_{iJt} &= F'_{im}(\tau_m\rho_{mm} + \tau_f\rho_{fm}) + F'_{if}(\tau_f\rho_{ff} + \tau_m\rho_{fm}) + x'_{iJt}\beta_c + \tilde{x}'_i(\tau_m\zeta_m + \tau_f\zeta_f) + \tilde{\epsilon}_{iJt}, \end{aligned} \tag{11}$$

where $\tilde{\epsilon}_{imt} = \epsilon_{imt} + \eta_{im}$, $\tilde{\epsilon}_{ift} = \epsilon_{ift} + \eta_{if}$ and $\tilde{\epsilon}_{ijt} = \epsilon_{ijt} + \tau_m\eta_{im} + \tau_f\eta_{if}$.

Our assumptions imply that $E(z'_i\tilde{\epsilon}_{ijt}) = 0$ are satisfied for $j = m, f, 1, \ldots, J$ and $t = 1, \ldots, T$, and these equations are essentially the moment conditions we use for the GMM estimation. However, because of the proliferation of moment conditions as the number of children increase (both the number of equations and the number of moment conditions increase rapidly), it is not feasible to use all children in the GMM estimation. In Section 7 we present estimates using a maximum of one, two and three children.[21] Appendix A.3 lists the variables used to obtain the estimates presented in each case.

The reduced-form representation (11) facilitates the discussion of identification. As usual, coefficients of variables which are constant within the family are not identified. In (11), these variables appear in both x_{ijt} and $\tilde{x}_i$, and the corresponding elements of β and ζ are therefore not separately identified. The coefficients on the famine dummies (11) are mixtures of stunting and nuisance parameters. This demonstrates the pitfalls of a single-equation approach. The system, however, identifies six reduced-form parameters, $(\alpha_m + \rho_{mm}, \rho_{mf}, \alpha_f + \rho_{ff}, \rho_{fm}, \tau_m\rho_{mm} + \tau_f\rho_{fm}, \tau_m\rho_{mf} + \tau_f\rho_{ff})$, which can be solved for the six "structural" parameters, namely the stunting effects (α_m, α_f) and the nuisance parameters $(\rho_{mm}, \rho_{mf}, \rho_{fm}, \rho_{ff})$.

The system therefore identifies the stunting effects. We may interpret ρ_{mm} and ρ_{ff} as selection effects and ρ_{mf} and ρ_{mf} as effects of assortative mating. This interpretation

[21] About 25% of the observations (family-year) have one child living at home, 37% have two children, 25% have three, and 12% have four or more.

requires the additional assumption that the error terms in (10) satisfy $E(\eta_{ij}|z_i) = 0$ for $j = m, f$ (which is stronger than $E(z_i'\eta_{ij}) = 0$).

As can be seen in (10), ρ_{mm} and ρ_{ff} represent the mean difference in g_{ij} between the famine cohorts and the control group, adjusted not only for age, birth year, etc., but also for the spouse's and the children's characteristics. If there are differences in the spouse's and children's variables between famine cohorts and the control group, part of the total effect of selection may be captured by the coefficients of these variables, ζ_m and ζ_f.

The estimates presented in Section 7 are two-stage estimates, where the weight matrix in the first stage is $\sum_i z_i z_i'$ and the usual estimate of the optimal weight matrix is used in the second stage.

7 Discussion

7.1 Results for the Rural Sample

The estimated stunting and selection effects for the rural sample are presented in Table 5. As discussed in Section 5, issues of multicollinearity between the mother's and the father's cohort dummies render t-tests unreliable and hence Wald tests are used to test statistical significance.

It is possible that families with many children living at home are an unrepresentative group, and we therefore report results for four different specifications related to the number of children used in the estimation.[22] It is reassuring that the GMM and within-estimators for all specifications provide similar estimates of the stunting effects. As expected, the standard errors of the GMM estimates are smaller than for the within-group estimates.

As predicted by our preliminary analysis, we find large and significant stunting of the young famine cohort for the rural population. For mothers, the estimated stunting effects range from 1.49 to 2.80cm while for fathers, the estimates are smaller ranging from 1.30 to 1.82cm. All specifications show joint significance at below the 1% level. The finding of

[22] The results are based on all families, but only the first child in each family, the first two children etc. are used in the estimation. Families with fewer than the maximum number of children are included using standard methods for unbalanced panels.

larger stunting effects for mothers is consistent with other evidence that females suffered more than males during the Great Famine (see Section 2).

For the older famine cohort, mothers are stunted between 0.46cm and 1.49cm and fathers between –0.04cm and 0.87cm. Although only the two-children and three-children GMM estimates are jointly significant at the 10% level, except for the single and insignificant estimate of +0.04 they are all of the right sign.

The estimates are reasonable in comparison with empirical evidence of the immediate impact of drought on height. Hoddinott and Kinsey (2001) find that a year of drought reduced growth of Zimbabwean children aged between 12 and 24 months by between 1.5 and 2.0cm. However, since they do not follow drought-affected children to full adulthood they do not provide evidence of long-term stunting.

We now turn to the estimates of the selection effects derived from the GMM estimates. For the young famine cohort, they are jointly significant at the 1% level. For young famine mothers', the estimates range from 1.29 to 2.71cm while the young famine fathers show smaller selection of between 0.69 and 1.23 cm. The estimated selection effects for the older famine cohorts range from 0.35 to 1.11cm for mothers and 0.75 to 0.87 for fathers, and they are jointly significant at the 11% level for the one-child specification and 1% for the specifications with two and three children.

7.2 Sensitivity Analysis of Rural Results

To check the robustness of our estimates, we re-estimate the model (using only the within-group estimator) dropping either the pre-famine cohort or the post-famine cohort from the control group. If our results are an artifact of inadequately controlling for birth-year effects, we would expect these alternative definitions to expose such an anomaly.

The results for dropping the pre-famine cohort are reported in the Table 6 of the paper, while the results for dropping the post-famine cohort are available upon request. The OLS results for mother and father's height (top panel of Table 6) demonstrate a very similar pattern to the full model. In particular, none of the famine cohorts exhibit any significant stunting.

Turning to the child-height equation, again the pattern is very similar to the full sample, young famine cohort have taller children with both mother and fathers famine cohorts being positive and jointly significant. The older famine cohorts loses significance, and the sign for the mother's coefficient in the all-children specification is reversed.

Therefore after excluding the pre-famine cohort from the control group, there is still no evidence of visible stunting in the mother's famine cohort. Moreover, children of young famine cohorts continue to be taller than the rest of the sample. If the results for the full sample is an artifact of incorrectly controlling for birth-year effects, one would expect that excluding the older part of the control group would change the results.

We next turn to the within-group estimates of this alternative control group (bottom panel of Table 6), which yields stunting for young famine mothers of between 0.91 and 2.01cm and between 2.08 to 2.88cm for young famine fathers. These are all jointly significant (p-values of 0.03 or less) and similar in size to the stunting estimates obtained for the full model.

For older famine mothers we now observe a positive differential for the two-children specification, while older fathers show large stunting effects. However, these estimates are individually and jointly insignificant as they were in the case of the full sample.

Results from omitting the post-famine cohort demonstrate a similar pattern. The within-estimates continue to show significant stunting in the young famine cohort and insignificant stunting in the old famine cohort.

We also perform an additional check to determine whether our results are robust to using a narrower control group. We restrict the control group to be those who were born five years immediately before 1948 and five years immediately after 1961. This reduces the sample size to 3,677.

The OLS estimates of parental height are reported in the top panel of Table 7. The results indicate that none of the famine dummy variables are statistically significant for either mothers or fathers. The results from the child-height equation are very similar to the full sample. Both famine cohorts have taller children and the coefficients for mothers and fathers are jointly significant.

The within-group estimates of the stunting effects are similar in size to the results using full control group (bottom panel of Table 7). Stunting of the young famine mothers ranges from 1.31cm to 1.72cm. For the young famine fathers, it ranges from 1.89 to 2.16cm. The young famine cohort continues to be jointly significant. The coefficients of the older fathers range from 0.51 to 0.87cm and for older mothers between 0.04 and 0.65cm. These are not jointly significant.

As yet another check on the results, we estimated the OLS parent- and child-height regressions as well as the within-group model on the full sample, but with a separate dummy for cohorts born before the famine. We then tested whether the coefficient on the pre-famine cohort was significantly different from the post-famine cohort. We found that we could not reject this hypothesis at the 10% level. This suggests that pooling the pre-famine and post-famine cohorts is justified.

In the analysis reported in Table 5 we have assumed $\tau_m = \tau_f = 1/2$, which is reasonable given that there is no evidence in the literature to suggest that the genes of either parent are more important in determining the height of their child. Nevertheless, it is useful to check how sensitive our results are to this assumption. Figure 2 shows the estimated stunting effects and 95% confidence bands plotted against τ_m (with $\tau_f = 1 - \tau_m$). The solid line is the estimate for the mother and the dashed line is for the father. We only report the all-children specification since one-child and two-children are similar.[23]

The figures confirm that for the older famine coefficients the stunting estimates are very robust to changes in τ_m and τ_f. The estimates are fairly constant over the range 0.3 to 0.7, and even outside this interval the variation is modest. In the case of the young famine cohorts the results are slightly different. Fathers' coefficients are stable, but for the young famine mothers, the stunting estimates becomes large as less weight is placed on their inheritability although this is not significant when the confidence band is taken into consideration.

[23] The exponential increase in the width of the confidence band for α_j as τ_j approaches 0 reflects the fact that α_j is not identified when $\tau_j = 0$, because the height of the children are not informative about the potential height of the parent in this case.

7.3 Results From the Urban Sample

We now turn to the results for the urban sample. Recall that while the urban famine cohorts appear taller than the control group (Table 3), we cannot reject the hypothesis that the children of these two groups have the same height (Table 4).

The estimated stunting effects from the within group specification are presented in Table 8. The within-estimates are negative for the young famine cohorts although they are jointly insignificant. In the case of the older cohorts, we find evidence of stunting only for famine mothers.

As discussed in Section 4, there are a number of problems with interpreting the coefficients for the urban sample. An additional problem is the small number of intermarriages between members of the famine cohort groups and non-famine groups in the urban sample (Table 1). These problems, together with the small size of the one-child sample (1,237 as opposed to 3,087 in the rural sample), explain why the GMM estimator for the urban sample does not yield sensible results, and we do not report them.

8 Conclusion

This paper studies the long-term effect of the Great Chinese Famine on health as measured by height. We disentangled the stunting from the selection effect of the famine using the children of the famine cohorts and a control group to identify selection bias. We discovered that far from there being complete catch-up in the height of victims, the famine caused a small but significant amount of stunting amongst rural females and males aged under five during the famine.

This finding is robust to differences in econometric specifications as well as alternative definitions of the control group. For people who were subject to the famine as older children, we uncovered some evidence of stunting, but this was less robust.

The results of this paper suggests that a cautious approach to the use of stature as a measure of well-being in a developing country or historical settings is warranted. While stature undoubtedly has a crucial role to play in documenting economic conditions, interpreting trends in height must be undertaken in light of information on morbidity and mortality. Indeed, when an economic shock has a catastrophic effect, we may fail to observe any secular trend in height.

A Technical appendix

A.1 Children's Age Splines

For children's height to be a good measure of their genetic height, it is important to control properly for their age. A preliminary data analysis suggested that the population average height-age relationship for children is very well modeled using cubic splines. For our final results we use

$$\begin{aligned}
\text{Age1} &= 1(\text{age} < 18)\left(\frac{1}{324}\text{age}^2 - \frac{1}{9}\text{age} + 1\right) \\
\text{Age2} &= 1(\text{age} < 18)\left(\frac{1}{864}\text{age}^3 - \frac{1}{24}\text{age}^2 + \frac{3}{8}\text{age}\right) \\
\text{Age3} &= 1(\text{age} < 10)\left(\frac{27}{4000}\text{age}^3 - \frac{27}{200}\text{age}^2 + \frac{27}{40}\text{age}\right) \\
\text{Age4} &= 1(\text{age} < 10)\left(-\frac{1}{1000}\text{age}^3 + \frac{3}{100}\text{age}^2 - \frac{3}{10}\text{age} + 1\right).
\end{aligned}$$

These variables correspond to a cubic spline with knots at age 10 and 18, restricted to be constant after age 18 and restricted to have a continuous first derivative. As defined the variables are scaled to range between 0 and 1. In the estimation we allow for different coefficients for boys and girls.

The splines capture the height-age relationship for children very well, as can be seen in Figure 3 which shows the average age-specific height (circles) and the predicted values obtained from regressing height on the four spline variables and a constant. The variability in the age-specific averages for children in their twenties and thirties is due to small sample sizes.

A.2 Preliminary Estimates Revisited

It is useful to consider the estimates in our preliminary analysis in Section 5 in the light of (8). The parents' equation (6) is identical to the parents' equations in (8), if $u_{ijt} = g_{ij} + \epsilon_{ijt}$. OLS estimation yields inconsistent estimates when (F_{ij}, x_{ijt}) and g_{ij} are correlated. However, if these variables were uncorrelated the OLS estimator of the stunting, α_j, would be consistent, as claimed in Section 5.

The children's equation (7) is slightly more complicated, because it includes the parents' famine dummies. Suppose there are linear mean relationships between the parents' genetic heights on the one hand and the famine dummies and the explanatory variables in the parents' or the children's equations on the other hand,

$$E(g_{ij}|F_{im}, F_{if}, x_{ikt}) = F'_{im}\xi_{jm} + F'_{if}\xi_{jf} + x_{ikt}\xi_{jx}, \quad j = m, f, \quad k = m, f, 1, \ldots, J. \quad (12)$$

Here ξ_{mm} and ξ_{ff} represent the selection effects of famine and ξ_{mf} and ξ_{fm} represent the effects of assortative mating based on on height. (A stunted person is more likely to marry a person who is short for genetical or environmental reasons, hence F_{im} is positively correlated with g_{if} and vice versa.) The children's equations in (8) together with (12) imply equation (7), where $\alpha^*_m = \tau_m\xi_{mm} + \tau_f\xi_{fm}$, $\alpha^*_f = \tau_m\xi_{mf} + \tau_f\xi_{ff}$, $\beta^*_c = \beta_c + \tau_m\xi_{mx} + \tau_f\xi_{fx}$, and where u^*_{ijt} is uncorrelated with the explanatory variables by construction. By the latter property, consistent estimates of the parameters in (7) can be obtained by OLS.

It follows that a test of $\alpha^*_m = 0$ and $\alpha^*_f = 0$ in (7) is a test of selection if (12) is a valid representation and if the effects of assortative mating are negligible under the null. We expect the linear model (12) to be a reasonable approximation to the true conditional mean. In view of the empirical finding in Section 5, absence of selection implies absence of stunting. Absence of stunting annihilates the effect of assortative mating. Hence, we expect assortative mating effects to be small under the null of no selection effects. The test is therefore informative about the selection effects of famine. If $\xi_{mf} = 0$, $\xi_{fm} = 0$ and $\tau_m = \tau_f = 1/2$, then the selection effects are $\xi_{mm} = 2\alpha^*_m$ and $\xi_{ff} = 2\alpha^*_f$.

A.3 Moment Conditions

As mentioned in Section 6, our assumptions imply that $E(z_i'\tilde{\epsilon}_{ijt}) = 0$ are satisfied for $j = m, f, 1, \ldots, J$ and $t = 1, \ldots, T$, where z_i represents all right-hand side variables in the model. The number of moment conditions available for estimation is very large as each person's variables generates a moment condition for each family member's equation, and it is necessary to restrict the number of variables included in z_i.

The following variables are included in z_i for the GMM estimates based on the first child only: constant, mother's famine dummies, father's famine dummies, province dummies, mother's age in 1997, total number of children observed in the family during 1989–1997 and that number squared, mother's maximum schooling, father's age in 1997, father's maximum schooling, age of the first child in 1997 (A97), 1(A97 < 10), A97∗1(A97 < 10), 1(10 ≤ A97 < 18), A97∗1(10 ≤ A97 < 18), sex of the first child, and the five age variables interacted with sex. The GMM estimates based on two children uses additional 11 moment conditions: the equivalent age and sex variables for the second child. The GMM estimates based on three children uses additional three moment conditions: the age of the third child in 1997, the sex of the third child, and the interaction between age and sex.

As the panel is heavily unbalanced some of the moment conditions are not useful. In order to reduce collinearity, moment conditions with few nonzero contributions were dropped. The number of moments actually matched is indicated in the tables.

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Table 1: Family Cohort Frequencies

Mother's Cohort	Father's Cohort: Pre-famine (1938–1947)		Old Famine (1948–1956)		Young Famine (1957–1961)		Post-famine (1962–1971)		Total	
Rural Population										
1938–1947	310	15%	22	1%	0	0%	0	0%	332	16%
1948–1956	161	8%	557	26%	34	2%	0	0%	752	36%
1957–1961	2	1%	135	6%	203	10%	29	1%	369	17%
1962–1971	0	0%	25	1%	105	5%	532	25%	662	31%
Total	473	22%	739	35%	342	16%	561	27%	2115	100%
Urban Population										
1938–1947	142	13%	13	1%	0	0%	0	0%	155	14%
1948–1956	85	8%	278	26%	16	1%	0	0%	379	35%
1957–1961	2	0%	90	8%	111	10%	10	1%	213	20%
1962–1971	0	0%	22	2%	73	7%	238	22%	333	31%
Total	229	21%	403	37%	200	19%	248	23%	1080	100%

Table 2: Summary Statistics

	Mothers		Fathers		Children	
	Mean	SD	Mean	SD	Mean	SD
Rural Population						
All						
Height (cm)	155.2	5.7	165.8	6.1	130.2	29.2
Age (years)	36.9	7.9	38.4	8.1	10.9	6.7
Schooling (years)	5.3	3.9	7.4	3.1		
Males (%)					54.0	0.5
Old Famine Cohort (1948–1956)						
Height (cm)	154.6	5.7	165.9	5.7		
Age (years)	39.8	3.8	40.0	3.9		
Schooling (years)	4.4	3.7	7.0	3.1		
Young Famine Cohort (1957–1961)						
Height (cm)	156.4	5.5	166.5	6.0		
Age (years)	33.3	3.2	33.5	3.2		
Schooling (years)	6.3	4.1	8.7	2.8		
Control Group						
Height (cm)	155.3	5.8	165.5	6.4		
Age (years)	35.6	10.6	38.8	11.0		
Schooling (years)	5.7	3.7	7.3	3.2		
Urban Population						
All						
Height (cm)	156.0	5.9	166.9	6.7	132.4	30.2
Age (years)	36.6	7.6	38.5	7.7	11.1	6.9
Schooling (years)	7.6	4.1	8.9	3.6		
Males (%)					53.1	0.5
Old Famine Cohort (1948–1956)						
Height (cm)	156.0	5.6	167.0	6.5		
Age (years)	39.9	3.8	39.9	3.8		
Schooling (years)	6.8	4.3	8.5	3.6		
Young Famine Cohort (1957–1961)						
Height (cm)	155.6	6.2	166.4	7.3		
Age (years)	33.7	3.2	34.0	3.1		
Schooling (years)	8.8	3.7	9.8	3.3		
Control Group						
Height (cm)	156.2	6.1	167.0	6.6		
Age (years)	34.7	10.2	39.1	10.8		
Schooling (years)	7.8	3.8	8.8	3.6		

SD denotes standard deviation. Averages over all respective individuals in all years with no adjustment for the unbalanced sample.

Table 3: Mother's Height and Father's Height OLS Results

	Old Famine Cohort (1948–1956)		Young Famine Cohort (1957–1961)	
	Mother	Father	Mother	Father
Rural Population	−0.07	0.50	0.43	−0.62
	(−0.27)	(1.80)	(1.35)	(−1.68)
Urban Population	0.64	0.93	−0.73	−0.98
	(1.64)	(2.31)	(−1.60)	(−1.87)

Robust t-statistics in parenthesis. Separate regressions for the mother and father. In addition to famine dummies, the set of regressors include age, birth-year, years of schooling, year dummies for 1989, 1993 and 1997, province dummies, and a constant. The complete estimation results are available from the authors upon request.

Table 4: Children's Height OLS Results

	Old Famine Cohort (1948–1956)			Young Famine Cohort (1957–1961)		
	Mother	Father	Wald	Mother	Father	Wald
All Children						
Rural Population	0.62	0.66	13.71	0.94	1.15	24.32
	(1.63)	(1.71)	[0.00]	(2.24)	(2.74)	[0.00]
Urban Population	0.09	0.56	1.74	−0.13	−0.53	1.11
	(0.16)	(1.04)	[0.42]	(−0.21)	(−0.83)	[0.58]
One Child						
Rural Population	0.81	0.69	11.62	1.37	0.82	16.83
	(1.78)	(1.48)	[0.00]	(2.81)	(1.64)	[0.00]
Urban Population	0.17	0.19	0.30	0.00	−0.56	0.78
	(0.27)	(0.31)	[0.86]	(0.00)	(−0.78)	[0.68]

Wald indicates the Wald statistic for the joint significance of mother's and father's cohort dummies; robust t-statistics in parenthesis and p-values in brackets. The results for one child are based on the oldest (observed) child in each family. The full set of regressors include four famine dummies for the mother and the father, a four-parameter cubic spline in age, a sex dummy and four interaction terms between sex and the age spline, the child's birth-year and that number squared, the birth-order of the child, the total number of children in the family and that number squared, the mother's and father's years of schooling, the mother's birth-year, year dummies for 1989, 1993 and 1997, province dummies, and a constant. The complete estimation results are available from the authors upon request.

Table 5: Summary of Estimates for the Rural Population

Estimator	Old Famine Cohort (1948–1956)			Young Famine Cohort (1957–1961)		
	Mother	Father	Wald	Mother	Father	Wald
Within-Group Stunting Estimates						
One Child	−1.19	−0.52	4.52	−2.22	−1.63	12.04
	(−1.42)	(−0.61)	[0.10]	(−2.31)	(−1.62)	[0.00]
Two Children	−0.46	−0.87	3.84	−1.49	−1.80	12.38
	(−0.65)	(−1.17)	[0.15]	(−1.69)	(−2.09)	[0.00]
Three Children	−0.65	−0.63	4.08	−1.62	−1.47	11.56
	(−0.96)	(−0.90)	[0.13]	(−1.88)	(−1.75)	[0.00]
All Children	−0.84	−0.40	3.98	−1.69	−1.51	13.08
	(−1.19)	(−0.55)	[0.14]	(−1.98)	(−1.80)	[0.00]
GMM Stunting Estimates						
One Child	−1.39	−0.18	3.56	−2.80	−1.30	15.65
	(−1.60)	(−0.23)	[0.17]	(−3.05)	(−1.46)	[0.00]
Two Children	−1.49	−0.04	10.44	−1.69	−1.79	23.38
	(−2.65)	(−0.07)	[0.01]	(−2.57)	(−2.73)	[0.00]
Three Children	−0.87	0.04	4.85	−1.60	−1.82	30.24
	(−1.82)	(0.08)	[0.09]	(−2.79)	(−3.13)	[0.00]
GMM Selection Estimates						
One Child	1.11	0.87	4.33	2.71	0.69	11.23
	(1.27)	(1.08)	[0.11]	(2.97)	(0.79)	[0.00]
Two Children	1.07	0.85	9.51	1.46	1.23	12.68
	(1.87)	(1.40)	[0.01]	(2.25)	(1.93)	[0.00]
Three Children	0.35	0.75	4.56	1.29	1.03	13.03
	(0.74)	(1.50)	[0.10]	(2.35)	(1.81)	[0.00]

Wald indicates the Wald statistic for the joint significance of mother's and father's cohort dummies; robust t-statistics in parenthesis and p-values in brackets. †Calculated as the OLS (overall) estimate minus the within-group (stunting) estimate. The set of regressors are the same as for the OLS estimates (see Tables 3 and 4), except that due to perfect multicollinearity between the transformed variables the following were dropped: mother's age, mother's year dummies, father's year 1991, parents' constant, mother's birth-year in children's equation, children's province dummies. The complete estimation results are available from the authors upon request.

Table 6: Summary of Estimates for the Rural Population
Omitting Pre-famine Cohorts

Estimator	Old Famine Cohort (1948–1956)			Young Famine Cohort (1957–1961)		
	Mother	Father	Wald	Mother	Father	Wald
OLS Estimates						
Parent's Height	−0.71	0.10		0.06	−0.80	
	(−0.93)	(0.12)		(0.12)	(−1.44)	
Children's Height, All	−0.17	1.20	3.44	0.61	1.43	13.48
	(−0.20)	(1.80)	[0.18]	(1.03)	(2.86)	[0.00]
Children's Height, One	0.52	0.49	1.11	1.29	0.73	8.69
	(0.53)	(0.63)	[0.57]	(1.98)	(1.27)	[0.01]
Within-Group Stunting Estimates						
One Child	−0.95	−1.52	1.15	−2.01	−2.08	7.23
	(0.46)	(−0.69)	[0.56]	(−1.48)	(−1.46)	[0.03]
Two Children	0.44	−3.01	2.36	−0.91	−2.84	8.41
	(0.25)	(1.49)	[0.31]	(−0.72)	(−2.22)	[0.01]
Three Children	−0.14	−3.00	2.85	−1.32	−2.68	9.49
	(−0.08)	(−1.53)	[0.24]	(−1.08)	(−2.16)	[0.01]
All Children	−0.20	−3.06	3.06	−1.37	−2.88	11.04
	(−0.11)	(−1.57)	[0.22]	(−1.12)	(−2.34)	[0.00]

See notes for Tables 3, 4 and 5. The complete estimation results are available from the authors upon request.

Table 7: Summary of Estimates for the Rural Population
Using Narrow Five-year Control Group

Estimator	Old Famine Cohort (1948–1956)			Young Famine Cohort (1957–1961)		
	Mother	Father	Wald	Mother	Father	Wald
OLS Estimates						
Parent's Height	−0.31	0.48		0.24	−0.63	
	(−0.93)	(1.48)		(0.72)	(−1.57)	
Children's Height, All	0.47	0.92	12.58	0.93	1.47	29.42
	(1.10)	(2.27)	[0.00]	(2.13)	(3.38)	[0.00]
Children's Height, One	0.30	0.83	5.29	1.13	1.10	15.33
	(0.60)	(1.70)	[0.07]	(2.24)	(2.14)	[0.00]
Within-Group Stunting Estimates						
One Child	−0.38	−0.51	0.94	−1.72	−1.96	9.88
	(−0.43)	(−0.58)	[0.63]	(−1.76)	(−1.87)	[0.01]
Two Children	−0.04	−0.87	1.77	−1.31	−2.16	12.78
	(−0.05)	(−1.13)	[0.41]	(−1.43)	(−2.41)	[0.00]
Three Children	−0.42	−0.77	2.76	−1.56	−1.89	13.35
	(−0.55)	(−1.05)	[0.25]	(−1.78)	(−2.18)	[0.00]
All Children	−0.65	−0.57	2.85	−1.66	−1.96	15.23
	(−0.83)	(0.74)	[0.24]	(−1.91)	(−2.25)	[0.00]

See notes for Tables 3, 4 and 5. The complete estimation results are available from the authors upon request.

Table 8: Summary of Estimates for the Urban Population

Estimator	Old Famine Cohort (1948–1956)			Young Famine Cohort (1957–1961)		
	Mother	Father	Wald	Mother	Father	Wald
Within-Group Stunting Estimates						
One Child	−1.07	1.06	1.19	−2.22	0.30	3.03
	(−0.88)	(0.94)	[0.55]	(−1.63)	(0.21)	[0.22]
Two Children	−0.68	0.43	0.45	−1.46	−0.34	2.42
	(−0.65)	(0.45)	[0.80]	(−1.24)	(−0.28)	[0.30]
Three Children	−0.79	0.45	0.62	−1.53	−0.51	3.03
	(−0.77)	(0.48)	[0.73]	(−1.32)	(−0.42)	[0.22]
All Children	−0.57	0.46	0.39	−1.42	−0.50	2.67
	(−0.55)	(0.50)	[0.82]	(−1.22)	(−0.41)	[0.26]

See notes for Table 5.

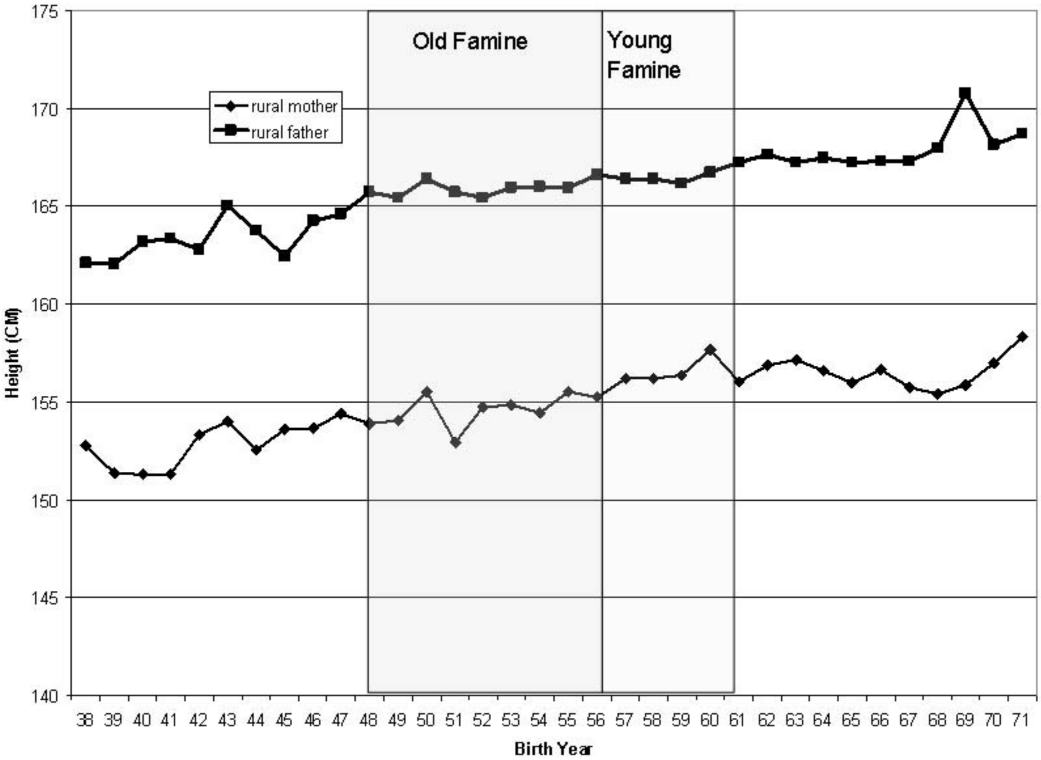
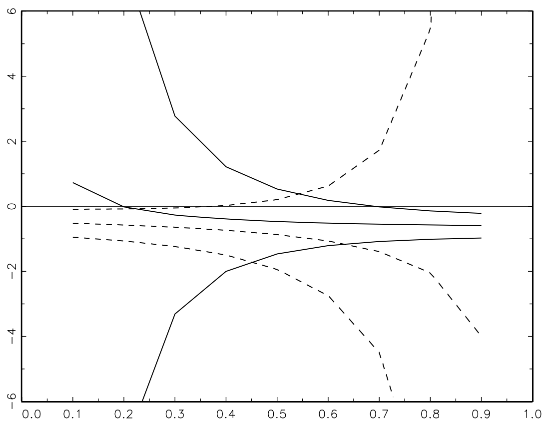
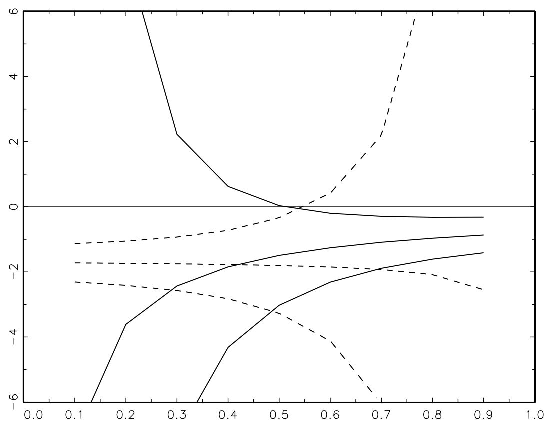


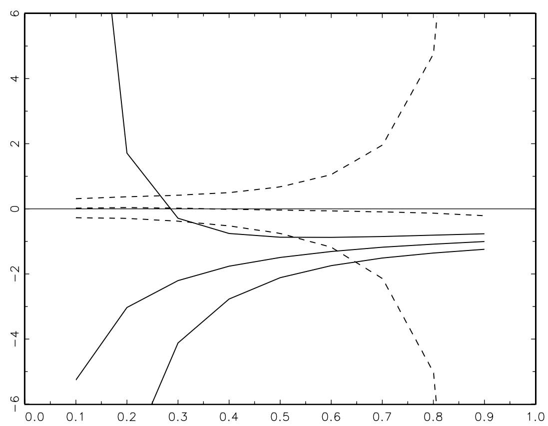
Figure 1: Height by birth year



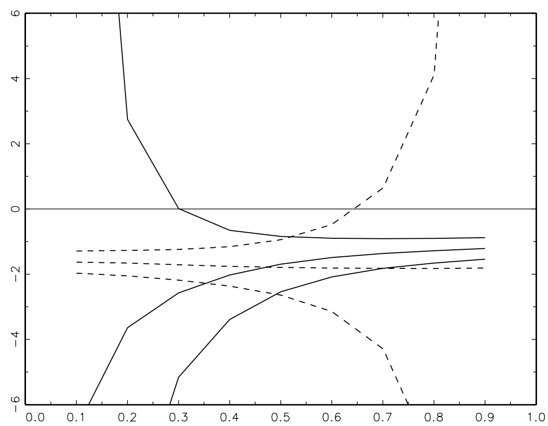
Within-Group Estimates
Old Famine Cohort



Within-Group Estimates
Young Famine Cohort

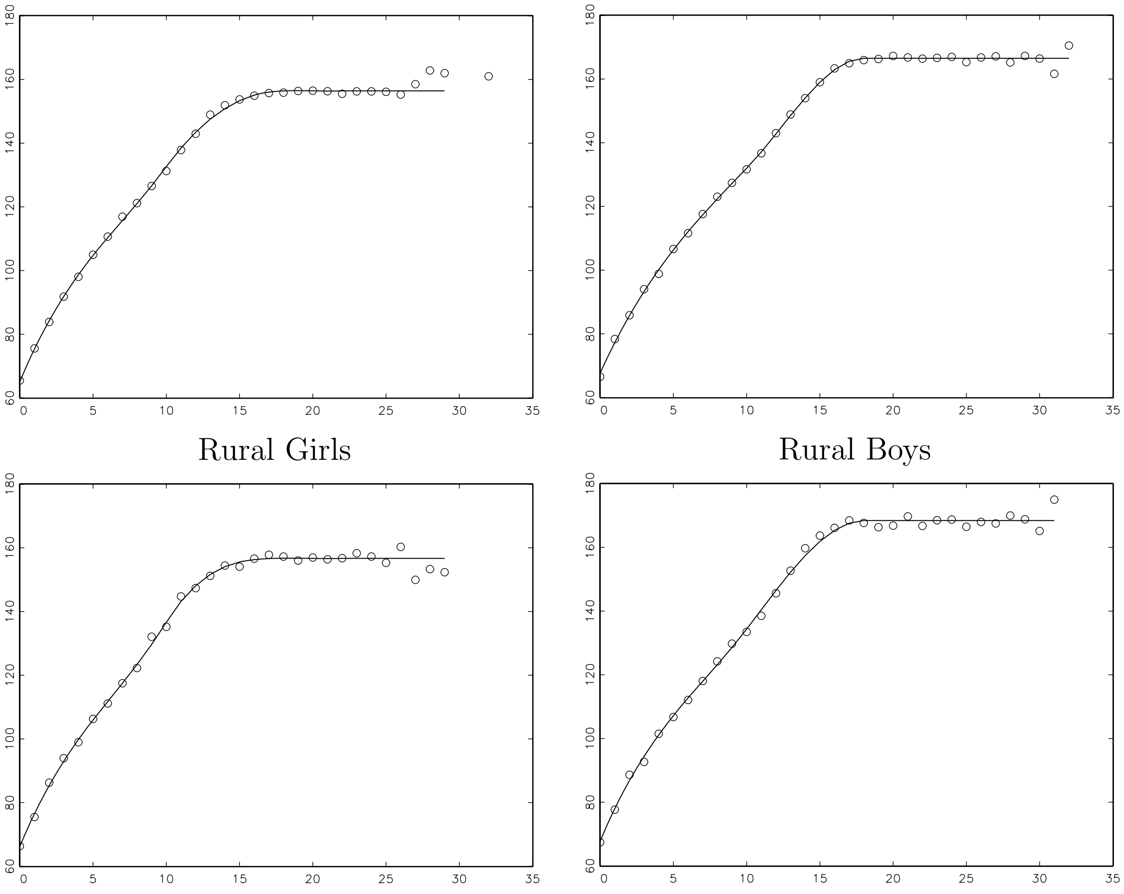


GMM Estimates
Old Famine Cohort



GMM Estimates
Young Famine Cohort

Figure 2: Stunting Effects Plotted Against τ_m (Two Children, Rural Population)



Rural Girls

Rural Boys

Urban Girls

Urban Boys

Figure 3: Height-Age Profiles For Children