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ORIGINAL RESEARCH

Gender Gap in Intergenerational Educational Persistence: Can Compulsory Schooling Reduce It?

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Abstract
We analyze the impact of an increase in compulsory schooling policy on the gender gap in intergenerational educational persistence using the nationally representative Turkish Adult Education Survey. Prior to the reform, there is a gender gap in the association of parents' educational attainment with their offspring's. Daughters' educational attainment is more dependent on their parents' education background. We show that the education reform that increased compulsory schooling from 5 to 8 years reduced the impact of parental education on completion of new compulsory schooling (8 years) and post-compulsory schooling (high school) for both sons and daughters. The gender gap in intergenerational education transmission has decreased by about 5 percentage points in the completion of new compulsory schooling level but remains unchanged at the post-compulsory schooling level after the reform. Heterogeneous effects of the reform indicate that mandating additional years of education is an ineffective intervention in the eastern regions with poorer economic conditions, larger rural population, and more traditional gender views in reducing the gender gap in educational mobility, even at the compulsory level of education.

Keywords Intergenerational education transmission · Gender equality · Compulsory schooling

JEL Classification I20 · I24 · J16 · J62

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Introduction

Educational achievements are influenced by parental human capital in almost all societies. When educational attainment is inherited from one generation to the next, chances of success are largely pre-ordained by the accident of birth, which goes against a basic notion of fairness. Low intergenerational mobility also leads to unrealized human potential and misallocation of resources, as talented individuals from disadvantaged families are excluded from opportunities that favor those born into greater privilege rather than those with the greatest potential (Björklund & Salvanes, 2011). Hence, the education reforms that aim to increase equality and efficiency by reducing the impact of family background have been the focus of recent policy debates.

Large regional differences in intergenerational educational persistence exist, with Latin American countries displaying the highest intergenerational correlations between parent's and offspring's education, and the Nordic countries the lowest (Hertz et al., 2008),[1] suggesting that public policies can play a role in increasing intergenerational mobility. Also documented are large gender differences in the intergenerational transmission of education as daughters exhibit more persistence than sons suggesting that parental attitudes toward sons and daughters can also play a role in intergenerational mobility (Emran & Shilpi, 2015; Glick & Sahn, 2000).

We investigate the impact of the compulsory schooling law on intergenerational educational persistence both at the compulsory and post compulsory levels in a developing country exhibiting persistent dependence of offspring's educational attainment on parental education even at the compulsory level. To our knowledge, ours is the first paper to examine the effect of compulsory schooling on gender gap in intergenerational transmission of education. We leverage a Turkish education reform which went into effect in 1997. For political reasons, the reform was implemented quickly and rather unexpectedly, and it increased the mandatory years of education from 5 to 8 years. Students who had completed the fifth grade in the Summer of 1997 have been exempted from the mandate of the reform, while younger cohorts were required to complete at least 8 years of schooling. Since the school starting age is 6 in Turkey, the reform affected cohorts of offspring born after 1986, while it had no effect on those born before. We analyze the effect of compulsory schooling policy on educational persistence using regression discontinuity design. There is a discontinuous jump at the time of policy in the offspring's education, and the relationship between the outcome variable (offspring's education) and the running variable (offspring's year of birth) is continuous.

We capture the transmission of education from parents to their offspring by the regression coefficient of parental education as a predictor of schooling in the next generation. Our identification strategy examines the relationship between parental education and offspring's educational attainment differentiated according to offspring's exposure to the compulsory schooling reform by exploring the

[1] Hertz et al. (2008) demonstrate that the global average correlation between parent and child's schooling has held steady at about 0.4 for the past 50 years.

heterogeneous effects of the policy by parental education. We capture the offspring's educational attainment by two binary variables: completion of (at least) new compulsory (8 years) and post-compulsory (high school) level of education. We are primarily interested in our results at the post-compulsory level of education, which shows the impact of greater access to education. However, we still investigate the effect of the policy at the new compulsory level of education for two reasons: First, the compliance with the old law was not perfect and even lower for girls than boys. Second, investigating the effect at the post-compulsory level is a meaningful question only if we can establish the impact at the new compulsory level.

Our results show that the compulsory schooling law reduced the impact of parental education (either parent's having at least 12-year high school degree) on completion of new compulsory schooling (8-year primary education) from 31 to 8 percentage points for men and from 46 to 17 percentage points for women. At the post-compulsory schooling level, the association between parent's and offspring's education has decreased by 8 percentage points for both men and women. Prior to the reform, the effect of parental education was higher for women than men both at the new compulsory and post-compulsory schooling levels. The gender gap in intergenerational persistence in education has significantly decreased by about 5 percentage points in the completion of new compulsory schooling level but remains unchanged at the post-compulsory schooling level after the reform. These results are robust to alternative sample specifications and time trend controls.

The association between the educational decisions of parents and their children could come from either causal (the direct influence of parent) or non-causal (genetic inheritance or sharing a similar environment) channels. Empirical studies use twin and adoptee studies (Behrman et al., 1994; Currie & Moretti, 2003; Rosenzweig & Wolpin, 1994; Sacerdote, 2002, 2004) and instrumental variables method where a variable is used as an instrument for parent's education to solve for the omitted variable bias due to unobserved family ability: Oreopoulos et al. (2006) in the USA, Chevalier (2004) in the UK, and Black et al. (2005) in Norway. However, these studies do not examine how this causal effect changes when children are exposed to extended compulsory schooling. Compulsory schooling laws are a common policy tool to achieve greater participation in education, particularly for marginalized groups. Hence, in that regard, our exercise is an important contribution to studies examining how an extension in compulsory schooling can mitigate the association of parents' education with offspring's education and thereby interrupt the persistence of marginalization in education outcomes across generations. While we do not study the causal effect of parental education on offspring's education, we examine the effect of exposing offspring to compulsory schooling on intergenerational educational persistence both at the compulsory and post compulsory levels where intergenerational educational persistence is defined as the association between parental education and offpsring's education. Studies in this literature are all from advanced economies with high levels of income equality and intergenerational educational mobility (Aakvik et al., 2010; Betthäuser, 2017; Meghir & Palme, 2005). Hence, it is a mostly unexplored question to what extent policy interventions can effectively increase educational mobility in low equality and low mobility society. One recent example by Urbina (2018) investigates the role of school expansion in Mexico, one of the countries shown to have the lowest mobility by Hertz et al. (2008). Her findings show that school expansion policy increases intergenerational mobility only at the lower levels of education, while mobility at the higher levels of education decreases. Turkey, an emerging economy with low-income equality, has lower intergenerational educational mobility compared to these developed countries and higher mobility than Latin American countries (Aydemir & Yazici, 2019). Hence, it provides a compelling environment to examine this question where it is unexplored yet.

Our work also contributes to the literature on gender inequality and the role of government policies in bringing about equality between men and women (Dayioglu et al., 2016; Fang et al., 2012; Spohr, 2003; Tsai et al., 2009;). Several studies identified differential preferences of parents for their sons and daughters as a source of women's lower educational attainment in Turkey (Caner et al., 2016; Rankin & Aytac, 2006), West Africa (Glick & Sahn, 2000), Brazil (Emerson & Souza, 2007), and India (Kingdon, 2002, 2005; Pal, 2004). Children of parents with low levels of education are more likely to be constrained by the monetary and non-monetary cost of education, in addition to the higher psychic cost for daughters than sons due to less gender-equal views regarding educational and labor market outcomes of their offspring (Akyol & Okten, 2019). Hence, it is an important policy question to what extent government policies can mitigate the adverse effects of the role parental education play in leading to low intergenerational mobility for daughters.

Our findings show that the effect of compulsory schooling on increasing intergenerational educational mobility is strong as the persistence has decreased substantially for both men and women at the new compulsory and post-compulsory level of education. The gender gap in the differential dependence on parental education for daughters also reduces with the policy at the new compulsory level, showing the effectiveness of the policy change on improving equality. However, positive spillover effects of the reform to post compulsory levels appear to be limited as the gender gap in intergenerational persistence in education remains unchanged at the post-compulsory schooling level alluding to the strong and enduring effect of parental education. This finding also provides support for the existing research which finds that the positive spillover effects of policies implemented to reduce gender gap tend to be limited, and continuous policy commitment is needed to bring about equality between men and women.[2]

We further examine the heterogeneous impact of parental education as we refine our parental education variable according to number of parents with secondary education and also whether a parent has tertiary education. The reform is equally effective in bringing about equality between offspring without any secondary educated parents and offspring with one parent or two parents with secondary education. In terms of closing the gender gap in education mobility at the compulsory level, the gains largely come from closing the gap between daughters without any secondary educated parents and daughters with a tertiary-educated parent. We also examine the heterogeneous effects of the reform across regions and find that mandating

[2] For details, see the paper by Duflo (2012) that surveys the literature suggesting the limited spillover effects of policies targeting gender equality.

additional years of education is an ineffective intervention in the eastern regions with poorer economic conditions, larger rural population, and more traditional gender views in reducing the gender gap in educational mobility, even at the compulsory level of education.

The remainder of the paper is organized as follows: "The Compulsory Schooling Reform in 1997" and "Conceptual Framework" sections introduce the compulsory schooling reform and the conceptual framework. Sections "Data" and "Identification" sections explain the data and empirical strategy. "Results" section presents and discusses the results and robustness checks. Finally, "Conclusion" section states the concluding remarks.

The Compulsory Schooling Reform in 1997

In Turkey, formal education has been managed by the central government since 1923, and the Ministry of National Education (MONE) is in charge of all structural reforms and education policies.

Prior to 1997, the school system in Turkey consisted of mandatory 5 years of primary education, followed by voluntary 3 years of lower secondary and 3 years of upper secondary education. Students could choose from either general, or vocational schools for the voluntary part. In Turkey, vocational schools include both technical vocational schools teaching subjects like electronics, machinery, wood craft, etc. and religious vocational schools. Religious vocational schools (also known as *Imam Hatip* schools) are formal education institutions that were founded in lieu of a vocational school to train government employed imams and religious spokespersons. In Imam Hatip schools, students are taught religious subjects alongside a more general education. Hence, prior to the reform, a student who completed primary school in 5 years could drop out (to either quit education completely or attend informal religious education) or continue studying in general or vocational schools. Turkey's law required education to be provided only in mixed-sex schools and prohibited wearing a headscarf. However, in practice, in religious vocational schools, girls were allowed to wear a headscarf. During the 1990s, political Islam gained substantial public support, leading to an Islamist party's (namely Welfare Party) win the general elections. Following the right wing's winning the 1995 national elections, the conflicts between secular and religious groups increased and led to the military intervention with a set of decisions aiming to prevent the spread of the Islamic movement in Turkey. The compulsory schooling law was among these decisions and enforced with Law 4306 by the end of August 1997 (Erten & Keskin, 2018). The law was enacted and prolonged the mandatory education from 5 years of primary education to 8 years by incorporating junior high school into compulsory education and delayed starting a vocational school including religious vocational schools by 3 years. In other words, vocational schools were no longer offered at the junior high school level and were available only at the high school level.

The amendment requires all children to complete 8 years of schooling without any exceptions for any group of children. Parents who do not comply with the law face monetary penalties and even possible incarceration in case of repeated noncompliance. However, despite these measures, noncompliance is quite common before and after the policy (Dayioglu et al., 2016).

This law restricts religious education in two ways. First, students could not attend religious vocational schools after completing 5 years of education and instead had to complete three more years in general schools. Second, students who attend informal religious education immediately upon completion of compulsory education had to delay their attendance for 3 years (Aydemir & Kirdar, 2017).

In order to make the law applicable in a short amount of time and to ensure that all students were schooled without a significant decline in education quality, the MoNE introduced a number of measures (Dulger, 2004; Aydemir & Kirdar, 2017; Dinçer et al., 2014) Güneş (2015) including increasing the number of classes and teachers in the existing schools, bussing children from rural areas to nearby schools, and establishing boarding schools for children living in distant rural areas. Students from low-income families receive free textbooks and school meals starting from 2004. In 1998–2000, the public investment in education was around 30%, while prior to compulsory schooling law, it was about 15% (Dulger, 2004; Dayioglu et al., 2016).

The compulsory education law came into force in August 1997, just before the 1997–1998 school year. Students who had not completed their 5-year education on this date and had not been awarded the primary school diploma are required to complete 8-years of schooling to earn a primary education diploma. Thus, students who started primary school in September 1993 or afterward were bound by the policy, whereas students who started school in September 1992 or earlier were exempt from it. The official school starting age in Turkey is six, but in practice many individuals start primary education at age seven. Hence, individuals born after 1986 are exposed and those born before 1986 are not exposed to the compulsory schooling reform, while the information on the year of birth alone is not conclusive on the exposure to the reform for those born in 1986. Since the primary school diploma is the first official proof of schooling one can earn, the incentive to finish the compulsory level of education is strong for students and their families.

The schooling law has increased school enrollment dramatically. Before the compulsory schooling policy, the schooling rate was 53% (1996–1997 school year) in the 8-year primary education and increased to 85% even in the first year of policy change (1997–1998 school year). In 3 years, the enrollment had increased to 95% (2000–2001 school year). Similarly, the secondary education enrollment rate had increased from 38% in 1997–1998 to 44% in 2000–2001 (MONE Education Statistics, 2006–2007).

We further present the change in school attainment with the nationally representative micro-level data from Turkish Adult Education Survey. Panel (a) of Fig. 1 displays the share of individuals who completed at least 8-year primary education by birth year separately for men and women. Separate linear polynomials are fitted on both sides of the cutoff, and the cohort born in 1986 is excluded due to the fuzziness in their treatment status. The fitted polynomials show significant discontinuities before and after the cutoff for both men and women. Mean educational attainment of those born in 1986 falls between the trends for pre- and post-policy cohorts, confirming their partial exposure. The increasing trends are similar for both genders

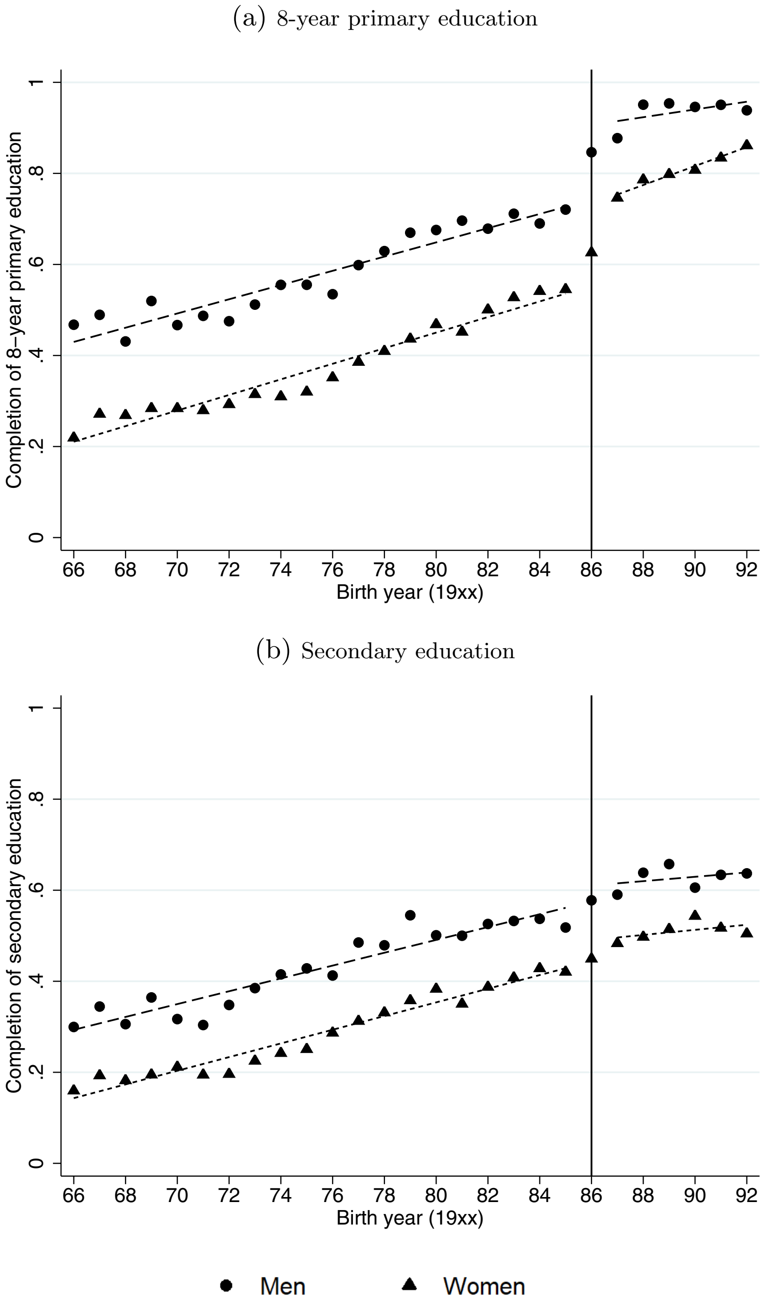

Fig. 1 Mean education completion rates by birth year. *Source* Adult Education Survey, 2012 and 2016. The points give the mean completion rates for each education level, 8-year primary and secondary education, across year of birth. Local linear polynomials are fitted on both sides of the cutoff

not exposed to the reform. For the cohorts exposed to the policy, men's educational attainment plateaus with their higher compliance with the law, and women's schooling exhibits a similar increasing trend similar to pre-reform cohorts in addition to the jump at the cutoff. Panel (b) of Fig. 1, which presents the fractions of birth cohorts completing at least high school for the same birth cohorts, shows the relatively smaller spillover effects of the policy on secondary education attainment for both sexes. The panel indicates around a five percentage point increase right before and after the cutoff, with a change in the trend for high school completion. Both men and women's time trend follow a flatter path after the jump at the cutoff.

Conceptual Framework

In this section, we first outline a simple model of optimal educational attainment decisions. Within this model, we point out the potential sources of inequalities by parental education and gender. Next, we discuss how the new compulsory schooling policy changes the cost and benefits of education and, therefore, the existing disparities.

An individual's educational attainment decision is determined by a trade-off between the discounted value of higher future earnings capacity and the direct and indirect costs of education in the present. The contribution of education to future earnings is modeled in two theories in the literature. The human capital theory suggests that education is an investment activity raising future earnings by increasing worker productivity (Becker, 1975; Mincer, 1974). According to the signaling theory, employers resolve the level of education signal sent by employees to infer the true worker productivity (unobserved by the employer), and therefore, the level of education raises the future earnings (Spence, 1978). Hence, there is an additional benefit of completing a certain level of education, known as *sheepskin effect*, in addition to its productivity increasing effect. The cost of education includes direct monetary costs like transportation costs and purchases of school supplies, and indirect costs in the form of the opportunity cost of school time and the psychic costs of sending children to school.

Both perceived benefits and incurred costs of education can affect the association between parent's and offspring's education in several ways. First, parents with higher education may be more aware of the returns to education, while lower educated parents underestimate the true value of higher future earnings to be gained through education.[3] Second, well-educated parents may provide better job opportunities for their children in the event that they attain higher levels of education via their social networks, whereas children born to low educated parents may face lower returns to education due to their family's lack of access to such networks. Finally,

[3] Jensen (2010) shows the underestimation of returns to education in the Dominican Republic and argues that students and their parents rely mainly on the earnings of the workers that they can observe in forming their expectations on returns to education. Therefore, the residential and social segregation by income could lead to underestimates of returns to education in low-income low education settings.

higher educated parents may have higher incomes, and their children may be less constrained with the direct and opportunity cost of education. These can be interpreted as the causal channels leading to lower educational attainment of offspring born to lower educated families. In addition to aforementioned causal effects of parents' education and income, the link between the schooling of parents and their children could be due to unobserved inherited family characteristics. Higher educated parents may have higher abilities, and so may their children. Unobserved and inherited family characteristics such as genetic endowment can influence both parents' and offspring's education levels, also known as ability bias in the literature.

We present completion rates of 5 years of compulsory education, 8 years of education and secondary education for pre-policy cohorts born in 1976–1985 according to their parental education levels in Table 1. While those with educated parents almost perfectly comply with 5 years of compulsory education, 8 percent of those with less educated parents do not comply with the law. The education gap between those with high and low educated parents is 46 percentage points in 8 years of education and rises to 53 percentage points at the secondary level of education. Note that statistics reported in this table are for cohorts not exposed to the reform. As such, both 8 years of education and secondary education are post-compulsory schooling levels for these cohorts. The significant differences in the schooling of offspring born to families with differential educational backgrounds provide evidence supporting the aforementioned channels through which parental education may affect offspring's education level.

Within this framework, we further argue that perceived value of future earnings is higher for boys and incurred costs of education are higher for girls favoring boys' education over girls'. First, parents might value future earnings of their daughters less than their sons due to low labor market expectations for women compared to men,[4] and thereby discounting their daughters' future income more. Similarly, the value of future earnings would be discounted more for girls as daughters are more likely to move away from their parents after marriage (Dayioglu et al., 2016).

While the direct monetary cost of education is not expected to differ by gender, the opportunity cost of sending daughters to school might be higher than sons (Rankin and Aytaç 2006), since parents might value their daughters' nonmarket time contributing to home production. Furthermore, the psychic cost of education might be especially high for girls due to conservative social norms. Due to the cost of traveling away from home to school at puberty and education carried out in mixed-sex schools in Turkey, socially conservative families may incur a higher psychic cost of educating daughters. Indeed, several studies, which examine data from three different decades, found lower educational attainment of girls in Turkey and its relation to the persistence of the view on gender roles in a conservative society, confirming higher psychic costs for daughters than sons. Rankin and Aytaç (2006) find that patriarchal family beliefs and practices discourage the education of girls beyond the compulsory level using the 1988 Turkish Family Structure Survey. Caner

[4] The female labor force participation in Turkey, at 28.9%, is the lowest among OECD countries and the 22nd lowest in the world (World Bank, 2018).

Table 1 Educational completion percentages of offspring by parental education

	Both genders		Men		Women	
	Neither parent has at least second. educ. (%)	Either parent has at least second. educ. (%)	Neither parent has at least second. educ. (%)	Either parent has at least second. educ. (%)	Neither parent has at least second. educ. (%)	Either parent has at least second. educ. (%)
Completing 5-year education	92	100	97	100	88	99
Completing 8-year education	48	94	60	96	39	92
Completing secondary education	35	88	43	89	28	87
Completing higher education	14	61	17	62	12	60
Observations	9894	1750	4185	809	5709	941

Authors' calculations using AES, 2012 and 2016. Education completion percentages are calculated for individuals born in 1976–1985 (10 cohorts not exposed to the reform). Three samples are determined by the offspring's gender as both gender, men, and women. Next, each of these three samples is divided into two mutually exclusive sub-samples by the parental education, as those whose neither parent has secondary education and those whose either parent has secondary education. Hence, six samples of interest are determined by the gender and the parental education. Each number in a cell gives the offspring's mean completion rates for the associated level of education, given in each row, on the sample, given in each column et al. (2016) similarly show that a traditional view on gender roles adversely affects the educational attainment of daughters by analyzing the Turkish Demographic and Health Surveys conducted in 1998 and 2003.

Low parental education may exacerbate aforementioned factors that result in favoring boys' education over girls. First, labor force participation of women increases with educational attainment in Turkey (Dayıoğlu & Kırdar, 2010). Families with higher educational backgrounds may discount their daughters' future incomes less by foreseeing a higher labor force participation of their daughters. Second, the foregone home production of girls which is an opportunity cost of education may be higher in low educated families since they tend to be larger in size (Schultz, 2002). Third, conservative gender views, which lead to lower demand for girls' education, are more prevalent in low educated families (Akyol & Okten, 2019) leading to higher psychic cost for girls born to low educated parents. Indeed, in AES data, the gap in the education completion rates of children born to high and low educated families is larger at all levels of education for women compared to men (Table 1). Among those not exposed to the extension of compulsory schooling (individuals born in 1976–1985), the completion rate of 5 years of education for men whose both parents are low educated is only 3 percentage points lower than men who had at least one educated parent. The same difference is 11 percentage points for women. Hence, women's educational attainment depends more on parental education even at the compulsory level due to incompliance with the law. The educational gap in the completion of 8 years of education by parental education is a striking 53 percentage points for women and 36 percentage points for men. The gap between offspring with no parent with a high school degree and the ones with at least one such parent also exists at the secondary level (59 percentage points for women and 46 percentage points for men), as shown in Table 1.

Extending mandatory years of schooling in Turkey increases the attainment of the new compulsory level of education by affecting both the benefits received by attendance and diminishing the cost of education. As discussed in the institutional context, with the enforcement of the policy, the government increased classroom capacities all around the country and reduced the cost of transportation and school supplies. Since the policy is additionally enforced with monetary penalties in case of noncompliance[5], it increased the cost of not attending the new compulsory level of education. Hence, these penalties can be interpreted as the negative cost of education, further reducing the cost of education at the compulsory level. This fall in the monetary cost of education is expected to increase the educational attainment of those with higher price elasticity of demand. Orazem and King (2007) state, by examining a large body of empirical studies, that the demographic groups having lower educational attainment have higher income and price elasticities. By this discussion, the

[5] Parents who do not comply with the law face only monetary penalties for the first two violations. In the third violation, there is the risk of additional penalties and incarceration. After four or more violations, parents can be sentenced to up to 6 months in prison. The local administrative body and teachers are responsible for tracking the compliance of the students. However, noncompliance with the law is common both before and after the policy (Dayioglu et al., 2016).

fall in the cost of education should diminish the existing disparities across genders and regions. However, the enforcement of compulsory schooling policy and the penalties associated with it might vary across regions depending on local administrators' and teachers' attitudes as they are responsible for tracking students' compliance with the law. On the other hand, if the psychic cost of sending a daughter at the age of puberty is higher than the psychic cost of sending a son to school at the same age, increased costs of not attending school in terms of penalties imposed for noncompliance may have a higher marginal positive effect on sons than daughters in terms of junior high school attendance which is the new compulsory level.

Furthermore, the compulsory schooling reform extended general (secular) education by 3 years and eliminated religious vocational schools at the junior high level. This might increase the psychic cost of education for girls. Religious vocational schools allow girls to get an education in same-sex classrooms and wear headscarves in classroom, and include both religious and general courses in their curriculum. Low educated parents who are also socially conservative might be less willing to send their daughters to mixed-sex schools at the age of puberty and therefore, not comply with the new policy. Nevertheless, if compliance with the compulsory schooling law is achieved, this can have positive spillover effects for higher levels of education. Parents who were not inclined to send their daughters to junior high school fearing the ill effects of a mixed-sex school environment may have a favorable experience, and therefore, their psychic cost of sending their daughters to school beyond the compulsory level might decrease. Daughters who are older can also influence their parents in continuing with their education.

Regarding the benefits, the compulsory schooling policy increases worker productivity besides substantially altering the sheepskin effects of education. The policy incorporates primary and lower secondary education under the new compulsory level of education and, therefore, abolishes the primary school diploma. This change eliminated the sheepskin effect of 5 years of education entirely. Since the primary school diploma is the first official proof of schooling one can earn in the absence of the policy, eliminating the sheepskin effect of it encourages students who would have finished only 5 years of education in the absence of the policy to complete at least 8 years of schooling to distinguish themselves from those without any official proof of education. The policy additionally diminishes the sheepskin effect of the 8 years of education since it becomes the lowest level of education that one can earn after the policy. A large share of boys and girls born to educated families completed 8 years of education even before the enactment of the law (with 96% for men and 92% for women). Hence, the transformed sheepskin effects are likely to increase the educational attainment of those born to lower educated families, with a possibly higher increase for boys since their labor market participation prospects are higher and psychic costs are lower.

These changes in sheepskin effects might also explain why many studies documented the spillover effects of the compulsory schooling change on the high school degree attainment (Dayioglu et al., 2016). For example, the individuals who would only complete 8 years of education in the absence of the policy, and therefore, gain a degree that is higher than the compulsory level of education, need to extend their education to eleven years to gain the marginal benefit of higher than compulsory level of education. Besides, after being compelled to complete another 3 years of education, some students might change their decision to attain secondary education due to improved information on returns to education and reduced marginal cost of completing high school. Children who are 14 years old instead of 11 when they make the decision to continue their education beyond the compulsory level maybe in a better position to persuade their parents to continue their education.

In sum, we expect compulsory schooling law to help children from low educated families to catch up with their peers from high educated families at least at the compulsory level with possible spillover effects at the post compulsory level. There is a gender gap in education especially for those born to low educated parents. If compliance with the law is achieved, we also expect gender gap in education for offspring born to low educated parents to close at the compulsory level. However, there are considerable opportunity costs and psychic costs of sending daughters to schools and these costs are likely to be higher for low educated families thereby constituting an impediment for girls from low educated families to catch up with their peers from high educated families not only at the post compulsory level but also at the compulsory level (Turkstat, 2012).

Data

The empirical analysis of the intergenerational transmission of education requires a dataset with information on the education of individuals and their parents. Here, we use the Adult Education Survey (AES) conducted in 2012 and 2016 by Turkstat, which provides information on adults aged 18 and older. The major advantage of these data is that it allows the matching of an individual's educational attainment to her parental educational outcome, even if parents and offspring do not share the same residence. This is the first study, to our knowledge, that explores the heterogeneous effects of the Turkish compulsory school reform on educational attainment according to parents' educational backgrounds. In addition to the educational attainment of individuals and their parents, the AES provides information on individual characteristics like age, gender, the region, and type of residence (only available in 2012 wave).[6]

AES provides schooling outcomes as the highest level of education that an individual has completed. Using this information, we have constructed four different dummy variables that show whether individuals have completed at least each of the four levels of education: 5-year primary, 8-year primary (junior high school), secondary (high school), and tertiary (college and above) education. We aim to analyze the changes in the association of parents' educational background with the education outcomes of their offspring due to the compulsory schooling law. Thus, we use completion of 8-year primary education (the new compulsory level) and secondary education (post-compulsory level) binary variables as dependent variables instead of years of schooling.

[6] We use observations from 2012 wave of the AES for our analysis relying on regional information.

The parental educational attainment is given in three categories in the AES data: 8-year primary education or lower, secondary education, and tertiary education. Using these categories, we construct several education variables to measure different aspects of parental education. First, we construct a dummy variable indicating whether father (mother) has completed at least secondary education attainment, called father's education (mother's education). Second, we construct a dummy variable indicating whether father (mother) has also completed tertiary education called father's tertiary education (mother's tertiary education). Third, we construct a composite measure called parental education, that takes the value of 1 if either parent has completed at least secondary education and 0 if neither parent has completed secondary or higher education. This binary parental education variable is our key explanatory variable in the rest of the paper.

To investigate the heterogeneous effects of parental education on offspring's education, we refine our parental education variable further and construct additional composite measures of parental education. First, we want to understand the differential effects of having one parent versus having two parents with secondary education compared to neither parent having secondary education. To that end, we construct a dummy variable that is equal to one if only one parent has completed high school and zero otherwise and is called one parent is educated and a second dummy variable that is equal to one if both parents have completed high school and is called both parents are educated. Note that parental education equals 1 if either one parent is educated or two parents are educated variables take value 1, and therefore, these two variables can be considered as subcategories of our key explanatory variable of interest. Second, we want to understand the differential effects of having one parent with tertiary education versus having one or two parents with only secondary education compared to neither parent having a secondary education. Hence, we construct a dummy variable called secondary education which takes value 1 if the offspring has one or two parents with at **most** secondary education and zero otherwise, and another dummy variable called tertiary education which takes value 1 if either parent has completed tertiary education. Similarly, parental education equals 1 if either secondary education or tertiary education variables take value 1, and so, these two variables are considered as subcategories of parental education.

Table 2 presents summary statistics for both genders in column (1) and also for men and women separately in columns (2) and (3) for cohorts born in 1980–1992 because the estimated bandwidths fall in this range. Panel A of Table 2, which summarizes the statistics for offspring's educational outcomes, indicates that 91% of these cohorts complete their 5-year education, the compulsory level of education until 1997 reform, while the completion rate was around 71% for the new compulsory level of education. Fifty percent of the same sample additionally completed high school. Column (4) tests for differences in the group means given in columns (2) and (3) for men and women, respectively. The share of those who completed old and new compulsory levels of education among women was 9 and 17 percentage points lower, respectively, than the same rates for men. The gender difference in high school completion is 12 percentage points, to the disadvantage of women. Overall, these statistics point out a gender gap in educational attainment at all levels of education including the compulsory level.

Table 2 Summary statistics for individuals born in 1980–1992

	All mean (SD) (1)	Men mean (SD) (2)	Women mean (SD) (3)	Difference (2)–(3) Estim. (SE) (4)
Panel A: education outcomes				
Completed 5-year education	0.908	0.961	0.868	0.093
	(0.289)	(0.194)	(0.338)	(0.005)
Completed 8-year education	0.710	0.809	0.636	0.173
	(0.454)	(0.393)	(0.481)	(0.008)
Completed secondary education	0.498	0.569	0.446	0.123
	(0.5)	(0.495)	(0.497)	(0.009)
Panel B: parental education outcomes				
Father's (secondary) education	0.174	0.190	0.161	0.029
	(0.379)	(0.393)	(0.368)	(0.007)
Father's higher education	0.066	0.074	0.060	0.014
	(0.248)	(0.263)	(0.237)	(0.004)
Mother's (secondary) education	0.076	0.084	0.071	0.013
	(0.266)	(0.277)	(0.257)	(0.005)
Mother's higher education	0.024	0.027	0.022	0.005
	(0.154)	(0.163)	(0.146)	(0.003)
Either parent has at least secondary education	0.188	0.206	0.175	0.031
	(0.391)	(0.405)	(0.38)	(0.007)
Both parents have at least secondary education	0.062	0.068	0.058	0.010
	(0.241)	(0.252)	(0.233)	(0.004)
Either parent has at least higher education	0.072	0.080	0.066	0.014
	(0.259)	(0.272)	(0.249)	(0.004)
Both parents have at least higher education	0.018	0.021	0.015	0.006
	(0.133)	(0.145)	(0.123)	(0.002)
Observations	13,807	5896	7911	

Authors' calculations using AES 2012 and 2016. The sample includes those born in 1980–1992. Columns 1–3 report means and standard deviations in parentheses. Column 4 reports differences in the group means between columns 2 and 3 with standard errors in parentheses

We present the statistics for parental education in Panel B. We observe that the completion rate for the same level of education is relatively lower for parents compared to children, with a completion rate of high school around 6–19% for parents and 50% among offspring. 17% of fathers and 8% of mothers completed high school, while these rates are 7% and 2%, respectively, for college completion, indicating a gender gap in the educational outcomes of the previous generation. Consistent with the educational gender gap observed for parents, the mean share of offspring with at least one high school graduate parent is only slightly higher than the share of those with a high school graduate father. This close relationship signals that the positive values for parental education are driven mainly by fathers' secondary educational attainment rather than that of mothers. In contrast, the share of offspring with both parents having secondary education is only slightly lower than the share of those with a secondary educated mother, since mothers' lower educational attainment than fathers' makes it the primary determinant of both parents' educational attainment.

An unusual feature observed in Table 2 is that parental education is systematically lower for daughters than sons. Though the available data do not allow us to further investigate the underlying reason for this phenomenon, we conjecture that this could be due to lower educated families favoring sons and having more female children until they have at least one son (Altindag, 2016).[7] We take this relation into consideration in our regression analysis by defining gender-specific covariates.

Identification

Students who do not have a primary school diploma by the beginning of the 1997–1998 school year are the ones affected by the policy change. Since school starting age is 6 in Turkey, we assume those born after 1986 must complete 8 years of schooling, while those born in 1986 or earlier could drop out after 5 years, as explained in the compulsory schooling reform section. We use this break in a regression discontinuity design to estimate the causal impact of extending mandatory years of schooling on reducing the association between parents' and offspring's educational achievements. In practice, starting school one year later at age 7 is common in Turkey, so the treatment status of those born in 1986 is ambiguous. Due to this fuzziness in the treatment status of those born in 1986, we drop this cohort from our sample. Therefore, our identifying assumption becomes that two cohorts born before and after 1986 do not exhibit any systematic difference in their educational attainments other than exposure to the compulsory schooling law. If this assumption holds, then assignment to treatment by birth year is as good as random.[8]

In our underlying model, we use a regression discontinuity design by exploiting discontinuity in the birth year and identify the stark increase in the outcome variable with the following specification:

[7] Altindag (2016) investigates the role of son preference on population and he shows that couples in Turkey exhibit son preference through son-biased differential stopping behavior that does not cause a sex ratio imbalance in the population. He finds that the demand for sons leads to lower (higher) ratios of boys to girls in large (small) families. We additionally confirm his finding using Household Labor Force Surveys conducted in 2004–2013, and find that the positive correlation between the number of children and the share of female children is higher for families with both parents having less than secondary education compared to families with at least one parent with secondary education.

[8] Several earlier studies (Erten & Keskin, 2018; Gulesci et al., 2020) using other datasets that the birth month is available assign the treatment status jointly by birth month and year. Due to the lack of data on the birth month, we rely on the year of birth and thereby a slightly cruder identification strategy similar to the studies by Dayioglu et al. (2016) and Aydemir and Kirdar (2017).

$$
\begin{aligned}
& y_i = \beta d_i + f(t_i) + \alpha X_i + \epsilon_i \\
& \forall t_i \in (-h, +h) \\
& d_i = \begin{cases} 1 & t_i > 0 \\ 0 & t_i \leq 0 \end{cases}
\end{aligned} \tag{1}
$$

where y_i is the binary educational attainment variable, d_i is the binary policy variable indicating the treatment status, t_i is the forcing variable of the treatment status, defined as the year of birth centered around the cutoff year 1986, and h is the bandwidth around the cutoff point. We determine the optimal bandwidth by the algorithm defined by Calonico et al. (2017), which is estimated to be 6 years around the cutoff year 1986.

The outcome variables in this study are two binary educational attainment variables: completion of 8-year primary education (new compulsory level) and completion of secondary education (high school). The former and the latter are employed to capture the direct and spillover effects of the reform, respectively. We do not focus on the spillover effects beyond the secondary level because our sample is relatively younger to observe tertiary education outcomes. Also, the earlier studies documented the absence of the impact of the 1997 reform on higher levels of educational attainment (Aydemir & Kirdar, 2017).

We visualize our identification strategy, in the first panel of Fig. 2, by showing the mean completion of 8-year primary education rates separately for men and women born 6 years before and after the cutoff. After fitting two distinct linear trends on both sides, we observe a significant jump at the cutoff for both men and women.

Next, we use older cohorts in AES to conduct a placebo experiment to examine whether there exists another structural increase in the absence of the reform. We shift the reform year by 7 and 14 years and show the graphical representation of a similar analysis for cohorts around the first hypothetical policy year 1979 in the second panel and the second hypothetical policy year 1972 in the third panel. The graphs on placebo experiments indicate no evidence of a jump in educational attainments for the older cohorts. Hence, the stark increase in the educational attainment of those born after 1986 is not likely to be due to an underlying feature but rather the outcome of the reform. These figures present a positive impact of the compulsory schooling policy on educational attainments at the new compulsory (8-year primary) level, which can be interpreted as the direct effect of the reform.

To account for time trends in the outcome variable, we use polynomial $f(t_i)$ of the running variable t_i. Based on the visual inspection of the increasing trend in Fig. 2, we include the centered year of birth variable around the cutoff year t_i and the interaction of it with policy dummy d_i and therefore capture the linear time trend that varies in the left and right hand side of the policy cutoff in our outcome variable. We also experiment with a second-order polynomial time trend, and our results are given in "Appendix."

The validity of our identification strategy requires that the timing of the policy be independent of the outcome variables, i.e., the education outcomes of offspring. The sudden and politically motivated nature of the policy strengthens the exogeneity of the reform to offspring's demand for education. The validity of our identification

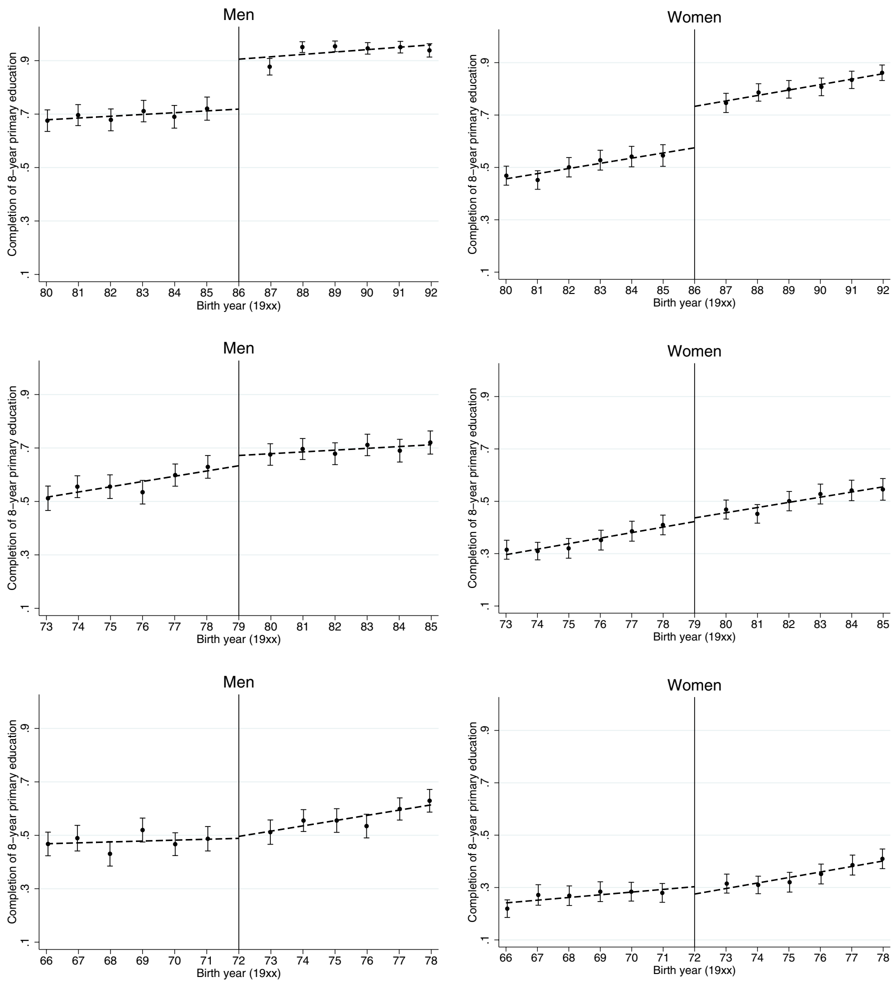

Fig. 2 Mean education completion rates by birth year in treatment and placebo samples. *Source* Adult Education Survey, 2012 and 2016. The points give the mean completion rates for 8-year primary education, across year of birth, and the 95% confidence intervals are given by error bar for each mean level. Local linear polynomials are fitted on both sides of the cutoff

strategy also requires that there be no other policy change at the same time that affects education demand. The other two notable educational programs in Turkey were implemented much later than 1997. The first was a conditional cash transfer program (called Haydi kizlar okula in Turkish) started in 2003 to increase girls' educational attainment in provinces with the lowest rate of women's education. Another NGO-driven conditional cash transfer program targeting girls only (called Baba Beni Okula Gonder) started in 2005. The effect of these policies, implemented about 5 to 7 years after the compulsory schooling policy, are expected to be trivial compared to the national-level policy of ours due to the limited number of program beneficiaries in these programs. In our identification, the effects of these programs and their interaction with the compulsory schooling policy would be captured by the time trend after the discontinuity.

We extend our underlying model to capture the relation of parental education with offspring's education by the regression coefficients of parental education. To that end, we include a parental education variable (father's, mother's or either parent) and its interaction with policy dummy indicating treatment status in Eq. (1) and thereby capture the differential effect of parental education by the compulsory schooling law on offspring's educational attainment by the following linear probability model estimated for male and female offspring separately:

$$y_i = \beta_1 p_i + \beta_2 d_i + \beta_3 p_i d_i + \alpha X_i + \epsilon_i \tag{2}$$

where p_i denotes the parental educational attainment. We include a constant, linear time trend heterogeneous across policy cutoff to the vector of control variables X_i.

Our parameters of interest β_1 and $\beta_1 + \beta_3$ capture the relation between parental education and the offspring's education outcomes before and after the compulsory schooling reform, respectively. The positive coefficient estimate for β_1 indicates the higher probability of completing the given level of education for those born to more educated families. The negative coefficient estimate for the interaction of parental education with policy dummy would indicate compulsory schooling policy is reducing the impact of higher parental education on offspring's educational attainment, and thereby reducing the educational attainment differences between those born to families with high and low levels of parental education.

In order to examine the differential effects of parental education across genders, we next pool the sample of male and female offspring and estimate a model where we include all variables in Eq. (2) and allow policy, the interaction of policy with parental education, in addition to a constant, time trends before and after the policy to vary by gender. In this specification, we will use various composite measures of parental education as mentioned above to examine having one versus two parents with secondary education and also having a tertiary-educated parent versus parents having at most secondary education compared to the omitted category of neither parent having at least secondary education. Hence, our linear probability model becomes

$$y_i = \gamma_0 w_i + \gamma_1 p_i + \gamma_2 w_i p_i + \gamma_3 d_i + \gamma_4 w_i d_i + \gamma_5 p_i d_i + \gamma_6 w_i p_i d_i + \theta_1 X_i + \theta_2 w_i X_i + \nu_i \tag{3}$$

where w_i is a gender indicating variable that takes value 1 if the offspring i is woman and all the other variables are as defined before. Similarly, we include a constant and time trend heterogeneous across policy to the vector of control variables X_i, and $w_i X_i$ denotes the interaction of all control variables with woman dummy that captures the gender varying differences summarized in Table 2.

In Eq. (3), the coefficient estimate of γ_1 captures the dependence of offspring's educational attainment on parental education for men, and $\gamma_1 + \gamma_2$ does the same for women. Hence, a positive coefficient estimate for γ_2 also indicates a higher dependence on parental education for women than men. γ_5 shows the change in the dependence of offspring's education on parental education for men, and $\gamma_5 + \gamma_6$ shows the same change for women. While the negative coefficient estimate for the interaction of parental education with the policy dummy indicates the relaxation of dependence on parental education for men, a similarly negative coefficient estimate for the three-way interaction of parental education with policy and woman dummy would suggest an additional decrease in the dependence on parental education for women compared to men.

Another important question that we can examine by estimating Eq. (3) is the gender gap in educational attainment of those born to low educated families before and after the policy. Recall that the gender gap in educational attainment mostly arises from education differences between men and women whose neither parent has completed secondary education (see Table 1). Here, the coefficient estimate γ_0 captures the gender gap among those born to low educated families and not exposed to policy, while the coefficient estimate for the interaction of woman dummy with policy dummy captures the change in gender gap for those born to low educated families as a result of the compulsory schooling policy. A positive coefficient estimates for the interaction term *woman*×*policy* indicates an improvement in the gender differences to the advantage of women from low educated families due to the policy change.

We calculate standard errors in two ways, following different strands in the literature. First, we cluster errors at the year of birth level following Bertrand et al. (2004), who suggest that clustering should be performed to capture time-trend-related autocorrelation in the sample. However, following this approach leads to few number of clustering problem that might be associated with biased estimation results (Cameron & Miller, 2015). Hence, we also use wild-cluster bootstrapping and report the *p* values by this approach instead of estimated standard errors by following Roodman (2020)[9] to increase the qualitative precision of our results.

Our main sample constructed based on the bandwidth estimation algorithm introduced by Calonico et al. (2017) covers those born between 1980 and 1992 (excluding the 1986 cohort, since their treatment status is ambiguous), which corresponds to 6 cohorts around the policy cutoff. The age window of our primary sample is 20–36. To check the robustness of our results to bandwidth selection, we define alternative samples by expanding the cohort window around the cutoff by two and four cohorts. We present our results for cohorts born in 1976–1996, 1978–1994 (all excluding the cohort 1986) in "Appendix."

To show that our results are not driven by the exclusion of cohort 1986, we include those born in 1986 in our sample and investigate the robustness of our results to this sample specification. Also, we note that some parents may send their children to school even later, especially in the east part of the country. In that case, the cohort born in 1985 might also be partially exposed to extending the mandatory years of schooling. Hence, we experimented with excluding cohort 1985 (one unit left of the policy cutoff) from our sample and additionally cohort 1987 (one unit right of the policy cutoff) to have a symmetric sample on both sides of the cutoff, in other words, expanding the policy cutoff one cohort on both sides.

[9] Roodman (2020) suggest not to use standard errors for inference leading to our choice of presenting *p* values.

Table 3 The effect of compulsory schooling on the intergenerational education transmission

	Men			Women		
Dependent var: completion of 8-year primary education						
Father's education	0.323***			0.494***		
	(0.008)			(0.020)		
	[0.0000]			[0.0000]		
Father's educ. × policy	− 0.249***			− 0.306***		
	(0.015)			(0.025)		
	[0.0000]			[0.0000]		
Mother's education		0.220***			0.413***	
		(0.050)			(0.038)	
		[0.0010]			[0.0000]	
Mother's educ. × policy		− 0.149**			− 0.217***	
		(0.052)			(0.039)	
		[0.0000]			[0.0000]	
Parental education			0.314***			0.458***
			(0.011)			(0.019)
			[0.0250]			[0.0000]
Parental educ. × policy			− 0.239***			− 0.284***
			(0.016)			(0.024)
			[0.0000]			[0.0000]
Observations	5401	5401	5401	7277	7277	7277
Dependent var: completion of secondary education						
Father's education	0.452***			0.548***		
	(0.016)			(0.022)		
	[0.0150]			[0.0130]		
Father's educ. × policy	− 0.089***			− 0.088***		
	(0.020)			(0.028)		
	[0.0030]			[0.0060]		
Mother's education		0.452***			0.590***	
		(0.021)			(0.020)	
		[0.0150]			[0.0180]	
Mother's educ. × policy		− 0.093***			− 0.098***	
		(0.023)			(0.022)	
		[0.0020]			[0.000]	
Parental education			0.450***			0.550***
			(0.014)			(0.023)
			[0.0220]			[0.0140]
Parental educ. × policy			− 0.084***			− 0.075**
			(0.020)			(0.028)

Table 3 (continued)

	Men			Women		
			[0.0030]			[0.0180]
Observations	5401	5401	5401	7277	7277	7277
Birth cohorts	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992

The sample includes the cohorts born between the years given for each column, and the cohort 1986 is excluded. A constant, policy dummy, time trend and its interaction with policy dummy, and survey year fixed effects are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$. The p values by wild-cluster bootstrapping are also given in the square brackets

Results

The Impact of Policy on Intergenerational Educational Persistence

We first estimate Eq. (2) separately for men and women with alternative parental education variables as our key explanatory variable and present our results in Table 3. In Panel A, our dependent variable is equal to 1 if offspring has completed (at least) 8 years of primary education (new compulsory level) and 0 otherwise. In Panel B, our dependent variable is equal to 1 if offspring has completed (at least) secondary education and zero otherwise. In our explanatory variables, we include father's education in the first and fourth columns, mother's education in the second and fifth columns, and finally the parental education (capturing the composite of both parents' education outcomes as it takes value 1 if either parent has completed secondary education or above, and 0 otherwise) in the third and sixth columns to capture the relation between parents' and offspring's education, and also the interaction of these variables with policy dummy to show how the dependence on parental education has been relaxed with the compulsory schooling policy. The significance patterns for the coefficient estimates are very similar for different parental education variables, with closer estimates for father's education and either parent's education as father's education and either parent's education variables are highly correlated. In other words, in our sample of interest, there are few observations (180 (1.4%) out of 12,678) where mother is educated and father is not. In reporting our results, we will focus on the binary parental education variable with the understanding that results are coming mostly from father's education. In further analysis, we will refine this variable according to number of parents having secondary education and level of parental education as to tertiary versus secondary levels.

Our coefficient estimates for the binary *parental education* variable show the relation between the educational attainment of parents and offspring prior to policy change. We observe that before the compulsory schooling law, men and women whose either parent has at least a high school degree are 31 and 46 percentage points more likely to complete at least 8-year primary education, respectively. The compulsory schooling has decreased the effect of parental education on the completion of primary education since the coefficient of the interaction term *parental education*× *policy* is negative and significant for all samples of men and women. The remaining persistence is estimated to be around 8 percentage points (found by adding the coefficient on parental education and the coefficient on the interaction term) (around 76% decrease) for men and 17 percentage points (around 62% decrease) for women. In other words, men (women) whose at least one parent have completed high school or above are more likely to complete primary education by about 8 (17) percentage points even after the change in the compulsory schooling law. Hence, though significantly diminished, parental education continues to play a role in offspring's completing compulsory school.

For the post-compulsory level of education, the relation between the completion of at least secondary education and parent's educational attainment prior to the policy change is higher in magnitude, compared to the new compulsory level for both men and women. Having a parent with at least a high school degree increases the probability that the offspring will also have at least a high school degree by about 45 percentage points for men and 55 percentage points for women before the compulsory schooling reform. We observe spillover effects of the policy on intergenerational educational transmission for both men and women, as indicated by the negative and significant coefficient of the interaction term of parental education with the policy dummy. Our results suggest that the intergenerational persistence after the policy, measured by the sum of the coefficient on parental education and the coefficient on the interaction term, is about 37 percentage points for men and 48 percentage points for women. These results suggest that compulsory schooling significantly decreases intergenerational persistence not only at the new compulsory level of 8 years of primary education but also at post-compulsory level of secondary education though the reduction appears to be higher at the 8-year new compulsory level.

In countries where laws are strictly enforced, it is not surprising that compulsory education positively influences educational outcomes, and eliminates the existing disparities. Based on our regression estimates from three different specifications relying on different parental education variables, we similarly conclude that the reduced cost of education and increased valuation of returns to education reduced the impact of parental education on the completion of both compulsory and post-compulsory level of education and, consequently, increased intergenerational mobility of both men and women in an institutional setting which lacks perfect compliance with the law. However, the differences in schooling outcomes of those born to families from different educational backgrounds are not entirely eliminated both at the compulsory and post compulsory levels.

The Impact of Policy on Gender Gap in Intergenerational Educational Persistence

Here, we examine whether the compulsory schooling law decreases the existing gender differences in intergenerational educational persistence and estimate Eq. (3) by using a pooled sample of men and women. Since this equation controls for interaction terms of all variables with woman (offspring) dummy variable, by including a triple interaction term of binary policy, parental education, and woman variables, we can estimate the differential effect of the policy on the association of parents' education with offspring's educational attainment according to offspring's gender.

Table 4 The effect of compulsory schooling on the gender gap in intergenerational education transmission

Dependent var	Completion of 8-year primary education		Completion of secondary education	
		Coef. for women by (1)		Coef. for women by (3)
	(1)	(2)	(3)	(4)
Woman	− 0.162***		− 0.099***	
	(0.023)		(0.015)	
	[0.0801]		[0.0921]	
Policy	0.248***	0.206***	0.101***	0.058**
	(0.036)	(0.016)	(0.030)	(0.023)
	[0.0000]		[0.0000]	
Woman × policy	− 0.042		− 0.044**	
	(0.032)		(0.018)	
	[0.2633]		[0.0621]	
Parental education	0.322***	0.496***	0.450***	0.550***
	(0.009)	(0.019)	(0.014)	(0.023)
	[0.0220]		[0.0280]	
Parental education × woman	0.174***		0.099***	
	(0.018)		(0.023)	
	[0.0000]		[0.0160]	
Parental education × policy	− 0.246***	− 0.300***	− 0.084***	− 0.075**
	(0.016)	(0.027)	(0.020)	(0.028)
	[0.0000]		[0.0010]	
Parental education × woman × policy	− 0.054**		0.009	
	(0.022)		(0.024)	
	[0.0460]		[0.7117]	
Observations	12,678		12,678	
Birth cohorts	1980–1992		1980–1992	

The sample includes the cohorts born between the years given for each column, and the cohort 1986 is excluded. A constant, time trend and its interaction with policy dummy, survey year fixed effects, and the interaction of all variables with the woman dummy are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$. The p values by wild-cluster bootstrapping are also given in the square brackets

We present our results in Table 4 for the new compulsory (column 1) and post-compulsory level of education (column 3). In the second and fourth columns, we additionally report the composite coefficients for women (found by adding the coefficient estimate of the given variable and the coefficient estimate of its interaction with woman dummy), based on our regression estimates in the first and third columns.

In column (1), where we examine the completion of 8-year primary education, the intergenerational persistence is higher for women prior to the policy change, compared to men, as indicated by the positive and significant coefficient estimates of *parental education×woman*. Prior to the policy change, the impact of a parent's secondary education attainment on his daughter's completion of 8-year primary education is 17 percentage points higher than the impact of a similar parent's secondary education attainment on his son's completion of the same level. Persistence for men has decreased with the policy suggested by the negative and significant coefficient estimates for the interaction of parental education with policy dummy. The decrease in persistence is even higher for women as the differential impact of the reform on the association of women's primary education attainment with parental education (as shown by the coefficient on the triple interaction term, *woman×policy×parental education*) is also estimated to be negative and significant. Hence, the relatively higher gap in attainment of 8-year primary education between women from educated and uneducated parents significantly decreased by 5 percentage points with the extension of mandatory years of schooling.

In our conceptual framework, we argue that the policy decreases the cost of 8 years of schooling. This reduction increases educational outcomes of those with higher price elasticity, who are likely to be those with lower educational attainment by Orazem and King (2007). The policy raises the new compulsory level of educational attainment for low educated parents' sons, as indicated by the positive and significant estimate for policy dummy. Daughters of low educated parents, the most disadvantaged group in our pre-policy sample, experience an increase in their educational outcomes similar to sons of parents with low educational backgrounds, indicated by the imprecise estimate for the interaction of policy with woman dummy. In other words, the policy improved the education outcomes of girls from low education families as much as it increased education levels of boys from low education families, thereby the gender gap for this group at the new compulsory level remained intact. The policy effect is statistically equal to zero for sons of educated families (found by adding the coefficient estimate for policy and the interaction of policy with parental education, given in column (1)). In contrast, the effect of policy on daughters of educated parents is negative (found by adding the composite coefficients of policy, 0.206, and the interaction of parental education for women, − 0.300 given in column (2)). These findings indicate that the policy reduced the educational attainment gap by parental education at the new compulsory level of education for sons by improving the educational attainment of boys from disadvantaged families. For women, the policy reduces disparities in 8-year primary education by parental education, both by raising the education of the daughters born to uneducated families and also by lowering the education of the daughters born to educated families, with a higher in magnitude effect of the former. We propose that these adverse effects on daughters of educated families might be due to eliminating religious vocational schools at the junior high school level and imposing secular education on daughters of conservative families. This is an intriguing explanation that warrants further examination in future research.

In column (3) of Table 4, we examine offspring's completion of secondary education as our outcome variable. Similar to our results at the compulsory level, prior to the reform, we observe a gender gap in intergenerational persistence in secondary education for all samples since the relation with the parents' education is 10 percentage points higher for daughters compared to sons. We find that for men, the policy has decreased intergenerational persistence in secondary education. Women exposed to the compulsory schooling reform have enjoyed a similar decrease in intergenerational persistence as their men peers. However, as suggested by the imprecisely estimated coefficients for the three-way interaction of woman, policy, and parental education variables, the reform did not reduce the gender gap in education transmission across generations at the post-compulsory level. In other words, the reform reduced the impact of parental education on both sexes' completion of secondary education in a similar magnitude.

The spillover effects of the reform on reducing the schooling differences by parental education remain strong for men. We conceptualize the spillover effects of the reform on reducing the impact of parental education on sons' schooling beyond the compulsory level with altered sheepskin effects and the reduced marginal cost of secondary education relative to the mandatory level of education. With their higher labor force participation, these channels especially appear to be effective for men. However, while the gender gap in intergenerational persistence closes at the compulsory level, it remains intact at the post compulsory level. These results may result from more discounted future returns for daughters due to women's lower labor force participation and/or higher opportunity and psychic costs of education for daughters born to low educated parents.

Our findings presented in Table 4 have additional insights on the gender gap for those born to low-educated families, which is shown to be the main driver of the overall gender gap (see Table 1). The coefficient estimate for the woman dummy captures the gender differences in educational attainment for those whose neither parent completed at least secondary education prior to the policy (parental education = 0). We find that these women without an educated parent are 16 percentage points less likely to complete 8 years of schooling than their male peers with similar uneducated parents prior to the policy. This gender gap remains statistically the same, indicated by the negative but statistically insignificant coefficient estimates for the interaction of woman and policy dummies. Similarly, girls without an educated parent are 10 percentage points less likely to complete high school than boys without an educated parent prior to the policy. The policy increased the educational attainment of both sons and daughters in low education families, indicated by positive and significant coefficient estimates for the policy dummy (0.101 in column 3) for men and (0.058 in column 4) for women. Since, the increase in women's high school attainment is less than the increase for men's high school attainment, the gender gap in post compulsory schooling increases for offspring from low educated parents.

Heterogeneity Analysis by Parental Education

In the previous section, we show that the dependence of offspring's education on parental education decreases with the extension of mandatory years of schooling for both men and women, with an additional decrease for women at the compulsory level. Here, we investigate the source of this decrease in dependence which actually means increase in educational mobility by two different refinements of parental education variables.

First, we aim to explore the differential effects of having one parent versus two parents with secondary education, in comparison to neither parent having secondary education. To that end, we define two binary variables, namely *one parent is educated* which is equal to one if one parent has at least secondary education and zero otherwise and *two parents are educated*, which is equal to one if both parents have at least secondary education and zero otherwise and replace the parental education in Eq. (3) with these two variables. Similar to our earlier specifications, we include the interaction of these variables with woman and policy dummies and also the three-way interaction of them with both woman and policy dummies. Table 5 presents our results in columns (3) at the new compulsory level and (7) at the post-compulsory level of education. We also show composite coefficient estimates for women in columns (4) and (8). Column (1) presents our baseline results for easy comparison with composite coefficient estimates for women presented in column (2).

For both men and women, we estimate a lower coefficient for one parent's education and a higher coefficient for two parent's education, compared to our baseline estimates. Parallel to higher dependence on two parent's being educated, the decrease in the dependence on parental education is also higher for those with two educated parents compared to one parent being educated, for both men and women. The differential decreases in the association of offspring's education with parents for women, captured by the three-way interaction terms, are negative and very close to each other for both parental education dummies, indicating the differential decrease comes from eliminating the gap between daughters without an educated parent and the others.

Our results for both men and women where the dependent variable is the completion of the post-compulsory level of schooling, similarly, show that the association between parent's and offspring's education is higher if both parents are educated than if only one parent is educated, both estimated compared to neither parent being educated. The policy reduces the educational attainment gaps between sons without an educated parent and with one or two educated parents. For women, we only find a negative and significant coefficient estimate for the interaction of policy with two parents are educated dummy, but not with one parent is educated. Therefore, for women, the reduction in the high school attainment disparities by parental education comes from the lower gap between daughters with two educated parents and daughters without an educated parent.

Next, we investigate whether the effect of parental education on offspring's educational attainment varies by the level of parental education. We continue with a composite measure of father's and mother's education and define two binary variables that collectively explains our baseline results with the binary parental education variable. The first one called *secondary education* takes value 1 if either parent has at most a high school degree, and the second one, called *tertiary education*, takes value 1 if either parent has a college degree. Similarly, we replace the parental education in Eq. (3) with these two variables and add a full set of interactions of these variables with policy and woman dummies. Table 6 presents our results in column (3) at the new compulsory level and (7) at the post-compulsory level of education. We similarly present composite coefficients for women in columns (4) and (8).

Table 5 The effect of compulsory schooling on the gender gap in intergenerational education transmission: Heterogeneity analysis by the number of educated parent

Dependent var	Completion of 8-year primary educ.				Completion of secondary educ.			
		Coef. for women by (1)		Coef. for women by (3)		Coef. for women by (5)		Coef. for women by (7)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Parental education	0.322***	0.496***			0.450***	0.550***		
	(0.009)	(0.019)			(0.014)	(0.023)		
	[0.0220]				[0.0280]			
Parental education × woman	0.174***				0.099***			
	(0.018)				(0.023)			
	[0.0000]				[0.0160]			
Parental education × policy	− 0.246***	− 0.300***			− 0.084***	− 0.075**		
	(0.016)	(0.027)			(0.020)	(0.028)		
	[0.0000]				[0.0010]			
Parental education × woman × policy	− 0.054**				0.009			
	(0.022)				(0.024)			
	[0.0460]				[0.7117]			
One parent is educated			0.304***	0.459***			0.409***	0.489***
			(0.012)	(0.019)			(0.015)	(0.028)
			[0.0250]				[0.0100]	
Two parents are educated			0.358***	0.563***			0.534***	0.660***
			(0.004)	(0.019)			(0.016)	(0.019)
			[0.0000]				[0.0100]	
One parent is educated × woman			0.154***				0.080**	
			(0.020)				(0.031)	
			[0.0000]				[0.0360]	
Two parents are educated × woman			0.205***				0.125***	
			(0.020)				(0.017)	
			[0.0000]				[0.0220]	

Table 5 (continued)

Dependent var	Completion of 8-year primary educ.				Completion of secondary educ.			
		Coef. for women by (1)		Coef. for women by (3)		Coef. for women by (5)		Coef. for women by (7)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
One parent is educated × policy			− 0.228***	− 0.278***			− 0.077***	− 0.052
			(0.017)	(0.025)			(0.025)	(0.035)
			[0.0000]				[0.0260]	
Two parents are educated × policy			− 0.282***	− 0.333***			− 0.104***	− 0.096***
			(0.017)	(0.025)			(0.020)	(0.023)
			[0.0000]				[0.0000]	
One parent is educ. × woman × policy			− 0.049*				0.024	
			(0.023)				(0.035)	
			[0.0871]				[0.5205]	
Two parents are educ. × woman × policy			− 0.051*				0.009	
			(0.026)				(0.027)	
			[0.0881]				[0.7548]	
Observations	12,678	12,678	12,678	12,678	12,678	12,678	12,678	12,678
Birth cohorts	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992

The sample includes the cohorts born between the years given for each column, and the cohort 1986 is excluded. A constant, policy dummy, time trend and its interaction with policy dummy, survey year fixed effects, and the interaction of all variables with the woman dummy are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$. The p values by wild-cluster bootstrapping are also given in the square brackets

Table 6 The effect of compulsory schooling on the gender gap in intergenerational education transmission: Heterogeneity analysis by the level of parental education

Dependent var	Completion of 8-year primary education				Completion of secondary education			
		Coef. for women by (1)		Coef. for women by (3)		Coef. for women by (5)		Coef. for women by (7)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Parental education	0.322***	0.496***			0.450***	0.550***		
	(0.009)	(0.019)			(0.014)	(0.023)		
	[0.0220]				[0.0280]			
Parental education × woman	0.174***				0.099***			
	(0.018)				(0.023)			
	[0.0000]				[0.0160]			
Parental education × policy	− 0.246***	− 0.300***			− 0.084***	− 0.075**		
	(0.016)	(0.027)			(0.020)	(0.028)		
	[0.0000]				[0.0010]			
Parental education × woman × policy	− 0.054**				0.009			
	(0.022)				(0.024)			
	[0.0460]				[0.7117]			
Secondary education			0.314***	0.459***			0.415***	0.487***
			(0.011)	(0.019)			(0.022)	(0.027)
			[0.0380]				[0.0090]	
Tertiary education			0.334***	0.560***			0.505***	0.657***
			(0.009)	(0.019)			(0.016)	(0.021)
			[0.0240]				[0.0060]	
Secondary education × woman			0.144***				0.072**	
			(0.023)				(0.032)	
			[0.0000]				[0.00851]	
Tertiary education × woman			0.226***				0.152***	
			(0.014)				(0.017)	
			[0.0000]				[0.0310]	

Table 6 (continued)

Dependent var	Completion of 8-year primary education				Completion of secondary education			
		Coef. for women by (1)		Coef. for women by (3)		Coef. for women by (5)		Coef. for women by (7)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Secondary education × policy			− 0.239***	− 0.284***			− 0.084**	− 0.059*
			(0.016)	(0.024)			(0.028)	(0.033)
			[0.0000]				[0.0220]	
Tertiary education × policy			− 0.257***	− 0.330***			− 0.088***	− 0.105***
			(0.018)	(0.025)			(0.020)	(0.025)
			[0.0000]				[0.0010]	
Secondary educ. × woman × policy			− 0.045				0.024	
			(0.028)				(0.033)	
			[0.1672]				[0.5385]	
Tertiary educ. × woman × policy			− 0.073***				− 0.017	
			(0.018)				(0.024)	
			[0.0040]				[0.5035]	
Observations	12,678	12,678	12,678	12,678	12,678	12,678	12,678	12,678
Birth cohorts	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992

The sample includes the cohorts born between the years given for each column, and the cohort 1986 is excluded. A constant, policy dummy, time trend and its interaction with policy dummy, survey year fixed effects, and the interaction of all variables with the woman dummy are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$. The p values by wild-cluster bootstrapping are also given in the square brackets

Men with a high school graduate parent and a tertiary-educated parent are more likely to complete 8 years of education than men without an educated parent. Similarly, women whose either parent completed secondary or tertiary education are more likely to finish 8 years of schooling compared to their peers with uneducated parents. The relative advantage of men whose at least one parent got a high school versus college degree is close to each other, with a 2 percentage points difference. In contrast, women who had a university graduated parent are 11 percentage points more likely to complete 8-year primary education than women who have only high school graduated parent(s). The compulsory schooling policy reduces relative advantages for both men and women from college and high school educated families compared to the ones born to families with lower educational backgrounds. The reduction in gender differences in the dependence on parental education comes from the fall in the gap between daughters from a family with a college graduate parent and those from a family without an educated parent.

At the completion of secondary education, the educational attainment gaps by parental education differ by the level of parental education for both men and women. The pre-policy relative advantage of both daughters and sons is higher for those with a tertiary-educated parent than those with a high school graduated parent, compared to the children without an educated parent. With the reform, the relative advantages for sons diminish in a similar way, resulting in the same ranking of average high school attainment probabilities for men from different educational backgrounds. The gap between daughters born to a family with a high school educated parent and without one slightly decreases as indicated by − 0.059 in column 8, while the fall in the gap between daughters of tertiary-educated parents and the uneducated ones is strongly significant and higher in absolute value as indicated by coefficient − 0.105 in the same column.

Heterogeneity Analysis: Regional Variations

Turkey displays a broad spectrum of economic and demographic differences across its regions. Due to these differences, the effect of the policy on eliminating the existing gender disparities in educational attainment by parental education might vary across regions. In this section, we investigate the possibly heterogeneous impacts of the policy in different areas by restricting our sample of interest to the observations from the 2012 wave of the AES, which includes the region of residence information for all survey participants.

The region of residence information is given at both NUTS-1 and NUTS-2 levels in AES, 2012. Based on NUTS-1 classification, we divide our sample into two, namely West (that includes south and west part of the country) and East (that includes the north and east part of the country)[10]. Figure 3 shows the West versus East classification on a Turkey map. Next, we estimate Eq. (3), first on the pooled sample of West and East (in other words, by restricting the sample to AES, 2012),

[10] According to NUTS-1 classification, TR1 to TR7 regions are determined as West, and TR8 to TRC regions are determined as East.

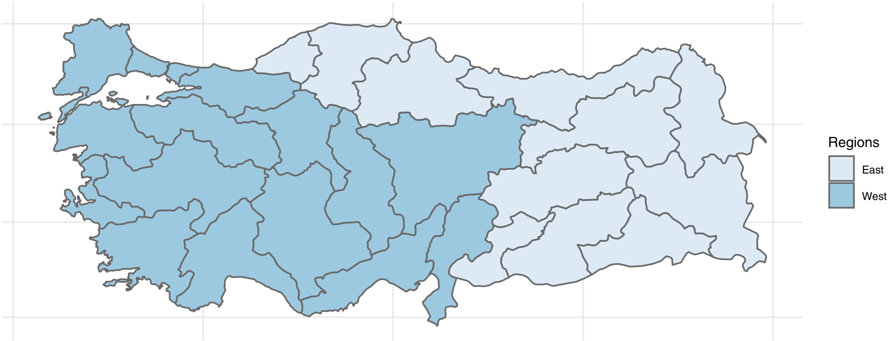

Fig. 3 Geographical classification: west vs. east

and then separately for two samples. Table 7 presents our result for the pooled sample in columns (1) and (4), for West in columns (2) and (5), and finally for East in columns (3) and (6).

The dependent variable in the first three columns is the new compulsory level of schooling. The first column reports our baseline results by restricting the sample to observations from the 2012 wave of AES. Columns (2) and (3) report the coefficient estimates for the same regression equation as (1) where the sample is restricted to those who live in the West and East parts of the country, respectively. We observe that the gender differences are less prominent in the Eastern regions, possibly due to lower educational attainment for both genders (completion rate of 8 years of schooling among cohorts in our sample not exposed to the policy is 45% in the East and 60% in the West). For cohorts not exposed to the policy, we do not find a gender gap among low educated families in the east part of the country, indicated by the imprecisely estimated coefficient estimate for the woman dummy. Similarly, the differential dependence on parental education for women compared to men captured by the interaction of parental education with woman dummy is only weakly significant in the East, while it is higher in magnitude and statistically significant in the West. However, the policy increased the gender disparities in the East: First, by increasing the gender gap among offspring with less educated parents, and also by not reducing the gender differences in the dependence on parental education. Hence, the reduced gender gap in intergenerational persistence in education, found in column (1) and also in our baseline results, is driven mostly by the improvements in the western regions of the country.

In the last three columns, we report our coefficient estimates for the same regression analysis and samples as in the first three columns by changing the dependent variable as the binary secondary education attainment variable. We observe that the gender gap in high school attainment among low educated families only exists in the eastern regions. While the dependence on parental education is higher for men residing in the East, it is lower for women in the East. Due to the spillover effects of extending mandatory years of schooling, the educational attainment gap between those born to low versus high educated families decreases with the policy for both men and women in the West. However, the policy only reduces the effect of parental

Table 7 Regional differences on the effect of compulsory schooling on the gender gap in intergenerational education transmission

Dependent var:	Completion of 8-year primary education			Completion of secondary education		
	All	West	East	All	West	East
	(1)	(2)	(3)	(4)	(5)	(6)
Woman	− 0.141***	− 0.150***	− 0.130	− 0.054**	− 0.040	− 0.211**
	(0.022)	(0.031)	(0.087)	(0.019)	(0.028)	(0.078)
	[0.0230]	[0.0541]	[0.3153]	[0.0180]	[0.1862]	[0.0470]
Policy	0.272***	0.268***	0.287***	0.073**	0.085**	0.031
	(0.037)	(0.033)	(0.067)	(0.027)	(0.034)	(0.041)
	[0.0000]	[0.0000]	[0.0350]	[0.0460]	[0.0120]	[0.6266]
Woman × policy	− 0.046	0.010	− 0.194**	− 0.035	− 0.029	− 0.029
	(0.030)	(0.034)	(0.076)	(0.029)	(0.036)	(0.057)
	[0.1992]	[0.8048]	[0.0060]	[0.3674]	[0.4945]	[0.6316]
Parental education	0.314***	0.298***	0.386***	0.450***	0.428***	0.550***
	(0.023)	(0.024)	(0.041)	(0.018)	(0.018)	(0.060)
	[0.0000]	[0.0000]	[0.0010]	[0.0030]	[0.0070]	[0.0090]
Parental education × woman	0.172***	0.183***	0.142*	0.084**	0.111***	− 0.028
	(0.034)	(0.037)	(0.073)	(0.033)	(0.029)	(0.089)
	[0.0100]	[0.0050]	[0.1502]	[0.0631]	[0.0430]	[0.7768]
Parental education × policy	− 0.281***	− 0.268***	− 0.335***	− 0.131***	− 0.113***	− 0.209**
	(0.027)	(0.027)	(0.051)	(0.020)	(0.024)	(0.067)
	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0000]	[0.0150]
Parental education × policy × woman	− 0.065*	− 0.102**	0.018	0.043	− 0.002	0.193*
	(0.032)	(0.037)	(0.078)	(0.036)	(0.036)	(0.098)
	[0.0841]	[0.0100]	[0.8168]	[0.2953]	[0.9650]	[0.1051]
Observations	7756	5883	1873	7756	5883	1873
Birth cohorts	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992

The sample includes the cohorts born between the years given for each column, and the cohort 1986 is excluded. A constant, policy dummy, time trend and its interaction with policy dummy, region fixed effects, and the interaction of all variables with the woman dummy are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$. The p values by wild-cluster bootstrapping are also given in the square brackets

education for men in the East, but not for women (found by adding the coefficient estimate for the three-way interaction term to the estimate for the interaction of parental education and policy), thereby leading to a gender gap in intergenerational mobility. Hence, the spillover effects of the policy on reducing the existing disparities by parental education are ineffective for women in the East.

In the conceptual framework, we emphasize the lower perceived returns and higher perceived cost of schooling for daughters than sons, especially for those with less educated parents lead to a gender gap in intergenerational mobility of education.

We further point out that the policy is likely to reduce these existing disparities by reducing the cost of education and altering the returns to education, and therefore, providing higher benefits for the education than its cost for exposed cohorts, with possibly spillover effects at the post-compulsory level. Hence, our findings on the different effects of the policy in the western and eastern parts of the country can be argued to be driven by the several important differences in the cost and benefits of education that varies across regions.

First, the perceived cost of education might differ due to varying income levels, enforcement, and associated penalties between the west and east parts of the country. Lower household income and larger household size lead to low income per capita in the east than in the west (according to authors' calculations using HLFS 2012, real household income per capita is about 35% lower in the east part of the country than the rest). The policy lowers the monetary cost of education for all with the public investment to support the quick implementation of the policy across the country; however, families still have to buy uniforms and school supplies. In addition, students had to buy all their textbooks until the 2003–2004 school year. Hence families living in the East may face more binding financial constraints in meeting school expenses compared to others, and so poorer economic conditions might explain the higher dependence on parental education in the East, even after the compulsory schooling policy. Also, enforcement of the policy through monetary (and other) penalties that alter the cost of education might vary in the eastern and western regions. The compliance of students with the policy is tracked by the teachers and local administrative bodies. Hence, the attitudes of the local population might contribute to the level of enforcement in the East, leading to a lower cost of not attending in the Eastern provinces.

The varying perceived cost of education by gender might also alter the effect of policy in the eastern regions due to higher opportunity and psychic costs. The opportunity cost driven by the foregone home production of daughters with their attendance in education is possibly more heightened in the East due to the larger household sizes (21% larger households in the East compared to the rest). Besides, the psychic cost of sending daughters to school is expected to be higher with the region's more prominent traditional gender views. To show the varying traditional gender views across regions, we draw on a new survey conducted by Konda, Research and Consulting company, in 2015 (KONDA, 2020), and define a gender view index to measure the conservativity in the East in comparison to West (The details of the construction of the gender view index and its mean levels across NUTS-1 regions are given in "Regional variation of traditional gender views in Turkey" section of Appendix). Our findings show a more conservative population in the eastern regions and confirm the higher psychic cost in these areas. Even though the policy reduces the monetary cost of education, if the opportunity and psychic cost remain to be a large component of the cost of education for daughters, then we might argue these non-monetary costs might contribute to the policy's failure to reduce the gender differences in the eastern regions.

The policy is likely to be less effective in reducing women's educational disadvantages in the East due to lower perceived returns to education for daughters in the region. Individuals living in rural areas constitute a larger part of the population in the East (41% in the East and 23% in the West), and female's paid employment in the region remains to be low compared to the rest of the country (%8 in the East, and %15 in the West, by HLFS, 2012) despite the larger rural population. Parents residing in rural areas may put less value on their offspring's educational attainment, as individuals from rural areas are more likely to work in the agricultural sector whose returns to education are also lower. With the lower paid employment rates of women, parents in the East might discount their daughters' future income more since the returns to education play a more important role in the paid employment than other types of employment, and so the policy might be less effective in eliminating existing gender differences.

Robustness Checks

Recall that our main sample constructed based on the bandwidth estimation algorithm introduced by Calonico et al. (2017) that includes six cohorts around the policy cutoff (excluding the 1986 cohort, since their treatment status is ambiguous). To check the robustness of our results to bandwidth selection, we define alternative samples by expanding the cohort window around the cutoff by two and four cohorts. We present our results for cohorts born in 1976–1996, 1978–1994 (all excluding the cohort 1986), in addition to the results on our main sample, in the first panel of Table 9 in Appendix. To further check that our results are not due to an underlying trend, we include the second-order polynomial of time trend instead of the linear one. We similarly allow for heterogeneous time trends across policy cutoff by interacting the trend variables with the policy dummy. In the second panel of Table 9, we report the coefficient estimates for the same samples as in the first panel with second-order time trend controls. The significance patterns are the same as our baseline results, confirming that our results are not driven by the bandwidth selection or the choice of trend fitted on both sides of the policy cutoff.

As explained in the identification section, we exclude those born in 1986 from our sample due to the uncertainty in their exposure to the reform. Here, to check the validity of our identification, we include those born in 1986 in our sample and investigate the robustness of our results to this sample specification. We also point out that some parents may send their children to school a year later than their peers, especially in the east part of the country. In that case the cohort born in 1985 might also be partially exposed to extending the mandatory years of schooling. Hence, we experimented with excluding cohort 1985 (one unit left of the policy cutoff) from our sample and additionally cohort 1987 (one unit right of the policy cutoff) to have a symmetric sample on both sides of the cutoff, in other words, expanding the policy cutoff one cohort on both sides. Table 10 presents our results on these samples with different cohort windows and shows that the significance patterns are the same under the alternative sample specifications, confirming our identification strategy.

The data for this study are a pooled sample of AES 2012 and 2016, but the two waves slightly differ in their sampling and sizes. To confirm that our results are not driven by a pattern only observed in one wave, we re-estimate our main equation by restricting the sample to AES 2012 and 2016. Table 11 presents our results for the pooled sample in the first and fourth columns, for the sample restricted to the 2012 wave in the second and fifth columns, and for the 2016 wave in the third and sixth columns. The point estimates are quite similar in the first three columns, where the dependent variable is the completion of 8-year primary education. The significance of the three-way interaction of parental education with policy and woman dummies vanishes for both the 2012 and 2016 waves, possibly due to the smaller sample size. However, the sign and magnitude of the coefficient estimate support the main findings of the paper. Similarly, the point estimates and the significance patterns are relatively close in the last three columns, where we examine the impact of policy at the post-compulsory level of education, except for the insignificant estimate for the interaction of parental education with the policy dummy in AES 2016, though the point estimate has the same sign as in the fourth and fifth columns, and therefore, confirm our results.

Conclusion

People who are well educated are likely to have children who are also well educated. Similarly, parents who have low education levels are likely to have children with low levels of education. This intergenerational transmission of education holds to varying degrees in most countries with Latin America displaying the highest, and the Nordic countries lowest (Hertz et al., 2008). Large gender differences in intergenerational education transmission also reported in several developing countries (Emran & Shilpi, 2015; Glick & Sahn, 2000). Daughters exhibit more intergenerational educational persistence than sons.

Turkey, an emerging economy, exhibits high intergenerational education persistence (Aydemir & Yazici, 2019) and a gender gap in intergenerational persistence to the detriment of women (Demirel & Okten, 2020; Tansel, 2015). In this study, we examine the impact of a change in compulsory schooling from 5 to 8 years in 1997 which affected some cohorts of offspring and not others, on intergenerational educational persistence and gender differences in intergenerational educational transmission in Turkey.

Our work contributes to the literature in two ways. First, we contribute to the small literature examining the impact of compulsory schooling policies on mitigating the inheritance of educational inequality. Existing studies are from advanced countries with relatively low levels of intergenerational persistence (Meghir & Palme, 2005; Aakvik et al., 2010; Betthäuser, 2017). We provide the first empirical evidence from an emerging country with high intergenerational educational persistence, even at the compulsory level. Second, our work, to our knowledge, is the first to examine the impact of compulsory schooling on the gender gap in the intergenerational transmission of education. Compulsory schooling laws are a common policy tool to achieve greater participation in education, particularly from marginalized groups. Our paper examines how an extension in compulsory schooling can mitigate the gender gap in the association of parents' education with offspring's education that exists to the detriment of women and thereby interrupt the persistence of marginalization in education outcomes across generations.

Prior to the reform, there is a gender gap in the association of parents' educational attainment with their offspring's. Daughters exhibit more intergenerational persistence than sons. We show that compulsory schooling law reduced the impact of parental education on completion of new compulsory schooling (8-year primary education) from 31 to 7 percentage points for men and from 46 to 17 percentage points for women. At the post-compulsory schooling, the association between fathers' and offspring' education has decreased by about 10 percentage points for both sexes. The gender gap in intergenerational persistence has decreased by 5 percentage points in the completion of new compulsory schooling but remains unchanged at the post-compulsory schooling level after the reform.

Our findings indicate that the effect of compulsory schooling on increasing intergenerational educational mobility is strong as the persistence has decreased substantially for both men and women at the new compulsory and post-compulsory level of education. The gender gap in the differential dependence on parental education for daughters also reduces with the policy at the new compulsory level, showing the effectiveness of the policy change on bringing about equality. However, the reform is not very effective in bringing gender equality in the transmission of educational outcomes across generations in the areas with poorer economic conditions, larger rural population, and more traditional gender views in reducing the gender gap in educational mobility, even at the compulsory level of education. Moreover, the positive spillover effect of the reform on the gender differences beyond the compulsory level is limited. Hence increasing mandatory years of schooling is an effective policy tool to eliminate the gender gap in the inheritance of education inequality as long as policymakers recognize that its effect will be limited to what is compulsory and even than its effect may be not be as strong in disadvantaged regions. Our results concur with Duflo (2012) in that continuous policy commitment is needed to bring about equality between men and women.

Appendix

Regional Variation of Traditional Gender Views in Turkey

To study the relation of the traditional gender views with the effect of compulsory schooling reform on providing gender equality in intergenerational persistence, we draw on a new survey conducted by Konda, Research and Consulting company, in 2015 (KONDA, 2020). The sample of Konda survey was collected according to the address-based population system with stratified sampling according to the 2011 General Election Results. The survey is representative of adults (18 years old or older), similar to our main dataset in this study, at the NUTS-1 regional level. There are 12 regions at the NUTS-1 level in Turkey. The primary purpose of the Konda survey was to gather views and opinions on gender roles and domestic violence.

The survey has several questions on gender views, which we use to examine which regions have more equal gender views. Survey respondents are asked whether they strongly disagree, disagree, agree, or strongly agree with the following statements: S1—The main responsibility of a woman is to raise children and run a household; S2—Women entering the labor force leads to unemployment among men; S3—Women cannot be good managers by nature; S4—Women are delicate, it is not appropriate for them to work in men's jobs; S5—Women should be careful about their outfit in the workplace. We construct five gender view dummy variables for each statement that takes a value of 1 if the respondent agrees or strongly agrees with the given statement and 0 otherwise. We, next, define a gender view index variable by the first principal component of all five dummies constructed by the questions asked in the Konda survey on gender views.

Table 8 summarizes the mean levels for five gender view dummies and gender view index in 12 NUTS-1 regions. The higher means in the east part of the country than the west indicates that more traditional gender views are held by people residing in the east (Tables 9, 10, 11).

Table 8 The gender view index at the NUTS-1 level

NUTS-1 Regions	Gender view dummy 1	Gender view dummy 2	Gender view dummy 3	Gender view dummy 4	Gender view dummy 5	Gender view index	Above median gender view index
TR1-Istanbul	0.579	0.205	0.252	0.649	0.620	1.023	No
TR2-West Marmara	0.528	0.132	0.153	0.625	0.681	0.938	No
TR3-Aegean	0.661	0.109	0.281	0.738	0.834	1.162	No
TR4-East Marmara	0.546	0.128	0.133	0.520	0.845	0.951	No
TR5-West Anatolia	0.607	0.171	0.291	0.649	0.769	1.104	No
TR6-Mediterranean	0.734	0.357	0.305	0.721	0.814	1.302	Yes
TR7-Central Anatolia	0.802	0.286	0.429	0.808	0.849	1.419	Yes
TR8-West Black Sea	0.706	0.398	0.365	0.733	0.795	1.329	Yes
TR9-East Black Sea	0.674	0.302	0.209	0.791	0.860	1.257	No
TRA-Northeast Anatolia	1.000	0.667	0.806	0.944	1.000	2.002	Yes
TRB-Centraleast Anatolia	0.802	0.304	0.337	0.824	0.911	1.424	Yes
TRC-Southeast Anatolia	0.718	0.302	0.373	0.814	0.940	1.386	Yes

Authors' calculations using Konda Survey on gender views and domestic violence, 2015

Table 9 Robustness checks: alternative bandwidths and time trend controls

Dependent var	Completion of 8-year primary education			Completion of secondary education		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: controls for linear time trend						
Woman	− 0.162***	− 0.173***	− 0.194***	− 0.099***	− 0.093***	− 0.106***
	(0.023)	(0.016)	(0.020)	(0.015)	(0.010)	(0.016)
Policy	0.248***	0.244***	0.222***	0.101***	0.162***	0.119***
	(0.036)	(0.028)	(0.029)	(0.030)	(0.044)	(0.038)
Woman × policy	− 0.042	− 0.032	− 0.002	− 0.044**	− 0.099***	− 0.064**
	(0.032)	(0.024)	(0.025)	(0.018)	(0.029)	(0.028)
Parental education	0.322***	0.331***	0.349***	0.450***	0.450***	0.457***
	(0.009)	(0.009)	(0.016)	(0.014)	(0.011)	(0.010)
Parental education × woman	0.174***	0.182***	0.177***	0.099***	0.115***	0.120***
	(0.018)	(0.016)	(0.015)	(0.023)	(0.021)	(0.018)
Parental education × policy	− 0.246***	− 0.267***	− 0.289***	− 0.084***	− 0.095***	− 0.109***
	(0.016)	(0.015)	(0.019)	(0.020)	(0.025)	(0.024)
Parental education × woman × policy	− 0.054**	− 0.062***	− 0.060***	0.009	− 0.009	− 0.019
	(0.022)	(0.019)	(0.017)	(0.024)	(0.027)	(0.026)
Observations	12,678	17,012	19,948	12,678	17,012	19,948
Panel B: controls for second-order polynomial of time trend						
Woman	− 0.182***	− 0.156***	− 0.134***	− 0.073***	− 0.080***	− 0.062***
	(0.031)	(0.037)	(0.035)	(0.015)	(0.016)	(0.019)
Policy	0.155***	0.236***	0.274***	0.068*	0.044*	0.166**
	(0.031)	(0.044)	(0.046)	(0.037)	(0.024)	(0.058)
Woman × policy	0.044	− 0.027	− 0.074	− 0.088**	− 0.015	− 0.104**
	(0.032)	(0.043)	(0.047)	(0.038)	(0.028)	(0.044)
Parental education	0.322***	0.331***	0.350***	0.450***	0.450***	0.458***

Table 9 (continued)

Dependent var	Completion of 8-year primary education			Completion of secondary education		
	(1)	(2)	(3)	(4)	(5)	(6)
	(0.009)	(0.009)	(0.016)	(0.014)	(0.011)	(0.010)
Parental education × woman	0.173***	0.182***	0.177***	0.100***	0.115***	0.120***
	(0.018)	(0.016)	(0.015)	(0.023)	(0.021)	(0.017)
Parental education × policy	− 0.246***	− 0.267***	− 0.290***	− 0.084***	− 0.095***	− 0.109***
	(0.016)	(0.016)	(0.020)	(0.020)	(0.025)	(0.025)
Parental education × woman × policy	− 0.055**	− 0.062***	− 0.060***	0.009	− 0.009	− 0.019
	(0.023)	(0.018)	(0.017)	(0.024)	(0.027)	(0.026)
Observations	12,678	17,012	19,948	12,678	17,012	19,948
Birth cohorts	1980-1992	1978–1994	1976–1996	1980-1992	1978–1994	1976–1996

The sample includes the cohorts born between the years given for each column, and the cohort 1986 is excluded. A constant, policy dummy, time trend and its interaction with policy dummy, and region fixed effects are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$

Table 10 Robustness check with cohort 1986 is included in the sample

Dependent var	Completion of 8-year primary education			Completion of secondary education		
	(1)	(2)	(3)	(4)	(5)	(6)
Woman	− 0.162***	− 0.212***	− 0.135***	− 0.099***	− 0.129***	− 0.103***
	(0.023)	(0.030)	(0.026)	(0.015)	(0.023)	(0.022)
Policy	0.248***	0.173**	0.322***	0.101***	0.072*	0.132***
	(0.036)	(0.057)	(0.013)	(0.030)	(0.034)	(0.015)
Woman × policy	− 0.042	0.011	− 0.110***	− 0.044**	− 0.011	− 0.053*
	(0.032)	(0.040)	(0.025)	(0.018)	(0.026)	(0.026)
Parental education	0.322***	0.295***	0.322***	0.450***	0.436***	0.443***
	(0.009)	(0.025)	(0.010)	(0.014)	(0.018)	(0.015)
Parental education × woman	0.174***	0.172***	0.185***	0.099***	0.101***	0.117***
	(0.018)	(0.017)	(0.017)	(0.023)	(0.019)	(0.019)
Parental education × policy	− 0.246***	− 0.219***	− 0.260***	− 0.084***	− 0.070***	− 0.088***
	(0.016)	(0.028)	(0.011)	(0.020)	(0.023)	(0.018)
Parental education × woman × policy	− 0.054**	− 0.052**	− 0.062**	0.009	0.008	− 0.003
	(0.022)	(0.021)	(0.023)	(0.024)	(0.021)	(0.022)
Observations	12,678	13,807	10,726	12,678	13,807	10,726
Sample						
Cohort 1985	Yes	Yes	No	Yes	Yes	No
Cohort 1986	No	Yes	No	No	Yes	No
Cohort 1987	Yes	Yes	No	Yes	Yes	No

The sample includes the cohorts born in 1980-1984 and 1988–1992 in addition to the cohorts indicated in the table for each column. A constant, time trend and its interaction with policy dummy, region fixed effects, and the interaction of all variables with the woman dummy are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$

Table 11 Robustness check with sample restricted to 2012 and 2016 waves

Dependent var	Completion of 8-year primary education			Completion of secondary education		
	AES 2012–16	AES 2012	AES 2016	AES 2012–16	AES 2012	AES 2016
	(1)	(2)	(3)	(4)	(5)	(6)
Woman	− 0.162***	− 0.151***	− 0.138***	− 0.099***	− 0.075***	− 0.092*
	(0.023)	(0.017)	(0.040)	(0.015)	(0.016)	(0.050)
Policy	0.248***	0.276***	0.201***	0.101***	0.080**	0.127**
	(0.036)	(0.036)	(0.051)	(0.030)	(0.027)	(0.044)
Woman × policy	− 0.042	− 0.051*	− 0.032	− 0.044**	− 0.043	− 0.034
	(0.032)	(0.025)	(0.051)	(0.018)	(0.028)	(0.062)
Parental education	0.322***	0.329***	0.311***	0.450***	0.464***	0.429***
	(0.009)	(0.019)	(0.017)	(0.014)	(0.018)	(0.012)
Parental education × woman	0.174***	0.184***	0.155***	0.099***	0.096***	0.109***
	(0.018)	(0.029)	(0.028)	(0.023)	(0.028)	(0.022)
Parental education × policy	− 0.246***	− 0.268***	− 0.216***	− 0.084***	− 0.120***	− 0.039
	(0.016)	(0.024)	(0.024)	(0.020)	(0.021)	(0.028)
Parental education × policy × woman	− 0.054**	− 0.053	− 0.046	0.009	0.050	− 0.047
	(0.022)	(0.030)	(0.038)	(0.024)	(0.034)	(0.032)
Observations	12,678	7756	4922	12,678	7756	4922
Birth cohorts	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992	1980–1992

The sample includes the cohorts born between the years given for each column, and the cohort 1986 is excluded. A constant, policy dummy, time trend and its interaction with policy dummy, and the interaction of all variables with the woman dummy are included in control variables. The standard errors clustered at birth year level are given in round parentheses with the patterns of p values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$

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Declarations

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