Families, Children and Human Capital Formation

Volume 60 Issue 2 June 2014
For checked items

view abstracts
download to citation manager

Articles

Abstract
Full Text (HTML)
Full Text (PDF)
Permissions
Select this article
Torberg Falch,
Ole Henning Nyhus,
and Bjarne Strøm
Performance of Young Adults: The Importance of Different Skills
CESifo Economic Studies (2014) 60 (2): 435-462 doi:10.1093/cesifo/ifu005
Abstract
Full Text (HTML)
Full Text (PDF)
Permissions

CESifo Economic Studies, Vol. 60, 2/2014, 257–279 doi:10.1093/cesifo/ifu008
Advance Access publication 26 February 2014

The Road to Egalitaria: Sex Differences in Employment for Parents of Young Children*

Kevin Milligan

Vancouver School of Economics, University of British Columbia,
e-mail: kevin.milligan@ubc.ca

Abstract

In 1985, Gary Becker predicted employment and childcare sex gaps may ‘disappear or be greatly attenuated in the near future.’ In this article, I examine trends in the employment gap between mothers and fathers of young children over the last 40 years. I review theoretical explanations for the gap, then proceed to analyse the gap empirically in data for Canada, the USA, the UK, and Germany. Substantial closing of the gap in the 1970s and 1980s was followed by stability since then. Evidence from Canada finds childcare subsidies have a bigger impact on the gap than parental leave. (JEL codes: J13, J16, J18, J21).

Keywords: employment, maternity leave, childcare, gender

1 Introduction

In the conclusion to a seminal paper on the economics of the family, Gary Becker looks to a future where childcare responsibilities might no longer be dominated by women.

> ‘The persistence of [childcare] responsibilities in all advanced societies may only be a legacy of powerful forces from the past and may disappear or be greatly attenuated in the near future . . . a person’s sex would no longer be a good predictor of earnings and household activities. It is still too early to tell how far Western societies will move in this direction.’
>
> Becker (1985, p. S55–6)

* This article was prepared as a keynote address for the CESifo Economic Studies and UCLS Conference on Families, Children, and Human Capital Formation held 19–20 October 2012, in Munich, Germany. Parts of this article were produced at the British Columbia Interuniversity Research Data Centre using data provided by Statistics Canada. This article represents the views of the author and does not necessarily reflect the views of Statistics Canada. Parts of this article use the British Labour Force Survey, provided by the UK Data Archive. The original data creators, depositors or copyright holders, the funders of the Data Collections (if different), and the UK Data Archive bear no responsibility for their further analysis or interpretation. Parts of this article build on joint work with Michael Baker and Jonathan Gruber.

For the purposes of this article, I will refer to Becker's hypothesized future as 'Egalitaria.' In Egalitaria, childcare responsibilities are no longer sexually asymmetric and employment among those with young children does not differ systematically for mothers and fathers.[1] Importantly, Becker's theory does not suggest that childcare would be split evenly within any particular household; his model features a knife's edge result in which one parent or the other would still have a comparative advantage in home production and that parent would specialize in home production. However, if the comparative advantages were distributed independently of sex, then on average across households in Egalitaria the employment rates of mothers and fathers would be equal.

In this article I explore one of the features of Becker's Egalitaria—the patterns in employment rates for parents of young children. I begin with a review of explanations for sexual asymmetries in the work of parents; why we do not today live in Egalitaria. I then ponder how far down the road to Egalitaria we have travelled in the years since Becker's speculation, bringing evidence from Canada, the USA, the UK, and Germany. Finally, I use Canadian evidence to assess the success of some common policies that might be used to accelerate the trip to Egalitaria.

I find a substantial gap between the employment of men and women with younger children in all four countries. This gap has more than halved over the last 40 years, but remains at a level of 3–4 years of work between the time a child is born and when the child reaches age 10. The evidence suggests that while maternity leave policies have a short-run impact while the child is very young, this impact fades away very quickly. In contrast, childcare subsidies appear to have a more permanent impact on the labour supply of mothers, and consequently on the gap in work between mothers and fathers.

2 Sources of sex differences in parental labour supply

There are vast differences through time and across countries in the labour market outcomes of mothers and fathers. In this section I discuss some potential sources of these differences; barriers blocking the road to Egalitaria. I start with a basic economic story involving relative productivities in housework and market work. Next, I discuss preferences, culture, and discrimination. Finally, some insights from biology are brought to bear on the question.

1 Of course, one could easily envision many other characteristics an idealized gender-balanced world might equalize. The definition of Egalitaria here is admittedly more narrow.

3 Marginal productivities

In a basic model of the division of labour in a household, Becker (1991) features couples who make choices about work inside and outside the home, given their accumulated human capital and given prices for selling labour to the market and prices for market-purchased goods. A key prediction of the model is specialization of one spouse inside and one spouse outside the home, resulting from the comparative advantage of one member of the couple in household production. The determinant of who specializes where is the ratio of productivities in the home and outside the home. These productivities are expressed as the amount of consumption afforded by a marginal hour applied to work. In this framework, various shocks to productivities can be imagined that would lead to differences in the employment of mothers and fathers.

For example, Cutler et al. (2003) argue that changes in the technology used within households—such as refrigeration and microwaves—have had a very large impact on the time needed for food preparation. This change in productivity within the home may lead to different decisions made by mothers and fathers about specialization. As another example, Cortes and Tessada (2011) find changes in household vs. market work decisions in the USA when low-wage immigrant childcare workers are more readily available. Moreover, technology shocks in the workplace can also affect choices. Håkanson (2013) develops the theoretical implications of continued technical improvement in the flexibility of the workplace (such as email; video conferencing) on labour supply decisions, and in particular on the intensive margin of how many hours to work and whether to take ‘career’ or less intensive jobs.

All of these technical improvements have tended to diminish the gap in the costs of work for mothers and fathers. If these costs of work act as a wedge between the marginal productivity and the wage received from market work, then decreases in costs would affect marginal conditions.[2] Shocks to relative productivities for work in and out of the household could contribute to how much and how quickly countries move in the direction of Egalitaria.

4 Preferences, discrimination, and culture

Another source of sex differences in the employment of parents is tastes and attitudes about work inside and outside the home. Beyond women’s

[2] In a similar way, relative shocks to market and household productivity could affect parental employment decisions through their impact on the intrahousehold bargaining position of men and women, in the style of Chiappori (1992).

own ideas of what they might like to do, the tastes and attitudes of others may intrude on the choices of women in the form of discrimination. Of course, all of these tastes and attitudes develop in the context of a society with a given history and set of economic institutions, which raises the potential role of culture.

Fortin (2005) documents and investigates the importance of gender attitudes in explaining cross-country patterns in the work of women. Using data from the World Values Survey, Fortin finds a substantial relationship between attitudes on the role of women in the workplace and in the home, and labour market outcomes such as employment rates and the gender pay gap.

The World Values Survey also contains a question on attitudes towards child-raising and work. The survey asks whether 'A working mother can establish just as warm and secure a relationship with her children as a mother who does not work.' In Figure 1, I plot the country-wide proportion of respondents who agree or strongly agree with the statement and compare it to the proportion employed (including self-employed). I select only married women between the ages of 20 and 40; a group for whom questions of work and children are most salient. There is a clear positive relationship between positive beliefs about children and working mothers and the proportion of married females age 20–40 who work. Of course, it is possible that many things underlie this relationship but it does suggest that attitudes about raising children may play a role in the division of work.

However formed, the tenacity of these attitudes towards family and work may run deep. Fernandez and Fogli (2009) find that the work and fertility of American women whose parents were immigrants can be predicted by labour market patterns in the country of origin. This evidence suggests that attitudes may be set early in life and do not respond quickly to new experiences and environments.

There is evidence of some evolution in attitudes through time, however, across cohorts. Using the General Social Survey from the USA, some long time-series on questions about work and family may be formed. I use the pooled 1972–2010 version of the GSS and select married women between the ages of 20 and 40.[3] In Figure 2, I show the proportion agreeing, disagreeing, or approving with a series of statements. Not all the variables are available for all years, but the pattern is clear. Through time, attitudes towards women working have grown more favourable. For example, one question asked in the GSS is very similar to the question analysed above in the World Values Survey. GSS respondents are asked to agree or disagree

[3] I use release 2 of the 1972–2010 GSS, available through the National Opinion Research Center at http://www3.norc.org/gss+website/.

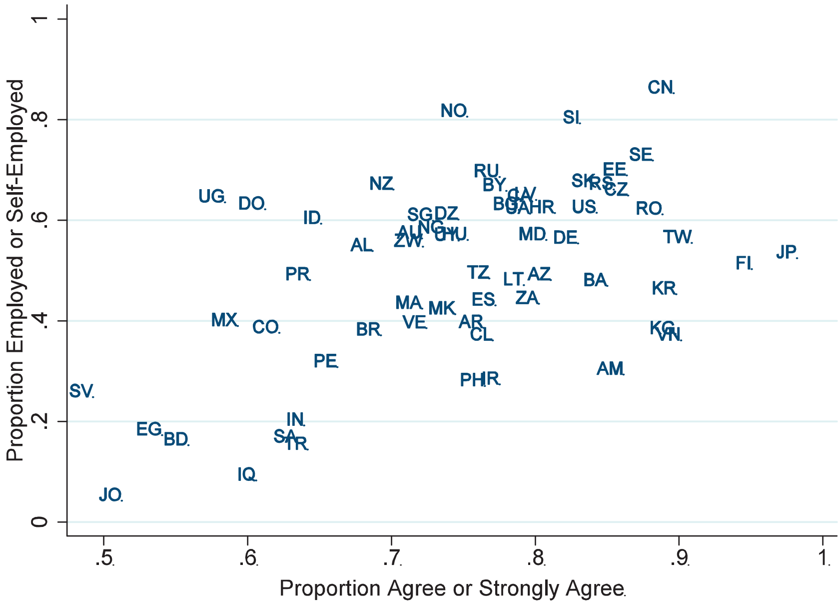


Figure 1 Can working mothers establish warm and secure relationships with their children?

Notes: Data taken from the World Values Survey, waves 3 and 4. The *x*-axis shows the proportion of respondents agreeing or strongly agreeing with the statement 'A working mother can establish just as warm and secure a relationship with her children as a mother who does not work.' The *y*-axis shows the proportion employed or self-employed. The sample includes all married women, age 20–40.

with the statement, 'A preschool child is likely to suffer if his or her mother works.' Figure 2 shows this proportion disagreeing or strongly disagreeing with that statement growing from 31.8% in 1977 to 64.9% in 2010.

How the future evolution of preferences and discrimination, and their expression through culture, continues is not an easy question to answer. Moreover, one can speculate that attitudes may reflect experience, making it very hard to tease causation out of any analysis of differences in preferences and work. However, reaching Egalitaria would likely be difficult in the absence of work and family attitudes converging towards sexual symmetry.

5 Biology

Biology contributes some part of the explanation for sexual asymmetry in work in and outside the home. Two channels through which biology

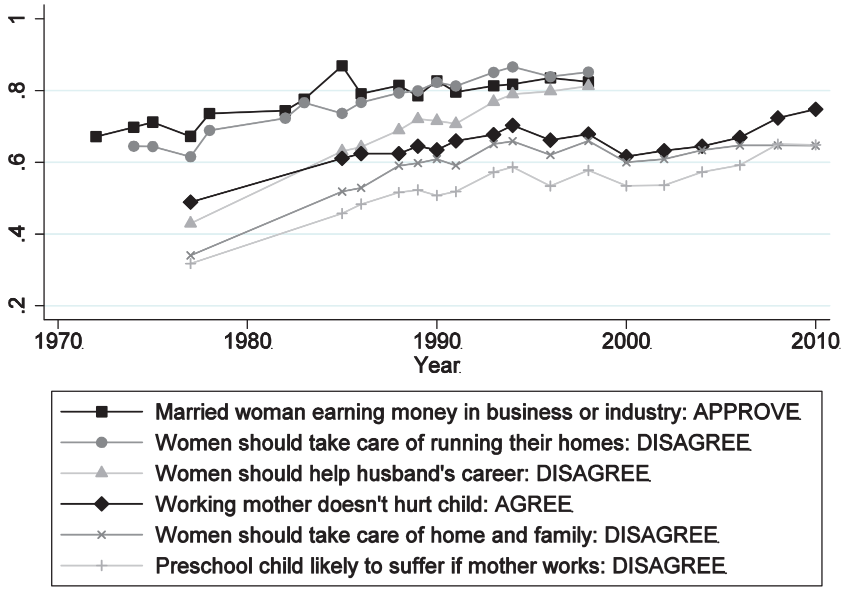


Figure 2 Attitudes about work and family in the US General Social Survey. *Notes*: Data taken from the 1972–2010 pooled General Social Survey. Graphed is the proportion of married women agreeing or disagreeing with each statement in each year.

matters are direct constraints on women's time and through evolutionary biological influences on behaviour.

Women are biologically different than men in ways directly related to the care of children. First, after fertilization the woman has sole responsibility to carry the fertilized egg through to birth. The difficulties of pregnancy increase the effort required to stay in the labour market. Second, the circumstances of birth require time both in the act of birth and in the recovery period. Again, this detracts from time available for market work. Finally, the feeding of the child through breast milk is biologically assigned to women.

Beyond the original biological assignment to women, the degree to which women devote effort to these tasks does vary. Nathoo and Ostry (2009) demonstrate this variance with the case of breastfeeding in Canada. Through the 20th century, breastfeeding moved from being very common (with over 80% initiation rates in the 1920s) to falling out of fashion (with rates below 50% in the 1950s) to again becoming very common (with initiation rates of 85% in the 1990s). These trends were driven in part by the technological innovation of formula-feeding. However, in the latter half of the 20th century, attitudes and practices around breastfeeding rebounded. In part, this may have been driven by medical evidence such as that summarized in an editorial in the journal *Pediatrics* in 1997 recommending at least 6 months of exclusive breastfeeding (American Academy of Pediatrics 1997.). In addition to medical motivations, Nathoo and Ostry argue that women today foregoing breastfeeding suffer social sanction in failing to be seen as a 'good mother', in addition to any negative health impact that may occur. In this way, Nathoo and Ostry emphasize that the degree of effort devoted to breastfeeding is more than just biological; it is also a socially determined practice. In this way, biology should not be taken as completely determinative.

The other channel where biology matters is through evolutionary influences. Trivers (1972) defines parental investment as 'any investment by the parent in an individual offspring that increases the offspring's chance of surviving (and hence reproductive success) at the cost of the parent's ability to invest in other offspring.' Women make large biological investments in ovulation, gestation, and giving birth. These biological efforts also involve large amounts of time. Men, on the other hand, are almost unlimited biologically in the number of offspring they can produce. Because women are much more limited in the number of chances they have to successfully produce an offspring, Trivers and others argued that the parental investment by mothers is higher in order to improve the likelihood that their 'rare' offspring survive.

Again, as with breastfeeding, biology should be considered a starting point and not a complete explanation of human behaviour. That is, finding a way to reach Egalitaria would likely require innovations both in technology and society that allow the existing difference in biological burdens to be minimized. However, even with small remaining biological differences, Becker's model can lead to a strong division of labour, since it relies on a knife's edge comparative advantage result. For this reason, even small persistent differences in biology present a strong barrier to reaching Egalitaria in the world of Becker's model.

6 Magnitudes of sex differences in parental labour supply

From his vantage point in 1985, Becker saw that changes were afoot in the factors influencing the sexual division of labour. The 1970s saw a great leap in the labour market participation of women in many industrial economies. Projecting those trends forward, it is not hard to see how Becker may have thought the end of the employment gap between mothers and fathers could be on the horizon. In the quarter century since Becker's observation, how much has the employment gap between mothers and fathers closed? In this section I define a particular measurement of the parental employment gap and implement it for Canada, the USA, the UK, and Germany through time.

Empirically distinguishing between the three mechanisms discussed in the previous section would be interesting, but also challenging. Relative productivities in the home and in the market present some measurement problems. Indications of market productivity can be picked up by observed wages, but the lack of market prices for home production makes home productivity difficult to measure. Preferences, on the other hand, are fairly well documented in surveys. But, preferences are difficult to distinguish from relative productivities because investments in home or workplace productivity may be influenced by attitudes and preferences in society. Finally, while biology changes only slowly through evolution, technology that relaxes biological constraints might be used to measure the impact of biology. However, this might be hard to distinguish from technological change that affects relative productivities as well.

In the empirical analysis here, I aim to compare simple employment rates across a set of countries. By calculating the gap between male and female employment rates across the ages of their children, an interesting view on the sex differences in employment emerges. I then aggregate these differences across ages to form a single summary measure of the employment gap between mothers and fathers.

7 Measurement

To construct employment rates for parents, I start by taking parents of sex $\theta \in \{\text{male}, \text{female}\}$ with children of age a. I define E_a^θ as the proportion employed for sex θ with at least one child of age a. Subtracting the female rate from the male rate gives δ_a, the difference in the employment rate for that age a:

$$\delta_a \equiv E_a^{\text{male}} - E_a^{\text{female}}.$$

I then take the δ_a terms and add them up over some values of a from 0 up to A to arrive at a measure I call the parental employment gap Δ_A.

$$\Delta_A \equiv \sum_0^A \delta_a.$$

In the mythical land of Egalitaria, the difference $\delta_a = 0\,\forall a$.

This measure has the advantage of being able to difference away any common cyclical aspect of employment rates that are common to men and women. That is, if a recession or a boom hits the labour market and affects men and women equally, it should have no impact on δ_a. Evidence on recessions, however, indicates that male employment is more volatile. Hoynes et al. (2012) find for the USA that male concentration in certain more volatile industries leads to higher employment volatility. This means that some cyclical trend in the parental employment gap may persist even after differencing.

8 Parental employment rates and gaps in four countries

I implement this parental employment gap measurement for Canada, the USA, the UK, and Germany using labour force data. These countries were chosen in part out of convenience—the calculations require precise information on family relationships and child age that are not always easily available in public-use datasets. The countries also span different policy environments and have historical differences in the employment patterns of women.

I use microdata on the labour force for all four countries. For Canada, I use the monthly Labour Force Survey, from which data are available monthly from 1976 to 2011. I form a dataset of the 'incoming rotation group' by keeping only those observations observed for the first month in their 6-month rotation in the survey. For the USA, I use the March Current Population Survey going back to 1968. In the UK, I use the Spring (April–June) waves of the quarterly Labour Force Survey and its antecedents, going back to 1975.[4] Finally, in Germany I use data from the German Socio-Economic Panel.[5]

I take a sample of all parents (married or single) and find the proportion of parents at each child age who are employed (or self-employed) and at work.[6] Because of parental leave, some people can be employed but not at work. My focus is on employed at work because I am interested in the actual time allocation with children rather than the legal status of employment. For Germany, the data do not allow employed and working to

4 The source of the UK data is: Office of Population Censuses and Surveys. Social Survey Division (various years).

5 I thank Michele Battisti for his assistance with assembling and analysing the German data. The sample sizes in the German data are much smaller than in the other countries, resulting in more variance in the graphs than observed for the other countries.

6 In Canada, I use the labour force status question for the reference week, which distinguishes between employed at work and employed but absent. For the USA, the CPS also provides a clear distinction for employment at work and absent for the reference week. In the UK Labour Force Survey, I use the question asking whether the respondent was working in the previous week. Finally, the German data do not allow me to consistently identify those who are employed and at work vs. not at work. So, for the case of Germany I analyse employment without distinguishing between those who are at work and those who are not.

be distinguished from employed and on leave. So, this means the employment rates for mothers in Germany may be overstated relative to the other countries. If a person has children of different ages, that person will appear in the age 'bins' corresponding to each of the children's ages. The sample sizes vary by country. In Canada, the USA, and the UK, the number of observations varies a bit by year, but lies mostly within the range of 15,000–20,000 per year for fathers and 20,000–25,000 per year for mothers. The German sample size is much smaller, at around 1,800 fathers per year and 2,000 mothers. There are more mothers because the number of single mothers is greater than the number of single fathers.

The employment rates E_a^{male} and E_a^{female} are graphed for Canada in Figure 3 for the years 1980, 1990, 2000, and 2010. In all years, the gap δ_a is positive at all values of a, meaning fathers work more than mothers. Across ages, E_a^{male} is quite flat, while E_a^{female} increases with a. There are two notable differences across the four time periods. First, there is a distinct 'hook' at age 0 that emerges through time and becomes quite sharp in 2010. This likely is a result of Canada's parental leave policies, as studied in Baker and Milligan (2008a,b, 2010, 2011). The provisions for job-protected maternity leave entitlements (which vary by province) moved from an average of 15.3 weeks in 1980 to 19.8 weeks in 1990; 34.9 weeks in 2000; 54.2 weeks in 2010. I will return to analysis of this policy in the next section, but the evidence here in Figure 3 is certainly suggestive. The second notable trend in Canada is an overall closing of the gap (with the exception of age 0, which moves in the opposite direction). The employment rate for mothers of 6-year olds E_6^{female} is 38.2% in 1980, 53.0% in 1990, 58.3% in 2000, and 60.8% in 2010.

The same graph appears for the USA in Figure 4. A similar flatness as seen for Canada for E_a^{male} is evident, as the employment participation of men seems invariant to the presence of young children. Women also show an upward slope of E_a^{female} with a, as in Canada. In contrast to Canada, however, there is no emerging dip at age 0—which is not surprising given the different policy development on maternity leave in the USA.[7] There is also little further closing of the gap after 1990. The employment rate for mothers of 6-year olds E_6^{female} in the USA is 44.5% in 1980, 54.4% in 1990, 60.0% in 2000, and 58.9% in 2010. At most ages, American women were working more in 1980 than Canadian women; by 2010 this had reversed.

[7] The Family and Medical Leave Act of 1993 does provide 12 weeks of job-protected but unpaid leave. However, eligibility requires a year of full-time employment at a large employer, meaning not all working mothers are eligible. Private employers often exceed this minimum in their compensation agreements with their employees. As well, several states have their own initiatives—such as New Jersey and California which have short paid-leave programs. See National Partnership for Women and Families (2012) for more detail.

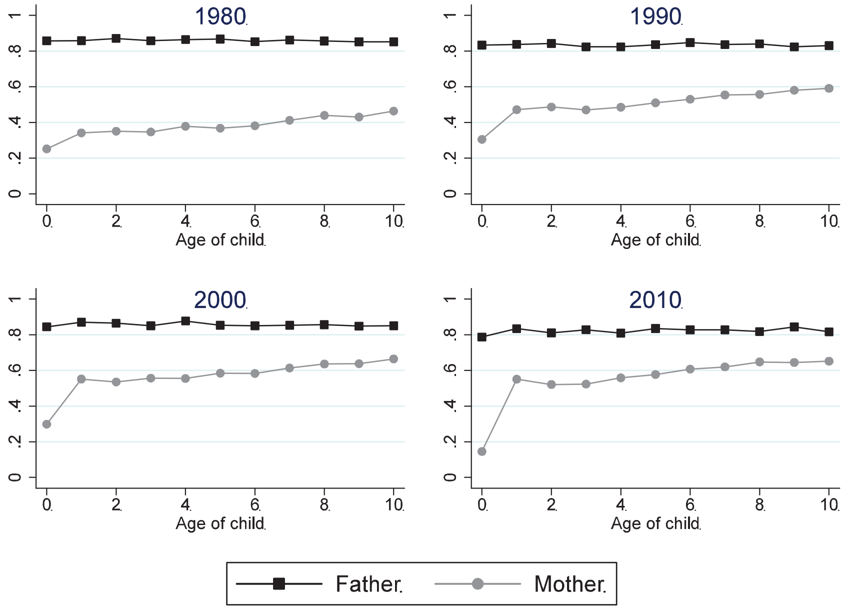


Figure 3 Employment rates of mothers and fathers, Canada.
Notes: The source is the Labour Force Survey. Parents with a child at each given age are included in the calculation for that age.

The analysis is repeated for the UK in Figure 5. Compared to Canada and the USA, the slope of the employment rate for women is much steeper in each of the years displayed.[8] The now-familiar dip at age 0 becomes more evident through time. By 2010 it is quite pronounced, more like Canada than the USA. This reflects the continued expansion of maternity leave starting in 1999 and carried out through the 2000s.

The German data in Figure 6 show several distinctions relative to the other countries. To start, sampling variability makes the lines less stable, with sample sizes in Germany about one-tenth that seen for the other countries. In the 1980s and 1990s, there was substantially less employment of women among those with older children than was the case in the other countries. As mentioned earlier, the inability to exclude those employed and on leave from this measure for Germany means that, if anything, these low employment rates are overstated compared to the other countries. The overall gap between the employment of men and women with

[8] I use 1979 instead of 1980 because there is no Labour Force Survey for 1980; it was biannual in those years.

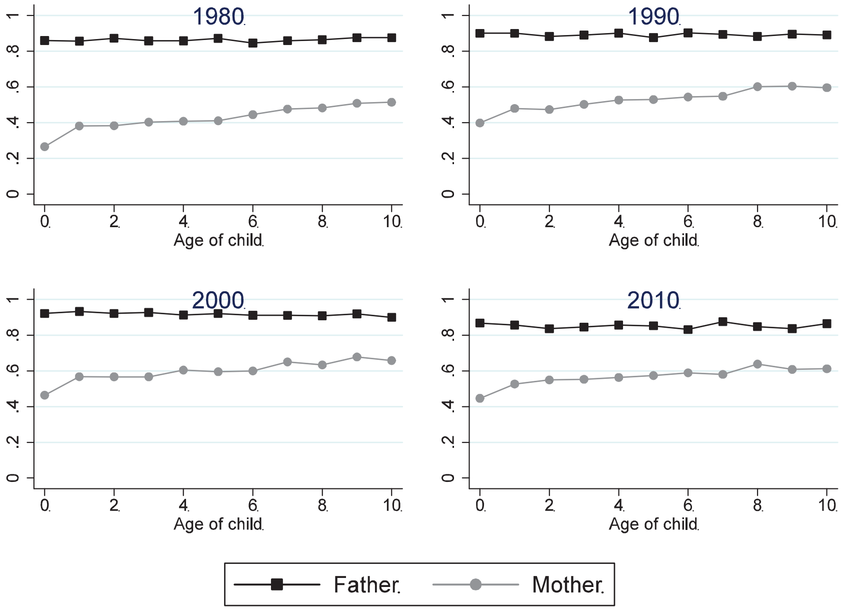


Figure 4 Employment rates of mothers and fathers, USA.
Notes: The source is the Current Population Survey. Parents with a child at each given age are included in the calculation for that age.

children has closed in Germany, but most of the change has happened among mothers of children age 5–10, with less change for those with younger children. Leitner (2010) reviews and analyses the development of policy in Germany, suggesting that in the 2000s there was a shift from a 'sequential model' of female employment towards a 'continuous employment' model. This manifested itself through large expansions of childcare in Germany in the 2000s and a reduced emphasis on parental leave.

The trends uncovered here permit a preliminary analysis, but further work could yield more insight. In particular, the changes in the employment rate gap through time likely vary across sub-groups within each country. As one example, the growth in female employment for women with young children in the USA has been stronger among high education women than for lower education women. Such analysis, while interesting, is left for future work.

9 Parental employment gap

The final calculation to assess the magnitude of the sex differences in parental employment is to add up the δ_a gaps across ages to find a

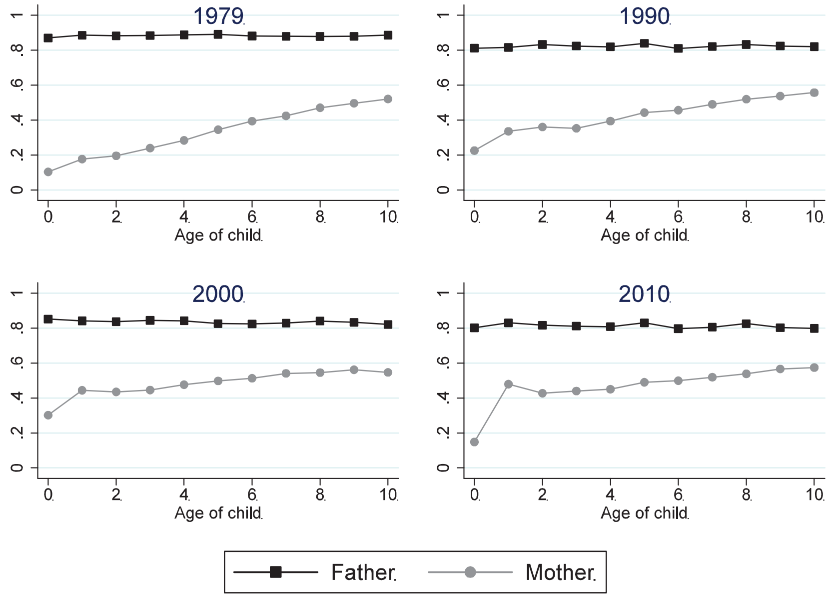


Figure 5 Employment rates of mothers and fathers, UK.
Notes: The source is the Labour Force Survey. Parents with a child at each given age are included in the calculation for that age.

summary number for the parental employment gap for each year up to age 10, Δ_{10}. The Δ_{10} for each year are graphed in Figure 7 for each of the four countries. The parental employment gap in the USA in 1968 was 6.97. This measure fell steadily until reaching 3.96 in 1987, for a drop of about 3 years of work. Over the next 20 years, however, it dropped only another 0.5. In contrast, the gap in Canada in the 1970s was higher than the USA, but crossed over in 1988. In the 2000s, the difference between Canada and the USA remained. The time path of the parental employment gap in the UK is quite similar to the USA and Canada. There was a steady downtrend until around 1990 when the gap flattened out around the level of 4 years of work. The country showing the biggest difference from the others is Germany, with the gap staying around 6 until breaking decisively downward in the 2000s. In all four countries, there appears to be some cyclical shifts in the gap relating to the financial crisis in 2009, as male employment fell more than female employment during this recent period of labour market upset. However, the cyclical trend is not strong in the earlier years.

Looking back at the prediction from Becker (1985) quoted at the beginning of this article, it is clear that from the vantage point of 1985 the gap between men and women seemed to be improving at a very steady pace.

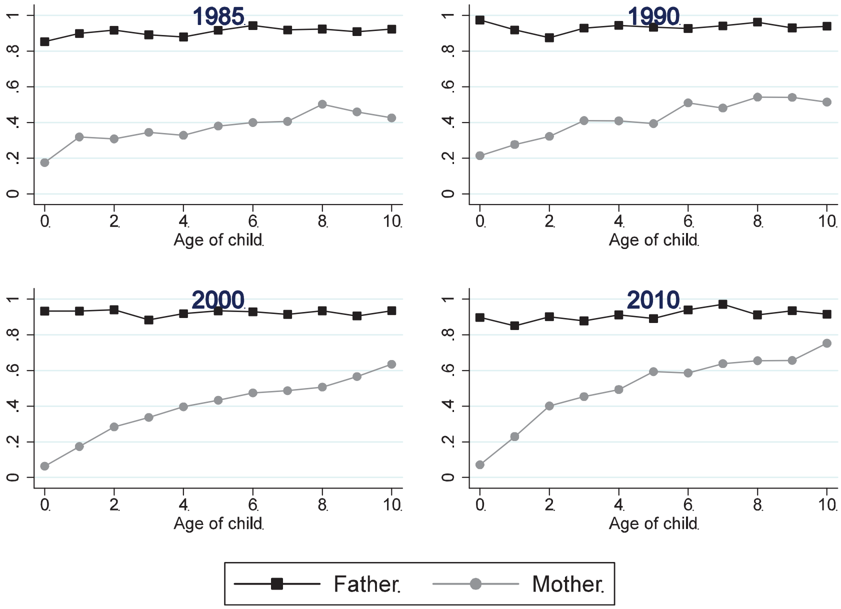


Figure 6 Employment rates of mothers and fathers, Germany.
Notes: The source is the German Socio-Economic Panel. Parents with a child at each given age are included in the calculation for that age.

Somewhat ironically, this pace flattened considerably just a few years after Becker's musings.

10 Effect of policy on parental employment gaps

If a country wished to hasten its pace along the road to Egalitaria, the earlier discussion on the sources of sex differences provides several possible levers for policy. A list of options might include:

- discouraging discrimination with pay that matches productivity;
- job-protected and/or paid maternity leave to ensure job continuity;
- lowering the price of childcare through subsidies, tax credits, or low-wage immigration; and
- attempting to change attitudes about gender roles in the household and workplace.

While I cannot evaluate each of these policy options, I do have available two policy reforms in Canada that prove useful to get a sense of how policy might influence the parental employment gap. In this section I

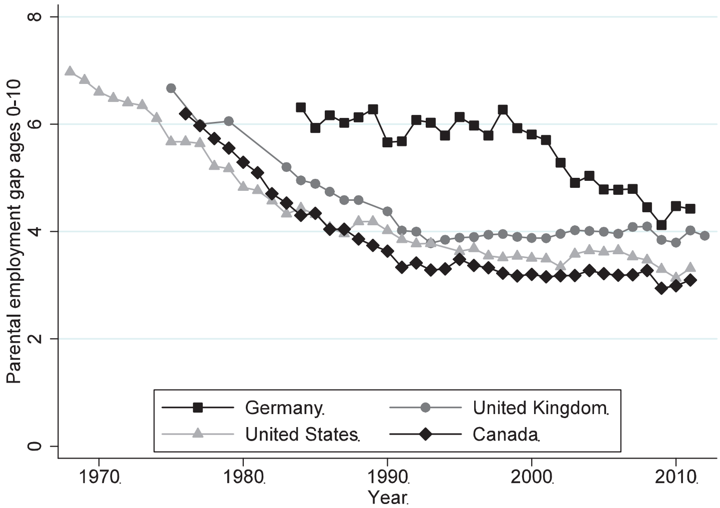


Figure 7 Parental employment gap, ages 0–10.
Notes: The data sources are the same as mentioned in Figures 3–6. The parental employment gap presented here sums the age-specific gaps seen in those figures.

estimate the impact of parental leave expansions and subsidized childcare on the parental employment gap.

The two policy changes I consider are an expansion of parental leave in 2000/2001 across Canada, and the introduction of heavily subsidized childcare in Quebec in the late 1990s. I describe each in turn. More detail on parental leave in Canada can be found in Baker and Milligan (2008a,b, 2010, 2011), while details on the Quebec childcare program are available in Baker et al. (2008). This time period in the late 1990s and early 2000s also saw the expansion of cash transfers to lower income families. Milligan and Stabile (2007) show this had a strong impact on employment for single-parent families, but was less important for two-parent families. Still, the cash transfers are a potentially confounding factor that may inhibit the causal interpretations of the parameters estimated here.

11 Parental leave reforms

Working conditions in Canada are regulated in most instances by provincial governments, with the exception of a few federally regulated industries. Job protection for maternity leaves comes under this mostly provincial jurisdiction. Leaves of 17–18 weeks were introduced in many provinces in the 1970s and 1980s. Many provinces expanded these leaves to the range of 29–35 in the early 1990s. Quebec stuck out with a jump to 52 weeks in 1991 and then to 70 weeks in 1997. In 2000 all remaining provinces jumped up to 52–54 weeks of job-protected leave.

In addition to the job protection, there is also a paid-leave element. The paid leave was organized through the federal Unemployment Insurance program, which was renamed Employment Insurance in the mid-1990s. The number of weeks of paid leave was expanded from 15 to 25 in 1990, with the extra 10 weeks also available to men for the first time. In 2001, the portion sharable between mothers and fathers moved from 10 weeks to 35 weeks, bringing total paid leave to 50 weeks.

The compensation under the Employment Insurance parental leave program is limited. The program replaces at most 55% of earnings, subject to a cap. The cap is set at a level of $45,900 in 2012, which is near the median full-time earnings level. Quebec launched its own expanded leave program in 2006 that allows for a higher replacement rate and a higher earnings cap, as well as dedicated leave for fathers.

Extended parental leave might be expected to have a negative impact in the short run on being employed and at work. More women on leave imply there are fewer at work. The leave no longer has a direct effect after the child turns age 1, as the mother would no longer qualify for leave. The enduring impact after the child turns 1-year old could go either way—women may have difficulty or less desire to re-enter the labour market after a more extended absence so they may stay out longer than the 1 year provided for by leave coverage. On the other hand, the guarantee of their job being held for them might facilitate a return to the labour market and allow women to re-enter their career where they left off.

For the purposes of the exercise conducted here, I will consider only the 2000–2001 reform, which almost simultaneously changed paid leave and job protection in most provinces to around 52 weeks. Children born on or after 31 December 2000 were eligible for the extended leave. I follow Baker and Milligan (2010) in removing Quebec from the analysis and focusing on two-parent families. As with Baker and Milligan (2010), I compare across year-of-birth cohorts born before the reform (1997 up to 2000) and those born after it (from 2001 to 2003). I implement the policy as a year-of-birth reform, assigning those born before 2001 a '0' for the policy variable POST and those born after a '1'. As I am interested in the difference between male and female employment, I include males with a 0 treatment and infer the impact of policy by an interaction between dummies for being female and treatment. I do this using the following specification:

$$\begin{aligned}\mathrm{Emp}_{ica} = \beta_0 + \mathrm{FEMALE}_{ica}\beta_1 + \mathrm{POST}_c\beta_2 + \mathrm{POST}_c * \mathrm{FEMALE}_{ica}\beta_3 \\ + X_{ica}\beta_4 + e_{ica}.\end{aligned}$$

Here, individuals are indexed by i, cohorts are indexed by c, and the current age of the child is indexed by a. Emp_{ica} is a binary variable for being employed and at work, FEMALE_{ica} is a binary variable for being female, POST_c is an indicator for being a member of a cohort born from 2001 onward, and X_{ica} is a vector of control variables for mother and father education and age. The estimation is implemented as an ordinary least squares model, which for a binary dependent variable makes it a linear probability model. Standard errors are adjusted using the robust adjustment, which accounts for the heteroskedasticity inherent in the linear probability model errors. The parameter of interest is β_3, the interaction of POST_c and FEMALE_{ica}, which picks up the differential effect of parental leave expansion on mothers.

The regression is run separately for each age from 0 to 10. If someone has multiple children in this age range, they appear in each of the corresponding regressions. The data employed are from the Labour Force Survey, as described earlier. Using the estimated coefficient β_3 for each age, a counterfactual path for female employment in 1999 can be constructed when the policy impact is added to the prevailing employment rates in the last year before the reform, 1999.

The results are presented graphically in Figure 8, which shows a line for men, another for women without the parental leave, and another for women with the predicted policy impact (and associated 95% confidence interval). The 1999 employment rates for mothers and fathers are graphed, along with the predicted impact of policy. The impact at age 0 is negative for employment, as was expected. More women taking leave means fewer employed and at work. In subsequent years, however, there is little strong impact. It is slightly negative over the first few years, then switches to slightly positive. This is consistent with the findings in Baker and Milligan (2011) where little long-run impact of leave expansion was found.

This analysis of extended parental leave suggests that, whatever its other merits during the first year of the child's life, there appears to be little lingering labour market impact. These regression results are consistent with the cross-country graphs for the UK which also show a kink developing at age 0 when maternity leave entitlement was expanded in the late 1990s and 2000s. Because of the short period in which it has impact, parental leave appears to have little effect on the parental employment gap measured over ages up to 10.

12 Subsidized childcare

In September of 1997, the province of Quebec began to implement a novel policy experiment. A new system of childcare centres featuring tightened

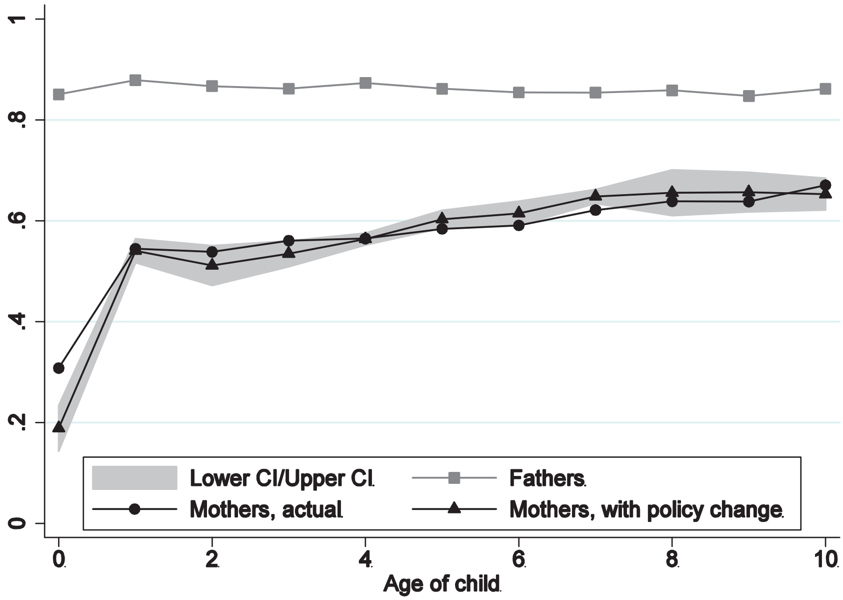


Figure 8 Impact of parental leave on employment rates, Canada excluding Quebec.

Notes: The source is the Labour Force Survey and the author's estimation. Shown are the employment rates for mothers and fathers in 2000, along with the estimated impact of the policy change for mothers. Parents with a child at each given age are included in the calculation for that age. The shaded area represents the 95% confidence interval around the estimated impact.

regulatory guidelines, a new curriculum, and heavy subsidies from the province became available for 4-year olds. Parents paid about one-seventh of the price of the care—$5 a day. The program was not means-tested in any way and proved very popular. There were previously subsidies for lower income families, so the program has its biggest impact on those with middle incomes. Over the next 3 years to 2000 the program was expanded first to 3-year olds, then 2-year olds, then finally to 0 and 1-year olds. There was no corresponding program in Canada outside Quebec.

As a first cut at the impact of the policy, a simple comparison of Quebec with Canada outside Quebec is shown in Figure 9. The parental employment gap for ages 0–10 was higher in Quebec than Canada outside Quebec until the mid-1990s. After 2000, Quebec takes a clear jump downward, coincident with the timing of the full implementation of the childcare program.

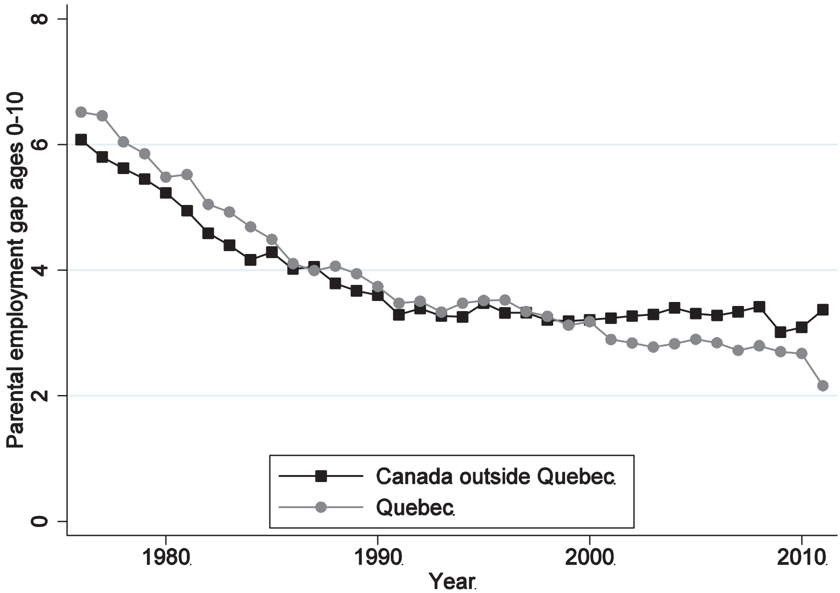


Figure 9 Parental employment gap, Quebec and Canada outside Quebec.
Notes: The source is the Labour Force Survey.

To look more carefully at the long-run impact of the Quebec childcare program on female employment, I compare birth cohorts in Quebec and Canada outside Quebec in a difference-in-differences format. Birth cohorts varied in their exposure to the program, with those born in the mid-90s and earlier having no exposure while those born from 2000 onward were eligible for all of their preschool years.

The empirical implementation uses the following equation:

$$\begin{aligned}\mathrm{Emp}_{icqa} = {} & \beta_0 + \mathrm{FEMALE}_{icqa}\beta_1 + \mathrm{QUE}_q\beta_2 + \mathrm{POST}_c\beta_3 \\ & + \mathrm{POST}_c * \mathrm{QUE}_q * \mathrm{FEMALE}_{icqa}\beta_4 + \mathrm{POST}_c * \mathrm{QUE}_q\beta_5 \\ & + \mathrm{POST}_c * \mathrm{FEMALE}_{icqa}\beta_6 + \mathrm{QUE}_q * \mathrm{FEMALE}_{icqa}\beta_7 \\ & + X_{icqa}\beta_8 + e_{icqa}.\end{aligned}$$

As with maternity leave, individuals are indexed by i, cohorts are indexed by c, and the current age of the child is indexed by a. In addition, observations are indexed for province of residence by q. The variables are the same as for maternity leave with the addition of a binary variable for Quebec residence, and its interaction with several variables. The policy parameter of interest is β_4 which picks up the differential impact of the

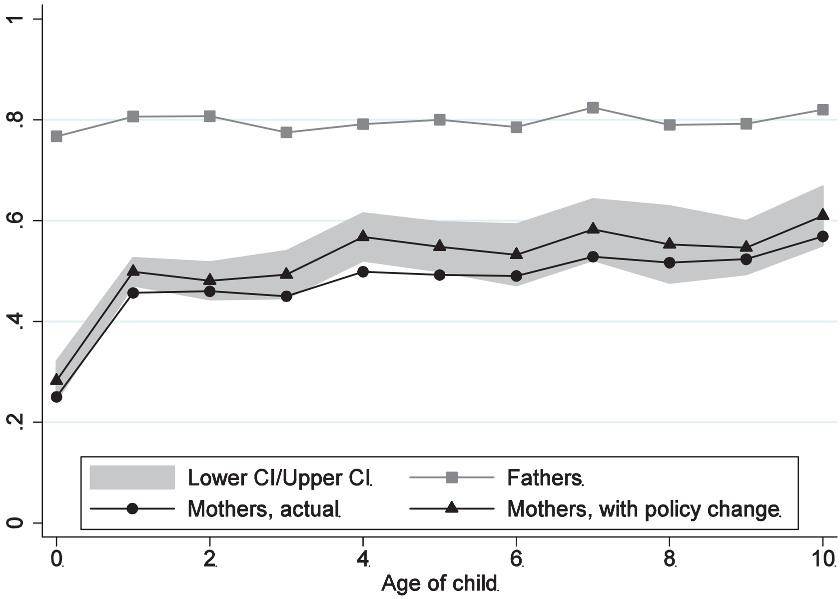


Figure 10 Impact of childcare subsidies on employment rates.
Notes: The source is the Labour Force Survey and the author's estimation. Shown are the employment rates for mothers and fathers in 1996, along with the estimated impact of the policy change for mothers. Parents with a child at each given age are included in the calculation for that age. The shaded area represents the 95% confidence interval around the estimated impact.

policy on females for affected cohorts in Quebec. I also include a full set of 'second order interactions' between Quebec residence, year-of-birth cohort, and being female. I again use the Labour Force Survey for the estimation. I exclude observations for the 1997–1999 transition years. Regressions are run separately by age of the child and the set of parameters recorded.

Figure 10 shows the results by graphing the actual 1996 employment rates for mothers and fathers and the predicted policy impact. Again, the shaded area represents the 95% confidence interval. The impact varies a bit by age, but the point estimates lie mostly in the range of 3–7% points. The statistical significance for many of the individual points is questionable, but the pattern of positive impacts is consistent across the ages. Interestingly, the impact of the policy is sustained after the children graduate out of the program at age 5 and enter school. Mothers of 10-year-old children who were exposed to the program show a 4.1 percentage point increase in being employed compared to those mothers who were not exposed.[9] When aggregated over the ages 0–10, the impact is about 0.46 years of work compared to untreated birth cohorts.

While not completely closing the gap between the employment of mothers and fathers, the Quebec childcare reform uncovers evidence that childcare subsidies may have a continued impact on maternal work even after the child graduates into regular school. This sustained and continued impact suggests that childcare subsidies may do more to close the parent employment gap than parental leave policies.

13 Conclusions

This article has sought to bring evidence to Gary Becker's future-looking speculation from 1985 about the coming of a time when sex would no longer predict employment or child-raising roles: the land of Egalitaria. After reviewing several barriers that stand in the way of reaching Egalitaria, I provided some descriptive evidence of the trends in the sex differences in employment of parents of younger children from the 1970s to 2011. In the USA, the UK, and Canada there was substantial closing of the parental employment gap until the late 1980s, when the gap flattened out. In contrast, the parental employment gap stayed relatively high in Germany until a shift in policy towards childcare in the 2000s. Finally, I provide an empirical investigation of two policies that have been put forward as beneficial to the employment circumstances of mothers. Using Canadian policy variation, I find stronger evidence for a lasting impact of childcare subsidies than for parental leave.

From this work, I draw two primary conclusions. First, Becker's intuition that the gap in the employment of mothers and fathers was closing was correct—and this gap fell by nearly half in a relatively short 20-year period during the 1970s and 1980s. Parental employment patterns are not biologically fixed; the evidence presented here shows maternal employment can swing strongly in a short time period. Second, policy by itself seems to have a modest impact. Maternity leave makes the gap larger while the leave is underway, but does not seem to be compensated by more work by mothers when the child is older. The evidence from Quebec and from Germany suggests that large childcare subsidies can have a more lasting impact, but a large parental employment gap still remains even with these programs in place.

The technological, social, and policy developments of the 1970s and 1980s that mixed together to generate decreases in the parental employment gap may not be reproducible. It is also possible that the 'easy gains'

9 This result is statistically significant with a p-value of 0.041.

have been exhausted and the parental employment gap has now settled at a level where it is tightly constrained by biology, preferences, and culture. However, it should be acknowledged that a lot has changed in the space of two short generations. It seems prudent to allow more time to pass before closing judgment on the likelihood of reaching the Egalitaria foreseen by Becker (1985).

Acknowledgements

I thank Frances Woolley for insightful comments on an early draft, and conference participants, the referees, and the editor for several helpful suggestions. I thank Anujit Chakrabarti for research assistance. Michele Battisti provided invaluable assistance in procuring and analysing the German data in this article.

References

American Academy of Pediatrics (1997), "Breastfeeding and the Use of Human Milk", *Pediatrics* **100**, 1035–9.

Baker, M. and K. Milligan (2008a), "Maternal Employment, Breastfeeding, and Health: Evidence from Maternity Leave Mandates", *Journal of Health Economics* **27**, 871–87.

Baker, M. and K. Milligan (2008b), "How Does Job-Protected Maternity Leave Affect Mothers' Employment?", *Journal of Labor Economics* **26**, 655–91.

Baker, M. and K. Milligan (2010), "Evidence from Maternity Leave Expansions of the Impact of Maternal Care on Early Child Development", *Journal of Human Resources* **45**, 1–32.

Baker, M. and K. Milligan (2011), "Maternity Leave and Children's Cognitive and Behavioral Development", NBER Working Paper No. 17105.

Baker, M., J. Gruber and K. Milligan (2008), "Universal Childcare, Maternal Labor Supply, and Family Well-Being", *Journal of Political Economy* **116**, 709–45.

Becker, G. S. (1985), "Human Capital, Effort, and the Sexual Division of Labor", *Journal of Labor Economics* **3** (**Part 2**), S33–58.

Becker, G. S. (1991), *A Treatise on the Family*, enlarged edn. Harvard University Press, Cambridge.

Chiappori, P.-A. (1992), "Collective Labor Supply and Welfare", *Journal of Political Economy* **100**, 437–67.

Cortes, P. and J. Tessada (2011), "Low-Skilled Immigration and the Labor Supply of Highly Skilled Women", *American Economic Journal: Applied Economics* **3**, 88–123.

Cutler, D. M., E. L. Glaeser and J. M. Shapiro (2003), "Why Have Americans Become More Obese?", *Journal of Economic Perspectives* **17**, 93–118.

Fernandez, R. and A. Fogli (2009), "Culture: An Empirical Investigation of Beliefs, Work, and Fertility", *American Economic Journal: Macroeconomics* **1**, 146–77.

Fortin, N. M. (2005), "Gender Role Attitudes and the Labour-Market Outcomes of Women Across OECD Countries", *Oxford Review of Economic Policy* **21**, 416–38.

Håkanson, C. (2013), *Changes in Workplaces and Careers*, Doctoral Dissertation, Department of Economics, Stockholm University. Chapter 5.

Hoynes, H., D. L. Miller and J. Schaller (2012), "Who Suffers During Recessions?", *The Journal of Economic Perspectives* **26**, 27–47.

Leitner, S. (2010), "Germany Outpaces Austria in Childcare Policy: The Historical Contingencies of the 'Conservative' Childcare Policy", *Journal of European Social Policy* **20**, 456–67.

Milligan, K. and M. Stabile (2007), "The Integration of Child Tax Credits and Welfare: Evidence from the Canadian National Child Benefit Program", *Journal of Public Economics* **91**, 305–26.

Nathoo, T. and A. Ostry (2009), *The One Best Way: Breastfeeding History, Politics, and Policy in Canada*, Wilfrid Laurier University Press, Waterloo, ON.

National Partnership for Women and Families (2012), *Expecting Better: A State-by-State Analysis of Laws that Help New Parents,* 2nd edn.

Office of Population Censuses and Surveys. Social Survey Division (various years), Labour Force Survey, (various years) [computer file]. UK Data Archive, Colchester, Essex. SN: 1756, http://dx.doi.org/10.5255/UKDA-SN-1756-1 (last accessed 20 February 2013).

Trivers, R. L. (1972), "Parental Investment and Sexual Selection", in B. Campbell, ed. *Sexual Selection and the Descent of Man 1871-1971*, Aldine Publishing Company, Chicago.

Rising Inequality and Intergenerational Mobility: The Role of Public Investments in Human Capital

November 4, 2013

One consequence of the rise in inequality witnessed over the past 40 years is its potentially negative impact on intergenerational mobility if parents at the bottom of the income distribution invest significantly less in their children's human capital. I consider whether public investments in children can potentially offset the inequality of private investments. Specifically, examining changes in public spending in 25 OECD countries over the period 2000-2009, I find that increases in spending on health are most strongly associated with reductions in the importance of family background and declines in inequality in the production of child human capital as measured by PISA test scores among 15 year-olds. Public spending on family support, housing and education are also moderately related. In contrast, increased spending on the elderly is associated with increases in the importance of parental background and inequality of child test scores. These results suggest that public investments in child human capital have the potential to offset the potentially negative impact of increasing income inequality on intergenerational mobility and inequality of the next generation. Further research firmly establishing a causal relationship is needed.

JEL Codes: I14, I18, I24, I38

Keywords: Inequality, Human Capital, Investments, Health, Education

I. Introduction

After decades of decline, income inequality began to rise in the mid-1970s, a trend that continues today. The European Commission has recently documented how earnings inequality, as measured by the 90/10 ratio, increased between 1979 and 2000 in 10 of 15 OECD countries and fell in only one (European Commission, 2010). Policy makers have voiced concern over the growth in income inequality as there exist strong associations between inequality and diminished growth, higher crime, drug use and persistent poverty (Wilkinson and Picket, 2009).

More recently, inequality has been linked to low levels of intergenerational mobility. Corak (2013) documents a strong cross-sectional correlation in income inequality and intergenerational mobility across OECD countries, referred to as the "Great Gatsby Curve". Corak (2013) includes a discussion of why greater inequality might cause a decline in intergenerational mobility. The rationale is rooted in economic models of investment in child human capital developed by Becker and Tomes (1979, 1986) and adapted by Solon (2004). These models predict that as inequality rises, so too will the difference in child human capital investments made by parents at the top and bottom of the income distribution. This would, in turn, lead to a decline in intergenerational mobility and an increase in the inequality of human capital (and therefore earnings) of the next generation. But while these models focus largely on private investments made in children, there is scope for public investments as well. Indeed, Corak (2013) continues "the reasons for the difference in the intergenerational elasticities across countries have to do with the different balances struck between the influence of families, the labor market and public policy in determining the life chances of children" (page 85), though he does not go on to examine this empirically.

In this paper I examine the potential role of public policy and specifically, public investments in human capital, in reducing both intergenerational elasticities and inequality of child human capital. The main hypothesis is that in societies in which public investments in child human capital are increasing, the relative importance of private investments in child human capital should decrease and we should see a decline in intergenerational elasticities as well as a decline in inequality of child human capital.

To examine this hypothesis, I use individual-level Program for International Student Assessment (PISA) test score data for 15 year olds in 25 OECD countries for the years 2000 and 2009 to generate three measures of intergenerational transmission and inequality of human capital. The first measure is defined as the elasticity between parental SES (parental income is not available) and a 15 year old child's test scores, calculated over all students in a country in a particular year. The second is the ratio of child test scores of those with parents in the top 25% of the SES distribution to those with parents in the bottom 25% of the SES distribution. The third measure is a measure of the inequality of child human capital and is defined as the distance between those children at the top of the cognitive test score distribution and those at the bottom (the 90:10 ratio) within a country and year.

Based on these three measures, I first document a strong positive correlation between inequality of parental SES, low levels of intergenerational mobility (as reflected in a strong elasticity between parental SES and child test scores), and inequality of child human capital (as reflected by the 90:10 ratio in child test scores). In an effort to control, at least in a crude way, for other confounding factors that might be correlated with both intergenerational elasticities and inequality, I focus on changes in these measures within a country over the past decade with a country fixed effect specification. I find that countries that experienced the biggest increases in inequality of parental SES witness the biggest increases in intergenerational elasticities and inequality of child test scores.

I follow this with an exploration of how changes in both inequality and intergenerational elasticities within a country over time might be associated with changes in spending on social programs including education, health, family support, housing and labor support programs. In considering which social programs to examine and what associations we might expect to find, I relied heavily on the existing micro-evidence on the effectiveness of different public programs in increasing child human capital and improving child well-being more generally as collected in a recent review of the literature (Aizer and Doyle, 2013).

I find that increases in spending on health are associated with the greatest reductions in the correlation between parental SES and child test scores and reductions in child test score inequality. Spending on housing, family support and education (the latter with respect to math scores) are more moderately associated with these outcomes. Not surprisingly, spending on the elderly has the opposite relationship: it is associated with increases in intergenerational correlations and increases in inequality of child human capital. This may reflect the fact that spending on the elderly likely crowds-out spending on the young.

There are two main contributions of this work. First, I provide empirical evidence to support the hypothesis that inequality affects intergenerational mobility by changing the distribution of child human capital. To do so, I showed that as parental SES becomes more unequal in a society, so too does the human capital distribution of the next generation. Second, I provide evidence that increased public spending on children has the potential to reduce any negative impact of increasing income inequality on intergenerational mobility and inequality of the next generation. These results, while based on analysis that include basic controls for potential confounding, should be viewed as largely suggestive. Future work establishing a causal relationship between public spending on child human capital and improvements in intergenerational mobility and inequality of the next generation is warranted.

II. Background

A. Human Capital Production: the Roles of Private and Public Investments

A child's human capital is determined by his or her initial endowment, private and public investments in the child, and luck. Parents affect both the initial endowment (through genetic heritance) and all private investments. Strong intergenerational correlations between parental income and children's human capital observed in micro-data provide some evidence that private investments are an important component of child human capital.[1] If there were no public investments and a child's human capital were driven exclusively by private (parental) investments, then inequality of child human capital would rise with inequality of parental income as the children of parents with more resources would witness increased investment in their human capital, while the children of the poor would not. But if public investments can substitute for private investments, then as public investments in children increase, the influence of parental SES on child education should diminish, resulting in both an increase in intergenerational mobility and a decrease in educational inequality (see Holter, 2011; Viaene and Zilcha, 2001 for more theoretical treatment).[2]

B. Empirical Work on Intergenerational Correlations and Inequality: Evidence from the Cross Section

Consistent with the above, there is a strong empirical link between intergenerational mobility and inequality across countries. Work by Corak (2013) previously mentioned, documents this, as does other work. For example, in a 2010 report, the OECD presents evidence that intergenerational social mobility is lower in more unequal European societies, based on the 2005 EU-SILC database. In fact, they document a 0.56 correlation between the Gini coefficient (the measure of inequality) and intergenerational wage persistence (as measured by the distance between the wages of men who father has tertiary schooling and the men whose father has less than secondary schooling).

The OECD report further argues that education is the key driver in intergenerational correlations in wages. The authors base their argument in part on the fact that once they control for a son's educational attainment, a father's educational attainment no longer affects the son's wages, with some exceptions.[3] Based on this, the OECD study concludes that "policies that facilitate access to education of individuals from disadvantaged family background promote intergenerational wage mobility" (OECD, 2010, page 18). The education policies they consider include: total resources, early education, tracking, vocational education, teacher salaries. They conclude that total resources for education matter less than specific policies such as tracking, early child care,

[1] This is also consistent with the importance of endowments which are inherited from parents.
[2] If public and private investments are complements, we would not expect a reduction in inequality.
[3] The exceptions include Ireland, Italy, Luxembourg, Netherlands, Spain and the UK.

and growth in teacher salaries. These conclusions are based on cross sectional differences across European countries at a single point in time.

In this paper, I go beyond the existing work by considering the role of both education policies (total spending, tracking and the share in private schools) and other social programs in increasing intergenerational mobility and reducing future inequality of human capital. To inform the analysis, I discuss the existing micro-evidence on the effectiveness of various public programs in terms of improving child well-being in the next section.

C. Micro Evidence on the Effectiveness of Public Programs

The argument that public investments in children can reduce intergenerational correlations and inequality in child human capital relies on the notion that public investments will increase the human capital of children at the bottom of the distribution who otherwise would receive fewer private (parental) investments. But what does the evidence suggest regarding the effectiveness of public programs in increasing child human capital? In a review of the literature, Aizer and Doyle (2013) examine the effects of various child welfare interventions. These include: foster care policies, family planning policies, income transfer programs, residential mobility interventions, educational interventions and public health programs. While the focus of that review chapter was to highlight the role that economic models and econometric techniques play in the evaluation of public programs, there were a number of relevant findings, which I summarize here.

<u>Foster Care</u>: Perhaps the most extreme type of public investment is the government's removal of a parent's custody rights when a child is found to be maltreated by the parents. Research that has sought to establish whether removal improves child outcomes have relied on propensity score (Berger et al, 2009) or instrumental variable (Doyle, 2007) techniques to identify the effect of out of home placement. In general, the research has found that removal has little positive effect and in some cases a negative effect on child well-being as measured by teen motherhood, juvenile delinquency, and ultimately, adult incarceration and employment and earnings.

<u>Family Planning policies</u>: There is evidence that access to family planning services reduces fertility (Bailey, 2011). What is less clear, however, is the causal impact of fertility on child quality. While the theory supports a clear negative relationship between family size and child human capital (Becker and Lewis, 1973), the empirical evidence is less consistent. There is clear evidence that lower fertility improves child well-being through a change in the composition of children born (eg, Gruber, Levine and Staiger, 1999). Less clear is whether reduced fertility improves the living circumstances among those born (eg, Black, Devereaux and Salvanes, 2010 and Angrist, Lavy and Shlosser, 2005).

<u>Anti-poverty programs</u>: There is in general, little evidence that income benefits children (Mayer, 1997 and Blau, 1999). Indeed, in the cross-section, participation in anti-poverty or welfare programs in the US is associated with worse child outcomes. There is some work showing that

once one controls for maternal characteristics (Levine and Zimmerman, 2000) or instruments for welfare receipt (Currie and Cole, 1993), welfare is no longer negatively associated with well-being. Almond, Hoynes and Schanzenback (2011) estimate the impact of the roll out of the food stamp program in the 1960s, which is essentially a means-tested cash transfer, and find that it improves birth outcomes.

There is greater evidence of positive effects of income transfers that operate outside traditional welfare programs. These include the EITC in the US and the child tax benefits in Canada. Dahl and Lochner study the impact of the EITC on child cognitive test scores in the US. They find that an additional $1000 in cash from the EITC raised test scores by 2 to 4 percent of a standard deviation. Milligan and Stabile (2011) show that the child tax benefit in Canada has substantial benefits in terms of child health, but less so in terms of improvements in child test scores.

Housing and neighborhood effects: There have been a number of studies based on strong identification strategies estimating the impact of housing mobility on child well-being. Those based on natural experiments include Oreopolous (2003) and Jacob (2004). Oreopolous studies placement in public housing in Canada and exploits a long wait list and the availability of different types of housing in different neighborhoods when one has finally proceeded to the top of the list. He finds that the type of housing and income of the neighborhood matters very little in terms of future earnings, income or welfare receipt. Likewise, Jacob (2004) exploits the destruction of high density public housing in Chicago and removal of families to lower density public housing in higher income neighborhoods on child outcomes and likewise finds no effects. Perhaps the best known housing mobility evaluation is the MTO experiment which randomized participants to one of two treatment groups or a control group. The treatment groups received either a housing voucher to move from public housing to private housing or a voucher to move with a stipulation that they move to a low poverty neighborhood. The children were followed up 10-15 years after the initial random assignment. The researchers concluded that housing mobility had small effects on traditional measures of child well being (health, education, juvenile delinquency) but some larger positive effects on mental health and happiness.

Educational interventions: Public education differs from most other public programs in two main ways: first, it's compulsory in all developed countries until the child reaches a certain age, typically 14-18 years, and second, eligibility is not means-tested. Moreover, depending on the country, there can be considerable heterogeneity in how the schools are administered and funded and the extent to which they rely on different inputs in the production of education.

Evidence regarding the causal impact of school funding on test scores is lacking. Most education researchers agree that more important than the amount of funding is the way the funds are used to purchase the two main inputs in an education production function: class size and teacher quality. There has been a substantial amount of work attempting to evaluate the effectiveness of these inputs in raising test scores. With respect to class size, there is conflicting evidence that smaller class sizes raise test scores. Studies including Rivkin, Hanushek and Kain (2005), Angrist and Lavy (1999), and Hoxby (2000) find small or no effects of reducing class size on test scores. Frederrriksson, Ockert and Oosterbeek (2013) using Swedish data find that reducing class size by seven students among 10-13 year olds improves cognitive achievement at age 16 by 15 of a standard deviation, educational attainment by one third of year and increases earnings by 4.2 percent in adulthood.

Teacher quality is more difficult to measure. Using traditional measures of teacher quality (educational attainment, credentialing, experience) there appears to be no effect of quality on test scores (Kane, Rockoff and Staiger, 2006 and Hanushek and Rivkin 2004 and 2006). However, as Hanushek (1992) and Hanushek and Rivkin (2010) show, there appears to be large variation in teacher quality if one simply looks at gains in test scores associated with individual teachers: with some teachers producing test score gains of only half a year of achievement, and others producing three times as much. Hanushek (2011) estimates that having a teacher one standard deviation above the mean in terms of teacher effectiveness would generate marginal gains of nearly half a million dollars in the future earnings of the student. Dobbie and Fryer (2011) present evidence that attending a "high quality" charter school has a positive impact on math, and English test scores. However, the particular intervention studied (Promises Academy in the Harlem Children's Zone) includes better teachers, longer school years, tutoring, early enrollment and wrap-around services, making it difficult to know which of these improvements in quality is responsible for the improved test scores.

Other factors that can influence the production of education include teacher and student incentives and peer effects. The strongest evidence in favor of teacher incentives come from randomized experiments conducted in developing countries which suggest that teacher incentives can significantly improve student test scores (Glewwe, Ilias and Kremer, 2010). One study based on US school children has shown that providing young students with financial incentives has some moderate effects, but only if structured in such a way as to reward not test performance, but performance on more "intermediate" products such as practicing reading (Fryer, 2010).

More recently, researchers have focused on understanding the role of early education. Cunha and Heckman (2007) develop a model that explains why early investments are more productive than later investments in producing human capital. There is growing empirical evidence consistent with this. For example, the Tennessee STAR experiment that randomized kindergarten students to classrooms that varied in terms of class size, teacher and peer quality showed large short run and long run gains associated with higher quality Kindergarten classes (Chetty et al, 2011). The Perry Preschool program which randomized families to high quality preschool has also been shown to have large effects. The Head Start program, a public preschool program for families below poverty, has been shown to have short term effects on cognitive achievement (Curre and Thomas, 1995 and 1999; Garces, Currie and Tomas, 2002) that seem to fade by grade 3, but more recently, researchers have found large positive long term benefits (Ludwig and Miller, 2007 and Deming, 2009.) While there is little evidence that subsidized, universal child care improves child outcomes, this could potentially reflect the fact that effects might be found in some parts of the distribution of children, but not others, as Havnes and Mogstad (2010) have found.[4]

Public Health interventions. Children's poor economic circumstances are strongly predictive of worse health, and more so as the child ages (Case, Lubotsky and Paxson, 2002). Child health is also strongly related to adult health and productivity (Currie, 2009). Together, these two facts suggest that improvements in child health among poor families could lower future inequality of earnings. There are two main ways in which public policy can affect child health. The first is through the provision of public health insurance. European countries largely rely on publicly provided health insurance for the majority of the population. In contrast, the US is one of the only developed countries that relies primarily on private insurance, but provides public insurance to poor families. The evidence regarding the beneficial impact of publicly provided health insurance is strong. Public health insurance expansions have been shown to improve birth outcomes among pregnant women (Currie and Gruber,1996a) improve use of preventive care and reduce child mortality (Currie and Gruber, 1996b), increase immunization rates (Joyce and Racine, 2005) and reduce avoidable hospitalizations (Aizer, 2007).

While the above studies provide strong evidence of the impact of public health insurance expansions on child health outcomes, they do not directly link public health insurance with other measures of human capital such as educational attainment or earnings. A second set of studies focusing on direct public health interventions, however, does. These studies include Bleakley (2007) who finds that a de-worming campaign in the American South resulted in improvements in literacy, educational attainment and adult income. This is consistent with work by Miguel and Kremer (2004) showing the positive benefits of a de-worming campaign in Kenya in terms of school attendance, and long term adult employment and earnings (Baird et al 2011).

In sum, evidence based on micro data suggests a positive impact of education (especially geared toward the young), income support and health interventions on child human capital, with the largest effects for health and no documented effects for housing mobility interventions.

[4] Havnes and Mogstad (2010) find that universal child care in Norway has positive effects on children in the lower part of the distribution and are not detected in analyses that focus on estimating the mean effect.

III. Data

The outcomes of interest are not years of schooling or earnings, but rather scores on the PISA reading test among 15 year olds.[5] The main reason for using this outcome is data availability: the full distribution of test scores is available for OECD countries for 2000 and 2009 and these data also include information on parental SES. There are PISA test scores for Math and Science, as well. However, the sample sizes for these two tests are half the size of the sample for the reading test. In the analysis that follows, I focus on reading test scores, but also provide results for math and science.

There are advantages and disadvantages of focusing on child test scores. On the one hand, the test scores are comparable across countries due to the common scale. Moreover, there is evidence that cognitive test scores in late adolescence generally, and PISA test scores specifically, are correlated with better employment outcomes such as lower unemployment and employment in higher status occupations (Lee and Newhouse, 2013). Also, test scores do not reflect inequalities in work opportunities, as earnings can. They are also less likely than measures of educational attainment (ie, years of schooling) to reflect individual choices regarding the pursuit of higher education that might be independently motivated, in part, by a family's resources, though test scores do likely reflect some ambition for higher education. A disadvantage of these measures is that they capture human capital at age 15 and not ultimate human capital attainment.

I calculate three main outcome measures based on the PISA data. The first is a measure of the importance of parental background in determining child human capital and is the estimated elasticity between an index of parental SES and a child's test scores. I refer to this as the measure of *intergenerational elasticity*. The second measure is related to the first, but more explicitly relates parental background to the inequality in child test scores. It is defined as the ratio of test scores of children whose parents are in the top 25% of the SES distribution and those whose parents are in the bottom 25% of the distribution. I refer to this as a measure of the *inheritance of inequality*. The final measure is a measure of the inequality of test scores and is defined as the ratio of the test score of the 90th percentile of the test score distribution to the 10th percentile within a country. I refer to this as a measure of the *inequality in child test scores*.

The data on PISA student test scores and parental SES for 2000 and 2009 are linked with data on public spending in different categories from the OECD SOCX and Educational spending database and available by country and year. The spending is categorized into spending on education, health, housing, family support, labor support, old age and survivors benefits. The values for spending are the average annual spending over the course of the previous 14 years in the life of a child. The only spending variable that cannot be measured over the previous 14 years

[5] The survey tests literacy in terms of "general competencies, that is, how well students can apply the knowledge and skills they have learned at school to real life challenges. PISA does not test how well a student has mastered a school's specific curriculum."

is education spending, which is measured only over the previous 3 years because of data limitations.[6] In the analysis, spending is converted to a percent of GDP. Table 1 lists the countries with consistent data for test scores and spending, the main condition for inclusion in this analysis.

The empirical investigation should be characterized as largely descriptive with reasonable attempts to control for important confounding factors. By looking at trends within a country over time (ie, including country and year fixed effects), one implicitly controls for anything that might be confounding the relationship between public spending and inequality or intergenerational mobility that is country-specific and fixed over time (eg, fixed attitudes towards inequality and redistribution, or the structure of the labor market). The inclusion of year fixed effects also controls for global trends in any factors that are common across countries (eg, global recessions or downturns in the economy). In addition to including country and year fixed effects, I also control for factors that change over time within a country and might be correlated with spending, as discussed below.

All test score and expenditure measures by country and year are presented in Table 1.

IV. Results

A. Measures of Intergenerational Transmission and Inequality: Trends and Correlations

First I show how the three measures of intergenerational transmission and inequality vary across countries and over time. The first measure, *intergenerational elasticity*, categorizes to the relationship between the family's socio-economic status and a child's score on the reading portion of the PISA exam. The measure is obtained be regressing the log of the parent's SES measure on the log of the PISA reading score of the 15 year old child for each country, including controls for child gender and nativity (whether born in the country or not). In 2000, this measure varied from 0.10 for Finland, the most mobile, to 0.23 for Germany, the least mobile.

The second measure, *intergenerational inequality*, is very much related to the first but explicitly reflects the difference in child test scores between those with parents at the top of the SES distribution and those at the bottom. This measure is obtained by calculating the average test scores of those with parents in the bottom and top 25% of the SES distribution within a country/year and calculating the difference. The difference in 2000 ranged from 49 points (Island) to 102 points (Germany), relative to an average test score of 500. These correspond to ratios of 1.23 (Germany) and 1.10 (Iceland). The difference is larger than many cross country differences in average test scores.

[6] Education spending at the country level is not available before 1998.

The third and final measure reflects the *inequality in child test scores* within a country. It is the ratio of the reading score for the 90th percentile to the 10th percentile: the higher the ratio, the greater the inequality. This measure varies from 1.52 for Finland, to 1.78 for Luxembourg and Poland (for 2000).

There appears to be no universal trend in these measures over the period 2000-2009. Figures 2A-2C plot these three measures by country for 2000 and 2009. Those countries above the 45 degree line exhibit an increase in the measure, while those below exhibit a decrease. It's clear from these figures that while many countries witnessed increases in these measures of intergenerational elasticity and inequality of human capital over time, just as many witnessed decreases.

However, within countries, these three measures do trend with both inequality of parental SES and with each other over time. Countries that witnessed the greatest increase in inequality of parental SES, as measured by the 90:10 ratio of parental SES, also witnessed the greatest increase in the inequality of child test scores, as measured by the 90:10 ratio of child test scores (Figure 3A). It is also true that those countries that witnessed the greatest increase in intergenerational elasticity and intergenerational inequality also witnessed the greatest increase in child test score inequality over the period 2000-2009 (Figures 3B and 3C).

We next turn to an examination of whether changes in public investments in social programs correlate with trends in intergenerational transmission and inequality of child human capital.

B. Are Greater Public Investments Associated with Declines in either Intergenerational Transmission or Inequality within a Country Over Time?

Over the period from 1985 to 2009, OECD countries generally witnessed an increase in spending on social programs as a percent of GDP (Figure 4). In Table 1 I present the share of GDP devoted to social programs by country. The numbers, presented in the second panel of the table, reflect not spending in 2000 or 2009, but rather the average spending over the life of a child who was 15 in 2000 or 2009. Spending on social programs as a share of GDP increased on average from 26.1 percent of GDP to 27.8 percent over this period, with the largest increase in spending in health, followed by smaller increases in family and "other" spending which includes old age, survivors and disability support (programs geared mainly to the elderly population).Spending on education as a percent of GDP actually fell slightly, on average, over this period.

I explore whether the increase in spending on social programs is related to student test scores in terms of 1) decreasing the importance of parental background in child test scores as measured by intergenerational elasticities or intergenerational inequality, 2) decreasing test score inequality or 3) increasing test scores at any or all points in the distribution of test scores. To do so, I estimate the following:

$$\Delta Y_c = \beta_1 Ln(\Delta Health)_c + \beta_2\, Ln(\Delta Family)_c + \beta_3\, Ln(\Delta Labor\ Support)_c + \beta_4\, Ln(\Delta Housing)_c + \beta_5\, Ln(\Delta Other\ Social)_c + \beta_6\, Ln(\Delta Education)_c + \beta_7 \Delta SES_c + \beta_8 \Delta Education\ Policies_c + \beta_9\, Ln(\Delta X)_c\ \varepsilon \quad (2)$$

Where ΔY_c is defined as the change or difference for each country between 2009 and 2000 in: 1) the intergenerational elasticity between parental SES and child test score, 2) the ratio in test scores of those with high and low parental SES, 3) child test score inequality (90:10), and 4) the change in test scores at the bottom 10%, top 10% and median of the distribution of test scores.

Spending is measured as the log of spending as a percent of GDP. The last category (Other Spending) includes old age, disability and survivors benefits and as such represents spending on the elderly, primarily. Defining spending as the ln (% of GDP) implicitly controls for the average GDP over the past 15 years. I also control for the median SES index of parents in the country as well as the 10^{th} and 90^{th} percentile of the SES index in the country, population size, the share young, and the share elderly, and the two variables that capture differences in education policies that we think might affect inequality: the share of students in schools with ability grouping (eg, tracking) and the share in private schools. Country and year (2009) fixed effects are included which essentially transforms all variables in the regression into first differences within a country. All regressions are weighted by country population and standard errors are corrected for within country correlations over time.

The results for reading test scores (Table 2) suggest that of the different types of social spending, spending on health is most strongly associated with reductions in the importance of parental background in determining child test scores as measured by the intergenerational elasticity between parental SES and child test scores (column 1) and the difference in test scores of children at the top and bottom of the SES distribution, intergenerational inequality, (column 2). These coefficient estimates imply strong relationships: an increase in health spending of one percentage points is associated with a nearly 70% reduction in intergenerational elasticities, and a 14 percent reduction in intergenerational inequality, evaluated at the overall mean. Spending on family support also seems to be associated with increases in intergenerational mobility, though the estimates, while statistically significant, are much smaller than the estimates for health. Interestingly, other social spending (spending on the elderly), is associated with an increase in intergenerational elasticity and inequality. This would be expected if spending on the elderly crowds out spending on the young. The lack of an association between education spending and intergenerational mobility may reflect the fact that there is little variation in education spending across countries over this period or the fact that the education spending measure does not reflect spending when the children were youngest (due to data limitations) where the micro evidence has found the largest effects.

We also examine how spending affects the 90:10 ratio of test scores (column 3). This measure differs from the previous measure in that it does not necessarily reflect the disparity in test scores of those with relatively rich or poor parents, only the overall variance or inequality in test scores in the population of children. Spending in health, family support and housing appear to be associated with reductions in the 90:10 ratio, with a much larger coefficient estimate on health spending. Again, other social spending which consists largely of spending on the elderly is associated with increases in the 90:10 ratio. Surprisingly, spending on labor support is also positively related to the 90:10 ratio (though it is uncorrelated with intergenerational elasticity or inequality).

When spending is associated with a decline in the 90:10 ratio, it is usually operating through a greater improvement in scores at the bottom of the distribution (health, family support and housing) than the upper part of the distribution (columns 4-6). In contrast, when spending is associated with an increase in the 90:10 ratio (as it is for other spending and, to a lesser extent, labor support) it is operating through a greater reduction in test scores at the bottom of the distribution relative to the top of the distribution.

I repeat the above for math and science test scores (Table 3) for which the patterns differ from reading. In particular, with respect to math test scores, education spending also appears to be associated with reductions in intergenerational elasticities and inequality (Table 3, columns 1-2) that are as large as the reductions associated with health spending. For the science test scores (Table 3, columns 4-6), the estimates are similar to those found for reading test scores, though slightly smaller and less precise.

Overall, the findings suggest that spending on the elderly is associated with both increases in the importance of parental background in determining child test scores and increases in inequality of child test scores. In contrast, spending on health is associated with particularly large improvements in intergenerational mobility and reductions in inequality of human capital of the next generation. Spending on family support and housing are also positively associated with improvements in intergenerational mobility and reductions in inequality, though less so. If one considers math test scores, education spending is also associated with improvements in intergenerational mobility and reductions in inequality. This may imply that the production functions for reading and math test scores differ significantly, with math responding to interventions geared specifically to acquiring math skills and reading scores responding more to interventions geared at general improvements in underlying human capital.

The finding that health spending exhibits the strongest association with intergenerational mobility and inequality warrants further investigation. In particular, unlike some of the other spending measures (education, family support) which are specifically geared toward children, the measure of health spending captures spending on the entire population, not just spending on children. Nor does it necessarily reflect changes in child health. In the next section I explore: 1) how changes in spending on health over time affect the quantity of inputs into the production of child health as well as child health, 2) whether and how changes in inputs/child health are related to changes in reading test scores.

C. Health Spending, Health Inputs and Child Health

1. Relationship between Health Spending and Child Health Inputs/Child Health

I begin with the question of whether and to what extent changes in health spending at the level of the country correspond to changes in child health inputs and health. The OECD measures of child health inputs that I use are the number of pediatricians per 1000 and the share of physicians who are pediatricians. A third measure of health inputs, though not exclusive to children, is the number of hospital beds per 1000. The OECD measure of child health used is infant deaths per 1000 live births. While the OECD currently collects additional measures of child health, most have not been collected for long enough or consistently enough to be included in this analysis.[7] I argue that infant deaths, while not perfect, are a reasonable measure of child health for two reasons. First, infant mortality rates are highly correlated with other measures of population health, more generally, such as the disability adjusted life expectancy (Weidpath and Allotey, 2003). Second, the leading cause of infant mortality is low birth weight (LBW) which has in turn been linked to important long term outcomes such as height (a marker of child nutrition and health), IQ and educational attainment (Black, Deveraux and Salvanes, 2005).[8]

As with the measures of health spending, I calculate 15 year averages over the period 1985-2000 and 1994-2009 for the three measures of health inputs. This is done because we are ultimately interested in how spending over the course of the child's life affects cognitive achievement as measured by the score on the PISA test which is administered to 15 year olds in 2000 and 2009. For the three measures of health inputs, the first two increase over this period by 10 percent, but the number of hospital beds declines nearly 20 percent, coincident with a nearly universal reduction in the average length of hospital stay over this period and a shift toward outpatient care.

Estimates of the extent to which changes in spending on health correspond to changes in child health inputs and health are presented in Table 4. I regress each of the above three measures of child health inputs (pediatricians per 1000, share of physicians who are pediatricians, hospital beds per 1000) and infant deaths/1000 live births on measures of spending on health and education as well as all controls previously included in Tables 2 and 3.

The results show that an increase in health spending is associated with an increase in health inputs geared toward children. A standard deviation increase in health spending leads to a doubling of the number of pediatricians per 1000 evaluated at the mean, while an increase in education spending has a small negative relationship with the number of pediatricians (Table 4, Column 1). To explore whether the estimated effect on pediatricians reflects an increase across

---

[7] Another potential measure of child health that has been collected for many years is the immunization rate for DTP (diphtheria, pertussis and tetanus) and measles. However, the rates are very high even at the beginning of the period (over 90%) so that there is little scope for improvement over time.

[8] One can reasonably assume that a reduction in the infant mortality rate is likely correlated with a reduction in the rate of LBW.

the board in the number of physicians or whether pediatricians are disproportionately affected by changes in public health spending, I present the estimates with respect to health spending and the share of physicians who are pediatricians in column 2 of Table 4. I find that over this period, increases in health spending did disproportionately affect spending on child health as measured by the share of physicians who are pediatricians: an increase in health spending of 1 percent of GDP leads to a 42 percent increase in the share of physicians who are pediatricians over this period. Our third measure of health inputs is the number of hospital beds per 1000 in the country, which, because it is not child-specific, differs significantly from the previous two measures of inputs. While an increase in spending is associated with an increase in the number of hospital beds, the estimate is very imprecise.

Next, I present results for the infant health measure. These estimates are imprecise, but display a generally positive relationship between health spending and child health: as spending increases, infant mortality declines (Table 4, column 4). The lack of precisions is not entirely surprising given the rarity of infant deaths. Future work should attempt to estimate this relationship with richer measures of child health than infant mortality.

Overall, the evidence shows that increases in health spending correspond to an increase in health inputs geared toward children, but the evidence with respect to improvements in health are less clear, though potentially suggestive of a positive relationship.

2. **Relationship between Child Health Inputs/Child Health and Reading Test Scores**

Next I explore the extent to which increases in child health inputs or improvements in child health are associated with reduced inequality of child human capital as measured by reading test scores. For the former, I estimate the extent to which increases in pediatricians and hospital beds correspond to reductions in intergenerational elasticities, intergenerational inequality and inequality of child human capital, as well as improvements in test scores throughout the distribution. To do so, I regress the usual outcomes (intergenerational elasticity, intergenerational inequality, test score inequality, and test scores at the 10%, 90% and median) on the above measures of health inputs as well as spending on education and the controls included in Tables 2 and 3.

I find that increases in the number of physicians per 1000 during the first 15 years of a child's life is very much positively related to reductions in the intergenerational correlations in human capital and inequality in test scores (Table 5, top panel, column 1). For example, a one standard deviation increase in pediatricians per 1000 leads to a 61% decline in intergenerational elasticities, and a 21 percent decline in intergenerational inequality, evaluated at the mean. When looking at associations across the distribution of test scores, the estimates are not precise but generally show that increases in the number of pediatricians are associated with reductions in child test score inequality, with greater improvements at the bottom of the distribution and small reductions at the top. Increases in the number of hospital beds are also associated with reductions in intergenerational correlations and inequality (Table 5, second panel). A standard deviation increase in the number of hospital beds reduces intergenerational correlations by 12 % and inequality by 5 % evaluated at their respective means.

Next I look at the relationship between improvements in infant mortality and reading test scores (Table 5, third panel). Reductions in infant deaths are associated with reductions in intergenerational transmission and inequality (significant at the 10% level) as well as declines in test score inequality, though the associations are moderate in size. A standard deviation decline in infant deaths is associated with a 6% -10% decline in intergenerational elasticity or inequality.

V. Conclusion

Policy makers and academics have become increasingly alarmed by the rise in inequality witnessed in developed countries over the past forty years. Of particular concern are its implications for intergenerational mobility. If rising income inequality results in reduced private investments in child human capital among parents with the fewest resources relative to those with the greatest resources, the result will be a reduction in upward mobility and an increase in the inequality of human capital of the next generation.

However, it is possible for public investments to offset the impact of rising inequality on the human capital of the next generation. In this paper I examined the potential for public spending to offset the unequal distribution of private resources among parents to equalize child human capital. To do so I used PISA test score data for 15 year olds in 25 OECD countries in 2000 and 2009 merged with spending on social programs (health, education, family support, housing assistance, labor support, and spending on the elderly). I find that increases in spending on the elderly are strongly associated with increases in the importance of family background in producing child human capital and increases in inequality of child human capital. In contrast, spending on health is most strongly associated with reductions in the relationship between parental resources and child test scores and reductions in the inequality of child test scores. Spending on housing and family support are more moderately associated with such improvement. This is particularly true for reading test scores. For math test scores, spending on education is also strongly related to improvements in intergenerational mobility and reductions in inequality of human capital.

Upon further inspection of the results for health spending, a clear pattern emerges between the quantity of health services for children (as measured by the number of pediatricians) in a country and test scores. As health services for children increase, test scores of all children rise, but more so for those at the bottom of the test score distribution, resulting in a decline in both inequality of test scores and the importance of parental background in determining test scores. The results suggest that public investments in child human capital, and particularly health, have the potential to offset the negative impact of rising income inequality on the mobility of the next generation, while spending on the elderly may have the opposite effect, most likely due to crowding out of spending on children.   These findings, based on trends over time within a country, while an improvement over an examination based cross-sectional relationships, are only suggestive. Further work establishing a causal relationship between public spending, and inequality of child human capital is needed.

References

Aizer, A. (2007). Public Health Insurance, Program Take-up and Child Health. *Review of Economics and Statistics*, 89(3), 400-415.

Aizer, A and J. Doyle (2013) "The Economics of Child Well-Being: Measuring the Effects of Child Welfare Interventions" in A. Ben-Arieh, F. Casas, I. Frones and J. E. Korbin (Eds.) Handbook of Child Well-Being

Almond, D., Hoynes, H. & Shanzenbach, D. (2011). Inside the War on Poverty: The Impact of Food Stamps on Birth Outcomes. *Review of Economics and Statistics,* 93(2), 387-403.

Angrist, J. D. & Lavy, V. (1999). Using Maimondides' rule to estimate the effect of class size on scholastic achievement. *Quarterly Journal of Economics,* 114(2), 533-575.

Angrist, Joshua D., Lavy, Victor & Schlosser, Analia. (2005). New Evidence on the Causal Link Between the Quantity and Quality of Children. *NBER Working Paper,* 11835.

Barro, Robert and Jong-Wha lee, 2000. International Data on Educational Attainment Updates and Implications

Bailey, M. (2012). Reexamining the Impact of US Family Planning Programs on Fertility: Evidence from the War on Poverty and the Early Years of Title X. *American Economic Journal: Applied Economics,* 4(2), 62-97.

Baird, Sarah, Hamory Hicks, Joan, Kremer, Michael, & Miguel, Edward. (2011). Worms at Work: Long-run Impacts of Child Health Gains. *Working Paper*.

Becker, Gary & Lewis, H. Gregg. (1973). On the Interaction between the Quantity and Quality of Children. *Journal of Political Economy,* 81, S279–S288.

Berger, L. M., Bruch, S. K., Johnson, E. I., James, S., & Rubin, D. (2009). Estimating the 'Impact' of Out-of-Home Placement on Child Well-Being: Approaching the Problem of Selection Bias. *Child Development,* 80(6), 1856-1876.

Black, SE, Devereaux PJ and KJ Salvanes (2007) "From the Cradle to the Labor Market? The Effect of Birth Weight on Adult Outcomes," *Quarterly Journal of Economics*.

Black, S. E., Devereux, P. J. & Salvanes, K. J. (2010). Small Family, Smart Family? Family Size and the IQ Scores of Young Men. *Journal of Human Resources,* 45(1), 33-58.

David M. Blau, 1999. "The Effect Of Income On Child Development," The Review of Economics and Statistics, MIT Press, vol. 81(2), pages 261-276, May.

Black, Sandra, Devereux, Paul and Kjell G Salvanes, 2007. "From the Cradle to the Labor Market? The Effect of Birth Weight on Adult Outcomes," The Quarterly Journal of Economics, vol. 122(1): 409-439.

Bleakley, H. (2007). Disease and Development: Evidence from Hookworm Eradication in the American South. *The Quarterly Journal of Economics*, 122(1), 73-117.

Case, A., Lubotsky, D. & Paxson, C. (2002). Economic Status and Health in Childhood: The Origins of the Gradient. *American Economic Review, 92(5),* 1308-1334.

Chetty, R., Friedman, J., Hilger, N., Saez, E., Whitmore, D., & Yagan, D. (2011). How does your Kindergarten Classroom Affect your Earnings? Evidence from Project Star. *NBER Working Paper*, 16381.

Cunha, F., & Heckman, J. (2007). The Technology of Skill Formation. *The American Economic Review*, 97(2), 31–47.

Currie, J. (2009). Healthy, Wealthy, and Wise? Socioeconomic Status, Poor Health in Childhood, and Human Capital Development. *Journal of Economic Literature,* 47(1), 87-122.

Currie, Janet & Cole, Nancy. (1993). Welfare and Child Health: The Link Between AFDC Participation and Birth Weight. *American Economic Review*, 83(4), 971-985.

Currie, J. & Gruber, J. (1996a). Health Insurance Eligibility, Utilization of Medical Care, and Child Health. *The Quarterly Journal of Economics*, 111(2), 431-466.

Currie, Janet & Gruber, Jonathan. (1996b). Saving Babies: The Efficacy and Cost of Recent Expansions of Medicaid Eligibility for Pregnant Women*.* *Journal of Political Economy*, 104(6), 1263-1296.

Currie, J. & Thomas, D. (1995). Does Head Start make a Difference? *The American Economic Review,* 85(3), 341-64.

Currie, J. & Thomas, D. (1999). Does Head Start Help Hispanic Children. *Journal of Public Economics,* 74(2), 235-62.

Deming, D. (2009). Early Childhood Intervention and Life-Cycle Skill Development: Evidence from Head Start. *American Economic Journal: Applied Economics,* 1(3), 111-134.

Dobbie, Will and Roland Fryer (2011). "Are High-Quality Schools Enough to Increase Achievement Among the Poor? Evidence from the Harlem Children's Zone" AEJ: Applied (3): 158-187.

Doyle, Joseph. (2007). Child Protection and Child Outcomes: Measuring the Effects of Foster Care. *American Economic Review,* 97(5), 1583-1610.

European Commission (2010). "Why Socio-Economic Inequalities Increase? Facts and Policy Responses in Europe."

Fredriksson, Peter, Ockert, Bjorn and Hessel Oosterbeek (2013) :Long-Term Effects of Class Size" Quarterly Journal of Economics. 128(1): 249-285.

Fryer, R. (2010). Financial Incentives and Student Achievement: Evidence from Randomized Trials. *NBER Working Paper,* 15898.

Fryer, Roland G and Lawrence Katz (2013). "Achieving Escape Velocity: Neighborhood and School Interventions to Reduce Persistent Inequality" American Economic Review Papers and Proceedings, 103(3): 232-237.

Garces, E., Currie, J. & Thomas, D. (2002). Longer Term Effects of Head Start. *American Economic Review,* 92(4), 999-1012.

Glewwe, P., Ilias, N. & Kremer, M. (2010). Teacher Incentives. *American Economic Journal: Applied Economics,* 2(3), 205-27.

Gruber, J., Levine, P., & Staiger, D. (1999). Abortion Legalization and Child Living Circumstances: Who is the "Marginal Child"? *Quarterly Journal of Economics,* 114(1), 263-291

Hanushek, E. A. (1992). The trade-off between child quantity and quality. *Journal of Political Economy*, 100 (1), 84-117.

Hanushek, E. A., & Rivkin, S. G. (2004). How to improve the supply of high quality teachers. In D. Ravitch (Ed.), *Brookings papers on education policy* (pp. 7-25). Washington, DC: Brookings Institution Press.

Hanushek, E. A., & Rivkin, S. G. (2006). Teacher quality. In E. A. Hanushek, & F. Welch (Eds.), *Handbook of the economics of education*, vol. 2 (pp. 1051-1078). Amsterdam: North Holland.

Hanushek, E. A., & Rivkin, S. G. (2010). Generalizations about using value added measures of teacher quality. *American Economic Review*, 100(2), 267-271.

Hanushek, E. (2011). The Economic Value of Higher Teacher Quality. *Economics of Education Review,* 20, 466-479.

Havnes and Mogstad (2010) "Is Universal Child Care Leveling the Playing Field? Evidence from non-linear Difference-in-differences" IZA DP No. 4978.

Holter, Hans (2011). "Accounting for Cross-Country Differences in Intergenerational Earnings Persistence: The Impact of Taxation and Public Education Expenditure" Working paper Social Science Research Network.

Hoxby, C. (2000). The Effect of Class Size on Student Achievement: New Evidence from Population Variation. *Quarterly Journal of Economics*, 115(4), 1239-1285.

Jacob, B. (2004). Public Housing, Housing Vouchers and Student Achievement: Evidence from Public Housing Demolitions in Chicago. *American Economic Review*, 94(1), 233-258

Joyce, Ted and Andrew Racine (2005) CHIP Shots: Association Between the State Children's Health Insurance Programs and Immunization Rates. *Pediatrics* Vol. 115 No. 5 May 1, 2005 pp. e526 -e534

Kane, Thomas., Rockoff, Jonah., & Staiger, Douglas. (2006). What does certification tell us about teacher effectiveness? Evidence from NewYork City." *NBER Working Paper*, 12155.

Lee, Jean N and David Newhouse (2013). "Cognitive Skills and Youth Labor Market Outcomes" Background Paper for the World Development Report 2013.

Levine, Phillip B. & Zimmerman, David J., (2005). Children's Welfare Exposure and Subsequent Development. *Journal of Public Economics,* 89(1), 31-56.

Ludwig, Jens & Miller, Douglas L. (2007). Does Head Start Improve Children's Life Chances? Evidence from a Regression Discontinuity Design. *Quarterly Journal of Economics*, 122(1), 159-208.

Mayer, Susan. 1997. What Money Can't Buy: Family Income and Children's Life Chances. Cambridge: Harvard University Press.

Miguel, Edward & Kremer, Michael. (2004). Worms: Identifying Impacts on Education and Health in the Presence of Treatment Externalities. *Econometrica*, 72(1), 159-217.

Milligan, Kevin & Mark Stabile, (2011). "Do Child Tax Benefits Affect the Well-Being of Children? Evidence from Canadian Child Benefit Expansions," American Economic Journal: Economic Policy, American Economic Association, vol. 3(3), pages 175-205, August

Pfeffer, Fabian (2008). "Persistent Inequality in Educational Attainment and its Institutional Context" European Sociological Review. vol 24(5): 543-565.

Oreopoulos, P. (2003). The Long Run Consequences of Living in a Poor Neighborhood. *Quarterly Journal of Economics*, 118(4), 1533-1575.

Reidpath, DD and P Allotey (2003) "Infant mortality rate as an indicator of population health" *Epidemiol Community Health* Vol 2 (57):344-346

Rivkin, Steven, Hanushk, Eric, & Kain, John. (2005). Teachers, Schools and Academic Achievement. *Econometrica*. 73(2), 417-458.

Viaene, Jean-Marie, and Itzhak Zilcha (2001) "Human Capital Formation, Income Inequality and Growth" CESifo Working Paper # 512.

Wilkinson, Richard & Kate Pickett. (2009) The Spirit Level: Why More Equal Societies Almost Always Do Better.US: Bloomsbury Press

Table 1: Averate Reading Scores and Social Spending, 2000 and 2009

	PISA Reading Scores for 15 Year Olds													
	2000							2009						
	IGT	IGE	IGE	50%	90:10	90%	10%	IGT	IGE	IGE	50%	90:10	90%	10%
AUS	0.18	83	1.17	534	1.70	657	386	0.18	76	1.16	515	1.69	634	375
AUT	0.21	72	1.15	505	1.66	615	369	0.20	94	1.22	478	1.76	598	341
BEL	0.22	99	1.21	530	1.74	636	365	0.23	110	1.24	519	1.70	630	371
CAN	0.15	68	1.14	528	1.62	645	398	0.14	53	1.11	515	1.62	629	389
CHE	0.19	91	1.20	502	1.70	619	364	0.16	59	1.13	499	1.63	608	374
DEU	0.23	102	1.23	505	1.75	627	358	0.21	85	1.18	504	1.68	614	366
DNK	0.17	75	1.16	504	1.65	616	372	0.16	78	1.17	485	1.62	591	366
ESP	0.15	63	1.13	501	1.57	596	379	0.15	71	1.16	493	1.63	594	365
FIN	0.10	51	1.10	554	1.52	655	431	0.12	48	1.09	536	1.54	640	414
FRA	0.17	81	1.17	509	1.63	617	379	0.20	95	1.21	506	1.76	624	355
GBR	0.20	94	1.19	528	1.67	651	390	0.20	79	1.17	495	1.67	614	369
GRC	0.14	70	1.16	478	1.74	594	341	0.17	83	1.19	491	1.65	603	365
HUN	0.22	87	1.20	485	1.65	597	362	0.23	91	1.20	504	1.60	610	381
IRL	0.16	75	1.15	532	1.59	645	406	0.19	80	1.18	507	1.63	612	375
ISL	0.11	49	1.10	514	1.62	619	383	0.12	53	1.11	507	1.66	617	372
ITA	0.16	69	1.15	494	1.63	603	369	0.16	76	1.17	498	1.65	605	366
LUX	0.20	100	1.25	450	1.78	563	316	0.26	113	1.27	485	1.79	602	337
NLD	0.17	74	1.15	552	1.55	649	420	0.17	69	1.14	520	1.56	627	401
NOR	0.19	70	1.15	514	1.74	634	364	0.18	65	1.14	506	1.62	620	382
NZL	0.18	79	1.16	536	1.72	657	382	0.23	83	1.17	532	1.68	648	385
POL	0.20	82	1.19	474	1.78	593	334	0.16	75	1.16	509	1.58	618	391
PRT	0.22	94	1.22	484	1.71	596	348	0.19	81	1.18	493	1.60	598	373
SWE	0.16	70	1.14	523	1.61	629	391	0.19	80	1.17	503	1.68	621	370
USA	0.19	89	1.19	500	1.76	629	358	0.19	82	1.18	500	1.67	622	371
Average, unweighted	0.18	78	1.17	508	1.67	621	372	0.18	78	1.17	503	1.65	614	372
Average, weighted	0.19	85	1.18	504	1.71	624	366	1.12	81	1.18	501	1.67	617	370

	Social Spending as a Percent of GDP													
	2000							2009						
	Total	Education	Health	Labor	Housing	Family	Other	Total	Education	Health	Labor	Housing	Family	Other
AUS	17.74	3.03	4.66	0.40	0.22	1.95	7.48	18.29	1.04	5.39	0.39	0.21	2.74	8.51
AUT	26.25	6.58	5.70	0.36	0.07	2.07	11.47	33.47	5.54	6.59	0.57	0.10	2.80	17.86
BEL	30.26	4.34	6.44	1.16	0.00	2.39	15.93	32.50	6.29	6.91	1.15	0.05	2.58	15.52
CAN	24.46	6.28	6.41	0.52	0.61	0.71	9.93	23.07	5.84	6.56	0.38	0.51	0.95	8.82
CHE	23.49	2.43	4.44	0.39	0.11	1.16	14.96	27.89	1.94	5.40	0.65	0.14	1.39	18.37
DEU	31.80	5.64	7.41	1.06	0.24	1.94	15.51	32.80	4.89	8.10	1.08	0.42	2.07	16.25
DNK	34.03	7.58	4.85	1.17	0.66	3.33	16.44	34.59	7.35	5.55	1.69	0.70	3.45	15.85
ESP	25.44	5.39	5.07	0.58	0.13	0.39	13.88	25.95	4.95	5.56	0.66	0.19	0.93	13.65
FRA	33.50	6.39	6.76	0.95	0.82	2.69	15.89	34.88	5.99	7.44	1.07	0.85	2.99	16.54
GBR	24.51	5.15	5.22	0.51	1.45	2.23	9.94	25.98	5.54	6.06	0.34	1.50	2.84	9.70
GRC	21.15	4.43	4.11	0.24	0.40	0.70	11.29	24.02	4.27	5.17	0.22	0.55	1.06	12.75
HUN	15.58	6.52	6.11	0.43	0.06	0.32	2.14	18.32	2.26	5.56	0.37	0.41	2.33	7.39
IRL	22.30	5.51	4.88	1.13	0.59	1.62	8.57	20.23	5.20	5.25	0.87	0.36	2.16	6.39
ISL	24.09	8.00	5.95	0.06	0.06	2.39	7.64	25.03	7.94	6.08	0.08	0.20	2.68	8.06
ITA	28.60	5.06	5.62	0.22	0.01	0.78	16.91	30.28	4.80	6.11	0.53	0.01	1.18	17.65
NDL	30.27	5.64	5.48	1.38	0.37	1.57	15.83	27.32	5.64	5.61	1.38	0.38	1.60	12.72
NOR	27.25	6.88	3.88	0.82	0.15	2.71	12.80	29.23	5.63	5.43	0.75	0.17	3.12	14.14
NZL	19.69	0.00	5.41	0.77	0.44	2.47	10.60	25.22	6.52	6.25	0.47	0.76	2.71	8.51
POL	25.33	3.40	4.21	0.37	0.10	1.44	15.80	27.15	5.70	4.23	0.40	0.14	1.10	15.59
PRT	20.15	5.85	4.08	0.41	0.00	0.72	9.09	26.01	5.48	6.15	0.58	0.00	1.05	12.76
SWE	38.11	6.67	6.90	2.17	0.83	4.02	17.52	36.20	6.32	6.51	1.64	0.66	3.29	17.77
USA	21.81	7.07	5.49	0.21	0.00	0.56	8.48	22.87	7.02	6.70	0.14	0.00	0.67	8.32
Average, unweighted	25.72	5.36	5.41	0.70	0.33	1.73	12.19	27.33	5.28	6.03	0.70	0.38	2.08	12.87
Average, weighted	26.10	5.91	5.67	0.51	0.28	1.27	11.78	27.76	5.79	6.48	0.51	0.31	1.50	11.95
Change 2009-2000	1.66	-0.12	0.81	0.00	0.03	0.23	0.16							

Table 2: Categories of Spending (as % GDP) and Reading Score Gains

	(1) Elasticity	(2) Inequality	(3) 90:10	(4) 10%	(5) 90%	(6) 50%
Ln(health spending as % GDP)	-0.197	-0.162	-1.007	263.0	69.00	153.0
	[0.0284]	[0.107]	[0.133]	[15.76]	[74.42]	[48.75]
Ln(family support spending as % GDP)	-0.0552	-0.106	-0.225	76.86	49.30	65.75
	[0.0127]	[0.0476]	[0.0595]	[7.031]	[33.20]	[21.75]
Ln(housing as % GDP)	-0.0109	-0.0757	-0.197	62.57	33.05	53.71
	[0.0170]	[0.0638]	[0.0797]	[9.420]	[44.47]	[29.14]
Ln(education spending as % GDP)	0.00217	-0.0483	0.00337	-23.97	-36.14	-36.18
	[0.0122]	[0.0459]	[0.0574]	[6.776]	[31.99]	[20.96]
Ln(labor support spending as % GDP)	0.0431	0.0345	0.470	-143.6	-69.98	-100.8
	[0.0128]	[0.0482]	[0.0603]	[7.124]	[33.63]	[22.03]
Ln(other spending as % GDP)	0.104	0.244	0.775	-271.5	-176.6	-236.9
	[0.0318]	[0.120]	[0.150]	[17.68]	[83.49]	[54.69]
SES Index- top 10%	0.00500	0.0320	0.0360	-8.857	-1.788	-6.230
	[0.00417]	[0.0157]	[0.0196]	[2.315]	[10.93]	[7.159]
SES Index - bottom 10%	0.00790	0.00627	-0.0271	10.44	7.912	8.883
	[0.00422]	[0.0159]	[0.0198]	[2.343]	[11.06]	[7.248]
SES Index - median	0.00183	-0.0188	0.0247	-8.121	-4.310	-5.607
	[0.00572]	[0.0215]	[0.0269]	[3.176]	[15.00]	[9.825]
Share of population >65	-1.176	0.473	-21.63	7,879	5,349	6,654
	[0.843]	[3.171]	[3.964]	[468.3]	[2,211]	[1,448]
Share of population <15	2.127	4.376	12.83	-4,177	-2,343	-3,288
	[0.409]	[1.538]	[1.923]	[227.1]	[1,072]	[702.5]
Country Population	0.00466	0.00567	0.0475	-16.91	-11.02	-13.68
	[0.00136]	[0.00513]	[0.00642]	[0.758]	[3.578]	[2.344]
Education tracking	-0.0579	0.147	0.175	-117.2	-138.8	-144.1
	[0.0445]	[0.167]	[0.209]	[24.70]	[116.6]	[76.40]
Share in private school	0.0713	-0.111	0.312	-137.0	-109.1	-125.5
	[0.0533]	[0.201]	[0.251]	[29.62]	[139.8]	[91.61]
Year=2009	0.0553	0.0366	0.306	-97.36	-50.68	-75.40
	[0.00883]	[0.0332]	[0.0415]	[4.902]	[23.15]	[15.16]
Observations	36	36	36	36	36	36
R-squared	0.999	0.992	0.997	0.999	0.989	0.992
Mean of dependent variable	0.181	1.18	1.17	366	624	500

Standard errors in brackets. All standard errors adjusted for clustering within Country.
Elasticity refers to intergenerational elasticity and is the coefficient on ln(parental SES) in a country*year specific regression in which ln(child test scores) is regressed on ln(parental SES) and controls for gender and nativity.
Inequality refers to the ratio of test socres for those with parents in the top 25% of the SES distribution to the test scores for those whose parents are in the bottom 25% of the SES distribution.
90:10 refers to the ratio of test scores in the 90th percentile to the 10th percentile
10% refers to the test score of the bottom 10% of the distribution, 90% refers to the test scores of the top 10% of the test score distribution and 50% refers to the median.

Table 3: Categories of Spending (as % GDP) and Gains in Math and Science Scores

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Math Test Scores			Science Test Scores			Averaged Test Score		
	Elasticity	Inequality	90:10	Elasticity	Inequality	90:10	Elasticity	Inequality	90:10
Ln(health spending as % GDP)	-0.0771	-0.0606	-0.687	-0.182	-0.131	-0.695	-0.171	-0.132	-0.858
	[0.0317]	[0.0204]	[0.259]	[0.0264]	[0.0994]	[0.0879]	[0.00147]	[0.0679]	[0.00427]
Ln(family support spending as % GDP)	-0.0359	-0.0724	-0.0874	-0.0894	-0.131	-0.334	-0.0524	-0.0929	-0.178
	[0.0141]	[0.00910]	[0.116]	[0.0118]	[0.0443]	[0.0392]	[0.000654]	[0.0303]	[0.00190]
Ln(housing as % GDP)	0.0106	-0.0500	-0.177	0.0176	-0.0392	-0.131	0.00241	-0.0592	-0.172
	[0.0190]	[0.0122]	[0.155]	[0.0158]	[0.0594]	[0.0525]	[0.000876]	[0.0406]	[0.00255]
Ln(education spending as % GDP)	-0.0223	-0.0683	-0.0465	0.0916	0.0660	0.0860	0.0205	-0.0253	0.0132
	[0.0136]	[0.00877]	[0.111]	[0.0114]	[0.0427]	[0.0378]	[0.000630]	[0.0292]	[0.00183]
Ln(labor support spending as % GDP)	0.0655	0.0558	0.496	0.0881	0.0835	0.398	0.0623	0.0481	0.445
	[0.0143]	[0.00922]	[0.117]	[0.0119]	[0.0449]	[0.0397]	[0.000663]	[0.0307]	[0.00193]
Ln(other spending as % GDP)	0.00989	0.121	0.435	0.155	0.288	0.834	0.0880	0.207	0.635
	[0.0356]	[0.0229]	[0.291]	[0.0297]	[0.112]	[0.0986]	[0.00164]	[0.0762]	[0.00479]
SES Index- top 10%	-0.00248	0.0219	0.00274	0.000599	0.0259	0.0313	0.000634	0.0262	0.0252
	[0.00466]	[0.00299]	[0.0381]	[0.00388]	[0.0146]	[0.0129]	[0.000215]	[0.00997]	[0.000627]
SES Index - bottom 10%	0.00358	0.00118	-0.0333	0.00903	0.00802	-0.0388	0.00732	0.00572	-0.0232
	[0.00471]	[0.00303]	[0.0386]	[0.00393]	[0.0148]	[0.0131]	[0.000218]	[0.0101]	[0.000634]
SES Index - median	0.00758	-0.00886	0.0365	0.0102	-0.0111	0.0348	0.00670	-0.0128	0.0268
	[0.00639]	[0.00411]	[0.0523]	[0.00533]	[0.0200]	[0.0177]	[0.000295]	[0.0137]	[0.000860]
Share of population >65	-2.608	-0.957	-26.13	-4.401	-2.875	-24.93	-2.600	-0.752	-22.16
	[0.942]	[0.606]	[7.705]	[0.785]	[2.954]	[2.610]	[0.0436]	[2.017]	[0.127]
Share of population <15	0.724	2.503	5.055	2.644	4.739	12.84	1.784	3.697	9.670
	[0.457]	[0.294]	[3.737]	[0.381]	[1.432]	[1.266]	[0.0211]	[0.978]	[0.0615]
Country Population	0.00308	0.00220	0.0333	0.00986	0.0104	0.0521	0.00553	0.00557	0.0405
	[0.00152]	[0.000980]	[0.0125]	[0.00127]	[0.00478]	[0.00422]	[7.05e-05]	[0.00326]	[0.000205]
Education tracking	-0.113	0.0348	0.0195	-0.109	0.104	0.332	-0.0955	0.0888	0.105
	[0.0497]	[0.0320]	[0.406]	[0.0414]	[0.156]	[0.138]	1.12	[0.106]	[0.00669]
Share in private school	0.0193	-0.126	0.211	0.367	0.227	0.603		-0.0330	0.347
	[0.0596]	[0.0383]	[0.487]	[0.0497]	[0.187]	[0.165]	[0.00275]	[0.128]	[0.00802]
Year=2009	0.0597	0.0468	0.301	0.0951	0.0696	0.267	0.0680	0.0487	0.284
	[0.00986]	[0.00634]	[0.0807]	[0.00822]	[0.0309]	[0.0273]	[0.000456]	[0.0211]	[0.00133]
Observations	36	36	36	36	36	36	36	36	36
R-squared	0.999	1.000	0.990	0.999	0.994	0.999	1.000	0.997	1.000

Standard errors in brackets. All standard errors adjusted for clustering within Country.
Elasticity refers to intergenerational elasticity and is the coefficient on ln(parental SES) in a country*year specific regression in which ln(child test scores) is regressed on ln(parental SES) and controls for gender and nativity.
Inequality refers to the ratio of test socres for those with parents in the top 25% of the SES distribution to the test scores for those whose parents are in the bottom 25% of the SES distribution.
90:10 refers to the ratio of test scores in the 90th percentile to the 10th percentile

Table 4: Changes in Health Spending and Health Inputs Over Time

	(1) Pediatrician/1000	(2) % Pediatricians	(3) Hospital Beds/1000	(4) Infant Deaths/1000
Ln(health spending as % GDP)	0.0772	1.647	5.388	-7.589
	[0.00978]	[0.631]	[4.281]	[6.613]
Ln(education spending as % GDP)	-0.0123	-0.532	-0.706	1.354
	[0.00671]	[0.644]	[0.992]	[2.674]
SES Index- top 10%	0.000875	0.133	0.0748	-0.214
	[0.00355]	[0.234]	[0.262]	[0.381]
SES Index - bottom 10%	-0.000943	-0.158	0.0557	0.241
	[0.00157]	[0.170]	[0.216]	[0.510]
SES Index - median	0.00201	0.0535	-0.164	0.478
	[0.00119]	[0.142]	[0.311]	[0.535]
Share of population >65	-0.257	-40.67	16.75	-81.67
	[0.227]	[42.19]	[54.65]	[83.18]
Share of population <15	0.167	-4.686	22.62	6.531
	[0.125]	[15.90]	[26.13]	[56.85]
Country Population	0.000283	0.00735	0.0116	0.0887
	[0.000117]	[0.0146]	[0.0324]	[0.0784]
Year=2009	0.00708	0.191	-1.828	-2.243
	[0.00581]	[0.370]	[1.201]	[1.240]
Observations	36	35	39	40
R-squared	0.999	0.998	0.966	0.951
Mean of dependent variable	0.11	3.8	5.97	8.7

Standard errors in brackets. All standard errors adjusted for clustering within Country.
All controls included in Table 2 also included.

Table 5: Changes in Health Inputs and Changes in Reading Test Scores

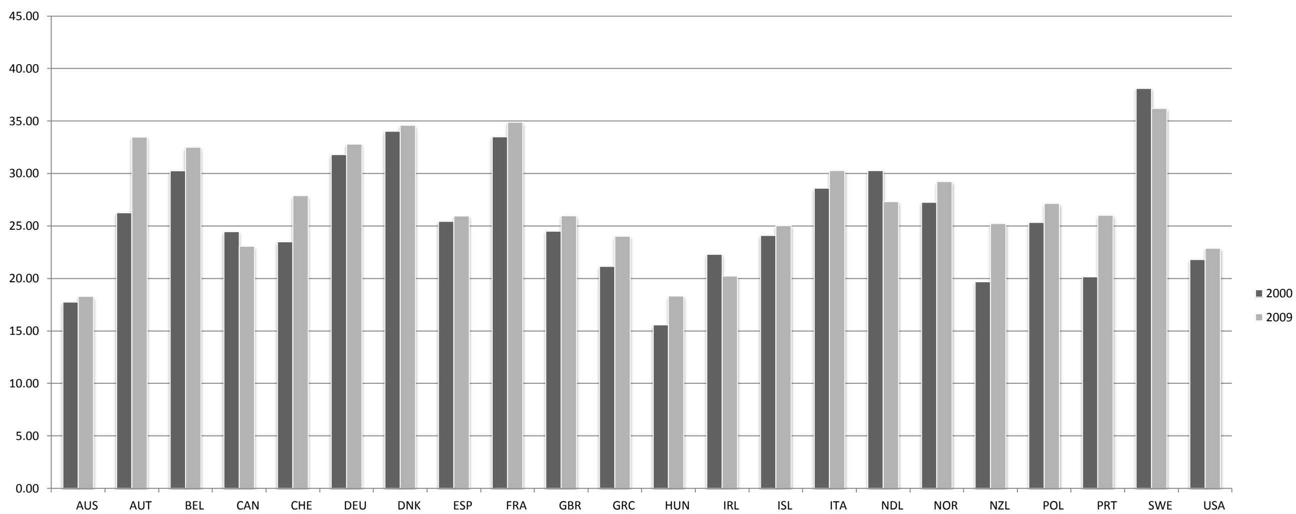
	(1) Elasticity	(2) Inequality	(3) 90:10	(4) 10%	(5) 90%	(6) 50%
Pediatricians per 1000, avg over past 15 years	-1.979	-1.883	-5.065	901.1	-374.6	116.3
	[0.465]	[0.766]	[3.004]	[756.0]	[391.9]	[427.4]
Observations	38	38	38	38	38	38
R-squared	0.981	0.970	0.881	0.912	0.976	0.974

	(1) Elasticity	(2) Inequality	(3) 90:10	(4) 10%	(5) 90%	(6) 50%
Total hospital beds per 1000, avg over past 15 years	-0.0100	-0.0118	-0.0364	6.986	-1.928	2.125
	[0.00271]	[0.00714]	[0.0102]	[2.645]	[3.522]	[3.372]
Observations	39	39	39	39	39	39
R-squared	0.979	0.918	0.891	0.879	0.933	0.904

	(1) Elasticity	(2) Inequality	(3) 90:10	(4) 10%	(5) 90%	(6) 50%
Infant deaths per 1000 live births	0.00501	0.00646	0.0294	-7.960	-2.516	-4.579
	[0.00301]	[0.00350]	[0.0112]	[1.842]	[2.139]	[1.630]
Observations	42	42	42	42	42	42
R-squared	0.931	0.911	0.817	0.861	0.936	0.923
Mean of dependent variable	0.181	1.18	1.17	366	624	500

Standard errors in brackets. All standard errors adjusted for clustering within Country.
Elasticity refers to intergenerational elasticity and is the coefficient on ln(parental SES) in a country*year specific regression in which ln(child test scores) is regressed on ln(parental SES) and controls for gender and nativity.
Inequality refers to the ratio of test socres for those with parents in the top 25% of the SES distribution to the test scores for those whose parents are in the bottom 25% of the SES distribution.
90:10 refers to the ratio of test scores in the 90th percentile to the 10th percentile
10% refers to the test score of the bottom 10% of the distribution, 90% refers to the test scores of the top 10% of the test score distribution and 50% refers to the median.
All controls from Table 2 also included.

Figure 1: Social Spending as a % GDP in 2000 and 2009



Spending consists of spending on education, health, family support, labor support, housing and other (disability, old age and survivors benefits) as indicated in the OECD SOCX and education databases.

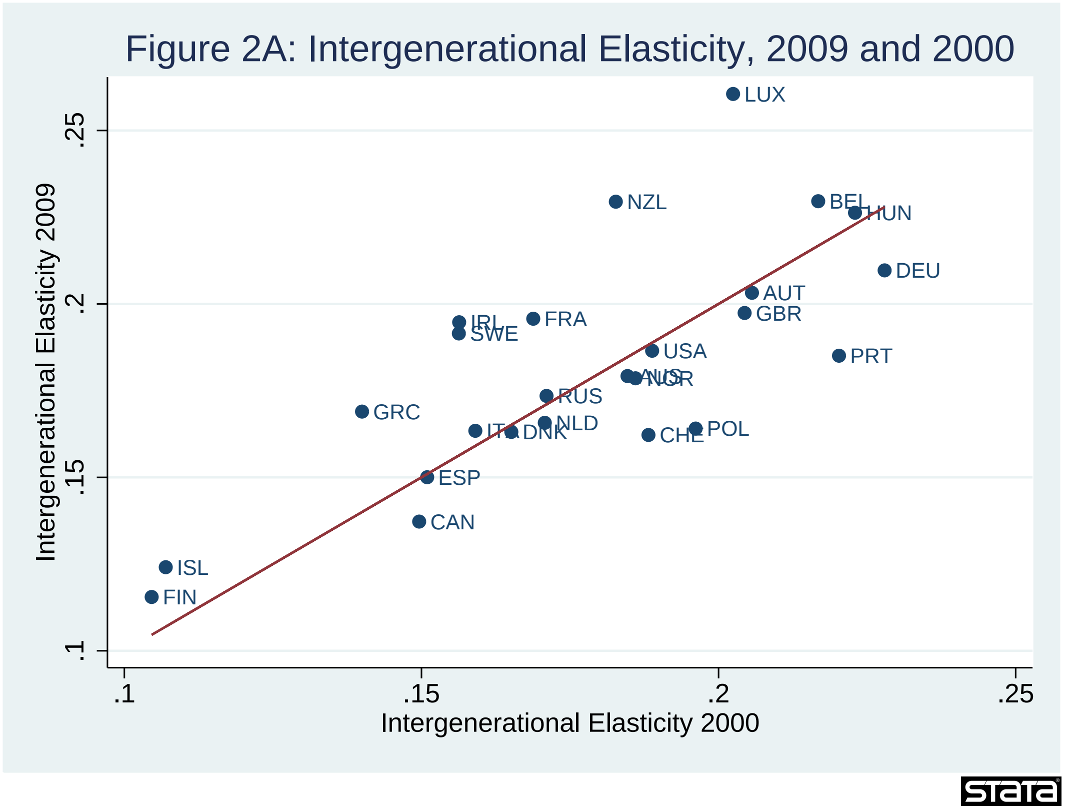


Figure 2A: Intergenerational Elasticity, 2009 and 2000
Intergenerational Elasticity 2009
Intergenerational Elasticity 2000
.1
.15
.2
.25
LUX
NZL
BEL
HUN
DEU
AUT
GBR
IRL
SWE
FRA
USA
PRT
AUS
NOR
RUS
GRC
ITA
DNK
NLD
CHE
POL
ESP
CAN
ISL
FIN

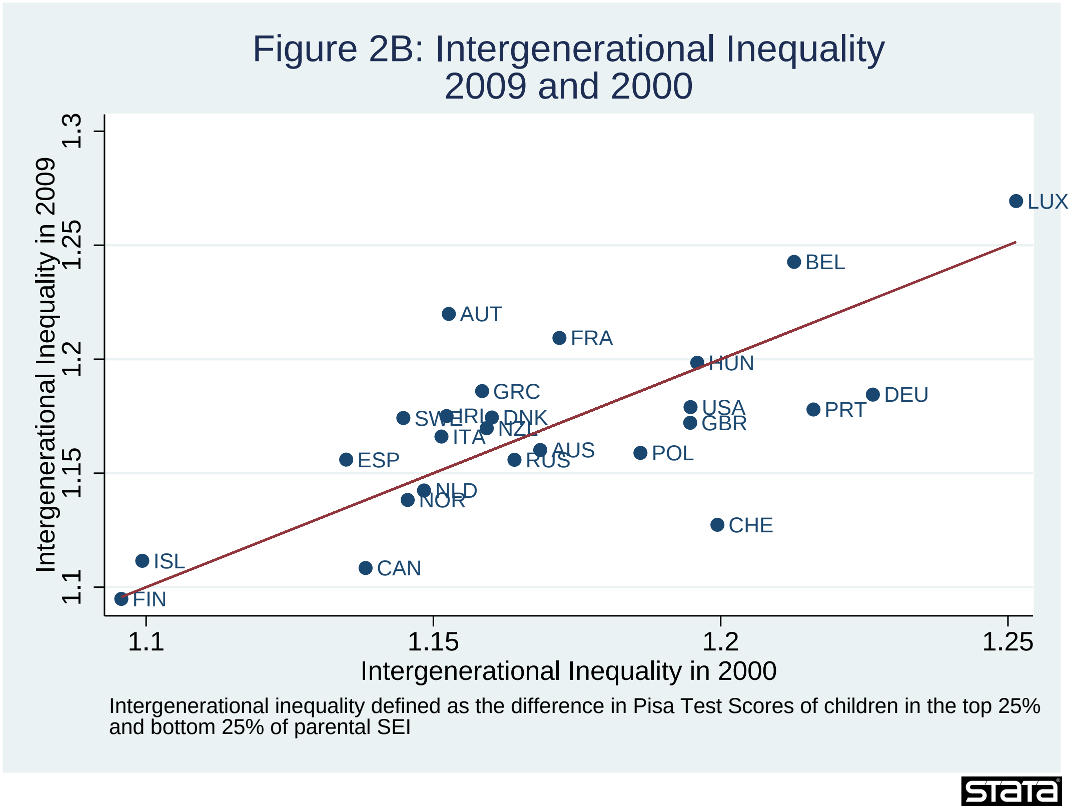


Figure 2B: Intergenerational Inequality
2009 and 2000
Intergenerational Inequality in 2009
Intergenerational Inequality in 2000
1.1
1.15
1.2
1.25
1.3
LUX
BEL
AUT
FRA
HUN
GRC
DEU
USA
PRT
GBR
SWE
IRL
DNK
NZL
ITA
AUS
ESP
RUS
POL
NLD
NOR
CHE
ISL
CAN
FIN
Intergenerational inequality defined as the difference in Pisa Test Scores of children in the top 25%
and bottom 25% of parental SEI

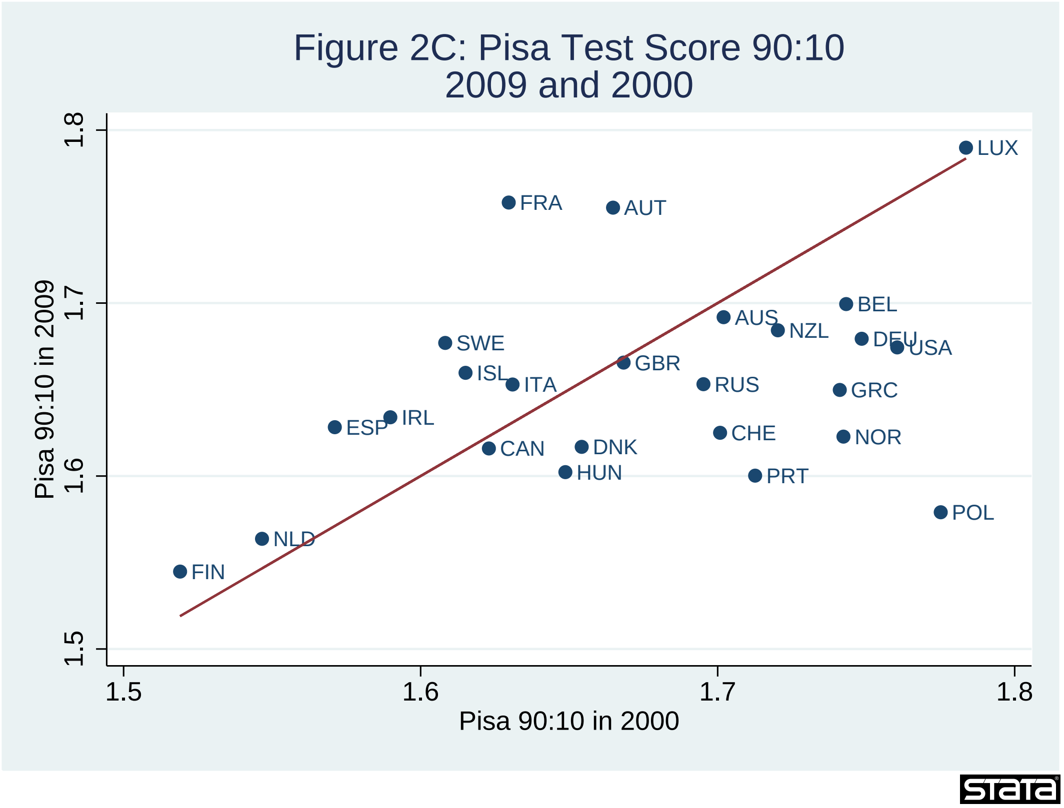


Figure 2C: Pisa Test Score 90:10
2009 and 2000
Pisa 90:10 in 2009
Pisa 90:10 in 2000
1.5
1.6
1.7
1.8
LUX
FRA
AUT
BEL
AUS
NZL
DEU
USA
SWE
GBR
ISL
ITA
RUS
GRC
IRL
ESP
CHE
NOR
CAN
DNK
HUN
PRT
POL
NLD
FIN

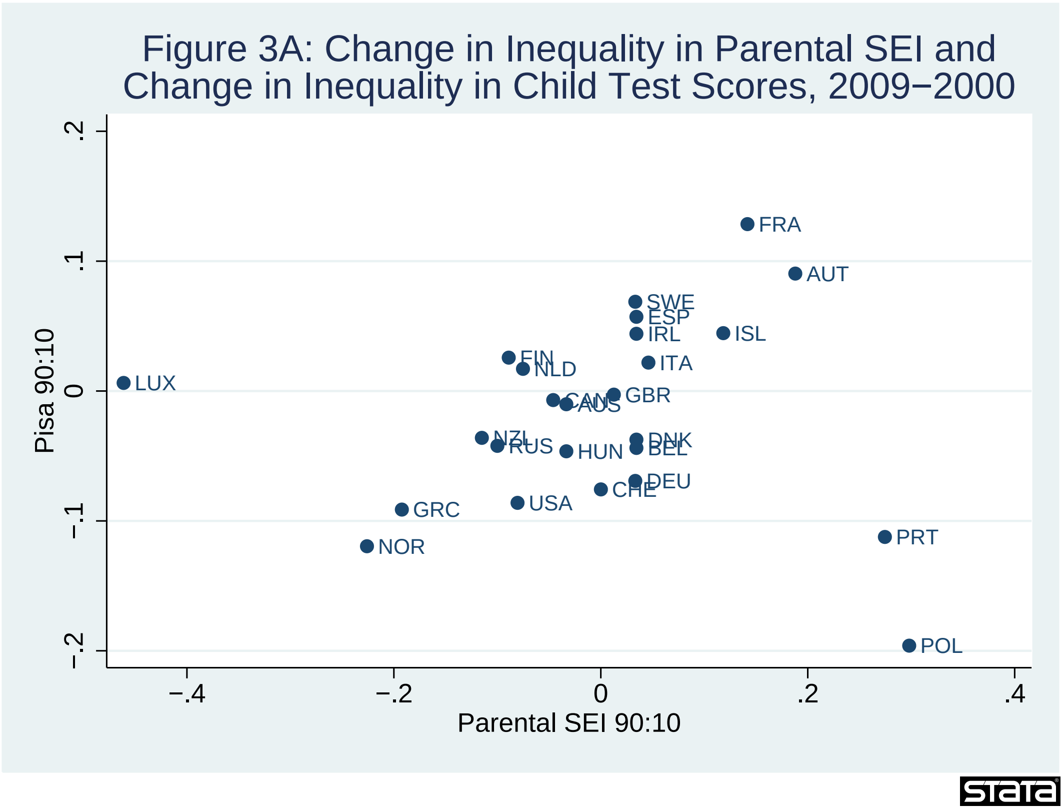


Figure 3A: Change in Inequality in Parental SEI and
Change in Inequality in Child Test Scores, 2009−2000
Pisa 90:10
Parental SEI 90:10
.2
.1
0
−.1
−.2
−.4
−.2
0
.2
.4
FRA
AUT
SWE
ESP
IRL
ISL
FIN
NLD
ITA
LUX
GBR
CAN
AUS
NZL
RUS
HUN
DNK
BEL
DEU
CHE
USA
GRC
PRT
NOR
POL

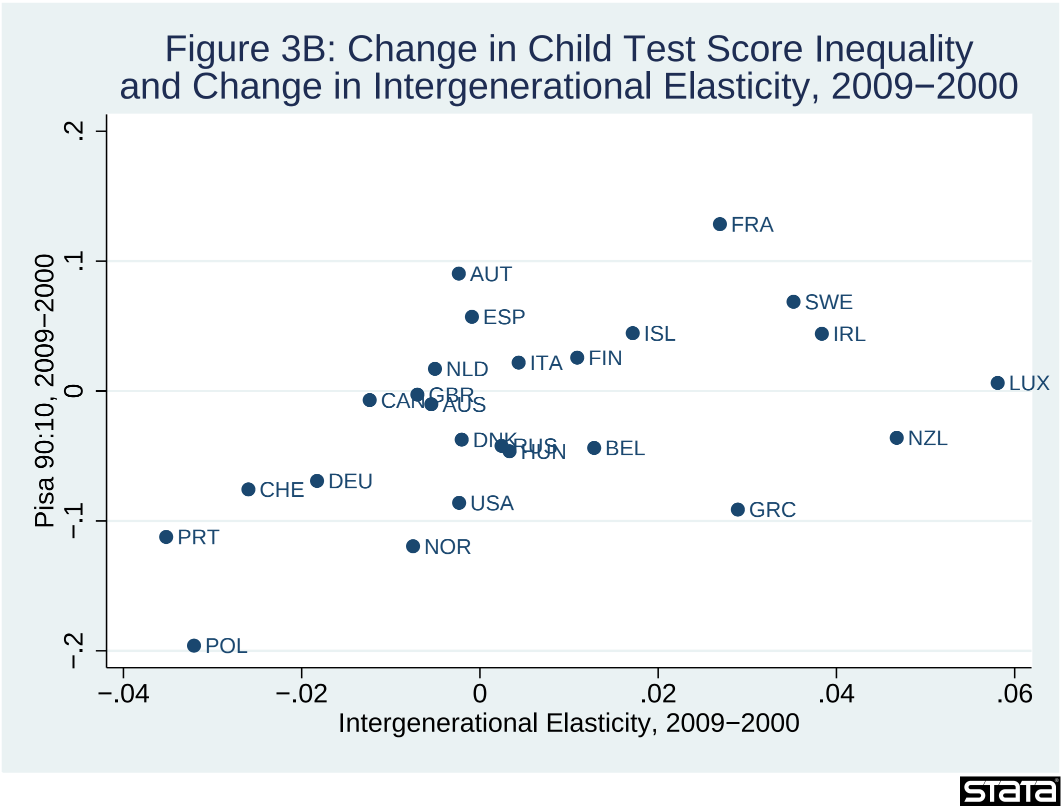
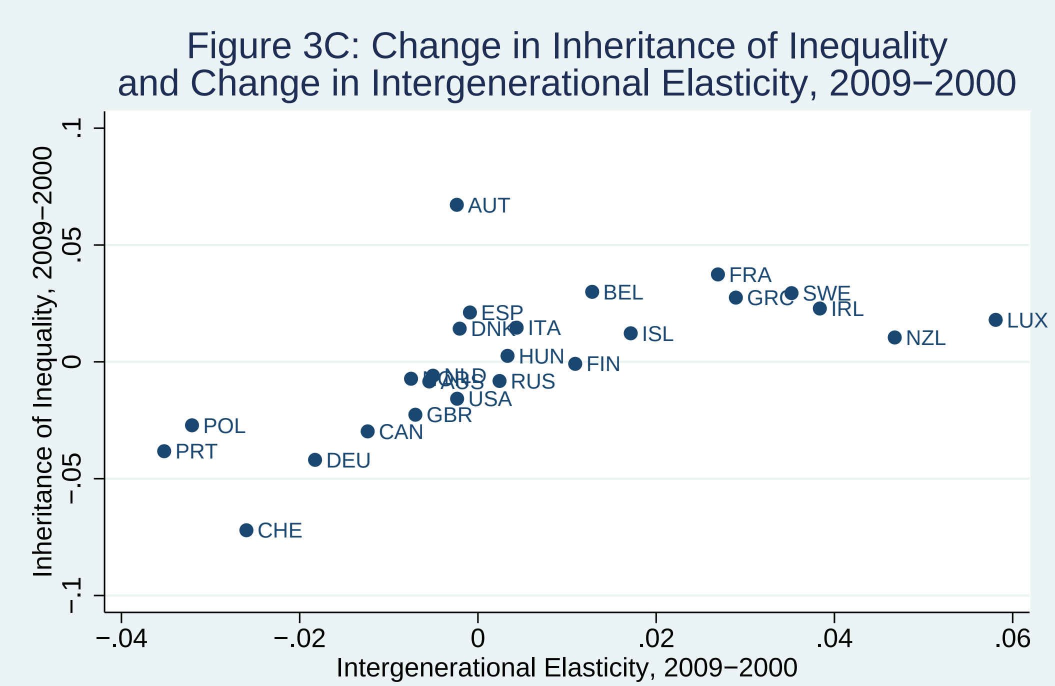


Figure 3B: Change in Child Test Score Inequality
and Change in Intergenerational Elasticity, 2009−2000
Pisa 90:10, 2009−2000
Intergenerational Elasticity, 2009−2000
−.2
−.1
0
.1
.2
−.04
−.02
0
.02
.04
.06
FRA
AUT
SWE
ESP
ISL
IRL
FIN
ITA
NLD
LUX
CAN
GBR
AUS
DNK
RUS
HUN
BEL
NZL
DEU
CHE
USA
GRC
PRT
NOR
POL



Inheritance of inequality defined as the ration of Pisa Test Scores of children in the top 25% and bottom 25% of parental SEI

CESifo Economic Studies, Vol. 60, 2/2014, 312–337 doi:10.1093/cesifo/ifu011
Advance Access publication 6 March 2014

Do Family Planning Programs Decrease Poverty? Evidence from Public Census Data

Martha J. Bailey*, Olga Malkova and Johannes Norling

*Department of Economics, University of Michigan, 611 Tappan Street, Ann Arbor, MI 48109, USA, e-mail: baileymj@umich.edu

Abstract

This article provides new evidence that family planning programs are associated with a decrease in the share of children and adults living in poverty. Our research design exploits the county roll-out of US family planning programs in the late 1960s and early 1970s and examines their relationship with poverty rates in the short and longer-term in public census data. We find that cohorts born after federal family planning programs began were less likely to live in poverty in childhood and that these same cohorts were less likely to live in poverty as adults. (JEL codes: I3, J13, J18)

With US income inequality soaring to its highest level in almost a century (Saez 2013), increasing the economic opportunities of poor children is a growing policy concern. Poor children are significantly more likely to experience delayed academic development, have health problems, live in more dangerous neighborhoods, and attend underperforming schools (Levine and Zimmerman 2010). In the longer-term, children from poorer households have lower test scores (Reardon 2011), are less likely to complete high school, enroll in college, and, conditional upon enrolling, complete college (Bailey and Dynarski 2011), which limits their earnings potential as adults. Ultimately, over 40% of children born to parents in the lowest quintile of family income remain in that income quintile as adults (Pew Charitable Trusts 2012).

This article explores the role of family planning programs as a public policy strategy to improve children's economic resources in childhood. The rationale that family planning programs would increase children's resources and opportunities was integral to their inclusion in US President Lyndon B. Johnson's War on Poverty, which began in 1964. Five years later, when campaigning for a national family planning program, President Richard Nixon asserted their more direct connection to children's economic disadvantage: "Unwanted or untimely childbearing is one of several forces which are driving many families into poverty or keeping them in that condition" (18 July 1969).

A long theoretical tradition in economics also rationalizes a causal link running from children's economic resources, to their lifetime opportunities, and ultimately to their adult outcomes.[1] This link occurs both through income and price channels. More affluent parents not only have more economic resources, but they may invest more in each child and have fewer children if the income elasticity of parental investments in children ("child quality") exceeds the income elasticity of child quantity (Becker and Lewis 1973; Willis 1973). Having fewer children, in turn, reduces the shadow price of child quality and further encourages investment in children. In addition, credit constraints may lead poorer families to underinvest in their children's formal human capital (Becker and Tomes 1979, 1986).

Family planning programs could increase investments in children through both income and price channels. First, they may induce greater parental investments in their children by reducing the relative price of child quality. Second, they may raise the incomes of the average parent, for instance by reducing the cost of delaying childbearing so that parents can themselves increase their human capital investments, find better partners, and, ultimately, earn higher wages (Christenson 2011; Rotz 2011; Bailey et al. 2012). Family planning programs could also raise the family income of the average child as they disproportionately allow poorer households to delay or avoid additional childbearing.

This article provides new empirical evidence on the relationship of family planning programs to child poverty rates, both in the short and long-run. Building on Bailey's (2012) research design, we exploit the roll-out of US federally funded family planning grants from 1964 to 1973. The first US family planning programs were quietly funded under the 1964 Economic Opportunity Act (EOA) and the program expanded under the Family Planning Services and Population Research Act (Public Law 91-572).[2] This legislation supported the opening of new clinics in

1 Thomas Malthus popularized the link between childbearing and poverty in his *Essay on the Principle of Population* (1798). Malthus argued that this link was rooted in the arithmetic growth of agricultural yields being outstripped by the exponential growth of population. Left unchecked, population growth would outstrip the growth in agricultural production and result in a subsistence economy.

2 Before 1965, US federal involvement and investments in family planning had been modest. This reflected the view expressed by President Dwight Eisenhower in 1959, who said that he could not "imagine anything more emphatically a subject that is not a proper political or government activity or function or responsibility... The government will not, so long as I am here, have a positive political doctrine in its program that has to do with the problem of birth control. That's not our business" (Tone 2001, p. 214). According to 1967 estimates, expenditure for family planning through the Maternal and Child Health programs (started in 1942; US Department of Health, Education, and Welfare [DHEW] 1974, p. 3, citing a 1942 memorandum from Surgeon General Thomas Parran to state health departments) and the Maternal and Infant Care programs under the 1963 Social Security Amendments were small (US DHEW 1974, p. 3, citing House Appropriations Committee hearings; US DHEW 1967, p. 988).

disadvantaged areas and, to a lesser extent, the expansion of existing family planning programs. Federal family planning dollars funded education, counseling, and the provision of low-cost contraceptives and related medical services; they did not fund abortion, which remained illegal in most states until 1973. Use of these programs was not explicitly means tested, but programs tended to benefit lower income women.

Our research design compares the poverty rates of individuals born in the years leading up to and just after federally funded family planning programs began. We draw upon several public-use datasets that measure individuals' ages and place of residence: the 1980 US decennial census observes the potentially affected cohorts as children and the 2000 census and 2005–2011 American Community Survey (ACS) observe the same cohorts as adults.

Our results show that federally funded family planning programs are associated with significant reductions in child poverty rates and, later, poverty rates in adulthood.[3] Individuals born 1–6 years after program funding were 4.2% less likely to live in poverty in childhood and 2.4% less likely to live in poverty in adulthood. Although both white and non-white children born after family planning programs began experienced large reductions in childhood poverty, white children experienced greater relative reductions in poverty rates in adulthood. Whites born after family planning programs began were 4.1% less likely to live in poverty in childhood and 6.1% less likely to live in poverty in adulthood. Non-whites born after family planning programs began were 8.2% less likely to live in poverty in childhood, but 2% less likely to live in poverty in adulthood.

In short, family planning programs may help break the cycle of poverty. Our results suggest that family planning programs reduce poverty among children and, ultimately, in adulthood. These findings complement a growing body of research that suggests that investments in children can have sizable effects on children's longer-term educational attainment, health, and labor market productivity (Cunha and Heckman 2007; Almond and Currie 2011).

1 The initiation and potential impact of US family planning programs

Margaret Sanger's zealous advocacy of what became known as "birth control" is often credited to her encounters with child poverty. Her work as a maternity nurse on the Lower East Side of New York City

3 Poverty rates in this article are defined using the official US measure.

took her to the residences of poor families with many children living in squalor. She also encountered women who died (or nearly died) from attempted abortions or debilitating contraceptive techniques.[4] The best medical recommendation of the day to prevent unwanted childbearing (as related in a letter to Sanger) was often to tell one's husband to "sleep on the roof."

1.1 The initiation of US family planning programs, 1964–1973

The introduction of the first oral contraceptive gave women and physicians much more reliable, safer, and enjoyable options. Its expense, however, prohibited many women from using it. Differences in access to "the Pill" led many to advocate for federal subsidies. Largely due to these efforts, federal grants for family planning began under the EOA (1964, Public Law 88-452), a key piece of President Johnson's War on Poverty.[5] Between 1965 and 1970, federal outlays for family planning through the OEO rose more than 20-fold, from 1.6 to 41 million (2008 dollars). This increase reflects two important sets of policy changes. The first was the 1967 Amendments to the EOA (Public Law 90-222, Title II, Section 222a), which designated family planning as a "national emphasis" program. The second was the increase in outlays under President Nixon, who became president in 1969. The November 1970 enactment of Title X of the Public Health Services Act allowed the Department of Health, Education, and Welfare (DHEW) to make grants to local organizations directly and prohibited the use of federal funds "in programs where abortion is a method of family planning" (§ 1008). After the enactment of Title X, federal outlays for family planning increased by another 50% by 1973.

Federally funded family planning programs provided access to birth control as well as related education and counseling services. These programs tended to open in locations whose residents had limited access to family planning services. In many locations, no program existed prior to the federal grant. In others, programs had existed but were much smaller in scale. Consequently, the federal grants significantly increased availability, reduced wait times, and increased the supply of free or low-cost

4 One letter to Margaret Sanger read, "I am the mother of two lovely little girls. I have been married fifteen years. I married at the age of fifteen to escape a home that was overcrowded with unloved and unwanted children, where there was never clothing or food enough to divide among the eight of us. . .I have been pregnant 15 times, most of the time doing things myself to get out of it and no one knows how I have suffered from the effect of it, but I would rather die than bring as many children into the world as my mother did and have nothing to offer them" (Sanger 1923).

5 According to 1967 estimates, expenditures for family planning through the Maternal and Child Health programs (started in 1942) and the Maternal and Infant Care programs under the 1963 Social Security Amendment were small (DHEW 1974, p. 3).

contraceptives in affected communities. Because federally funded programs did not require an explicit means test, they may have also reduced the costs of visits and supplies at private providers in the area.

Less is known, however, about these programs' day-to-day operations. In the 1960s, programs were subjected to little oversight from the federal government. Information was sparse on all federal programs, and officials rarely spoke about this largely taboo topic. In an evaluation of the War on Poverty, Sar Levitan (1969, p. 209) wrote that, "Contrary to the usual OEO tactic of trying to secure the maximum feasible visibility for all its activities, OEO prohibited [family planning] grantees from using program funds to "announce or promote through mass media the availability of the family planning program funded by this grant.""[6] The implication is that the treatment effect of these grants can be understood as one of increasing federal funding for "family planning," rather than the effect of a particular, homogeneous intervention.

Figure 1 presents the roll-out of the first federal family planning grants from 1965 to 1973. Counties that received federal grants in this period (shaded on map; we call these counties "funded") were more likely to be in cities and, consequently, differed in a number of their observable dimensions (Bailey 2012: Table 1). Data from the 1960 census indicates that roughly 60% of the US population of women ages 15–44 lived in funded counties. Funded counties were more urban, had more elderly residents, and were more educated and affluent than were unfunded counties. Interestingly, funded and unfunded counties had a similar share of residents under age 5 in 1960, suggesting little difference in fertility rates in these areas before the passage of the EOA. To account for time-invariant, area-level differences, our analysis includes area fixed effects.

The different shades of gray in Figure 1 represent variation in the timing of each county's first federal family planning grant. Counties in the lightest shade of gray first received grants between 1965 and 1967; counties in the next darkest shade of gray first received grants between 1968 and 1969; counties shaded in black first received grants from 1970 to 1973. Although counties in each of the lower 48 states (i.e., excluding Alaska and Hawaii) received grants, the timing of program start dates varied considerably within states: in 43 states, programs were first funded in at least two different years; counties in 41 states first received funding in at least four

[6] The fact that the OEO might fund birth control was contentious before the EOA passed. For instance, on 18 April 1964, Eve Edstrom in the *Washington Post* (p. A4) reported the controversy on this topic between Representative Phil M. Landrum (D-Ga.), the House sponsor of the EOA, and Republican members of the special House Education and Labor subcommittee.

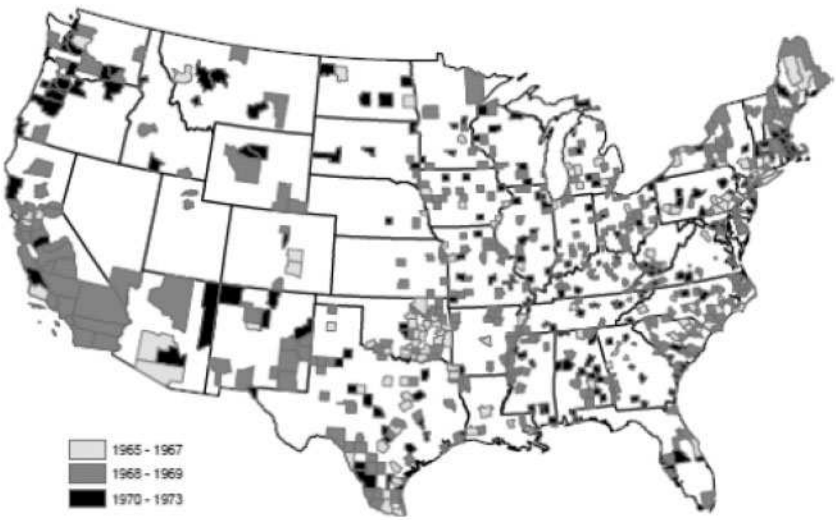


Figure 1 The date of the first federal family planning grant, 1965–1973.
Note: Dates are the year that the county first received a federal grant. Counties not receiving a family planning grant between 1965 and 1973, including communities that received funding but with an unknown starting date, are not shaded.
Source: NACAP, NAFO, and OEO (1969, 1971, and 1974).

different years; and, in more than half of all states, counties were first funded in at least five different years of the period considered.

1.2 The expected effects of family planning programs on outcomes

The potential effects of these family planning grants on children operate through several channels, each relating to their effects on fertility rates. By providing cheaper, more reliable contraception and more convenient services, family planning should reduce ill-timed and unwanted childbearing. Additionally, reductions in the price of averting births should increase the number of births that parents *choose* to avert or delay.[7] Standard economic models and related empirical work motivate the following expected relationships between family planning policies and poverty rates.

First, holding constant other uses of parents' time, fewer children in a household at a given point in time implies an increase in the availability of parental time and economic resources per child. Fewer children in a

[7] Potentially offsetting this effect is the fact that cheaper and more reliable contraception should reduce precautionary undershooting as well (Michael and Willis 1976). Estimates presented later suggest that reductions in childbearing have dominated empirically, so that greater access to cheaper and more reliable contraceptives tends to reduce family size.

household should mechanically reduce poverty rates as a family with a given income is less likely to fall below the poverty threshold.

Second, family planning programs may directly increase household income, thus reducing poverty rates. Cheaper and more reliable contraception reduces the immediate and expected costs of delaying childbearing, freeing up resources for investment in the parents' human capital. Delaying parenthood, even for just a year or two, could allow soon-to-be parents to get more education, work experience, and job training, and thus increase their lifetime earnings. The results of empirical studies of teen access to the birth control pill are consistent with the claim that delaying childbearing has value. Bailey et al. (2012) show that earlier access to the Pill increased women's investment in their careers and, ultimately, their wages. Hock (2008) shows that early access to the Pill increased men's educational attainment as well. Family planning also reduces the price of delaying marriage (Goldin and Katz 2002) and can improve spousal matching; thereby reducing subsequent divorce rates (Christensen 2011; Rotz 2011). However, delaying childbearing does not necessarily yield economic benefits for mothers. Hotz et al. (2005) show that women who became mothers as teenagers have slightly *higher* subsequent levels of employment and earnings than women of the same age who miscarried as teenagers.

Third, family planning programs may affect the composition of parents by benefitting the lower income population. Because higher income households could afford services at private medical providers, federally subsidized services may have disproportionately benefitted poorer families. Consistent with this claim, Torres and Forrest (1985) document that, in 1983, family planning programs served almost 5 million Americans. In the same year, roughly 83% of family planning patients had incomes below 150% of the poverty line, and 13% were recipients of Aid to Families with Dependent Children (AFDC, the principal cash welfare program at the time). Jaffe et al. (1973) report that 90% of all patients in organized family planning programs had household incomes of no more than 200% of the federal poverty line. If poorer families elected to postpone childbearing or have fewer children, children born following the introduction of the programs would enjoy, on average, greater economic resources.

Finally, parents' investments in children may also be complemented by decreases in children's cohort size. Smaller cohorts could increase the public resources available per child and decrease competition for these limited resources (Easterlin 1978). In schools, for instance, a decrease in cohort size might decrease class sizes, increase the likelihood of getting attention from teachers, and reduce classroom disruptions. Changes in cohort size are unlikely to be accommodated fully by universities, and a larger share of these smaller cohorts may be admitted to and complete college (Bound and Turner 2007). Smaller cohort sizes may also affect the scale of markets for illicit drugs and other social "bads" and thereby reduce the incidence of related crimes (Jacobson 2004). Finally, smaller cohorts may reduce aggregate labor supply, decrease workers' competition for firms' resources, increase capital–labor ratios, and tend to raise wages.

In summary, by increasing adults' pre-childbearing human capital and by benefitting lower income families, family planning programs may increase children's economic resources and decrease child poverty rates. Under standard quality–quantity formulations, these changes would tend to increase parental investment in their children (Becker and Lewis 1973). To the extent that family planning increases parental investment in children, it may improve their lifetime opportunities and labor market outcomes as adults. Cohort-size effects tend to reinforce the positive effects of family planning.

Note that these labor market channels—in addition to the within-household spillovers in family income and reductions in the price of child quality—suggest that the consequences of family planning may extend *beyond* the children immediately affected. Access to family planning may benefit slightly older or younger children in the affected households, children in unaffected households in the same cohort, and children in slightly older or younger cohorts in the same labor market. Because our research design compares the outcomes of children who were born in the years leading up to and just after the first funding for federal family planning programs began, this framework implicitly treats the older siblings of children born just before the family planning program as part of the comparison group. We expect, therefore, that our results understate the effects of family planning programs.

2 Data and research design

Our analysis integrates the approach of Gruber et al. (1999), who study the impact of legalizing abortion on children's economic resources, and Bailey (2012), who studies the impact of funding family planning programs on fertility rates. We use three separate datasets to document effects at different stages by race: Vital Statistics data on fertility rates by race; the 1980 decennial census which contains information on poverty rates among the affected cohorts in childhood; and a pooled sample of the 2000 decennial census and 2005-11 ACS which contains information on poverty rates among the affected cohorts in adulthood. Our data have been collapsed to birth year × area × year of observation cells, indexed as t, j, and c, respectively. Geographic area is defined either as a county (in the Vital Statistics data), county group of residence (in the 1980 decennial census), or a Public Use Microdata Area of residence (PUMA, in the 2000 census and 2005-11 ACS).

Our research design compares poverty outcomes in childhood and adulthood between cohorts born before and after their area of birth/residence was first funded within the following linear difference-in-differences specification,

$$Y_{j,t,c} = \tau \text{PostFP}_{j,t} + \boldsymbol{X}'_{j,t,c}\boldsymbol{\beta} + \theta_{j,c} + \gamma_{s(j),t} + \varepsilon_{j,t,c}, \quad (1)$$

where Y is a poverty rate and $\text{PostFP}_{j,t} = 1(t > T_j^*)$ is equal to 1 for areas observed after the first fiscal year family planning programs were funded (T_j^*).[8] Other covariates include either area × year fixed effects (in the 2000 census and 2005-11 ACS) or area fixed effects (Vital Statistics and 1980 decennial census), θ, to account for within year, area-level differences; a set of year fixed effects or state-by-birth-cohort fixed effects that capture changes in state policies such as the staggered legalization of abortion and the state-level roll-out of Medicaid, γ. $\boldsymbol{X}$ is a set of covariates which are discussed in later sections.

The estimates of interest, τ, capture the average change in outcomes between individuals whose mothers would have had access to a family planning program before childbirth and individuals in the same area whose mothers would have conceived them before federal family planning grants began. In all specifications, estimates are unweighted to minimize the importance of measurement error due to mobility (migration in and out of cities is much higher than in smaller areas). (See also Solon et al. 2013). Additionally, we present cluster-robust standard errors, which account for an arbitrary covariance structure within each area across birth years (Arellano 1987; Bertrand et al. 2004).

2.1 Support for key identifying assumptions

A central assumption of this article's research design is that the roll-out of family planning programs is unrelated to other determinants of childbearing or child outcomes. Evidence for this assumption comes from both historical accounts and quantitative evidence. According to oral histories, the "wild sort of grant-making operation" during the period provides a plausibly exogenous shock to the availability of local family planning services (Gillette 1996, p. 193). Bailey (2012) also provides quantitative support for this assumption. She shows that, although family planning programs were funded earlier in areas with greater urban populations,

8 For simplicity in our later exposition, we refer to the year family planning programs were funded as the date they began. The date of the first grant is not technically the date these clinics began operating, but the date of the first grant serves as a close proxy.

neither 1960 census characteristics, 1964 fertility levels, 1960–1964 fertility changes, nor a rich set of 1965 measures of sexual behavior, birth control use, and childbearing predict *when* federal family planning programs began. She also shows that the timing of the first family planning grant appears unrelated to changes in the funding for other War on Poverty programs.

Another key assumption underlying this article's empirical strategy is that federal funding of family planning meaningfully increased the *use* of family planning services in the affected areas. This assumption is difficult to test explicitly, but administrative reports suggest that the number of users of federally funded family planning services increased from 0 in 1965 to around 1.2 million in 1969 and nearly 5 million in 1983.

Further evidence of these programs' relevance comes from their relationship to reductions in local fertility rates. Bailey's main findings also support this claim. Before federal funding of family planning programs, the trend in the general fertility rate was similar in counties that would eventually receive funding and in those that would not (the pre-treatment differences are close to 0 and individually and jointly statistically insignificant). However, fertility rates fell sharply in the funded counties after the family planning grants began. Within 3 years of the grant, the general fertility rate had fallen by roughly 1 birth per 1,000 women of childbearing age in these counties on average. By years 6–10, it had fallen by an average of 1.5 births per 1,000 women. Fifteen years after an organization received its first federal family planning grant, the fertility rate in funded counties remained 1.4–2% lower than in the year of first grant receipt, net of declines in fertility in other counties in the same state and after adjusting for observable county-level characteristics. These findings are robust to variations in the specification: omitting unfunded counties, not weighting the regressions, and including county-level linear time trends. In addition, the effects are similar for programs funded before and after Title X began in 1970.

Using Vital Statistics birth certificate records that report mother's county of residence, we provide further evidence on the fertility effects of family planning grants by crude race categories consistently available in this period: white and non-white. Due to incomplete reporting of fertility rates by race in the early 1960s, our sample begins in 1968 with the natality microdata files (NCHS 2003). For our fertility analyses, we drop counties that received their first family planning grant before 1968, so our post-grant estimates capture changes in fertility rates for a consistent group of counties. Our overall sample, which aggregates across racial groups, includes 2,633 counties, 514 of which received a federal family planning grant (we call these "funded counties"). The subsample of these counties that allows disaggregation by race (white and non-white in this period) consists of 1,481 counties, 197 of which were funded. The Vital Statistics contain information on county of mother's residence for each birth, which makes it possible to compare the results for different estimators and samples.

In practice, θ_j in equation (1) consists of a set of county fixed effects, and X includes county covariates for the number of abortion providers, which account for within-state changes in the provision of abortion from 1970 to 1979 and annual information on per capita measures of government transfers from the Bureau of Economic Analysis Regional Information System (REIS 2009) (cash public assistance benefits such as AFDC, Supplemental Security Income, and General Assistance; medical spending such as Medicare and military health care; and cash retirement and disability payments). In addition, X includes 1960 county covariates interacted with a linear trend.[9] Finally, $\text{PostFP}_{j,t}$ is replaced with dummy variables for three birth cohort categories: cohorts born 5–1 years before the family planning program began; cohorts born 1–15 years after funding began, and cohorts born 16–20 years after funding began. The sample consists of a balanced set of counties, while the control group consists of the cohort born at the time of first grant in funded counties and all cohorts in unfunded counties. We report estimates of the effect of federally funded family planning on cohorts born 1–15 years after the family planning program was first funded.

Table 1 shows the relationship between family planning grants and fertility rates (τ) for all individuals (panel A), whites (panel B) and non-whites (panel C). Columns labeled (1) use a sample of all counties and include county, year, and state-by-year fixed effects; columns labeled (2) add county-level covariates to the samples in columns labeled (1). The results for all individuals suggest a relationship between family planning programs and fertility rates similar to those reported in Bailey (2012), even though programs funded before 1968 are dropped and the sample only covers years 1968–1988 (not 1959–1988). One to fifteen years after counties first received federal family planning funding, fertility rates remained 2.3 births lower per 1,000 women of childbearing age—a reduction of

9 The interactions of county covariates are identical to those in Almond et al. (2011) and include share of population in urban area, non-white, under age 5, over age 64; share of households with income under \$3000; and the share of the county's land that is rural or a farm. We are grateful to Doug Almond, Hilary Hoynes, and Diane Schanzenbach for providing the REIS data and to the Guttmacher Institute and Ted Joyce for providing the data on abortion providers. Because information on abortion providers is not available at the county level before 1973, we follow Joyce et al. (2013) in assuming the number of providers in 1970–1972 in states that legalized before Roe v. Wade are identical to the number observed in 1973. Note that changes in the distance to states providing legal abortion before 1970 are accounted for in the state-by-birth-year fixed effects.

Table 1 The effect of family planning on fertility rates, by race

	Dependent variable: fertility rate (births per 1,000 women ages 15–44)					
	A. All individuals		B. White		C. Non-white	
	(1)	(2)	(1)	(2)	(1)	(2)
Mean in funded counties before funding began	*90*	*90*	*83*	*83*	*122*	*122*
After family planning program funding began	–2.75	–2.26	–1.96	–1.73	–1.28	–1.72
	[0.43]	[0.40]	[0.47]	[0.45]	[1.63]	[1.63]
R^2	0.56	0.57	0.52	0.53	0.30	0.31
Counties	2,633	2,633	1,481	1,481	1,481	1,481
Observations	55,293	55,293	31,101	31,101	31,101	31,101
County FE	Yes	Yes	Yes	Yes	Yes	Yes
Birth year FE	Yes	Yes	Yes	Yes	Yes	Yes
State $\times$ birth year FE	Yes	Yes	Yes	Yes	Yes	Yes
County characteristics		Yes		Yes		Yes

Notes: The unit of observation is county by year, and estimates of τ are presented using equation (1). The results use the funded and unfunded sample of counties. Estimates are not weighted. Columns labeled (1) include county, year, and state-by-year fixed effects, while columns labeled (2) add county covariates (1960 county covariates interacted with a linear trend, number of abortion providers, and REIS controls). Panel A presents results for both races, panel B presents results for whites only, and panel C presents results for non-whites only. Heteroskedasticity-robust standard errors clustered by county are presented beneath each estimate in brackets. *Source*: Vital Statistics.

2.5% over the pre-program mean in funded counties and the overall mean for unfunded counties.[10]

Panels B and C of Table 1 present the relationship between family planning programs and fertility rates by race. For both whites and non-whites, the introduction of family planning is associated with declines in fertility rates. Using the column (2) specification, the white fertility rate was about 2.1% lower in the 15 years after first federal funding of family planning programs, and the non-white fertility rate was about 1.4% lower. For non-whites, however, these estimates are imprecise and not statistically different from zero.

In summary, these results support previous findings that the introduction of federally funded family planning programs—and the increase in the availability of family planning services they engendered—is associated with reduced fertility rates. Next, our analysis examines the relationship between family planning programs and child poverty.

3 Poverty rates among affected cohorts in childhood

We use measures of child poverty from the 5% 1980 Integrated Public Use Microdata Series (IPUMS, Ruggles et al. 2010) sample of the US decennial census. These data have several advantages for the purposes of our analysis. First, they provide large sample sizes and allow us to compute for each area and birth cohort and race the share of children in families below 100% and 200% of the poverty line. A second advantage is that information on county group in the 1980 census (the lowest level of geographic identification in the IPUMS files) allows us to link the location of family planning programs to individuals in areas smaller than states.[11]

These data, however, also have limitations for the purposes of this analysis, because they only provide geographic information at the county group level. County groups in the continental US are typically contiguous agglomerations of counties, but some counties are split between different county groups or are non-contiguous. This limits our ability to link covariates to county groups and match them to family planning grant information. For this reason, we restrict our sample to county groups that consist only of contiguous counties and that do not contain split counties.

[10] Restricting the sample to funded counties only, however, reduces the magnitudes of these estimates and they become statistically insignificant. Although the estimates remain negative, they are a fraction of the size in Table 1A, which suggests that using funded only counties (as we do in subsequent analyses) may understate the overall impact of the program.

[11] We link *county-level* introduction of family planning to census *county groups* using a cross-walk generously provided by Elizabeth Cascio.

Ongoing work by Bailey et al. (2013) uses the 1970 and 1980 restricted census samples that consist of 16% and 20% samples of the population and include the county of residence information. This allows them to provide more precise estimates of the effect of first family planning program grants and to link all households to family planning grants based on their county of residence.

A further limitation of the geographic information in the public files is that county group at the time of the census may not accurately measure mothers' county group around the time of conception. This source of measurement error is empirically important: Bailey et al. (2013) find that migration-induced measurement error in access to family planning is greater in cities and increases in funded areas (relative to unfunded areas) after the first federal family planning grant. They demonstrate that using unweighted regressions and limiting the sample to funded areas generates similar implied reductions in fertility rates in the census as in the Vital Statistics data (compare to this article's Table 1) as a result of family planning program funding. To reduce measurement error in access to family planning in our analysis, we also use unweighted regressions and limit the sample to funded county groups. Out of 1,154 overall county groups, our final sample consists of 251 county groups that do not contain split or non-coterminous counties and that receive their first federal family planning funding at some point before 1974. Of these county groups, only 154 have sufficient observations on non-whites for inclusion.[12]

The final limitation of the 1980 IPUMS census derives from the fact that the unit of observation is a household. The census does not measure outcomes of children not *residing* with their parents. Because children often leave home around age 18, we limit our analysis to individuals under age 18, or birth cohorts born from 1963 to 1979. The practical implication of this limitation is that our pre-trend in the 1980 census is very short and begins only 2 years before the first family planning grant.

The data available in the 1980 public census files necessitate that we estimate a restricted version of equation (1). Only one census year is used, so c is 1980 for all individuals, and θ_j is a set of county group fixed effects. $\boldsymbol{X}$ includes county group covariates for the number of abortion providers and annual information on per capita measures of government transfers from REIS (cash public assistance benefits such as Aid to Families with Dependent Children, Supplemental Security Income, and General Assistance; medical spending such as Medicare and military health care;

[12] We also exclude Virginia from the analysis, because so many of its counties changed boundaries over the 1970s making it difficult to merge county groups with appropriate covariates.

and cash retirement and disability payments). Finally, $\text{PostFP}_{j,t}$ is replaced with dummy variables for three birth cohort categories: cohorts born 10–3 years before the family planning program began and cohorts born 1–6 years and 7–14 years after the family planning program began. The comparison group in this analysis is the cohort born in event years -2 to 0, which is observed for all county groups in the analysis. We report coefficients for the 1–6 years post-funding category, because they are based on a balanced set of county groups.

Access to affordable family planning may lead to lower poverty rates by permitting families to adjust their childbearing decisions in a way that raises their family income. Table 1 shows that family planning grants allowed women to defer childbearing. As we discussed previously, the share of children in poverty may decrease following the introduction of a family planning program due to smaller family sizes, parents' accumulation of more human capital, work experience, higher earning mates, or a change in the income composition of parents.

Table 2 presents the estimated relationship between funding for family planning and child poverty rates. Panel A shows the share of children living in families below the poverty line and panel B shows the share of children living in families below twice the poverty line. The results suggest that children born after family planning programs were funded were less likely to live in poverty. Children born 1–6 years after funding were 0.76 percentage points less likely to live in poverty than the children born before the federal funding began—a reduction of 4.2% (from a mean poverty rate of 18.2% for children born 0–2 years before funding began). These results are robust across specifications that include county group, year and state-by-year fixed effects (column (1)) and the addition of county group level controls (column (2)).

Federal family planning programs expanded access to and affordability of family planning particularly to disadvantaged individuals. Whether white or non-white children experienced greater reductions in poverty depends on how family planning influenced parents' use of their services and also how parents using these services changed their economic circumstances. Different relationships between family planning and poverty rates by race may also result from differences in access to education, job training, or spousal matching for mothers, for instance. To examine these differences, we perform our analysis by crude categories for race to correspond to those categories available in the Vital Statistics data on births. Although both white and non-white children were significantly less likely to live in poverty, the reduction was largest among non-white children. Column (3) shows that white children are 0.56 percentage points less likely to live in poverty, a reduction of 4.1% from a mean of 13.7%.

Table 2 The effect of family planning on next generation childhood poverty, by race

	All individuals		White	Non-white
	(1)	(2)	(3)	(4)
A. Dependent variable: percent with family income < poverty line				
Mean in funded counties before funding began	*18.2*	*18.2*	*13.7*	*38.7*
After family planning program funding began	–0.81	–0.76	–0.56	–3.16
	[0.31]	[0.32]	[0.30]	[1.22]
R^2	0.27	0.28	0.26	0.30
B. Dependent variable: percent with family income < two times the poverty line				
Mean in funded counties before funding began	*42.9*	*42.9*	*37.0*	*69.7*
After family planning program funding began	–0.45	–0.50	–0.40	–2.09
	[0.41]	[0.42]	[0.44]	[1.16]
R^2	0.47	0.47	0.47	0.31
County group FE	Yes	Yes	Yes	Yes
Birth year FE	Yes	Yes	Yes	Yes
State × birth year FE	Yes	Yes	Yes	Yes
County characteristics		Yes	Yes	Yes
County groups	251	251	251	154
Observations	4,267	4,267	4,267	2,618

Notes: The unit of observation is county group by year, and estimates of τ are presented using equation (1). The results use the funded only sample. We classify as 'white' all individuals in the census who list their race as 'white', while 'non-white' comprises all other individuals. We drop county groups where fewer than fifty non-white children were born in any year in the analysis. We drop non-coterminous county groups and county groups that contain split counties. We define the share in poverty as the share of children who live in families whose income is below the poverty threshold, we also compute the share of children who live in families whose income is below 200% of the poverty threshold. Column (1) presents results for both races and includes county group, birth year, and state-by-birth-year fixed effects; column (2) adds county characteristics (number of abortion providers and REIS controls) to column (1); column (3) presents results for whites only and includes county group, birth year, state-by-birth-year fixed effects, and county characteristics; column (4) presents results for non-whites only and adds the same controls as column (3). Panel A presents results when using the share of children living in families whose income is below 100% of the poverty line as a dependent variable. Panel B presents results when using the share of children living in families whose income is below twice the poverty line as the dependent variable. Estimates are not weighted. Heteroskedasticity-robust standard errors clustered by county are presented beneath each estimate in brackets. *Source*: 1980 Integrated Public Use Microdata Series.

Column (4) shows that non-white children are 3.2 percentage points less likely to live in poverty, a reduction of 8.3% from a mean of 38.7%.

A second (and related) hypothesis is that family planning programs would affect more disadvantaged families more, because they are substantially more likely to gain from access to affordable contraception. Consistent with this hypothesis, the relative reductions in the share of children below two times the poverty line are generally smaller than the reductions in the share of children living below the poverty line. Family planning programs are associated with a reduction in the share of children living near poverty, particularly among non-white children. Panel B shows that the share of children below two times the poverty line also fell. The relative reductions for all, white and non-white children are smaller than the reductions in the share of children living in poverty and the estimates are no longer statistically significant. Compared to white children, the reduction in the share of non-white children living near poverty is both absolutely and relatively larger. Non-white children born after family planning programs began were 3.0% less likely to live below two times the poverty line while white children were 1.1% less likely to live below two times the poverty line.

4 Poverty rates among affected cohorts in adulthood

A final analysis investigates the long-run relationship between a mother's access to family planning services and the adult outcomes of the affected children. Children born after the funding of family planning programs may have been part of smaller families and cohorts, were less likely to grow up in poverty, and, consequently, may have benefitted from greater parental and societal investments. The accumulation of these changes in childhood circumstances suggests these cohorts may have been less likely to live in poverty as adults.

We use the 5%, public use sample of the 2000 decennial census and the 2005-11 ACS (Ruggles et al. 2010) to investigate this hypothesis. An advantage of these data for the purposes of our analysis is that they allow the inclusion of a long pre-trend of cohorts, as information on poverty status exists even if individuals do not live with their parents. Our sample, therefore, includes individuals born from 1946 to 1980 who were ages 20–59 when observed. We choose these age limits to capture the labor market outcomes of workers after they have left home and before they have retired.

A disadvantage of these data is that they do not contain information on the county in which individuals were born. As in the analysis of the 1980 IPUMS data, we proxy for county of birth using the Public Use Microdata

Area (PUMA) of residence at the time of observation.[13] The role of misclassification error induced by this data limitation is difficult to assess without national data on lifetime migration. In the absence of systematic changes in migration, we expect that misclassification error introduced by using PUMA of residence should tend to work against finding results. On the other hand, using PUMAs rather than counties for longer-term outcomes may reduce misclassification error if, for instance, using a slightly larger area improves the assignment of mothers' access to family planning (i.e., more of the individuals remain in the PUMA of birth than lived in their county of birth). As in the analysis of the 1980 census, we estimate unweighted regressions and include only the 1,269 PUMAs that received a family planning grant before 1974 to limit the role of misclassification error.[14]

Our specification of equation (1) is similar to the analysis using 1980 IPUMS data with several exceptions. First, we use multiple survey years, so c equals 2000, 2005, 2006, ..., 2011. Pooling multiple years yields observations on the same cohorts at different ages, so we include age and age squared as covariates in $\boldsymbol{X}$. Second, due to the difficulty of mapping county characteristics onto PUMAs, we cannot include other covariates in the analysis. Third, $\text{PostFP}_{j,t}$ is replaced with dummy variables for three birth cohort categories: cohorts born 27–14 years before family planning programs began; cohorts born 1–7 years and cohorts born 8–15 years after family planning programs began. We omit cohorts born 13–0 years before family planning programs began, so this category becomes our comparison group. Estimates for the first and last categories are suppressed in the presentation in Table 3, because they are estimated using only a subset of cohorts.

Table 3 shows that within cohort changes in funding of federal family planning programs are associated with significant reductions in adult poverty rates among cohorts born after the programs began.[15] Many individuals in cohorts born before first funding of family planning programs transitioned out of poverty between childhood and adulthood: 18% of these cohorts lived in poverty in childhood, while 12% lived in poverty in adulthood. We provide evidence that this transition was significantly

13 PUMAS are the finest consistent geographic detail available for all individuals in the publically available versions of these data. There are 2,069 distinct PUMAs, each with a population of 100,000 or more, and, unlike county groups, PUMAs do not cross state borders.

14 Some PUMAs overlap multiple counties. The count of PUMAs that contain funded programs exceeds that of counties because we treat each PUMA that overlaps with a funded county as having received a family planning grant in the same year as the county.

15 We borrow from the US census the definition of poverty that uses a family income threshold that depends on the number of overall family members and the number of children (Dalaker and Proctor 2000). For instance, the poverty threshold for the annual income of a household of four is $23,550 in 2013 dollars.

Table 3 The effect of family planning on next generation adult poverty, by race

	All Individuals		White	Non-white
	(1)	(2)	(3)	(4)
A. Dependent variable: percent with family income < poverty line				
Mean in funded counties before funding began	*11.5*	*11.5*	*8.18*	*16.4*
After family planning program funding began	–0.28	–0.28	–0.50	–0.32
	[0.12]	[0.18]	[0.14]	[0.28]
R^2	0.05	0.06	0.02	0.03
B. Dependent variable: percent with family income < two times the poverty line				
Mean in funded counties before funding began	*27.9*	*27.9*	*20.4*	*38.1*
After family planning program funding began	–0.68	–0.68	–0.97	–0.76
	[0.18]	[0.18]	[0.21]	[0.34]
R^2	0.13	0.13	0.06	0.05
PUMA $\times$ observation year FE	Yes	Yes	Yes	Yes
Birth year FE	Yes	Yes	Yes	Yes
State $\times$ birth year FE	Yes	Yes	Yes	Yes
Age and age^2		Yes	Yes	Yes
PUMAs	1,268	1,268	1,268	1,268
Observations	328,403	328,403	320,634	298,216

Notes: We classify as 'white' all individuals recorded in the census as belonging to no other racial group and not being Hispanic, while 'non-white' comprises all other individuals. There were 2,072 PUMAs in the fifty US states in 2000. Following population displacement in Louisiana due to Hurricane Katrina, three PUMAs (1801, 1802, and 1905) were combined, and we merge these PUMAs together throughout the entire 2000–2011 sample period. Additionally, we drop PUMA 5423 in Los Angeles because it has few white residents, for none of whom poverty status is recorded. Our final sample consists of 1,268 PUMAs whose boundaries include all or part of county in which a family planning grant began between 1965 and 1973 and in which poverty status was measured for at least one white and at least one non-white resident age 20–59 and born 1946–1980 in each of the 8 years of observation (yielding 10,144 unique combinations of PUMA $\times$ year of observation). This figure of 1,268 PUMAs exceeds the tally of 654 counties with a grant because, while a single PUMA may span several counties, so too may a single county span several PUMAs. Finally, we average poverty status across all individuals, and separately by race for those who reside in the same PUMA, share the same year of birth, and are observed in the same year. The units of analysis are 328,403 PUMA $\times$ year of birth $\times$ year of observation cells. Not every cell contains both white and non-white individuals for whom poverty status is recorded, so the actual number of units is slightly smaller for the race-specific specifications (3) and (4). Heteroskedasticity-robust standard errors clustered by PUMA and observation year are presented beneath each estimate in brackets. The mean in funded counties before funding began is the average across individuals born 2 years prior to funding to those born in the year of funding. Estimates are not weighted. *Source*: 2000 US Decennial Census and 2005–2011 ACS.

greater among cohorts born after family planning programs began. Table 3 shows that the share of adults in poverty (panel A) and the share of adults with family income below two times the poverty line (panel B) fell significantly for the affected cohorts. Relative to individuals born in the years prior to when family planning programs began, individuals born in the seven subsequent years were 0.28 percentage points less likely to live in poverty as adults, a reduction of 2.4% over the pre-program mean of 11.5%. This result is unaltered with the inclusion of age and age-squared controls in column (2).

Following our analysis of child poverty, we also examine reductions in near poverty. The effect of funding family planning programs on the share of adults living near poverty is similar to the effect on the share of adults living in poverty. Panel B of Table 3 shows that cohorts born after family planning programs were funded were 2.4% less likely to live below two times the poverty line as adults, relative to cohorts born before funding began but residing in the same PUMA. In addition, we find that the mean long-run effects are slightly stronger (though not statistically so) among whites. White cohorts born after the introduction of family planning were 4.8% (0.97 percentage points) less likely to live below two times the poverty line. The same statistic was 2% among non-white cohorts. This striking relationship between family planning programs and poverty rates decades later suggests that family planning programs may reduce poverty rates, both in the short and longer-term.

5 Conclusions

In 2012, approximately one in five US children lived below the official poverty line, only slightly lower than in 1965. The persistence of child poverty and its potentially negative consequences for children's opportunities has made reducing child poverty a public policy concern. While the majority of Americans have higher incomes than their parents, children with parents in the lowest income quintile experience the lowest absolute increase in income through adulthood (Pew Charitable Trusts 2012). In fact, 43% of all children and 50% of black children with parents in the bottom income quintile remain in the bottom income quintile as adults.

Our findings suggest the potential of family planning programs to disrupt this cycle of disadvantage. Individuals born after family planning programs began were 4.2% less likely to live in poverty in childhood and were 2.4% less likely to live in poverty as adults, than individuals born just before family planning programs began and residing in the same location.

A simple calculation relies on our estimates to approximate some of the costs and benefits of spending on family planning programs. On the benefit side, we multiply the number of children in funded county groups in 1980 who were born after family planning programs were funded by our estimate in Table 2 in panel A of column (2). This calculation implies that 79,800 fewer children (0.0076×10.5 million) lived below the poverty line in 1980 than would have in the absence of the program. To approximate the number of adults who escaped poverty as a result of these programs, we multiply the number of adults ages 20–59 living in funded PUMAs in 2000 who were born after program funding by the coefficient in Table 3 in panel A of column (2) which yields 46,760 adults (0.0028×16.7 million). Between 1964 and 1973, the federal government spent approximately $2.6 billion (in 2010 dollars) on family planning grants. This implies that the per-child reduction in poverty cost approximately $32,581, while the long-run cost of each adult lifted out of poverty was $55,603.

Of course, these calculations likely misstate the effects of family planning for several reasons. First, siblings and slightly older and younger cohorts may also benefit from the programs and they contaminate the comparison group. Second, the mismeasurement of family planning status of parents (due to migration) should lead us to misstate the relationship of interest, and understate it if measurement error is unrelated to access to family planning. Finally, using only changes in poverty rates ignores many of the other consequences of family planning programs, which extend to population growth and labor supply, higher education, labor force participation, and wages (Bailey 2013). Nevertheless, even these conservative estimates of the cost per child or adult exiting poverty suggest that family planning programs could improve economic outcomes over the longer-term.

Funding

The collection of data on US family planning programs was supported by the University of Michigan's National Poverty Center (NPC) and Robert Wood Johnson Health and Society Programs; the University of Michigan Population Studies Research Center's Eva Mueller Award; and the National Institute of Health (HD058065-01A1), and the University of Kentucky Center for Poverty Research (UKCPR) through the US Department of Health and Human Services, Office of the Assistant Secretary for Planning and Evaluation (5 UO1 PE000002-05). Work on various aspects of this project was generously supported by the Small Grants Program at the University of Michigan's National Poverty Center, the University of California-Davis Center for Poverty Research (1H79AE000100-1 from the US Department of Health and Human

Services, Office of the Assistant Secretary for Planning and Analysis, which was awarded by the Substance Abuse and Mental Health Services Administration), and the Elizabeth Caroline Crosby Fund and Rackham Graduate School at the University of Michigan. During work on this project, Malkova was supported by the Rackham Merit Fellowship funded by the University of Michigan and Norling was supported by the NICHD (T32 HD007339) as a University of Michigan Population Studies Center (PSC) Trainee. We gratefully acknowledge the use of the services and facilities of the Population Studies Center at the University of Michigan (funded by NICHD Center Grant R24 HD041028). The opinions and conclusions expressed herein are solely those of the authors and should not be construed as representing the opinions or policy of any of these funders or any agency of the federal government.

Acknowledgements

We are grateful to Doug Almond, Hillary Hoynes, and Diane Schanzenbach for sharing the Regional Economic Information System (REIS) data for the 1959–1978 period and to the Guttmacher Institute and Ted Joyce for sharing information on abortion providers from 1973 to 1979. We are grateful for comments and insights on early versions of this article from Brad Hershbein, Melanie Guldi, David Lam, Zoë McLaren, and participants at the CESifo Economic Studies and UCLS Conference on Families, children, and human capital formation in Munich, Germany. Outstanding research assistance was provided by Anna Erickson.

References

Almond, D., H. W. Hoynes and D. W. Schanzenbach (2011), “Inside the War on Poverty: The Impact of Food Stamps on Birth Outcomes”, *Research in Economics and Statistics* **93**, 387–403.

Almond, D. and J. Currie (2011), “Human Capital Development Before Age Five”, in O. Ashenfelter and D. Card, eds., *Handbook of Labor Economics*, vol. 4B, Maryland Heights, MO, Elsevier.

Arellano, M. (1987), “Computing Robust Standard Errors for Within-Group Estimators”, *Oxford Bulletin of Economics and Statistics* **49**, 431–4.

Bailey, M. J. (2012), “Reevaluating the Impact of U.S. Family Planning Programs on Fertility: Evidence from the War on Poverty and the Early Years of Title X”, *American Economic Journal: Applied Economics* **4**, 62–97.

Bailey, M. J. and S. Dynarski (2011), "Inequality in College Entry and Completion", in G. J. Duncan and R. J. Murnane, eds., *Whither Opportunity? Rising Inequality and the Uncertain Life Chances of Low-Income Children*, Russell Sage, New York, New York.

Bailey, M. J., B. J. Hershbein and A. Miller (2012), "The Opt-In Revolution? Contraception, Fertility Timing and the Gender Gap in Wages", *American Economic Journal: Applied Economics* **4**, 225–54.

Bailey, M. J., O. Malkova and Z. McLaren (2013), "Does Family Planning Increase Children's Opportunities?".

Bailey, M. J. (2013), "Fifty Years of Family Planning: New Evidence on the Long-Run Effects of Increasing Access to Contraception", *Brookings Papers on Economic Activity*, Spring.

Becker, G. S. and H. G. Lewis (1973), "On the Interaction Between the Quantity and Quality of Children", *Journal of Political Economy* **81**, S279–88.

Becker, G. S. and N. Tomes (1979), "An Equilibrium Theory of the Distribution of Income and Intergenerational Mobility", *Journal of Political Economy* **87**, 1153–89.

Becker, G. S. and N. Tomes (1986), "Human Capital and the Rise and Fall of Families", *Journal of Labor Economics* **4**, S1–39.

Bertrand, M., E. Duflo and S. Mullainathan (2004), "How Much Should We Trust Differences-in-Differences Estimates?", *Quarterly Journal of Economics* **119**, 249–75.

Bound, J. and S. E. Turner (2007), "Cohort Crowding: How Resources Affect Collegiate Attainment", *Journal of Public Economics* **91**, 877–99.

Christensen, F. (2011), "The Pill and Partnerships: The Impact of the Birth Control Pill on Cohabitation", *Journal of Populations Economics* **25**, 29–52.

Cunha, F. and J. Heckman (2007), "The Technology of Skill Formation", *American Economic Review* **97**, 31–47.

Dalaker, J. and B. D. Proctor (2000), *Poverty in the United States: 1999*, U.S. Census Bureau, Current Population Reports, Series P-60, No. 210.

Department of Health, Education and Welfare (DHEW), Center for Family Planning Program Development (Planned Parenthood-World Population) (1974), *Family Planning, Contraception, and Voluntary Sterilization: An Analysis of Laws and Policies in the United States, each State and Jurisdiction (as of September 1971): A Report of the National Center for Family Planning Services, Health Services and Mental Health Administration*, publication no. 74-16001, Rockville, MD, U.S. DHEW.

Easterlin, R. (1978), "What Will 1984 Be Like? Socioeconomic Implications of Recent Twists in Age Structure", *Demography* **15**, 397–421.

Edstrom, E. (1964), "Anti-Poverty Bill Doesn't Extend to Birth Control, Landrum Says: Monsignor Testified", *Washington Post*, April 18, p. A4.

Gillette, M. L. (1996), *Launching the War on Poverty: An Oral History*, Twayne Publishers, New York, N.Y.

Goldin, C. and L. Katz (2002), "The Power of the Pill: Oral Contraceptives and Women's Career and Marriage Decisions", *Journal of Political Economy* **110**, 730–70.

Gruber, J., P. Levine and D. Staiger (1999), "Abortion Legalization and Child Living Circumstances: Who is the "Marginal Child?", *Quarterly Journal of Economics* **114**, 263–91.

Hock, H. (2008), "The Pill and the Educational Attainment of American Women and Men", Florida State University Working Paper.

Hotz, V. J., S. W. McElroy and S. G. Sanders (2005), "Teenage Childbearing and Its Life Cycle Consequences: Exploiting a Natural Experiment", *Journal of Human Resources* **40**, 683–715.

Jacobson, M. (2004), "Baby Booms and Drug Busts: Trends in Youth Drug Use in the United States", *Quarterly Journal of Economics* **119**, 1481–512.

Jaffe, F. S., J. G. Dryfoos and M. Corey (1973), "Organized Family Planning Programs in the United States: 1968-1972", *Family Planning Perspectives* **5**, 73–9.

Joyce, T., R. Tan and Y. Zhang (2013), "Abortion Before & After *Roe*", *Journal of Health Economics* **32**, 804–15.

Levine, P. B. and D. J. Zimmerman (eds.) (2010), *Targeting Investments in Children: Fighting Poverty When Resources are Limited*, University of Chicago Press, Chicago.

Levitan, S. A. (1969), *The Great Society's Poor Law: A New Approach to Poverty*, Johns Hopkins Press, Baltimore.

Malthus, T. R. (1798), *An Essay on the Principle of Population*, John Murray, London.

Michael, R. T. and R. J. Willis (1976), "Contraception and Fertility: Household Production under Uncertainty", in: Household Production and Consumption, *NBER Chapters*, pp. 25–98. National Bureau of Economic Research, Inc.

National Archives Records about Community Action Program Grants and Grantees (NACAP) [Electronic Records]; Grantee Organization

Masterfile, 7/1/1964 - 6/30/1971; Grantee Organization Master File, 7/1/1971 - 9/30/1981; Funded Program Account Master File, 7/1/1964 - 6/30/1971; Funded Program Account Master File, 7/1/1971 - 9/30/1977; Funded Program Account Master File, 10/1/1977 - 9/30/1981; Record Group 381 (National Archives at College Park, College Park, MD).

National Archives Records, Federal Outlays System Files, FY1968-FY1980 (599052). (NAFO) [Electronic Record]; Records of the Community Services Administration, Record Group 381 (National Archives at College Park, College Park, MD).

National Center for Health Statistics (NCHS), Natality Detail File, 1968-1988: [UNITED STATES] [Computer file]. U.S. Dept. of Health and Human Services, Hyattsville, MD. National Center for Health Statistics [producer], 1981. Inter-university Consortium for Political and Social Research [distributor], Ann Arbor, MI.

Nixon, R. (1969), "Special Message to the Congress on Problems of Population Growth", *The American Presidency Project*, www.presidency.ucsb.edu/ws/?pid=2132 (last accessed 13 November 2013).

Office of Economic Opportunity (1969), *The Need for Subsidized Family Planning Services: United States, Each State and County, 1968*, Government Printing Office, Washington, D.C.

Office of Economic Opportunity (1971), *The Need for Subsidized Family Planning Services: United States, Each State and County, 1969*, Government Printing Office, Washington, D.C.

Office of Economic Opportunity (1974), *The Need for Subsidized Family Planning Services: United States, Each State and County, 1971*, Government Printing Office, Washington, D.C.

Pew Charitable Trusts, "Pursuing the American Dream: Economic Mobility across Generations", Washington and Philadelphia, www.pewstates.org/uploadedFiles/PCS_Assets/2012/Pursuing_American_Dream.pdf (last accessed 13 November 2013).

Reardon, Sean. (2011), "The Widening Academic Achievement Gap between the Rich and the Poor: New Evidence and Possible Explanations", in G. J. Duncan and R. J. Murnane, eds., *Whither Opportunity? Rising Inequality and the Uncertain Life Chances of Low-Income Children*, Russell Sage Foundation, New York, New York.

Regional Economic Information System (REIS), 1969-1989 (2009), *Bureau of Economic Analysis*, U.S. Department of Commerce, CIESIN, http://www.ciesin.org/datasets/reis/reis-home.html (last accessed 12 February 2014).

Rotz, D. (2011), "Why Have Divorce Rates Fallen? The Role of Women's Age at Marriage", Social Science Research Network Working Paper.

Ruggles, S., J. T. Alexander, K. Genadek, R. Goeken, M. B. Schroeder and M. Sobek (2010), *Integrated Public Use Microdata Series: Version 5.0 [Machine-readable Database]*, University of Minnesota, Minneapolis.

Saez, E. (2013), "Striking it Richer: The Evolution of Top Incomes in the United States (Updated with 2012 Preliminary Estimates)", University of California-Berkeley Working Paper, http://elsa.berkeley.edu/~saez/saez-UStopincomes-2012.pdf (last accessed 19 January 2014).

Sanger, M. (1923), "Prevention or Abortion—Which? Letters Showing the Dilemma Faced by Many Mothers", *Birth Control Review* **7**.

Solon, G., S. J. Haider and J. M. Wooldridge (2013), "What are We Weighting For?", Mimeo Michigan State University, https://www.msu.edu/~solon/WAWWFfeb2013.pdf (last accessed 19 January 2013).

Tone, A. (2001), *Devices and Desires: A History of Contraceptives in America*, Hill and Wang, New York.

Torres, A. and J. D. Forrest (1985), "Family Planning Clinic Services in the United States, 1983", *Family Planning Perspectives* **17**, 30–5.

Willis, R. J. (1973), "A New Approach to the Economic Theory of Fertility Behavior", *Journal of Political Economy* **81**, S14–64.

Parental Unemployment and Child Health[*]

by

Eva Mörk, Anna Sjögren, Helena Svaleryd[+]

2013-10-29

Abstract

We combine Swedish hospitalization data for 1992-2007 with register data on unemployment and analyze to what extent health outcomes of Swedish children, measured as overnight stays at hospitals, are worse among children whose parents are or become unemployed. In order to assess the extent to which parents who become unemployed are not a random sample of parents, we use an individual fixed effects approach. We find that the children of unemployed parents are much more likely to be hospitalized than other children. Although much of the difference is driven by selection our results suggest that unemployment matters even if it is difficult to establish causality. We also find that the impact of unemployment on child health varies across child age and gender as well as if it is the mother or the father who is unemployed.

Keywords: Parental unemployment, child health, human capital
JEL-codes: I12, J13

[*]This research has benefitted greatly from comments and discussion with Marcus Eliasson, Katarina Huttunen, Rudolf Winter-Ebmer, Björn Öckert and seminar participants at the ELE workshop in Faro, CESifo in Munich and IFAU, The authors are grateful for financial support from The Swedish Research Council.
[+]Eva Mörk, Uppsala University, UCLS, UCFS, CESifo, IZA and IEB eva.mork@nek.uu.se; Anna Sjögren, IFAU and UCLS, anna.sjogren@ifau.uu.se; Helena Svaleryd, Uppsala University, UCLS, UCFS helena.svaleryd@nek.uu.se.

1 Introduction

Many children are exposed to parental unemployment during childhood. For example, Lovell and Isaacs (2010) note that one out of nine American children has an unemployed parent as a result of the current recession. There are reasons to believe that children of unemployed parents fare worse than other children with respect to a number of different outcomes. For example, UNICEF recently reported that the fraction of deprived children in jobless households in rich countries is on average five times higher than the corresponding fraction in the population as a whole[1] Part of this deprivation is likely to be related to underlying factors that increase the likelihood of parents to become unemployed as well as lead to poor child health. However, it is also possible that unemployment per se may worsen family conditions affecting child health. Understanding the role of parental unemployment in shaping the human capital and well-being of children is important not only in order to estimate the full societal costs of unemployment, but also to guide the formulation of adequate human capital policies aiming to bridge and prevent permanent consequences of childhood disadvantage. Also, recent research point to the importance of shocks early in life for cognitive development and later success on the labor market (see e.g. Cunha and Heckman, 2007, 2008; Almond and Currie, 2011). Parental unemployment is likely to make up such a shock given that it may lead to reduced financial resources as well as social exclusion. Therefore it is important to understand how parental unemployment affects the children.

There are a number of studies that investigate the correlation between parental unemployment and child outcomes. Christoffersen (2000) using Danish data finds that children hospitalized for abuse more often have unemployed parents than other children. Christoffersen (1994) interviewed a sample of children of long term unemployed parents at age 25 and finds that these persons were more likely to have vocational training, being unemployed and to suffer from psychological problems. Similar patterns are found on UK data, see, e.g. Madge (1983). Pedersen et al (2005), using survey data from the Nordic countries find that children in families with at least one parent without paid work fare worse when it comes to chronical illnesses and psychosomatic symptoms, but do not use more prescribed medicine. They also show that controlling for the family's financial conditions only slightly reduce the associations between

[1] According to the UNICEF deprivation index a child is deprived if it lacks 2 or more of fourteen listed items including three meals per day, books in the home, etc.

parental unemployment and children's health outcomes. Also relying on survey data, Ström (2002) finds a positive correlation between parental unemployment and child accidents.[2]

Although analyzing interesting correlations, these studies tell us little about the causal effect of unemployment on child outcome, since unemployment does not hit workers randomly. In absence of any exogenous variation in unemployment, a large literature has been analyzing effects plant closures on outcomes for those losing their job. This literature has found that plant closures have negative consequences for worker's health, mental well-being, economic status and marriage stability, all of which influence the parents' capacity to invest in and care for the well-being and human capital of their children (Jacobsen, Lalonde and Sullivan, 1993, Stevens, 1997, Sullivan and von Wachter, 2009, Lalive and Zweimuller, 2007, Eliason and Storrie, 2009, and Eliason 2011b). Also, evidence from the plant-closure literature suggest intergenerational consequences of parental job-loss on long run outcomes such as earnings and employment for disadvantaged families (Page, Huff and Lindo, 2007; Oreopoulos, Page and Stevens, 2008). There are also a few studies which find immediate effects on children's educational outcomes of parents' experience of parental job-loss (Coelli, 2010; Stevens and Schaller, 2010; and Rege, Telle and Voturba, 2009). We will return to those studies in the next section.

Since a large proportion of the job-losers are likely to find a new employment relatively soon, these studies do not capture the effects of unemployment, although unemployment is likely to be one of the consequences of job loss. The purpose of this paper is to further analyze the relationship between parental unemployment and child health outcomes. Using register data, we analyze the overall health disadvantage of children exposed to parental unemployment. Besides measuring the relationship between parental unemployment and child health outcomes, we will also explore the panel dimension of our data and estimate models with children specific fixed effects. Doing this, we argue that we are able to control for selection into unemployment, i.e. that some parents are more likely both to become unemployed and to have children with bad health.

[2] See also the references within these studies for more correlation studies.

We also explore differences depending on child age and sex and depending on whether it is the mother or the father who is unemployed. An important contribution is to assess the extent to which this disadvantage is due to selection. This is particularly important when designing policies to address childhood disadvantage since the degree of selection will be informative on whether adequate policies should focus on reducing unemployment or alleviating the immediate negative consequences of unemployment or be directed towards improving the situation for children in vulnerable environments.

We combine Swedish hospitalization data for 1992-2007 with register data on unemployment and other labor market outcomes and analyze to what extent health outcomes of Swedish children, measured as overnight stays at hospitals, are worse among children whose parents become unemployed. In order to take selection into account, i.e. that parents who become unemployed are not a random sample of parents, we use an individual fixed effects approach. Thus we can compare cross section estimates to individual fixed effects estimates where the latter use within-child variation in parental unemployment.

Our fixed effects approach allows us to handle and assess the importance of selection. However, to the extent that the health consequences of parental unemployment develop slowly or if they are long lasting, this approach risks overestimating selection and underestimating the strength of the association between adverse labor market outcomes and child health. To remedy this problem we also study the effect of the first time that the parent is unemployed on health in all following years. We are however reluctant to draw strong conclusions regarding causality. First, it is possible that the causality runs in the opposite direction, i.e. from child health to parental unemployment. That children's health status may affect parental labor supply is supported by, e.g., Powers (2001) and Heck and Makuc (2000) who find that parents to children with disabilities or special needs are likely to work fewer hours. Second, absent true exogenous variation in parental unemployment and child health it is not possible to rule out the presence of unobserved factors or shocks that may influence both parental unemployment and child health.

We find that the children of unemployed parents are much more likely to be hospitalized than other children. Comparing the incidence of hospitalization for any diagnosis of the children whose parents are unemployed to children whose parents are employed, the former are 17 percent more likely to be hospitalized. However, we find that much of these raw differences are driven by selection. Using the child-fixed effect approach we find that the immediate effect is a 1 percent increase in hospitalization and the long-run effects about 5 percent increased likelihood of hospitalization. Studying the impact of maternal and paternal unemployment spells separately shows that the mother being unemployed seem to be more detrimental to child health than the father's employment status. This is the opposite pattern to the one found in the plant closure literature which has studied school outcomes.

We also find some interesting heterogeneous effects depending on parental characteristics. For example, although parental unemployment among families with low education level is correlated with worse child health, we find no effect of unemployment once we control for child-fixed effects. For families where at least one parent has some higher education, however, unemployment have an effect on the probability to being admitted to hospital. Another interesting result is that parental unemployment seems to be more harmful for children whose parents are born abroad.

The paper is organized in the following way: First we outline a theoretical framework for thinking about the consequences of parental unemployment for child health and discuss empirical evidence on the effects of unemployment on parents and children. We also discuss the issue of reverse causality, i.e. that poor health of children may make it hard for parents of work. In section 3 we present the empirical strategy and discuss its limitations. We present the data and the institutional setting in section 4. Section 5 presents the main results and section 6 concludes.

2 Consequences of parental unemployment on child health

In this section we will first formulate a simple production function for child health and discuss how the arguments in the production function are affected when a parent becomes unemployed. We will thereafter discuss earlier empirical evidence on the effect of parental unemployment on child outcomes.

2.1 A production function for child health

In order to organize ideas on how parental unemployment may affect child health, it is helpful to start with a simple production function for child health. [3] The main elements of this production function are family consumption of market goods and parental care, where the latter is a function of parental time and parental human capital. Further elements are publicly provided goods and care, such as preventive health care programs and other forms of publicly provided health investments in school or otherwise, the child's previous human capital which is a function of both previous health condition, genetic disposition and other cognitive and non-cognitive skills that may influence health outcomes. There is of course also an element of luck, or bad luck in the case of bad health shocks.

$$\begin{aligned} &child\ health_t \\ &= H(\ market\ goods_t, parent\ care_t, publicly\ provided\ goods_t, publicly\ provided\ care_t, \\ &\qquad child\ human\ capital\ _{t-1}\ , health\ shock_t) \end{aligned}$$

Parental unemployment is likely to affect several components in this production function. First, and perhaps most direct, unemployment implies lost earnings, which can lead to a reduction in both quantity and quality of market goods. To some extent, lost earnings are compensated with benefits from the unemployment insurance, but even if parents receive UI-benefits, these typically do not fully compensate for lost earnings.[4] Also, there is typically a limit on the time that UI-benefits can be received. [5] Loss of income could potentially also lead to a reduction in the family consumption of goods and activities that are hazardous, such as alcohol and cigarettes[6] Swedish evidence, however suggests that job-loss leads to more alcohol related diseases both for men and

[3] Inspiration for the proposed health production function comes from Gronau (1974) and Rosenzweig and Schultz (1983).
[4] A persons entitled to the income related benefit receives 80 percent of lost income, up to a ceiling, for the first 200 days and thereafter 70 percent for an additional of 100 days. Thus, becoming unemployed does mean lower income but most workers receive benefits which make up for 80-50 percent of the income loss.
[5] To receive unemployment benefit from unemployment insurance you need to fulfill a working requirement which implies that you need to have worked at least 80 hours per month for 6 months within the last 12 months, or a total 480 hours (min 50 hours per month) during uninterrupted 6 month period to qualify for basic benefits. Second you need to have been a member of an unemployment insurance fund for at least 12 months to qualify for income related benefits.
[6] The earlier evidence on adult health effects of economic recessions and down turns, often using aggregate data, show elements of counter cyclicality in health (Ruhm, 2000 and Ruhm and Black, 2002). One explanation for this pattern is that the consumption of hazardous goods decreases.

women (Eliason and Storrie, 2009). We might at first suspect that changes in consumption patterns following income loss would have negative consequences mostly for child health in poor households, where nutrition levels are critical. However, lower or altered consumption patterns may also involve sports activities or other health promoting activities for the children that middle income families no longer can prioritize if they experience a drop in income.

Second, unemployed parents arguably have more time for their children since they do not spend time at work. However, to the extent that job search and home production of goods services that previously could be outsourced or bought in the market require time, the increase in time available for child health investments need not be all that important. Moreover, if the unemployed parent suffers from status loss, stress or poor health as a result of the job loss, as is shown to happen in Kuhn, Lalive and Zweimuller (2007) and in the Swedish case in Eliason and Storrie (2009), or if the job loss leads to a deterioration of the home environment due to parental conflict, the quality adjusted time spent with children need not improve[7] Eliasson (2011b) finds that the risk of marriage dissolution increases by 13 percent in Sweden as a consequence of husband's job loss[8] That parents' well-being is important for child outcomes is also supported by Adda, Björklund and Holmlund (2011) who analyze the effect of parental deaths on children's cognitive and non-cognitive skills. They find that both mothers and fathers are important, but mothers are somewhat more important for cognitive skills and fathers for non-cognitive skills.

Publicly provided goods and care may also change as a result of parental unemployment. Although local economic conditions are likely to be correlated with parental unemployment, and although local child health promoting spending in schools and in childcare may decline as many parents lose their jobs in a region, these changes are likely to be of smaller magnitude than the direct effect on parental resources and

[7] While Kuhn, Lalive and Zweimuller (2009) find that expenditures on medical treatments in general are not strongly affected by job displacement they find that job loss significantly increases expenditures for antidepressants and related drugs, as well as for hospitalizations due to mental health problems for men (but not for women) although the effects are economically rather small. They also find that sickness benefits strongly increase due to job loss. In a study on Danish data, Browning, Dano and Heinsen (2006) find no health effects of job loss. Job displacement is found to increase mortality in Sweden, Norway and the US (Eliason and Storrie (2009), Rege Telle and Voturba (2009) and Sullivan and von Wachter (2009)) However, Martikainen, Maki and Jäntti (2007) find no effect for Finland. In particular, Eliason and Storrie (2009) study the consequences of job displacement during a 12-year period and find that job loss significantly increases the risk of hospitalization due to alcohol-related conditions, among both men and women, and due to traffic accidents and self-harm, among men only.

[8] Huttunen and Kellokumpu (2012) find similar evidence for Finland.

care, in particular for children in daycare and at school. However, if children who were previously enrolled in childcare instead are cared for at home, publicly provided care and goods investments in these children will decline. It is possible, but not certain, that time and resources invested in child health at home make up for the difference. Also, to the extent that access to publicly provided health investments, such as immunization programs, check-ups and other forms of preventive care, require time investments from parents, unemployed parents may in some situation have better access to these resources, which could lead to improvements in child health. To summarize, the theoretical prediction of in which parental unemployment affects child health outcomes is ambiguous, implying that it is an empirical question.

2.2 Empirical evidence – Child outcomes

There are a few studies on the long run consequences of parental job loss due to plant closure on children. Oreopoulos, Page and Stevens (2008) study the intergenerational cost of negative employment shocks on child earnings. Using Canadian administrative data that follows more than 39,000 father-son pairs from 1978 to 1999, they find that children whose fathers were displaced have annual earnings about nine percent lower than similar children whose fathers did not experience an employment shock. These children are also more likely to receive unemployment insurance and social assistance as young adults. The estimates are driven by the experiences of children whose family income was at the bottom of the income distribution.

Page, Huff Stevens and Lindo (2007) find no evidence of intergenerational effects of parental job loss on the average child on US data. However, when they analyze disadvantaged children (defined by family income or race), they find evidence of negative effects of parental displacement on income, earnings, and completed education. In particular, they find large, statistically significant negative effects on the next generation's income and earnings for all children when displacements include layoffs, but not when they are restricted to firm closures. They interpret this to imply that individuals who are selected for layoffs may have unobserved characteristics that are correlated with their children's outcomes. Although the findings suggest that firm closings have no intergenerational effects on average, there is evidence that such events impose long-term costs on disadvantaged children. Moreover, an interesting finding is that the effects of exogenous income shocks (from business closings) are largest among children who were younger than 7 at the time of the income shock. In a similar study on Norwegian data, Bratberg et al (2008) find that although displaced parents experience significant reductions in both earnings and employment, there are no significant effects on earnings of the next generation. This results contrast from the studies on North America which found negative effects at the lower end of the income distribution.

Studies of the long run consequences for children do not provide information as to why parental job loss affects the children. A recent literature has attempted to study the immediate effects of parental job loss on children's schooling outcomes. Stevens and Schaller (2010) study the effect of parental job loss on grade retention. They find a fifteen percent increase when controlling for child fixed effects. The effect is driven by children whose parents have a high school education or less. There is no evidence of significantly increased grade retention prior to the job loss, suggesting a causal link between the parental employment shock and children's academic difficulties. In a study of Canadian youth, Coelli (2010) finds that parental job joss from layoff and business failure that occur when youth are in the process of completing high school, leads to drops in College enrollment by ten percent. The effect comes from main bread winner job loss – and not spousal job loss. This is interpreted to indicate that the main channel is the loss of income. It is indeed shown that parental job losses are followed by significant falls in parental income.

Using Norwegian register data, Rege, Telle and Voturba (20011) estimate how children's school performance is affected by their parents' exposure to plant closure. The estimates suggest that paternal job loss has a negative effect on children's school performance. Maternal job loss is instead associated with a non-significant increase in school performance. The study explores and finds that the negative effect of paternal job loss appears to be unrelated to its effect on father's income, father's employment status, shifts in maternal time towards employment, marital dissolution, and residential relocation.

3 Empirical strategy

There are two issues we aim to analyze in this paper. First, we are interested in the correlation between parental unemployment and child health outcomes, i.e. do children to unemployed experience worse health outcomes than children whose parents are working. We believe that these correlations are of own interest. Second, we would like to control for selection into unemployment, i.e. the fact that some families are more likely to experience both unemployment and bad health due to some underlying, unobservable characteristics. We will handle the issue of selection by including children fixed effects in the estimation, which implies that we compare the health of a child when the parent is unemployed to the health of the same child when the parent is working.

We will estimate the following econometric model both without and with the children specific fixed effect:

$$health_{it} = \alpha + \beta\, parent\ \ unemployed_{it} + \delta_i + \boldsymbol{X}_{it,(t-1)} + year_t + \varepsilon_{it}$$

where the outcome *health* is an indicator variable taking the value 1 if an individual *i* has been admitted to hospital at least once in year t. Our variables of interest are *parent unemployed* which is a dummy variable taking the value one if the mother or the father is unemployed and zero if both parents are employed. The year fixed effects capture calendar year variation in hospitalization and possible changes in coding practice that affect all admittance in a given year, regardless of age of the child. X_{it} is a vector of time varying (and fixed) parent and child characteristics, in particular age and sex of the child, parental age and education, parental health in previous period, family disposable income in the previous period as well as the local unemployment rate. ε is the usual error term. Standard errors are clustered on maternal level since there may be random shocks to the family creating correlation in illness across siblings.

This approach arguably allows us to handle and assess the importance of selection. However, to the extent that the health consequences of parental unemployment develop slowly or if they are long lasting this approach risks underestimating the strength of the association between adverse labor market outcomes and child health. We will therefore conduct an analysis where we try to capture long run effects of parental unemployment by defining the variable *unemployment long* that takes the value zero until one of the parents become unemployed and the value one thereafter.[9] Since our data starts in 1992 we cannot observe the unemployment history of children that are born before 1992, and

[9] We define the variable as missing if any parent is outside the labor force before the first unemployment spell, but once a parent has experienced unemployment it takes the value one thereafter even if any parent leave the labor force.

we will therefore focus on the cohorts born in 1992 and later for this part of the analysis.

Even though we believe that our approach is able to handle both selection and long run consequences of parental unemployment, there are several reasons why it may be unable to capture causal effects. First, as we have argued, there are reasons to believe that causality can run in both directions, such that parental unemployment can affect child health and that poor child health can cause parental unemployment or withdrawal from the labor market. [10] Second, both unemployment and child health may be affected by an outside event that we are unable to measure. The child specific fixed effects are only able to capture different type of family characteristics that are constant over time, and hence unable to control for time-varying factors that might be of importance. We are hence reluctant to draw strong conclusions about the direction of causality from our analysis.

4 Data and variables

The data analyzed in this paper comes from a number of official registers covering the total Swedish population. We will focus on children aged 3-18 during the years 1992-2007 and their biological parents.[11] For simplicity we will limit our analysis to children with both biological parents alive. This gives us approximately 1,3 million observations (children) each year.

We have chosen to study the effect of unemployment of the biological parents rather than, for example, the adults living in the same household as the child according to the register. The main motivation for this choice is that it is common in Sweden that parents share custody of children when separating, implying that the child live with both the mother and the father. In the registers the child can only be assigned to one household making it impossible to know exactly how the time is divided (if it is divided) between

[10] Furthermore, it is not obvious how to sort out the exact timing of events event in monthly register data, both regarding unemployment spells and hospitalization In the case of unemployment, a parent is likely to be given notice well in advance of actual registration at the unemployment office. Labor market contracts will typically dictate different lengths of the legal notification period, both depending on the type of job and on tenure. Since entitlement to benefits requires registration, it is however likely that those who become unemployed eventually register when they need benefits. Moreover, being registered as unemployed requires the individual to actively seek work. A parent, who has become unemployed because of the need to care for a sick child, may hence have to postpone registering as unemployed to when the child is getting better. Determining the timing of child health shocks is also problematic. It is likely that the child in many cases have been ill already a number of days before hospitalization.

[11] The reason for excluding children younger than age 3 is that parents to larger extent stay home with parental leave benefits.

parents. Moreover, the child may be affected by parental unemployment although the child does not live with the parent since it may imply lower payments to the care of the child. Finally, due to data limitations, we cannot always observe whether an additional adult live in a separated parent's household or not, unless the adults are married or have common children.

Data on health outcomes are taken from the National Patient Register that contains information about all in-patient care in Swedish hospitals, including information about the length of the stay as well as both main diagnosis and secondary diagnoses. Our dependent variable is a dummy variable indicating whether the child has been in in-patient care for any diagnose during the year. Obviously, hospitalization is not always the first sign of bad health and in most cases children do not need hospital care at all. Thus, a limitation with our measure of child health is that hospitalization data only pick up severe health problems. The advantage with using register data is that it is a fairly objective measure of health. Since Sweden has a universally provided, publicly funded health care system of good quality and free health care for children, admittance to hospital should reflect the need of health care rather than the financial resources of the parents.

Data on children's health outcomes are linked to data on parents from the administrative registers LOUISE from Statistics Sweden. LOUISE have information on parental income, education and age. Information on unemployment comes from the Swedish Public Employment Service. In the data, we observe whether the parent has been registered at the employment office during the observational year. A parent is defined as unemployed if he/she is registered as openly unemployed or participates in a labor market program at any occasion during the year, and as employed if he/she is not registered as openly unemployed/participating in a labor market program and has an income from paid work or self-employment which exceeds the Income Base Amount[12]

By imposing an earnings requirement for being employed we restrict the analysis to children of parents participating in the labor force and exclude parents who are not registered as unemployed without earnings. The motivation for excluding parents outside the labor force is that we suspect that one reason for not participating in the

[12] The Income Base Amount is set every year by the Swedish Government and is depends on the development of wages in the economy. Among other things it is used to determined amount paid to the public pension system.

labor force might be that parents take care of a sick child.[13] By excluding these parents we limit the risk of capturing reversed causality. In addition we avoid the risk of having the estimated relation between unemployment and child health affected by the possibility that parents who are out of the labor force to care for a sick child register as unemployed when the child gets well. Such behavior would imply that health improvements induce parental unemployment. Besides taking care of a sick child, there are a number of other potential reasons for being out of the labor force, e.g. being a full time student, staying at home taking care of (healthy) children. It is therefore likely that being out of the labor force affects children quite differently than parental unemployment.

How good is our measure of unemployment, or put in another way, is there a risk that we miss people that are actually in the labor force searching for jobs, but have chosen not to register at the Employment Services? We believe that this risk is limited, since there are strong incentives for unemployed to register at the Employment Services. First of all, only registered unemployed are eligible for unemployment benefits. Second, access to training and coaching requires registration.

Column i) in Table 1 shows summary statistics for sickness prevalence, a number of child and parental characteristics as well as unemployment for the sample used in the estimations. Approximately 38.6 children out of 1,000 have at least one hospital stay during the year. Furthermore, 30.5 percent have at least one parent that experiences unemployment during the year, whereas only 6.8 percent experience that both parents are unemployed during the year.[14] It is also somewhat more common that the mother is unemployed than that the father is unemployed. Finally, we note that hospitalization is considerably larger among mothers than among fathers, something we think is due to women experiences spells of hospitalizations in connection with child births.

In the next three columns we have divided the sample depending on parental unemployment status. Children whose parents are unemployed during the year are more likely to have at least one hospital stay. They are slightly younger, which is also true for

[13] However, in Sweden all working parents get compensation from the public insurance system when they temporarily need to stay home from work to care for a sick child under the age of 12, and in special circumstances until age 16. During the 1990’s and early 2000’s, mothers took about 65 percent of the total number of days. In case of longer illnesses lasting more than 6 months, parents are entitled to a special care allowance.

[14] Not that the unemployment spells of the mother and father do not need to occur on the same time.

the parents, and live in families with lower disposable income. Moreover, unemployed parents are more likely to have worse health.

In sum, from simple summary statistics it does seem that children whose parents experience unemployment have worse health outcomes than other children. However, this may be due to the fact that these families typically are younger or that parents have worse health. In addition, there might be other, unobservable family characteristics that affect both the likelihood of the parent being unemployed and the likelihood of the child experiencing bad health. The next section will deal with these issues.

Table 1 Summary statistics – annual observations on children age 3-18 for the years 1992-2007.

	Parents in labor force	Any parent unemployed	Both parents unemployed	No parent unemployed
Hospitalization	38.55	43.78	46.92	36.25
Age	10.50	9.79	9.12	10.82
Girl	0.486	0.485	0.484	0.49
Swedish-born parents	0.901	0.812	0.631	0.940
Age, mother	39.12	37.13	35.39	39.99
Age, father	41.84	40.25	38.97	42.54
Years of educ. mother	11.76	11.04	10.62	12.06
Years of educ. father	11.72	11.13	10.86	11.97
Sick, mother	78.83	103.63	129.97	67.92
Sick, father	47.00	60.33	75.20	41.15
Disp. Income	271,854	214,092	178152	297,280
Any parent unemp.	0.305	1	1	0
Both parents unemp.	0.068	0.224	1	0
Mother unemp.	0.210	0.688	1	0
Farther unemp.	0.164	0.536	1	0
No of obs	21,109,926	6,445,896	1,444,610	14,664,030

As mentioned in Section 3, we will in addition to analyzing the direct links between parental unemployment and child health also analyze more long run consequences of parental unemployment by focusing on a sample for which we observe parental outcomes over the child's , by focusing on children born 1992 and later. Table 2 shows summary statistic for this reduces sample.

Table 2 Summary statistics – annual observations on children born 1992 or later, for the years 1992-2007.

	All children	Any parent unemployed	No parent unemployed
Hospitalization	36.28	37.38	33.10
Age	7.16	7.53	6.75
Girl	0.487	0.487	0.488
Swedish-born parents	0.851	0.831	0.957
Age, mother	36.53	35.98	37.73
Age, father	39.43	38.93	40.15
Y. of educ., mother	11.72	11.37	12.44
Y. of educ., father	11.67	11.39	12.26
Sick, mother	107.56	110.92	80.12
Sick, father	45.38	49.10	33.60
Disp. Income	273,351	247,440	338,954
Any parent unemp.	0.671	1	0
No of obs	9,384,169	5,641,407	2,771,080

5 Results

5.1 Graphical analysis

Before turning to the results from the estimations, we will start with a graphical analysis. Figure 1 below shows how children's health varies through childhood and by parental unemployment status. The number of children out of a thousand who are hospitalized during a year at a particular age whose parents are unemployed is depicted in the dashed line and the corresponding number for children whose parents are not unemployed is depicted by the full drawn black line. The left figure shows the health profiles for boys and the right figure shows the profiles for girls.

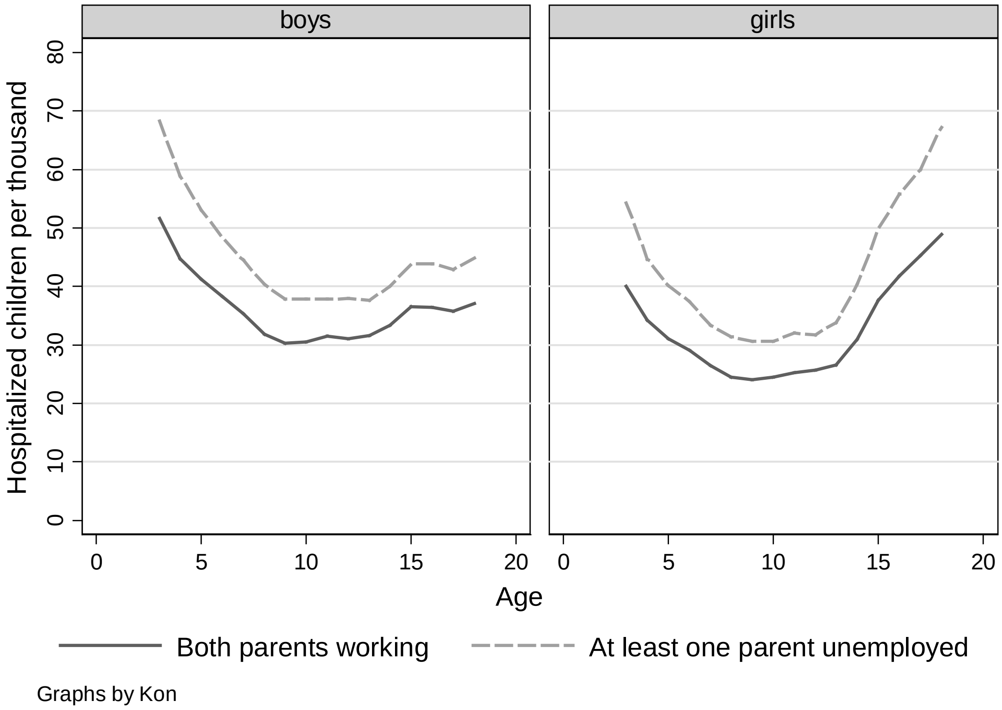


Figure 1 Hospitalization for illness or injury per 1000 children by parent's employment status for boys and girls, ages 3-18 in 1992-2007.

It is evident that the probability to be admitted to hospital is largest for very young children. As the children grow older the incidence decreases until the age of 9 were the curve flattens out. During adolescence the curve turns upward again. It is also interesting to note that preschool and primary school boys have a higher incidence of hospitalization than girls, but that teenage girls are more likely to be hospitalized than teenage boys. This can be seen more readily in Figure2 which shows the ration between the two curves.

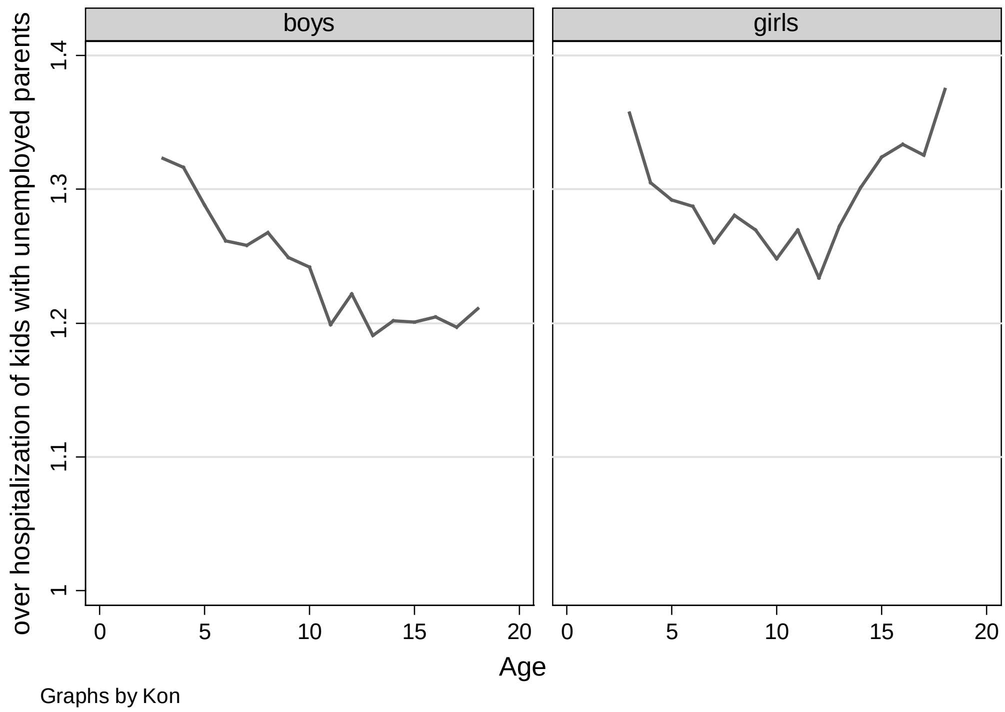


Figure 2 Ratio of hospitalization for illness or injury per 1000 children. Children with unemployed parents/children with employed parents for boys and girls, ages 0-18 in 1992-2007

The next figure show the relative hospitalization rate for children with unemployed parents compared to employed parents for children which experienced the first unemployment at different ages. Thus, the full line show the relative hospitalization rate for children for which one parent was unemployed at the age 0-2. The comparison group for all categorizes are children with parents who always are employed.

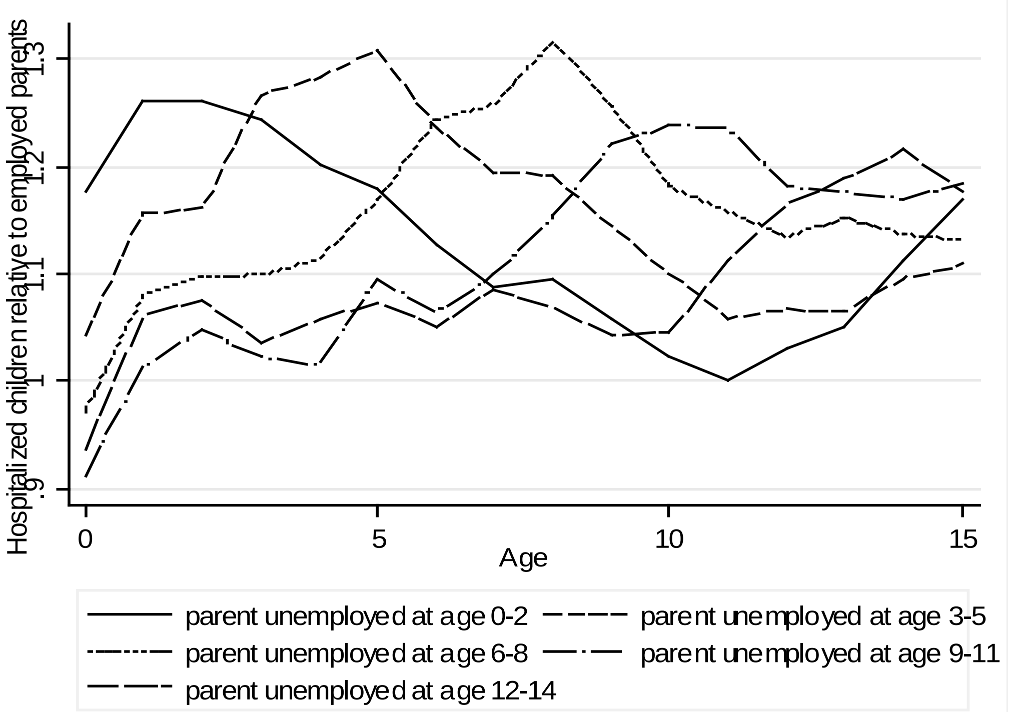


Figure 3 Hospitalization for illness or injury per 1000 children by parent's employment status and age of parent's first unemployment spell ages 3-18 in 1992-2007

As discussed before, this pattern of worse health for the children on the unemployed can be due to selection, i.e. that there is some underlying reason why a parent both is unemployed and has sick children. But, it is also possible that parental unemployment affects child health or that the causality goes the other way, such that child health affects parental employment status. Next we therefore turn to the econometric analysis

5.2 Estimation results

5.2.1 Short run

Next we turn to the empirical results, where we will analyze how unemployment status of parents co-varies with children's hospitalization. The results are displayed in Table 3. Column (i) shows the correlation between parental unemployment (measures as having at least one parent unemployed) and children's hospitalization, only controlling for age and sex of the child as well as year-effects. There seems to be a strong positive correlation between having a parent that is unemployed and spending at least one night at the hospital. The estimate implies that 6.6 more children per thousand children are hospitalized at least one night if at least one of their parents is unemployed during the year. If we compare these figures to the mean incidence of hospitalization which is 38.55 per thousand children, is corresponds to an increase with 17 percent if any of the parents is unemployed. In column (ii) we also control for parental characteristics such as parental age, education, past hospitalization, and whether the parents are born in Sweden or not in order to will handle at least some of the potential selection into unemployment. Doing this reduced the parameter estimate somewhat to 5.1. In column (iii) we also control for lagged family income[15], which further reduces the point estimate somewhat. In column (iv) we also include municipal unemployment, resulting in a point estimate of 4.2.[16]

Hence, there is a strong correlation between parental unemployment and children's health outcomes. It is also clear that much of the correlation is due to selection: the initial estimates are considerably reduced as controls are introduced. In a further attempt to handle selection, we include child specific fixed effects, that will capture any genetic pre-position or family or child specific characteristics that are constant over time. Such an approach builds on individual observations that change status over time, i.e. children whose parents work in some years and not in others. It is worth noting that the identifying variation comes from the sample of parents that change employment status during the child's upraising. As a result we cannot say anything about the effects of parental unemployment for children whose parents always or never are unemployed. The results in column (v) in Table 3 show that the point estimate diminishes considerably, but that we still find a statistical significant effect of unemployment on child hospitalization. The point estimate implies that 0.37 more children per thousand children are hospitalized at least one night if at least one of their parents is unemployed during the year. This corresponds to an increase with 1 percent.

Table 3. Parental unemployment status and child hospitalization, short run

	i	*ii*	*iii*	*iv*	*v*
Any parent	6.617***	5.144***	4.841***	4.238***	0.367**

[15] In this specification, we include lagged disposable income in the mother's family in percentiles. We have also experimented with including the family income in levels as well as defining families differently. This does not alter the results to any large extent.
[16] We have also included municipality fixed effects, but that it turns out that this gives the same result as when including municipal unemployment.

Unemployed	(0.114)	(0.121)	(0.124)	(0.125)	(0.151)
Year, age	Yes	Yes	Yes	Yes	Yes
Parental controls	No	Yes	Yes	Yes	Yes
Family income	No	No	Yes	Yes	Yes
Municipal unemployment	No	No	No	Yes	Yes
Child fixed effect	No	No	No	No	Yes
No of observations	21,109,926	20,180,064	20,132,611	20,132,611	21,023,720
No of individuals					2,929,595

Clustered (on mother) robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1.

In the analysis in Table 3 it was enough that one of the parents was unemployed. Does it matter if both parents are unemployed? In Table 4 we investigate this. Not surprisingly, we find that the correlations between having both parents unemployed and being hospitalized, but when controlling for child-fixed effects, we do not find any additional effect of having both parents unemployed.

Table 4 Both parents unemployed

	i	ii	iii	iv	v
Any parent	5.902***	4.569***	4.377***	3.856***	0.382**
Unemployed	(0.120)	(0.126)	(0.129)	(0.129)	(0.152)
Both parents	3.261***	3.057***	2.686***	2.240***	-0.171
Unemployed	(0.228)	(0.250)	(0.253)	(0.253)	(0.287)
Year, age	Yes	Yes	Yes	Yes	Yes
Parental controls	No	Yes	Yes	Yes	Yes
Family income	No	No	Yes	Yes	Yes
Municipal unemployment	No	No	No	Yes	Yes
Child fixed effect	No	No	No	No	Yes
No of observations	21,109,926	20,180,064	20,132,611	20,132,611	21,023,720
No of individuals					2,929,595

Clustered (on mother) robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1.

5.2.2 Long run

In the analysis above we investigated the association between parental unemployment and child hospitalization taking place in the same year. However, there are reasons to believe that this analysis may underestimate the detrimental effects of parental unemployment. It probably takes some time before a parent's unemployment start to affect child health. Especially since we are analyzing a quite serious indication of bad health; hospitalization. In Table 5 we therefore estimate a long-run model where the variable "Any parent unemployed" takes the value one all first year that any parent is unemployed as well as all consecutive years. Since we only observe the full employment history for parents whose children are born in 1992 and thereafter, we limit the sample to those children.

Looking at the results in Table 5 we find evidence of a strong correlation between parental unemployment and child hospitalization, that decreases when we control for more and more covariates. Most interesting however, is the fact that the point estimate in column (v), i.e. when controlling for child-specific fixed effects, is considerably larger than the corresponding point estimate in **Table 3**. The point estimate of 1.8 indicates that a child is 5 percent more likely to be hospitalized the years following parental unemployment which must be considered to be a rather large effect.

Table 5 Parental unemployment status and hospitalization: Long run

	i	ii	iii	iv	v
Any parent unemployed	5.785*** (0.167)	3.987*** (0.183)	3.927*** (0.188)	3.157*** (0.188)	1.812*** (0.504)
Year, age	Yes	Yes	Yes	Yes	Yes
Parental controls	No	Yes	Yes	Yes	Yes
Family income	No	No	Yes	Yes	Yes
Municipal unemployment	No	No	No	Yes	Yes
Child fixed effect	No	No	No	No	Yes
No of observations	8,412,487	7,672,022	7,663,944	7,663,944	8,390,901
No of individuals					1,239,557

Clustered (on mother) robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. Children born 1992 and later

5.2.3 Does it matter whether it is the mother or the father that is unemployed?

So far we have treated maternal and paternal unemployment symmetrically. There are however some empirical evidence that the effects might differ depending on which parent that experiences unemployment. In Table 6 we therefore separate between maternal and paternal unemployment. In columns (i) and (ii) we estimate short run effects with and without child-specific fixed effects, and in columns (iii) and (iv) we estimate long run effects.

From the results it is evident that maternal unemployment matters more. In the fixed-effect estimations, paternal unemployment does not have a statistical significant effect. This suggests that the mother being unemployed is more detrimental to child health.

Table 6 Maternal or paternal unemployment status and hospitalization

	i	ii	iii	iv
	Short run		Long run	
Maternal unemp.	4.087***	0.379**	3.116***	1.207**
	(0.144)	(0.171)	(0.186)	(0.540)
Paternal unemp.	2.753***	0.0430	1.765***	0.821
	(0.158)	(0.195)	(0.193)	(0.557)
No of obs.	20,132,611	21,023,720	7,551,504	8,265,745
No of individuals		2,929,595		1,239,748
Year, age	Yes	Yes	Yes	Yes
Parental controls	Yes	Yes	Yes	Yes
Family income	Yes	Yes	Yes	Yes
Municipal unem.	Yes	Yes	Yes	Yes
Child fixed effect	No	Yes	No	Yes

5.2.4 Does children's age matter?

To be included

5.2.5 Heterogenous effects

The next thing we are interested in analyzing is whether the associations between parental unemployment and child hospitalization differs with respect to the child's sex, parental country of birth and parental education. In order to be able to put the estimates into perspective, Table 8 presents summary statistics of hospitalization and parental unemployment in these subgroups. From the table we can note that boys are more often hospitalized than girls, and that hospitalization decreases with parental education. Also, unemployment is much more common among parents born outside Sweden, and also decreases with parental education.

Table 7 Summary statistics, hospitalization and parental unemployment for the different subgroups

	Sample, short run		Sample, long run	
	Hospitalization	Any parent unemployed	Hospitalization	Any parent unemployed
Girls	37.65	0.305	32.07	0.670
Boys	41.33	0.306	40.29	0.671
Swedish-born parent	39.98	0.275	36.61	0.639
No Swedish-born parents	36.62	0.579	34.41	0.889
Low education	45.15	0.490	41.54	0.867
Medium education	40.41	0.313	37.02	0.673
High education	35.52	0.240	33.39	0.626

In Table 9 we estimate our models for girls respectively boys separately. The overall message from the table is that there are no large differences between girls and boys. If anything, the long run effects are stronger for girls (see column (iv)). Note also that we see from Table 8 that the prevalence of hospitalization is lower for girls than for boys (37.6 compared to 41.3), so in also in relative terms the effects are larger for girls than for boys.

Table 8 Parental unemployment status and hospitalization: Heterogenous effects wrt sex

	i	ii	iii	iv
	Short run		Long run	
		Girls		
Any parent unemployed	4.582*** (0.174)	0.247 (0.210)	3.301*** (0.253)	2.637*** (0.676)
No of obs.	9,796,497	10,226,743	3,732,182	4,086,040
No of individuals		1,424,422		603,004
		Boys		
Any parent	3.913***	0.493**	3.028***	1.051

unemployed	(0.175)	(0.215)	(0.273)	(0.742)
No of obs.	10,336,114	10,796,977	3,931,762	4,304,861
No of individuals		1,505,173		636,553
Year, age	Yes	Yes	Yes	Yes
Parental controls	Yes	Yes	Yes	Yes
Family income	Yes	Yes	Yes	Yes
Municipal unem.	Yes	Yes	Yes	Yes
Child fixed effect	No	Yes	No	Yes

Table 10 instead separate between children depending on whether at least one of their parent is born in Sweden or not. The results in the top-panel (those with at least one parent born within Sweden) resemble the results in Table 3 a lot, which is natural given that the majority of parents are born in Sweden. The result in the bottom panel indicates somewhat lower associations when child-specific fixed effects are not controlled for (columns (i) and (iii)), but larger effects once child-specific fixed effects are included. Our interpretation of this pattern is that once we have taken care of selection, unemployment per se is more harmful for children with both parents born outside Sweden. The point estimate of 5.4 implies that children are almost 15 percent more likely to be hospitalized in a year when at least one of the parents is unemployed.

Table 9 Parental unemployment status and hospitalization: Heterogenous effects wrt parental country of birth

	i	ii	iii	iv
	Short run		Long run	
	At least one parent born in Sweden			
Any parent	4.317***	0.217	3.148***	1.280**
unemployed	(0.132)	(0.161)	(0.193)	(0.522)
No of obs.	18,620,048	18,978,109	7,067,838	7,342,045
No of individuals		2,574,811		1,066,131
	Both parents born outside Sweden			
Any parent	3.288***	0.997**	2.310***	5.401***
unemployed	(0.394)	(0.444)	(0.863)	(2.036)

No of obs.	1,512,563	2,045,611	596,106	1,048,856
No of individuals		368,898		178,432
Year, age	Yes	Yes	Yes	Yes
Parental controls	Yes	Yes	Yes	Yes
Family income	Yes	Yes	Yes	Yes
Municipal unem.	Yes	Yes	Yes	Yes
Child fixed effect	No	Yes	No	Yes

Finally, Table 11 investigates whether the estimates differ with respect to parental education. We have divided parents into three categories depending on the number of years of education. We define parents to have *low education* if both parents have no more than compulsory schooling, to have *high education* if any of the parents has some higher education, and to have *medium education* otherwise.

From the table we see that the associations are higher for children to low and medium educated parents, whereas the estimates when controlling for child-fixed effects are larger for children with medium or highly educated parents. The results in the top panel first column show that children with unemployed parents are 10 percent more likely to be hospitalized among families where the parents have a low education level. For children from families with higher education (results in the bottom panel) the association show a 5 percent increase. According to the estimates in the second column 2 percent of the increase is due to unemployment. Our interpretation of this is that selection is not an issue among highly educated parents, but when unemployment hits, the damage is worse.

Table 10 Parental unemployment status and hospitalization: Heterogenous effects wrt parental education

	i	ii	iii	iv
	Short run		Long run	
	Low education			
Any parent unemployed	4.680***	-0.815	4.951***	-1.065
	(0.470)	(0.607)	(1.146)	(3.070)
No of obs.	1,246,213	1,246,213	402,846	402,846

No of individuals		200,082		58,151
		Medium education		
Any parent	4.602***	0.405**	3.559***	1.917***
unemployed	(0.143)	(0.176)	(0.215)	(0.590)
No of obs.	14,636,997	14,636,997	5,910,160	5,910,160
No of individuals		1,963,371		836,653
		High education		
Any parent	1.888***	0.653**	1.578***	1.757*
unemployed	(0.303)	(0.332)	(0.402)	(1.015)
No of obs.	4,249,401	5,140,510	1,350,938	2,077,895
No of individuals		766,142		344,753
Year, age	Yes	Yes	Yes	Yes
Parental controls	Yes	Yes	Yes	Yes
Family income	Yes	Yes	Yes	Yes
Municipal unem.	Yes	Yes	Yes	Yes
Child fixed effect	No	Yes	No	Yes

6 Conclusions

There is vast anecdotal and correlation evidence that children of unemployed parents fare worse than children whose parents are employed. However, a careful empirical analysis of the role of parental unemployment for child health is more scarce. Using rich register data on child hospitalizations and parental labor market outcomes for all Swedish families over the period 1992-2007, we analyze how parental unemployment, measures as being registered at the unemployment office is related to hospitalization of children aged 3-18. In order to take selection into account we have also adapted a fixed-effect approach where we use the within child variation in parental employment, i.e. compare the health of a child in a year when his/her parent is unemployed with the health of the child when his/her parent is employed.

We confirm that there is a strong correlation between parental unemployment and children's hospitalization: having an unemployed mother or father is on average associated with 17 percent higher likelihood of having to stay at least one night at a hospital. We find that a large part of the correlation seems to be driven by selection. However, even after controlling for child-fixed effects, we find that unemployment lead to a 1 percent increase in hospitalization in the short run and 5 percent increase in the long run. In contrast to earlier literature on the impact of parental exposure to plant closure on school achievement, we find maternal unemployment to be more detrimental to child health than paternal unemployment. Moreover, we find that the higher likelihood of children with unemployed parents with low education to be hospitalized is due to selection, whereas for children with parents with higher education parental unemployment has a negative effect on health.

Our overall conclusion is that parental unemployment hurts child health, but that policies directed towards improving the health of children need to address not only the consequences of temporary parental unemployment, but also the long term vulnerability of children growing up in families with weak labor market attachment.

References

Adda, J, A Björklund and H Holmlund (2011) "The Role of Mothers and Fathers in Providing Skills: Evidence from Parental Deaths", IZA Discussion Paper No. 5425.

Almond, Douglas and Currie, Janet, (2011) "Human Capital Development before Age Five," *Handbook of Labor Economics*, Elsevier.

Bratberg, Espen, Øivind Anti Nilsen, Kjell Vaage, "Job losses and child outcomes", *Labour Economics* 15: 591–603

Browning, Martin, Anne Moller Dano and Eskil Heinesen, "Job displacement and stress-related health outcomes," *Health Economics*, 15(10): 1061-1075.

Christoffersen, Mogens Nygaard (1994), "A follow-up study of longterm effects of unemployment on children: loss of self-esteem and self-destructive behavior among adolescents." *Childhood* 2.4, 212-220.'

Christoffersen, Mogens Nygaard (2000), 2Growing Up with Unemployment A Study of Parental Unemployment and Children's Risk of Abuse and Neglect Based on National Longitudinal 1973 Birth Cohorts in Denmark.2 *Childhood* 7.4 (2000): 421-438.

Cunha, Flavio and James Heckman (2007) "The technolgy of skill formation". *American Economic Review*, 97(2):31-47.

Cunha, Flavio and James J. Heckman (2008) "Formulating, Identifying and Estimating the Technology of Cognitive and Noncognitive Skill Formation". *J. Human Resources*, 43(4):738-782.

Coelli (2010) "Parental job loss and the education enrollment of youth", *Labour Economics,* 18: 25–35.

Eliason, Marcus (2011a) "Income loss: the role of the family and the welfare state" *Applied Economics*, 43(5): 603-618.

Eliason, Marcus (2011b) "Lost jobs, broken marriages". *Journal of Population Economics*, 25(4):1365–1397.

Eliason, Marcus and Donald Storrie (2009) "Does Job Loss Shorten Life?" *Journal of Human Resources*, 44(2).

Huttunen and Kellokumpu (2012) The Effect of Job Displacement on Couples' Fertility Decisions, *IZA DP* No. 6707

Heck, K and D Makuc (2000) "Parental Employment and health insurance coverage among school-aged children with special needs", *American Journal of Public Health* 90, 1856-1860

Kuhn, Lalive and Zweimuller (2009) "The public health costs of job loss", *Journal of Health Economics* 28: 1099–1115.

Lindo, Jason (2010) "Are Children Really Inferior Goods? Evidence from Displacement-driven Income Shocks", *The Journal of Human Resources*, 45.

Madge, N. (1983) "Unemployment and its effects on children" *Journal of Child Psychology and Psychiatry*, 24: 311–319.

OECD (2011) OECD Family database www.oecd.org/social/family/database

Oreopoulos, P., Page, M. E., & Stevens, A. H. (2008) "The intergenerational effects of worker displacement" *Journal of Labor Economics*, 26: 455–483.

Page, Marianne, Ann Huff Stevens Jason M. Lindo (2009) "Parental Income Shocks and Outcomes of Disadvantaged Youth in the United States" in *An Economic Perspective on the Problems of Disadvantaged Youth*, Jonathan Gruber, ed. NBER, University of Chicago Press

Pedersen, C R, M Madsen and L Köhler (2005), "Does financial strain explain the association between children's morbidity and parental non-employment", *Journal of Epidemiol Community Health*, 59, 316-321.

Pekka Martikainen, Netta Mäki, and Markus Jäntti (2007). "The effects of unemployment on mortality following workplace downsizing and workplace closure: a register-based follow-up study of Finnish men and women during economic boom and recession". *American Journal of Epidemiology* 165(9): 1070–1075.

Powers, E. T (2001) "New Estimates of the Impact of Child Disability on Maternal Employment" *American Economic Review* 91, 135-139.

Rege, M, K. Telle and M. Votruba, 2011, "Parental Job Loss and Children's School Performance", *Review of Economic Studies* 78 (4): 1462-1489.

Rege, M., K. Telle and M. Votruba, 2009, "The Effect of Plant Downsizing on Disability Pension Utilization", *Journal of the European Economic Association*, 7(4): 754-785.

Gronau, R. (1974) "The Effect of Children on the Housewife's Value of Time," NBER Chapters, in: *Economics of the Family: Marriage, Children, and Human Capital*, ed Theodore Schultz, National Bureau of Economic Research, Inc, University of Chicago Press.

Rosenzweig and Schultz (1983) "Estimating a Household Production Function: Heterogeneity, the Demand for Health Inputs, and Their Effects on Birth Weight" *Journal of Political Economy*, 91(5): 723-746.

Ruhm, Christopher J. (2000) "Are recessions good for your health?" *Quarterly Journal of Economics*, 115(2): 617–650.

Ruhm, Christopher J. and Black, William E., (2002) "Does drinking really decrease in bad times?," *Journal of Health Economics*, 21(4): 659-678.

Stevens and Schaller (2010) "Short-run effects of parental job loss on children's academic achievement" *Economics of Education Review* 30: 289–299.

Stevens, A. H. (1997) "Persistent effects of job displacement: The importance of multiple job losses" *Journal of Labor Economics*, 15: 165–188.

Ström, Sara (2002) "Keep out of the reach of children: parental unemployment and children's accident risks in Sweden 1991–1993", *International Journal of Social Welfare*, 11(1)40-52

Sullivan, D. G., & von Wachter, T. (2009). "Job displacement and mortality: An analysis using administrative data" *Quarterly Journal of Economics*, 124(3): 1265–1306.

UNICEF (2012) New league tables of child poverty in the world's rich countries, Innocenti Reseach center , Report card 10, UNICEF

Young Children and Unemployment Posted: 09/27/11 03:50 PM ET John Irons: Young Children and Unemployment Huffington post

CESifo Working Papers

www.cesifo.org/wp

Performance of Young Adults: The Importance of Different Skills

Torberg Falch
Ole Henning Nyhus
Bjarne Strøm

CESifo Working Paper No. 4124
Category 5: Economics of Education
February 2013

CESifo *Working Paper No. 4124*

Performance of Young Adults: The Importance of Different Skills

Abstract

This paper uses teacher assessments at age 16 in Norwegian comprehensive schools to measure different types of skills. While we follow the literature and interpret test scores in Mathematics and Science as proxy for cognitive skills, we use a novel measure for another type of skills: Performance in behavioral and practical subjects. Using individual register data, we find fairly strong and equal effects of the two types of skills on high school graduation probabilities. However, we find that "non-cognitive" skills has a much larger impact than "cognitive" skills on the probability to receive welfare benefits or being inactive (NEET) at age 22, while the findings are the opposite for the probability of college enrollment.

JEL-Code: I210, J240.

Keywords: skills, grades, high school graduation, NEET, welfare benefits.

Torberg Falch
Department of Economics
Norwegian University of Science and Technology
Trondheim / Norway
Torberg.Falch@svt.ntnu.no

Ole Henning Nyhus
Department of Economics
Norwegian University of Science and Technology
Trondheim / Norway
Ole.Nyhus@sof.ntnu.no

Bjarne Strøm
Department of Economics
Norwegian University of Science and Technology
Trondheim / Norway
Bjarne.Strom@svt.ntnu.no

January 2013
We thank Flavio Cunha, Erik Plug, Helena Svaleryd, and participants at the CESifo Economic Studies & UCLS Conference on Families, Children and Human Capital Formation for insightful comments.

1. Introduction

A number of studies in economics have found that students’ cognitive ability as measured by test scores in Mathematics and Science are important predictors of future earnings and other individual outcomes, see Hanushek (2002) for a review of the evidence. Moreover, recent cross-country studies suggest that aggregate measures of test scores are important determinants of economic growth and development (Hanushek and Woessmann, 2008). However, such measures can indirectly also capture personality traits such as motivation and conscientiousness. This has initiated a growing literature that emphasizes in more detail the skill formation process and in particular the role of non-cognitive skills, see Cunha and Heckman (2007). The evidence indicates that non-cognitive skills explain as much of the variation in earnings and employment prospects as cognitive skills. While cognitive skills traditionally are measured by some test score, measures of non-cognitive skills are usually based on self-reported survey data. Heckman et al. (2006) and Carneiro et al. (2007) estimate effects of indicators for loss of control and self-esteem, while a popular concept in the psychology literature is the five-factor model shaping human personality (Digman, 1990, Mueller and Plug, 2006, Borghans et al. 2008).

The reliance in the literature on self-reported information, combined with inherent difficulties in distinguishing conceptually between cognitive and non-cognitive skills, suggests that evidence based on other types of data can enhance the understanding of the role of different types of skill. The present paper has a different approach. Instead of relying on survey information, we use detailed information from transcript of records at the end of compulsory education at age 16 in Norway. Thus, we estimate the effect of the skills that are regarded as important by the school system and evaluated objectively by teachers. We distinguish between skills reflected in achievement in Mathematics and Science and skills reflected in performance in “behavioral and practical” subjects such as Arts and crafts and Physical education. By this approach, the observed skills cannot be directly compared to cognitive and non-cognitive skills as measured in the existing literature. For simplicity, however, we denote the former for “cognitive skills”, in line with several other studies, and the latter for “non-cognitive skills”. Teacher assessed grades in the latter subjects arguably reflect characteristics such as conscientiousness and openness to experience to a larger extent than other subjects. For example, teacher assessed grades in Physical education reflects engagement, preferences for following rules and motivation rather than just particular performance in sports.

We relate the skill measures to high school graduation, college enrollment, labor market attachment and the probability to receive welfare benefits for young adults using data for the cohorts leaving compulsory education in 2002–2004. We find that performance of young adults is strongly associated with both types of skills. Further, the estimated effect of one type of skill is biased by exclusion of the other type.

The empirical models include a rich set of individual characteristics, and the results are robust to the inclusion of different sets of school and neighborhood fixed effects. Although the results seem robust, we do not claim that estimated effects can be interpreted causally in the sense that an intervention increasing these skills has the same impact. This is a weakness our analysis shares with the literature on cognitive and non-cognitive skills in general.

The paper is organized as follows: Section 2 reviews related literature. Section 3 presents relevant institutional settings in Norway, the data, and the empirical specification. Empirical results are reported in Section 4, while Section 5 contains concluding remarks.

2. Related literature

Cognitive skills are associated with intelligence and the ability of problem solving. A number of papers have investigated the impact of such skills measured by test scores in Mathematics and Science on earnings and to some extent also on other individual outcomes. To take a few representative studies; Bishop (1989), Murnane et al (1995), and Altonji and Pierret (2001) all find that achievement measures are important determinants of individual earnings for given educational attainment and observed individual and family characteristics. Koedel and Tyhurst (2012) use a resume-based field experiment and find that stronger mathematical skills improve labor market outcomes.

Motivated by the micro-econometric evidence, some recent studies assess the role of cognitive skills for economic growth and development. Hanushek and Kimko (2000) and Hanushek and Woessmann (2008, 2009) all find that cognitive skills as measured by aggregate scores on international comparable student tests in Mathematics and Science have a sizeable positive effect on economic growth.

The literature on the role of cognitive skills has been challenged by authors arguing that some of the estimated effect of cognitive skills in reality captures the impact of non-cognitive skills. Non-cognitive skills are much more difficult to define and measure than cognitive skills. A popular taxonomy of non-cognitive skills is given by the five-factor model shaping human personality: agreeableness, conscientiousness, emotional stability, extraversion and autonomy. Extensive discussion of these concepts is given in Digman (1990), Mueller and Plug (2006) and Borghans et al. (2008).

Psychologists and sociologists have a long tradition in studying the role of non-cognitive skills in shaping individual behavior and outcomes using survey data. Jencks (1979) found that personal traits as leadership, industriousness, and perseverance had substantial impact on individual earnings and educational attainment holding family characteristics and cognitive skills constant. Recently, a number of papers by Heckman and coauthors have brought the role of non-cognitive skills to the forefront in the economics of education and skill formation literature. Heckman and Rubinstein (2001) provide an instructive example of the potential role played by non-cognitive skills. They show that recipients of degrees from the general education development program (GED) had lower wages, and less schooling than ordinary high school graduates, and comparable or even worse outcomes than high school dropouts, holding cognitive skills constant. Heckman et al. (2006) use US national Longitudinal Survey of Youth (NLSY) to estimate the impact of different skills on earnings, schooling and occupational choice within a structural latent factor model. They find that non-cognitive skills measured by self-reported indicators of loss of control and self-esteem strongly influence schooling decisions and wages. Carneiro et al. (2007) find similar results for the UK.

The studies above use self-reported survey data on non-cognitive skills. In addition to possible measurement error, self-reported measures may themselves be interpreted as outcomes. Two recent papers address this concern by using data on non-cognitive skills based on external evaluations. Lindquist and Vestman (2011) exploits that data from the Swedish military enlistment include a measure of non-cognitive skills based on an evaluation conducted by psychologists using individual interviews. They find that while cognitive skills measured by an IQ test is generally the most important determinant of male wages, non-cognitive skills turns out to be more important for low skilled workers and earnings below the median. Further, non-cognitive skills are more important than cognitive skills for the probability to receive unemployment support and social assistance.

Segal (2011) uses NELS data from the US and studies the impact of premarket teacher reported student misbehavior in eight grade (tardiness, absence, disruptiveness etc.) on male labor market outcomes and the probability to obtain a post-secondary degree. She finds that, controlling for test scores in mathematics and reading, educational attainment for males is negatively correlated with misbehavior. Her results are consistent with the findings in Lindquist and Vestman (2011) on the relationship between earnings and different types of skills.

Lindquist and Vestman (2011) and Segal (2011) only consider the impact of cognitive and non-cognitive skills on outcomes for males. Our data include the total student population, and allow us to investigate whether the return to different skill types differs by gender.

Our study is also related to the literature on the association between participation in physical activity and sports and individual outcomes, e.g., Barron et al (2000), Pfeifer and Cornelissen (2010) and Rees and Sabia (2010). Most of the studies find that participation in such activities increase school performance, years of schooling and future earnings even when controlling for cognitive skills. These results may reflect that non-cognitive skills are important factors explaining participation in physical activity and sports or that participation increase non-cognitive skills. A recent study by Rooth (2011) uses data from fictitious applications to real job openings in Sweden and finds that applicants signaling sports skills had a significantly higher callback rate. In the literature on returns to physical fitness it is typically attributed to non-cognitive skills.

3. Institutions, data, and empirical specification.

3.1. Institutions

The Norwegian school system consists of ten compulsory years. Students are normally enrolled the year they turn six years, and there is no possibility to fail a class. All students finish compulsory education 10 years after enrollment.

At graduation the students receive a diploma containing the different grades set by teachers and exam results, although for some of the weakest students grades may be missing in some subjects. Table 1 gives an overview of the relevant subjects. The grading system consists of a scale from one to six, where six is the highest grade. Teacher grades are based on the achievement throughout the 10th school year, but with largest weight on the latest tests and performance. They shall at the outset reflect skills and not effort. Thus, in subjects such as Mathematics and Science, the grades are to a large extent based on written tests conducted within the school year. Regarding other subjects, such as "behavioral and practical" subjects, casual evidence (e.g., Prøitz and Borgen, 2010) clearly indicates that effort and behavior matter for grading in addition to skills.

After the end of compulsory education, students can choose to leave school or to enroll in high school education. High school education consists of 15 different study tracks. Three of the study tracks qualify for higher education (academic tracks) and 12 study tracks give a certificate for work in a broad amount of occupations (vocational tracks). The academic tracks consist of three years, while the vocational study tracks normally consist of two years in school plus two years as apprentice.

Students have a legal right to five consecutive years of high school education after finishing compulsory school, and the government uses graduation within five year as the measure of the graduation rate in official statistics. Therefore, we use a five-year window in the empirical study of high school graduation below.

About 95 percent of the cohorts enroll high school the year they finish compulsory education. Students have to rank three different study tracks when applying for enrollment. All students have a legal right to be enrolled in one of these three tracks, but the actual study track and school they enroll into depends on achievement in compulsory education measured by their average grade. Despite the high initial enrolment rate, only around 70 percent of each cohort graduate within five years. A large fraction of the students drop out of high school education, which clearly is an important political concern.

Public schools have a common curriculum and the same number of teaching hours in each subject.[1] The 430 municipalities are responsible for compulsory education, while the 19 counties are responsible for high school education. The municipalities use about one-fifth of their budget on education, while the counties spend over 50 percent on education. Enrollment into compulsory schools is based on catchment areas, while the

[1] Few students enroll in private schools. About two and five percent of a cohort enroll in private compulsory schools and high schools, respectively.

counties have major leeway on enrollment rules for high schools. They determine the capacity of the individual schools and study tracks according to local needs and student demand. Some counties use catchment areas for the individual study tracks, other counties have free school choice within certain regions, while some do not have any restrictions on school choice.

3.2. Data

We use register data from Statistics Norway covering all students that finished compulsory education in the years 2002-2004. The 2002-cohort is the first cohort with grade information in the registers. To make the sample more homogeneous we only include students that turn 16 years of age the year they finish compulsory education in the empirical analysis.[2] In addition, we only include students with grade information on all relevant subjects and information on which compulsory school they graduated from. The data reduction is presented in Table 2. The analytical sample consists of 88.8 percent of the population, amounting to 154,515 observations.

We apply the teacher assessed grades to classify two different measures of skills. Following the economics literature, we denote the average grade in Science and Mathematics as "cognitive" skills. In "practical and behavioral" subjects, traits as conscientiousness, openness to experience, engagement and motivation are valued. We calculate the average of the grades in the subjects Food and health, Arts and crafts, Physical education, and Music, and for simplicity we denote this variable "non-cognitive" skills. Recognizing that the same types of skills might improve the grades in several different subjects, isolating the impact of one type of skill, say, cognitive skills, requires that the model condition on other types of skills. Analyses that only include one skill measure might overestimate the importance of this type of skill.[3]

---

[2] Since no students fail any grade in Norwegian compulsory education, one could expect that all students turn 16 years of age the year they finish compulsory education. However, there are some exceptions. If a child is not considered to be mature enough, the parents together with the school and psychologists can postpone enrollment one year. It is also possible to start one year ahead the birth cohort. In addition, some older students return to improve their grades, and immigrants are often over-aged.

[3] The correlation coefficient between our skill measures is 0.74, which clearly indicates that there are some common characteristics important for performance in both classes of skill. An alternative classification of skills could be based on a principal component analysis. However, in our view, relying on averages of observable skill measures regarded as important by the school

Mean values and standard deviations for the subjects are presented in Table 1. The average grade is lowest in Mathematics and highest in Physical education. While 23 percent of the students obtain grade 1 or 2 in Mathematics, that is the case for less than 6 percent in Physical education, Food and health, Arts and craft, and Music. Thus, the standard deviation of the mean grade in Mathematics and Science is larger than the standard deviation of mean grade in the "non-cognitive" subjects. To facilitate interpretation, we use standardized values with mean zero and standard deviation equal to unity in the empirical analyses. The distributions of the standardized variables are presented in Figure 1. While the distribution of "cognitive" skills is close to the normal distribution, some individuals have "non-cognitive" skills more than three standard deviations below the mean.[4] The lower part of Figure 1 present separate distributions for females and males. Measured by grades, the distributions of skills for females are to the right of the distributions for males, and the difference is most pronounced for "non-cognitive" skills. [5]

Table 3 gives a description of educational and labor market outcomes. About 57 percent graduate within expected time, while additionally about 14 percent graduate delayed but within 5 years after the end of compulsory education. There are only small differences across the cohorts. Figure 2 present the distribution of skills for graduates and dropouts. Panel A shows that the distribution of "cognitive" skills for individuals that graduate high school within five years is clearly to the right of the distribution for dropouts. Panel B shows a similar picture for "non-cognitive" skills.[6]

We measure enrollment in higher education in the fall five years after the individuals' finished compulsory education. Since the expected time to graduate high school in an academic track that qualify for higher education is three years, the individuals might

---

system makes interpretation easier than using measures based on a purely data based principal component analyses. In addition, by this approach our results are comparable to previous studies on the impact of cognitive skills.

[4] Since the distribution of our measure of "non-cognitive" skills is skewed to the left, the standardization might in principle affect the estimated effects. We show below that this is not the case in the present analysis.

[5] The skill variables are discrete, but the presentation of the distributions in the lower part of Figure 1 and Figures 2-3 below is smoothed by the choice of bandwidth. Mean values of standardized "cognitive" skills are equal to 0.11 and -0.11 for females and males, respectively. The corresponding numbers for "non-cognitive" skills are 0.25 and -0.24.

[6] Mean values of standardized "cognitive" skills are 0.32 and -0.79 for graduates and dropouts, respectively. The corresponding numbers for "non-cognitive" skills are 0.31 and -0.76.

do military services or other activities between graduating high school and our measure of higher education participation. Table 3 shows that the enrollment share in higher education at age 21 is slightly increasing in the empirical period and is on average 37 percent.

Regarding labor market attachment, we follow the students up to age 22. Since employment is registered on a daily basis, we measure inactivity on a specific day (October 15th). Inactivity is defined as not registered in employment, education or training (NEET).[7] Table 3 shows that 16.3 percent is NEET this day. The distributions of skills for both inactive and active individuals are presented in Figure 3, which shows the same pattern as in Figure 2. The distribution of both types of skills for non-NEET individuals is clearly to the right of the distribution for NEET individuals.

In addition to the inactivity outcome, we study the probability of receiving welfare benefits. We utilize information on the number of months receiving benefits during the calendar year and use the share of the months with benefits in the analysis. Table 3 shows that on average 1.9 percent are on welfare in a random month the year they turn 22. During this year, about five percent of the sample receive welfare benefits at least one month.[8]

The last two columns in Table 3 present mean values of the outcome variables separately for females and males. Females perform better than males on all outcomes. They are more likely to graduate high school and to enroll higher education, and less likely NEET and to receive welfare benefits.

Individual characteristics and family background are well documented as factors affecting individual outcomes. In the empirical analysis we include variables for gender, immigrant status, birth quartile, domestic mobility, whether the student need support related to diseases and disabilities, parental education, parental income, parental employment status, and parental marital status. The dummy variable for domestic mobility is defined as living in different municipalities at age 7, 13 or 16.

[7] The data available for this project do not include education data for the fall 2010. Thus, we cannot analyze the NEET-outcome for the 2004 cohort.

[8] The mean values of the two variables for labor market attachment differ markedly across high school graduates and high school dropouts. On average, 36.7 percent of the dropouts are not registered in employment or education at October 15 the year they turn 22 years of age, compared to 9.9 percent of the graduates. Regarding welfare participation, the corresponding numbers are 6.9 and 0.4 percent.

Parental education is classified into four levels (only compulsory education; graduated from high school; bachelor degree; master or PhD degree), and our measure is based on the education for the parent with the highest education level. Parental income is represented by the level of taxable income and its square. For marital status, two dummy variables are included in the model; one if parents are married when the student graduate from compulsory education and one if the parents are divorced at that time. 61.5 percent of the parents were registered as married, 12.5 percent were registered as divorced and 26 percent had never been married. Descriptive statistics are presented in Appendix Table A1.

3.3. Empirical specification

We use the regression model in equation (1) to estimate the impact of different skills on the following outcomes; high school graduation, enrolling higher education, NEET, and welfare participation. The outcome is a binary variable, Y_{ijc}, for student *i* from compulsory school *j* in cohort *c*. The variable *cog* is defined as the mean grade in Mathematics and Science, and *noncog* is the mean grade in Physical education, Food and health, Arts and craft, and Music. X_{ic} is a vector of individual characteristics. In addition, the model includes interaction between fixed effects for cohort (θ_C) and the compulsory school from which the student graduated (γ_j). These fixed effects control for all systematic differences in grading practices between schools and cohorts as well as other unobserved school and cohort characteristics that potentially affect high school graduation. Below, we also present results for alternative sets of fixed effects, including detailed neighborhood fixed effects. ε_{ic} is a random error term.

$$(1) \quad Y_{ijc} = \alpha + \beta_1 cog_{ic} + \beta_2 noncog_{ic} + X_{ic}'\delta + \gamma_j * \theta_c + \varepsilon_{ic}$$

As robustness checks, we provide estimates using more flexible model formulations. We estimate non-parametric specifications and models with skills represented by grades in separate subjects rather than average grades in subject categories.

4. Empirical results

We begin with an analysis of the probability to graduate from high school, with the main emphasis on graduation within five years after the end of compulsory education. We also decompose this outcome into graduation on-time and delayed and conduct separate analysis on each component. The last part of the section presents results for the labor market attachment at age 22.

4.1. Educational outcomes

Table 4 presents the estimated effects of skills on the probability to graduate from high school within five years after the end of compulsory education (the year the individuals turn 21). The first column presents the correlation between skills and graduation. Increasing "cognitive" skills with one standard deviation is associated with 14.3 percentage points (20 percent of the sample mean) higher graduation probability, while a similar change in "non-cognitive" skills is associated with 11.7 percentage points (17 percent) higher graduation probability. These are large effects.

Column (2) in Table 4 includes socioeconomic characteristics. Even though several of the socioeconomic characteristics are strongly related to the probability to graduate (see below), the effects of the skill measures are only marginally reduced when they are included.

However, the skill variables are correlated. Column (3) shows that the effect of "cognitive" skills increases to 20.5 percentage points when "non-cognitive" skills is excluded from the model, an increase of 55 percent. This result clearly indicates that estimates of cognitive skills as measured by achievement in mathematics and science are biased in models not taking other kinds of skills into account. Similarly, excluding "cognitive" skills from the model as in column (4) increases the effect of "non-cognitive" skills by 87 percent.[9]

One concern is that grading standards can vary across schools and introduce biased estimates of the coefficients. In this case we would expect the coefficients in column (2) in Table 4 to be underestimated. This concern is probably of less importance for our measure of cognitive skills since external exit exams in Mathematics provide a check on grading practices. Systematic differences in grading standards might be a larger concern for our measure of "non-cognitive" skills. If variation in grading practices is merely across schools rather than across classrooms within schools, school fixed effects will reduce the potential bias. The model in column (5) includes school fixed

[9] In models with cohort by school fixed effects (model specifications similar to the model in column 6 in Table 4), the biases are of similar size (62 and 80 percent for cognitive and non-cognitive skills, respectively).

effects, which increases the effect of non-cognitive skills, but does not change the effect of cognitive skills. On the other hand, if grading practices varies substantially across classrooms within schools, we would expect cohort by school fixed effects to further increase the estimated effects since Norwegian teachers usually teaches the same group of students in several years. However, the results when including cohort by school fixed effects (column 6 in Table 4) are basically identical to the previous ones. Finally, in order to control for unobserved neighborhood effects, the model in column (7) in Table 4 includes cohort specific ward fixed effects. This model specification accounts for detailed neighborhood characteristics since, in each cohort, the average number of students in the ward is only 4.8. Again, the effects of the skill variables are similar to those obtained in the simpler specifications.

The full results for the model in column (6) in Table 4 are reported in Column (2) in Appendix Table A1. The estimated effects of the socioeconomic characteristics are mainly as expected. Conditional on skills, individuals with married and working parents with some post-compulsory education have a higher probability to graduate. The effect of income is positive except for the students with the very highest parental income. Females and individuals born late in the year also have a higher graduation probability, where the latter effect must be interpreted as a catch-up effect since those born late appears to have lower school achievement at younger ages (Bedard, 2006).[10] In addition, mobility during compulsory education and disability status is negatively related to graduation, while immigrants have a higher graduation probability than native Norwegians. The latter effect is critically dependent on the inclusion of parental characteristics in the model.

The effects of skills may be non-linear. To explore this issue, Figure 4 presents estimated skill effects from a model where the continuous skill variables are replaced by dummy variables for the observed values of these discrete variables. Otherwise the model is identical to the model in column (6) in Table 4. The figure presents estimated skill effects relative to students with normalized skills equal to zero (mean skills).[11] For both skill variables, the effects are small in both tails of the skill distribution, and close to linear for skill levels with the highest density; in the range -2 to 1 in the standardized distribution of skills.

[10] This is the case also for both our measures of skills in the present paper.

[11] The confidence intervals are small and not shown in the figure. Due to the small confidence intervals, null hypotheses of linear effects are clearly rejected.

One concern is that the "non-cognitive" skill measure contains more information in the lower end of the distribution than the "cognitive" skill measure. The combination of this fact and the use of standardized measures may to some extent explain the slightly larger effect of "non-cognitive" than "cognitive" skills on the graduation. However, sensitivity analyses suggest that this is not the case. If we replace the standardized variables with the average grades, the effects in terms of standard deviations are identical to the results reported in Table 4. In addition, if we disregard the detailed information in the left tail of "non-cognitive" skills, we also get identical results.[12]

Another issue is that the classification of skills into two categories implies that we do not utilize all the information available in the transcript of records at the end of compulsory education. Table 5 presents estimates using more flexible models and evaluates to what extent grades in individual subjects matter for high school graduation. The first two columns presents results for models allowing for independent effects of each of the subjects used to calculate our measures of skills above. Standardized values are used for all grades.

In the model without fixed effects (column (1)), the effect of Science (7.5 percentage points) is slightly larger than the effect of Mathematics (6.5 percentage points). Notice that in this conditional model the implicit scale is different from the model above since increasing the grade in Mathematics by one grade, holding the grade in Science constant, only increases our measure of "cognitive" skills by 0.5. Thus, the sum of the effects of Mathematics and Science in Table 5 is close to the effect of "cognitive" skills in Table 4. For every "non-cognitive" subject, the effect is smaller than for the "cognitive" subjects. Since the effects are very precisely estimated, we formally reject the hypothesis of equal effects of all four "non-cognitive" subjects even though the effects are not very different in numerical terms. Column (2) in Table 5 shows that taking unobserved heterogeneity across schools and cohorts into account increases

[12] We disregard the detailed information in the left tail of the "non-cognitive" skills (*noncog*) distribution in two ways; (i) excluding the 1,743 observations with *noncog* below the minimum value of "cognitive" (*cog*) skills (which is equal to -2.56) and (ii) replacing all values of *noncog* for these observations with the lowest value of *cog*. In addition, the results are also robust to excluding all outliers defined as values of *cog* or *noncog* larger than two in absolute value (10,652 observations). In this case, the effects of *cog* and *noncog* in the model formulation with cohort by school fixed effects (column 6 in Table 4) are 0.136 and 0.114, respectively, which are changes of less than five percent compared to Table 4.

the effects of all grades except the grade in Mathematics, which is the only of these subjects with an external written exit examination.

The models in columns (3) and (4) in Table 5 include the grades in all subjects in the transcript of record. In addition to the subjects above, that includes three different grades in Norwegian language, two grades in English language, Religious and ethical education, and Social studies (including history).[13] The effects of both oral and written English are insignificant, and the effects of the three grades in Norwegian language are small. Language skills do not seem to be important for the probability to graduate high school, conditional on the other grades. The grades in Religious and ethical education and Social studies, however, have significant effects, and in particular the performance in Social studies has predictive power on the probability to graduate. In our view, it is, however, hard to classify the skills inherent in these subjects compared to the other subjects. To simplify the exposition of the paper, we thus restrict the succeeding analyses to “cognitive” and “non-cognitive” skills as defined above. Notice that by this approach the impact of skills inherent the other subjects is partly taken into account by the positive correlation between the grades in the different subjects. That is visualized in Table 5, by the fact that the effects of the “cognitive” and “non-cognitive” subjects decline by 18-47 percent when the grades in the other subjects are included.

The graduation outcome used so far may be decomposed into graduating on time and graduating delayed, but within five years after compulsory education. Most students graduate on-time (three or four years after compulsory education, depending on study track). Panel B in Table 6 presents results for this outcome. Compared to the results for graduation within five years (replicated in panel A in the table), “cognitive” skills seem slightly more important and “non-cognitive” skills slightly less important, but the differences are small. Panel C in Table 6 shows results for the probability to graduate delayed. In this case we restrict the sample to individuals not graduating on-time. Both skill measures clearly increase the probability to graduate also in this case. The estimated effects in percentage points are smaller than the effects for graduating on-time. Increasing “cognitive” skills by one standard deviation increases the probability to graduate on time and delayed by 16 and 11 percentage points, respectively. However, since the share of students graduating on time is much higher than the share graduating delayed, the effect on graduating delayed is largest in relative terms (the

[13] Since information is missing for some individuals for some of these additional subjects, the number of observations declines by 1.3 percent compared to the former models.

effects are 27 and 36 percent, respectively, evaluated at sample means). For both graduation on-time and graduation delayed, the bias by excluding one skill variable from the model is similar to the models above.[14]

The final education outcome is enrollment in higher education five years after the end of compulsory education. Panel D in Table 6 shows that the effect of "cognitive" skills is almost four times larger than the effect of "non-cognitive" skills for college enrollment. Although the effect of the latter skill variable is relatively small, it is significant at one percent level. While this result is as expected, the effects are clearly non-linear as shown in Panel B in Figure 4. There is no effect of "cognitive" skills for the 15 percent of the individuals with lowest skills (weaker than one standard deviation below mean), while the effect of "non-cognitive" skills is concentrated to individuals with skills in the range ± 1 standard deviation from mean. We note that this group of students is likely to be on the margin of enrolling into higher education or not as the share of the cohort enrolling is 37 percent.

An interesting issue is to what extent the skill effects on the education outcomes vary by gender. The last two columns in Table 6 estimate separate effects for females and males. Although the gender differences are relatively small, it turns out that the effect of "cognitive" skills is higher for males than for females for all outcomes. The opposite is the case for "non-cognitive" skills, except for delayed graduation. While these patterns are interesting, our data does not enable us to further explore why the skill effects on these outcomes differ between males and females.

4.2. Labor market outcomes

Table 7 presents the estimated effects of skills on the probability of receiving welfare benefits and inactivity (NEET). Regarding the probability of being on welfare, the effect of "non-cognitive" skills is 4-5 times larger than the effect of "cognitive" skills, in contrast to the results for enrollment in higher education. The results in the fixed effects specification imply that, on average, increasing "non-cognitive" skills by one standard deviation decreases the probability of being on welfare by 1.96 percentage points, i.e., 102 percent of the average rate of welfare participation. This is indeed a

[14] For the models without school fixed effect (column 1 in Table 6), the bias in the effect of "cognitive" skills of excluding "non-cognitive" skills from the model is 43 percent for both graduation on-time and graduation delayed. For the model including cohort by school fixed effects (column 2), the bias is 46 and 48 percent, respectively.

large effect, and it is nonlinear as shown in Panel A in Figure 5. The small effect when "non-cognitive" skills exceed -0.5 is likely to reflect the fact that very few individuals with high skills receive welfare benefits. If we disregard the detailed information in the left tail of this skill variable, the estimated coefficients change slightly, but the qualitative results remain. Using the same robustness checks as above, the largest change is for the model where we exclude all observations with skill measures larger than two in absolute value. In this case the effect of "cognitive skills" increases from -0.0038 to -0.0047 and the effect of "non-cognitive" skills decreases from -0.0196 to -0.0133.[15]

Panel B in Table 7 shows that the effect of "non-cognitive" skills is larger than the effect of "cognitive" skills also for the probability of being inactive (NEET). The difference in effect size is, however, smaller than for being on welfare. In the fixed effects model, the effects of increasing skills by one standard deviation are -5.5 and -4.0 percentage points (34 and 24 percent), respectively. These relationships are close to linear as shown in Panel B in Figure 5.[16]

The change in the estimated effect of "cognitive" skills of excluding the variable for "non-cognitive" skills from the model is larger for the outcomes in Table 7 than for high school graduation (not reported in the table). This must be related to the finding that the latter skill variable is much more important for labor market attachment than the former.[17]

The last two columns of Table 7 presents separate models for males and females. The results are qualitatively similar to the gender differences for the educational

[15] If we replace the standardized variables with the average grades, the effects in terms of standard deviations are identical to the results reported in Table 7. When excluding the 1,743 observations with "non-cognitive" skills (*noncog*) below the minimum value of "cognitive" skills (*cog*), or replace these values with the minimum value of *cog*, the effect of *cog* increases by about 20 percent and the effect of *noncog* decreases by about 15 percent, which imply that the effect of *noncog* is about 3.5 times larger than the effect of *cog*.

[16] The estimated coefficients are insensitive to the differences in the distribution of the skill variables. Using the same approaches as above, the effect of *cog* varies from -0.0404 to -0.0413 and the effect of *noncog* varies from -0.0474 to -0.0551.

[17] The change in the estimated effect of "cognitive" skills of excluding "non-cognitive" skills from the model is in the order of 350 and 100 percent for being on welfare and NEET, respectively. On the other hand, the change in the estimated effect of "non-cognitive" skills of excluding "cognitive" skills from the model is much smaller, about 15 and 50 percent for being on welfare and NEET, respectively.

outcomes. In general, the gender differences are relatively small, but the effect of "cognitive" skills is higher for males than for females and the effect of "non-cognitive" skills is higher for females than for males.

Overall, both classes of skill seem to be important for labor market attachment for young adults. The finding that "non-cognitive" skills are more important than "cognitive" skills is similar to the Swedish evidence for males in Lindquist and Vestman (2011). It is interesting to note that we reach a similar conclusion using clearly different measures of skills compared to their study. In addition, we find a similar pattern for females as for males.

5. Concluding remarks

This paper investigates the impact of different types of skills on educational outcomes and labor attachment for young adults. We use detailed grade transcripts from compulsory education in Norway at age 16, and measure "cognitive" skills by average grades in Mathematics and Science and "non-cognitive" skills by average grades in practical and behavioral subjects.

We find that both classes of skill are of roughly equal importance for the probability to graduate from high school, and that the effect sizes depend on whether the measures of both classes of skill are included in the model or not. "Cognitive" skills are of much more importance than "non-cognitive" skills for college enrollment, while the opposite is the case for labor market attachment. Interestingly, this pattern is in accordance with results from other studies using measures of cognitive and non-cognitive skills very different from ours. It seems like cognitive skills are most important for educational outcomes, while non-cognitive skills are most important for the probability of exclusion. In addition, the evidence in this paper indicates that the impact of cognitive skills in general is larger for males than for females, while the impact of non-cognitive skills in general is larger for females than for males.

References

Altonji, J. and C. P. Pierret (2001): Employer learning and statistical discrimination. *Quarterly Journal of Economics* 116, 313-50.

Barron, J. M., B.T. Ewing and G. R. Waddell (2000): “The effects of high school athletic participation on education and labor market outcomes”. *Review of Economics and Statistics 82*, 409-421.

Bedard, K. and E. Dhuey (2006): The persistence of early childhood maturity: International evidence of long-run age effects, *Quarterly Journal of Economics, 121* 1437-1472.

Bishop, J. H. (1989): Is the test score decline responsible for the productivity growth decline? *American Economic Review* 79, 178-97.

Borghans, L., A. Duckworth, J. Heckman and B. ter Weel (2008): The economics and psychology of personality traits. *Journal of Human Resources 43*, 972-1059.

Brunello, G. and M. Schlotter (2011): Non cognitive skills and personality traits: Labour market relevance and their development in E&T systems. EEENE Analytical Report no 8. European Commission, Education and Culture.

Carneiro, P., C. Crawford and A. Goodman (2007): The impact of early cognitive and non-cognitive skills on later outcomes. CEE DP.92.

Cunha, F. and J. Heckman (2007): The technology of skill formation. *American Economic Review 97*, 31-47.

Digman, J.M. (1990): Personality structure: Emergence of the five-factor model. *Annual Review of Psychology 41*, 417-440.

Hanushek, E. A. (2002): Publicly provided education. In Auerbach, A. J. and M. Feldstein (eds): *Handbook of Public Economics, Volume 4,* Elsevier Science, North-Holland.

Hanushek, E. A. and D. Kimko (2000): Schooling, labor-force quality and the growth of nations. *American Economic Review 90*, 1184-1208.

Hanushek, E. A. and L. Woessmann (2008): The role of cognitive skills in economic development. *Journal of Economic Literature 46*, 607-668.

Heckman, J. and Y. Rubinstein (2001): The importance of noncognitive skills: Lessons from the GED testing program. *American Economic Review 91*, 145-149.

Heckman, J., N. Stixrud and S. Urzua (2006): The effects of cognitive and noncognitive abilities on labor market outcomes and social behavior. *Journal of Labor Economics 24*, 411-482.

Jencks, C. (1979): *Who gets ahead? The determinants of success in America*. New York: Basic Books.

Koedel, C. and E. Tyhorst (2012): Math skills and labor-market outcomes: Evidence from a resume-based field experiment. *Economics of Education Review 33*, 131-140.

Lindquist, E. and R. Vestman (2011): The labor market returns to cognitive and noncognitive ability: Evidence from the Swedish Enlistment. *American Economic Journal: Applied Economics 3*. 101-128.

Mueller, G. and E. Plug (2006): Estimating the effect of personality on male and female earnings. *Industrial and Labor Relations Review 60*, 3-22.

Murnane, R. J., J. B. Willett and F. Levy (1995): The growing importance of cognitive skills. *Review of Economics and Statistics 77*, 251-66.

Pfeifer, C. and T. Cornelissen (2010): “The impact of participation in sports on educational attainment-New evidence from Germany”. *Economics of Education Review 29*, 94-103.

Prøitz, T. S. and J. S. Borgen (2010): «Rettferdig standpunktvurdering – det (u)muliges kunst? Læreres setting av standpunktkarakterer i fem ulike fag i grunnopplæringen». NIFU STEP Report 16/2010

Rees, D. I. and J. J. Sabia (2010): “Sports participation and academic performance: Evidence from the National Longitudinal Study of Adolescent Health”. *Economics of Education Review 29*, 751-759.

Rooth, D-A. (2011): Work out or out of work-The labor market return to physical fitness and leisure sports activities. *Labour Economics 18*, 399-409.

Segal, C. (2011): Misbehaviour, education, and labor market outcomes. *Journal of the European Economic Association*, forthcoming.

Table 1. Description of subjects in compulsory education, and descriptive statistics on grades

Subject	Description	Mean value (Std. dev.)
Mathematics		3.49 (1.12)
Science	Science and the environment	3.95 (1.11)
Physical education	Gymnastics, sports, etc.	4.35 (0.96)
Food and health (Home economics)	Food and lifestyle, food and culture, and food and consumption	4.35 (0.84)
Arts and crafts	Visual communication, design, art and architecture	4.23 (0.91)
Music	Making music, composition, and listening	4.22 (0.97)

Table 2. Data reduction

	Observations	Percent
Finish compulsory education in 2002-2004	174,067	100.0
Not turning 16 years the year finishing compulsory education	10,059	5.8
Missing grade information*	8,883	5.1
Missing compulsory school identifier	610	0.4
Analytical sample	154,515	88.8

* Missing information for at least one of the six subjects used to calculate our measures of cognitive and non-cognitive skills.

Table 3: Descriptive statistics for high school graduation and labor market attachment

	2002	2003	2004	All	Females	Males
Graduating within 5 years, percent	70.8	70.6	71.2	70.8	75.6	66.3
On-time graduation, percent	57.3	56.8	57.0	57.1	64.5	49.9
Graduating delayed but within 5 years, percent	13.4	13.7	14.2	13.8	11.0	16.4
Enrolled in higher education at age 21, percent	36.5	36.6	37.8	37.0	45.8	28.5
NEET October 15th at age 22, percent	15.9	16.7	-	16.3	15.1	17.4
Share of month on welfare at age 22, percent	1.66	2.01	2.11	1.93	1.81	2.05
Observations	49,056	50,884	54,575	154,515	75,778	78,737

Table 4. The effects of skills on high school graduation within 5 years

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Cognitive skills	0.143*	0.132*	0.205*	-	0.129*	0.130*	0.132*
	(0.0019)	(0.0019)	(0.0014)		(0.0018)	(0.0017)	(0.0018)
Non-cognitive skills	0.117*	0.106*	-	0.198*	0.116*	0.117*	0.115*
	(0.0018)	(0.0019)		(0.0014)	(0.0017)	(0.0017)	(0.0018)
Socioeconomic characteristics	No	Yes	Yes	Yes	Yes	Yes	Yes
School FE (no of groups)	0	0	0	0	1,186	-	-
Cohort x school FE (no of groups)	0	0	0	0	0	3,349	-
Cohort x ward FE (no of groups)	0	0	0	0	0	0	32,319
R-squared	0.285	0.300	0.277	0.264	0.297	0.299	0.289
Observations	154,515	154,515	154,515	154,515	154,515	154,515	154,447

Note. Standard errors in parentheses are clustered by compulsory school, * denotes significance at one percent level.

Table 5. Subject specific effects. Dependent variable is graduation from high school within 5 years

	(1)	(2)	(3)	(4)
Mathematics	0.065*	0.059*	0.046*	0.039*
	(0.0019)	(0.0019)	(0.0019)	(0.0019)
Science	0.076*	0.082*	0.041*	0.045*
	(0.0021)	(0.0020)	(0.0023)	(0.0022)
Physical education	0.043*	0.044*	0.035*	0.037*
	(0.0015)	(0.0014)	(0.0015)	(0.0014)
Food and health (Home economics)	0.026*	0.032*	0.019*	0.024*
	(0.0017)	(0.0015)	(0.0017)	(0.0015)
Arts and crafts	0.034*	0.038*	0.027*	0.031*
	(0.0016)	(0.0016)	(0.0016)	(0.0016)
Music	0.032*	0.035*	0.017*	0.019*
	(0.0018)	(0.0016)	(0.0018)	(0.0017)
Norwegian, written			0.006*	0.009*
			(0.0021)	(0.0020)
Second Norwegian language, written			0.001	-0.002
			(0.0019)	(0.0018)
Norwegian, oral			0.016*	0.017*
			(0.0020)	(0.0020)
English, written			-0.000	-0.001
			(0.0020)	(0.0020)
English, oral			-0.003	-0.003
			(0.0021)	(0.0020)
Religious and ethical education			0.031*	0.032*
			(0.0023)	(0.0022)
Social studies			0.044*	0.044*
			(0.0021)	(0.0020)
Socioeconomic characteristics	Yes	Yes	Yes	Yes
Cohort x school FE (no of groups)	0	3,349	0	3340
R-squared	0.300	0.299	0.305	0.304
Observations	154,515	154,515	152,468	152,468

Note. Same model socioeconomic characteristics as in the models in Table 4. In all models the sample is restricted to individuals with grade information in all the 13 subjects, except the second Norwegian language for which the model include an indicator for missing value. Standard errors in parentheses are clustered at the school level. * denotes significance at one percent level.

Table 6. The effect of skills on educational outcomes

	(1)	(2)	(3)	(4)
Sample	All	All	Females	Males
A. Graduation within 5 years				
Cognitive skills	0.132*	0.130*	0.114*	0.144*
	(0.00192)	(0.00171)	(0.0024)	(0.0023)
Non-cognitive skills	0.106*	0.117*	0.128*	0.110*
	(0.00185)	(0.00170)	(0.0025)	(0.0023)
Observations	154,515	154,515	75,778	78,737
B. On-time graduation				
Cognitive skills	0.156*	0.157*	0.145*	0.165*
	(0.0019)	(0.0018)	(0.0026)	(0.0024)
Non-cognitive skills	0.097*	0.106*	0.125*	0.094*
	(0.0020)	(0.0018)	(0.0027)	(0.0024)
Observations	154,515	154,515	75,778	78,737
C. Delayed graduation, but within 5 years				
Cognitive skills	0.110*	0.112*	0.099*	0.121*
	(0.0030)	(0.0029)	(0.0050)	(0.0037)
Non-cognitive skills	0.064*	0.073*	0.070*	0.075*
	(0.0024)	(0.0025)	(0.0042)	(0.0032)
Observations	66,348	66,348	26,877	39,471
D. College enrollment				
Cognitive skills	0.173**	0.178**	0.175**	0.179**
	(0.0017)	(0.0017)	(0.0027)	(0.0023)
Non-cognitive skills	0.048**	0.052**	0.079**	0.030**
	(0.0017)	(0.0017)	(0.0028)	(0.0022)
Observations	154,515	154,515	75,778	78,737
Socioeconomic characteristics	Yes	Yes	Yes	Yes
Cohort x school FE	No	Yes	Yes	Yes

Note. The same socioeconomic characteristics as in the models in Table 4 are included. Standard errors in parentheses are clustered at the school level, * denotes significance at one percent level.

Table 7. The effect of skills on lack of labor market attachment at age 22, 2002-cohort

	(1)	(2)	(3)	(4)
Sample	All	All	Females	Males
A. On welfare				
Cognitive skills	-0.0044*	-0.0038*	-0.0035*	-0.0039*
	(0.0004)	(0.0004)	(0.0006)	(0.0006)
Non-cognitive skills	-0.0180*	-0.0196*	-0.0210*	-0.0190*
	(0.0006)	(0.0007)	(0.0010)	(0.0009)
Observations	154,515	154,515	75,778	78,737
B. Inactive				
Cognitive skills	-0.0418*	-0.0399*	-0.0423*	-0.0368*
	(0.0019)	(0.0019)	(0.0028)	(0.0026)
Non-cognitive skills	-0.0484*	-0.0549*	-0.0602*	-0.0520*
	(0.0020)	(0.0021)	(0.0031)	(0.0029)
Observations	99,940	99,940	49,003	50,937
Socioeconomic characteristics	Yes	Yes	Yes	Yes
Cohort x school FE	No	Yes	Yes	Yes

Note. The same socioeconomic characteristics as in the models in Table 4 are included. Standard errors in parentheses are clustered at the school level, * denotes significance at one percent level.

Figure 1. The distribution of cognitive and non-cognitive skills

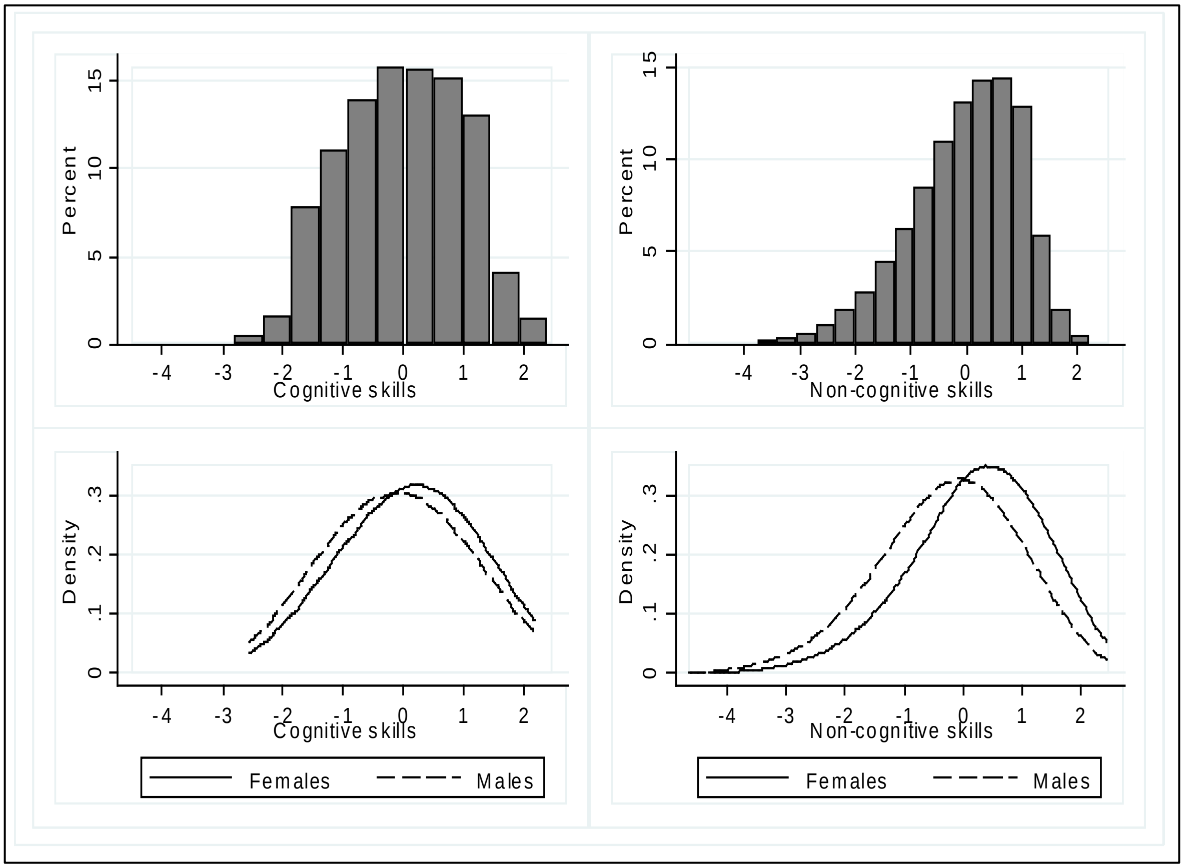
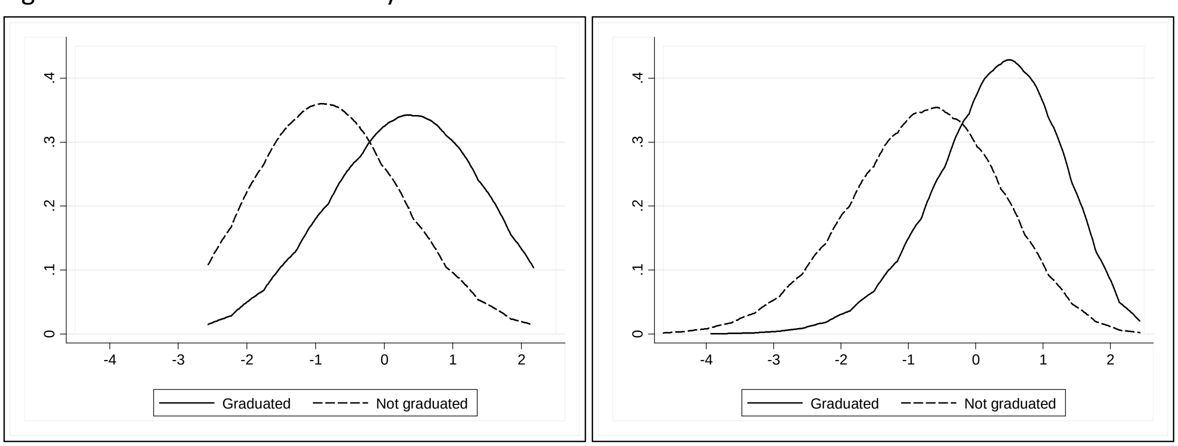


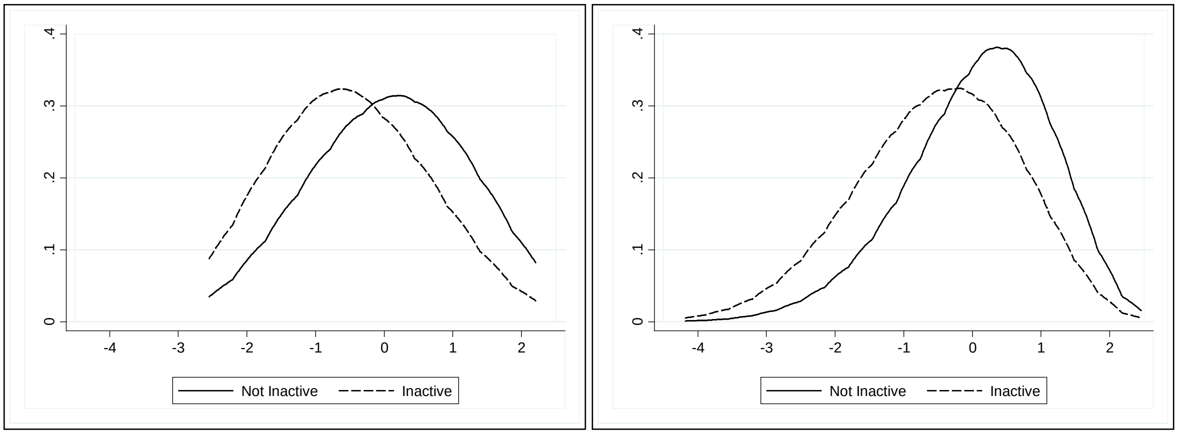
Figure 2. Graduation within five years and skills



Panel A. Cognitive skills

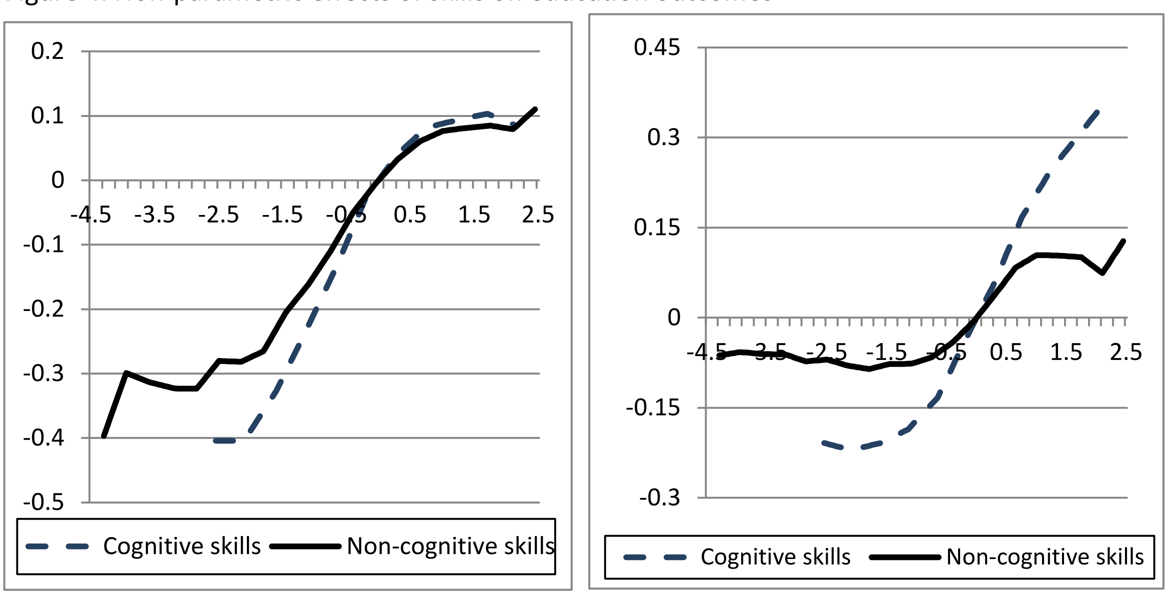
Panel B. Non-cognitive skills

Figure 3. Inactivity and skills



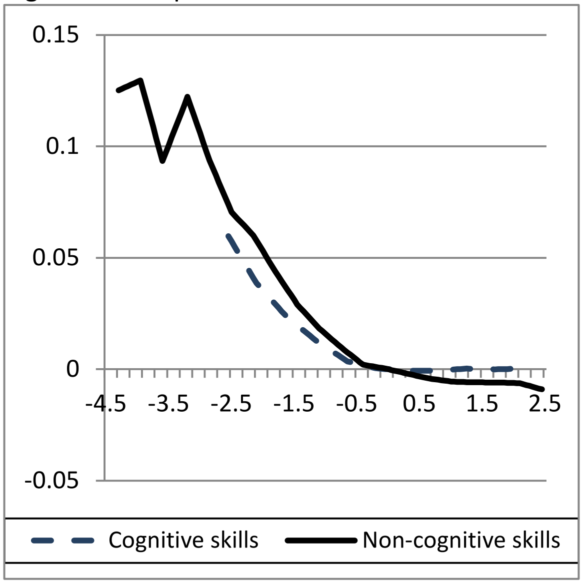
Panel A. Cognitive skills Panel B. Non-cognitive skills

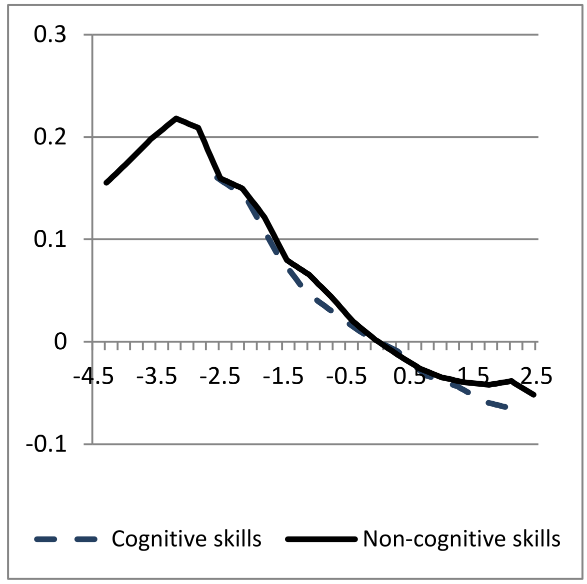
Figure 4. Non-parametric effects of skills on education outcomes



Panel A. High school graduation Panel B: College enrollment

Figure 5. Non-parametric effects of skills on labor market attachment





Panel A. On welfare

Panel B: NEET

Appendix Table A1. Descriptive statistics and full model results

	(1) Mean values	(2) Graduation	(3) On welfare	(4) NEET
Cognitive skills	0.00 [1.00]	0.130*	-0.0038*	-0.0399*
		(0.00171)	(0.0004)	(0.0019)
Non-cognitive skills	0.00 [1.00]	0.117*	-0.0196*	-0.0549*
		(0.00170)	(0.0007)	(0.0021)
Female	0.49	0.00587*	0.0080*	0.0151*
		(0.00226)	(0.0006)	(0.0027)
First generation immigrant	0.03	0.0454*	-0.0133*	-0.0029
		(0.00723)	(0.0025)	(0.0091)
Second generation immigrant	0.02	0.0515*	-0.0182*	-0.0130
		(0.00825)	(0.0019)	(0.0100)
Parents' highest educational level is	0.47	0.0437*	-0.0117*	-0.0230*
high school education		(0.00358)	(0.0013)	(0.0044)
Parents' highest educational level is	0.29	0.0577*	-0.0095*	-0.0219*
bachelor degree		(0.00385)	(0.0013)	(0.0047)
Parents' highest educational level is	0.1	0.0418*	-0.0059*	-0.0173*
master or PhD		(0.00452)	(0.0013)	(0.0055)
Benefits due to disease before the	0.02	0.0108	0.0005	0.0142
age of 18		(0.00835)	(0.0030)	(0.0112)
Benefits due to disabilities before the	0.02	-0.0557*	0.0127*	0.0661*
age of 18		(0.00774)	(0.0029)	(0.0105)
One parent employed	0.24	0.0399*	-0.0258*	-0.0295*
		(0.00546)	(0.0026)	(0.0071)
Both parents employed	0.71	0.0706*	-0.0337*	-0.0582*
		(0.00555)	(0.0025)	(0.0072)
Parental income in 100,000 NOK	6.05 [3.98]	0.00120*	-0.0004*	-0.0014*
		(0.000328)	(0.0001)	(0.0004)
Parental income in 100 NOK squared	52.5 [995.2]	-0.0023*	0.0000*	0.0000
		(7.72e-07)	(0.0000)	(0.0000)
Married parents	0.61	0.0511*	-0.0113*	-0.0196*
		(0.00262)	(0.0008)	(0.0031)
Divorced parents	0.12	0.00619	-0.0035*	-0.0029
		(0.00349)	(0.0012)	(0.0045)
Mobility	0.11	-0.0439*	0.0143*	0.0264*
		(0.00344)	(0.0013)	(0.0040)
Mobility unknown	0.02	-0.00931	-0.0008	0.0280*
		(0.00879)	(0.0028)	(0.0108)
Born second quartile	0.27	0.00912*	-0.0007	-0.0060
		(0.00268)	(0.0007)	(0.0033)
Born third quartile	0.26	0.0228*	-0.0011	-0.0098*
		(0.00274)	(0.0007)	(0.0031)
Born fourth quartile	0.23	0.0297*	-0.0028*	-0.0098*
		(0.00289)	(0.0008)	(0.0033)
Observations	154,515	154,515	154,515	99,940
R-squared	-	0.299	0.070	0.073
Cohort x school FE (no of groups)	-	3,349	3,349	2,230

Note. Standard deviations in brackets. Standard errors in parentheses are clustered by compulsory school. * denotes significance at 1 percent level.