Intergenerational Social Mobility and Assortative Mating in Britain

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Abstract

This paper investigates the links between the socio-economic position of parents and the socio-economic position of their offspring and, through the marriage market, the socio-economic position of their offspring's parents-in-law. Using the Goldthorpe-Hope score of occupational prestige as a measure of status and samples drawn from the British Household Panel Survey 1991-1999, we find that the intergenerational elasticity is around 0.2 for men and between 0.17 and 0.23 for women. On average, the intragenerational correlation is lower, and of the order of 0.15 to 0.18, suggesting that the returns to human capital, which is transmitted across generations by altruistic parents, contribute more to social status than assortative mating in the marriage market. Substantially higher estimates are reported when measurement error is accounted for. We also find strong nonlinearities, whereby both inter- and intra-generational elasticities tend to increase with parental status. We offer four possible explanations for this finding, three of which – one based on mean-displacement shifts in the occupational prestige distribution, another based on life-cycle effects and the third based on differential measurement errors – do not find strong support in our data. The fourth explanation is based on the notion of intergenerational transmission of social capital and intellectual capital. The evidence supports the idea that richer parents are likely to have a larger and more valuable stock of both social capital and intellectual capital to pass on to their children.

JEL Classification: J12, I20, D31, D64

Keywords: Intergenerational links, marriage market, assortative mating, Goldthorpe-Hope occupational prestige index, social and intellectual capital

Non-technical summary

This paper investigates the links between the socio-economic position of parents and the socio-economic position of their offspring and, through the marriage market, the socio-economic position of their offspring's parents-in-law. On one hand, that the family plays a crucial role in shaping income inequality – through behaviour that forges intergenerational links between parents' and children's wealth – is not surprising. On the other hand, for centuries and in several countries, marriage has represented one of the primary institutions by which socio-economic mobility and social stratification took place. For example, consider the development of a national marriage market in London and Bath in the second half of the eighteenth century. This 'market' greatly widened the pool of potentially satisfactory spouses from the point of view of upper-class parents, because it dramatically increased the number of potential spouses who would meet the necessary financial and social.

In joining the intergenerational mobility and the marital sorting literatures, this paper develops an economic model that relies on utility-maximising behaviour by all agents (parents and children), in line with Becker and Tomes (1979, 1986), Becker (1991) and Mulligan (1997, 1999). Not only does the model provide us with a new richer interpretation of the intergenerational elasticity parameter than mechanical models typically do, but it also allows us to improve our understanding of the role of the socio-economic status of parents-in-law in a somewhat more formal way than in previous studies.

For the empirical analysis, we use information on the Goldthorpe-Hope score of occupational prestige as a measure of status and samples of about 2,400 men (sons) and 2,300 women (daughters) drawn from the British Household Panel Survey 1991-1999. We find that the intergenerational elasticity is around 0.2 for men and between 0.17 and 0.23 for women. On average, the intragenerational correlation is lower, and of the order of 0.15 to 0.18, suggesting that the returns to human capital, which is transmitted across generations by altruistic parents, contribute more to social status than assortative matching in the marriage market.

Higher estimates are reported when measurement error issues are taken into account. But even an intergenerational elasticity of 0.2 implies low social mobility when the measure of socio-economic status is the Goldthorpe-Hope score of occupational prestige. In fact, we find that an individual whose father's status is at 25th percentile has a 0.29 chance of remaining in the bottom quartile, a 0.45 chance of rising above the median and a 0.08 change of reaching the top decile. But the child of a poorer father, whose status is at the tenth percentile, will have a 0.34 chance of being the bottom quartile, only a 0.40 chance of rising above the median and only a 0.06 chance of reaching the top decile. Further, for an individual whose father is at 95th percentile, these three probabilities are 0.17, 0.60 and 0.15 respectively. Naturally, the occupations at the bottom quartile of the fathers' distribution are very different from the occupations in the top decile. In the bottom quartile, 42 percent of the fathers are involved in such occupations as truck drivers, coal mining labourers and workers, farm labourers and security guards; while 70 percent of the fathers in the top decile are managers (marketing, sales, bank, service industries and transport) and professionals (engineers, architects, university professors, medical doctors, solicitors and chartered accountants). Using the earnings data on children, those occupational differences are also reflected in substantial pay differentials. For example, the 1999 average monthly earnings is £1,069 for men and women in the lower-level occupations, while men and women at the top-level occupations earn £1,497, approximately 40 percent more.

We also find strong nonlinearities. Both intergenerational and intragenerational elasticities tend to increase with parental status, producing an asymmetry such that upward mobility from the bottom is more likely than downward mobility from the top. Again, in most cases, the intergenerational elasticity is larger than the intragenerational correlation. But for a few groups in the population, the opposite is true. This is the case of men in the second quartile of the parent's status distribution, 55 percent of which correspond to occupations such as decorators, carpenters, plumbers, roofers and lineworkers. It is also the case of slightly better-off women whose parents are in the third quartile of the prestige distribution, 50 percent of which correspond to occupations such as non-professional engineers, lower-rank managers, local government workers and primary school teachers. For these groups of individuals, 'marrying up' may still be an important mechanism to occupy a specific social position. In other words, their socio-economic mobility is greatly affected by their marriage market decisions, possibly to a greater extent than by their labour market behaviour.

We offer four possible explanations for the finding that upward mobility from the bottom is more likely than downward mobility from the top. Three of such explanations – one based on mean-displacement shifts in the occupational prestige distribution, another based on life-cycle effects and the third based on differential measurement errors – do not find strong support in our data. The fourth explanation is based on the notion of intergenerational transmission of social capital and intellectual capital. The evidence, which we derive from friends' occupational prestige, and own organisation membership and activity and education, supports the idea that richer parents are likely to have a larger and more valuable stock of both social capital and intellectual capital to pass on to their children.

1. Introduction

This paper examines how the socio-economic position of parents may influence the socio-economic position of their offspring. By considering sorting in the marriage market, the paper also analyses the extent to which the socio-economic position of parents and parents-in-law are correlated. This correlation, in turn, may be relevant for the offspring's economic success through the investment that parents-in-law have made in their partner's human capital and subsequent financial transfers from parents-in-law. In economics, there is an extensive literature on the transmission of economic success from generation to generation, and there is a smaller and mainly theoretical literature on the marital choices that match individuals in the marriage market. Rarely have these two literatures been joined (Pencavel, 1998).

That the family plays a crucial role in shaping income inequality has long been recognised by economists. For instance, Knight (1935) identified the family as the principal social institution that fosters income inequality through behaviour that forges intergenerational links between parents' and children's wealth (see also Parsons, 1975). Ever since the contributions of Becker and Tomes (1979, 1986) and Loury (1981), economists have increasingly paid attention to the issue of income (or earnings) inequality within families over generations.[1] Most of the discussion has focused on getting a precise estimate of the degree of intergenerational mobility, which typically entails stringent data requirements (Mazumder (2001), and references therein).[2] However, besides the early works mentioned above, relatively few recent studies have contributed to enhance our understanding of the processes that lie beneath the intergenerational elasticity parameter. Notable exceptions are the studies

[1] An excellent recent survey of the (primarily empirical) literature on intergenerational mobility is in Solon (1999). Other surveys of the theoretical literature can be found in Mulligan (1997) and Behrman (1997).

[2] The usual approach has been to estimate a log linear regression of economic status of the child in family *i*, y_i^{child}, on the same measure of economic status for his/her parent(s), y_i^{parent}, of the form:

$$y_i^{child} = a + by_i^{parent} + \varepsilon_i,$$

where ε_i is a white-noise error term. The slope coefficient *b* is the intergenerational elasticity parameter and measures the degree of regression towards the mean in economic stature.

by Mulligan (1999), Checchi, et al. (1999), Altonji and Dunn (2000), and Han and Mulligan (2001). An aim of our paper is indeed to unravel some of the behavioural forces that determine the intergenerational mobility parameter. For this purpose we extend the standard theoretical model proposed by Becker and Tomes (1979, 1986) to include marital sorting considerations. This leads us to the other relevant literature.[3]

The seminal work by Becker (1973, 1974) and its subsequent extensions (Becker, 1981; Lam, 1988) provide an important foundation for the economic theory of the family.[4] Of direct salience to our study is the notion of assortative mating on spouses' traits, which has received considerable empirical attention by many social researchers, sociologists in particular (e.g., Mare, 1991; Kalmijn, 1994).[5] For centuries and in several countries, marriage has represented one of the primary institutions by which socio-economic mobility and social stratification took place (Goody, 1983). To substantiate this point, we provide a few distinctively British examples, but examples for other countries can be found in Outhwaite (1981). Elliott (1981) documents that, in the early seventeenth century, London kinship networks were a highly effective medium for promoting upward social mobility for farmers' daughters: 40 percent of these women married gentry and high-status tradesmen husbands, compared to only 28 percent of otherwise similar non-migrant women living in the rural areas around London. In this case, geographic mobility (from the countryside to London) and kin were key in the marriage market. Another example is the development of a national marriage market in London and Bath in the second half of the eighteenth century, which greatly widened the pool of potentially satisfactory spouses from the point of view of upper-class parents, because it dramatically increased the number of potential spouses who would meet the necessary financial and social qualifications (Stone, 1977). Analysing a group of 200

[3] In our empirical work, however, the measurement error issue will receive special attention.
[4] Recent extensive surveys of this literature are in Bergstrom (1997) and Weiss (1997).
[5] An economic analysis of the interactions between martial sorting and the macroeconomy (in particular, inequality, fertility differentials and per capita output) is in Fernandez et al. (2001).

working-class families sampled in London between 1943 and 1946,[6] Slater and Woodside (1951) report high and significant correlation coefficients between husbands and wives on a number of characteristics (e.g., age at marriage, stature and intelligence) and lower coefficients on other measures (e.g., social background, pre-marital sex experience and clinical rating of neurotic disposition). Thus, besides undoubtedly important considerations of reciprocity, intimacy and love, many generations of individuals seem to have used their marriage to trade up or down their inherited socio-economic position. In particular, there appears to be a systematic relationship between the socio-economic position that individuals (or their parents) occupy and the marital choices that match those individuals in the marriage market.[7]

In joining the intergenerational mobility and the marital sorting literatures, this paper develops an economic model that relies on utility-maximising behaviour by all agents (parents and children), in line with Becker and Tomes (1979, 1986), Becker (1991) and Mulligan (1997, 1999). Not only does the model provide us with a new richer interpretation of the intergenerational elasticity parameter than mechanical models typically do (see Goldberger (1989) and the discussion in Mulligan (1999)), but it also allows us to improve our understanding of the role of the socio-economic status of parents-in-law in a somewhat more formal way than in previous studies (e.g., Lam and Schoeni, 1993).

A brief review of what we know about Britain is in order. There is a relatively small empirical economic literature on intergenerational mobility.[8] Using a sample of father-son

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[6] The three commonest types of occupation among the men were in various sectors of the building trade (carpenters, painters and decorators); transport workers and railwaymen; and in clerical jobs. They accounted for by about one-third of men in the sample. Two-thirds of women were clerks, domestics, factory hands, shop assistants and workers in tailoring and dressmaking.

[7] A growing number of contributions extend Becker's (1973) neoclassical marriage market model by analysing marriage markets with frictions, whereby match creation is time consuming and matching precludes further search (Shimer and Smith, 2000; Burdett and Coles, 2001). In the model developed below, we also depart from the neoclassical assignment model with no search frictions, but these will be incorporated in a simpler fashion.

[8] There is, however, a larger sociological literature on the transmission of economic (dis)advantage from generation to generation, and especially on the degree of mobility in education and occupational status. See, among others, Halsey et al. (1980), Goldthorpe (1980) and Erikson and Goldthorpe (1992). Bjorklund and Jantti

pairs with fathers coming from the 1950 Rowntree inquiry in York and sons subsequently traced (in the late 1970s), Atkinson et al. (1983) extend the earlier work by Atkinson (1981) and find that the intergenerational earnings elasticity is of the order of 0.4 to 0.5. On the other hand, the regression coefficient relating the living standards of parents and children is between 0.15 and 0.20 (the latter making allowance for possible measurement error).[9] Dearden et al. (1997) analyse the National Child Development Survey (NCDS), which follows a cohort of all individuals born in Britain in a week in March 1958. They find that the extent of intergenerational mobility is overall limited. Their regression estimates suggest that the intergenerational mobility parameter is of the order of 0.4 to 0.6 for men and 0.45 to 0.7 for women. Blanden et al. (2001) extend the work by Dearden et al. (1997) by exploring the extent to which the degree of intergenerational mobility has changed over time. For this purpose, they compare the 1958 NCDS cohort to another cohort of individuals drawn from the British Cohort Survey (BCS), which is a longitudinal survey of all children born in a week in April 1970. Their results suggest a sharp fall in intergenerational mobility between these two cohorts, even though they are only twelve years apart. Their highest estimates from the BCS are, however, between 0.18 and 0.25 for men and between 0.17 and 0.23 for women,[10] which are substantially lower than those reported in Dearden et al. (1997). They argue that it is the use of family income (needed to make the cross-cohort comparison possible) rather than father's earnings as the independent variable of interest that produces the different magnitudes in the estimated intergenerational elasticities.

Contemporary evidence on assortative mating in Britain is even scarcer. Using data from the British Household Panel Survey, Chan and Halpin (2000) find that educational

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(2000) offer an interesting review of the existing sociological literature, which measures mobility in class and status, and its links to the economic literature, which measures mobility in earnings and income.

[9] "Living standards" are measured by total family income net of tax payments, other deductions (e.g., superannuation payments) and housing costs, while family income is defined by husband's and wife's earnings, state income and other income (e.g., interest and capital income).

[10] Their lower estimates from the NCDS are, at most, between 0.12 and 0.13 for women and men, respectively.

homogamy (i.e., the correlation between the wife's and husband's schooling) has remained fairly stable over the second half of the twentieth century, with about 40 percent of men and women marrying a partner with the same educational level. As a consequence of the expansion of further education which has seen a dramatic increase particularly in women's schooling, their results also suggest that men are decreasingly likely to marry down and women decreasingly likely to marry up. For a sample of married individuals drawn from the British Cohort Study, Fernandez (2001) reports a schooling homogamy (i.e., the correlation between wife's and husband's schooling) of 0.5, with homogamy increasing as inequality (or segregation) increases.

The rest of the paper has the following structure. In Section 2 we present a model of intergenerational transmission of education and assortative mating in education. Section 3 describes the data used to estimate the relationships of interest, and discusses the main measurement problems and the econometric method. The estimates presented in Section 4 indicate that the intergenerational elasticity is about 0.2 for men and 0.17-0.23 for women. The correlation between own parents' and parents-in-law's social status is slightly lower, of the order of 0.15 to 0.17 for men and 0.16 to 0.18 for women. This suggests that labour market characteristics (such as returns to education) and parental altruism, which are behind the intergenerational mobility parameter, are on average more important than assortative matching in the marriage market in shaping people's socio-economic position. After allowing for plausible measurement error, the estimates of the intergenerational elasticity reach their highest values around 0.3-0.35 for men and around 0.3-0.4 for women, whereas the assortative matching correlation ranges between 0.2 and 0.3 for men and between 0.25 and 0.35 for women. There are strong nonlinearities in these estimates. But, as found also in other studies, our results suggest that the intergenerational elasticity is greater as parental social status increase, the opposite of our (and Becker and Tomes's) conjecture. In Section 5, we discuss this result and provide four possible interpretations, one of which – based on the notion of intergenerational transmission of "social" and "intellectual" capital – is new. Section 6 concludes.

2. Model

This section presents a simple model of intergenerational transmission of human capital and assortative mating on education. The objective of this model is to illustrate to what extent parental income (or socio-economic position) interacts with the socio-economic position of the parents-in-law in determining the child's subsequent socio-economic position.[11] As discussed in the Introduction, we draw both from the intergenerational mobility literature – particularly Becker and Tomes (1979, 1986) – and from the assortative mating literature – particularly Boulier and Rosenzweig (1984). All parents are altruistic and, besides caring for their own consumption, invest in their offspring's human capital. Upon reaching adulthood, the child endowed with given levels of attractiveness and market luck and provided an optimal level of human capital, chooses an optimal search time to select a partner in the marriage market.[12] Thus, the model is composed of two sub-problems which are solved backwards. First, the child (whose generation is denoted by t) selects an "optimal" mate. Second, the parents (generation t-1) determine the "optimal" schooling level for their offspring, with knowledge of the offspring's behaviour in the marriage market. In formalising this idea, we assume that all parents act as if they maximise a single utility function (i.e., consensus parental preferences) and that each family has one child only. Following most of the literature, we also assume that each parent-child pair plays a noncooperative game and

[11] In this context, the terms "socio-economic position" and "income" are interchangeably used. In Section 3, we shall define the measure of socio-economic position used in our empirical analysis.
[12] "Marriage", "cohabitation" and "live-in partnership" are all synonymous in the context of this model, as long as they refer to an ex-ante durable long-term union.

reaches a Nash equilibrium (which is efficient), that is, each agent maximises its objective function taking the action of the other as given.

2.1 The child's problem

Let τ_t be marital search time after schooling completion and H_t be educational attainment (both measured in years of age), and let y_t^s denote the post-school per-period "single" earnings and y_t^m the per-period marital earnings. The young adult child, endowed with a level of attractiveness ε_t and "market luck" e_t, will choose an optimal search time (i.e., an "optimal" mate) to maximise

$$(\tau_t - H_t)y_t^s + (T - \tau_t)y_t^m$$

subject to

(1) $$y_t^s = \mu_1 H_t + \mu_2 e_t + \mu_3 \varepsilon_t$$

and

(2) $$y_t^m = \gamma_1 H_t + \gamma_2 e_t + \gamma_3 \varepsilon_t + \gamma_4 \tau_t H_t^p,$$

where τ_t lies between 0 and T, T is the child's generation life span (measured in years), and H_t^p is the potential partner's human capital (in years). Incomes in both marital states (single and married) are assumed to be related to own schooling, attractiveness and market luck. For given H_t and H_t^p, equation (2) also assumes that a greater search time will lead to a greater marital income as long as γ_4>0. This is consistent with the arguments outlined in Becker et al. (1977), Keeley (1977) and Weiss and Willis (1997). For a given H_t, a smaller τ_t means a lower age at marriage. That is, people marrying young are less likely to be informed about themselves, their potential mates and the marriage market. They face therefore lower returns to marriage through a lower search time. Because of search frictions, a longer search time increases the "weight" of y_t^s in the child's objective function (τ_t – H_t), but y_t^s is unaffected by τ_t. On the other hand, a higher τ_t increases y_t^m (with γ_4>0) but shortens the time period over which marital income is enjoyed.[13]

Another central proposition of the theory of marriage formulated to explain the set of marital matches that occur in an economy is that there is a non-random assortative mating of agents with respect to complementary characteristics and non-wage incomes (Becker, 1973, 1974; Mare, 1991; Behrman et al., 1994; Weiss, 1997). We incorporate this idea in the following matching function:

(3) $$H_t^p = \beta_0 + \beta_1 H_t,$$

where the parameter β_1 captures the degree of assortative mating in education: with $\beta_1 > 0$, the greater level of schooling for one spouse the greater the level of schooling for his/her partner (educational homogamy). As in Boulier and Rosenzweig (1984) and Behrman et al. (1994), equation (3) may be augmented to allow for market luck and attractiveness (and other endowments) to affect the schooling of the spouse. The main implications of the model are, however, unaffected by their inclusion. Moreover, our data do not allow us to control for unobserved endowments in estimation, and thus we keep (3) as simple as possible.

The optimal (interior) search time depends on all the technological parameters of the earning functions (1) and (2), the assortative matching function (3) as well as the time horizon *T*, and is given by

(4) $$\tau_t = \frac{T}{2} - \frac{(\gamma_1 - \mu_1)H_t + (\gamma_2 - \mu_2)e_t + (\gamma_3 - \mu_3)\varepsilon_t}{2\gamma_4(\beta_0 + \beta_1 H_t)}.$$

[13] In this model parents are assumed to be too poor to make monetary transfers to their children (i.e., the marginal utility of their private consumption is greater than the marginal utility of transfers). In the case in which parents are not financially constrained (i.e., they give monetary transfers to their children), the income equations (1) and (2) must allow for some curvature in H_t so that their optimisation problem has an interior solution. We do not consider this case because the parents' investment decision is the same as that obtained in the existing models of intergenerational mobility, whereby the child's educational level does not depend on parents' income.

The optimal search time clearly increases with the length of the life cycle. Provided that the returns to human capital, market luck and attractiveness are greater in the married state than in the single state, expression (4) reveals that the optimal search time decreases in e_t and ε_t, and increases in H_t as long as $\beta_1[(\gamma_2-\mu_2)e_t+(\gamma_3-\mu_3)\varepsilon_t]>(\gamma_1-\mu_1)\beta_0$.

In equilibrium, only marital income is affected by marital search time and is

(5) $$y_t^m=\delta_0+\delta_1 H_t+\delta_2 e_t+\delta_3\varepsilon_t,$$

where $\delta_0=\beta_0\gamma_4 T/2$, $\delta_1=(\gamma_1+\mu_1+\beta_1\gamma_4 T)/2$, and $\delta_j=(\gamma_j+\mu_j)/2$ for j=2,3.

2.2 The parents' problem

Parents care about their own consumption, C_{t-1}, and the anticipated human-capital-dependent income of their (adult) child, y_t. This is the weighted sum of per-period incomes in the post-school unmarried and married states, the weights being the (ex-ante) probability of staying single and the (ex-ante) probability of getting married, respectively:

(6) $$y_t=(1-\phi_t)y_t^s+\phi_t y_t^m,$$

where ϕ_t is the probability that the child will marry, and y_t^s and y_t^m are given by (1) and (5) respectively.[14] The parents' consensus utility is assumed to be Cobb-Douglas in child's income and own consumption

(7) $$(1-\alpha)\log C_{t-1}+\alpha\log y_t,$$

with $\alpha\in[0,1]$ reflecting the relative preference for child income as against own consumption. Parents then choose H_t to maximise (7) subject to the expected income of their child

See, among others, Becker and Tomes (1986), Becker (1989), Mulligan (1999), Han and Mulligan (2001) and Ermisch and Francesconi (2001).

[14] For simplicity, we assume that ϕ_t is exogenous and equal to a constant, ϕ. The main results are unchanged if we allow ϕ_t to depend on e_t and ε_t. The results are also unchanged if ϕ is assumed to depend on parents' own market luck and attractiveness, e_{t-1} and ε_{t-1}, respectively.

(expression (6))—which depends on the solution of the child's problem (equations (1) and (5))—and to their budget constraint

(8) $$y_{t-1} = p_c C_{t-1} + p_h H_t,$$

where p_c and p_h are the unit prices of parental consumption and child's human capital. As in all the models that do not explicitly consider the marriage market, the solution to this problem shows that the optimal level of child's human capital is positively determined by parents' income through parental preferences. In fact, letting $\eta_0 = -(1-\alpha)\phi\delta_0 / D$, $\eta_1 = \alpha / p_h$, $\eta_j = -(1-\alpha)(\mu_j + \phi\delta_j)/D$, for j=2,3, $D = (1-\phi)\mu_1 + \phi\delta_1$, and assuming p_c=1, the solution will be

(9) $$H_t = \eta_0 + \eta_1 y_{t-1} + \eta_2 e_t + \eta_3 \varepsilon_t.$$

Using (9) to compute y_t^s and y_t^m in (1) and (5) respectively, the anticipated income of the child given in (6) is, in equilibrium, a function of parents' income, that is,

(10) $$y_t = \pi_0 + \pi_1 y_{t-1} + u_{1t},$$

where $u_{1t} = \pi_2 e_t + \pi_3 \varepsilon_t$, $\pi_0 = \eta_0 \kappa_1$, $\pi_1 = \eta_1 \kappa_1$, $\pi_j = \eta_j \kappa_1 + \kappa_j$, for j=2,3 and $\kappa_j = (1-\phi)\mu_j + \phi\delta_j$, for j=1,2,3.[15] The parameter π_1 is called intergenerational correlation (or intergenerational transmission parameter), and will play a crucial role in our subsequent analysis. The potential spouse's parents will make the same human capital investment in their child as that given in (9),

(9a) $$H_t^p = \eta_0^p + \eta_1^p y_{t-1}^p + \eta_2^p e_t^p + \eta_3^p \varepsilon_t^p.$$

On the other hand, H_t^p can be expressed as a function of y_{t-1} through the assortative mating equation (3) and the optimal human capital investment (9):

[15] According to equation (10), intergenerational mobility will be higher (lower π_1) if: (a) there are lower returns to education for children in either the single state (μ_1 is lower) or the married state (δ_1 is lower); or (b) children's human capital is less sensitive to parental social status (η_1 is lower). This last condition occurs when the degree

(9b) $H_t^p = \beta_0 + \beta_1(\eta_0 + \eta_1 y_{t-1} + \eta_2 e_t + \eta_3 \varepsilon_t)$.

Equating (9a) to (9b) yields

(11) $y_{t-1}^p = \lambda_0 + \lambda_1 y_{t-1} + u_{2t}$,

where $u_{2t} = \lambda_2 e_t + \lambda_3 \varepsilon_t + \lambda_4 e_t^p + \lambda_5^p \varepsilon_t^p$, $\lambda_0 = (\beta_0 + \beta_1\eta_0 - \eta_0^p)/\eta_1^p$, $\lambda_j = \beta_1\eta_j/\eta_1^p$, for *j*=1,2,3, and $\lambda_{j+2} = -\eta_j^p/\eta_1^p$, for *j*=2,3. The parameter λ_1 captures what we call intragenerational correlation in socio-economic status. Equations (10) and (11) constitute the basis of our empirical analysis, whose main focus is therefore on the reduced-form parameters π_1 and λ_1. Both parameters are nonlinear functions of complex deep parameters, such as preferences (degree of altruism), the technological parameters of the income and matching function equations, and other parameters relevant in the marriage market (probability of marrying). In the case of liquidity-constrained families, equation (10) shows that parents' income directly affects children's socio-economic position, while equation (11) reveals that parent's income is also related to the parents-in-law's socio-economic position via the marriage market. This intragenerational correlation is relevant for the offspring's socio-economic success to the extent that parents-in-law have invested in the (potential) partner's human capital and make transfers to the couple.

Using (10) to express y_t^p as a function of y_{t-1}^p, and using again (9a) and (9b) to solve for y_{t-1}^p, we can also link y_t^p and y_{t-1} as follows:

(12) $y_t^p = r_0 + r_1 y_{t-1} + u_{3t}$,

where $u_{3t} = r_2 e_t + r_3 \varepsilon_t + r_4 e_t^p + r_5 \varepsilon_t^p$, $r_0 = \pi_0^p + \pi_1^p(\beta_0 + \beta_1\eta_0 - \eta_0^p)/\eta_1^p$, $r_j = \beta_1\eta_j\pi_1^p/\eta_1^p$, for *j*=1,2,3, and $r_{j+2} = \eta_j^p\pi_1^p/\eta_1^p$, for *j*=2,3. Equation (12) provides direct information about the impact of the marriage market on socio-economic position through r_1. But both r_1 in (12) and

of altruism α is lower. These characteristics of the intergenerational mobility process are similar to those formalised in Solon (2001).

λ_1 in (11) have the same informational content as long as $\pi_1 = \pi_1^p$, because $r_1 = \lambda_1 \pi_1$. So, π_1 from (10) and λ_1 will give us all is needed to unpack the intergenerational and marriage market links. In addition, it is immediate to see that $r_1 < \pi_1$, if $\pi_1 = \pi_1^p$, $\eta_1 = \eta_1^p$ and $\beta_1 < 1$. Thus, under these conditions, it is always the case that the intergenerational correlations within a family line contribute more to the socio-economic position of an individual than the intergenerational correlations determined via the marriage market. In our empirical work, however, we shall estimate (11) rather than (12). As compared to the estimation of r_1, the estimation of λ_1 does not involve any loss of information, but has the advantage of comparing the socio-economic distributions of two groups of individuals (parents and parents-in-law) from the same generation. This is likely to minimise the problems induced by intergenerational changes in the status distributions (or variances), which can be relevant when inequality varies across generations (Solon, 1992). As Section 3 illustrates, this is the case for our measure of socio-economic status.

Keeping this comparability problem in mind, it is interesting to see whether it is the intergenerational transmission parameter or the intragenerational correlation arising in the marriage market that contributes more to one's socio-economic position. Parental income will have a greater impact on y_t than on y_{t-1}^p if, ceteris paribus, $\pi_1 > \lambda_1$. Assuming that $\eta_1 = \eta_1^p$ and $\phi = 1$, this inequality is satisfied when[16]

$$\alpha\delta_1 / p_h > \beta_1,$$

[16] Notice that, for a given horizon T, the parameter π_1 may be negative as long as μ_1 is sufficiently negative, even if both the return to human capital in the married state, γ_1, and the return to marital search, γ_4, are positive. The parameter λ_1 is positive (negative) if there is positive (negative) assortative mating in education in equation (3), i.e., β_1>0 (β_1<0).

that is, when the real return to human capital in the married state weighted by the parental relative preference for child income is larger than the degree of assortative mating in education in the marriage market.[17]

To the extent that u_{1t} and u_{2t} in equations (10) and (11) are unobservable, we may ask whether our model places any restriction on the covariance structure of these two error terms that can be tested on data. Although we cannot bound the error correlation to specific values within the unit circle, the model does allow us to sign it under some assumptions. For convenience, we impose that $\mathrm{E}(e_t) = \mathrm{E}(\varepsilon_t) = \mathrm{E}(e_t \varepsilon_t) = \mathrm{E}(e_t \varepsilon_t^p) = \mathrm{E}(e_t^p \varepsilon_t) = 0$. The covariance between u_{1t} and u_{2t} is then given by

$$\text{(13)} \qquad \mathrm{Cov}(u_{1t}, u_{2t}) = \pi_2 \lambda_2 \sigma_e^2 + \pi_3 \lambda_3 \sigma_\varepsilon^2 + \pi_2 \lambda_4 \, \mathrm{Cov}(e_t, e_t^p) + \pi_3 \lambda_5 \, \mathrm{Cov}(\varepsilon_t, \varepsilon_t^p),$$

where σ_x^2 is the variance of x, with x=e,ε. Clearly the sign of (13) can be either positive or negative depending on the sign and magnitude not only of $\mathrm{Cov}(e_t, e_t^p)$ and $\mathrm{Cov}(\varepsilon_t, \varepsilon_t^p)$ but also of all the other relevant parameters. If (a) $\eta_j = \eta_j^p$, for j=1,2,3, (b) the attractiveness of an individual is positively correlated with the attractiveness of his/her partner (i.e., $\mathrm{Cov}(\varepsilon_t, \varepsilon_t^p) > 0$),[18] and (c) $\alpha(\mu_j + \phi\delta_j) > \phi\mu_j$, for j=2,3 (which is satisfied when the parents' preference for child income is sufficiently high and when the attractiveness and market luck returns in the married state are greater than in the single state), then Cov(u_{1t},u_{2t}) is positive as long as the labour market lucks of partners are positively correlated. But if matches in the marriage market occur with risk sharing motives (e.g., Rosenzweig and Stark, 1989), then the correlation between e_t and e_t^p is likely to be negative, and the sign of Cov(u_{1t},u_{2t}) becomes ambiguous even when conditions (a)-(c) are satisfied.

[17] It is clear that our model allows the intergenerational mobility parameter π_1 (as well as λ_1) to be less than zero and greater than one, values which would be hard to justify in a purely mechanical model (Goldberger, 1989).
[18] This is in line with the argument presented in Burdett and Coles (2001).

3. Data

3.1 Samples and socio-economic status measures

The model presented in Section 2 implies that to examine the relationship between intergenerational mobility in socio-economic status and assortative matching requires data that provide information on the socio-economic position of individuals, their parents and their parents in-law. The data used in our estimation come from the first nine waves of the British Household Panel Survey (BHPS) collected over the period 1991-1999.[19] While no data on parents' incomes or earnings are provided, the BHPS elicits respondents to provide information about their parents' occupations when they were aged 14. This information is then used to compute an index of occupational prestige according to the technique proposed by Goldthorpe and Hope (1974), which was based on the rankings of occupations made by a random sample of individuals interviewed throughout England and Wales in 1972.[20] Thus matching partners[21] in all available waves of BHPS data provides us with longitudinal (on-going) information for each partner's Goldthorpe-Hope (GH) score, and childhood information (at age 14) about his/her parents' and his/her parents-in-law's GH scores (both father and mother).[22]

[19] In Autumn 1991, the BHPS interviewed a representative sample of 5,500 households, containing about 10,000 people. The same individuals have been re-interviewed each successive year. If they leave their original household to form new households, then all the adult members (i.e., those aged 16 or more) of the new households are also interviewed as part of the survey. Children in the original households are also interviewed once they reach the age of 16. Some 88 percent of the original BHPS sample were re-interviewed for the second wave (1992) and the response rates from the third wave onwards have been consistently higher than 95 percent. The BHPS data are, therefore, unlikely to suffer from any serious bias resulting from attrition. This means that the sample used in this analysis remained broadly representative of the population of Britain as it changed through the 1990s. Other information about the BHPS can be found at http://www.iser.essex.ac.uk/bhps/docs.

[20] Goldthorpe and Hope (1974) suggest that the scale which results from their occupational prestige grading exercise should not be viewed as a grading of social status *stricto sensu*, i.e., as tapping some underlying structure of social relations of "deference, acceptance and derogation" (p. 10). It should instead be viewed as "a judgement which is indicative of what might be called the 'general goodness' or … the 'general desirability' of occupations" (pp. 11-12).

[21] The analysis focuses both on people who are legally married and on people who are in live-in partnerships or cohabitations.

[22] The information on parents' occupation was collected only in waves 1, 8 and 9 (1991, 1998 and 1999 respectively). An alternative measure of socio-economic position would be permanent incomes of children, parents and parents-in-law. But the BHPS does not provide any retrospective information on (own and parents') incomes. Similarly, to date, the panel is too short to allow us to follow individuals over their entire childhood

The GH scale ranges from 5 to 95. It has been long documented that it is highly correlated with earnings.[23] Between 1991 and 1999 (when the BHPS collects data also on respondents' incomes), the BHPS reveals a correlation between the log of gross monthly earnings and the GH index of 0.70 for men and 0.75 for women. Thus, to the extent that labour income represents a substantial fraction of total income,[24] the GH scale of occupational prestige is likely to be a good measure of people's socio-economic position. Furthermore, it turns out that in Britain the position of individuals in the occupational hierarchy is relatively stable over time. That is, the occupational position of workers at the time when they permanently enter the labour force is a reliable predictor of their occupational position when they are at the end of their working career.[25] Therefore, the GH scale is also an adequate measure of people's permanent socio-economic status.

There are, however, three problems in the way in which y_{t-1} is measured in our study. First, we only have a single GH score observation for parents. As shown by Solon (1992) and Zimmerman (1992), this is likely to produce downward biased estimates. Second, our data on parents are obtained from adult children (and not from the parents themselves) and refer to the

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(when we observe their parents' occupations and incomes) and into their adulthood (when they report their own occupations and incomes and, through matching in the marriage market, those of their partners).

[23] Using data from the British New Earnings Survey, Phelps Brown (1977) reports a strong relationship between median gross weekly earnings by occupation and the GH scale, with a rise of 1 unit in the scale of occupational status being associated with an increase of 1.031 percent in earnings. Nickell (1982) also reports a correlation between the GH index and the average hourly earnings by occupation of 0.85 using data from the National Training Survey.

[24] Using data from the Family Expenditure Survey, Goodman et al. (1997) document a declining contribution of earnings to household income, from over three-quarters in the early 1960s to about 60 percent in 1993. Most of the reduction seems to be accounted for by lower levels of employment particularly between 1979 and 1984. From the mid 1990s, the income share made up by labour income has further increased.

[25] To document this, we use the employment history file collected in the 1993 wave of the BHPS and obtain a measure of the GH score at the time of the permanent entry into the labour market for all respondents who provided full information at that wave. We then collect the GH score for workers at the end of their careers (ages 56-65 for men and ages 55-60 for women) during the survey period (1991-99). We find that the correlation between an individual's position in the occupational hierarchy (as measured by the GH scale) on entry into the labour market and at end of the working career is almost 0.60 for men and 0.62 for women, respectively. Averaging the GH scores over those ages for each worker produces similar results, that is, 0.57 for men and 0.63 for women. Nickell (1982) argues that the best single predictor of a man's occupational position when he is 60 is his occupational position when he first entered the labour market. Indeed, using a ranking of occupations based on the average hourly earnings within each of the 396 occupations contained in the KOS classification (Key Occupations for Statistical Purposes), Nickell finds a correlation between the entry position in the labour market and the position at age 60 of 0.62.

parents' occupation when the child was aged 14. This retrospective questioning is likely to be affected by recall error, which may be more or less severe depending on the child's age at the interview. The third limitation arises when a new occupation is observed or an old occupation disappears.[26] This will clearly generate a misclassification error. Both recall and misclassification errors will again bias our estimates downward. Given that the bias produced by these three types of error has the same sign, our estimates of π_1 and λ_1 will therefore be downward biased. This should be kept in mind while interpreting the results below, although our main goal is to compare π_1 and λ_1, and, to the extent that they are contaminated by the same sort of error, this comparison may suffer less from this problem.

We perform our analysis on two samples, each of which uses a different definition of y_{t-1} and y_{t-1}^{p}. In the first ('Sample 1'), these variables refer to father's and father-in-law's GH scores, respectively. It includes 2,046 women (daughters) and 2,151 men (sons). In the second sample ('Sample 2'), y_{t-1} refers to the average of the father's and mother's GH scores for the men and women who have both parents' information, and to the father's or mother's GH score for those who have only one parent's information. The definition of y_{t-1}^{p} is analogous. Sample 2 includes a total of 2,266 women and 2,382 men.

Table 1a reports quartile transition matrices for children's GH score and fathers' and fathers-in-law's GH scores for the individuals in Sample 1. Looking at the numbers on the main diagonal of father-son and father-daughter pairs, we see that the biggest proportions of children who remain in the same quartile as their fathers are either in the top (i.e., highest GH

[26] Goldthorpe and Hope (1974) do not consider this issue. However, they describe (see their chapter 6) some of the problems that are relevant in collecting and coding occupational data, which in turn are of crucial importance to apply the GH scale. In particular, the use of the scale requires that occupations be coded according to official statistics procedures, because the occupations should be identifiable in the official statistics index of occupational titles. But "uncertainty arising from incomplete occupation descriptions can be resolved only by the invention of ad hoc rules" (p. 69) over and above those embodied in the GH scale. To the extent that such measures are used, the new GH scale may not be expected to reflect the occupational prestige that it originally intended to capture. The emergence (disappearing) of a new (old) occupation can presumably generate the same sort of uncertainty.

score) or in the bottom (lowest GH score). As found by Zimmerman (1992) using US data drawn from the National Longitudinal Survey, upward and downward mobility are essentially the same for men (sons): 39 percent remain in the top quartile if their fathers were in the top quartile of their GH score distribution, and another 38 percent remain in the bottom quartile had their fathers been in the bottom quartile too. The corresponding percentages for women (daughters) are 41 percent (top) and 35 percent (bottom), which suggests that upward mobility from the bottom of the distribution is more likely than downward mobility from the top. This asymmetry in women's mobility is in line with the results presented in Dearden et al. (1997). The transition rates from father-in-law's GH score to men or women's GH score again indicate low mobility both at the top and at the bottom of the GH distribution, with upward and downward mobility being relatively unlikely. Table 1b reports the quartile transition matrices in GH scores of men and women and their parents and parents-in-law for the individuals in Sample 2. The figures in this table confirm the high persistence (low mobility) both at the bottom and at the top of the GH distribution. This time the asymmetry of upward mobility from the bottom being more likely than downward mobility from the top emerges for both parents/son pairs and parents/daughter pairs. The larger persistence at the top is also evident in the case of parents-in-law/daughter pairs, while upward mobility from the bottom and downward mobility from the top are quite similar for parents-in-law/son pairs.

In Figure 1 we plot the GH score distributions for fathers, parents and children by child gender. The child distributions are bimodal, but more clearly so in the case of women (daughters). The spike on the left (which corresponds to lower occupational prestige) is also more pronounced for women. This may reflect the fact that women are more concentrated in low-prestige occupations and can be found in fewer occupations than men (Heath and Payne, 2000). More importantly, the figure documents a clear example of mean displacement, with the GH score distribution for fathers (or parents) being skewed to the right, and the child's distributions being bimodal, or (as in the case of sons) even skewed to the left (the skewness coefficient is about –0.043). The mean displacement is associated with an increase in the average GH score over time as well as with higher dispersion (see also Table 2). This reveals that the GH score distribution is not stable across generations, and, in particular, low-prestige occupations that characterised the parents' distribution have increasingly disappeared in the child's distributions. This feature of the GH score distribution may turn out to be important for the interpretation of the estimates in Section 4.

3.2 Attractiveness, market luck and estimation method

The term u_{1t} in equation (10) is a linear function of own attractiveness (ε_t) and market luck (e_t). The term u_{2t} in (11) is also a function of the partner's attractiveness and market luck (ε_t^p and e_t^p). We assume that there exists a vector of observable characteristics, which may partially reflect ε_t, e_t, ε_t^p, and e_t^p, as follows:

(14) $$u_{1t} = \pi_{21}\mathbf{X}_t + \omega_{1t}$$

(15) $$u_{2t} = \lambda_{21}\mathbf{X}_t + \lambda_{22}\mathbf{X}_t^p + \omega_{2t},$$

where $\mathbf{X}_t$ contains region of residence, local labour market conditions,[27] industry, union coverage and employing sector, and $\mathbf{X}_t^p$ includes the partner's industry, union coverage and sector (clearly, region of residence and local labour market conditions are the same for both partners). These variables are meant to approximate characteristics such as geographic proximity, common interests and values, which are typically posited to be crucial factors in mate selection (e.g., Adams, 1971) and are likely to be associated with attractiveness in the marriage market and luck in the labour market. For example, partner's industry may be directly related to market luck and may also signal a particular type of partner. The terms ω_{1t}

and ω_{2t} are disturbances with zero mean and finite variance, and are assumed to be correlated. This is because the processes that generate y_t and y_{t-1}^p may share common unobservables. Thus, we jointly estimate equations (10) and (11) augmented with (14) and (15) using a seemingly unrelated regression (SUR) method. Both y_{t-1} and all the variables in $\mathbf{X}_t$ and $\mathbf{X}_t^p$ are taken as exogenous.

To align the age and experience profiles of socio-economic status across generations, both $\mathbf{X}_t$ and $\mathbf{X}_t^p$ are augmented with a quartic polynomial in age and quadratic polynomials in part-time and full-time experience. These variables may also reflect individual aspects that are potentially correlated to attractiveness and market luck. Summary statistics of all these variables by sample and gender are presented in Table 2.

4. Results

4.1 Benchmark estimates

Table 3 presents the SUR estimates of π_1 and λ_1 for men and women in Samples 1 and 2 obtained using two different specifications. Specification [1] does not include any control for market luck and attractiveness nor does it attempt to eliminate time-varying life-cycle components.[28] Specification [2] does include such controls as in equations (14) and (15).

Looking at the results obtained with Sample 1, the estimates of π_1 and λ_1 from specification [1] are respectively 0.255 and 0.198 for men (sons), and 0.247 and 0.210 for women (daughters). Even after $\mathbf{X}_t$ and $\mathbf{X}_t^p$ are included in estimation, π_1 is always significantly greater than λ_1 for men, while, in the case of women, they are not statistically

[27] Local labour market conditions are measured by the ratio of unemployment stock to vacancies stock. The geographic unit of this measure is given by 306 matched job centres (providing information on the vacancies stock) and travel-to-work areas (providing information on the unemployment stock).

[28] All regressions for women are selectivity corrected with the semiparametric two-step procedure described in Vella (1998). They include a cubic polynomial of the single index function that governs the selection into employment. The variables included in this selection equation are listed in the note of Table 3.

different. This may occur because the real rate of return to human capital investments weighted by parents' relative preference for child income is greater than the degree of educational homogamy for married men, but not for married women. It should be noted, however, that the estimates obtained from Sample 2 reveal a somewhat different story. In fact, for both men and women, we find that π_1 is always significantly greater than λ_1, regardless of the specification. Therefore, although our estimates are below their true values because of measurement error, it is likely that the returns to education, which is passed on through intergenerational investments of altruistic parents, contribute more to socio-economic position than assortative matching in the marriage market.

Part of the reason why we obtain different results from the two samples could be that Sample 2 includes mothers' and mothers-in-law's GH scores. Not only do we fail to account for mother's selection into employment (due to lack of instruments), but also it is likely that the transmission of socio-economic position from mother to son differs from the same transmission from mother to daughter.[29] Restricting the definition of y_{t-1} and y_{t-1}^{p} to include only mothers' and mothers-in-law's GH scores and performing again the same analysis as that reported in Table 3 (specification [2]), we find that π_1 is smaller than λ_1 for men (0.132 versus 0.152), but it is greater for women (0.149 versus 0.136). In neither case, however, is this difference statistically significant at conventional levels. Bearing in mind the measurement error problem, these results suggest that men's and women's socio-economic position interact differently with their mothers'. Men's occupational prestige seems to be more strongly correlated with their mothers-in-law's than their own mothers', but is also more strongly associated with their own fathers' than their fathers-in-law's occupational prestige. Women's

[29] Blanden et al. (2001) argue that the changing influence of mothers' earnings may underpin some of the observed fall in intergenerational mobility between the 1958 NCDS cohort and the 1970 BCS cohort. They too find that the effect is stronger for daughters than sons.

prestige, on the other hand, displays the opposite pattern of correlations (stronger with their own mother's and father-in-law's prestige).

Another way to see the potentially different impact of mothers' and fathers' occupational prestige on sons' and daughters' is to re-estimate (10) and (11) with both mother's and father's GH scores as separate independent variables. The estimates of the corresponding intergenerational elasticities are reported in the Appendix Table A1. For men, we can reject the hypothesis that mother's and father's prestige scores have the same impact on their sons, with the effect of fathers' score being significantly greater than the effect of mothers' at the 3 percent level. However, we cannot reject the hypothesis that the two effects are equal in the case of the fathers-in-law equation. We find the opposite result for women. That is, the effects of mothers' and fathers' prestige on their daughters' are statistically identical, whereas in the case of the fathers-in-law equation they differ at the 7 percent level. Investigating such gender and parental differences is not the focus of this paper as that would require the development of a more complex model than that in Section 2, where parental investments in child human capital are allowed to differ for sons and daughters and where father's and mother's incomes are endogenously determined. We leave this for future research.

In all cases, the values of π_1 (and λ_1) are small in comparison to what Atkinson et al. (1983) suggest as a benchmark, i.e., Galton's regression of fathers' height on sons' height which produced an intergenerational transmission coefficient of about 0.5 (Galton, 1886). But is their impact actually small? To obtain suggestive results, we follow the approach proposed by Solon (1992) and assume that long-run status (i.e., the permanent component of log GH score) is normally distributed across generations. We then calculate the probability that a child's socio-economic position lies in different intervals of the status distribution as a function of the percentile of the father's or parents' status. The results of this exercise suggest that, conditional on father's position, the extent of occupational mobility is relatively small even if π_1=0.2. For example, an individual whose father's status is at the 25th percentile has a 0.29 chance of remaining in the bottom quartile, a 0.45 chance of rising above the median and a 0.08 chance of reaching the top decile. The child of an even poorer father, whose status is at the tenth percentile, will have a 0.34 chance of being in the bottom quartile, a 0.40 chance of rising above the median and only a 0.06 chance of reaching the top decile. Further, for an individual whose father is at the 95th percentile, these three probabilities are 0.17, 0.60 and 0.15 respectively. Naturally, the occupations at the bottom quartile of the fathers' distribution are very different from the occupations in the top decile. In the bottom quartile, 42 percent of the fathers are involved in such occupations as truck drivers, coal mining labourers and workers, farm labourers and security guards; while 70 percent of the fathers in the top decile are managers (marketing, sales, bank, service industries and transport) and professionals (engineers, architects, university professors, medical doctors, solicitors and chartered accountants). Using the earnings data on children, those occupational differences are also reflected in substantial pay differentials. For example, the 1999 average monthly earnings is £1,069 for men and women in the lower-level occupations, while men and women at the top-level occupations earn £1,497, approximately 40 percent more.

Finally, notice that Cov(ω_{1t},ω_{2t}) in Table 3 is always positive, regardless of sample, specification and gender. This may suggest that the risk sharing motive in the marriage market is either inconsequential or offset by other characteristics of the labour and marriage markets that are not included in $\mathbf{X}_t$ and $\mathbf{X}_t^p$, e.g., large sample dispersion in unobserved market luck and attractiveness, and high correlation in partners' unmeasured attractiveness (see Burdett and Coles, 2001). The *p*-value of the Breusch-Pagan test for independent equations is always zero, suggesting that the error terms associated with y_t and y_{t-1}^p, ω_{1t} and ω_{2t} respectively, are indeed correlated (which makes the OLS estimates of (10) and (11) inefficient).

4.2 Adjusting for measurement error

With the estimates of π_1 being around 0.2 for men and between 0.17 and 0.23 for women, our intergenerational elasticities are not only below Galton's benchmark but are also smaller than those presented in the earlier British studies of intergenerational earnings mobility, notably Atkinson et al. (1983) and Dearden et al. (1997). Instead, our estimates are closer to (and possibly higher than) those shown in Atkinson et al. (1983), when they use living standards rather than earnings as their variables of interest, and in Blanden et al. (2001), where they regress the log of children's earnings on the log of parental income. In this paper, we use occupational prestige as our dependent and independent variables of interest, again not individual earnings, so the comparison with previous results may not be appropriate. Clearly, the variables used to measure socio-economic status matter. Our estimates of π_1 and λ_1 are nonetheless below their true values, because y_{t-1} is likely to be measured with error. Unfortunately, the BHPS data do not allow us to use any standard method to eliminate or reduce this bias, e.g., averaging parental GH scores over time (because we only have information on parents' GH score when the child was aged 14) or using instrumental variables techniques (because we do not have information on parents' education or other family background variables that are uncorrelated with the child's socio-economic position).

Generalising the studies by Atkinson et al. (1983), Solon (1992) and Zimmerman (1992), Mazumder (2001) shows that the "attenuation" factor which biases π_1 toward zero is given by

$$\frac{\sigma_*^2}{\sigma_*^2 + m_1\sigma_{transitory}^2 + m_2\sigma_{error}^2},$$

where σ_*^2 denotes the variance of the true y_{t-1}, $\sigma_{transitory}^2$ is the variance of transitory fluctuations of y_{t-1} and σ_{error}^2 is the variance of the measurement error in y_{t-1}. The terms m_1 and

m_2 account for serial correlation in transitory fluctuations of occupational prestige and averaging $y_{t\text{-}1}$ over time. Assuming that $\sigma_{transitory} = s_1\sigma_*$ and $\sigma_{error} = s_2\sigma_*$, with s_1 and s_2 are proportions or multiples of the standard deviation of the true $y_{t\text{-}1}$, the attenuation factor becomes $1/(1+m_1s_1^2+m_2s_2^2)$. We therefore adjusted our estimates of π_1 multiplying them through by $(1+s_1^2+s_2^2)$, after having set m_1=m_2=1 since we have information on $y_{t\text{-}1}$ for only one year (i.e., when the child was aged 14).[30]

Table 4 reports the results from this adjustment for three values of s_1 (0, 1/3 and 2/3) and seven values of s_2, ranging from 0 to 1.5 at intervals of 0.25 length. Because of the high persistence in the occupational hierarchy that characterises British workers (see Section 3 and Nickell (1982)), we expect s_1 to be relatively small, but we have no priors on the size of s_2. It should be noticed that around the mean value of the observed father's GH score reported in Table 2, the three modal occupations are metal-working production occupations, plumbing and welding and other semi-skilled manual occupations (they account for almost 50 percent of the occupations). A score reduced by one standard deviation is associated with occupations such as farm and farming-related workers and unskilled builders and porters (which comprise 40 percent of the total workers with this reduced score); while an increase of one standard deviation from the mean corresponds to the GH score of teachers (in secondary and primary schools), welfare, community and social workers, and nurses (which account for approximately 70 percent of the workers with this higher score). A change in occupational prestige by one-and-half standard deviation leads to even larger changes in the associated occupations: 90 percent of the workers with one-and-half standard deviation lower score are waiters, bar assistants, caterers, care and educational assistants and cleaners, while 45 percent of the workers with one-and-half standard deviation higher score are civil servants, local

[30] Using data on earnings histories from Social Security records and the Survey of Income and Program Participation, Mazumder (2001) finds that the intergenerational elasticity in earnings between fathers and sons is of the order of 0.6 or higher.

government administrators and middle-rank managers, and another 15 percent are teachers and nurses. These changes are unarguably large and probably beyond the actual incidence of measurement error.[31] In the case of men from Sample 1, Table 4 shows that, with s_1=0, the estimates of π_1 will increase between 0.204 and 0.624, while with s_1=2/3, the estimated intergenerational elasticity will range between 0.277 and 0.709. Similar dramatic variations in π_1 are found also for women and for Sample 2. Measurement error and error induced by transitory fluctuations in occupational prestige do, therefore, have an impact on the estimated mobility. But they bring our estimates close to those reported in Dearden et al. (1997) for Britain or in Mazumder (2001) for the United States only under highly implausible values of s_1 (=2/3 or higher) and s_2 (=1 or higher).

Setting s_1=2/3 and s_2=1.5 yields the largest values of λ_1 at about 0.621 for men and 0.661 for women (both from Sample 2). It is difficult to ascertain whether or not these are believable estimates, because we cannot refer to any existing evidence on the correlation between own parents' and parents-in-law's social status. However, we are confident that the extent of mobility in terms of assortative mating is likely to be higher (i.e., lower values of λ_1) even after accounting for measurement error. First, as in the case of π_1, the values of s_1 and s_2 are untenably high. Second, we may derive a rough estimate of β_1 in equation (3), one of the key deep parameters behind λ_1, by regressing partner's education on own education (both expressed in years)[32] while keeping region of residence, local labour market conditions, own

[31] A drop in GH score by half standard deviation leads to scores which are associated to various occupations in the building sector (decorators, carpenters and plumbers) and other semi-skilled and unskilled occupations, accounting for about 40 percent of workers with these lower score. A correspondent increase leads to occupations such as lower-level managerial occupations (sales, office and service managers), which account for almost 60 percent of the workers. Although there is always an element of discretionary judgement, the changes in GH scores induced by an error of half standard deviation (or less) seem to be more plausible.

[32] In the BHPS, respondents report their highest qualification level, which we map into years of education by considering how long it usually takes to obtain it. If an individual indicates no qualification, 9 years of education were assigned. Individuals with apprenticeship and other lower commercial qualifications but short of O levels were assigned 10 years; those with GCSEs and O levels or equivalent qualifications were assigned 11 years; those with advanced vocational qualifications and A levels were assigned 13 years; those with a degree level of further education (such as nursing, teaching and other higher education diplomas) were assigned 15 years; those

age, union coverage, sector and industry constant. The OLS estimates of β_1 from these regressions are 0.338 (t-ratio=18.985) for men and 0.374 (t-ratio=24.840) for women. These estimates are substantially lower than the correlation reported in Fernandez (2001) – the difference might be due to the different cohorts under analysis and the inclusion of controls – but are in line with those found by Chan and Halpin (2000). The measurement-error-adjusted estimates of λ_1 in Sample 2 reach such values with s_1=2/3 and s_2=0.75.

4.3 Results by quartile of parental GH score

The model in Section 2 imposes an extreme form of capital market imperfection. In fact, the only transfer parents can make in favour of their offspring is in the form of human capital investment. No other type of intergenerational transfer (e.g., cash, services and bequests) is admitted. There is, however, substantial empirical evidence of "inter vivos" gifts and some evidence of bequests at death, motivated either by altruism or by exchange.[33] But, even if altruism is widespread or exchange motives are deeply grounded in the economy, for many households binding nonnegativity constraints may block their manifestation.

For this reason, we partition each sample by quartile of the GH score of the father (Sample 1) or of the parents (Sample 2). To check whether this partition is useful, we analyse savings and housing tenure patterns by quartile of parental occupational status. The fifth wave (1995) of the BHPS gathered information on personal financial circumstances. We use this information to compute the amount of money each individual holds in total personal savings.[34] Interestingly, those individuals whose father was in the top quartile of his GH score

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with a university (or polytechnic) first degree were assigned 16 years; and finally those with higher university degrees were assigned 18 years. For a relatively similar assignment procedure, see Fernandez (2001).

[33] In this context, deciding whether the transfer is accidental, egoistic, altruistic or motivated by exchange is irrelevant. For a review of the theoretical and empirical literature on intergenerational transfers, see Laitner (1997).

[34] Savings are defined as the sum of savings accounts (in bank, Post Office and building society) and investments (from national saving certificates, premium bonds, unit trusts, personal equity plans, shares and national savings/building society insurance bonds).

distribution when they were aged 14 hold an amount of money in savings that is about 75 percent greater than that held by individuals whose father was in the bottom quartile (16,500 pounds versus 9,500 pounds in 1995). We obtain the same figures when we partition the joint GH score distribution of parents. In addition, using the whole nine waves of BHPS data, only 9 percent of men and women whose father/parents was/were in the top quartile live in local authority housing as opposed to nearly 30 percent of those whose father/parents was/were in the bottom quartile. To the extent that both lower savings and social housing tenure capture a greater exposure to liquidity constraints, the partitioning of the father's or parents' GH score distribution by quartile seems to separate out the families that are likely to be financially constrained from those that are not. Of course, any ability or taste heterogeneity will make it more difficult to identify the presence of families that are borrowing constrained (Han and Mulligan, 2001).

Tables 5a and 5b report the SUR estimates of π_1 and λ_1 for men and women by quartile of father's (Sample 1) and parents' (Sample 2) GH score, respectively. These estimates are obtained using specification [2] described in the previous sub-section. The estimates from Sample 1 reveal a high degree of upward mobility from the bottom. Indeed, π_1 and λ_1 are not significantly different from zero for individuals whose father was in the bottom half of the distribution of occupational prestige (except for λ_1 in the case of women in the bottom quartile of the father's GH score distribution). The extent of intergenerational mobility is instead limited at the top. For both men and women whose father was in the third quartile of the GH score distribution, we find that λ_1 is greater than π_1, that is, educational homogamy may play a more substantial role than the joint contribution of the degree of parental altruism and the returns to human capital investments in shaping people's socio-economic position. We find the opposite result (i.e., $\pi_1 > \lambda_1$) for individuals whose father was in the top quartile of the GH score distribution, although for women the difference between the two estimates is not statistically significant at conventional levels.

When we look at the impact of the parents' occupational prestige on children's and parents-in-law's GH scores (Sample 2), Table 5b provides us with results that are in part similar to and in part different from those reported in Table 5a. They are similar because, again, the intergenerational elasticity is lower for individuals who come from families characterised by low occupational prestige. Indeed, both π_1 and λ_1 are lowest for men and women whose parents were in the bottom quartile of their GH score distribution. The results in Table 5b are also different because they highlight different patterns of socio-economic mobility. The highest estimated value of π_1 is reported for men and women with parents in the second lowest quartile (0.367 and 0.471 respectively), although even those in the top half of the parents' GH score distribution show a somewhat limited downward mobility (particularly women). In the case of men, the largest value of λ_1 is once more detected in the second quartile of the parents' prestige distribution, while in the case of women λ_1 peaks in the third quartile (0.752 and 0.615 respectively). Perhaps as a consequence again of the different transmission mechanism that links the socio-economic positions of mothers to sons' and daughters', the estimates in Table 5b suggest that the largest fraction of the inter- and intra-generational correlations is within the interquartile range of the parents' GH score distribution. They also document the absence of any correlation at the bottom and some limited downward mobility from the top of the distribution.[35]

In sum, the relationship between π_1 and λ_1 is far from straightforward. Although the results in Table 3 show that π_1 is generally greater than λ_1, this is not always the case when individuals are stratified by percentile of their fathers' or parents' occupational prestige.

[35] There is again little evidence in support of a risk sharing motive in the marriage market. Notice also that the correlation of the residuals between equations is always significantly different from zero at the 5 percent level in all cases, except for women in the top quartile of parents' GH score (Table 5b).

Depending on the sample (i.e., the measure of y_{t-1}), some groups of men and women are characterised by values of λ_1 greater than π_1 (e.g., women in the third quartile of fathers' and parents' GH score, women at the bottom quartile of the father's GH score, and men in the second quartile of parents' GH index). These individuals' social position (and mobility) may therefore be greatly affected by their marriage market decisions, possibly to a greater extent than by their labour market behaviour. Furthermore, despite the differences across the two samples, the estimates in Tables 5a and 5b consistently point at the result that π_1 and λ_1 increase (perhaps non-monotonically) as parental social status increase. This is the opposite of our (and Becker-Tomes's) conjecture, because parents in higher socio-economic positions are less likely to be financially constrained. Several earlier studies have reported this problem, e.g., Behrman and Taubman (1990) and Mulligan (1999) for the United States,[36] Corak and Heisz (1999) for Canada, and Dearden et al. (1997) for Britain. Various hypotheses can be offered to justify this result. In the next section, we will explore four explanations in detail.

5. Why is there more upward mobility from the bottom than downward mobility from the top?

5.1 Shifting GH score distributions

The first explanation is based on the hypothesis that the GH score distributions systematically shift away from the bottom over time. This is in part due to the way in which the GH index of occupational prestige is constructed and to the misclassification error issues discussed in Section 3. The idea is that if people whose parents were in the bottom quartile of the GH score distribution moved up simply because low-prestige occupations have disappeared, then the results of Tables 5a and 5b would not be surprising. In this world, we would expect to observe high upward mobility from the bottom only because, in the child's generation, there are no

[36] Zimmerman (1992) finds high intergenerational elasticities at both extremes of the status (earnings, wage, Duncan index) distributions. On the contrary, Mazumder (2001) reports evidence of lower intergenerational

longer low-prestige occupations to be employed in and not because there is a genuine improvement in the social position of people coming from the bottom of the prestige distribution. Children of low-occupational-prestige parents may improve upon their parents' absolute position but still be at the bottom of their GH score distribution, relative to children of high-occupational-prestige parents.[37]

To test this hypothesis we need to partition the children's GH score distribution using cut-off values that are not determined on the fathers' (or parents') distribution, exactly because of the mean displacement that is at work across generations. We therefore partition the children's distribution by quartile using the children's distribution itself, and re-estimate equations (10) and (11) using specification [2] on Samples 1 and 2 for men (sons) and women (daughters) separately. It is easy to show that, as a result of this partition on the dependent variable, both π_1 and λ_1 are downward biased, but the bias is larger in the case of π_1. The results of this exercise are reported in Table 6. As expected, the estimates of π_1 are extremely low and do not show any clear pattern by quartile. In the case of women, however, the highest values of this parameter are again found for those individuals in the top quartile, confirming that upward mobility from the bottom is indeed stronger than downward mobility from the top. The estimates of λ_1 are significantly higher, despite the downward bias generated by the partitioning on y_t. Interestingly, as we move from the bottom to the top quartile of the child's GH score distribution, we observe a positive gradient in λ_1, particularly in the case of men. A higher correlation between parents' and parents-in-law's socio-economic position at the top may simply imply that a replication of status is more likely at the top than in other parts of the child's status distribution (through the marriage market). Although a substantial mean displacement in the GH score distribution is at work across generations, these results therefore

mobility for families with lower net worth.

[37] The quartile transition matrices reported in Tables 1a and 1b are not useful here because the quartiles are determined on parents' and children's distributions separately. However, Figure 1 documents a secular shift of the GH score distribution away from the bottom.

suggest that upward mobility from the bottom is higher than downward mobility from the top, even when we partition on the child's prestige distribution. Of course, different results may emerge if we use some other procedure to partition the child's distribution, and so we cannot generalise our findings. However, with the estimates reported in Table 6, we have greater confidence in the results of Tables 5a and 5b.

5.2 Ability and life-cycle effects

The second explanation relies on a life-cycle argument. In an economy with "equal opportunities" – in which children with equal abilities have equal options in their human resource investments and their expected income (see Behrman and Taubman, 1990) – and capital market imperfections –whereby human capital is a poor collateral to lenders – high-ability children who inherit a low social position from their parents may go through a career progression which takes them from low-level to high-level occupations. Naturally, the occupational gradient for high-ability children who already start at a relatively high social position is much less steep. It is possible, instead, that low-ability children of high-status parents move down in the occupational ladder, but only gradually. If this explanation were valid, we would expect different patterns of π_1 by birth cohort, with older (younger) cohorts displaying lower (higher) intergenerational elasticities, particularly at the bottom of the social status distribution.

Tables 7a and 7b report the SUR estimates of π_1 and λ_1 by cohort for Samples 1 and 2, respectively. We distinguish three broad birth cohorts, which identify three groups of children roughly similar in size.[38] At the bottom quartile of the father's GH score distribution (Table 7a), the estimated intergenerational elasticity is 0.011 for men born before 1950, 0.130 for men born between 1950 and 1959 and –0.096 for men born after 1959. These estimates are not statistically different from each other at the 5 percent level. In the case of women, none of the estimates of π_1 by cohort are significantly different from zero. At the other end of the prestige distribution, we find weak correlations for the youngest and the oldest male cohorts, but the estimated π_1 for the 1950-59 cohorts of men is large and precisely measured (0.472 with a standard error of 0.096). For women at the top of the social status, both the youngest and the oldest cohorts show high intergenerational correlations. The picture we get from the estimates in Table 6b is qualitatively similar. These results clearly fail to provide an empirical support for the hypothesis based on ability and life-cycle effects.[39]

A further piece of evidence may come from the relationship between and individual's socio-economic status on entry into the labour market and his/her socio-economic status at different points of the life cycle. If our hypothesis based on life-cycle effects were at work, we would expect to see stronger correlations earlier rather than later in the life cycle for all individuals, and especially for those whose parents were at the bottom of the occupational distribution. We use the 1993 employment history file of the BHPS, which provides information on all jobs (including occupation and GH score) from the time an individual left full-time education to September 1990. We regress the GH score averaged over three periods of the life cycle (ages 36-45, 46-55 and 56-65) on the GH score at entry into the labour market by quartile of parental occupational position (constant schooling). Table 8 reports the results from such regressions for two definitions of parental position in the occupational hierarchy (echoing what we did earlier with Samples 1 and 2) and for men and women separately. Contrary to our life-cycle effects argument, there is little evidence of a declining pattern of correlations for individuals from low-status families (except for men, when we use the

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[38] For the sake of space, R^2 statistics, number of individuals and person-wave observations and estimates of corr(ω_{1t}, ω_{2t}) are not reported, but can be obtained from the authors.

[39] Both tables also report the SUR estimates of π_1 and λ_1 by cohort for the entire sample (in the columns labelled 'Mean'). From these estimates, we find that, regardless of the measure of y_{t-1}, intergenerational mobility has increased for men, but stayed relatively stable for women over time (the decrease in π_1 from 0.197 to 0.175 in

parents' GH score). Indeed, the correlations between an individual's position in the occupational prestige hierarchy on entry into the labour market and at different points of his/her life cycle are fairly stable. With only one other exception (i.e., women from the third quartile of parents' GH score), we cannot find any substantial decline in these occupational correlations. All these results suggest that life-cycle considerations may not be appropriate to explain why, in our data, there is more upward mobility from the bottom than downward mobility from the top.

5.3 Differential measurement errors

Another explanation rests on the hypothesis that different portions of the distribution of $y_{t\text{-}1}$ are affected by different measurement errors. For this hypothesis to be able to explain our problem, measurement error should be larger at the bottom than at the top of the distribution.

Our benchmark is given by the estimates in Tables 5a and 5b. As we did in sub-section 4.1, we multiply them through by $(1+s_{1j}^2+s_{2j}^2)$, where j=1,…,4 indexes the quartile of the parental GH score (in ascending order). Setting $s_{11}=0.5$, it is easy to see that the value of π_1 for men whose father was in the bottom quartile of his prestige distribution (Sample 1) will increase to 0.473 if $s_{21}=5$ (Table 5a). This new value of π_1 is twice as large as that found for men in the top quartile, but still lower than that of men in the third quartile of the GH score distribution with no adjustment for measurement error. Although the values of s_{11} and particularly s_{21} are implausibly high and $y_{t\text{-}1}$ is assumed to be measured with no error in the top half of the distribution, the extent of upward mobility from the bottom is at best equivalent to the degree of downward mobility from the top. We can reach a similar

Table 6a is not significant at the 5 percent level). On the other hand, women experienced the steepest decline in λ_1. The decline for men from 0.173 to 0.149 in Table 6b is not significant at standard levels.

conclusion also for the other estimates of π_1 in Tables 5a and 5b.[40] Therefore, an explanation based solely on differential measurement errors must rest on unbelievably large errors at the bottom of the parental GH score distribution. We find this explanation unpersuasive.

5.4 Intellectual capital versus social capital

Our last explanation is based on the notions of “intellectual heritage” and “social heritage”. These notions have been developed in previous studies by Sjogren (1998) and Hassler and Rodriguez Mora (2000). To see how this works, we rewrite equations (10) and (14) in the form

(16) $$y_t = \pi_1 y_{t-1} + \omega_{1t},$$

from which the intercept term π_0 and the $\mathbf{X}_t$ vector are dropped for expositional convenience.

Let $\xi_{t\text{-}1}$ denote parental “social capital”, that is the social advantages that result from a particular upbringing, and $\psi_{t\text{-}1}$ denote the parents’ “intellectual capital”, or innate ability and genetic heritage. We may think of ω_{1t} and $y_{t\text{-}1}$ as being linearly determined by both $\xi_{t\text{-}1}$ and $\psi_{t\text{-}1}$ as follows

(17) $$\omega_{1t} = d_1\xi_{t-1} + d_2\psi_{t-1} + v_t$$

(18) $$y_{t-1} = \rho_1\xi_{t-1} + \rho_2\psi_{t-1} + \varphi_{t-1},$$

where v_t and $\varphi_{t\text{-}1}$ are uncorrelated disturbances with zero mean and finite variance. Assume for simplicity that $E(\xi_{t\text{-}1})=E(\psi_{t\text{-}1})=Cov(\xi_{t\text{-}1},\psi_{t\text{-}1})=0$. Substituting expressions (17) and (18) into (16), it is straightforward to see that

(19) $$\mathrm{cov}(y_t, y_{t-1}) = \rho_1(d_1 + \pi_1\rho_1)\sigma_\xi^2 + \rho_2(d_2 + \pi_1\rho_2)\sigma_\psi^2 + \pi_1\sigma_\varphi^2,$$

[40] With s_{11}=1 and s_{21}=1.5, the value of π_1 for men in the bottom quartile of the parents’ GH score (Sample 2) will increase from 0.102 to 0.434, which is higher than the estimates of π_1 for men in the other quartiles. The magnitudes of s_{11} and s_{21} are, however, still untenable.

with $\sigma_j^2 = \text{var}(j)$ and $j = \xi_{t-1}, \psi_{t-1}, \varphi_{t-1}$. Clearly, the intergenerational mobility parameter π_1 is now just one of the ingredients of the intergenerational correlation given in (19). If richer parents have a larger and more valuable stock of social capital to pass on to their children, then it is reasonable to assume that σ_ξ^2 increases with y_{t-1}, while σ_ψ^2 is the same over the entire occupational distribution. On the other hand, if also intellectual capital is largely determined by the parents' social position through, say, genetic transmission, then σ_ψ^2 increases with y_{t-1} too. Under either of these assumptions, $\text{cov}(y_t, y_{t-1})$ is greater for parents with higher values of y_{t-1}.

To test this explanation, we use information on social support networks and organisation membership and activity from various waves of the BHPS. Glazear et al. (2000) use the number of organisation memberships to proxy social capital for a sample of individuals drawn from the General Social Surveys between 1972 and 1998. An additional measure of social capital may be the number of organisations in which an individual is active, because these are likely to capture features of social organisation that facilitate cooperation and coordination for mutual benefit (Putnam, 1993). Furthermore, Pahl (2000) claims that friends provide a powerful form of social capital through trust, mutually reciprocal obligations and their labour market networks. Our proxies of ξ are, therefore, the number of organisations to which an individual belongs, the number of organisations in which he/she is active, and the occupational prestige (GH score) of the best friend.[41] The proxy of ψ is instead education (measured in years).

[41] The BHPS does not collect information on the number of close friends, but only asks for the employment status of the three closest friends. In what follows, the GH score takes value of zero if the best friend is reported to be out of the labour force or unemployed. Our results are unchanged if, for these cases, we impute the best friend's GH score with the GH score of the respondent. The BHPS collects information on the best friend occupation only in waves 2, 4 and 8 (1992, 1994 and 1998), whereas the data on the number of organisation memberships and activities are available for waves 1 through 5 and wave 7.

Before discussing the results of this test reported in Table 9, it is worthwhile looking at the average (first-order) relationship between these measures of social capital and parental social status. The average number of organisations to which men aged 16 or more belong is 1.60 if their fathers were from the bottom quartile of the GH score distribution and 1.84 if their fathers were from the top quartile. Similarly, the average number of organisation in which men are active is 1.44 at the bottom and 1.59 at the top of the GH distribution; whereas the mean GH score of the men's best friend is 44.8 if their father was at the bottom quartile and 54.79 is their fathers were at the top quartile of the occupational prestige distribution. The differences in these means are significant at conventional levels. After controlling for education and a quadratic polynomial in year of birth, Poisson regressions reveal that the number of organisations to which men of the highest social status belong is 0.16 (*t*-ratio=9.63) higher than the number of organisation to which men of the lowest status belong to, and the number of organisations in which men of high status are active is 0.15 (*t*-ratio=7.75) higher than the number of organisation in which men of low status are active. With the same set of controls, OLS regressions show that the GH score of the best friend of men in the top quartile of father's prestige is 5.3 points (*t*-ratio=11.14) higher than the GH score of the best friend of men in the bottom quartile of father's occupational prestige. Similar findings emerge for women and when the measure of parental status is the occupational prestige of both parents. These results suggest that children of richer parents do have a larger and potentially more valuable stock of social capital.

Our hypothesis rests, however, on variances. Table 9 reports the estimated standard deviations in the number of organisations to which men and women belong, the number of organisations in which they are active, and their best friend's GH score by quartile of parental occupational prestige. It is striking that the variation in these three measures of social capital for individuals whose fathers or parents were at the top quartile of the GH score distribution is always larger than the corresponding variation for individuals whose fathers or parents were at the bottom quartile. For example, the standard deviation in the number of organisations to which an individual belongs is 1.239 for men whose fathers were at the top quartile and 1.022 for men whose fathers were at the bottom quartile of the prestige distribution. As the bootstrap standard errors show, this difference is statistically significant at the 5 percent level. Similar results emerge for the other measures of social capital (organisation activity and best friend's occupation), the other measure of parental prestige and for women. Thus, each dimension of social capital analysed here seems to suggest that σ_{ξ}^2 indeed increases with y_{t-1}, probably reflecting the fact that richer parents have a larger stock of social capital to transfer to their children.

Finally, Table 9 reports also the standard deviation in years of education.[42] While the average number of years of education increases monotonically as parental social status increases,[43] their dispersion does not. In the case of men at the top quartile of parental occupational prestige, σ_{ψ} is not the highest, but it is always significantly greater than σ_{ψ} for men at the bottom quartile. Instead, in the case of women at the top quartile of social status, the standard deviation is significantly the highest. Taken all together, these findings suggest that $\mathrm{cov}(y_t, y_{t-1})$ is greater for parents with higher values of y_{t-1}, because both σ_{ξ}^2 and σ_{ψ}^2 increase with the socio-economic position people are in. This is because richer parents are likely to have a larger and more valuable stock of social capital *and* intellectual capital to pass on to their children.

[42] In assigning years of education to qualification levels in the BHPS, we use the same procedure described in section 4.2.

[43] For example, for men whose father was in the bottom quartile of the GH score distribution, the average number of yeas of education is 11.2, while for men whose father was in the top quartile it is 13.7. Differences of the same order of magnitude are also found for women and the other measure of parental prestige. All these differences are statistically significant.

6. Conclusions

This paper presents a model of intergenerational transmission of human capital and assortative mating on education, which allows us to underpin how the socio-economic positions of parents and parents-in-law operate to shape social status across generations. Because earnings or income variables for children, parents and parents-in-law are not typically available in general representative household datasets, our measure of socio-economic status is the Goldthorpe-Hope index of occupational prestige (or general desirability of occupations). Using data from the British Household Panel Survey collected annually between 1991 and 1999, we find that the intergenerational elasticity, π_1, is around 0.2 for men and between 0.17 and 0.23 for women, whereas the correlation between own parents' and parents-in-law's social status, λ_1, is slightly lower, of the order of 0.15 to 0.17 for men and 0.16 to 0.18 for women. This suggests that labour market characteristics (such as returns to education) and parental altruism, which are behind the intergenerational mobility parameter, are on average more important than assortative matching in the marriage market in shaping people's socio-economic position. Although both π_1 and λ_1 are likely to be downward biased because of measurement error, their effects are not negligible as they represent occupational changes which are associated with monthly earnings changes up to 30 percent. After adjusting for transitory fluctuations in socio-economic status and plausible measurement error, π_1 increases to 0.3-0.35 for men and 0.3-0.4 for women, while λ_1 increases to 0.2-0.3 for men and 0.25-0.35 for women, respectively.

There are strong nonlinearities. In particular, both the intergenerational elasticity and the assortative matching correlation tend to increase with parental social status, producing an asymmetry such that upward mobility from the bottom is more likely than downward mobility from the top. This asymmetry is not consistent with existing models (including ours) grounded on parental altruism and borrowing constraints. We offer four possible explanations for this finding, three of which – one based on mean-displacement shifts in the occupational prestige distribution, another based on life-cycle effects and the third based on differential measurement errors – do not find strong support in our data. The fourth explanation is based on the notion of intergenerational transmission of social capital and intellectual capital. The evidence, which we derive from friends' occupational prestige, and own organisation membership and activity and education, supports the idea that richer parents are likely to have a larger and more valuable stock of both social capital and intellectual capital to pass on to their children.

This paper provides one of the first attempts to join some aspects of assortative matching in the marriage market to the intergenerational transmission of human capital investment. In this respect, it follows Pencavel's (1998) suggestion and complements the studies by Fernandez (2001) and Fernandez et al. (2001). Several extensions of this work would be desirable. First, the model may be extended to incorporate search frictions in the marriage market more formally, in the spirit of Shimer and Smith (2000) and Burdett and Coles (2001). This will generate multiple equilibria in the marriage market which may in turn affect the parents' human capital decision and the reduced-form model to be estimated. Another extension of the model is to allow parental inputs to differ between sons and daughters, which may help explain the observed differences by gender in the estimates of π_1 and λ_1 (Behrman et al., 1986). A third extension is to endogenise parents' income by modelling their labour supply decisions. In the British context, this is likely to be relevant not only for mothers but also for fathers (Dickens et al., 1999). Finally, different testable predictions are likely to be generated when we allow children to make choices in the labour market – and not only in the marriage market – as in Fernandez (2001) and Fernandez et al. (2001).

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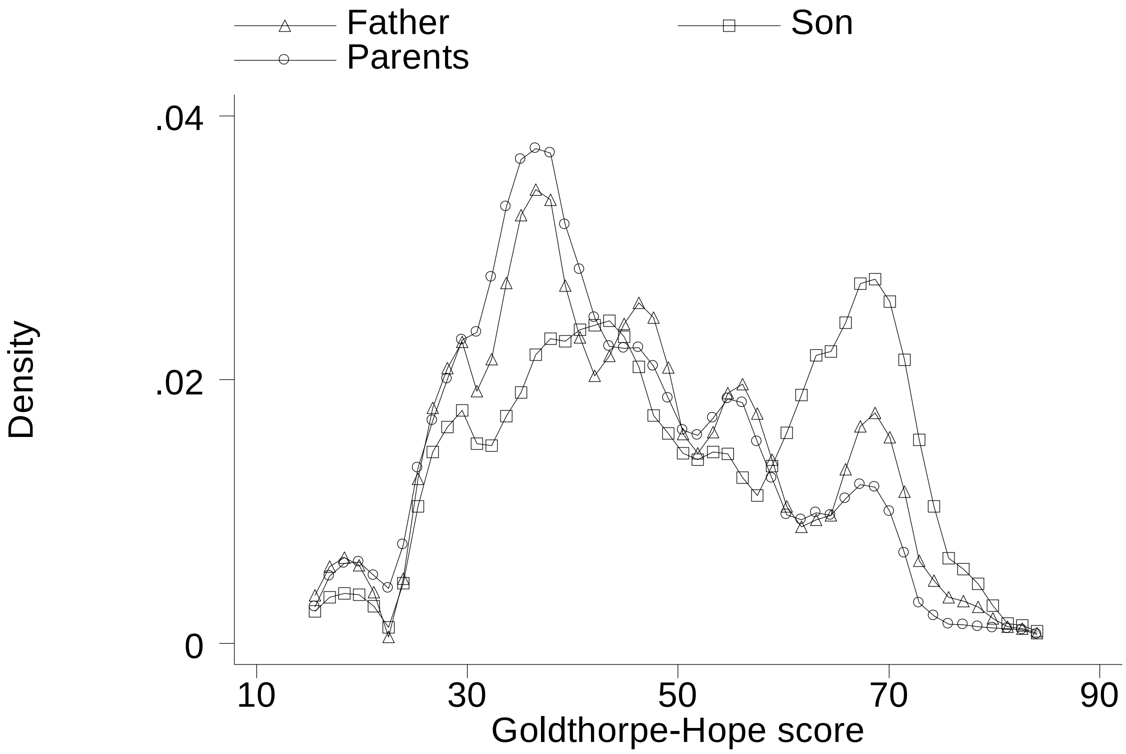
Figure 1: Kernel density estimates of the father's, parents' and child's GH score distributions by child gender

Men (sons)



Women (daughters)

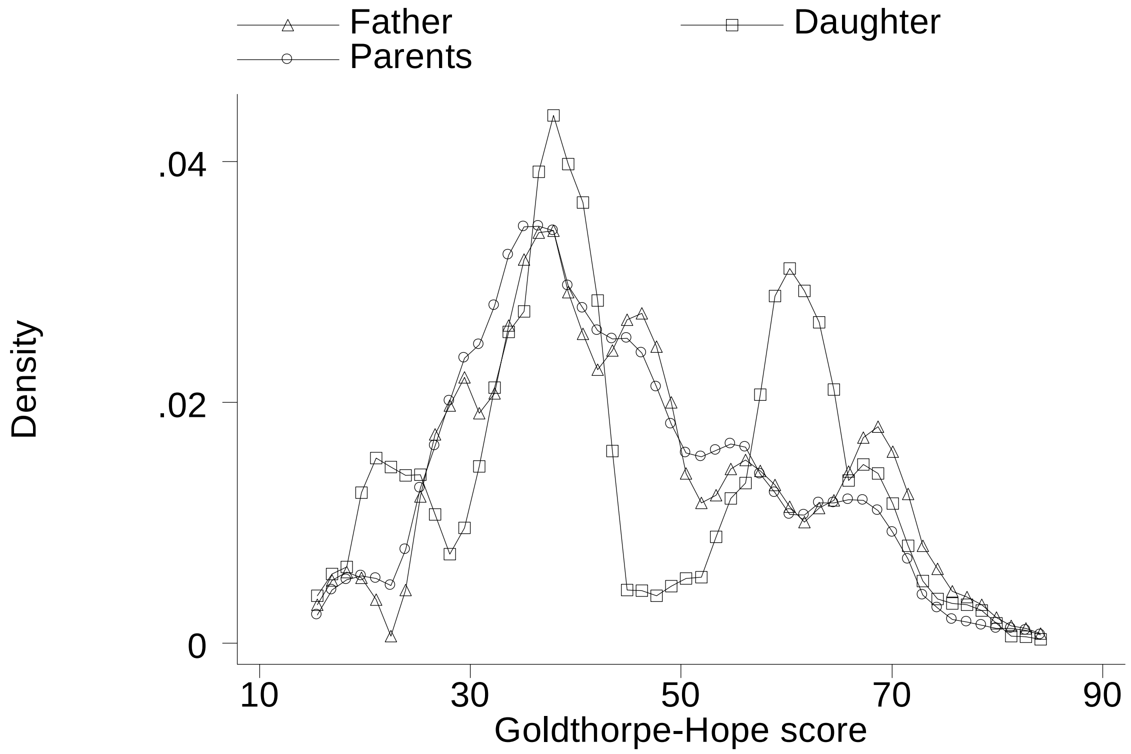


Table 1a
Quartile transition matrices in the GH score across generations (Sample 1)

Father's and father-in-law's quartile (child aged 14)	Child's quartile			
	Bottom	Second	Third	Top
Father/son (*N*=12,357)				
Bottom	0.384	0.271	0.221	0.144
Second	0.310	0.349	0.244	0.219
Third	0.176	0.237	0.227	0.245
Top	0.130	0.144	0.309	0.392
Father-in-law/son (*N*=12,357)				
Bottom	0.366	0.307	0.224	0.181
Second	0.283	0.311	0.260	0.263
Third	0.203	0.214	0.225	0.220
Top	0.149	0.169	0.291	0.336
Father-daughter (*N*=10,646)				
Bottom	0.346	0.310	0.232	0.137
Second	0.240	0.254	0.220	0.194
Third	0.269	0.261	0.261	0.258
Top	0.145	0.175	0.287	0.411
Father-in-law/daughter (*N*=10,646)				
Bottom	0.335	0.279	0.241	0.168
Second	0.223	0.225	0.260	0.234
Third	0.281	0.317	0.251	0.251
Top	0.161	0.179	0.248	0.348

Note: *N* is number of pair-wave observations.

Table 1b
Quartile transition matrices in the GH score across generations (Sample 2)

Parents' and parents-in-law's quartile (child aged 14)	Child's quartile			
	Bottom	Second	Third	Top
Parents/son (*N*=13,687)				
Bottom	0.353	0.252	0.214	0.135
Second	0.283	0.300	0.209	0.182
Third	0.212	0.285	0.256	0.275
Top	0.152	0.163	0.320	0.408
Parents-in-law/son (*N*=13,687)				
Bottom	0.334	0.286	0.188	0.150
Second	0.273	0.265	0.254	0.220
Third	0.201	0.260	0.251	0.294
Top	0.192	0.190	0.307	0.336
Parents-daughter (*N*=11,816)				
Bottom	0.341	0.268	0.196	0.128
Second	0.254	0.282	0.243	0.167
Third	0.231	0.255	0.245	0.300
Top	0.175	0.195	0.316	0.405
Parents-in-law/daughter (*N*=11,816)				
Bottom	0.316	0.267	0.190	0.158
Second	0.260	0.234	0.272	0.219
Third	0.237	0.280	0.247	0.236
Top	0.187	0.219	0.284	0.388

Note: *N* is number of pair-wave observations.

Table 2
Descriptive statistics by sample

Variable	Sample 1 Men	Sample 1 Women	Sample 2 Men	Sample 2 Women
GH score	51.238	46.065	51.048	45.904
	(15.238)	(15.608)	(15.750)	(15.664)
Father's GH score	45.969	46.694		
	(14.757)	(14.939)		
Father-in-law's GH score	46.621	45.467		
	(14.881)	(14.493)		
Parents' GH score			43.885	44.256
			(13.641)	(13.775)
Parents-in-law's GH score			44.281	43.355
			(13.770)	(13.330)
Age	40.274	38.932	40.339	39.024
	(9.607)	(9.742)	(9.659)	(9.773)
Part-time experience	0.356	6.484	0.348	6.531
Full-time experience	21.743	11.949	21.892	12.049
Ethnic groups				
Black	0.005	0.005	0.006	0.005
Indian	0.016	0.014	0.015	0.013
Bangladeshi and Pakistani	0.002	0.001	0.002	0.001
Other (including Chinese)	0.009	0.006	0.008	0.006
White [=base]	0.968	0.974	0.969	0.975
Marital status				
In a live-in partnership	0.134	0.147	0.135	0.147
Married [=base]	0.866	0.853	0.865	0.853
Year of first partnership				
Before 1970 [=base]	0.248	0.281	0.257	0.286
1970-1979	0.270	0.297	0.263	0.290
1980-1989	0.303	0.268	0.304	0.271
1990 and after	0.179	0.154	0.176	0.153
House tenure				
House owner (outright)	0.101	0.109	0.101	0.110
House owner (mortgage)	0.754	0.749	0.752	0.746
In social housing	0.085	0.082	0.085	0.082
In rented accommodation [=base]	0.060	0.060	0.062	0.062
Education				
No educational qualification [=base]	0.149	0.149	0.152	0.160
Less than O level (or equiv.)	0.077	0.106	0.079	0.105
GCSE/O level (or equivalent)	0.180	0.260	0.182	0.256
A level (or equivalent)	0.132	0.106	0.132	0.105
Higher vocational degree	0.305	0.255	0.305	0.252
University and higher degree	0.157	0.124	0.150	0.122
Number of dependent children aged:				
0-2	0.119	0.080	0.119	0.080
	(0.340)	(0.279)	(0.339)	(0.278)

3-4	0.122	0.087	0.123	0.088
	(0.341)	(0.290)	(0.342)	(0.292)
5-11	0.417	0.361	0.422	0.363
	(0.741)	(0.697)	(0.742)	(0.699)
12-15	0.225	0.226	0.224	0.225
	(0.504)	(0.507)	(0.505)	(0.505)
16-18	0.047	0.047	0.046	0.046
	(0.221)	(0.220)	(0.220)	(0.219)
Region of residence				
Greater London [=base]	0.085	0.085	0.085	0.088
Rest of South East	0.210	0.206	0.205	0.202
South West	0.090	0.087	0.093	0.087
East Anglia	0.045	0.040	0.043	0.039
East Midlands	0.096	0.093	0.096	0.093
West Midlands conurbation	0.033	0.032	0.033	0.032
Rest of West Midlands	0.064	0.066	0.062	0.063
Greater Manchester	0.031	0.035	0.032	0.037
Merseyside	0.019	0.017	0.019	0.017
Rest of North West	0.048	0.053	0.050	0.053
South Yorkshire	0.030	0.031	0.029	0.031
West Yorkshire	0.031	0.033	0.032	0.033
Rest of Yorkshire	0.036	0.037	0.036	0.034
Tyne & Wear	0.018	0.020	0.018	0.020
Rest of the North	0.046	0.043	0.045	0.044
Wales	0.042	0.041	0.041	0.040
Scotland	0.076	0.081	0.081	0.087
Occupation				
Managerial	0.218	0.093	0.216	0.095
Professional	0.131	0.116	0.127	0.115
Technical	0.103	0.127	0.101	0.125
Clerical	0.062	0.296	0.061	0.291
Craft	0.205	0.024	0.212	0.025
Protection/personal	0.052	0.135	0.054	0.142
Sales	0.043	0.097	0.041	0.093
Plant/machine operatives	0.135	0.035	0.136	0.034
Other unskilled [=base]	0.051	0.077	0.052	0.080
Trade union covered	0.463	0.497	0.460	0.497
Sector				
Civil service	0.040	0.043	0.040	0.041
Local government	0.086	0.177	0.085	0.184
Other public administration	0.048	0.113	0.047	0.114
Non-profit	0.027	0.052	0.028	0.053
Private [=base]	0.799	0.615	0.800	0.608
Industry				
Primary [=base]	0.028	0.011	0.027	0.010
Energy & water	0.031	0.006	0.030	0.006
Extraction & manufacturing	0.046	0.019	0.047	0.019
Metal goods	0.144	0.042	0.139	0.041
Other manufacturing	0.098	0.063	0.104	0.063
Construction	0.097	0.010	0.098	0.009

Distribution, hotel & catering	0.134	0.214	0.134	0.212
Transport and communication	0.082	0.025	0.085	0.027
Banking, finance & insurance	0.142	0.141	0.138	0.138
Other services	0.198	0.469	0.198	0.475
Number of person-wave observations	12,357	10,646	13,687	11,816
Number of individuals	2,151	2,046	2,382	2,266
Years in sample	5.7	5.2	5.7	5.2

Note: Figures are means (standard deviations) computed on the number of person-wave observations. ‘Sample 1’ is the sample of individuals who have information on their fathers’ (fathers-in-law’s) GH scores. ‘Sample 2’ is the sample of individuals who have information on their parents’ (parents-in-law’s) GH scores. Parents’ GH score is the average of mother and father’s GH scores for those who have both parents’ information, and father’s or mother’s GH score for those who have only one parent’s information.

Table 3
The impact of fathers' (or parents') GH score on children's GH score and fathers-in-law's (or parents-in-law's) GH score – SUR estimates by sex (Standard errors in parentheses)

	Men		Women	
Estimating sample	[1]	[2]	[1]	[2]
Sample 1 (Father's GH score)[a]				
π_1	0.255	0.192	0.247	0.167
	(0.009)	(0.009)	(0.010)	(0.010)
λ_1	0.198	0.150	0.210	0.162
	(0.009)	(0.009)	(0.009)	(0.009)
Corr(ω_{1t},ω_{2t})	0.116	0.087	0.119	0.045
	[0.0000]	[0.0000]	[0.0000]	[0.0000]
R^2	0.063, 0.040	0.203, 0.111	0.051, 0.044	0.238, 0.135
Number of individuals	2,151		2,046	
Number of person-wave observations	12,357		10,646	
Sample 2 (Parents' GH score)[b]				
π_1	0.274	0.208	0.278	0.234
	(0.009)	(0.009)	(0.010)	(0.010)
λ_1	0.215	0.168	0.225	0.179
	(0.008)	(0.009)	(0.009)	(0.009)
Corr(ω_{1t},ω_{2t})	0.124	0.099	0.142	0.093
	[0.0000]	[0.0000]	[0.0000]	[0.0000]
R^2	0.065, 0.046	0.205, 0.106	0.057, 0.052	0.189, 0.139
Number of individuals	2,382		2,266	
Number of person-wave observations	13,687		11,816	

Note: The figure in square brackets is the *p*-value of the Breusch-Pagan test of independence for the two equations. The numbers in the R^2 row refer to the R^2 statistics obtained from the children's equation and the fathers-in-law's equation, respectively. Both equations under specification [1] include father's GH score as the only regressor. For women, the regressions in specification [1] are selectivity corrected with a cubic polynomial of the index function from the selection equation (see below). Under specification [2], the children's equation includes also a quartic polynomial in age, quadratic polynomials in part-time experience and full-time experience, ratio of local unemployment stock to local vacancies stock, and dummy variables for trade union coverage, region of residence (16), industry (9) and sector (4); the fathers-in-law's equation includes quartic polynomial in own age and partner's age, quadratic polynomials in own part-time and full-time experience and partner's part-time and full-time experience, ratio of local unemployment stock to local vacancies stock, and dummy variables for own trade union coverage and partner's trade union coverage, region of residence (16), own and partner's industry (18), and own and partner's sector (8). The women's regressions are selectivity corrected, using the semiparametric two-step procedure illustrated in Vella (1998). The labour force participation selection equation (probit regression) contains: age, number of dependent children by age groups (five age groups: 0-2, 3-4, 5-11, 12-15, 16-18), and dummy variables for race (4), marital status, year of first partnership (3), house tenure (3), region of residence (16), and highest educational achievement (5).

[a] The terms π_1 and λ_1 are the estimated coefficients of father's GH score in the equation for the child's GH score and in the equation for the father-in-law's GH score, respectively.

[b] Parents' GH score is the average of father's and mother's GH score when both scores are reported and the father's or mother's GH score when only one of the two is reported. The terms π_1 and λ_1 are the estimated coefficients of parents' GH score in the equation for the child's GH score and in the equation for the parents-in-law's GH score.

Table 4
Measurement-error adjusted estimates of π_1

Gender, samples, and values of s_1	Values of s_2						
	0	0.25	0.5	0.75	1	1.25	1.5
Men – Sample 1							
$s_1 = 0$		0.204	0.240	0.300	0.384	0.492	0.624
$s_1 = 1/3$	0.213	0.225	0.261	0.321	0.405	0.513	0.645
$s_1 = 2/3$	0.277	0.289	0.325	0.385	0.469	0.577	0.709
Men – Sample 2							
$s_1 = 0$		0.221	0.260	0.325	0.416	0.533	0.676
$s_1 = 1/3$	0.231	0.244	0.283	0.348	0.439	0.556	0.699
$s_1 = 2/3$	0.300	0.313	0.352	0.417	0.508	0.620	0.768
Women – Sample 1							
$s_1 = 0$		0.177	0.209	0.261	0.334	0.428	0.543
$s_1 = 1/3$	0.186	0.196	0.227	0.279	0.352	0.446	0.561
$s_1 = 2/3$	0.241	0.252	0.283	0.335	0.408	0.502	0.610
Women – Sample 2							
$s_1 = 0$		0.249	0.293	0.366	0.468	0.600	0.761
$s_1 = 1/3$	0.260	0.275	0.318	0.392	0.494	0.626	0.787
$s_1 = 2/3$	0.338	0.353	0.396	0.470	0.572	0.704	0.865

Note: Original estimates of π_1 come from Table 3, specification [2]. The terms s_1 and s_2 denote the error due to transitory fluctuations in y_{t-1} and measurement error (expressed as a proportion or multiple of the standard deviation of the true y_{t-1}).

Table 5a

The impact of fathers' GH score on children's and fathers-in-law's GH scores by quartile of the father's GH score distribution (Sample 1) – SUR estimates by sex (Standard errors in parentheses)

	Quartile of father's GH score			
	Bottom	Second	Third	Top
Men				
π_1	0.018	-0.030	0.498	0.218
	(0.030)	(0.063)	(0.083)	(0.062)
λ_1	0.021	0.054	0.795	-0.005
	(0.028)	(0.062)	(0.093)	(0.066)
Corr(ω_{1t},ω_{2t})	0.071	0.083	0.064	0.119
	[0.0001]	[0.0000]	[0.0008]	[0.0000]
R^2	0.159, 0.166	0.181, 0.154	0.195, 0.168	0.215, 0.120
Number of individuals	552	622	459	518
Number of person-wave observations	3,176	3,487	2,733	2,961
Women				
π_1	-0.038	-0.106	0.211	0.359
	(0.032)	(0.119)	(0.075)	(0.083)
λ_1	0.120	-0.076	0.369	0.251
	(0.028)	(0.110)	(0.072)	(0.085)
Corr(ω_{1t},ω_{2t})	0.068	0.044	0.045	0.079
	[0.0003]	[0.0272]	[0.0170]	[0.0001]
R^2	0.161, 0.183	0.194, 0.187	0.173, 0.175	0.226, 0.178
Number of individuals	581	454	548	463
Number of person-wave observations	2,873	2,456	2,797	2,520

Note: Each regression includes all the variables of specification [2] listed in the note of Table 3. See that note for details.

Table 5b

The impact of parents' GH score on children's and parents-in-law's GH scores by quartile of the parents' GH score distribution (Sample 2) – SUR estimates by sex (Standard errors in parentheses)

	Quartile of parents' GH score			
	Bottom	Second	Third	Top
Men				
π_1	0.102	0.367	0.320	0.250
	(0.032)	(0.109)	(0.067)	(0.049)
λ_1	0.064	0.752	0.136	0.352
	(0.029)	(0.105)	(0.071)	(0.050)
Corr(ω_{1t},ω_{2t})	0.086	0.100	0.087	0.092
	[0.0000]	[0.0000]	[0.0000]	[0.0000]
R^2	0.141, 0.114	0.176, 0.151	0.214, 0.141	0.206, 0.158
Number of individuals	607	593	600	582
Number of person-wave observations	3,446	3,405	3,458	3,378
Women				
π_1	0.021	0.471	0.415	0.354
	(0.036)	(0.114)	(0.089)	(0.056)
λ_1	-0.070	0.273	0.615	0.215
	(0.030)	(0.097)	(0.078)	(0.054)
Corr(ω_{1t},ω_{2t})	0.071	0.105	0.108	0.031
	[0.0001]	[0.0000]	[0.0000]	[0.0891]
R^2	0.172, 0.164	0.155, 0.159	0.212, 0.171	0.189, 0.215
Number of individuals	581	561	569	555
Number of person-wave observations	3,010	2,902	2,953	2,951

Note: Each regression includes all the variables of specification [2] listed in the note of Table 3. See that note for details.

Table 6
Estimates by quartile of the child's GH score distribution – SUR estimates by sex and sample
(Standard errors in parentheses)

	Quartile of child's GH score			
	Bottom	Second	Third	Top
Men				
Sample 1				
π_1	0.015	-0.016	0.010	0.006
	(0.012)	(0.010)	(0.004)	(0.003)
λ_1	0.066	0.096	0.147	0.163
	(0.018)	(0.018)	(0.019)	(0.017)
Sample 2				
π_1	0.017	-0.018	0.009	0.006
	(0.014)	(0.015)	(0.004)	(0.002)
λ_1	0.094	0.142	0.148	0.184
	(0.018)	(0.018)	(0.019)	(0.017)
Women				
Sample 1				
π_1	0.004	0.024	0.021	0.056
	(0.002)	(0.004)	(0.005)	(0.012)
λ_1	0.104	0.133	0.155	0.191
	(0.017)	(0.019)	(0.018)	(0.024)
Sample 2				
π_1	0.005	0.023	0.021	0.080
	(0.002)	(0.004)	(0.005)	(0.013)
λ_1	0.151	0.146	0.183	0.189
	(0.017)	(0.018)	(0.018)	(0.024)

Note: Each regression includes all the variables of specification [2] listed in the note of Table 3. See that note for details.

Table 7a
Estimates by birth cohort – Sample 1 (Standard errors in parentheses)

	Quartile of parents' GH score				
	Bottom	Second	Third	Top	Mean
Men					
1935-1949					
π_1	0.011	0.004	0.773	0.049	0.229
	(0.047)	(0.107)	(0.141)	(0.130)	(0.015)
λ_1	0.092	0.020	0.762	0.025	0.174
	(0.041)	(0.096)	(0.163)	(0.139)	(0.015)
1950-1959					
π_1	0.130	0.104	0.377	0.472	0.183
	(0.073)	(0.119)	(0.159)	(0.096)	(0.017)
λ_1	-0.077	-0.284	0.603	-0.162	0.114
	(0.064)	(0.117)	(0.161)	(0.103)	(0.017)
1960-1979					
π_1	-0.096	0.199	0.379	0.163	0.155
	(0.048)	(0.116)	(0.146)	(0.099)	(0.013)
λ_1	-0.165	0.373	0.590	0.189	0.119
	(0.047)	(0.122)	(0.154)	(0.104)	(0.014)
Women					
1935-1949					
π_1	-0.068	-0.427	0.233	0.319	0.197
	(0.053)	(0.227)	(0.129)	(0.155)	(0.018)
λ_1	0.154	-0.219	-0.005	0.252	0.229
	(0.044)	(0.214)	(0.117)	(0.144)	(0.016)
1950-1959					
π_1	-0.105	0.419	0.361	-0.267	0.215
	(0.064)	(0.212)	(0.136)	(0.156)	(0.019)
λ_1	0.125	0.573	0.914	0.180	0.108
	(0.050)	(0.179)	(0.129)	(0.139)	(0.016)
1960-1979					
π_1	-0.072	0.098	0.223	0.622	0.175
	(0.056)	(0.207)	(0.123)	(0.133)	(0.016)
λ_1	0.062	-0.398	0.332	0.426	0.137
	(0.055)	(0.211)	(0.121)	(0.150)	(0.016)

Note: Each regression includes all the variables of specification [2] listed in the note of Table 3. See that note for details.

Table 7b
Estimates by birth cohort – Sample 2 (Standard errors in parentheses)

	Quartile of parents' GH score				
	Bottom	Second	Third	Top	Mean
Men					
1935-1949					
π_1	0.128	-0.172	0.293	0.405	0.221
	(0.046)	(0.202)	(0.123)	(0.097)	(0.014)
λ_1	0.159	0.722	0.348	0.442	0.173
	(0.039)	(0.185)	(0.137)	(0.096)	(0.014)
1950-1959					
π_1	0.120	0.580	-0.009	0.255	0.210
	(0.071)	(0.201)	(0.117)	(0.081)	(0.017)
λ_1	0.146	0.568	-0.066	-0.195	0.178
	(0.066)	(0.174)	(0.127)	(0.087)	(0.017)
1960-1979					
π_1	-0.038	0.214	0.697	0.145	0.175
	(0.053)	(0.183)	(0.113)	(0.085)	(0.014)
λ_1	-0.261	0.925	0.265	0.594	0.149
	(0.051)	(0.182)	(0.117)	(0.078)	(0.015)
Women					
1935-1949					
π_1	0.108	-0.289	0.517	0.726	0.238
	(0.062)	(0.231)	(0.176)	(0.104)	(0.018)
λ_1	0.112	0.855	0.979	0.307	0.238
	(0.050)	(0.197)	(0.150)	(0.096)	(0.016)
1950-1959					
π_1	-0.277	0.927	0.859	-0.063	0.221
	(0.072)	(0.220)	(0.160)	(0.093)	(0.019)
λ_1	-0.251	0.459	0.668	0.040	0.117
	(0.055)	(0.174)	(0.142)	(0.094)	(0.016)
1960-1979					
π_1	0.068	0.281	-0.148	0.497	0.232
	(0.056)	(0.165)	(0.137)	(0.092)	(0.015)
λ_1	-0.088	0.133	0.278	0.252	0.163
	(0.050)	(0.159)	(0.119)	(0.089)	(0.014)

Note: Each regression includes all the variables of specification [2] listed in the note of Table 3. See that note for details.

Table 8
Correlation between GH score at entry in the labour market and GH score at different points of the life cycle (ages) by quartile of parental GH score – OLS estimates (Standard errors in parentheses)

	Men			Women		
	36-45	46-55	56-65	36-45	46-55	56-65
Quartile of father's GH score						
Bottom	0.388	0.496	0.399	0.320	0.324	0.378
	(0.063)	(0.062)	(0.084)	(0.064)	(0.066)	(0.089)
Second	0.311	0.402	0.499	0.374	0.352	0.386
	(0.062)	(0.066)	(0.093)	(0.058)	(0.067)	(0.104)
Third	0.404	0.278	0.289	0.304	0.215	0.291
	(0.050)	(0.065)	(0.095)	(0.054)	(0.070)	(0.114)
Top	0.309	0.315	0.367	0.253	0.245	0.247
	(0.061)	(0.088)	(0.095)	(0.055)	(0.069)	(0.122)
Quartile of parents' GH score						
Bottom	0.441	0.392	0.271	0.326	0.398	0.538
	(0.051)	(0.055)	(0.076)	(0.058)	(0.054)	(0.071)
Second	0.335	0.510	0.518	0.306	0.302	0.355
	(0.060)	(0.070)	(0.096)	(0.053)	(0.066)	(0.093)
Third	0.331	0.356	0.433	0.350	0.285	0.187
	(0.055)	(0.063)	(0.091)	(0.056)	(0.071)	(0.113)
Top	0.363	0.250	0.322	0.210	0.184	0.170
	(0.051)	(0.074)	(0.080)	(0.051)	(0.065)	(0.129)

Note: Dependent variable is the individual GH score averaged over the ages 36-45, 46-55 and 56-65. The figures in the table are the coefficients on the GH score at entry into the labour market. Regressions include also five dummy variables for education.

Table 9
Standard deviation in three measures of social capital and standard deviation in years of education by gender and quartile of parental GH score
(Bootstrap standard errors in parentheses)

	Men				Women			
	Membership	Activity	Best friend's GH score	Years of education	Membership	Activity	Best friend's GH score	Years of education
Quartile of father's GH score								
Bottom	1.022	0.885	13.841	2.466	0.940	0.851	14.354	2.246
	(0.012)	(0.014)	(0.174)	(0.014)	(0.013)	(0.015)	(0.184)	(0.015)
Second	1.099	0.914	14.280	2.708	1.011	0.933	15.026	2.478
	(0.015)	(0.014)	(0.179)	(0.013)	(0.012)	(0.013)	(0.163)	(0.014)
Third	1.089	0.918	14.044	2.582	1.094	1.005	14.733	2.474
	(0.014)	(0.014)	(0.160)	(0.012)	(0.015)	(0.017)	(0.162)	(0.013)
Top	1.239	1.005	14.446	2.540	1.312	1.148	15.074	2.621
	(0.016)	(0.017)	(0.159)	(0.015)	(0.016)	(0.015)	(0.143)	(0.010)
Quartile of parents' GH score								
Bottom	1.012	0.869	13.481	2.431	0.933	0.833	14.177	2.252
	(0.012)	(0.014)	(0.185)	(0.014)	(0.013)	(0.014)	(0.179)	(0.017)
Second	1.079	0.899	14.308	2.624	0.997	0.923	14.938	2.443
	(0.014)	(0.014)	(0.169)	(0.012)	(0.013)	(0.013)	(0.166)	(0.013)
Third	1.095	0.931	14.342	2.632	1.105	0.997	14.561	2.492
	(0.013)	(0.013)	(0.148)	(0.011)	(0.014)	(0.015)	(0.155)	(0.012)
Top	1.248	1.060	14.443	2.610	1.300	1.152	14.968	2.667
	(0.016)	(0.016)	(0.152)	(0.014)	(0.016)	(0.016)	(0.139)	(0.010)

Note: Bootstrap standard errors are obtained with 1,000 bootstrap replications. "Membership" refers to the number of organisations to which an individual belongs. "Activity" refers to the number of organisations in which an individual is active.

Appendix Table A1
Intergenerational elasticities of children's/fathers-in-law's GH scores to fathers'/mother's GH scores – SUR estimates by sex (Standard errors in parentheses)

	Men	Women
π_1		
Father	0.150	0.117
	(0.013)	(0.013)
Mother	0.106	0.117
	(0.011)	(0.012)
Equality test	5.21	0.00
[*p*-value]	[0.0225]	[0.9783]
λ_1		
Father	0.053	0.104
	(0.014)	(0.013)
Mother	0.062	0.067
	(0.012)	(0.011)
Equality test		3.30
[*p*-value]	[0.6606]	[0.0691]
Number of individuals	1,083	733
Number of person-wave observations	6,149	6,154

Note: 'Equality test' is the χ^2 statistic of test of equality between mother's and father's coefficients. In square brackets we report the *p*-value of such a test. Each regression includes all the variables of specification [2] listed in the note of Table 3. See that note for details.