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ABSTRACT

The Impact of Teenage Motherhood on the Education and Fertility of their Children: Evidence for Europe*

This paper estimates the causal effect of being born to a teenage mother on children's outcomes, exploiting compulsory schooling changes as the source of exogenous variation. We impose external estimates of the direct effect of maternal education on child outcomes within a plausible exogeneity framework to isolate the transmission from teen motherhood per se. Our findings suggest that the child's probability of post compulsory education decreases when born to a teenage mother, and that the daughters of teenage mothers are significantly more likely to become teenage mothers themselves.

JEL Classification: I2, J13, J62

Keywords: teenage motherhood, education, fertility, children, instrumental variables, compulsory schooling laws

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1 Introduction

Over the last four decades, governments in many developed countries have increasingly targeted the reduction of teenage pregnancies and births. This reflects a widespread belief that teenage motherhood is linked to adverse socio-economic outcomes for the mother and the child (Kearney, 2010; Micklewright and Stewart, 1999, pp. 53).

The identification of the effect of teenage motherhood on outcomes is complicated by a number of factors. The most critical is that there are likely to be common unobservable factors that influence both the probability of giving birth as a teenager and other socio-economic outcomes. The sources of unobserved heterogeneity include omitted individual and family or background effects that lower the opportunity cost to early motherhood. This gives rise to an endogeneity problem leading to biased and inconsistent estimates of the causal effect of being born to a teenage mother on outcomes. One solution to this potential endogeneity problem is to adopt an instrumental variable approach. Here, we implement such an approach using comparable data across European countries.

There is large variation in the extent to which teenage motherhood is viewed as a problem across countries. This variation might be explained by a variety of factors such as differences in traditional attitudes to marriage, the extent to which teenage motherhood occurs within it,[1] and large country

[1]A study carried out by UNICEF (2001) suggests that traditional values still explain the low rates of teenage motherhood in Italy. Moreover, the fact that most of the cases

differences in teenage motherhood rates.[2] Such heterogeneity in views has led to a large range of policy responses which, in part, accounts for the varying degrees to which reductions in teenage birth rates have been achieved.

The existing literature on the causal effect of teenage motherhood on their own outcomes does not suggest important adverse effects on the mother. However, little is known about the causal effects of teenage motherhood on children's socio-economic outcomes. This is particularly noteworthy because any effect on children would be an important channel through which disadvantage is transmitted across generations. Simply applying OLS we find a modest but significant 2 percentage points difference in the school drop out rate for the children of teenmums relative to the 43% drop out rate for the 20-25 mothers. We also find a significant 4 percentage points difference in the teen motherhood rate of the daughters of teen mothers compared to a 1.4% rate for the daughters of mothers who gave birth age 20-25.

The contribution of this paper is to provide the first causal evidence on the effects of teenage motherhood across Europe on their children's outcomes.[3] In particular, our goal is to examine the causal effects of being born to a teenage mother on the probability of her child dropping out of school soon after compulsory education, and on the probability of her daughter becoming

occur within marriage in Greece and Portugal may contribute to the perception that it is not an important social problem in those countries.

[2]The lowest (5%) and highest (26%) teenage birth rates in Europe in 2006 are found in the Netherlands and the UK, respectively (EUROSTAT, 2009).

[3]The majority of the literature covers the US and UK experience, partly reflecting the relatively high teenage birth rates in these two countries. Only Francesconi (2008) has attempted to address this issue for UK data.

a teenage mother herself. To that end, we adopt an IV framework, exploiting compulsory schooling changes as the basis of exogenous variation in teenage motherhood.[4] A third order polynomial for mother's birth year controls for trends regarding access to contraception, aspirations towards education and family preferences affected by recessions, allowing RoSLA to act as a regression discontinuity. We view this choice of instrument as being particularly appropriate for identifying the causal effect of teen motherhood for those who are quite likely to become teenmums. That is, we feel that we identify an effect that is particularly relevant for policy. Of course, changes in maternal schooling affect the schooling of her children through intergenerational transmission. Thus our IV is contaminated - it has a direct effect on the outcomes through maternal education as well as an effect via teenage motherhood. Fortunately, Conley et al. (2012) provide a methodology to deal with exactly this kind of contamination. Thus, we relax the complete exogeneity condition that the IV framework would normally impose, to make inferences about the causal effect of teenage motherhood **per se** on children's outcomes.

Thus, our contribution is twofold: we provide estimates that exploit a novel methodology that allows us to use an instrument that might normally be regarded as legitimate; and we provide estimates that are relevant to a group of greatest relevance for policy.

Typically the literature has assumed that teen is defined as below 20 (or sometimes 19). We focus on identifying the effects of teen motherhood for

[4]Black et al. (2008) demonstrate such an effect in US and Norwegian data.

the more conservative definition of giving birth at or below 19 years old. For the sake of demonstrating the robustness of our results, we also provide estimates for a narrower definition of teenage motherhood (≤ 16 years old). We consider that a child is born to a teenage mother if the mother was a teenager when the given child was born, not if the mother was a teenager at the age of her first birth. We discuss this later.

We apply our methodology to data from the European Community Household Panel (ECHP) spanning the period 1994 to 2001.[5] The benefit of using cross-country data that is based on a harmonised and comparable dataset is that it allows us to exploit policy reforms across countries and it enables direct comparisons across countries.

Our preferred estimate from our empirical results suggests that the probability of not continuing school after compulsory education for children born to a teenage mother by age 19 is 3.4% higher than the 56% cross country average rate of dropping out at the minimum school leaving age for children born to a mother whose first birth was at the age of 20-25. When we drop the complete exogeneity assumption we find a significant and slightly larger effect of 3.7%. We also observe that the daughters of teenage mothers are 4.4% points more likely to give birth as teenagers themselves compared to a cross country average teenage motherhood rate of 1.4% for children whose mother's were aged 20-25 at first birth. However, when we relax the complete

[5] Berthoud and Robson (2003) also use ECHP data to present correlations for teenage motherhood and mother's and household outcomes, but not for children.

exogeneity condition of our IV estimator we find that this large fertility effect remains around 4.3% for the daughters of teen mothers. Note that these effects are larger than the OLS estimates which reflect the fact that they relate to the group of women who were likely to become teenmums - women who would have liked to drop out of school even earlier than currently allowed and so are affected by our minimum school leaving age instrument. It seems likely that our compliers are more vulnerable to shocks than average. While this is a selected group, this is the group of greatest relevance for policy.

The remainder of the paper is organized as follows. The next section provides a brief discussion of previous research in this area. In Section 3, the identification strategy is discussed and we show that the raising of school leaving age (henceforth RoSLA) leads to a reduction in the likelihood of giving birth as a teenager. Section 4 describes the data set employed. Section 5 presents the results of the econometric analysis where we show that offspring's education and fertility outcomes seem to be adversely affected by being born to a teenage mother. The final section concludes.

2 Previous literature on mother's and children's outcomes

The existing literature on the impact of teenage motherhood on second generation outcomes is primarily non-causal. The most common approach adopted is to compute correlation coefficients between teenage motherhood and children's outcomes. Typically, these estimated coefficients are large and negative, indicating poor performance of the offspring of teenage mothers. For instance, a large negative correlation of being born from a teenage mother and education outcomes of their children has been found in the US (Card, 1981), the UK (Pevalin, 2003) and New Zealand (Jaffee et al., 2001). Furthermore, there is evidence that young adults born to a teenage mother are more likely to become teenage parents themselves in the UK (Bonell et al., 2006; Botting et al., 1998; Ermisch and Pevalin, 2003; Kiernan, 1995; Manlove, 1997; Pevalin, 2003) and the US (Card, 1981; Hardy et al., 1998; Kahn and Anderson, 1992). The above results can be attributed, at least in part, to nonrandom selection into teenage motherhood caused by factors - such as prior disadvantage - that lower the opportunity cost of early childbearing (Wolfe et al., 2001).

The empirical literature dealing with the possible endogeneity of teenage motherhood on second generation outcomes is thin. An important exception is the study of Francesconi (2008), which suggests worse adult outcomes in the children born from a teenage mother in the UK with respect to education, labour market, inactivity, earnings, teenage childbearing, and health. However, his study employed an inevitably very small sample of sisters from the British Household Panel Survey (BHPS) dataset so as to provide siblings fixed effects estimates. For the US, Geronimus and Korenman (1993) also analyze children's health, by exploiting sibling fixed effects across sisters, and find an effect only for low-income black women in their twenties. However, Geronimus et al. (1994) find no effect of teenage motherhood on test scores of cousins where one mother was a teen mum and her sister was not.

In the case of studies examining mothers' outcomes, these have employed a variety of innovative methods - treating teenage motherhood as an evaluation problem - to control for unobserved characteristics influencing selection into teenage motherhood (see Ashcraft and Lang, 2010, for a review). Some of these studies suggest that the negative causal effects of giving birth as a teenager on mother's outcomes are insignificant or negligible.[6] That is, the poor performance of teenage mothers can be mainly attributed to prior disadvantage rather than early motherhood per se.

However, the methodologies employed in those studies may underestimate the true effects of teenage motherhood because of factors such as economies of scale when giving birth to twins versus a single teenage birth, misreporting and non-randomness in the case of miscarriages as an IV, and non-representativeness if the elderly sibling left the household when estimating sisters fixed effects (Hoffman, 1998). Moreover, there is imprecision in the resulting estimates because the implied treatment and control groups both

[6]For the US, Geronimus and Korenman (1992) exploit family fixed effects employing sisters that gave birth at different ages. Bronars and Grogger (1994) identify the effect from comparing giving birth to twins versus a single child as a teenager, and Brien et al. (2002) observe test scores before and after a woman has a child and use difference in difference estimates so as to examine how having a child as a teen affects the cognitive development of young women. An alternative approach is to use miscarriages as a mechanism for exogenously delaying age at first birth. For the US, higher labour market earnings and hours of work are found for teenage mothers (Hotz et al., 1997, 2005) as well as a small but negative effect on subsequent schooling (Ashcraft and Lang, 2010). For the UK, no statistically significant effect of teenage motherhood is found on household worklessness using RoSLA and mothers' month of birth as instruments (Walker and Zhu, 2009).

turn out to be very small in the datasets being used.

This leads to a second, but smaller, body of research that has been more successful in finding effects of teenage births on the mother's own outcomes. Using propensity score matching techniques Levine and Painter (2003) in the US and Goodman et al. (2004) in the UK, find a reduction in education achievement. Fletcher and Wolfe (2009) use miscarriages as an IV to deal with contamination by controlling for community fixed effects in the US, and find reductions in the probability of receiving a high school diploma, annual income and years of schooling. Within the same IV framework, Goodman et al. (2004) find that teenage mothers in the UK are less likely to be in work, work fewer hours, earn a lower hourly wage and tend to have partners with lower education qualifications and labour market status.[7] Qualitatively similar results are found in Chevalier and Viitanen (2003) where the instrumental variable is age at menarche.

The literature of teenage motherhood on own outcomes finds negative effects but a consensus still has not been reached on whether these negative effects are large enough to be important. There are different mechanisms through which teenage motherhood can affect child's outcomes. If teenage motherhood has a negative causal effect on the mother's subsequent schooling, this may relegate her to a lower lifetime wage and earnings trajectory than she would otherwise have followed. It can likewise influence other sources of household earnings, either because of assortative mating patterns,

[7]Similar findings are in Pevalin and Ermisch (2005).

or by influencing the probability and stability of partnering. Another mechanism through which teenage motherhood can affect child outcomes is maturity. The developmental psychology literature has shown that teenagers have different brain chemistry leading them to reckless and impulsive behaviour. Teenage mothers are more likely to adopt risky behaviour: being less likely to breastfeed, and more likely to drink or smoke during pregnancy (Botting et al., 1998). Therefore, this can have consequences on their children's outcomes and subsequent behaviour.

3 Methodology

It is well known that IV methods provide estimates which do not necessarily correspond to the average effect of the treatment on the treated. Consequently, it may well be the case that different studies are estimating different things. In this paper, we exploit changes in schooling laws as an instrument for teenage motherhood so as to estimate the causal effects of teenage childbearing on children education and fertility outcomes. It seems likely that our estimates will reflect the effects of teen motherhood on the children of mothers who were likely to have been teen mothers: women who would have left school early. Our approach is motivated by the findings of Fort (2007), which indicate an effect of compulsory schooling laws on fertility behaviour in Italy, and Black et al. (2008) who illustrate that raising the minimum school leaving age reduces the probability of teenage pregnancy in the US and Norway.[8] This works through both an incarceration and a human capital effect. The former reduces the time out of school available to have a child, while the latter increases the current and expected future human capital which has a corresponding impact on delaying fertility through raising the associated opportunity cost. This provides both an empirical and a theoretical justification for using RoSLA as an instrumental variable for teenage motherhood. The usual identifying assumption is that these affect the outcomes for the child only through their effect on teenage motherhood. Of course, it is likely that there are indirect effects, for example because teen motherhood affects the mother's education. Thus, such estimates capture both the indirect effect through maternal education as well as the teen motherhood effect per se.

It is usual to presume that the chosen instrumental variable, compulsory schooling laws, should be relevant and yet validly excluded. The relevance condition requires that there is correlation between the instrument Z and the likelihood of being a teenage mother. As far as the validity condition is concerned, it is usual to presume that changes in compulsory schooling laws that affect mothers then affect the outcomes of children only through teen motherhood and not through any other transmission mechanism.

RoSLA has been proved a successful instrument for schooling in the returns to education literature across a number of countries within our analy-

[8] Monstad et al. (2011) employ a Norwegian educational reform that impacted on the elder sister's teenage fertility to analyse the teenage childbearing of their younger sister. They find peer effects in teenage motherhood, such effects are larger where siblings are close in age and for women from low resource households.

sis, including the UK (Harmon and Walker, 1995; Oreopoulos, 2007; Grenet, 2010), Germany (Pischke and von Wachter, 2008), France (Grenet, 2010), Italy (Brunello and Miniaci, 1999), Ireland (Denny and Harmon, 2000), the Netherlands (Levin and Plug, 1999), Portugal (Vieira, 1999), and Spain (Pons and Gonzalo, 2002). Changes in compulsory schooling laws are also used to generate exogenous variation in schooling when analysing the effect of education on other individual outcomes.[9]

An appealing feature of our cross-country setting is that we can exploit the wide natural variation in the data caused by exogenous changes in compulsory schooling laws in 13 European Union countries since 1959. Table 1 lists the reforms that have taken place in each of the countries studied.[10]

(Insert Table 1)

In Section 5, we present OLS and 2SLS estimates of the effects of teenage childbearing on children outcomes based on the following model:

$$\mathrm{Y}_i = \alpha_i + \beta \mathrm{T}_i + \gamma_i \mathrm{Z}_i + \delta \mathrm{X}_i + \varepsilon_i, \tag{1}$$

$$\mathrm{T}_i = \mu_i + \theta \mathrm{RoSLA} + \phi(\mathrm{BIRTHYEAR})_i + \lambda \mathrm{X}_i + \upsilon_i \tag{2}$$

[9] They include lifetime wealth, health, unemployment and overall happiness in the US, Canada and the UK (Oreopoulos, 2007), as well as political interest and involvement in Britain and the US (Milligan et al., 2003). Moreover, these laws have been found to lower the likelihood of committing crime or ending up in jail (Lochner and Moretti, 2004), to increase life expectancy (Lleras-Muney, 2005), productivity (Chevalier et al., 2004) and to have an effect on schooling and its subsequent effect on teenage childbearing (Silles, 2011).

[10] Fort (2006) provides a survey of these changes as part of her review of the effects of individual's qualification levels on earnings.

Note that $\gamma_i = 0$ is the usual IV assumption.

The first equation examines the effect of being born to a teenage mother, T_i, on the children's adult outcomes, Y_i, where T is defined by either 16 or 19, and Y_i is either the probability of the child dropping out of school after compulsory education, or the probability (for the sample of daughters) of giving birth as teenagers. Z is the matrix of excluded instruments where $\gamma_i = \gamma + u_i$ is the parameter that measures the plausibility of the exclusion restriction with $\gamma_i = 0$ being the usual IV condition. Equation (2) relates the probability of having a first birth as a teenager, to RoSLA which defines the changes in the compulsory level of education for each country (see Table 1). The effect of (BIRTHYEAR) on T_i is assumed to be captured by a third order polynomial on **mother's** year of birth which controls for smooth social trends so as to allow RoSLA to act as a regression discontinuity. X is a matrix of exogenous regressors that may include a third order polynomial on **child's** year of birth, maternal education, year and country dummies (when pooling data across countries).

We are concerned that one of our instruments, changes in compulsory schooling laws, affects the children's outcomes not only via its effect on teenage motherhood and thus violates the exclusion restriction, $E[Z'\varepsilon_i] \neq 0$. Thus, we also attempt to exploit the recent development of "plausible exogeneity" in Conley et al. (2012)[11] to make inferences about the effect of early

[11]ltz STATA code to implement the method is available on Christian Hansen's website http://faculty.chicagogsb.edu/christian.hansen/research/.

motherhood without the assumption that the exclusion restriction, $\gamma_i = 0$, holds exactly (henceforth PE). In our context $\gamma_i \mathrm{Z}_i$ is $\eta \times \rho$RoSLA + u_i. $\gamma = \eta \times \rho$ where ρ indicates the effect of Raising the Minimum School Leaving Age on the probability of mother's postcompulsory schooling, and η is either the intergenerational transmission of education parameter when Y_i is the probability of children dropping out of school after compulsory education, or η is the effect of the mother's having at least post compulsory schooling on their daughters probability of giving birth as teenagers.

The children of teenage mothers can have poorer outcomes not because their mothers gave birth as teenagers but because they have not been exposed to the same education opportunities. For instance, mother's affected by RoSLA are likely to be in the labour market and thus have higher earnings. Therefore, their children can drop out of school, and their daughters can be more likely to become teenage mothers, not because they are born to a teen mother but to a mother that is doing badly in education or the labour market. If this is the case then our instrument will not be exogenous to children outcomes. But we can still provide inference about β if we allow for the direct effect of RoSLA on children's outcomes.

We are interested in checking that other variables, than the probability of treatment and mother's educational attainment, do not change discontinuously at the effective dates of the new laws. The cubic trend in mother's birth year will be picking up whether such laws were passed in response to an overall trend in economic opportunities, or structural changes in the economy, or concerns about such things as a high rate of juvenile delinquency. But if the laws were part of a larger policy aimed at improving young people's prospects, it is more difficult to know all the ways in which the children of interest might be affected. For example, imagine that compulsory schooling was passed simultaneously with programs that provided better food and health care for disadvantaged young people. Then, having a healthier mother might affect a child in different ways than having a more educated mother- and both would affect the child directly as well as through the effect of not having a teenage mother. The survey responses for the mother's generation correspond to the period when their children are already adults, and thus, we do not have information of the mother's generation at the moment they were giving birth or when those laws were implemented. We can expect that if there was a specific policy that improved health status at the same time as compulsory schooling laws were implemented, mother's affected by a RoSLA in their schooling years will be healthier than those not affected by a schooling reform. We observe some variables which indicate the health of the individual in general, whether the individual has any chronic physical or mental health problem, illness or disability, and their body mass index. Although these variables are observed once their children are young adults, we would expect to see the effects of a health policy not immediately but later on time. In Table A.1 in the Appendix we observe that in most of the countries there is not a statistically significant difference between the mean health status, mean health problem and body mass index for pre and post

RoSLA mothers. If we look at the weighted mean tests for all the sample countries, we observe that we can not reject the null hypothesis where mean differences are equal to 0 for the health status and the BMI variables, with p-values of 0.7030 and 0.0397, respectively.

Finally, the structure of our data can also be used to reduce the contaminating effect of unobserved family influences because we can include siblings. The ECHP is a longitudinal survey that provides details about the fertility of family members and the relationships between them that enables us to match parents with their children and one sibling with another. First differences or fixed effects estimation in a sample of siblings then allows us to control for time-invariant unobserved family characteristics that could be affecting children's outcomes directly and not just through being born to a teenage mother. The first drawback of this method is that it does not deal with individual unobserved characteristics that are potentially correlated with children's outcomes. Moreover, Holmlund (2005) study of teenage motherhood in Sweden, shows that with detailed individual-specific information, the siblings approach is no more informative than a traditional cross-section. A compelling practical problem is that sample sizes of matched siblings are inevitably small.

4 Data

Our empirical analysis employs data from the European Community Household Panel – a longitudinal survey conducted by Eurostat from 1994 to 2001 for the countries of the EU 15.[12] The appealing feature of this data set is that it enables us to identify mothers and their biological children living in the same household and, thus, allows us to observe the correlates of being a teenage mother and how this relates to their child's outcomes. Moreover, it allows us to make cross-country comparisons regarding the negative effects of being born to a teenage mother on children.

For our analysis we select mothers and their biological children, aged from 16 to 18 years old,[13] keeping multiple observations of siblings in the same household.

One problem, for us and most of the studies, is that our data does not have a complete fertility history for the mothers. We infer teen motherhood from comparing child age with maternal age. Thus, we fail to capture teen mums whose first child is no longer in the household. The fact that we do not observe children who were born when the mother was a teenager but have left the household implies that the mother's age at first birth may be over-estimated. And thus, the estimated effect obtained is a lower-bound

[12] Austria and Finland joined the survey in 1995 and 1996, respectively.

[13] Selecting children up to 18 years old ensures that the pattern of children cohabiting with their parents at that age does not differ across European countries (Manacorda and Moretti, 2006; Chiuri and Del Boca, 2010).

estimate of the effect of being born to a teenage mother.[14]

A second problem, for us and many other studies, is that children who are particularly adversely affected by being born to a very young mother may be more likely to leave the household at an early date and so be attriters. This would imply that our estimates are a lower bound of the true effect, because we omit the most adversely affected children. The probability of children dropping out of the sample conditional to being in it for the previous year increases at 2.4% per annum with the age of the child. Moreover, the probability that mothers-children pairs leave the sample is 8.54% if the mother gave birth to that child below 17 years old and 2.3% if teenage motherhood occurred below 19 years old. It seems likely that the attriters will exhibit worse outcomes because many of them will have left the parental home. For example, a young mother who is ejected from the family home is likely to have worse unobservables and find providing for a young child more difficult than would be the case for those who remain in the family home where grandparental care and advice might provide excellent support. Thus, this is an additional reason why our estimates might well be regarded as an underestimate of the true effect in the population, even for those most at risk.

A key issue in this framework is to specify an appropriate control group to act as a counterfactual. We choose young adults who were born to a mother who was 20-25 years old, rather than the more extensive definition of

[14]We test that our results are robust to using only oldest sibling.

all non teens that is usually employed. Furthermore, we use both a narrow and a broad definition of "teen mother" for which the same control group is employed. The narrow definition corresponds to any female who give birth at 16 years old or earlier, while the broad definition is the usual one which corresponds to giving birth at 19 years old or earlier. By varying the definition of teenage motherhood we observe whether childbearing in the early adolescent years affects child outcomes more than in the later teens.[15]

In Figure 1 we observe teenage motherhood rate trends since 1950s for the European countries that we are analyzing. We can see that teenage birth rates have started to decrease in all these countries since the 1970s reflecting the shift in aspirations of young women towards education and improved access to contraception leading to the delay in women's childbearing decisions. At the same time there has been a decline in all births partly explained by women's participation in education and the labour market. In the UK although the conception rate for women under 18 has started to decrease from 1998 (two decades later) the gap that emerged relative to the rest of the countries has not closed yet (ONS, 2010). Table 2 reports the number of observations for the estimation sample for the pool of countries and, for each country separately. In total there are 7891 mothers who had a first birth by the age of 25. 1.10% of these gave birth at the age of 16 or below, 3.24% at 17, 7.01% when they were 18 years old and 13.03% of them gave birth before

[15]A tighter definition implies that the treatment group becomes smaller so that the precision of the estimates may fall.

their twentieth birthday.

(Insert Figure 1 and Table 2)

The dependent variables are: whether the child completed various education levels defined by International Standard Classification of Education (ISCED); and daughter's teenage motherhood. The latter can be deduced from the dataset by combining information regarding all births within the household during the period and the identity of the recipient of the associated child and maternity benefits. In Figure 1, we observe that the probability of continuing school after compulsory education is lower for children born from a teenage mother than for those born to a mother that was 20-25 when she gave birth. This is also true in most of the countries separately, except in France, Germany, Ireland and Italy. Children's education outcomes are very different for children born to teenage mothers versus children born to mothers that gave birth at 20-25 in Portugal, Spain and the UK, whereas these differences are smaller in Austria, Belgium, Greece and the Netherlands (see Table 3). We can also see that the probability of giving birth in their teens is higher for daughters born to a teenage mother for the sample of the pool of countries in Figure 1 and in each of the countries in Table 3. Differences between daughter's teenage motherhood when born to a teenager or to a mother that was 20-25 when she gave her birth are higher in Germany, Ireland, Italy and Spain.

(Insert Figure 2 and Table 3)

We examine the stability of estimates to the inclusion of control variables for the mother and the child. They consist of education of the mother and a third order polynomial for mother's year of birth. The remaining explanatory variables include a third order polynomial for child's year of birth, year dummies and country fixed effects. In Table 4 we observe that the proportion of teenage mothers who have only primary education is higher in all countries than the percentage of mothers with primary education attainment that gave birth at the age of 20-25. The percentages of mothers that reached secondary or higher education levels in each of the countries is higher for mothers that gave birth in their early twenties than for teenage mothers.

(Insert Table 4)

We now turn to the description of the raising of school leaving age. The RoSLA dummy takes the value of unity if the mother belongs to a birth cohort that was subject to extended compulsory schooling in that particular country and zero otherwise. This dummy is expected to pick up the effects of the reform indicated by column 2 of Table 1. While, column 1 of the same table indicates the year of birth of the first cohort affected by the reform in each of the countries. However, we might expect RoSLA to have an effect that is delayed until after the reform because of the nine month gestation period. Thus, the variable that is likely to pick up the effect on reducing teenage childbearing is the lead, RoSLA_{t+1}.

5 Results

5.1 OLS and 2SLS estimates of the effects of teenage motherhood on children outcomes

Figure 3 shows the probability of becoming a teenage mother at age 19 by distance of the cohort from year of reform implementation. We observe a very clear discontinuity in teenage motherhood at the first cohort that is affected by the RoSLA reform. This is confirmed latter on by the strong robust F-tests of the instrument.

(Insert Figure 3)

The results from the first stage, from Linear Probability Model (LPM) estimates of teenage motherhood on the IV (RoSLA_{t+1}), are reported in Table 5. A conclusion that emerges is that raising the minimum school leaving age reduces the probability of giving birth as a teenager a year later. In particular, compulsory schooling laws decrease the probability of giving birth at or below 19 years old by 2% compared to a cross country average teenage motherhood rate of 13.03%. The corresponding figure for giving birth at 16 years old or earlier is 1.5% compared to a teenage motherhood rate of 1.10%. This is true whether we control for maternal education or not in the first and second stage.

We would like to make sure that timing of compulsory schooling laws is exogenous. For this purpose, in Table A.2 in the Appendix, we investigate the effect of a placebo reform for the year prior and two years prior to the

RoSLA on the probability of teenage motherhood. Estimates from these models revealed no effect of the lagged reform on the likelihood of becoming a teenage mother.

(Insert Table 5)

Table 6 reports our regression estimates of the effect of teenage child-bearing for each of the 13 countries in our sample using a simple Linear Probability Model (LPM). It shows estimated coefficients between adult children education outcomes and teenage motherhood by the age of 16 and 19 years old. Table 7 shows the probability of daughter's giving birth in their teens. Both Tables correspond to a specification, which includes a third order polynomial for child's year of birth so as to take into account cohort effects. The rows labeled OLS show estimates under the assumption that teenage motherhood is exogenous. Small sample sizes make the results using a teen definition of 16 very imprecise. Overall, our pooled OLS estimates of the percentage continuing school after compulsory education decreases by 2.3% (SE=3.8), and 2.3% (SE=1.1) when born to a teenage mother at age 16 and 19 respectively, compared to being born to mothers between 20 and 25 years old. Large significant negative education effects are found for Spain and the UK whereas they are very small in Denmark and Finland, although we cannot say why these teens in these countries are less affected.

The rows labeled 2SLS present results from an instrumental variable procedure, which instruments teenage motherhood by RoSLA_{t+1}.[16] In this case,

[16]We also employ as exclusion restrictions a third order polynomial for mother's year of

we take into account the bias due to omitted variables and yet find a larger negative effect of teenage motherhood on adult children education outcomes: 3.4% (SE=1.4) for children born to mother's that gave them birth before 20. Looking at the results by country, it appears that the effect is significant and smaller in France 5.2% (SE=2.5), than in Greece 9.0% (SE=4.5), Spain 13.6% (SE=4.0) and the UK 11.5% (SE=6.6). Although one might have expected IV to be lower than OLS coefficients, our results are consistent with a LATE interpretation that is usually associated with using an IV such as RoSLA.

The fact that we find higher IV than OLS is consistent with a Local Average Treatment Effect (LATE) as the RoSLA instrument does not affect the whole population but affects the behaviour of just a small percentage of those exposed to the instrument (see Oreopoulos, 2006). That is, the average effect that we are capturing is that for those children born to a teenage mother for young adults whose mother has delayed childbearing as a result of the reform. An alternative explanation is that RoSLA and teenage motherhood may be weakly correlated. In this case, IV estimates may be vulnerable to a weak instrument problem and may be even more biased than OLS (Bound et al., 1995). However, inspection of the first stages suggests that this is not the case.

PE estimates were obtained by assuming that γ was 0.0068 in the case of education outcomes and -0.00068 in the case of fertility outcomes. These

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birth to take into account mother's cohort effects.

coefficients come from $\gamma = \rho \times \eta$ where $\rho = 0.068$ is the effect of RoSLA on mother's probability of reaching post-compulsory schooling levels.[17] That is, changes in compulsory schooling laws increase the probability of maternal post-compulsory education by 6.8%. Whereas, $\eta = 0.10$ corresponds to the intergenerational transmission of education parameter: that is, the probability of children post-compulsory schooling increases by 10% if the mother has reached post-compulsory education levels herself. Thus, the direct effect of changes in compulsory schooling laws affecting the mother's generation is 0.68% on the children's probability of continuing in education. The direct effect of changes in compulsory schooling laws affecting the generation of the mother on daughter's probability of becoming a teenage mother is $\gamma = \rho \times \eta = -0.00068$ where $\eta = -0.01$ is the effect of mother's post-compulsory schooling on daughter's probability of becoming a teenage mother. While the education outcome estimate of a young teen mother remains insignificant under PE at 16 years old, the broader teenmum effect is now somewhat larger. Across countries, we observe that the magnitude of the PE estimated effect is higher in Denmark, Spain and the UK compared to the IV case where $\gamma = 0$; whereas it remains of the same size in France and Greece.

Our findings provide evidence that, while there may be a role for previous disadvantage on the impact of teenage motherhood, our estimates of causal effects are large and significant. One should note that the adverse evidence

[17] The resulting ρ and η are our own estimates using ECHP data.

obtained is conditional on not attriting. Thus, if the children that do worse are born to teenage mothers and those pairs are more likely to attrit, we will still be underestimating the adverse effects of teenage motherhood.[18]

(Insert Tables 6 and 7)

Turning back to Table 7, we observe the probability that daughters of teenage mothers become teenage mothers themselves from the age of the mother at her first birth. Starting with the OLS rows, we note that the effect of being born to a teenage mother on daughters giving birth at 16 or below is 5.8% (SE=1.9). That is, the effect is worse for a tighter definition of being a teenage mother. However, it decreases to 4.0% (SE=0.7) for the usual broader definition of teenage motherhood. The 2SLS estimates of teenage motherhood on second generation fertility behavior exhibit a similar pattern with the OLS case. In particular, the causal effects decrease from 7.5% (SE=3.0), to 4.4% (SE=0.9). Overall, the daughters of teenage mothers are significantly more likely to also give birth as teenagers.[19] These results are very close to those of sister fixed effect estimates that Francesconi (2008)

[18] Following the literature on mothers' outcomes we provide estimates of sisters fixed effects as we have available different siblings within the same household during the same year (Geronimus et al., 1994). Comparing the outcomes of sisters, where the mother was ≤ 19 at the birth of one, and > 19 at the birth of the other, we aim to mitigate the endogeneity problem that is produced by unobserved family characteristics that are common within siblings or sisters. The results in Table A.3 in the Appendix suggest that daughters of teenage mothers are 8.8% more likely to become teenage mothers themselves compared to our 4.4% IV estimate. Whereas the effect on children's post-compulsory education is not significantly different from zero.

[19] Hardy et al. (1998) find that the effect is weaker for sons than for daughters. Our data doesn't allow us to observe the incidence of teenage fatherhood among sons, as we rely on who is getting the maternity benefits and the likelihood of those babies being co-resident.

finds for the UK (2.7% likelihood of becoming a teenage mother) but not to the correlations reported in Ermisch and Pevalin (2003) and Pevalin (2003) which suggest that daughters of teenage mothers are 2 to 4 times more likely to give birth as teens.

We now move to the comparison of the results, displayed in Table 7, across countries. Initially, we focus on daughters' of mothers that gave birth before 20. In Italy 5.6% (SE=2.3) and Spain 4.8% (SE=1.8), the behaviour of daughters born to a teenage mother seems to be close to the average of the pool of countries. Whereas, in Belgium 25.4% (SE=7.3), Ireland 12.3% (SE=5.8) and the UK 11.9% (SE=5.9), second generation girls seem more likely to imitate their mother's past fertility behaviour by 19 years old. This is also true for Italy 16.7% (SE=6.4) and Portugal 12.4% (SE=6.6) when we use the mother's tighter definition of teenage motherhood. Irish and British daughters behave in a similar way if born to mothers that give birth by 19 years old, they do not give birth by 19 years old if the mother gave birth to them earlier. The failure to detect statistically significant coefficients in some countries such as Denmark and the Netherlands is likely to be due to the small sample sizes in these countries as teenage birth rates are very low; or to national policies that are effective in counteracting the effect of being born to a teen mother.

In order to shed light on the stability of our coefficients, in Tables 8 and 9, we conduct a sensitivity analysis based on another specification that takes into account child's cohort effects. The first specification has no covariates apart from teenage motherhood, year dummies and country dummies. Specification 2 includes also a third order polynomial for child's year of birth. The fact that our PE coefficients are hardly affected after introducing maternal education suggests that the effect of teenage motherhood on children outcomes does not arise through a pathway via mother's education. Thus, our study supports the economics literature mentioned in Section 2 that finds that teen motherhood has little influence on mother's education. Including a larger number of controls does not appear to have an impact on the coefficient and its statistical significance. Overall, we find that there is an adverse causal effect of being born to a teenage mother on children education and fertility outcomes which is robust to a number of different specifications. Moreover, we find that teen motherhood has an adverse effect which is bigger for daughters fertility, the earlier it occurs.

OLS, 2SLS and PE estimates, as well as, the number of observations in each of the samples, the Partial R^2, the F-test for exclusion restrictions and the corresponding p-values, and the Sargan test of over-identification for all the specifications for the pool of countries are also shown in Tables 8 - 9. Our findings for Europe as a whole are qualitatively similar with those of the ISER research for the UK (see, e.g., Ermisch and Pevalin, 2003; Pevalin, 2003; Francesconi, 2008): children of teenage mothers tend to have lower educational attainment and exhibit a higher risk of becoming teenage parents themselves. The former result holds also for the remaining specifications.[20]

[20]We also introduce abortion rates for the different years and countries from the Histor-

We should note the rows labelled PE in Tables 8 and 9. They show that the point estimates of the effect of teenage motherhood allowing for violation of the complete exogeneity condition in an IV framework are very stable across specifications.

(Insert Tables 8 and 9)

Finally, in order to facilitate comparison of our findings with the empirical literature on teenage motherhood on different dependent variables and datasets, that come from other IV and alternative methods. We standardize the results by dividing the coefficient obtained in the different studies by the standard deviation of the corresponding dependent variable.

Starting with child outcomes: the effect of being born to a teenage mother ranges from 350% to 878.13% of an SD increase in age at first birth of second generation in Manlove (1997). In contrast, it increases age at birth by 6.13% of an SD in Card (1981). In our case, the probability of daughters giving birth in their teens if born to a teenage mother at 19 years old increases by 31.26% of an SD (59.60% using the 16 definition of teenage motherhood).

The existing literature on mother's outcomes also features a wide range of standarized coefficient values even within the same study.[21]

ical Abortion Statistics in http://www.johnstonsarchive.net/policy/abortion/index.html so as to predict the likelihood of teenage motherhood. We find that they have a positive instead of a negative effect (as has been shown for the US) on teenage motherhood. Our results are slightly smaller than without abortion rates although do not change significantly.

[21]For instance, Bronars and Grogger (1994) find a decrease from 1.29% to 916.36% of an SD in mother's and family earnings respectively and a reduction of 1590.60% of an SD in years of education. Ribar (1999) adopts a DD framework and shows a reduction between 3.69% and 419.04% of an SD. The above numbers correspond to log family income

6 Conclusions

This is the first paper to provide evidence of a statistically significant causal effect of teenage motherhood on second generation education and fertility outcomes across Europe. First, we illustrate that the likelihood of giving birth as a teenager is reduced with the raising of the minimum school leaving age. Second, by employing OLS and 2SLS techniques we find firstly, that the probability of continuing school after compulsory education decreases when born to a teenage mother even when we account for the potential omitted variable bias. Secondly, we find that the daughters of teenage mothers are significantly more likely to give birth as teenagers themselves, and the magnitude of this effect is even larger, when we take into account unobserved factors. Finally, we can still make inferences about the causal effect of teenage

accounting for family fixed effects and to completed years of education when using sisters IV, respectively. These studies, although classified in the literature within those that fail to find a causal effect of teenage motherhood, seem to find an effect which is economically significant when looking at standardized results. In Goodman et al. (2004), where miscarriages are used as an IV, we observe a reduction of 21.23% for log equivalised family income and 9.93% for the age they left full-time education. The former is closer to our OLS results for the probability of second generation continuing school after compulsory education (10.95% and 6.42% if born to a mother that was 18 and 19 years old when she gave them birth). Furthermore, in Fletcher and Wolfe (2009) a lower bound decrease of 6.47% and 51.39% with and without fixed effects of an SD of the years of education is obtained as a result of having given birth as a teenager. Our IV estimates move in a range that goes from 41.69% to 18.46% of an SD if born to a mother that was 16 and 19 years old in our first specification, although this last number is reduced to a 6.69% of an SD if we consider our fourth specification. An interesting study is that of Brien et al. (2002) for which we estimate a reduction of the education scores by 2.54% to 22.19% of an SD for the sample of non-blacks, and an improvement across some of the different outcomes for the sample of blacks. Finally, for the study of Chevalier and Viitanen (2003) the largest impact corresponds to a 73.97% reduction of an SD in labour participation whereas in Goodman et al. (2004) that was calculated to be a 34.6%.

motherhood on children outcomes when we relax the complete exogeneity assumption in the IV framework. We find that for the pool of countries, the causal effects remain the same when we allow for RoSLA to directly affect children outcomes through maternal education.

We find a long-term effect of teenage motherhood in some specific countries. There are clearly differences in the incidence of teen motherhood across countries. The existing literature suggests how this might be addressed. Here we find differences in the effect of teen motherhood. We do not know why these effects differ but it seems likely that some countries have welfare programmes that make living standards more robust to teen motherhood.

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Table 1: Schooling Reforms in European Countries

Country	First cohort	Year Reform	Reform	References
Austria	1947	1962	14 → 15	
Belgium	1969	1983	14 → 18	
Denmark	1957	1971	14 → 16	Bingley et al, 2005
Finland	1961	1972-1977	13 → 16	
France	1953	1959	14 → 16 (14 in 1967)	Grenet, 2010
Germany	1953	1949-1967	14 → 15 (14 in 1967)	Pischke and Von Watcher, 2008
Greece	1952	1964	12 → 15	
Ireland	1958	1972	14 → 15	Denny and Harmon, 2000
Italy	1949	1963	11 → 13 (14 in 1963)	
Netherlands	1959	1975	15 → 16	
Portugal	1952	1964	12 → 14	
Spain	1958	1970	12 → 14	Pons and Gonzalo, 2002
UK	1958	1973	15 → 16	Harmon and Walker, 1995

Note: Information in this table is taken from (Fort, 2006) and Eurybase (2009): the Eurydice database on education systems in Europe. First column indicates the first cohort that experienced the RoSLA, Year Reform corresponds to the year the reform was implemented and Reform includes the exact increase in compulsory schooling years.

Table 2: Number and proportion of mothers in the corresponding age groups

Country	20-25	TeenMum16			TeenMum17			TeenMum18			TeenMum19		
	N	N	Total	%	N	Total	%	N	Total	%	N	Total	%
Austria	416	9	425	2.12	17	433	3.93	36	452	7.96	66	482	13.69
Belgium	214	1	215	0.47	3	217	1.38	7	221	3.17	17	231	7.36
Denmark	172	1	173	0.58	2	174	1.15	9	181	4.97	20	192	10.42
Finland	363	1	364	0.27	8	371	2.16	16	379	4.22	36	399	9.02
France	477	1	478	0.21	2	479	0.42	18	495	3.64	45	522	8.62
Germany	826	6	832	0.72	26	852	3.05	57	883	6.46	137	963	14.23
Greece	666	20	686	2.92	43	709	6.06	89	755	11.79	170	836	20.33
Ireland	337	2	339	0.59	6	343	1.75	17	354	4.80	39	376	10.37
Italy	825	5	830	0.60	25	850	2.94	54	879	6.14	99	924	10.71
Netherlands	374	2	376	0.53	6	380	1.58	14	388	3.61	25	399	6.27
Portugal	701	18	719	2.50	43	744	5.78	85	786	10.81	158	859	18.39
Spain	880	8	888	0.90	33	913	3.61	72	952	7.56	122	1,002	12.18
UK	538	2	540	0.37	15	553	2.71	39	577	6.76	86	624	13.78
Total	6,863	76	6,939	1.10	230	7,093	3.24	517	7,380	7.01	1,028	7,891	13.03

Note: Teen Mum 16, 17, 18 or 19 stands for the age of the mother at the child's birth.

Table 3: Proportion of children continuing school after compulsory education and daughter's giving birth in their teens by country

	Children's Education		Daughter's Teenage Motherhood	
Country	TeenMum 19	Mother 20-25	TeenMum 19	Mother 20-25
Austria	7.576	9.135	1.515	0.962
Belgium	11.765	13.084	11.765	0.000
Denmark	0.000	3.488	0.000	0.000
Finland	0.000	1.377	2.778	1.928
France	62.222	60.377	0.000	0.210
Germany	4.380	2.542	2.920	0.484
Greece	22.353	23.273	0.000	0.000
Ireland	23.077	20.475	7.692	1.187
Italy	10.101	7.879	2.020	0.364
Netherlands	8.000	8.289	0.000	0.000
Portugal	5.063	6.419	7.595	2.568
Spain	17.213	29.205	2.459	0.227
UK	18.605	30.669	5.814	1.859

Table 4: Descriptive statistics of explanatory variables

	PrimEducMother		SeconEducMother		HighEducMother	
	TeenMum	Mother 20-25	TeenMum	Mother 20-25	TeenMum	Mother 20-25
Austria	0.500	0.327	0.500	0.651	0.000	0.022
	(0.504)	(0.470)	(0.504)	(0.477)	(0.000)	(0.146)
Belgium	0.647	0.430	0.353	0.355	0.000	0.215
	(0.493)	(0.496)	(0.493)	(0.480)	(0.000)	(0.412)
Denmark	0.600	0.291	0.300	0.407	0.100	0.302
	(0.503)	(0.455)	(0.470)	(0.493)	(0.308)	(0.461)
Finland	0.333	0.207	0.528	0.471	0.139	0.322
	(0.478)	(0.405)	(0.506)	(0.500)	(0.351)	(0.468)
France	0.778	0.639	0.133	0.237	0.089	0.124
	(0.420)	(0.481)	(0.344)	(0.426)	(0.288)	(0.330)
Germany	0.358	0.270	0.577	0.534	0.066	0.196
	(0.481)	(0.444)	(0.496)	(0.499)	(0.249)	(0.397)
Greece	0.853	0.649	0.135	0.263	0.012	0.089
	(0.355)	(0.478)	(0.343)	(0.440)	(0.108)	(0.284)
Ireland	0.846	0.659	0.128	0.288	0.026	0.053
	(0.366)	(0.475)	(0.339)	(0.453)	(0.160)	(0.225)
Italy	0.818	0.655	0.172	0.319	0.010	0.027
	(0.388)	(0.476)	(0.379)	(0.466)	(0.101)	(0.161)
Netherlands	0.720	0.663	0.240	0.302	0.040	0.035
	(0.458)	(0.473)	(0.436)	(0.460)	(0.200)	(0.183)
Portugal	0.943	0.902	0.051	0.066	0.006	0.033
	(0.233)	(0.298)	(0.220)	(0.248)	(0.080)	(0.178)
Spain	0.828	0.776	0.148	0.131	0.025	0.093
	(0.379)	(0.417)	(0.356)	(0.337)	(0.156)	(0.291)
UK	0.779	0.602	0.093	0.154	0.128	0.243
	(0.417)	(0.490)	(0.292)	(0.362)	(0.336)	(0.430)

Note: Means (standard deviations in parentheses). TeenMum corresponds to giving birth below 20 years old. PrimEducMother, SeconEducMother and HighEducMother correspond to the mother's reaching Primary, Secondary or Higher Education levels.

Table 5: First stage: Effect of RoSLA_{t+1} on the probability of first generation teenage motherhood (LPM)

	TeenMum 16	TeenMum 19
RoSLA_{t+1}	-0.015^{c}	-0.020^{b}
	(0.004)	(0.010)
Birthyear	0.830^{c}	-0.179^{b}
	(0.073)	(0.084)
Birthyear^2	-0.036^{c}	0.006^{a}
	(0.003)	(0.003)
Birthyear^3	0.001^{c}	0.000
	(0.000)	(0.000)
Austria	-0.015^{c}	-0.093^{c}
	(0.005)	(0.013)
Belgium	-0.005	0.003
	(0.007)	(0.018)
Germany	-0.007^{b}	-0.007
	(0.003)	(0.010)
Denmark	-0.008	-0.011
	(0.007)	(0.018)
Spain	-0.021^{c}	-0.053^{c}
	(0.005)	(0.012)
Finland	-0.012^{b}	0.012
	(0.006)	(0.016)
France	0.002	0.019
	(0.003)	(0.012)
Greece	0.009^{a}	0.015
	(0.005)	(0.011)
Netherlands	-0.007	-0.017
	(0.006)	(0.015)
Ireland	-0.014^{b}	-0.025^{a}
	(0.006)	(0.015)
Portugal	0.008^{a}	0.012
	(0.005)	(0.010)
UK	-0.016^{c}	-0.014
	(0.006)	(0.015)
1995	-0.014^{c}	-0.103^{c}
	(0.005)	(0.011)
1996	-0.022^{c}	-0.208^{c}
	(0.007)	(0.013)
1997	-0.037^{c}	-0.295^{c}
	(0.008)	(0.014)
1998	-0.051^{c}	-0.384^{c}
	(0.010)	(0.015)
1999	-0.072^{c}	-0.503^{c}
	(0.012)	(0.016)
2000	-0.101^{c}	-0.620^{c}
	(0.013)	(0.017)
2001	-0.159^{c}	-0.795^{c}
	(0.017)	(0.018)
Constant	-6.304^{c}	1.405^{b}
	(0.562)	(0.687)
Observations	6941	7893

Note: a, b and c indicate statistical significance at the 10%, the 5 % and the 1% levels, respectively. Robust standard errors clustered at the household level in parentheses. Omitted category is Italy, 1994.

Table 6: Effect of being born to a teenage mother on adult children's education outcomes (LPM)

	POOL	AU	BE	DK	FI	FR	GER	GR	IRE	IT	NT	PO	SP	UK
TeenMum 16														
OLS	-0.023	0.031	-0.172	-0.012	-0.017	0.025^{a}	-0.010	-0.082	-0.124^{c}	0.101	-0.076^{b}	0.035	-0.040	-0.381^{c}
	(0.038)	(0.120)	(0.157)	(0.049)	(0.020)	(0.015)	(0.009)	(0.074)	(0.043)	(0.189)	(0.035)	(0.080)	(0.136)	(0.056)
2SLS	0.016	0.106	-0.665^{a}	-0.060	0.012	-0.147^{a}	0.020	-0.174	0.232	0.056	0.082	0.224	-0.105	0.334
	(0.063)	(0.112)	(0.400)	(0.360)	(0.062)	(0.078)	(0.041)	(0.182)	(0.420)	(0.194)	(0.150)	(0.167)	(0.113)	(0.512)
PE ($\gamma = 0.0068$)	0.014	0.106	-0.665^{a}	-0.078	-0.014	-0.148^{b}	0.016	-0.174	0.226	0.056	0.067	0.223	-0.110	0.320
	(0.059)	(0.112)	(0.400)	(0.358)	(0.061)	(0.075)	(0.039)	(0.182)	(0.437)	(0.194)	(0.139)	(0.175)	(0.115)	(0.503)
Observations	6941	425	215	173	364	478	832	686	339	830	376	719	888	540
TeenMum 19														
OLS	-0.023^{b}	-0.019	-0.053	-0.027^{a}	-0.012^{a}	-0.009	0.015	-0.025	0.057	0.014	0.004	-0.014	-0.077^{b}	-0.119^{c}
	(0.011)	(0.039)	(0.091)	(0.015)	(0.006)	(0.021)	(0.018)	(0.033)	(0.072)	(0.032)	(0.052)	(0.020)	(0.034)	(0.046)
2SLS	-0.034^{b}	0.032	-0.137	-0.034	0.008	-0.052^{b}	0.021	-0.090^{b}	0.130	-0.002	-0.020	0.009	-0.136^{c}	-0.115^{a}
	(0.014)	(0.039)	(0.107)	(0.029)	(0.014)	(0.025)	(0.020)	(0.045)	(0.090)	(0.040)	(0.040)	(0.032)	(0.040)	(0.066)
PE ($\gamma = 0.0068$)	-0.037^{c}	0.032	-0.137	-0.042	0.001	-0.052^{b}	0.019	-0.091^{b}	0.125	-0.002	-0.027	0.009	-0.141^{c}	-0.128
	(0.014)	(0.039)	(0.107)	(0.028)	(0.015)	(0.025)	(0.020)	(0.046)	(0.090)	(0.040)	(0.040)	(0.031)	(0.039)	(0.078)
Observations	7893	482	231	192	399	522	963	836	376	924	399	859	1002	624

Note: a, b and c indicate statistical significance at the 10%, the 5 % and the 1% levels, respectively. Robust standard errors clustered at the household level in parentheses. Country dummies are included in the equations for the pool estimation but not reported. Each cell represents a separate regression. 2SLS: instrumental variables with RoSLA_{t+1} and a third order polynomial for mother's year of birth as exclusion restrictions. PE corresponds to the Plausible Exogeneity approach that uses $\gamma = \rho \times \eta = 0.0068$ where $\rho = 0.068$ is the effect of RoSLA on mother's probability of reaching post-compulsory schooling levels and $\eta = 0.10$ corresponds to the intergenerational transmission of education parameter.

Table 7: Effect of being born to a teenage mother on daughter's fertility outcomes (LPM)

	POOL	AU	BE	DK	FI	FR	GER	GR	IRE	IT	NT	PO	SP	UK
TeenMum 16														
OLS	0.058^c	-0.024	0.000	0.000	-0.095	-0.000	-0.008	0.000	-0.071	0.238^c	0.000	0.117^b	-0.003	0.000
	(0.019)	(0.064)	(0.000)	(0.000)	(0.195)	(0.066)	(0.050)	(0.000)	(0.154)	(0.049)	(0.000)	(0.049)	(0.036)	(0.000)
2SLS	0.075^b	-0.024	0.000	0.000	-0.138	0.014	-0.021	0.000	-0.156	0.167^c	0.000	0.124^a	-0.030	0.000
	(0.030)	(0.091)	(0.000)	(0.000)	(0.205)	(0.155)	(0.067)	(0.000)	(0.304)	(0.064)	(0.000)	(0.066)	(0.081)	(0.000)
PE ($\gamma = -0.00068$)	0.077	0.000	0.000	0.002^b	0.147	0.010	-0.036	0.000	-0.085	0.000	0.003^c	0.209	-0.024	0.240
	(0.048)	(0.000)	(0.000)	(0.001)	(0.556)	(0.012)	(0.045)	(0.000)	(0.151)	(0.000)	(0.001)	(0.193)	(0.035)	(0.256)
Observations	3363	207	101	92	192	239	409	328	160	402	187	347	407	259
TeenMum 19														
OLS	0.040^c	0.011	0.242^c	0.000	0.011	-0.005	0.035^b	0.000	0.137^c	0.039^b	0.000	0.052^b	0.044^c	0.100^b
	(0.007)	(0.027)	(0.045)	(0.000)	(0.045)	(0.015)	(0.015)	(0.000)	(0.045)	(0.017)	(0.000)	(0.023)	(0.014)	(0.040)
2SLS	0.044^c	0.008	0.254^c	0.000	0.055	0.006	0.017	0.000	0.123^b	0.056^b	0.000	0.044	0.048^c	0.119^b
	(0.009)	(0.035)	(0.073)	(0.000)	(0.057)	(0.021)	(0.021)	(0.000)	(0.058)	(0.023)	(0.000)	(0.029)	(0.018)	(0.059)
PE ($\gamma = -0.00068$)	0.043^c	0.000	0.000	0.001^c	0.116	0.015	-0.022	0.000^a	0.039	0.000	0.001^c	0.058	-0.010	0.094
	(0.011)	(0.000)	(0.000)	(0.000)	(0.090)	(0.016)	(0.020)	(0.000)	(0.051)	(0.000)	(0.000)	(0.061)	(0.012)	(0.084)
Observations	3809	236	109	101	212	259	472	393	178	443	195	415	466	294

Note: a, b and c indicate statistical significance at the 10%, the 5 % and the 1% levels, respectively. Robust standard errors clustered at the household level in parentheses. Country dummies are included in the equations for the pool estimation but not reported. Each cell represents a separate regression. 2SLS: instrumental variables with RoSLA_{t+1} and a third order polynomial for mother's year of birth as exclusion restrictions. PE corresponds to the Plausible Exogeneity approach that uses $\gamma = \rho \times \eta = -0.00068$ where $\rho = 0.068$ is the effect of RoSLA on mother's probability of reaching post-compulsory schooling levels and $\eta = -0.01$ is the effect of mother's post-compulsory schooling on daughter's probability of becoming teenage mothers.

Table 8: Effect of being born to a teenage mother on adult children's education outcomes

Outcomes	Children's Education	
Specifications	(1)	(2)
TeenMum 16		
OLS	-0.020	-0.023
	(0.038)	(0.038)
2SLS	-0.151[b]	0.016
	(0.070)	(0.063)
PE ($\gamma = 0.0068$)	-0.154[b]	0.014
	(0.064)	(0.060)
Observations	6941	6941
Partial R^2	0.3403	0.3538
F-test for excl.IVs	86.37	97.00
Prob > F	0.0000	0.0000
Sargan	30.49	14.16
Chi-sq	0.0000	0.0027
TeenMum 19		
OLS	-0.027[b]	-0.023[b]
	(0.011)	(0.011)
2SLS	-0.078[c]	-0.034[b]
	(0.014)	(0.014)
PE ($\gamma = 0.0068$)	-0.081[c]	-0.037[b]
	(0.015)	(0.015)
Observations	7893	7893
Partial R^2	0.5280	0.5406
F-test for excl.IVs	1486.60	1531.43
Prob > F	0.0000	0.0000
Sargan	24.92	19.10
Chi-sq	0.0000	0.0003

Note: a, b and c indicate statistical significance at the 10%, the 5 % and the 1% levels, respectively. Robust standard errors clustered at the household level in parentheses. Each cell represents a separate regression. Specification (1) has no controls and (2) includes a third order polynomial for child's year of birth.

Table 9: Effect of being born to a teenage mother on daughter's fertility outcomes

Outcomes	Daughter's Teenage Motherhood	
Specifications	(1)	(2)
TeenMum 16		
OLS	0.059^c	0.058^c
	(0.019)	(0.019)
2SLS	0.070^b	0.075^b
	(0.031)	(0.030)
PE ($\gamma = -0.00068$)	0.075	0.077
	(0.050)	(0.048)
Observations	3363	3363
Partial R^2	0.3716	0.3836
F-test for excl.IVs	493.43	519.98
Prob > F	0.0000	0.0000
Sargan	5.23	4.77
Chi-sq	0.1433	0.1898
TeenMum 19		
OLS	0.040^c	0.040^c
	(0.007)	(0.007)
2SLS	0.042^c	0.044^c
	(0.009)	(0.009)
PE ($\gamma = -0.00068$)	0.042^c	0.043^c
	(0.011)	(0.011)
Observations	3809	3809
Partial R^2	0.5228	0.5316
F-test for excl.IVs	1036.53	1074.65
Prob > F	0.0000	0.0000
Sargan	4.90	4.51
Chi-sq	0.1792	0.2117

Note: a, b and c indicate statistical significance at the 10%, the 5 % and the 1% levels, respectively. Robust standard errors clustered at the household level in parentheses. Specification (1) has no controls and (2) includes a third order polynomial for child's year of birth.

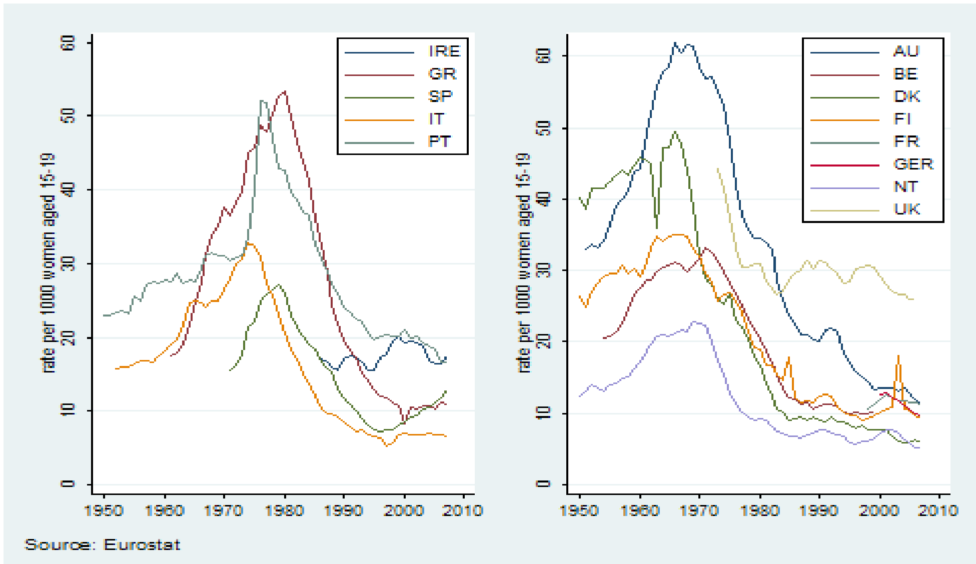

Figure 1: Teenage motherhood rate trends in the last decades in Europe

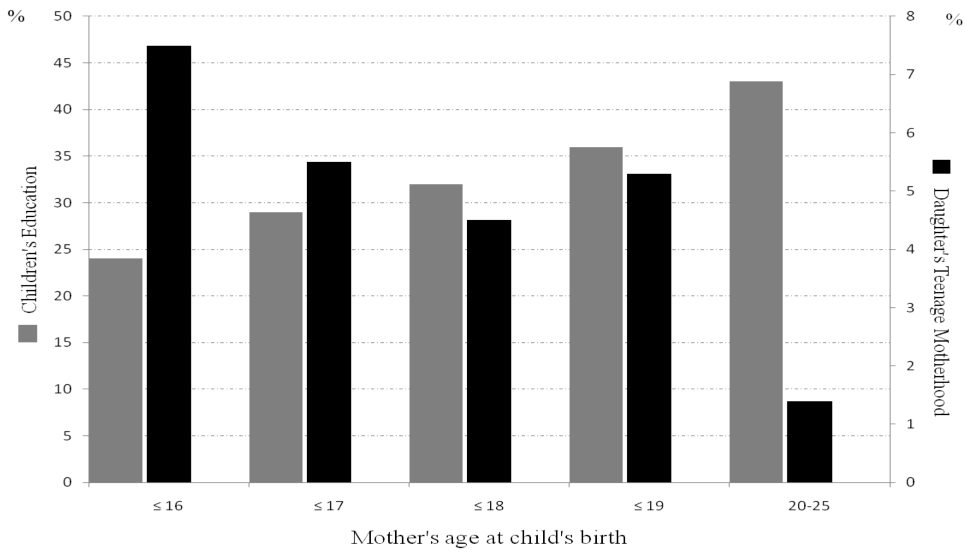

Figure 2: Children with post-compulsory education; Daughter's becoming teenage mothers

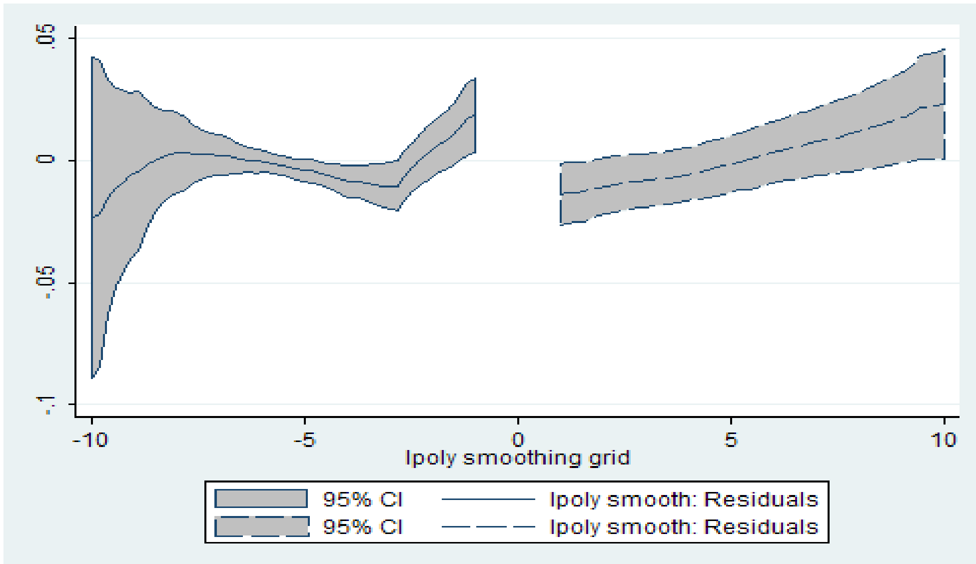

Figure 3: Effect of RoSLA$_{t+1}$ on the probability of first generation teenage motherhood

Appendix

Table A.1: Mean differences of health status, health problems and BMI between pre and post RoSLA mothers

H_0: diff=0	Health status			Health problem			BMI		
Country	diff	t	$P > \|t\|$	diff	t	$P > \|t\|$	diff	t	$P > \|t\|$
Austria	0.26877	3.0577	0.0023	-0.07864	-2.1821	0.0293	2.367951	1.7309	0.0843
Denmark	0.045688	0.4748	0.6351	-0.0302	-0.5886	0.5565	0.249713	0.3998	0.6897
Finland	0.104848	1.0462	0.2957	-0.00375	-0.0569	0.9547	-0.74382	-0.9142	0.3611
France	0.078709	1.5732	0.1159	-0.04559	-2.0682	0.0388			
Germany	0.122265	2.8962	0.0038	-0.07672	-2.8666	0.0042			
Greece	0.277963	6.2095	0.0000	0.006083	0.2997	0.7644	0.837244	2.0423	0.0415
Ireland	0.018787	0.3032	0.7618	0.003676	0.1294	0.8971	0.740306	1.8328	0.0674
Italy	0.165027	4.4035	0.0000	-0.03866	-2.0665	0.0389	0.972199	2.7638	0.0058
the Netherlands	-0.14775	-2.1597	0.031	0.088038	2.1597	0.0311			
Portugal	0.228137	6.1396	0.0000	-0.10362	-4.6624	0.0000	1.112587	2.8325	0.0047
Spain	0.186925	4.6753	0.0000	-0.08941	-4.6142	0.0000	0.845684	2.7794	0.0056
UK	-0.38607	-6.3582	0.0000	0.037155	1.114	0.2655			
Total	0.004967	0.3809	0.7033	-0.04659	-3.21	0.001	0.343028	0.85	0.397

Note: Health status corresponds to the question how is your health in general, Health problem to the question Do you have any chronic physical or mental health problem, illness or disability? and BMI to the mother's body mass index.

Table A.2: First stage: Effect of Placebo reforms on the probability of first generation teenage motherhood (LPM)

	RoSLA_{t-1}		RoSLA_{t-2}	
	TeenMum 16	TeenMum 19	TeenMum 16	TeenMum 19
RoSLA	-0.004	-0.008	-0.001	-0.008
	(0.003)	(0.009)	(0.003)	(0.008)
Birthyear	0.827^c	-0.181^b	0.830^c	-0.176^b
	(0.074)	(0.084)	(0.073)	(0.084)
Birthyear^2	-0.036^c	0.006^a	-0.036^c	0.005
	(0.003)	(0.003)	(0.003)	(0.003)
Birthyear^3	0.001^c	0.000	0.001^c	0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Austria	-0.014^b	-0.091^c	-0.014^b	-0.091^c
	(0.005)	(0.013)	(0.005)	(0.013)
Belgium	0.006	0.016	0.009	0.016
	(0.006)	(0.017)	(0.006)	(0.017)
Germany	-0.005	-0.004	-0.005	-0.004
	(0.003)	(0.010)	(0.003)	(0.010)
Denmark	-0.001	-0.003	0.001	-0.003
	(0.006)	(0.017)	(0.006)	(0.017)
Spain	-0.013^c	-0.044^c	-0.012^c	-0.044^c
	(0.004)	(0.011)	(0.004)	(0.010)
Finland	-0.001	0.025	0.002	0.026^a
	(0.005)	(0.016)	(0.004)	(0.016)
France	0.004	0.021^a	0.005	0.021^a
	(0.003)	(0.012)	(0.003)	(0.012)
Greece	0.010^b	0.017	0.010^b	0.017
	(0.005)	(0.011)	(0.005)	(0.011)
the Netherlands	0.002	-0.007	0.004	-0.006
	(0.005)	(0.014)	(0.005)	(0.013)
Ireland	-0.005	-0.015	-0.003	-0.014
	(0.004)	(0.013)	(0.004)	(0.013)
Portugal	0.010^b	0.014	0.010^b	0.014
	(0.005)	(0.010)	(0.005)	(0.010)
UK	-0.007	-0.004	-0.005	-0.003
	(0.005)	(0.013)	(0.004)	(0.013)
1995	-0.014^c	-0.103^c	-0.014^c	-0.103^c
	(0.005)	(0.011)	(0.005)	(0.011)
1996	-0.023^c	-0.209^c	-0.023^c	-0.209^c
	(0.007)	(0.013)	(0.007)	(0.013)
1997	-0.037^c	-0.296^c	-0.037^c	-0.295^c
	(0.008)	(0.014)	(0.008)	(0.014)
1998	-0.051^c	-0.384^c	-0.051^c	-0.383^c
	(0.010)	(0.015)	(0.010)	(0.015)
1999	-0.072^c	-0.502^c	-0.072^c	-0.502^c
	(0.012)	(0.016)	(0.012)	(0.016)
2000	-0.102^c	-0.621^c	-0.102^c	-0.621^c
	(0.013)	(0.017)	(0.013)	(0.017)
2001	-0.159^c	-0.795^c	-0.159^c	-0.795^c
	(0.017)	(0.018)	(0.017)	(0.018)
Constant	-6.278^c	1.423^b	-6.297^c	1.380^b
	(0.570)	(0.687)	(0.563)	(0.682)
Observations	6941	7893	6941	7893

Note: a, b and c indicate statistical significance at the 10%, the 5 % and the 1% levels, respectively. Robust standard errors clustered at the household level in parentheses. Omitted category is Italy, 1994.

Table A.3: Effect of Teenage Motherhood on daughter's probability of giving birth in their teens. Sisters fixed effects estimation

Countries	Observations	Teenage Motherhood
POOL	1074	0.088^c
	675	(0.022)
AU	70	0.000
	44	(0.000)
BE	24	1.000
	16	(0.000)
DK	28	0.000
	18	(0.000)
FI	44	0.000
	30	(0.277)
FR	64	0.000
	41	(0.000)
GER	130	0.167^b
	79	(0.067)
GR	123	0.000
	70	(0.000)
IRE	71	0.187
	47	(0.160)
IT	123	0.000
	84	(0.091)
NT	49	0.000
	32	(0.000)
PO	131	0.119^a
	79	(0.070)
SP	154	0.064
	98	(0.047)
UK	60	0.200
	35	(0.122)

Note: First row in the Observations column corresponds to the number of sisters whereas the second row is the number of different households.