

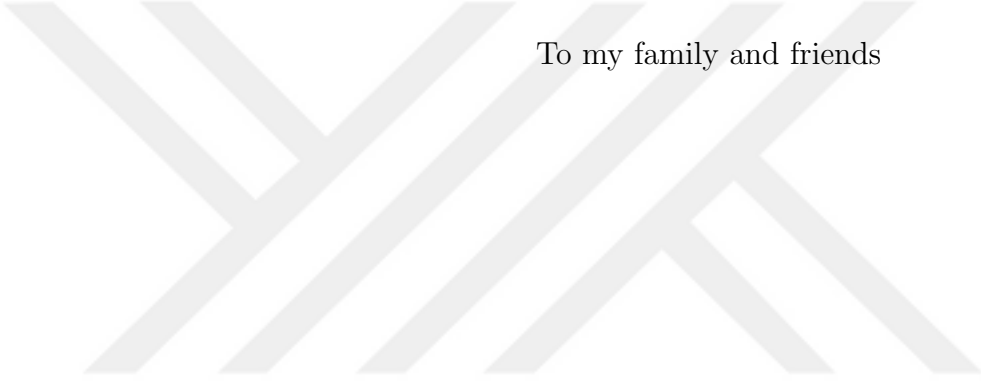
THREE ESSAYS ON THE EFFECTS OF EDUCATION
POLICIES ON EDUCATION AND LABOR MARKET
OUTCOMES IN TURKEY



A Ph.D. Dissertation

by
MERVE DEMİREL

Department of Economics
İhsan Doğramacı Bilkent University
Ankara
September 2022



To my family and friends

THREE ESSAYS ON THE EFFECTS OF EDUCATION
POLICIES ON EDUCATION AND LABOR MARKET
OUTCOMES IN TURKEY



The Graduate School of Economics and Social Sciences
of
İhsan Doğramacı Bilkent University

by

MERVE DEMİREL

In Partial Fulfillment of the Requirements for the Degree of
DOCTOR OF PHILOSOPHY IN ECONOMICS

THE DEPARTMENT OF ECONOMICS
İHSAN DOĞRAMACI BİLKENT UNIVERSITY
ANKARA

September 2022

THREE ESSAYS ON THE EFFECTS OF EDUCATION POLICIES ON EDUCATION AND LABOR MARKET OUTCOMES IN TURKEY

by Merve Demirel

I certify that I have read this thesis and have found that it is fully adequate, in scope and in quality, as a thesis for the degree of Doctor of Philosophy in Economics.

Çağla Ökten
Supervisor

I certify that I have read this thesis and have found that it is fully adequate, in scope and in quality, as a thesis for the degree of Doctor of Philosophy in Economics.

Şaziye Pelin Akyol
Examining Committee Member

I certify that I have read this thesis and have found that it is fully adequate, in scope and in quality, as a thesis for the degree of Doctor of Philosophy in Economics.

Nur Asena Caner
Examining Committee Member

I certify that I have read this thesis and have found that it is fully adequate, in scope and in quality, as a thesis for the degree of Doctor of Philosophy in Economics.

Emin Karagözoğlu
Examining Committee Member

I certify that I have read this thesis and have found that it is fully adequate, in scope and in quality, as a thesis for the degree of Doctor of Philosophy in Economics.

Belgi Turan
Examining Committee Member

Approval of the Graduate School of Economics and Social Sciences

Refet Soykan Gürkaynak
Director

ABSTRACT

THREE ESSAYS ON THE EFFECTS OF EDUCATION POLICIES ON EDUCATION AND LABOR MARKET OUTCOMES IN TURKEY

Demirel, Merve

Ph.D., Department of Economics

Supervisor: Assoc. Prof. Dr. Çağla Ökten

September 2022

This thesis consists of three essays on the effects of two education policies on education and labor market outcomes in Turkey. The first essay examines the causal effects of the drastic expansion in Turkish higher education, which occurred in 2006-2008 with the establishment of new universities and increasing the educational opportunities both in the new and established universities, on the educational attainment disadvantage of women in higher education. The expansion increases the attainment rates of both genders but does not significantly reduce the gender gap after controlling for the time trends. Studying the mechanisms shows that the expansion in social sciences, which constitutes more than half of the additional slots, benefited men and women evenly; but the expansion in engineering, about 25% of additional slots, benefited men more.

The second essay investigates the causal impact of the same expansion policy in higher education on college wage premium in Turkey. The expansion increases the share of college graduates significantly and decreases the college wage premium for young men exposed to the policy. The expansion has lower but significant adverse effects on older age groups, whose human capital accumulation is not affected by the expansion policy. This result provides causal evidence that the increase in the supply of skilled workers reduces the wage advantages enjoyed by these workers.

The third essay analyzes the impact of extending mandatory years of schooling from 5 to 8 years in 1997 on the gender gap in intergenerational educational persistence. Prior to the reform, there was a gender gap in the association of parents' educational attainment with their offspring's. Daughters' educational attainment is more dependent on their parents' educational background. The education reform reduces the impact of parental education on the 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 decreases in the completion of new compulsory schooling level but remains unchanged at the post-compulsory schooling level after the reform.

Keywords: education reform, difference-in-differences, gender equality, intergenerational education transmission, compulsory schooling, Turkey.

ÖZET

TÜRKİYE’DE EĞİTİM POLİTİKALARININ EĞİTİM VE İŞGÜCÜ PİYASASI ÇIKTILARI ÜZERİNDEKİ ETKİLERİ ÜZERİNE ÜÇ MAKALE

Demirel, Merve

Doktora, İktisat Bölümü

Tez Yöneticisi: Doç. Dr. Çağla Ökten

Eylül 2022

Bu tez, Türkiye’de iki farklı eğitim politikasının eğitim ve işgücü piyasası çıktıları üzerindeki etkileri üzerine üç makaleden oluşmaktadır. İlk makale, Türkiye’de yüksek öğrenimde, 2006-2008 yılları arasında yeni üniversitelerin açılması ve hem yeni hem de eski üniversitelerde olanakların arttırılmasıyla gerçekleşmiş olan genişleme politikasının kadınların yüksek öğrenime katılımlarındaki dezavantajı üzerindeki nedensel etkilerini incelemektedir. Genişleme politikası, her iki cinsiyetin eğitime katılımlarını benzer şekilde arttırarak cinsiyet farkını azaltmaz. Mekanizmaları inceleyerek, kontenjan artışının yaklaşık yarısını oluşturan sosyal bilimlerdeki artışların, kadın ve erkeklere eşit fayda sağladığı; ancak yeni kontenjanların %25’ini oluşturan mühendislik alanlarındaki artışın erkeklerin eğitimini daha fazla arttırdığı gösterilir.

İkinci makale, yüksek öğretimde genişleme politikasının Türkiye’de üniversite mezunlarının lise mezunlarına kıyasla elde ettiği ücret avantajları üzerindeki etkisini araştırır. Genişleme politikası, üniversite mezunlarının payını önemli ölçüde artırırken politikadan etkilenen genç erkekler için üniversite ücret avantajlarını azaltmaktadır. Genişleme, beşeri sermaye birikimi genişleme politikasından etkilenmeyen daha büyük yaş grupları üzerinde daha düşük ancak belirgin olumsuz etkilere sahiptir. Bu sonuç, vasıflı işçi arzındaki artışın bu işçilerin yararlandığı ücret avantajlarını azalttığına dair nedensel kanıtlar sunmaktadır.

Üçüncü makale, zorunlu eğitim süresinin 1997’de 5 yıldan 8 yıla çıkarılmasının nesiller arası eğitim aktarımındaki cinsiyet farkları üzerindeki etkisini analiz ediyor. Reformdan önce, ebeveynler ile çocuklarının eğitim çıktıları arasındaki ilişkide cinsiyet farklılıkları vardı. Kızların eğitimi erkeklere kıyasla ebeveynlerinin eğitime daha fazla bağlı idi. Eğitim reformu, hem erkek hem de kız çocukları için yeni zorunlu eğitimin (8 yıl) ve zorunlu sonrası eğitimin (lise) tamamlanmasında ebeveyn eğitiminin etkisini azaltır. Nesiller arası eğitim aktarımındaki cinsiyet farklılıkları, yeni zorunlu eğitim seviyesinde azalırken, zorunlu sonrası eğitimde belirgin bir değişiklik görülmez.

Anahtar Kelimeler: eğitim reformu, farkların farkı, cinsiyet eşitliği, nesiller arası eğitim aktarımı, zorunlu eğitim, Türkiye.

ACKNOWLEDGEMENTS

This dissertation would never have been possible without all the support, inspiration, and encouragement of many great people who accompanied me on my walk through this long and hard journey. I would like to thank them all.

First and foremost, I am extremely grateful to my supervisor, Çağla Ökten, for her excellent guidance, encouragement, patience, and immense knowledge. I will eternally be indebted to her support throughout my time in Bilkent and view her as my lifetime role model.

I am also thankful to Asena Caner for her motivation, guidance, and support throughout my Ph.D. studies. She has always been there for me whenever I need her guidance. Moreover, I am also grateful to Pelin Akyol, for being a member of my dissertation monitoring committee, and for her invaluable comments, contributions, and constructive suggestions.

I would like to express my deepest gratitude to Emin Karagözoğlu for his excellent guidance during my time in Bilkent, and also for being my thesis defense committee member. It is one of my greatest chances in Bilkent the time he devoted to me whenever I lose my way or motivation to succeed. I would like to extend my thanks to my final thesis defense committee member, Belgi Turan, for her valuable comments and suggestions.

I am grateful to Refet Gürkaynak for his trust and for providing me with financial support throughout my graduate degree. In addition, I am grateful to Banu Pakel for her invaluable comments on my studies and her contributions to my academic career. I owe special thanks to all my professors in the Department of Economics for always making me feel like a part of this amazing family. I need to mention

Özlem Demiral and Nilgün Çorapçıoğlu for making their support available whenever I need it.

I am also thankful to the Turkish Statistical Institute (TurkStat) for granting permission to use Household Labor Force and Adult Education Surveys. Also, the financial support of TUBITAK during my studies is gratefully acknowledged.

I am beholden to my dearest friend, companion, and life saver İdil Ayberk for her support in coping with many obstacles we face throughout this journey that we walk together. Her never-ending trust in me has been and will continue to be the reason why I keep walking through the ups and downs. Also, I am grateful to Ceren İnci for her everlasting friendship over the years. In addition, I would like to thank my friends from graduate studies, Asu İşbilen, Betül Akar, Zeynep Yılmaz, and Sıla Erikçi, for our delightful conversations about my research topics and their support during my Ph.D. studies. Finally, I am grateful to Ömer Faruk Akbal for his final support in writing my thesis.

The deepest gratitude should be reserved for the people who have shared the struggles and achievements with me. I would like to thank my mother, Neşe Demirel, my father, Metin Demirel, and my sister/best friend, Melda Demirel, for their unlimited support, encouragement, patience, and love throughout my life. Without my family's understanding and patience, I would not have completed my Ph.D. thesis. Finally, and most importantly, I am deeply grateful to my beloved husband Erdem for always believing in me and supporting me in all my pursuits. He is the person who, more than anyone else, has shared with me all ups and downs of this long journey, and the one with whom I want to share all future ones.

TABLE OF CONTENTS

ABSTRACT	iii
ÖZET	v
ACKNOWLEDGEMENTS	vii
LIST OF TABLES	xiv
LIST OF FIGURES	xvii
1 INTRODUCTION	1
2 ATTAINMENT AND GENDER EQUALITY IN HIGHER EDUCATION: EVIDENCE FROM A LARGE SCALE EXPANSION	6
2.1 Introduction	6
2.2 Institutional Context	11
2.3 Conceptual Framework	16
2.4 Data	22
2.5 Empirical strategy	23

2.5.1	Measure of educational attainment	24
2.5.2	Measure of exposure to the expansion policy	25
2.5.3	Other explanatory variables	27
2.5.4	Sample	27
2.6	Identification	28
2.7	The expansion policy on the gender gap in higher education attainment	35
2.7.1	The effect of expansion policy by gender	35
2.7.2	Potential channels	36
2.8	Heterogeneity analysis	42
2.8.1	Heterogeneous effects by the pre-reform density of higher education institutions	42
2.8.2	Heterogeneous effects by regional income prior to expansion .	43
2.9	Robustness Checks	44
2.9.1	Compulsory schooling	46
2.9.2	Relaxation of the headscarf ban	47
2.9.3	Relaxing the parallel trend assumption	48
2.9.4	Specification and sample checks	49
2.9.5	Controlling for intensity in private universities	51
2.9.6	Controlling for time-varying regional covariates	53
2.9.7	Controlling for heterogeneous treatment effects	54
2.10	Conclusion	58

3	HIGHER EDUCATION EXPANSION AND COLLEGE WAGE PREMIUM: EVIDENCE FROM TURKEY	61
3.1	Introduction	61
3.2	Higher education expansion	65
3.3	Conceptual Framework	68
3.4	Data	72
3.5	Empirical strategy	78
3.5.1	Sample	80
3.6	Identification	81
3.6.1	Exogeneity of expansion	82
3.6.2	The expansion policy on the share of college graduates	83
3.6.3	Parallel trend analysis	84
3.7	Results	88
3.7.1	The expansion policy on college wage premium	88
3.7.2	The effect of the expansion policy by quantile regression	90
3.8	Robustness Checks	92
3.8.1	Alternative measure of exposure	93
3.8.2	Alternative sample specifications	93
3.8.3	Controlling for job and workplace characteristics	96
3.9	Conclusion	98

4	GENDER GAP IN INTERGENERATIONAL EDUCATIONAL PERSISTENCE: CAN COMPULSORY SCHOOLING REDUCE IT?	100
4.1	Introduction	100
4.2	The Compulsory Schooling Reform in 1997	104
4.3	Conceptual Framework	108
4.4	Data	114
4.5	Identification	118
4.6	Results	124
4.6.1	The impact of policy on intergenerational educational persistence	124
4.6.2	The impact of policy on gender gap in intergenerational educational persistence	127
4.6.3	Heterogeneity analysis by parental education	131
4.6.4	Heterogeneity analysis: Regional variations	135
4.6.5	Robustness Checks	139
4.7	Conclusion	141
5	CONCLUSION	143
	APPENDICES	157
A	ADDITIONAL RESULTS FOR CHAPTER 2	158
A.1	Returns to education in Turkey	158

A.2	Expansion policy on higher education enrollment: Public vs. private universities	159
B	ADDITIONAL RESULTS FOR CHAPTER 4	169
B.1	Regional variation of traditional gender views in Turkey . . .	169



LIST OF TABLES

2.1	Available slots and enrollment statistics on four-year higher education programs	15
2.2	Higher education statistics (four-year, two-year, and open-education programs)	16
2.3	The effect of expansion policy on educational attainment	32
2.4	The effect of expansion policy on gender gap	36
2.5	Decomposition of policy intensity on gender gap	41
2.6	The effect of expansion policy on gender gap: Heterogeneous impacts by the pre-reform density of higher education institutions	43
2.7	The effect of expansion policy on gender gap: Heterogeneous impacts by the regional wealth prior to the expansion	45
2.8	The effect of expansion policy: Heterogeneous cohort effects across five regions	49
2.9	The effect of expansion policy: Specification and sample checks	51
2.10	The effect of expansion policy: Controlling for demand driven supply change	52
2.11	The effect of expansion policy: Controlling for time-varying regional covariates	53

3.1	Higher education statistics	67
3.2	Average faculty member per student over years	70
3.3	Distributional evolution of wages by age group and educational attainment	75
3.4	The exogeneity of expansion policy	82
3.5	The effect of expansion policy on the share of college graduates	84
3.6	The effect of expansion policy on hourly wages	88
3.7	Robustness check with public exposure	93
3.8	Robustness check with alternative sample specifications	96
3.9	Robustness check with occupation and workplace controls	97
3.10	Robustness check with sector controls	98
4.1	Educational completion percentages of offspring by paternal education	109
4.2	Summary statistics for individuals born in 1980-1992	116
4.3	The effect of compulsory schooling on the intergenerational education transmission	126
4.4	The effect of compulsory schooling on the gender gap in intergenerational education transmission	128
4.5	The effect of compulsory schooling on the gender gap in intergenerational education transmission: Heterogeneity analysis by the number of educated parent	132
4.6	The effect of compulsory schooling on the gender gap in intergenerational education transmission: Heterogeneity analysis by the level of parental education	134

4.7	Regional differences on the effect of compulsory schooling on the gender gap in intergenerational education transmission	136
A.1	Returns to higher education by gender	159
A.2	The expansion policy on public and private enrollment at the NUTS-3 level	160
A.3	Higher education statistics across cohorts	161
A.4	Programs by field of study	161
A.5	Descriptive statistics	163
A.6	The exogeneity of expansion policy intensity	166
A.7	The orthogonality of migration probability to expansion policy intensity	167
A.8	The effect of compulsory schooling law on higher education attainment	167
A.9	Orthogonality of the expansion policy to regional conservativeness .	168
B.1	The gender view index at the NUTS-1 level	170
B.2	Robustness checks: Alternative bandwidths and time trend controls .	171
B.3	Robustness check with cohort 1986 is included in the sample	172
B.4	Robustness check with sample restricted to 2012 and 2016 waves . . .	172

LIST OF FIGURES

2.1	The number of universities across years.	11
2.2	The enrollment rate of 18-20-year-olds in higher education in years 2004-2012.	13
2.3	he share of women in higher education graduates by major and the share of majors in total available slots: Comparison of before and after expansion policy	18
2.4	The spatial distribution of intensity in 2013.	26
2.5	The higher education attainment across cohorts: High vs. low- intensity regions.	29
2.6	The gender gap in higher education attainment across cohorts: High vs. low intensity regions	31
2.7	The decomposition of intensity in 2013 across old and new universities.	37
2.8	The decomposition of intensity in 2013 across college majors.	39
2.9	The sample by cohorts and exposure to education policies.	47
2.10	Treatment effects by event study method	55
2.11	Treatment effects by Callaway Sant’Anna method	57
3.1	The number of universities across years.	66

3.2	Kernel density estimates of university entrance exam scores	69
3.3	Distribution of faculty members	71
3.4	Evolution of college wage premium	74
3.5	The spatial distribution of exposure across years.	77
3.6	Event study analysis of the expansion policy	85
3.7	Event study analysis of the expansion policy with pre-trend control .	87
3.8	Distributional effects of expansion policy by unconditional quantile regression analysis	91
4.1	Mean education completion rates by birth year	107
4.2	Mean education completion rates by birth year in treatment and placebo samples	120
A.1	The spatial distribution of intensity across years.	162
A.2	The higher education attainment across cohorts: High, Low and medium intensity regions in years 2009, 2011 and 2013.	164
A.3	The gender gap in higher education attainment across cohorts: High, Low and medium intensity regions.	165
A.4	Variations in the mean of intensities across years: High, medium and low intensity regions.	168
B.1	Geographical classification: West vs. East	170

CHAPTER 1

INTRODUCTION

The accumulation of human capital through education has long been considered crucial for economic progress (Lucas 1989, Mankiw et al. 1992), besides its documented importance for economic and social outcomes, such as higher income (Card 1999), healthier lives (Oreopoulos 2007), less involvement with crime (Lochner & Moretti 2004), etc. Many countries, seeking higher growth in addition to the documented economic and social returns, implemented education policies to increase access to education and raise the average educational attainment. The recent literature investigates the impact of education and related policy interventions on inequality, particularly family background and economic and social disadvantages experienced over the life course. This thesis aims to contribute to this literature and examine the effects of two major education policies on Turkey's education and labor market outcomes. In particular, the first essay examines the causal effects of the drastic expansion in Turkish higher education, which occurred in 2006-2008 with the establishment of new universities and increasing the educational opportunities both in the new and established universities, on the educational attainment disadvantage of women in higher education. The second essay investigates the causal impact of the same expansion policy in higher education on college wage premium in Turkey. Finally, the third essay analyzes the impact of extending mandatory years of schooling from 5 to 8 years in 1997 on the gender gap in intergenerational educational persistence.

Turkish higher education expansion occurred in 2006-2008 with the establishment of new public universities, and increasing the available slots in existing and newly established universities. In particular, from 2005 to 2009, 40 new public universities were established, and the total number of available slots

increased by 60%. The expansion was exogenous and unanticipated from the perspective of university candidates. It occurred in a short period and did not evolve endogenously with the changing patterns in demand across time. Hence the first-order effects of the expansion are expected to be on the supply side of the education market. Furthermore, the higher education expansion was politically motivated and driven by requests from members of the parliament (Arap 2010) prior to the 2007 Turkish general elections. The drastic expansion exposed students to an increase in the supply of higher education in varying degrees depending on their cohort and region of residence.

The first essay identifies the causal effect of policy on educational attainment and the gender gap by exploiting the region and cohort variations in the exposure intensity to the expansion policy by using Turkish Household Labor Force Surveys (HLFS). We further examine the possible channels through which varying exposure intensity might have affected the gender gap: The first channel, increased proximity to new universities, is likely to reduce financial and psychic costs of going to college and increase knowledge on the returns to higher education. Proximity to new universities may benefit women more if informational, financial, and psychic constraints are bigger obstacles for daughters than for sons (Caner et al. 2016, Rankin & Aytaç 2006). The second channel, distribution of slots across fields of study, may affect men and women differently if majors offered are more preferable or reachable to one gender than the other (Caner & Okten 2010).

This work contributes to two strands in the literature. First, it contributes to a small but growing literature that examines the effects of increasing supply of higher education on education outcomes and also on socio-economic inequalities in students' backgrounds (Bratti et al. 2008, Oppedisano 2011, Blanden & Machin 2004, Reimer & Pollak 2010, Di Pietro & Cutillo 2008). It also contributes to this literature on equality of opportunity in higher education by investigating the impact of the expansion policy on gender differences in educational attainment.

Second, this work contributes to the literature that analyzes the role of student preferences in explaining the gender gap in Science, Technology, Engineering, and Mathematics (STEM) majors. Even in advanced countries, where the education level of women has increased dramatically relative to men, women are still greatly underrepresented in STEM majors (Delaney & Devereux 2019a, Turner & Bowen 1999). It is further hypothesized that the scale of expansion across STEM and

non-STEM majors may have different effects on the gender gap in higher education attainment.

A higher education expansion similar to that experienced in Turkey occurred in all countries throughout the past century, but at different times, resulting that the number of adults with higher education has increased twenty-fold since the 1900s. While the US experienced a vast increase in the number of college graduates in the 1960s and 1970s, the same phenomenon occurred in other advanced countries, such as Canada, France, Spain, and Nordic countries in the 1990s. Many developing countries, seeking higher growth through increased human capital (Lucas 1989, Mankiw et al. 1992), experienced similar inflation of college graduates in the last two decades. This transformation in tertiary education attainment is bound to impact the structure of wage differentials.

Many contributions in the literature have linked the demand for skilled labor and labor market institutions with the evolution of the educational wage differentials of the tertiary-educated labor force. Several studies analyzing the educational wage differentials in the US labor market provide evidence on the growing college wage premium in the 1960s and 70s parallel to the increase in the supply of college graduates. One prominent view on the underlying causes for the rising college wage premium is the skill-biased technological change (SBTC). On the other hand, Card & DiNardo (2002) and Lee (1999) argue that the changes in the labor market institutions, such as a decline in unionization and stagnant minimum wages, explain a substantial part of the wage differentials by educational attainment. In the UK (Walker & Zhu 2008) and several other European countries (Crivellaro 2014), the wage differential of college to high school graduates followed a flat path along with the increase in college graduates. (Crivellaro 2014) argues that even with a higher demand for skilled workers, compressed and regulated labor markets in Europe have led to lower educational wage inequality.

Even though these studies focus on the educational wage differentials at a time of relative supply increase of college graduates, the role of the supply in determining the wage differentials did not receive much attention. Turkey's recent higher education expansion policy, an exogenous reform that leads to a large increase in the supply of college graduates over a short period of time, offers an excellent case to identify the role of supply increase in wage differentials, whose results are presented in the second essay of this thesis.

This work identifies the causal impact of the increase in the supply of college graduates induced by higher education expansion on the additional wages earned by young college graduates using the difference-in-differences methodology on a sample of employed men constructed from Household Labor Force Surveys. Since the expansion reform started in all regions simultaneously and increased gradually over time, the identification strategy relies on region and year variations in the level of exposure to capture the causal relationship between the additional slots introduced in the expansion period and the changes in the college wage premium.

This study makes an important contribution to the literature on the effects of education policies on labor market outcomes. Few studies examine the effect of higher education expansion on returns to education (Devereux & Fan 2011*a*, Kyui 2016) or labor market advantages of young graduates (Walker & Zhu 2008, Li et al. 2014, Xing et al. 2018). This work differs from existing studies in two ways: First, as opposed to the studies estimating the local average treatment effect of the expansion policy, as in Devereux & Fan (2011*a*) and Kyui (2016), it examines the impact of a large-scale expansion on all labor market participants and hence identifies how it bounds to structure the labor market over time. Second, the identification strategy that relies on the regional exogenous variation in the supply of skilled labor distinguishes this study from those that rely on the variations in the educational attainment of different birth cohorts that have limited power to exclude demand-driven variations, as in Devereux & Fan (2011*a*), Walker & Zhu (2008).

Another aspect of inequality explored in the literature is the dependence on parental education, also known as the intergenerational transmission of education. 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.

The third essay of this thesis leverages another Turkish education reform that

extends the mandatory years of schooling from 5 to 8 years and went into effect in 1997. For political reasons, the reform was implemented quickly and unexpectedly, and students who had completed the fifth grade in the Summer of 1997 were exempted from the reform's mandate, while younger cohorts were required to complete at least eight years of schooling. This study explores 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 offspring's education. Studies in this literature are all from advanced economies with high levels of income equality and intergenerational educational mobility (Meghir & Palme 2005, Aakvik et al. 2010, Betthäuser 2017). Hence it is a mostly unexplored question to what extent policy interventions can effectively increase educational mobility in low equality and low mobility society. 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.

This study contributes to the literature on gender inequality and the role of government policies in bringing about equality between men and women (Spohr 2003, Fang et al. 2012, Tsai et al. 2009, Dayioglu et al. 2016). Several studies identified differential preferences of parents for their sons and daughters as a source of women's lower educational attainment in Turkey (Rankin & Aytac 2006, Caner et al. 2016), 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.

CHAPTER 2

ATTAINMENT AND GENDER EQUALITY IN HIGHER EDUCATION: EVIDENCE FROM A LARGE SCALE EXPANSION

2.1. Introduction

The accumulation of human capital through education has long been considered crucial for economic progress (Lucas 1989, Mankiw et al. 1992, Ehrlich et al. 2018). Yet, in many developing countries, higher education attainment rates remain low; moreover, young women are still under-represented in tertiary education (Barro & Lee 2013).¹

In recent years, several advanced and emerging countries, including Italy, France, England, Russia, and Turkey, have implemented reforms aimed at promoting higher education attainment and increasing equality of opportunity in higher education (Bratti et al. 2008, Blanden & Machin 2004, Kyui 2016, Devereux & Fan 2011*b*). In this study, we identify the causal effect of the drastic expansion in the Turkish higher education system on tertiary education attainment and gender gap in attainment. The expansion was fast and rather unexpected. From 2005 to 2009, 40 new public universities were established and the total number of available slots increased by 60%. Prior to the reform, in 2006, among 25-34 year-olds, the share of tertiary education graduates in Turkey was 14 percent, much lower than the Organization of Economic Cooperation and Development (OECD)

¹According to Barro & Lee (2013), the population over age 15 is estimated to have 11 years of schooling on average in high-income economies, compared to 7.1 years in developing countries. In Sub-Saharan African and South Asian countries attainment rates are the lowest at 5.2 years on average and women are greatly under-represented in higher education.

average of 33 percent at the time (OECD 2020a).² Furthermore, unlike in most OECD countries, there was a gender gap to the detriment of women as only 12 percent of women were higher education graduates compared to 15 percent of men in the 25-34 age group.

The higher education expansion in Turkey was exogenous and unanticipated from the perspective of university candidates. It occurred in a short period and did not evolve endogenously with the changing patterns in demand across time. Prior to the expansion, demand for higher education was higher than available slots (COHE 2004, 2007). Indeed, in 2005, around 2 million applicants took the university entrance exam, but only 10% were able to enroll in a four-year higher education program (COHE 2007). Therefore, a centralized competitive examination was and still is applied to ration excess demand for higher education. Hence the first-order effects of the expansion are expected to be on the supply side of the education market. Furthermore, the higher education expansion was politically motivated and driven by requests from members of the parliament (Arap 2010) prior to the 2007 Turkish general elections. The drastic expansion exposed students to an increase in supply of higher education in varying degrees depending on their cohort and region of residence.

We identify the causal effect of policy on educational attainment and the gender gap by exploiting the region and cohort variations in the exposure intensity to the expansion policy. Using Turkish Household Labor Force Surveys (HLFS), we focus on individuals who were 18 years old—and expected to start higher education—in the years 1999-2013. Spatial and time variation in the exposure intensity measure allows us to control for both region and cohort effects. Hence our estimates disentangle the effect of the expansion from time trends and time-invariant regional differences in educational attainment and can be interpreted as the causal impact of the reform.

For the purposes of our study, careful and detailed work was required in the construction of a new dataset with regional and cohort-specific variables on the exposure intensity. We collect the data on available slots in four-year programs in higher education from the Measuring, Selection and Placement Center (OSYM)

²Organisation for Economic Cooperation and Development (OECD) is currently comprised of 36 members, including 21 European Union member states and other countries including the USA, the UK, Mexico, and Turkey.

Almanacs for each region and year. We measure the exposure intensity, similarly to Card (1992), Duflo (2001), and Berlinski et al. (2009), by the increase in available slots per 18-year-old population. In particular, the intensity for cohort t is defined as the increase in available slots in public universities in a given region from 2005 to year t , per 18-year-old in year t . Hence, our exposure intensity measure varies across cohorts and regions.

Two studies examined the 2006-2008 higher education expansion in Turkey from different perspectives. Yilmaz (2014) used aggregate city-level data to examine the effect of the increase in the number of universities on gender gap in higher education and found that the expansion reduced the gender gap. Since the expansion is defined as a policy dummy variable and data are aggregated at the city level, it is not possible to take advantage of time-specific and region-specific effects of the reform in that study. Polat (2017) analyzed the impact of the expansion on education outcomes and how outcomes varied according to family income by comparing higher education enrollment of 18-25-year-olds in HLFS 2004 to that in HLFS 2012 and observed a rise in enrollment. As a before-and-after study using two data points in time, his study also did not control for time-specific and region-specific effects.

An increase in the supply of higher education may not necessarily reduce existing inequalities in access to education if the initially disadvantaged groups do not benefit from the increased supply. We consider the structure of the reform in terms of exposure intensity across regions and fields of study and conceptualize how it might affect the gender gap in attainment. Our rich dataset enables us to investigate the impact of exposure intensity in more detail and to study the underlying reasons for the observed changes in educational attainment for men and women.

We examine the possible channels through which intensity might have affected the gender gap: The first channel, increased proximity to new universities, is likely to reduce financial and psychic costs of going to college and increase knowledge on the returns to higher education.³ Card (1995) notes that proximity to college has a strong effect on completed college education, even after controlling for parental

³Several studies indicate that the perceived returns to education are lower than the actual returns and this leads to lower investment in schooling (Attanasio & Kaufmann 2009, Kaufmann 2014, Nguyen 2008, Jensen 2010).

education, region, and IQ. Frenette (2004) and Spiess & Wrohlich (2010) shows the relation between the distance to college and higher education attainment in Canada and Germany, respectively. The expansion in Turkey increased the supply of and reduced the distance to higher education. In 2005, prior to the expansion, the number of cities (provinces) that lacked a university was 41 out of 81; by 2009 there were only 2 cities without a university. Proximity to new universities may benefit women more if informational, financial, and psychic constraints are bigger obstacles for daughters than for sons (Caner et al. 2016, Rankin & Aytaç 2006).

The second channel, distribution of slots across fields of study, may affect men and women differently if majors offered are more preferable or reachable to one gender than the other (Caner & Okten 2010). We develop these channels in more detail in the conceptual framework section.

Our work contributes to two strands in the literature. First, we contribute to a small but growing literature that examines the effects of increasing supply of higher education on education outcomes and also on socio-economic inequalities in students' backgrounds (Bratti et al. 2008, Oppedisano 2011, Blanden & Machin 2004, Reimer & Pollak 2010, Di Pietro & Cutillo 2008). We contribute to this literature on equality of opportunity in higher education by investigating the impact of the expansion policy on gender differences in educational attainment.

Second, our work contributes to the literature that analyzes the role of student preferences in explaining the gender gap in Science, Technology, Engineering, and Mathematics (STEM) majors. Even in advanced countries, where the education level of women has increased dramatically relative to men, women are still greatly underrepresented in STEM majors (Delaney & Devereux 2019a, Turner & Bowen 1999). We hypothesize that the scale of expansion across STEM and non-STEM majors may have different effects on the gender gap in higher education attainment. Ours is the first paper that analyzes the impact of higher education expansion in different fields of study in addition to the impact of the overall expansion. In particular, we examine, how expansion in engineering and social science slots has affected the gender gap in higher education attainment.

Our results show that, in the overall, expansion policy increased the higher education attainment of both men and women, but did not reduce the gender gap in attainment above and beyond the time trend. Furthermore, we find the

intriguing result that the increase in available slots from newly established universities benefited men more than women thereby having an increasing effect on the gender gap. In order to examine this perplexing result, we examine fields of study offered in these new universities as well as in existing ones. While the increase in available slots is similar in old (existing) and new universities, the highest growth in available slots occurred in the fields of social sciences and engineering. We find that the expansion in social sciences benefited men and women almost evenly; therefore, it had no effect on the gender gap. However, the expansion in engineering in newly established universities, benefited men more than it did women, as it is a major traditionally preferred by men. These results support our main findings. In addition, we do a number of robustness checks including alternative samples and specifications. Our results are robust to these checks.

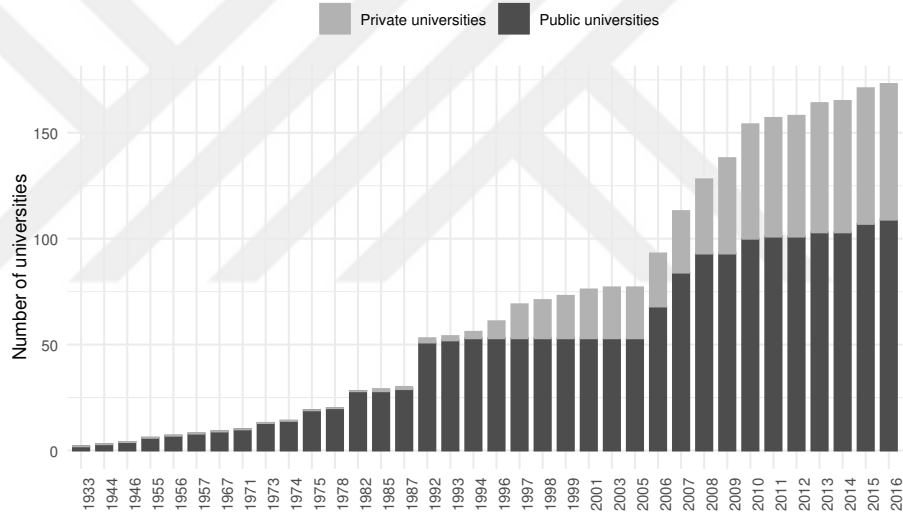
The literature on the effects of expansion policies on the socio-economic gap in higher education generally finds that expansion policies do not decrease the socio-economic gap (Blanden & Machin 2004, Reimer & Pollak 2010). In a similar vein, we show that the expansion policy did not reduce the gender gap in higher education in Turkey above and beyond the time trend. Hence, our results provide support for the argument that the expansion policies as they have been designed may not have been the best tools to address inequalities across different socio-economic and demographic groups. Our findings on how the allocation of the additional slots to different fields of study has affected men versus women inform us that an expansion policy with a specific target to reduce inequality needs to consider how the preferences of compliers of the policy may vary across different socio-economic and demographic groups. As such, our findings can inform future expansionary policies in parts of the world where women are still under-represented in tertiary education.

The remainder of the paper is organized as follows: Section 2 introduces the institutional context and expansion in Turkish higher education system; section 3 introduces the conceptual framework; Section 4 describes the data used in the analysis. Sections 5 and 6 introduce the empirical strategy and identification to analyze the causal effect of the expansion policy on higher education attainment, and Sections 7 and 8 present the results and the heterogeneity analyses, followed by the robustness checks of our results in Section 9. Finally, Section 10 states the concluding remarks.

2.2. Institutional Context

This section introduces the education system in Turkey and the changes in education policy in the years that are relevant to our study. In Turkey, the Ministry of National Education (MONE) is in charge of all structural reforms and education policies for primary and secondary education. Since 1981, all universities are affiliated with the Council of Higher Education (COHE), which is an independent entity of the central government, and COHE regulates the tertiary (higher) education.

Figure 2.1: The number of universities across years.



Source: Council of Higher Education, Higher Education Statistics.

Prior to 1997, formal education consisted of five years of primary school, three years of lower secondary education (four years if a preparatory year was required), three years of upper secondary education (four years if a preparatory year was required), and higher education. Primary education was compulsory for all citizens. In 1997, the government increased mandatory schooling from five to eight years by merging primary and lower secondary education under the umbrella of primary school. According to the law, students who had not yet graduated from the 5th grade in the summer of 1997 were supposed to finish eight years of compulsory schooling. Upper secondary education is provided in general, vocational, or technical high schools and takes three or four years. Since 2005, students starting high school were required to study for four years. All levels of education up to

higher education are free in public schools.

High school students choose a broad field of study in their second or third year of high school, such as quantitative, equally weighted, verbal, foreign languages, or art. In the curriculum, the hours spent on the subjects related to chosen field of study are determined to be higher by the Ministry of Education. While students who chose quantitative field spend most of their time on science and mathematics, those who choose verbal field spend most time on Turkish and social science courses.

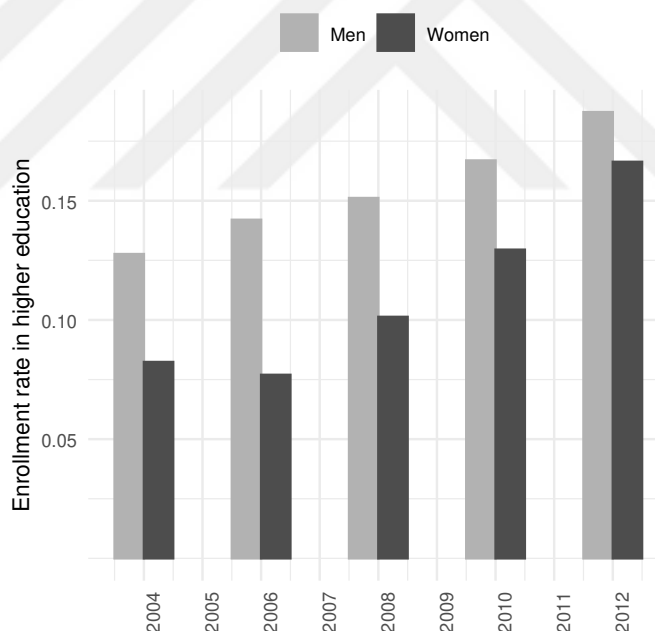
Demand for undergraduate education has been greater than supply in Turkey (COHE 2004, 2007). In 2003, around 1.5 million people applied to enroll in a higher education institution, of which only 12% and 10% were eligible to start 4-year and 2-year higher education programs, respectively (COHE 2004). The number of applicants increased to 2 million in 2005. As a result, the share of applicants who were able to enroll in a 4-year higher education program dropped to 10% due to demand for 4-year college programs exceeding the available slots for such programs (COHE 2007). To ration the excess demand, a centralized competitive examination is applied since 1974. Although the questions that students face in the university exam change over time, the exam is mainly composed of the five main subjects that student study during their high school education: mathematics, Turkish, science, social sciences, and foreign language. In addition to the curriculum differences in their high school education, students face some additional rules in the way their university entrance exam score is calculated based on their chosen field of study in high school. Students select and answer questions in different sections of the exam based on their field of study in high school. To be used in applying to different majors in college, a set of raw scores is calculated for each applicant by assigning a higher weight to sections that are more closely related to the targeted major in college. The raw exam scores are further adjusted for high school performance. In this adjustment, students are incentivized to choose a major that is closely related to their high school field.

The placement of exam takers to higher education programs is based on a centralized matching algorithm that takes into consideration the applicant's ranked preferences for higher education programs (at most 24 programs can be submitted in the preference list) and the applicant's scores that are suitable to the programs. For example, students who apply for a social science program are ranked according

to the ("equally-weighted") scores that assign higher weights to mathematics and Turkish language sections of the exam. Whereas students who apply for an engineering program will be ranked according to the ("quantitative") scores that assign higher weights to mathematics and science sections of the exam. Students are expected to start the higher education program they are matched to in year of the exam, or else are penalized with a reduced high school performance score in the following year's exam.

Higher education in Turkey includes two-year vocational colleges, four-year undergraduate programs, in addition to post-graduate level education. In this study, we are interested in four-year undergraduate programs.⁴

Figure 2.2: The enrollment rate of 18-20-year-olds in higher education in years 2004-2012.



Source: HLFS, 2004-2013.

Notes: Authors' calculations.

There are two types of universities in Turkey, namely public and private (non-profit foundation) universities. Public and private universities offer three types of training: normal education, evening education, and open (distance)

⁴The overwhelming majority of programs take four years. There are only a few exceptions that take longer: Schools of Dentistry, Pharmacy, and Veterinary take 5 years; School of Medicine takes 6 years of education. These are also included in our study.

education. Students who study in normal or evening education receive education at the university, whereas those who attain an open education program receive distance education and are only obliged to pass centralized exams offered at designated exam centers. Normal and evening education programs have limited quotas determined by the COHE, but many programs of open education do not have quota restrictions.

The first public university was introduced in 1933, and there was a gradual increase in the number of public universities until 1987. Turkish government introduced the first phase of higher education expansion in 1992-1993. During the two years of the expansion, 24 new public universities were established in 23 cities that did not have a university before. Between 1992 and 2005, the number of public universities remained relatively constant, and the number of private universities slightly increased, as shown in Figure 3.1. In the second phase of the higher education expansion, as in the first phase, cities without universities were the target of the expansion policy and 40 new public universities were established in these cities between 2006 and 2008. As a result, the number of public universities increased from 53 in 2005 to 93 by 2008, and the number of available slots increased by more than 60%. The rise in enrollment in higher education started in 2006, i.e., as soon as the new slots are introduced in the first year of the expansion, and continued even after the expansion period. The enrollment rates of 18-20-year-old men and women increased from 13% and 8% in 2005 to 17% and 13% in 2010, respectively (Figure 2.2). This second wave of higher education expansion is the focus of our study.

As quotas expand, students are not automatically assigned to these new quotas; instead, the slots have to be preferred by the students who enter the centralized matching algorithm of exam takers and higher education slots. We provide statistics, collected from Measuring, Selection, and Placement Center (OSYM) Almanacs, on the evolution of the available slots and the share of these slots not taken by that years' prospective students in Table A.3 in the Appendix. These statistics show that as the additional slots are gradually introduced over the years, the share of unused slots in newly established universities exceeds the share in existing universities, showing the role of student preferences in the higher education admission system.

The higher education expansion was politically motivated and driven by

Table 2.1: Available slots and enrollment statistics on four-year higher education programs

	Slots in four-year programs	Empty (unused) slots in four-year programs	Share of empty slots in old universities	Share of empty slots in new universities
2002	173,147	7,532	4.35%	-
2003	177,208	1,293	0.73%	-
2004	185,180	302	0.20%	-
2005	188,899	74	0.04%	-
2006	187,889	15,100	8.20%	5.34%
2007	192,049	4,343	2.37%	0.98%
2008	257,111	1,216	0.41%	0.99%
2009	299,505	13,373	4.11%	6.80%
2010	339,745	18,499	4.42%	9.32%
2011	357,275	20,187	4.96%	8.31%
2012	387,747	37,806	8.6%	13.5%
2013	405,802	23,473	4.81%	8.18%

Source: OSYM Almanacs.

Notes: These statistics are generated by adding up the quota and enrollment statistics for all four-year undergraduate programs in the OSYM almanacs for 2002 and later years. We could not access data on enrollments and quotas for 2002 and earlier years.

requests from members of the parliament to establish universities in their cities (Arap 2010). The Turkish general election was to be held in 2007 to elect 550 members of parliament. Before this, in 2005, 25 new public universities were planned with the stated purpose of providing economic benefits to the regions where they were to be established. Due to pressures from members of the parliament who represented different cities and sought reelection 15 more public universities were opened by 2008 than initially planned.

In 2005, prior to the expansion, the number of cities that lacked a university was 41 out of 81, whereas in 2009 there were only 2 cities without a university. The number of cities with more than one university increased from 6 to 8. In many cities, the number of available slots in higher education more than doubled with the establishment of new universities and an increase in slots in existing universities.

The COHE serves as the head of all higher education institutions since 1981 and monitors available slots for each university and each academic track. In Turkey, demand for four-year undergraduate programs exceeds the number of available slots, whereas the opposite is true for two-year vocational colleges (COHE 2007). Therefore, the expansion targeted primarily four-year undergraduate programs. Between 2005 and 2009, the available slots in four-year and two-year programs increased by 62% and 56%, and the number of new students increased by 56% and 29% percent, respectively. Even though there is no quota constraint in most of the

Table 2.2: Higher education statistics (four-year, two-year, and open-education programs)

	Slots in four-year programs	New students in four-year programs	Slots in two-year programs	New students in two-year programs	New students in open educ. programs
2005	190,655	194,516	195,667	216,608	266,122
2006	188,393	177,258	193,541	235,033	256,138
2007	193,886	193,541	211,460	236,881	256,818
2008	259,748	264,088	260,155	287,547	387,660
2009	309,167	300,029	305,354	280,016	282,918
2010	343,583	322,961	308,980	232,939	202,513

Source: OSYM News Bulletins.

Notes: OSYM publishes these statistics annually in its news bulletins; the format of the bulletins may vary. The statistics presented in this table also include the quotas reserved for students graduating with the highest rank in their high school; hence the difference from the statistics presented in Table 2.1.

open education programs, the increase in new students in open education is only 6% (See Table 2.2 for further details). Thus, we conclude that the expansion in four-year programs was the primary component of the expansionary policy on the supply side.

When the reform was adopted at the end of 2005, the Council of Higher Education opposed the draft by saying that there were insufficient human resources and infrastructure for the establishment of the new universities. Indeed, despite the fact that a total of 29 new universities were established in 2006 and 2007, the increase in the number of new students in the universities and available slots in higher education programs was only gradual. Table 2.2 presents higher education statistics across years in the country. It shows that the main increase in available slots and new students in higher education is not observed in 2006, so in this paper, we determine the post-expansion years as 2007 and later. This is also consistent with the political background of the expansion policy. The COHE president who opposed the expansion was replaced by his successor in 2007. The new COHE president embraced and implemented the expansion policy immediately after he took office (Arap 2010).

2.3. Conceptual Framework

In the absence of credit constraints and rationing of higher education, individuals make privately optimal higher education decisions. However, in Turkey, the number of available slots has been much lower than the number of applicants for higher education (COHE 2004, 2007). Prior to the expansion, 41 of the 81 cities did not

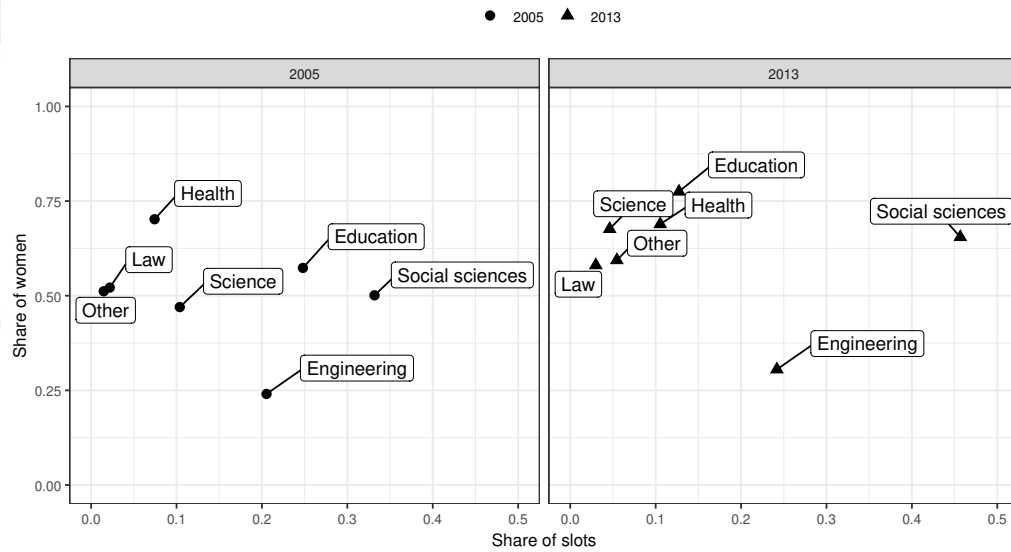
have any universities. Hence, many high school graduates lived at a considerable distance from a university. Although tuition has been very low in public universities, distance to university can be a significant factor in increasing both monetary and psychic costs of going to college. Card (1995) notes that proximity to college has a strong effect on completed college education, even after controlling for parental education, region, and IQ. There are substantial monetary costs of going to a college away from home. In the Turkish context, there can also be significant psychic costs for parents of sending their children far away for education. Students, as well as their parents, especially in less developed countries, may lack information and underestimate the monetary and social returns to education. Furthermore, perceived monetary and social returns to education may vary across genders, making higher education more or less compelling based on a student's gender. For example, in Turkey, while higher education increases the likelihood of employment for women, higher educated women are still less likely to be employed than their male counterparts although there are no significant wage differences between men and women with higher education.⁵ The proximity to universities may increase information available to students and their parents on the monetary and social returns to higher education. Hence, perceived benefits and costs of higher education attainment might both be affected by proximity to universities.

In order to understand how higher education expansion in Turkey affects the educational attainment by gender, we need to understand how the characteristics of the expansion will interact with gender differences in higher education attainment. Here, we focus on two factors: College proximity in relation to higher psychic costs of/lower employment prospects for women; majors supplied in relation to preferences for majors across genders.

Parents in Turkey, especially those with a conservative view are less likely to support their daughters' education beyond the compulsory level (Rankin & Aytac 2006, Caner et al. 2016). They incur higher psychic costs from sending their daughters away to college, compared to their sons, due to social norms (Caner et al. 2016, Dayioglu et al. 2016). Furthermore, they may perceive that returns to education are also lower for daughters as women with higher education are less likely to be employed than their male counterparts. Therefore, proximity to new

⁵We estimate the differential employment and wage returns to education across genders by using the Mincer model presented in Table A.1 in the Appendix. We explain the details of the analysis and discuss the results in Section A.1 in the Appendix.

Figure 2.3: The share of women in higher education graduates by major and the share of majors in total available slots: Comparison of before and after expansion policy



Source: HLFS, 2016-2017 and OSYM almanacs, 2005 and 2013.

Notes: Authors' calculations. The figure in the left-hand side gives the field of major distributions for cohorts 2003-2005 (pre-expansion) and the distribution of slots across majors in 2005. The figure in the right-hand side gives the field of major distributions for cohorts 2011-2013 (post-expansion) and the distribution of slots in 2013. In HLFS, the information on the majors in college is available only for graduates. The cohorts 2011-2013 might be still in education, hence the graph on the right hand side presents the share for graduates, but not enrolled students in HLFS 2016-2017.

universities may help narrow the gender gap in higher education by reducing these psychic and monetary costs of sending daughters away, and hence, more parents may now find that benefits exceed monetary and psychic costs when they make higher education decisions for their daughters. In addition, the establishment of new universities might induce demand for higher education overall, even in the universities away from their hometown, by revealing information on returns to college or eliminating the social and cultural barriers to education.

Another consideration is that being close to universities may influence the anticipation and motivation of high school students towards obtaining higher education. Compared to students in provinces without any universities, high school students in provinces with at least one university may have an advantage in human capital accumulation when they take the entrance exam. Therefore, students living in those provinces with higher anticipation of attending college may benefit more from the additional slots, limiting the expected positive impact of the college proximity in the short run. This assumption implies that the effect of the reform may depend on whether a university already exists or not where the reform takes place. In the heterogeneity analysis part of our paper, we consider the concerns on whether the reform effect depends on pre-reform density of higher education institutions.

It is well documented in the literature that there are significant differences in college major choices of men and women (Delaney & Devereux 2019*b*, Duta 2016). Caner & Okten (2010) show that there are important differences between men and women’s career choices in the Turkish context. They use a large micro data set that includes university applicants’ socio-demographic characteristics, major preferences and university entrance exam scores. Figure 2.3 presents evidence on college major preferences in HLFS 2016-2017 data separately for cohorts exposed and not exposed to the expansion policy, by exploring the relation between the share of women (for those expected to start tertiary education in 2003-2005 on the left, and in 2011-2013 on the right⁶) and share of available slots in 2005 and 2013 across majors, obtained from OSYM almanacs. Among pre-expansion cohorts presented on the left, while men appear to dominate the engineering field and

⁶In HLFS, the information on the field of study (major) in college is available only for graduates. The cohorts exposed to the expansion (took the university exam in 2007-2013) might be still in education, hence the graph on the right hand side of Figure 2.3 presents the share for graduates, instead of enrolled students in HLFS 2016-2017.

women dominate the health field, men and women seem to be equally represented in social sciences and other majors (See Table A.4 in the Appendix for the program list of each field of study). The share of women has increased in all majors in post-reform cohorts with more noticeable increases in education, science and social science majors than in law, health, and engineering majors. Women are underrepresented in engineering majors even after the reform, although the share of slots in 2013 appears to be the second highest in engineering majors.

A gender difference in the selection of students into college majors might be due to an indirect barrier on the demand side resulting from a gender difference in the type of human capital accumulated in the years before the university entrance exam. As argued by many in the literature, the gender differences in college major choice might be related to either the comparative advantage and/or ability distribution of men and women in different subjects, or the variations in job and workplace preferences of men and women. However, as emphasized by Delaney & Devereux (2019*a*), gender differences in comparative advantage might also be considered as a preference oriented outcome, since students make many decisions about their human capital accumulation up until the college entrance age, and these decisions collectively determine their comparative advantage in certain majors.

For the years in our study, several papers provide essential evidence showing that men and women might be distinguished by their academic choices and performance in their transition from high school to college. First, Saygin (2016), based on an administrative micro dataset of university entrance exam participants in 2008, points out that in high school women are more likely to select courses that require abilities in language and mathematics, while men tend to select more technical and science courses. Since the exam score used in the placement algorithm varies by major and field of study in high school, women’s comparative advantage might be lower than men’s in quantitative majors such as engineering. Similar to their field choice in high school, average university exam scores are found to be slightly higher in equally-weighted and verbal subjects for women, and in quantitative subjects for men (Saygin 2016, Kara & Okten 2019). Secondly, Saygin (2020) also shows, using the same dataset, that while females outperform males in high school performance in every subject and every quantile, males outperform females in the university entrance exam score in almost all quantiles conditional on the high school performance, with the largest magnitude in the quantitative subjects and highest quantiles. This evidence indicates two potential barriers for women’s enrollment

into higher education. First, since new universities are initially questionable in their quality, supported by the high share of empty slots as shown in Table 2.1, all students are more likely to prefer old and prestigious universities, *ceteris paribus*. The outperformance of men in the highest quantiles makes women more susceptible to the distribution of slots offered in new universities across majors. Second, women's lower performance in quantitative subjects makes them less likely to benefit from the additional slots in quantitative fields, such as engineering.

In summary, there is no evidence of unambiguous outperformance of either gender in the competition for slots in colleges. However, the differences in human capital accumulation of men versus women might generate a gender gap in higher education attainment, if the newly introduced slots are not allocated across majors to reflect such differences.

In our analyses, we consider the slots introduced in different fields of study as imperfect substitutes to each other and assume that preferences across majors remain relatively stable in our period of analysis. Reimer & Pollak (2010) study supports our assumption. Reimer & Pollak (2010) analyze the evolution of the choices for the field of study across different social groups in Germany during 1980-2000, at a time when enrollment in tertiary education increased rapidly. They find that the educational strategies of students from different social backgrounds remained mainly unchanged despite the ongoing expansion of tertiary education.

Hence, the allocation of the increase in available slots to different majors becomes an essential component of the expansion policy that can affect higher education attainment and the gender gap. For example, if most of the expansion occurred in the supply of engineering majors, then the expansion in the overall might benefit men more than women, whereas the expansion in social sciences might benefit the two genders pretty evenly. Note that, as we mentioned in the previous section, the distribution of extra slots during the expansion across majors has depended on supply-side constraints, since infrastructure and human resources such as instructors were limited and the expansion occurred in a short period. Hence, it is reasonable to assume that it was mostly the supply-side constraints that determined the majors offered, rather than the preferences of university applicants.

2.4. Data

Our main aim is to explore the impact of expansion policy on educational attainment. To that end, we match region-level data on the expansion policy to the Turkish Household Labor Force Surveys (HLFS). The HLFS is one of the largest household surveys available in Turkey and it is carried out annually by the Turkstat. It collects demographic and labor market data from roughly 390,000 individuals living in 150,000 households. For each individual, the region of residence is known at the NUTS-2 level.⁷ The survey has data on completed education, whether the individual is currently in education, and, if so, the level of education currently enrolled in and other demographic data, such as age and gender. In our analyses, we use the 2016 and 2017 waves of HLFS to estimate our main regressions. We also make use of earlier waves of the HLFS (2005 and 2007) to show that the intervention to higher education is exogenous to pre-existing trends in education.

We construct our key variable to measure the supply of higher education in a given region and year by collecting data on the number of available slots in four-year higher education programs from the Almanacs of the OSYM. Each year the OSYM publishes data on the available slots for all programs (majors) in all universities and the number of students placed in these programs. We collected these data for the 2005-2013 period. The location identifiers (province (NUTS-3), and NUTS-2 regions) of each university are also known. We aggregate the available slots at the NUTS-2 level to merge with our HLFS data. The number of available slots in higher education programs is an indicator of the supply of higher education in a given year and region and ours is the first study that merges this data to HLFS data.

Our second region-level data is the population of 18-year-olds at the province level, obtained from the Address Based Population Registry System, available since 2007. We aggregate the 18-year-old population information in 2007-2013, at the NUTS-2 level to merge with our HLFS data. The number of 18-year-olds represents the size of the population that benefits from the higher education opportunities in a particular region since an individual is typically 18-year-old when he graduates from high school. Centralized university exam which is based on high school

⁷There are 26 NUTS-2 regions in Turkey.

curriculum makes it very difficult if not impossible for students to take a break or work for a while before continuing to higher education since they could forget the material that the exam is based on. We use the number of 18-year-olds to normalize the available slots in order to control for any scale differences across regions.

We also use several auxiliary datasets to augment our region-level data. We use the gross domestic product (GDP), public investment in all industries except for education, and population data obtained from Turkstat and Ministry of Development at the province (NUTS-3) level to include in some specifications. Also, to demonstrate the orthogonality of expansion policy to demand, we collect data on the number of university entrance exam takers from OSYM almanacs and the value-added and total employment across industries to calculate labor productivity at the province level. All data are aggregated at the NUTS-2 level to merge with the microdata from HLFS. We also use the Barometer survey conducted by KONDA (a private research and consultancy company) to show orthogonality of expansion to the level of conservativeness across regions which might be related to the extent of other educational interventions' impact on education outcomes.

2.5. Empirical strategy

In our empirical analyses, cohort and region of residence at the university entrance age jointly determine exposure to higher education expansion. We measure the causal effect of the exposure to expansion by exploiting the spatial and temporal variations in the exposure to the expansion policy. Our empirical approach is similar to Card (1992), Duflo (2001), Berlinski et al. (2009), and Müller & Wrohlich (2020).

We introduce a generalized regression framework of the identification with the following linear probability model:

$$h_{irt} = \beta_1 s_{irt} + \beta_2 w_i s_{irt} + \alpha_0 + X_{irt} \alpha_1 + w_i X_{irt} \alpha_2 + X_{rt} \alpha_3 + w_i X_{rt} \alpha_4 \quad (2.1)$$

$$+ \gamma_{m,r} + \gamma_{w,r} + \mu_{m,t} + \mu_{w,t} + \epsilon_{irt}$$

where h_{irt} represents our dependent variable and s_{irt} represents our explanatory

variable of interest, exposure to treatment for individual i in cohort t and region r . These variables will be described in detail in subsections below. w_i is a dummy variable that is 1 for a woman. X_{irt} and X_{rt} represent individual-level and region-level controls, respectively. The inclusion of region effects, $\gamma_{m,r}$ for men and $\gamma_{w,r}$ for women, control for time-invariant unobserved factors that might be correlated with the dependent variable. The cohort fixed effects, $\mu_{m,r}$ for men and $\mu_{w,r}$ for women, capture any time-related differences observed in the dependent variable at the country-level. By the inclusion of two-way fixed effects, our parameters of interest β_1 and $\beta_1 + \beta_2$ capture the correlation between the exposure to reform and within region variations over time in the dependent variable. In this framework, our estimates can be interpreted as the causal effect of the expansion policy.

Abadie et al. (2017) suggest that clustering should be treated as a research design issue and that the standard errors should be clustered if the treatment assignment mechanism is clustered. Since the higher education expansion policy (i.e., our treatment variable) is determined at the region level, in our analyses the estimated standard errors allow within-region clustering.

2.5.1. Measure of educational attainment

We define our dependent variable *higher education attainment* as a binary variable which is equal to 1 if the individual is a graduate or is currently enrolled in a university (rather than a binary variable based on graduation only). The motivation behind this definition is to balance the treatment and control groups in attainment rates. In Turkey, the mean age of college graduation is 25 (OECD 2017). Individuals in the post-expansion group belong to 2007 and later cohorts and they are 21-28 years old in our sample. Thus, in our sample, the post-expansion cohorts are more likely to be continuing their higher education than the pre-expansion group. To elaborate on our point, we compare the averages of higher education graduation rates and higher education attainment (graduates plus currently enrolled ones) rates by cohorts in our sample (Table A.3 in the Appendix). The graduation rates are lower in younger cohorts, since many members of those cohorts are still students. For example, in 2005 and 2009 cohorts, graduates constitute 92% and 86% of those who attained higher education; yet, in

2012 and 2013 cohorts, the shares are only 60% and 43%. Therefore, we use *higher education attainment* as our dependent variable instead of *higher education graduation* to analyze education outcomes of the expansion policy. We should note that graduation rates in tertiary education conditional on enrollment are quite high in Turkey (93% among 25-64-year-olds by Adult Education Survey, 2016 conducted by Turkstat).

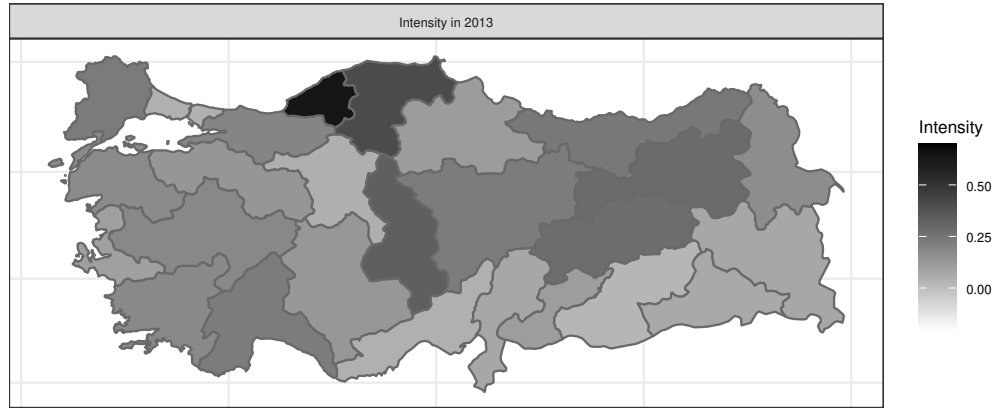
2.5.2. Measure of exposure to the expansion policy

We measure the exposure to higher education expansion, referred to as *intensity* for a given cohort, in a given region, by the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the number of 18-year-olds living in that region.⁸ The increase in public slots represents the increase in the level of supply shock in the higher education. In the definition of *intensity* the increase in slots is normalized by the population of 18-year-olds to control for the scale differences across regions. Thus, for example, for an individual in cohort 2010 (an individual who was 18 years old in 2010) living in Konya region, exposure intensity is measured as the increase in available public slots from 2005 (the year that the expansion policy was initiated) to 2010 in Konya region, divided by the number of 18-year-olds in 2010 in Konya region. Note that the intensity variable is zero for cohorts which belong to the pre-expansion period, similar to earlier studies in the literature (Duflo 2001, Berlinski et al. 2009).

Figure 2.4 depicts the geographical distribution of intensity in 2013 and provides evidence on the idiosyncratic nature of the expansion policy. The exposure intensity varies considerably -between 0.03 and 0.65- across regions. Figure A.1 in the Appendix is a more detailed demonstration of the evolution of the spatial distribution of intensity over time; it shows that intensity has significantly varied over time, as well as across regions. Some regions started with a much higher intensity than others in the early years of the expansion. There are several regions with low intensities at the beginning of the expansion, that had reached the highest intensities in the later years of the expansion. In other words, the expansion does not proceed at a uniform tempo across regions.

⁸We choose not to define *intensity* as a stock variable, since such a variable could be endogenous to our dependent variable, higher education attainment.

Figure 2.4: The spatial distribution of intensity in 2013.



Source: OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: Authors' calculations. Intensity (for cohort 2013), in a given region, is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the number of 18-year-olds living in that region.

2.5.3. Other explanatory variables

The region fixed effects, included in the equation (2.1), capture all time-invariant regional differences in higher education attainment. Therefore, we only include time-varying regional variables in the set of regional controls, X_{rt} , in equation (2.1). The first of those is the number of 18-year-olds in the region for each cohort—the denominator of the exposure intensity variable—in order to account for any variation in the intensity measure that originates from changes in demographics, such as a change in the population due to migration. In our heterogeneity analyses, we use regional categorical variables by gross domestic product (GDP) per capita, and the density of higher education institutions prior to expansion. In some specifications, public investment (in all industries except for education) and GDP per capita are added to further control for any regional variation that could increase higher education attainment besides the expansion policy. Another regional control, included in some specifications, is the increase in available slots per 18-year-old in private universities, referred to as *private intensity*, to capture the impact of possible demand-driven supply changes, relative to the supply increase in public universities, in the higher education. The set of individual controls, X_{irt} , includes survey year fixed effects, and in some specifications, we control for age and age squared instead of survey year fixed effects.

2.5.4. Sample

Our main sample includes the individuals who are expected to have started university education between 1999 and 2013 in the 2016-2017 waves of HLFS. There is no information on which year an individual graduated from high school or took the university entrance exam in the HLFS. Therefore, we assume that individuals start higher education at age 18 when students typically graduate from high school. Thus, we restrict our sample to individuals who were 18 years old in 1999-2013; simply called the cohorts of 1999-2013 in the rest of the paper. In addition, we exclude cohort 2006 from the post-expansion group, since the regional information on the number of 18-year-olds is not available for that particular year. Therefore, our main sample includes individuals in cohorts of 1999-2005 and 2007-2013, seven cohorts each in the pre-expansion and post-expansion groups. Table A.5 in the Appendix gives the descriptive statistics of our sample by higher

education attainment and gender. Among pre-expansion cohorts, 13,212 (28%) out of 47,478 men, and 12,214 (24%) out of 51,614 women attained higher education, and the numbers increased to 16,872 (39%) out of 43,279 for men and 17,189 (39%) out of 43,951 for women in the post-expansion period. Since we use the 2016 and 2017 waves of the HLFS, the age range of the individuals in our main sample is 21-36 (with the mean age around 32-33 for pre-expansion cohorts and 24-25 for post-expansion cohorts).

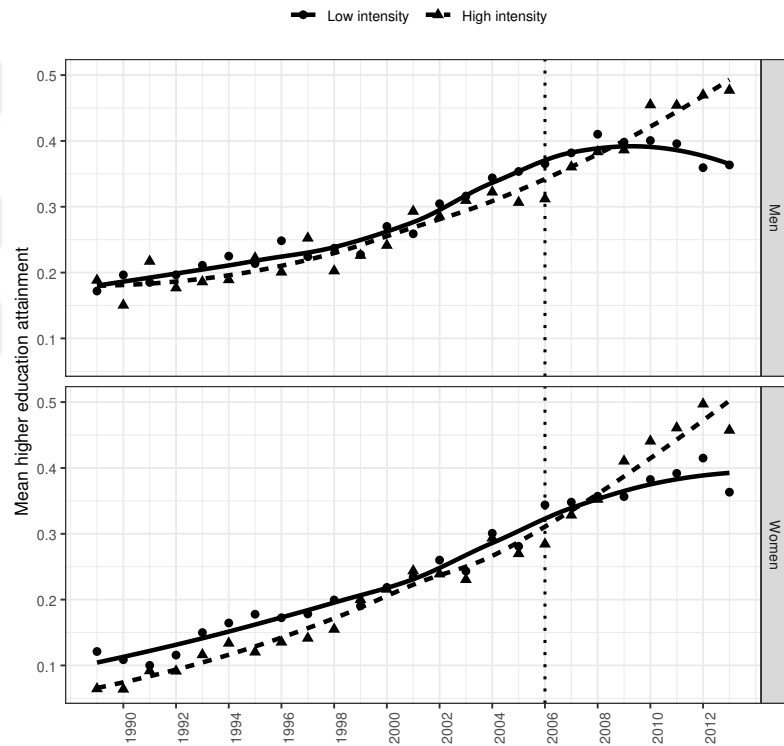
2.6. Identification

In this section, we first verify the validity of our identification strategy by providing evidence on each assumption embedded in the interpretation of the estimates from equation (2.1) as the causal effects of the reform. Then we experiment with two samples to show that an increase in educational attainment is observed in our sample of interest but not in a sample of individuals who are older and therefore not exposed to the expansion policy.

The crucial identification assumption embedded in these models is that the trends in enrollment among higher-intensity regions would have been the same as the trends in the comparison group of lower-intensity regions in the absence of the reform. In order to test the validity of this assumption, we check whether the high intensity and low intensity regions had parallel time-trends in educational attainment prior to the reform. We divide the regions into quartiles based on their exposure intensity (use the increase in slots from 2005 to 2013) and call the regions that are in the bottom and top quartiles as low and high intensity regions, respectively. We provide graphical evidence in Figure 2.5, to show that attainment rates in the pre-reform years (before 2006) were indeed moving together in the high and low-intensity regions. After 2006 a gap emerged between the two groups, in favor of high-intensity regions for both men and women. Additional graphical evidence on higher education attainment rates of high, low, and medium (those in the middle two quartiles in their ranking by exposure intensity) intensity regions determined by their intensity in the year 2009 and 2011, in addition to the year 2013 are given in Figure A.2 in the Appendix.

We further provide graphical evidence on the evolution of mean higher education attainment by cohorts separately for high and low intensity regions, in Figure 2.6.

Figure 2.5: The higher education attainment across cohorts: High vs. low-intensity regions.



Source: HLFS, 2016-2017, OSYM Almanacs and TurkStat Address Based Population Registry System.

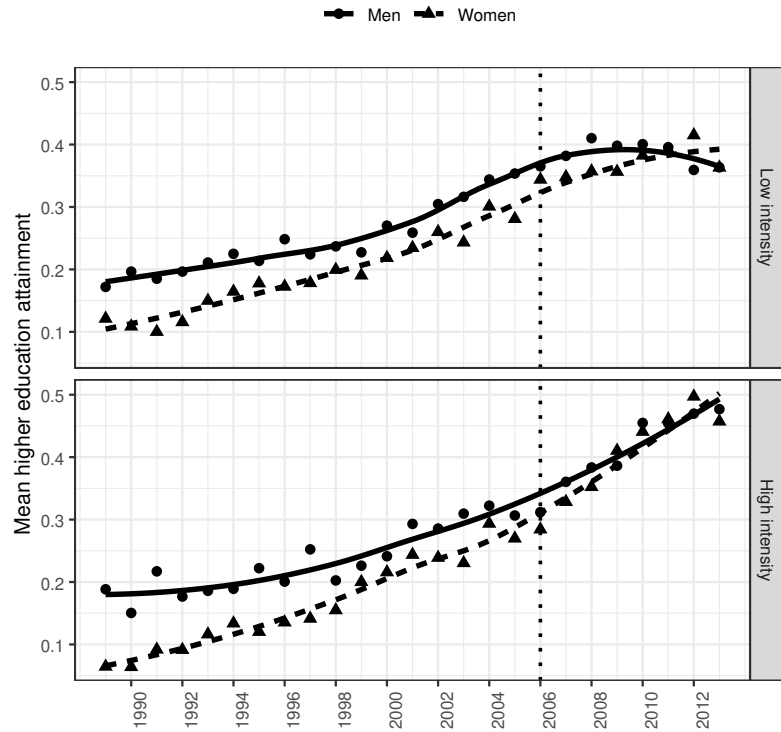
Notes: Authors' calculations. Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. High intensity and low intensity regions are defined as the ones whose increase in available slots per 18-year-old population from 2005 to 2013 is in the top and bottom quartile, respectively.

These figures intend to show that the gender gap in attainment has been narrowing over time across the country even before the expansion. The distance between two lines can be considered a simple measure of gender gap since the population is balanced in the two genders. The steeper trend for women residing in high intensity regions reduces the larger gender gap in these regions observed in the earlier cohorts. After the expansion, we note that gender differences vanish in the high intensity regions. Besides, women living in low intensity regions reach higher levels of educational attainment than men in the post-expansion period. Therefore, although both men and women living in highly-treated regions experience an increase in their educational attainment, we observe a greater reduction in the gender gap (eventually generating higher attainment rates for women than men) in regions treated with lower intensity. We take into account the gender differences in the time trend by including gender-specific region and cohort effects in our regressions. In addition, as a robustness check later in the paper, we relax the parallel trend assumption and allow cohort effects to vary across groups of regions.

In estimating the impact of a program, another problem that researchers face is identifying the effects of a compensatory intervention (Rosenzweig & Wolpin 1986) and the concern that the intervention may not be entirely exogenous to pre-existing trends in education. In order to address this concern, we introduce several region-specific indicators and examine the higher education outcomes and demand prior to the expansion. First, we calculate the university exam takers per 18-year-old, and next, the higher education graduation rates of different age groups (25-34 to capture young adults, and 25-44 and 25-64 to capture a more comprehensive picture for the working-age population) in 2005 HLFS (before the expansion) across regions. Next, we use the number of higher education institutions and their capacities, specifically the number of (public) universities and available (public) slots per 18-year-old in 2005, to examine whether the distribution of new educational opportunities is associated with the initial conditions. Finally, we use the gross domestic product (GDP per capita), labor productivity, and labor productivity in the industry (by excluding the product in agriculture and service sectors) to capture the relation between the expansion and the initial economic and labor market conditions across regions.⁹ We find that there is no statistically significant relationship between our predictors and intensity variables, suggesting that the initial (pre-reform) conditions do not violate our exogeneity assumption

⁹The information on the 18-year-old population available since 2007, hence, we fill in the missing

Figure 2.6: The gender gap in higher education attainment across cohorts: High vs. low intensity regions



Source: HLFS, 2016-2017, OSYM Almanacs and TurkStat Address Based Population Registry System.

Notes: Authors' calculations. Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. High intensity and low intensity regions are defined as the ones whose increase in available slots per 18-year-old population from 2005 to 2013 is in the top and bottom quartile, respectively.

(see Table A.6 in the Appendix).

Having performed these tests on identification, we move on to establish that higher education attainment increased as a result of the expansion. We estimate equation (2.1) for men and women separately.

Table 2.3: The effect of expansion policy on educational attainment

<i>Dependent variable: Higher education attainment</i>						
	Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)
<i>Panel A: Experiment of interest</i>						
Intensity	0.393*** (0.124)	0.364*** (0.125)	0.266** (0.120)	0.271*** (0.097)	0.260** (0.102)	0.254* (0.136)
Observations	90,757	77,050	63,469	95,565	81,266	67,393
<i>Panel B: Placebo experiment</i>						
Intensity	-0.003 (0.047)	0.027 (0.066)	-0.023 (0.083)	0.115 (0.071)	0.114 (0.077)	0.135 (0.086)
Observations	96,750	83,973	70,701	104,073	90,207	75,761
Cohort window	±7 cohorts	±6 cohorts	±5 cohorts	±7 cohorts	±6 cohorts	±5 cohorts

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. A constant, number of 18-year-olds, region, cohort and survey year fixed effects are included as control variables. The standard errors clustered at region level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

Panel A in Table 2.3 presents our results for several cohort windows around the policy cutoff 2006. Our largest sample includes ±7 cohorts around the policy cutoff 2006 and the narrowest sample includes ±5 cohorts. The coefficient estimates of *intensity* are statistically significant for both men and women in all samples. Our results in columns (1)-(6) suggest that a one-unit increase in the intensity variable (an additional slot in a public university per 18-year-old) raises higher education attainment by 27-39 and 25-27 percentage points for men and women, respectively. These estimates are both statistically and economically significant.

We also conduct a placebo test to explore the causal interpretation of our results against possible omitted time-varying region-level factors observed before the expansion. We test whether the intensity is correlated with the higher education

population data of 18-year-olds in 2005 by assuming that the population growth rate among 18-year-olds in 2005-2007 was equal to the population growth rate for the population of all ages, i.e.,

$$Pop18_{2005} = Pop18_{2007} \left(\frac{Pop_{2005}}{Pop_{2007}} \right)$$

where Pop_t and $Pop18_t$ represents the population of all ages and 18-year-olds in year t, respectively.

attainment of older cohorts who were not exposed to the reform. The lack of an association between intensity and higher education attainment of cohorts unaffected by the reform would support our causal interpretation of intensity on higher education attainment of younger cohorts exposed to the policy. However, a positive association of increase in supply with educational attainment of older cohorts would suggest that the estimated effects could be (at least partially) driven by changes in other region-level factors that were correlated with the policy. Hence, to validate our identification strategy, we need positive and statistically significant correlation between intensity and higher education attainment in our main sample, but not in a sample of cohorts that belong to the pre-reform period.

To implement the placebo test we introduce the new sample of cohorts 1989-2003 from HLFs 2016-2017. This sample includes the older cohorts in ages 31-46, and these cohorts could not have been exposed to the expansion policy.¹⁰ We add 10 years to their cohort variable to create false cohort variables. Next, we assign the intensity of the expansion and the policy cutoff based on these false cohorts and estimate the link between these assigned variables and the higher education attainment of these older individuals. To construct a sample that is symmetric to our main one, we exclude cohort 1996 (2006 according to the false cohort). Panel B of Table 2.3 presents the results of our placebo test. We find very small coefficient estimates of intensity for men, which are statistically zero. For women, we cannot reject the null hypothesis that the correlation between the intensity and higher education attainment of cohorts 1989-2003 is statistically zero. The larger coefficient estimates for women than men support our observation that the gender gap had been narrowing prior to expansion, especially in high intensity regions. However, since the coefficient estimates for intensity in the placebo experiment are statistically insignificant after controlling for region and cohort effects, we can conclude that the parallel trend assumption holds for both men and women. All evidence supports that our findings in Panel A of Table 2.3 are not due to the time and region trends that existed prior to the expansion.

In our identification strategy, an individual's exposure level to the expansion in

¹⁰The youngest cohort we use in the placebo test is 2003. Individuals in 2004-2005 cohorts could be partially exposed to the expansion policy by retaking the university exam in the post expansion years. Krishna et al. (2018) state that in Turkey retaking the university exam decreases with the number of attempts in the university exam. Indeed, 49% of university exam takers are first-time exam takers, whereas the students who make their second and third attempts in the exam accounted for 25% and 13% of all students, respectively.

higher education is determined by the individual's cohort and current region of residence. Ideally, we would measure the level of exposure to the policy by the region where the individual is living when he/she makes the educational attainment decision. However, HLFS does not have information on the region of residence when the individual is 18 year olds; it instead provides information on the current region of residence. Due to the lack of information, we assume that individuals' current region of residence is the same as the region of residence when they made their higher education decision. We can check the validity of this assumption to some extent. The HLFS asks individuals whether they have been living in their current city of residence since they were 16 years old. We observe that in our main sample, 72% of men and 67% of women are in the same city as when they were 16 years old (i.e., before making their higher education decision). Since there are several cities in a given region and some migrations take place within the region, the actual share of individuals living in their current region of residence can be even higher than these estimates show. In order to further address concerns about selection to migration, we examine whether the probability of migration since age 16 (before the higher education decision is made) is correlated with the level of exposure, i.e., the migration probability is orthogonal to treatment (see Table A.7 in the Appendix). In this table, dependent variable is 1 if individual has migrated since age 16 and 0 otherwise. In addition to exposure intensity, we control for number of 18-year-olds, region, cohort and survey year fixed effects as control variables and estimate the linear probability model for men and women separately for several different windows of cohorts. In all regressions, intensity measure is statistically insignificant. Hence, the migration decision does not seem to be correlated with exposure intensity. Since we control for observable factors, we will argue that unobservables that may affect migration are also uncorrelated with exposure intensity. Therefore, in our estimation of equation (2.1), we suppose that if there are any unobservables that affect both migration decision and higher education decision, these variables are orthogonal to our key explanatory variable, exposure intensity.

2.7. The expansion policy on the gender gap in higher education attainment

2.7.1. The effect of expansion policy by gender

Before the expansion, women were disadvantaged in higher education attainment compared to men. The average higher education attainment among pre-expansion cohorts of our main sample was 29% for men and 24% for women. In this section, we investigate the effect of the expansion policy on the gender gap in higher education attainment by comparing the differential effects of the reform for men and women. In Table 2.4, we present the estimation results of equation (2.1) for the full sample where we pool the samples for men and women. The equation includes a constant, the number of 18-year-olds, region, and cohort, and survey-year fixed effects. Moreover, we control for the interaction of all variables with the woman dummy to account for gender-varying observable factors aside from the expansion policy.

In panel A, the results presented in columns (1), (3), and (5), obtained by estimating equation (2.1) on the samples of cohorts including ± 7 , 6, and 5 cohorts around the policy cutoff, are parallel to the earlier results in Table 2.3. An increase in the public slots per 18-year-old in a region has increased educational attainment in that region by 26-33 percentage points on average.

The results presented in columns (2), (4) and (6), obtained by estimating equation (1) by controlling for the interaction term of intensity with woman dummy, show us whether the effects of the increase in public slots per 18-year-old vary across genders. Our results suggest that the effect was statistically the same for men and women (the 95% confidence interval (CI) for the *Intensity* $\times$ *Woman* term includes zero for all samples and the p-values of the tests for the coefficients being equal to zero are 0.182 in the largest sample and 0.942 in the narrowest sample). Hence, the increase in public slots at the regional level did not translate into a reduction in the gender gap in the region. In Panel B of the table, we repeat the placebo experiment of Section 6, but on the entire sample (as expected, the coefficient estimates for men and women are the same as those in Table 2.3 Panel B, since all control variables in the regression equation are gender-specific). We confirm that our findings on the gender gap in attainment are not due to the time

and region trends that existed prior to the expansion.

Table 2.4: The effect of expansion policy on gender gap

<i>Dependent variable: Higher education attainment</i>						
	(1)	(2)	(3)	(4)	(5)	(6)
<i>Panel A: Experiment of interest</i>						
Intensity	0.332*** (0.102) [0.122,0.542]	0.393*** (0.124) [0.138,0.647]	0.312*** (0.100) [0.106,0.518]	0.364*** (0.125) [0.105,0.622]	0.260** (0.103) [0.0487,0.471]	0.266** (0.120) [0.0187,0.513]
Intensity × Woman		-0.122 (0.089) [-0.305,0.061]		-0.103 (0.111) [-0.331,0.125]		-0.011 (0.156) [-0.332,0.309]
Intensity + Intensity × Woman		0.271*** (0.097) [0.072,0.470]		0.260** (0.102) [0.051,0.470]		0.254* (0.136) [-0.027,0.535]
Observations	186,322	186,322	158,316	158,316	130,862	130,862
<i>Panel B: Placebo experiment</i>						
Intensity	0.059 (0.041) [-0.024,0.143]	-0.003 (0.047) [-0.101,0.094]	0.073 (0.053) [-0.036,0.182]	0.027 (0.066) [-0.108,0.163]	0.059 (0.061) [-0.066,0.184]	-0.023 (0.083) [-0.194,0.149]
Intensity × Woman		0.118 (0.090) [-0.031,0.260]		0.087 (0.097) [-0.044,0.273]		0.158 (0.118) [-0.042,0.312]
Intensity + Intensity × Woman		0.115 (0.071) [-0.0678,0.304]		0.114 (0.077) [-0.114,0.288]		0.135 (0.086) [-0.0858,0.401]
Observations	200,823	200,823	174,180	174,180	146,462	146,462
Cohort window	±7 cohorts	±7 cohorts	±6 cohorts	±6 cohorts	±5 cohorts	±5 cohorts

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. A constant, number of 18-year-olds, region, cohort and survey year fixed effects and the interaction of all variables with woman dummy are included as control variables. The standard errors clustered at region level are given in parentheses. 95% confidence intervals are given in square brackets. *p<0.1 **p<0.05 ***p<0.01.

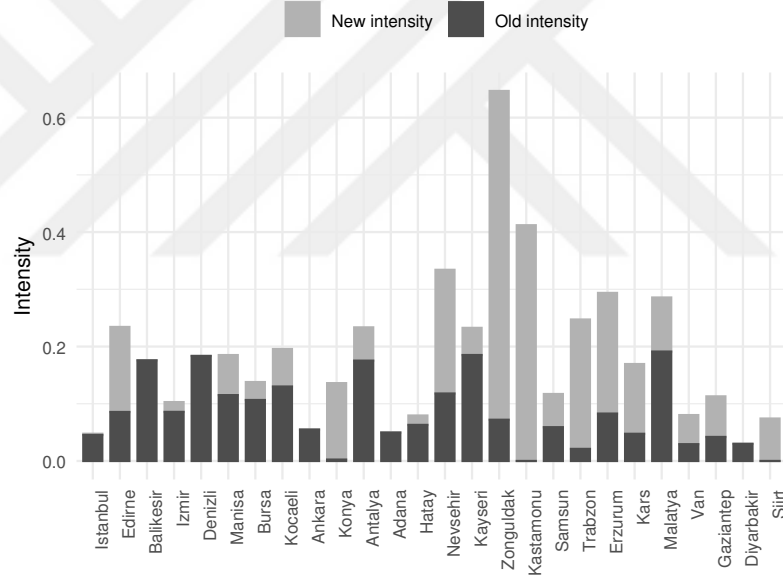
The estimates presented in Panel A indicate that the reform did not reduce the gender gap in attainment. Such a result could be possible if the additional slots were not used more by women than by men. In the next section, we will examine why the treatment effect did not vary across genders by analyzing how factors outlined in the conceptual framework section contributed to our results.

2.7.2. Potential channels

In the conceptual framework section, we developed a general framework to understand how the characteristics of the expansion will interact with gender differences in higher education attainment. Specifically, college proximity may have differential effects across genders due to larger psychic costs and lower perceived returns to higher education for women and majors supplied in the expansion may affect the two genders differently due to gender differences for college majors.

In order to understand the proximity channel we decompose the intensity variable into two components as *old intensity* and *new (local) intensity*, defined as the increase in public slots introduced in the universities established before and after 2005, divided by the number of 18-year-olds, respectively. Figure 2.7 shows the regional variation in these variables in 2013. The graph clearly shows that there is large regional variation in the increase of these slots. In order to examine whether exposure intensity in existing versus new (local) universities had different effects across genders we replace intensity with new intensity and old intensity variables in equation (2.1), and estimate their impacts on higher education attainment separately.

Figure 2.7: The decomposition of intensity in 2013 across old and new universities.



Source: OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: Authors' calculations. The dark and light gray bars show the old and new intensity levels in a region. Old intensity is defined as the increase in available slots in already existing (before the expansion policy) public universities from 2005 to 2013, divided by the 18-year-old population. New intensity is the number of available slots in 2013 in public universities established after 2005, divided by the 18-year-old population. Intensity is the sum of old intensity and new intensity in a region.

If college proximity lowers psychic costs incurred by parents from sending daughters to college, then we would expect new (local) intensity to have a higher impact on the educational attainment of women compared to men and hence help narrow the gender gap. However, old intensity should not necessarily have a differential effect across genders.

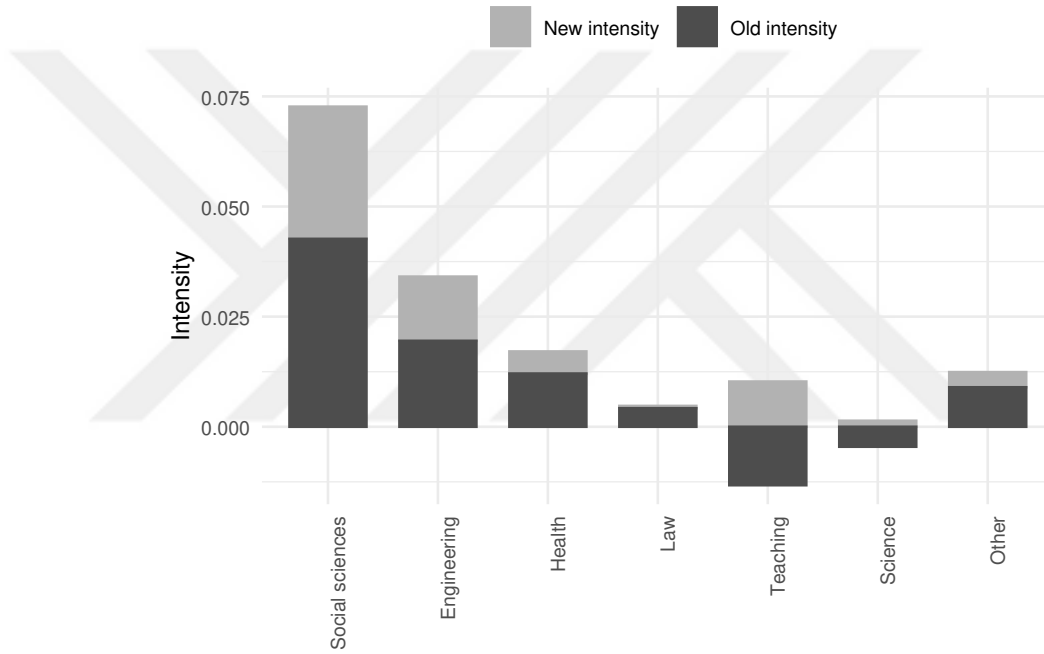
Our conceptual framework also emphasizes that preferences for majors vary across genders. Indeed, in the pre-expansion period, engineering graduates were mostly men, whereas social science majors were more gender-balanced. In order to examine this mechanism through which the expansion may affect men and women differently, we further decompose the exposure intensity variable across fields of study and present the decomposition of intensity in 2013 in Figure 2.8. We observe that there are significant differences in the magnitude of the expansion across fields of study. By far, the highest increase is in the field of social sciences. The second highest increase is in the field of engineering, followed by much lower increase in the field of health. In order to examine how expansion in social sciences and engineering, the two fields that constitute 80% of the increase in public slots, may have affected the gender gap in higher education, we decompose the intensity variables into social science, engineering, and other majors. We examine how the differences in the scale of expansion across majors affected the gender differences in attainment by estimating equation (2.1) with the decomposition of intensity instead of controlling for a single aggregate intensity variable.

If the preferences for college majors channel is at work, then we would expect the higher intensity in engineering to be associated with a wider gender gap. On the other hand, higher intensity in social sciences would affect the two genders similarly and keep the gender gap unchanged.

Next, we present our results on how the two possible aspects of gender inequality in higher education, namely college proximity and college majors, have contributed to the evolution of the gender gap in response to the reform.

We examine the college proximity hypothesis by testing whether the expansion in public slots in newly established universities and existing universities have a differential impact on women's higher education attainment. Column (1) of Table 2.5 shows the separate effects of the two intensities on average attainment in our main sample. A one-unit increase in old and new intensity increases the probability of higher education attainment by 41 and 30 percentage points, respectively, after controlling for region and cohort fixed effects. The coefficient estimates for old and new intensity are statistically the same. Next, in column (2) of the table, we estimate the heterogeneous effects of expansion policy by including the interactions of old and new intensity with woman dummy. We find that the interactions of old and new intensity with woman dummy have negative coefficients; the former is

Figure 2.8: The decomposition of intensity in 2013 across college majors.



Source: OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: Authors' calculations. The dark and light gray bars show the old and new intensity levels of a major. Old intensity is defined as the increase in available slots in already existing (before the expansion policy) public universities from 2005 to 2013, divided by the 18-year-old population. New intensity is the number of available slots in 2013 in public universities established after 2005, divided by the 18-year-old population. Negative values of old intensity indicate a decrease in available slots. Intensity is the sum of old intensity and new intensity for each major.

imprecisely estimated and the latter is only weakly significant. 95% CI, also shown in the table, include zero. Moreover, the p-values of the tests for the coefficients for the interaction term being equal to zero are 0.753 for the old intensity and 0.095 for the new intensity. Therefore, the evidence do not support the hypothesis that the newly established universities have a higher positive impact on women than on men. In fact, on the contrary, we find a somewhat perplexing result that intensity in new universities has benefited men more than women.

One possible explanation for this perplexing result can be found by examining more closely how slots in new and existing public universities are distributed across majors. Caner & Okten (2010) show that there are important differences between men and women’s career choices in the Turkish context, by using a large micro data set that includes university applicants’ socio-demographic characteristics, major preferences and university entrance exam scores. We observe gender differences in college major choices in our dataset as well (Figure 2.3). Hence, we investigate how the expansion in slots across different majors affect the gender differences by decomposing each intensity variable into three categories by majors: old and new intensity in engineering, in social sciences, and in other majors.

Our results, presented in column (3) of Table 2.5 suggest that additional slots in social sciences in the existing universities and additional slots in engineering in the new universities significantly increased educational attainment. Column (4) shows that the coefficient estimate of new intensity in engineering is positive and significant, whereas the interaction of new intensity in engineering and woman dummy is negative and statistically significant. These findings suggest that new intensity in engineering increases men’s education but not women’s, whereas the effect of old intensity in social sciences is statistically the same across genders. These results are consistent with our expectations: Men are more likely than women to prefer majoring in engineering; an increase in slots in engineering benefits men more than it does women. An increase in slots in social sciences benefits men and women equally, since these majors are preferred by both men and women. Our results show that the variation in the scale of expansion across majors has important consequences for higher education attainment and the gender gap. In the overall, the finding that the effect of preferences across majors is larger than the effect of college proximity suggests that policies promoting higher education attainment should take the heterogeneity in individual preferences into account.

Table 2.5: Decomposition of policy intensity on gender gap

	<i>Dependent variable: Higher education attainment</i>				
	(1)	(2)	Women's coef. by (2)	(3)	(4)
					Women's coef. by (4)
Old intensity	0.411** (0.154) [0.094,0.728]	0.439** (0.196) [0.035,0.843]	0.384** (0.156) [0.063,0.705]		
Old intensity×Woman		-0.055 (0.173) [-0.412,0.301]			
New intensity	0.304*** (0.103) [0.091,0.516]	0.377*** (0.130) [0.110,0.644]	0.228** (0.086) [0.045,0.411]		
New intensity×Woman		-0.149* (0.086) [-0.326,0.028]			
Old intensity in engineering				0.010 (0.385) [-0.783,0.802]	0.011 (0.479) [-0.976,0.998]
Old intensity in engineering×Woman					0.005 (0.390) [-0.798,0.808]
New intensity in engineering				0.432** (0.168) [0.0875,0.777]	0.690*** (0.245) [0.185,1.196]
New intensity in engineering×Woman					-0.542* (0.265) [-1.088,0.004]
Old intensity in social sciences				0.956*** (0.194) [0.555,1.356]	0.969*** (0.269) [0.415,1.524]
Old intensity in social sciences×Woman					-0.029 (0.237) [-0.517,0.459]
New intensity in social sciences				0.434 (0.317) [-0.220,1.087]	0.436 (0.419) [-0.427,1.299]
New intensity in social sciences×Woman					0.003 (0.399) [-0.819,0.826]
Old intensity in other				-0.154 (0.202) [-0.571,0.262]	-0.158 (0.288) [-0.752,0.436]
Old intensity in other×Woman					0.007 (0.400) [-0.818,0.832]
New intensity in other				-0.020 (0.161) [-0.351,0.311]	-0.026 (0.231) [-0.503,0.450]
New intensity in other×Woman					0.009 (0.311) [-0.631,0.649]
Observations	186,322	186,322	186,322	186,322	186,322

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Old intensity (for a given major) is defined as the increase in available slots in public universities that existed before the expansion relative to the number of slots in the same universities in 2005, divided by 18-year-old population. New intensity (for a given major) is defined as the increase in available slots in public and newly established universities with the expansion policy, divided by 18-year-old population. The column titled Women's coef. by (2) presents the estimated coefficients for women based on the estimates in column (2). Similarly, Women's coef. by (4) presents the estimated coefficients for women based on the estimates in column (4). Sample includes cohorts 1999-2005 and 2007-2013. A constant, number of 18-year-olds, region, cohort and survey year fixed effects and the interaction of all variables with woman dummy are included as control variables. The standard errors clustered at region level are given in parentheses. 95% confidence intervals are given in square brackets. *p<0.1 **p<0.05 ***p<0.01. *p<0.1 **p<0.05 ***p<0.01.

2.8. Heterogeneity analysis

In this section, we conduct two heterogeneity analyses. First, we analyze the impact of expansion by the density of higher education institutions in the pre-expansion period to examine the proximity channel from a different perspective. Next, we investigate the heterogeneous impact of expansion by the regional income in the pre-reform period to explore how the expansion has contributed to the regional convergence.

2.8.1. Heterogeneous effects by the pre-reform density of higher education institutions

To reveal the possible channels of effect, we divide the total intensity variable into intensity in old (existing) versus new universities with the aim to separate out the additional educational opportunities that significantly reduce the distance to college. This approach might overlook two concerns: First, high school students living in the provinces with at least one university before the expansion may have an advantage in human capital accumulation (i.e., may have higher anticipation of attainment and have higher motivation) when they take the entrance exam, hence they may have benefitted more from the additional slots. As explained in the conceptual framework, reducing distance to university is expected to increase educational attainment more for women than for men. However, the potential university students live in a province with no university may be less prepared for higher education in terms of their human capital accumulation and aspiration compared to others who live in a province with a university. In that case, density of existing universities prior to the expansion may be an important confounding factor that affects our findings on new universities. Second, the establishment of new universities might induce demand for college in both old and new universities (by revealing information on returns to college or eliminating the social and cultural barriers to education). Therefore, the reform effect in a region may be related to the pre-reform density of higher education institutions in the region.

To distinguish the regions by pre-reform density of higher education institutions, we define a binary variable *one or more provinces without university*, that is 1 for regions where at least one province was without a university before the expansion

and 0 for regions where every province had a university. We added the interaction of our variable of interest, intensity, with this newly defined variable to our baseline analysis to investigate heterogeneous effects. The estimates are given in Table 2.6. We find that the impact of the reform is higher for men in the regions that had a university in all of its provinces (i.e., when the *one or more provinces without university* dummy is 0) prior to the expansion. We also find that the gender gap increases in these regions —although statistically significant only in our widest sample (column 2). The three-way interaction term of intensity with woman and one or more provinces without university is positive but statistically insignificant. Overall, these results are consistent with our finding that the net effect of the expansion was not to reduce gender differences in attainment, as other mechanisms discussed in the conceptual framework section were also at play.

Table 2.6: The effect of expansion policy on gender gap: Heterogeneous impacts by the pre-reform density of higher education institutions

	<i>Dependent variable: Higher education attainment</i>					
	(1)	(2)	(3)	(4)	(5)	(6)
Intensity	0.633*** (0.162)	0.791*** (0.180)	0.587*** (0.156)	0.706*** (0.168)	0.496*** (0.156)	0.586*** (0.165)
Intensity×One or more province without university	-0.282*** (0.094)	-0.372*** (0.108)	-0.258** (0.102)	-0.321*** (0.115)	-0.224** (0.105)	-0.303** (0.109)
Intensity×Woman		-0.310** (0.139)		-0.234 (0.165)		-0.173 (0.216)
Intensity×Woman×One or more province without university		0.175 (0.104)		0.121 (0.116)		0.152 (0.143)
Observations	186,322	186,322	158,316	158,316	130,862	130,862
Cohort window	±7 cohorts	±7 cohorts	±6 cohorts	±6 cohorts	±5 cohorts	±5 cohorts

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. One or more provinces without a university is a time-invariant binary variable that is 1 for regions where at least one province was without a university before the expansion. A constant, number of 18-year-olds, region, cohort and survey year fixed effects and the interaction of all variables with woman dummy are included as control variables. The standard errors clustered at region level are given in parentheses.*p<0.1 **p<0.05 ***p<0.01.

2.8.2. Heterogeneous effects by regional income prior to expansion

Turkey is a country with high inequality of income and large variation in development across its regions. As the reform reduced the distance to higher education and thereby the cost of education, it is noteworthy to examine whether the reform has increased student access to education in poorer regions. Hence, in this section, we investigate the heterogeneity in the effect of intensity variable across richer and poorer regions of the country. We include the interaction of our variable of interest, intensity, with four dummy variables indicating the rank of the

region by GDP per capita in 2005 (prior to expansion policy). Quartile 1 equals 1 if a region is in the first quartile (bottom 25 percent) of the GDP distribution and 0 otherwise, and similarly for Quartile 2 (second quartile), Quartile 3 (third quartile), and Quartile 4 (top quartile). Instead of using a continuous GDP variable we choose to use indicator variables to avoid endogeneity of GDP of a region to college attainment, since there may be a two-directional relation between college attainment and GDP per capita. Although we confirm the orthogonality of GDP per capita in 2005 to our measure of intensity (to avoid endogeneity), unobserved factors may still lead to an endogeneity problem by affecting both college attainment and GDP per capita. A dummy variable approach by defining quartiles is less prone to this problem, since a region is less likely to switch quartiles in such a short period as a result of an increase in college attainment.

The higher education attainment rates in the quartiles were as follows for men and women, respectively: In the first quartile 23.1% and 13.8%, in the second 27.2% and 22.8%, in the third 27.0% and 23.1%, and in the top quartile 33.2% and 32.7%. These averages document that in the poorer regions, educational attainment is lower and gender gap is wider.

Our results given in Table 2.7 show that, in the lowest quartile, the policy was not effective in either increasing educational attainment or reducing the gender gap. In the top quartile, where attainment rates were high for both genders, the policy was not effective either. The educational attainment rates were similar in the middle (second and third) quartiles. In these regions, the policy increased educational attainment of both men and women with lower coefficients for women, thereby slightly increasing the gender gap while enabling the middle-income regions to catch up with the regions in the top quartile.

2.9. Robustness Checks

Turkey experienced two other changes in education policy that overlapped with the expansion policy in higher education. In this section, we show that these policies alone cannot explain the observed changes in higher education attainment. We also show that our results are robust to restricting the sample to natives; to those who live in smaller cities, and to using alternative specifications of the regression equation. We further check whether our results are robust to including

Table 2.7: The effect of expansion policy on gender gap: Heterogeneous impacts by the regional wealth prior to the expansion

<i>Dependent variable: Higher education attainment</i>						
	(1)	(2)	(3)	(4)	(5)	(6)
Intensity×Quantile 1	0.258 (0.167)	0.210 (0.179)	0.271* (0.148)	0.204 (0.135)	0.248 (0.158)	0.105 (0.104)
Woman×Intensity×Quantile 1		0.088 (0.089)		0.123 (0.113)		0.277 (0.225)
Intensity×Quantile 2	0.329*** (0.100)	0.399*** (0.123)	0.308*** (0.103)	0.373*** (0.130)	0.251** (0.106)	0.285** (0.127)
Woman×Intensity×Quantile 2		-0.143* (0.073)		-0.131 (0.090)		-0.066 (0.113)
Intensity×Quantile 3	0.481*** (0.150)	0.507*** (0.171)	0.441*** (0.145)	0.454*** (0.158)	0.381** (0.139)	0.321** (0.153)
Woman×Intensity×Quantile 3		-0.054 (0.088)		-0.026 (0.100)		0.117 (0.134)
Intensity×Quantile 4	0.363* (0.186)	0.342 (0.209)	0.264 (0.186)	0.271 (0.178)	0.229 (0.172)	0.204 (0.162)
Woman×Intensity×Quantile 4		0.041 (0.145)		-0.016 (0.218)		0.048 (0.279)
Observations	186,322	186,322	158,316	158,316	130,862	130,862
Cohort window	±7 cohorts	±7 cohorts	±6 cohorts	±6 cohorts	±5 cohorts	±5 cohorts

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. Quartile dummies are time-invariant binary variables indicating the region's quartile in the ranking of regional GDP per capita in 2005 (before expansion). A constant, number of 18-year-olds, region, cohort and survey year fixed effects and the interaction of all variables with woman dummy are included as control variables. The standard errors clustered at region level are given in parentheses.*p<0.1 **p<0.05 ***p<0.01.

time-varying regional covariates and the supply increase in private universities. Finally, we check the robustness of our findings with alternative estimation methods.

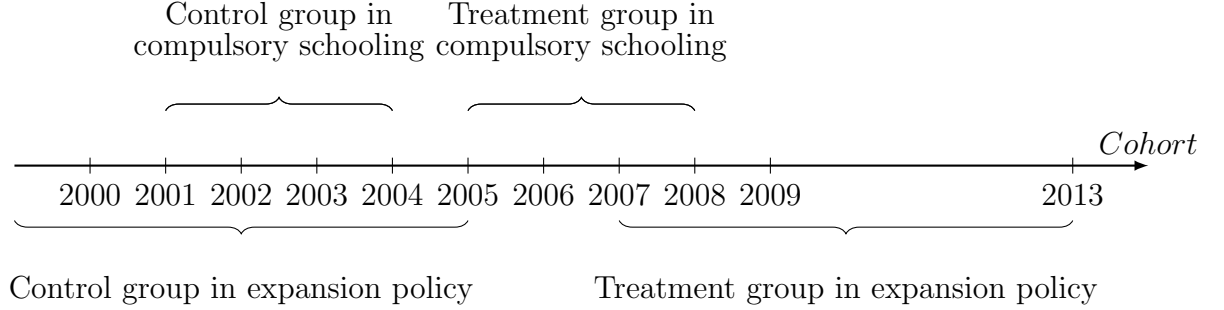
2.9.1. Compulsory schooling

Prior to 1997, the education system in Turkey consisted of five years of compulsory primary education, followed by three years of lower secondary and three years of upper secondary (high school) education. In 1997, the lower tiers were merged, and compulsory schooling was extended from five to eight years. The new policy covered all children who did not already hold a primary school diploma at the beginning of the 1997-98 school year. Considering that children typically start school at age 6, we can say that children born in or after January 1987 were affected by the policy. In our sample, this corresponds to the cohorts of 2005 and later. The treatment cohorts in our analyses (2007-2013 cohorts) coincide with those who were exposed to compulsory schooling extension. In this section, we address the concern that our earlier results may have been affected by the 1997 reform. We show that the compulsory schooling law alone cannot explain the results we present earlier.

First of all, since the compulsory schooling policy was implemented at the country level, we do not expect it to influence our estimates of the effect of *intensity*, which varies by region and cohort, in a regression where we control for region, cohort, and survey year fixed effects. Even so, we would want to know whether the 1997 reform resulted in a structural increase in higher education attainment even among those who were not treated by the higher education expansion. For this purpose, we build a sample of individuals who were affected by the 1997 reform but not by the higher education expansion. Hence, we restrict attention to cohorts earlier than 2009.

We know that 2004 and earlier cohorts were not exposed to the 1997 reform (2.9), so they can be used as the control group for the 1997 reform. 2005 and later cohorts are the treatment group for the 1997 reform. The only cohort exposed to the 1997 reform but not to the expansion policy is the 2005 cohort. Moreover, 2006-2008 cohorts were marginally affected by the expansion policy; therefore, they can also be useful in our analysis.

Figure 2.9: The sample by cohorts and exposure to education policies.



We experiment with two sub-samples: 2002-2007 cohorts (which includes three control and three treatment cohorts) and 2001-2008 cohorts (which includes four control and four treatment cohorts). In these estimations, we control for the first-order polynomial of time trend (and not higher orders to avoid overfitting), in addition to region and survey year fixed effects. In Table A.8 in the Appendix, we show that the 1997 compulsory schooling policy did not increase the probability of higher education attainment of men or women. Evidence in the literature supports our results. Using earlier waves of the HLFS, Aydemir & Kirdar (2017a) find that the 1997 reform increased the years of schooling, but not the probability of higher education attainment. We confirm their results in our sample. Therefore, we conclude that the 1997 reform in compulsory schooling cannot by itself be generating our findings.

2.9.2. Relaxation of the headscarf ban

Before 2008, women who covered their heads were banned (known as the *headscarf ban*) from attending universities in Turkey. In 2008, the practice of headscarf ban was relaxed in higher education. This policy change overlaps with the expansion in the higher education system, and it might have positively affected the demand for higher education among women who wore a headscarf. However, this policy alone cannot explain our results, since we find a positive and significant effect of the expansion policy on men, who should not be affected by the headscarf ban. Here, we additionally examine whether there could be regional differences in the impact of the relaxation of the headscarf ban due to regional variations in the social norms, a crucial point on our claims for the identification of the causal effect of the expansion. In our baseline analysis given in Table 2.4, we find that the expansion

did not reduce the gender gap. If the headscarf ban liberated more women in the higher exposure regions, then our finding would be even stronger, indicating that the gender gap did not close even with the additional incentives for women. On the other hand, if the relaxation of headscarf ban liberated more women in low exposure regions, then our findings can be confounded with the effects of headscarf policy. To eliminate such a concern, we show that our intensity measure of the level of exposure to the reform is orthogonal to several measures expected to be correlated with the extent a region would be affected by the relaxation of the headscarf ban using KONDA Barometer survey on gender views. Our choices on the measure of exposure to headscarf policy are the share of women wearing headscarf, the share of individuals who declare themselves as conservative when they are asked about their lifestyle, and the share of individuals who self-report themselves as religious in response to a question on their beliefs. In Table A.9 presented in Appendix, we present our results on the correlation between the level of exposure to headscarf policy by the given three measures and the intensity in years between 2007 and 2013. Statistically insignificant coefficients for all regressions confirm that our results are not due to this policy change that happened in about the same years as the expansion. In addition, we would like to note that we find differential effects of the intensity variable across majors on the higher education attainment of men and women: An increase in engineering slots benefited men more than it did women. Relaxation of the headscarf ban cannot explain these results that seem to be driven by students preference for majors.

2.9.3. Relaxing the parallel trend assumption

In our baseline analysis given in equation (2.1), we assume that the evolutions of higher education attainment in the higher and lower intensity regions are similar. Graphical evidence presented in Figure 2.5 visually supports this assumption separately for men and women. To investigate whether the visual evidence is backed by regression analysis, we present results where we relax the parallel trend assumption across regions by allowing cohort effects to vary across regions¹¹, as in the seminal paper Stephens Jr & Yang (2014). The results, presented in Table 2.8, show that with these additional controls for region specific cohort effects, the effect

¹¹In the choice of regional grouping, we follow Aksu et al. (2018a), and divide the country in five regions as done in the Turkish Demographic and Health Surveys.

of the expansion policy is estimated to be lower in magnitude for both men and women. While the impact on men is strongly significant, the statistical significance vanishes for women in all samples of different cohort windows. Hence, our finding that the expansion did not reduce the gender gap is still valid with the additional controls that relax the parallel trend assumption.

Table 2.8: The effect of expansion policy: Heterogeneous cohort effects across five regions

<i>Dependent variable: Higher education attainment</i>						
	(1)	(2)	(3)	(4)	(5)	(6)
Intensity	0.252** (0.093)	0.291*** (0.093)	0.236** (0.097)	0.270*** (0.094)	0.215* (0.110)	0.202** (0.082)
Intensity×Woman		-0.079 (0.080)		-0.068 (0.120)		0.024 (0.187)
Intensity + Intensity×Woman		0.213 (0.109)		0.205 (0.131)		0.226 (0.184)
Observations	186,322	186,322	158,316	158,316	130,862	130,862
Cohort window	±7 cohorts	±7 cohorts	±6 cohorts	±6 cohorts	±5 cohorts	±5 cohorts

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. A constant, number of 18-year-olds, region, cohort and survey year fixed effects and the interaction of all variables with woman dummy are included as control variables. The five regions of residence presented here are West (NUTS-1 regions 1 to 4), Central (NUTS-1 regions 5 and 7), South (NUTS-1 region 6), North (NUTS-1 regions 8 and 9) and East (NUTS-1 regions 10 to 12). This is also the classification used in the Demographic and Health Surveys of Turkey. The standard errors clustered at region level are given in parentheses.*p<0.1 **p<0.05 ***p<0.01.

2.9.4. Specification and sample checks

Here, we investigate whether our results are robust to restricting the sample to sub-samples where the gender gap in educational attainment might differ from the full sample and whether alternative specifications change the qualitative power of our results.

In our analyses, the *intensity* measure is based on the available slots in the region of residence (at the time of the survey) in the year that the individual was 18 years old. Ideally, we would like to know the region where the individual resided at age 18 to measure exposure to policy change. What is available in the HLFS is how long the individual has been living in her current city. We argue that the lack of region information at age 18 does not change our results on the impact of expansion policy, because the probability of migrating since age 16 is not correlated with our ‘intensity’ measure. However, if the migration movements differ by

gender, then our baseline gender gap results might change when the sample is restricted to those who have been living in the same city since age 16, which we call *natives*. Since the natives represent those who have not migrated for higher education, in this sample *intensity* has no measurement error. Among natives, higher education attainment of men and women prior to the expansion were 19% and 16%, compared to full sample rates of 28% and 24%, respectively.

Another concern about our baseline results may be that the expansion have no effect in metropolitan areas, where men and women are already on an almost equal footing in higher education attainment. In the three largest cities of Turkey, namely Istanbul, Ankara, and Izmir, the higher education attainment rates before the expansion for men and women were 35% and 34%, while in all other cities, the attainment rates were 26% and 21%. One could argue that the equalizing effect of the reform would be stronger in smaller cities. Hence, we exclude the three largest cities, restrict our sample to *smaller cities*, and examine whether the reform helped reduce the gender gap in attainment in this sample.

As a further robustness check, we restrict the sample to natives living in smaller cities, which had higher education attainment rates of 16% for men and 12% for women prior to the expansion policy. By the arguments that we stated above, we expect the impact of the expansion policy on educational attainment and on gender gap to be stronger in the *natives smaller cities* sample.

We present the estimates from equation (2.1) with the aforementioned robustness checks in Panel A of Table 2.9. In Panel B, we replace survey year fixed effects with controls for age and age squared. The baseline (full sample) results are presented in columns (1) and (2) for easy reference. The *natives* sample results are in columns (3)-(4), where the coefficient estimate of *intensity* is slightly lower for men and slightly higher for women, compared to our baseline results. Moreover, the gender gap in the educational attainment of natives decreased only slightly (and statistically insignificantly) after the reform.

In column (5), we estimate the impact of expansion in the *smaller cities* sample. The coefficient estimate is statistically significant and slightly higher in magnitude compared to our baseline result in column (1). In column (6), where we consider the differential impact of the expansion policy, we find a negative and imprecisely estimated coefficient estimate for the interaction of intensity with woman dummy.

In columns (7)-(8), we restricted the sample to natives in smaller cities. Similar to earlier findings, the higher education reform increases the probability of attainment; yet, we observe no differential effect of the reform on men and women. Hence, our results are qualitatively the same as our baseline results. Moreover, comparing the results in Panels A and B, we observe that both the magnitude and the significance level of all coefficients are very similar.

Table 2.9: The effect of expansion policy: Specification and sample checks

<i>Dependent variable: Higher education attainment</i>								
	Main sample		Natives sample		Smaller cities sample		Natives in smaller cities sample	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
<i>Panel A: Survey year fixed effect control</i>								
Intensity	0.332*** (0.102)	0.393*** (0.124)	0.328*** (0.110)	0.326*** (0.116)	0.406*** (0.093)	0.498*** (0.117)	0.416*** (0.112)	0.427*** (0.123)
Intensity×Woman		-0.122 (0.089)		0.003 (0.114)		-0.186 (0.119)		-0.023 (0.151)
Observations	186,322	186,322	129,766	129,766	146,283	146,283	103,434	103,434
<i>Panel B: Age and age squared control</i>								
Intensity	0.333*** (0.102)	0.394*** (0.124)	0.328*** (0.110)	0.326*** (0.116)	0.407*** (0.094)	0.499*** (0.117)	0.416*** (0.112)	0.427*** (0.123)
Intensity×Woman		-0.123 (0.089)		0.003 (0.114)		-0.187 (0.119)		-0.022 (0.151)
Observations	186,322	186,322	129,766	129,766	146,283	146,283	103,434	103,434

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. Sample includes cohorts 1999-2005 and 2007-2013. A constant, number of 18-year-olds, region and cohort fixed effects, additionally survey year fixed effects in Panel A, age and age squared in Panel B, and the interaction of all variables with woman dummy are included as control variables. The standard errors clustered at region level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

2.9.5. Controlling for intensity in private universities

Our intensity measure includes the increase in available slots in only the public universities. The sudden and politically-motivated nature of the expansion policy in public higher education strengthens the exogeneity of the reform to individual demand in higher education. The argument behind restricting our policy variable to public higher education is that the expansion in private universities may not necessarily be a part of the expansion policy and that the changes in private university slots might be more demand-driven. However, one could argue that correlation between the change in private university slots and attainment and correlation between the changes in public and private university slots could lead to

an omitted variable bias in our baseline estimations. To address this concern and take the demand-driven variations in the education supply into account, we control for the increase in private slots in (2.1) and present our results in Table 2.10.

Table 2.10: The effect of expansion policy: Controlling for demand driven supply change

	<i>Dependent variable: Higher education attainment</i>					
	(1)	(2)	(3)	(4)	(5)	(6)
Intensity	0.332*** (0.102)	0.393*** (0.124)			0.362*** (0.103)	0.431*** (0.123)
Intensity×Woman		-0.122 (0.089)				-0.137 (0.084)
Private intensity			0.439 (0.365)	0.509 (0.411)	0.669** (0.313)	0.801** (0.374)
Private intensity×Woman				-0.142 (0.344)		-0.258 (0.342)
Observations	186,322	186,322	186,322	186,322	186,322	186,322

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. Private intensity is intensity defined for private universities. Sample includes cohorts 1999-2005 and 2007-2013. A constant, number of 18-year-olds, region, cohort and survey year fixed effects and the interaction of all variables with woman dummy are included as control variables. The standard errors clustered at region level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

Columns (1) and (2) present our baseline results from Table 2.4 for reference. In the regressions in columns (3) and (4) we replace *intensity* with the increase in private slots per 18-year-old, referred as *private intensity*. We observe that the estimated standard errors are too large despite sizable coefficient estimates. Next, we estimate the regressions where the impacts of additional slots in both public and private universities are estimated. In columns (5) and (6), the coefficient estimates of *intensity* are very similar to the baseline results. As before, the increase in available slots in public universities per 18-year-old increases higher education attainment without reducing the gender gap, controlling for the increase in available slots in private universities. Another question is whether students enrolled in private universities switched to public universities as the expansion increased slots in public universities. In Section A.2 in the Appendix, we examine how enrollment in public and private universities were affected by the expansion in public university slots, using data at NUTS-3 level. We find that the expansion was positively associated with enrollment in public universities, but negatively (though significant at 10 percent) associated with enrollment in private universities.

2.9.6. Controlling for time-varying regional covariates

Our identification strategy includes controlling for two-way fixed effects —at region and cohort level— which vary also by gender. Region effects capture any time invariant differences among regions, and cohort fixed effects capture any country level changes over time. Hence, a possible threat to our identification strategy is the existence of any other macro-level factors that vary across regions and time, other than intensity, and affect educational attainment.

Table 2.11: The effect of expansion policy: Controlling for time-varying regional covariates

<i>Dependent variable: Higher education attainment</i>						
	(1)	(2)	(3)	(4)	(5)	(6)
Intensity	0.353*** (0.090)	0.433*** (0.122)	0.355*** (0.071)	0.438*** (0.095)	0.356*** (0.079)	0.439*** (0.108)
Intensity×Woman		-0.162 (0.103)		-0.167* (0.095)		-0.166 (0.100)
Observations	113,437	113,437	113,437	113,437	113,437	113,437
<i>Controls for time varying regional covariates:</i>						
GDP per capita	No	No	Yes	Yes	No	No
Public investment in non-education industries	No	No	No	No	Yes	Yes

Notes: Higher education attainment is a binary variable equal to 1 if the individual has graduated from or is currently enrolled in a higher education program. Intensity for a given cohort and region is defined as the increase in the available slots in public universities relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. Sample includes cohorts 2004-2005 and 2007-2013 due to lack of data in time varying regional controls. A constant, number of 18-year-olds, region, cohort and survey year fixed effects and the interaction of all variables with woman dummy are included as control variables. The standard errors clustered at region level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

Here, we question whether excluding other possibly time-varying regional trends from our main specification would change our results due to omitted variable bias. With this purpose, we control for GDP per capita and public investment in all industries except for education.

In these regressions we restrict our sample to cohorts 2004-2013, since GDP per capita information is available since 2004. We present our baseline results for this sample in columns (1) and (2) of Table 2.11. Our main findings on the impact of reform on educational attainment and the gender gap are similar in the new sample to baseline results. Next, we repeat our regression analysis by controlling for GDP per capita in columns (3)-(4) and public investment in non-education industries in (5)-(6). Similar to our earlier result, the expansion policy significantly increases the average educational attainment of both men and women. Looking into the differential impact by gender, we find that the coefficient estimate of the

interaction term is negative and weakly significant in column (4), while it is negative and statistically insignificant in column (6). Hence, we confirm our earlier results that the expansion policy did not reduce the gender gap in higher education attainment, which exists to the detriment of women, after controlling for time-varying regional covariates.

2.9.7. Controlling for heterogeneous treatment effects

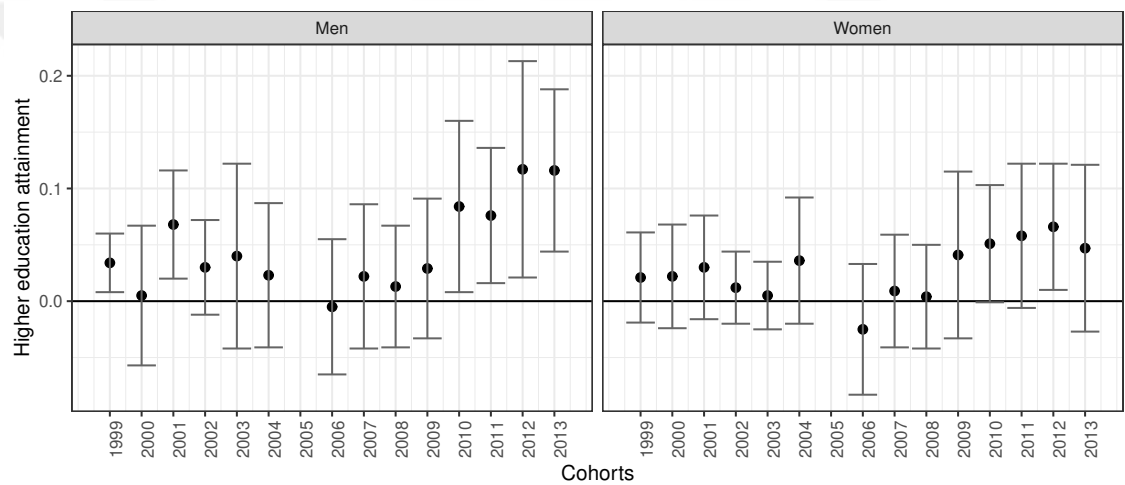
An emerging literature documents the bias in the estimates by two-way fixed-effect (TWFE) models due to the heterogeneous treatment effects for binary staggered treatment designs and proposes alternative estimators.¹² However, the case of continuous treatment, that we employ in our identification strategy, got little attention. One recent study by Callaway et al. (2021) focused on the TWFE models with continuous treatment and discuss that a potential selection bias problem arises from comparing the groups treated with low versus high doses leading to biased estimates¹³, unless a stronger parallel trend (that is stated to be not testable in Callaway et al. (2021)) and constant treatment effects across groups treated with different doses is assumed. However, the existing literature neither provides a test that we can use to determine whether our results are subject to this bias, nor proposes an alternative unbiased estimator of the treatment effect. Hence, we discretize our intensity variable into binary treatment variables in two different approaches to check the qualitative robustness of our results.

In the first approach, we aim to define a binary and time-invariant treatment variable that shows whether a region is treated with the expansion policy or not. To that end, we first observe the evolution of mean exposure intensities for three groups of regions, namely low, medium, and high intensity regions, as presented in Figure A.4 in the Appendix, which shows that high intensity regions are quite different from the other two in their mean intensities, especially in the later years of expansion. Therefore, we define our treatment variable based on whether the individual is residing in a high intensity region (the top quartile in their ranking by

¹²This literature includes Callaway & Sant’Anna (2021), De Chaisemartin & d’Haultfoeuille (2020), Goodman-Bacon (2021), and Sun & Abraham (2021). Our discussion on this literature is based on reviews by De Chaisemartin & D’Haultfoeuille (2022) and Roth et al. (2022)

¹³The selection bias in the estimated coefficient is not due to selection of different groups into different doses of treatment, but instead it comes from the heterogeneity in gains from the treatment.

Figure 2.10: Treatment effects by event study method



Source: HLFS, 2016-2017, OSYM Almanacs and TurkStat Address Based Population Registry System.

Notes: Dependent variable higher education attainment is defined as a binary variable equal to 1 if the individual is a graduate or is currently enrolled in a higher education program. Treatment variable takes value 1 if the individual is residing in a high intensity region. Each point represents the coefficient estimate for the interaction of the given cohort dummy with the time-invariant binary treatment variable, and the error bars extending from each point represent the 95% confidence intervals. The omitted cohort is 2005, the last cohort not exposed to treatment. A constant, time-invariant binary treatment dummy, full set of cohort dummies, region, and survey year fixed effects are included as the control variables.

intensities in 2013).¹⁴ This variable allows us to rely on the differential increase in the educational attainment of those residing in treated versus untreated regions, which serve as a control group. Another feature of our intensity measure, also summarized in Figure A.4 is that the education supply captured by intensity increases over time for all groups (both treatment and control), and therefore, instead of a canonical difference in differences model, we allow for dynamic treatment effects in our design and estimate the event study differences in differences analysis of the expansion policy by including the complete set of interactions of cohort dummies with binary treatment variable. We add region and survey-year fixed effects to this specification and estimate the coefficients in our main sample that includes seven cohorts before and after the policy cutoff.

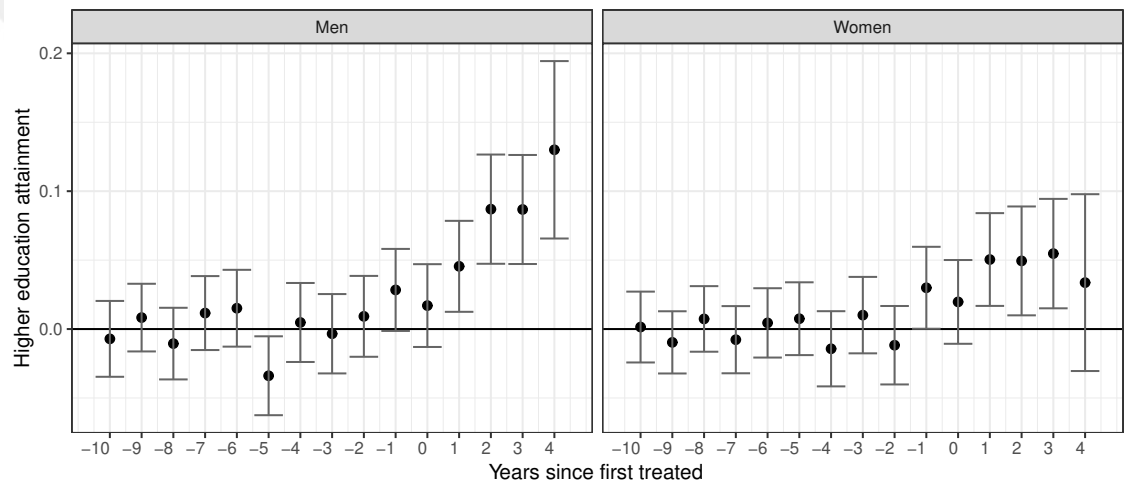
Figure 2.10 presents the coefficient estimates for the interaction of treatment dummy with cohort dummies capturing the differential increase in the educational attainment of each cohort living in treated regions compared to the same cohorts from control regions, separately for men and women. The coefficient estimates for the pre-expansion cohorts are close to zero and statistically insignificant for most cohorts confirming the parallel trend assumption. We find that point estimates for the post-expansion cohorts are positive, becoming statistically significant in 2010 and later cohorts for men. Although the point estimates are positive for women, they are not statistically different than zero for most post-expansion cohorts. In other words, the rising differences over time between high intensity and other regions lead to significantly higher educational attainment for men in treated versus untreated regions, but not for women. Although the new point estimates are slightly different than before, our results are qualitatively same, in the sense that the expansion did not reduce the gender gap in attainment, and if anything it increased it.

In the second approach, we slightly modify our treatment dummy to allow for a time-varying treatment status, and define a binary treatment variable that is 1 if the region is in the top quartile when non-zero intensities are ranked across regions and cohorts (for seven post-expansion years and 26 regions). This approach allows us to have a treatment dummy that varies across years, as well as regions.¹⁵ Callaway & Sant’Anna (2021) documents that unless one assumes

¹⁴Note that the ranking of the regions by their intensities does not change much across years due to the staggered nature of the expansion policy, validating this time-invariant treatment variable.

¹⁵Note that this variable is 1 for several regions starting from cohort 2009 but not before (because

Figure 2.11: Treatment effects by Callaway Sant'Anna method



Source: HLFS, 2016-2017, OSYM Almanacs and TurkStat Address Based Population Registry System.

Notes: Dependent variable higher education attainment is defined as a binary variable equal to 1 if the individual is a graduate or is currently enrolled in a higher education program. Each point represents the point estimates for the treatment effects by Callaway and Sant'Anna (2021) method, and the error bars extending from each point represent the 95% confidence intervals. Treatment variable takes value 1 the available slot per 18-year-old (intensity) that the individual is exposed to is in the top quartile. Cohort 2005 is omitted since it is the last cohort not exposed to the treatment.

homogeneous treatment effects across regions and cohorts, TWFE estimation using the binary and time-varying treatment variable leads to biased estimations because of using early treated regions as controls in the later stages. Hence, instead of estimating this model with a traditional TWFE model, we follow the method proposed by Callaway & Sant’Anna (2021), which first estimates dynamic treatment effects for units treated at each time period, then combines them by weighting by sample share rather than treatment variance. This method also purges the effect of potentially problematic comparison of late-treated to early-treated-as-control from the identifying variation. Note that we do not conduct Goodman-Bacon (2021) decomposition, because it is designed for TWFE models with a single treatment variable, rather than for event-study models such as our case with many treatment variables.

Our results by Callaway & Sant’Anna (2021) method are presented in Figure 2.11. Their method does not provide treatment effects by cohort but instead by the first treatment year. Since the earliest treated year is 2009 by our time varying treatment variable, the latest treatment effect is estimated for four years after the first treatment. Similar to our baseline findings, we find that the educational attainment increases with the treatment for both men and women, with a larger impact for men. Both approaches validate the robustness of our findings on the impact of expansion on the gender differences.

2.10. Conclusion

The expansion in Turkish tertiary education during 2006-2008 was drastic. In a short period, the higher education system witnessed the introduction of 40 new public universities and a 60% increase in the number of available slots. In 2005, prior to the expansion, the number of cities that lacked a university was 41 out of 81. By comparison, in 2009 there were only 2 cities without a university.

This is the first paper that lays the groundwork for the analysis of the consequences of increased supply in tertiary education in an emerging country where men’s educational attainment was greater than that of women’s. As a European Union (EU) candidate country, Turkish higher education attainment rates have been increasing, similar to the case in the group of EU-28 countries.

of the lower levels of intensity in the early years of expansion (see Figure A.4).

However, higher education attainment is still lower in Turkey than in the comparison group. Moreover, unlike in the EU-28, where a larger share of women attain higher education than men in young cohorts since mid-1990s, in Turkey the gender gap to the disadvantage of women among young adults has existed prior to the reform.

In this paper, we examine the impact of the large and politically-motivated expansion in higher education on attainment and the gender gap in higher education outcomes. We introduce an exposure intensity measure that varies across regions and cohorts. Our intensity measure shows the extent to which individuals are exposed to the expansion at the time that they were making decisions for attending higher education. We identify the causal effect of policy on overall attainment and the gender gap in higher education using the spatial and time variations in the level of exposure to the policy.

We show that for both men and women, the expansion policy increased higher education attainment. Yet, the policy did not reduce the gender gap in attainment above and beyond the time trend. Our unique dataset allows us to investigate the possible reasons behind this result. We examine how the scale of the expansion varied across fields of study and we observe that the highest growth in available slots occurred in the fields of social sciences and engineering. We find that the expansion in social sciences in the existing universities benefited men and women almost equally; therefore, it did not reduce the gender gap. On the other hand, the expansion in engineering (a STEM major traditionally preferred by men) in newly established universities benefited men more than it did women; therefore, it increased the gender gap.

From a policy point of view, our findings are important. First, we show that the policy of increasing supply in higher education was successful in raising attainment rates. Second, we show that the expansion policy did not reduce the gender gap in higher education above and beyond the time trend. In conclusion, our results suggest that the expansion policies as they have been designed may not have been the best tools to address inequalities across different socio-economic and demographic groups. The allocation of the additional slots across fields of study affected the attainment decisions of men and women differently. Hence, we conclude that an expansion policy with a specific target to reduce inequality needs to consider how the preferences of compliers of the policy may vary across different

socio-economic and demographic groups. As such, our findings can inform future expansionary policies in developing countries where women are still under-represented in tertiary education, such as those in sub-Saharan Africa and Southern Asia (Barro & Lee 2013) and even in advanced economies where women are still under-represented in STEM majors.



CHAPTER 3

HIGHER EDUCATION EXPANSION AND COLLEGE WAGE PREMIUM: EVIDENCE FROM TURKEY

3.1. Introduction

A striking feature of the last hundred years has been the enormous expansion in tertiary education. In 1900, only about one in a hundred young people in the world were enrolled in universities, but over the course of the twentieth century, this rose to about one in five (Schofer & Meyer 2005). The large increase in tertiary education attainment occurred at different times in different countries. While the US experienced a vast increase in the number of college graduates in the 1960s and 1970s, the same phenomenon occurred in other advanced countries, such as Canada, France, Spain, and Nordic countries in the 1990s. Many developing countries, seeking higher growth through increased human capital (Lucas 1989, Mankiw et al. 1992), experienced similar inflation of college graduates in the last two decades. Specifically, in Turkey, Slovakia, Czechia, Poland, Latvia, Slovenia, Portugal, and Hungary, the higher education attainment rate of 25-34-year-olds increased by more than 100% between 2002 and 2017 (OECD 2020*b*). This transformation in tertiary education attainment is bound to have an impact on the structure of wage differentials.

Many contributions in the literature have linked the demand for skilled labor and labor market institutions with the evolution of the educational wage differentials of the tertiary-educated labor force. Several studies analyzing the educational wage differentials in the US labor market provide evidence on the growing college wage premium in the 1960s and 70s parallel to the increase in the

supply of college graduates. One prominent view on the underlying causes for the rising college wage premium is the skill-biased technological change (SBTC). Bound & Johnson (1995), Katz & Murphy (1992), and Taber (2001), in addition to many others, argue that the increase in demand for skilled workers due to recent technological changes explains the increasing wage premiums of college-educated. On the other hand, Card & DiNardo (2002) and Lee (1999) argue that the changes in the labor market institutions, such as a decline in unionization and stagnant minimum wages, explain a substantial part of the wage differentials by educational attainment. In the UK (Walker & Zhu 2008) and several other European countries (Crivellaro 2014), the wage differential of college to high school graduates followed a flat path along with the increase in college graduates. The different findings between the US and European countries have been attributed to the different labor market institutions and policies (Crivellaro 2014). Even with a higher demand for skilled workers, compressed and regulated labor markets in Europe have led to lower educational wage inequality.

Even though these studies focus on the educational wage differentials at a time of relative supply increase of college graduates, the role of the supply in determining the wage differentials did not receive much attention. Fortin (2006) explores the impact of supply by exploiting interstate variation in the college graduates and shows the faster growth of college graduates suppresses the growth of college wage premium in states that experienced a higher growth. Card & Lemieux (2001) explore the role of supply in a period of college graduate stagnation and identifies the supply as an essential determinant of college wage premium by associating the rise in the premia of younger workers with the stagnated growth of college graduates. However, both of these studies fall short on identifying the impact of the supply increase since their analysis overlaps with the evolution of labor demand due to skill-biased technological change, and hence increase in demand for skilled workers is a confounding factor in their analysis. Kawaguchi & Mori (2016) compare the college wage premium in two developed countries, the US and Japan, where demand for skills due to recent technological progress have followed similar paths, while the supply of college graduates followed different trends as the supply of graduates grew substantially faster in Japan than in the US. They find that Japan and the US's contrasting wage premia trends can partly be explained by the faster-growing supply of college graduates in Japan than in the US in the same time period.

Though the association between supply and the educational wage differentials has been an important feature in many theoretical models¹, the empirical evidence on the causal relationship between the supply on the wage inequality is absent. It has been a challenge to examine the causal link between the supply of college graduates and educational wage differentials as demand for skills due to skill-biased technological change and supply of skilled workers increased and evolved in advanced countries. Turkey's recent higher education expansion policy, an exogenous reform that leads to a large increase in the supply of college graduates over a short period of time, offers an excellent case to identify the role of supply increase in the wage differentials.

In Turkey, the number of universities increased from 77 in 2005 to 138 in 2009. The number of available slots increased by 60% during the same period. By 2012, there were 158 universities in the country, and the number of available slots was two-fold of its 2005 level. The timing and spatial distribution of the policy were driven by political concerns, particularly by the requests of the members of the parliament seeking re-election in the 2007 general elections. Moreover, the implementation of the reform took place over a short period of time and was unanticipated from the perspective of university candidates.

Due to the expansion policy, the share of college graduates among young adults tripled (OECD 2020*b*), and the characteristics and skill accumulation of Turkish youth with higher education have also changed over this same period of time. As opposed to the existing trends in advanced economies, in Turkey, an emerging country with less rigid labor market institutions than EU member countries, the college wage gap has substantially narrowed down in the last 15 years (Eren et al. 2018). In this paper, we aim to investigate the role of the supply of college graduates on the evolution of the educational wage differentials in Turkey by exploiting a large-scale expansion policy, which was exogenous to the labor market and macroeconomic conditions. This question has wide relevance for other countries that have recently increased or are planning to expand higher education, especially for those where labor demand for skilled workers remains relatively stable along with the pace of technological change.

We identify the causal impact of the increase in supply of college graduates induced by higher education expansion on the additional wages earned by young

¹The pioneering work by Goldin & Katz (2009) forms the basis of many of these studies.

college graduates using the difference-in-differences methodology. Since the expansion reform started in all regions simultaneously and increased gradually over time, we use region and year variations in the level of exposure to capture the causal relationship between the additional slots introduced in the expansion period and the changes in the college wage premium. Two-way variation in the exposure measure allows us to control for both region and year fixed effects. Hence our estimates disentangle the effect of the expansion from country-level time trends and time-invariant regional differences in the educational wage differentials.

We find that the expansion in tertiary education has significantly reduced the additional wages enjoyed by the 25-29-year-old college graduate men over their high school graduate peers in the same age group. Our results indicate that the average fall in the college wage premium due to reform was 9% that varies in 3-57% in the last year of our sample. This can be due to the labor market equilibrium effect of the expansion as the supply of college graduates increased while demand remained constant. It can also be due to lower productivity of new graduates as the expansion happened over a short period, at a large scale when resources in terms of faculty and infrastructure were limited and the sorting power of the university entrance exam according to ability has decreased.

To explore whether the channel of the supply increase in skilled workers is at work, we further explore the causal impact of the expansion on the wage premiums of the older age groups, who were not subject to the policy when they were most likely to make their college entry decision. These workers were not exposed to quality issues related to newly established universities or lower selectivity of the university admission process that younger cohorts were exposed to. Therefore, we expect only labor market equilibrium effects for these cohorts; moreover, such effects are likely to be limited since younger and older graduates are not close substitutes due to their different experience profiles in the labor market. Our findings indicate a negative, statistically significant, but decreasing effect with age, providing evidence on the labor market equilibrium channel. To disentangle the impact through lower productivity channel from labor market equilibrium effects, we examine the additional wage that college graduates earn compared to their high school graduate peers at different quantiles of the wage distribution. We find that college wage premium decreases similarly along the entire wage distribution for younger age groups, while the negative impact vanishes for older age groups on the upper tail of the wage distribution.

Our paper makes an important contribution to the literature on the effects of education policies on labor market outcomes. Studies in the literature are mostly on the extension of compulsory schooling at lower levels of education (Angrist & Keueger 1991, Harmon & Walker 1995, Oreopoulos 2006, Devereux & Hart 2010, Pischke & Von Wachter 2008, Aydemir & Kirdar 2017*b*). Few studies examine the effect of higher education expansion on returns to education (Devereux & Fan 2011*a*, Kyui 2016) or labor market advantages of young graduates (Walker & Zhu 2008, Li et al. 2014, Xing et al. 2018). Our paper differs from existing studies in two ways: First, as opposed to the studies estimating the local average treatment effect of the expansion policy, as in Devereux & Fan (2011*a*) and Kyui (2016), our study examines the impact of a large-scale expansion on all labor market participants and hence identifies how it bounds to structure the labor market over time. Second, the identification strategy that relies on the regional exogenous variation in the supply of skilled labor distinguishes our study from those who rely on the variations in the educational attainment of different birth cohorts that has limited power on excluding demand-driven variations, as in Devereux & Fan (2011*a*), Walker & Zhu (2008).

The next section provides background information on the higher education expansion in Turkey. This is followed by a section that introduces the conceptual framework. Section 4 describes the data used in the analysis. Section 5 and 6 explain the empirical strategy and identification to analyze the causal impact of the reform on college wage premium. Section 8 and 9 present the results and the robustness of our results. Finally, section 9 states the concluding remarks.

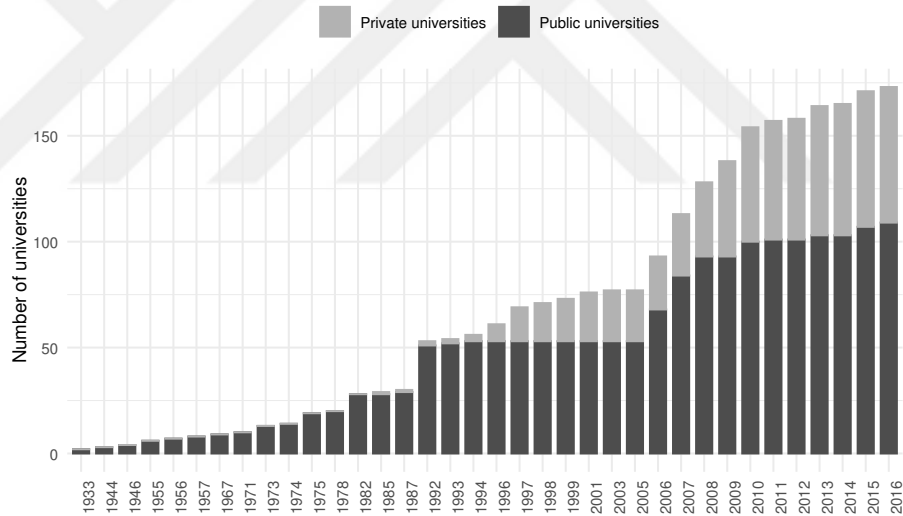
3.2. Higher education expansion

Higher education is offered by two types of universities in Turkey, namely public and private (non-profit foundation) universities. Since 1981, all universities are affiliated with the Council of Higher Education (COHE), an independent entity of the central government, and COHE is responsible for the strategic planning of higher education, in addition to establishing and maintaining quality assurance mechanisms. Public and private universities offer three forms of training: normal education, evening education, and open (distance) education. Students who study in normal or evening education receive education at the university, whereas those who attain an open education program are only obliged to pass centralized exams.

Normal and evening education programs have limited quotas determined by the COHE, but many open education programs do not have quota restrictions.

Higher education in Turkey includes vocational colleges, undergraduate programs², in addition to post-graduate level education. In this study, we are primarily interested in completing an undergraduate program. Demand for undergraduate education has been greater than supply in Turkey (parencitecohe2004, cohe2007). Therefore, a centralized competitive examination is applied to ration excess demand since 1974. Although the university exam structure may change over time, it is mainly composed of the four main subjects students study during their high school education: mathematics, Turkish, science, and social sciences.

Figure 3.1: The number of universities across years.



Source: Council of Higher Education, Higher Education Statistics.

The first university in the country was introduced in 1933, and there was a gradual increase in the number of public universities until 1987. Turkish government introduced the first wave of higher education expansion in 1992-1993. During the two years of the expansion, 24 new public universities were established in 23 cities that did not have a university before. Between 1992 and 2005, the

²Vocational colleges generally take two years, while the overwhelming majority of undergraduate programs take four years. There are only a few exceptions that take longer: Schools of Dentistry, Pharmacy, and Veterinary take 5 years; School of Medicine takes 6 years of education. These are also classified as undergraduate programs.

number of public universities remained relatively constant, and the number of private universities slightly increased, as shown in Figure 3.1. In the second wave of the higher education expansion, 50 universities (40 public and 10 private) were established between 2006 and 2008. As a result, the number of universities increased from 97 in 2005 to 138 by 2009 and 168 by 2012, and the number of available slots increased by around 60% by 2009 and 105% by 2012 compared to the slots in 2005. This second wave of higher education expansion is the focus of our study.

Table 3.1: Higher education statistics

Higher education enrollment of 18-20-year-olds			Higher education graduation of 23-25-year-olds		
Year	Men	Women	Year	Men	Women
2004	15.6	10.6	2009	15.7	15.7
2005	17.5	11.6	2010	16.8	17.7
2006	20.1	12.6	2011	18.3	19.8
2007	20.9	13.5	2012	21.4	23.3
2008	21.2	15.8	2013	22.6	26.1
2009	23.2	19.1	2014	24.3	27.5
2010	25.0	21.1	2015	25.1	30.3
2011	26.4	23.0	2016	27.1	32.2
2012	28.8	26.8	2017	28.1	35.0
2013	30.3	28.0	2018	28.3	36.9

Source: HLFS, 2004-2018.

Notes: Authors' calculations. HLFS only provides information on the educational attainment, but not on where the education had been attained (open education, public or private institutions, etc.). On the other hand, in the 2004-2013 waves of HLFS, the information on the enrollment into open and regular education is provided separately. Even though the focus of our paper is the four-year higher education in regular higher education institutions, to have comparable statistics, we calculate the mean enrollment rates as the mean enrollment into both regular and open education institutions at the higher education level.

According to education statistics derived from Turkish Household Labor Force Surveys, presented in Table 3.1 in the Appendix, the rise in enrollment in higher education started in 2006, the first year of the expansion, and continued even after 2008. The enrollment rates of young men and women of age 18-20 increased from 20% and 13% in 2006 to 29% and 27% in 2012, respectively. The individuals enrolled in higher education are assumed to graduate in about five years (Valero & Van Reenen 2019). Hence, to examine whether the enrolled students translated into graduates in five years, we also examine the higher education graduation rates of the same birth cohorts whose enrollment rates are given. To that end, we also present the higher education graduation rates of 23-25-year olds (individuals five years older than 18-20-year-olds) in the waves 2009-2018 (five years later than waves 2006-2013) in Table 3.1. The increase in the higher education attainment of 23-25-year-olds, same birth cohorts whose enrollment statistics are given in 2006 and 2012, from 18% and 20% in 2011 to 28% and 35% in 2017 supports this

argument by Valero & Van Reenen (2019).³

The higher education expansion was politically motivated and driven by requests from members of the parliament to establish universities in their cities (Arap 2010). The Turkish general election was to be held in 2007 to elect 550 members of parliament. Before this, in 2005, 25 new public universities were planned to open with the primary purpose of providing economic benefits to the regions where they were established. When the central government adopted the reform at the end of 2005, the Council of Higher Education (COHE) opposed the draft by saying that there were insufficient human resources and infrastructure to establish the new universities. However, due to pressures from members of the parliament who represented different cities and sought re-election, by 2008, 15 more public universities were opened in addition to those 25 universities initially planned.

3.3. Conceptual Framework

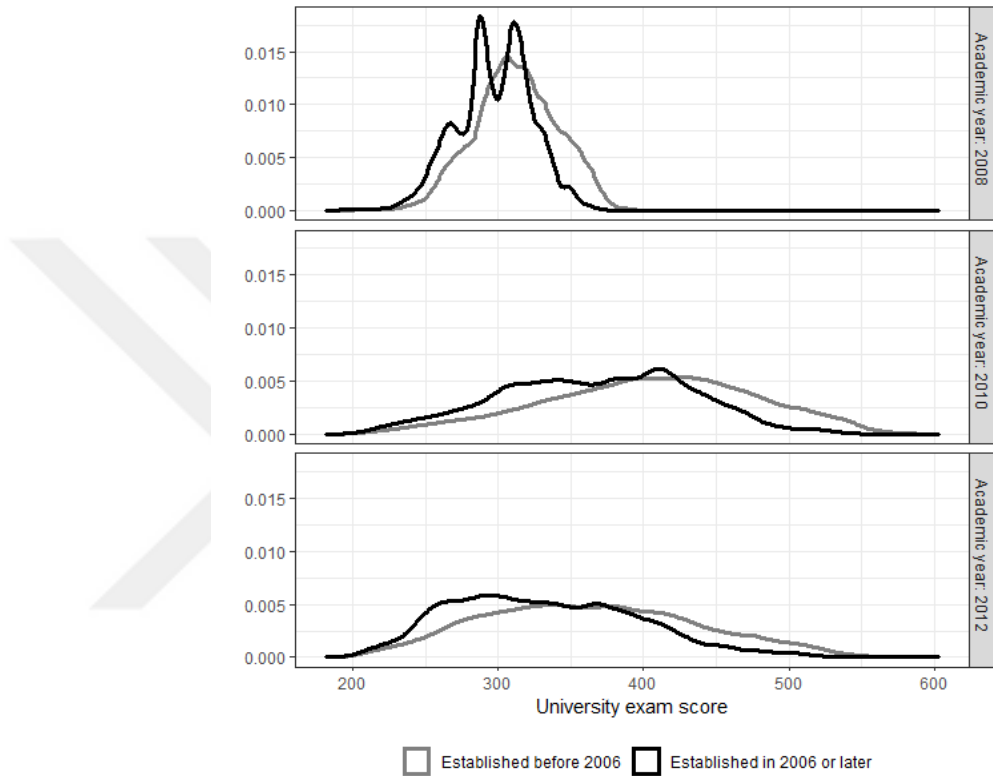
Following the conventional approach introduced by Goldin & Katz (2007), we conceptualize the changes in educational wage differentials first by using a framework of the interaction of the supply of skills (driven by changes in the educational attainment of the workforce) and the demand for skills (driven by skill-biased technological change). If we assume a CES aggregate production function with two labor inputs, namely high (college graduate) and low (high school graduate) skilled workers, with a fixed elasticity of substitution between two types of labors, then in the canonical model, the evolution of the wage differential depends on the balance between relative demand for and supply of high skilled and low skilled workers. Hence, in the absence of institutional constraints, the relative wage, called college wage premium, goes up or down depending on whether demand grows faster than supply or vice versa.

In our case, the expansion in the supply of college graduates happened in a short period. Hence, during this period, the demand for high skilled workers can be assumed to have been relatively stable. Under this assumption, we would expect the increase in labor supply beyond demand to result in a decrease in the market

³HLFS does not cover students living in dormitories. The systematically lower enrollment of women compared to their graduation rates is most probably due to a high share of women living in dormitories while studying in college.

wage of graduates in about five years after the educational expansion based on the assumption that the students graduate and participate in the local labor market five years later than their enrollment (Valero & Van Reenen 2019).

Figure 3.2: Kernel density estimates of university entrance exam scores



Source: OSYM Almanacs.

Notes: Authors' calculations. OSYM Almanacs provide the number of students admitted to all programs and the lowest and highest scores of the students for each program in all universities. The score distributions are generated by assuming that scores are uniformly distributed between the lowest and highest scores within each program.

Another possible effect of the expansion is the decline in the productivity of graduates via two channels. First, the expansion reduced the selectivity of the university entrance exam and allowed lower ability individuals to have access to higher education. Hence, the innate ability distribution of college graduates might have changed in the post-expansion period compared to the earlier years. In particular, the ability distribution of post-expansion graduates may have a lower mean and a higher variance with more observations at the low end of the distribution, compared to pre-expansion graduates. Such a change would lower the signaling value of higher education attainment, as the expected productivity of a university graduate is now lower, resulting in a lower average wage for a university

graduate. To provide empirical evidence for our claim, we compare the distributions of university entrance exam scores of students in already existing and newly established universities using data on the lowest and highest exam scores in all programs in these universities (Figure 3.2). Here, we assume that the scores in a program are distributed uniformly between the lowest and highest scores. We observe that the score distribution in newly established universities falls slightly to the left of the distribution in the existing universities (established before the reform). Hence, employers may perceive the newly established universities as less selective and pay lower wages to their graduates.

Table 3.2: Average faculty member per student over years

Year	Established before 2006	Established in 2006 or later
2001	0.037	-
2002	0.036	-
2003	0.036	-
2004	0.036	-
2005	0.034	-
2006	0.035	0.019
2007	0.038	0.021
2008	0.037	0.020
2009	0.035	0.021
2010	0.034	0.027
2011	0.032	0.031
2012	0.030	0.034

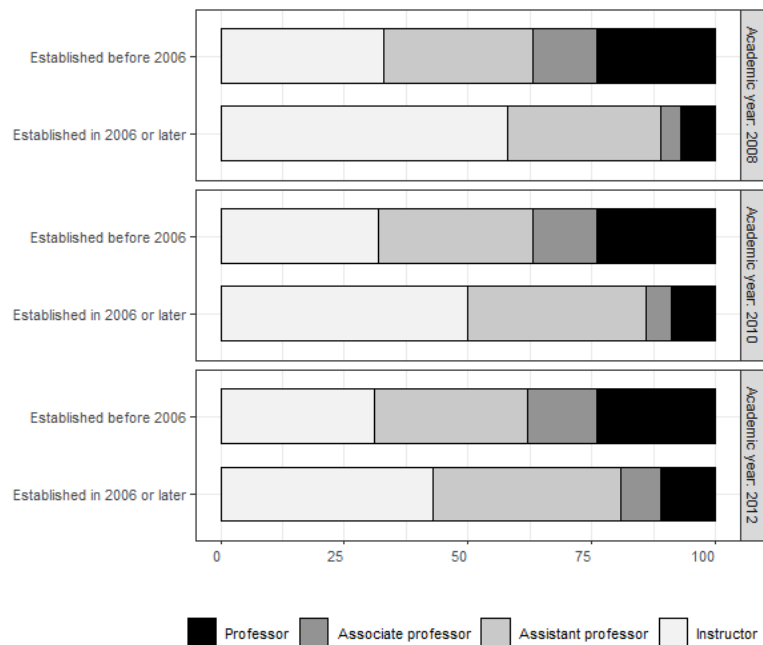
Source: Council of Higher Education statistics.

Notes: Authors' calculations. The faculty members include professors, associate professors, assistant professors and instructors. Research assistant are excluded from the faculty members since they are not allowed to teach in Turkish higher education system. The universities with more than 500,000 students are also excluded since the numbers include students having open education.

Second, the value-added of education could be lower in the universities established in the post-expansion period. With these universities, there are concerns that resources in terms of faculty (size and experience) and infrastructure fail to meet the demands of increasing number of available slots. Indeed, according to the Council of Higher Education statistics presented in Table 3.2, the average number of faculty per student is significantly lower in the newly established universities in the early years of the expansion (2006-2009) compared to the same statistic in the universities established prior to the expansion. The average number of faculty per student increased in the later years of the expansion (2010-2012), and even exceeded the rate in the pre-expansion group. Yet, in these later years, the distribution of faculty members by rank in the newly established universities is quite different from the distribution in the already existing universities. Figure 3.3 shows that instructors (the faculty members who do not necessarily hold a doctoral

degree) constitute more than half of the faculty members in the universities established in 2006 and later, in the early years of the expansion. Over time, the shares of professors and associate professors have increased. However, even in the latest year shown in the figure, the combined share of professors and associate professors is less than one-quarter of faculty in the new universities, while their share in the already existing universities is around one-third of the instructing faculty. Hence, the quality concerns about the newly established universities could limit the college wage premium enjoyed by recent graduates. In other words, the expected value added (skills gained) from a university degree may have been reduced. Moreover, the dispersion of value-added among college graduates may have increased. In this conceptual framework, we would expect a lower average wage and a higher wage dispersion among younger graduates.

Figure 3.3: Distribution of faculty members



Source: Council of Higher Education statistics.

Notes: Authors' calculations. The faculty members include professors, associate professors, assistant professors and instructors. Research assistant are excluded from the faculty members. The universities with more than 500,000 students are also excluded since the numbers include students having open education.

Ideally, to identify a decline in the innate ability of graduates and to measure the value-added by higher education, we would require data on the ability of students—such as the university entrance exam scores—as well as data that match students to the university they studied at. Since such detailed data do not exist, we examine

the impact of the quality channel indirectly by analyzing the effect of expanding educational opportunities on the heterogeneity in the wage premiums. We claim that the productivity channel at work is less relevant for individuals belonging to the upper tail of the wage distribution. In other words, wages of those who attained education in more prestigious existing universities that are more selective in the university exam and have a higher quality of academics and infrastructure remained considerably stable over time. In comparison, the channels described above reduced the relative wage advantage of individuals in the lower tail, those who received education in new universities only thanks to the higher education reform. Hence the impact of the reform might have been smaller in the upper tail of the ability distribution, but larger in the lower tail of the ability distribution.

3.4. Data

The micro dataset used in this study comes from the Turkish Household Labor Force Surveys (HLFS). The HLFS is one of the largest household surveys available in Turkey, and it is carried out annually by the Turkish Statistical Institute (Turkstat). It collects demographic, education, and labor market data from roughly 400,000 individuals living in 150,000 households each year. For each individual, the region of residence is available at the NUTS-2 level.⁴ The survey has data on age, gender, and marital status among the demographics; enrollment into education, highest educational degree, and the field of study in higher education among education outcomes; employment status, the type of employment, occupation, earnings, working hours, and the industry of the workplace among the labor market outcomes.

Our analysis using individual-level data from HLFS builds on the extensive literature that estimates the returns to schooling, beginning with the seminal work of Mincer (1974), what has become known as the Mincer model. This model estimates the rate of return by the coefficient on schooling in a log wage equation that controls for a quadratic polynomial of potential experience. Earlier literature points out that allowing for non-linearities associated with degree completion, known as the sheepskin effect, instead of assuming the same increase in earnings with each year of schooling, is essential to estimate the returns to schooling accurately (Hungerford & Solon 1987, Jaeger & Page 1996). By following these

⁴There are 26 NUTS-2 regions in Turkey.

insights, we restrict our HLFs sample to at least high school graduates and introduce a binary higher education variable that indicates whether a high school graduate had also attained higher education or not, based on the highest educational degree information available at the individual-level. We focus on the additional wage enjoyed by college graduates compared to the high-school graduates, known as college wage premium in the literature, as our variable of interest. We estimate the premium of a college degree by the coefficient of the higher education dummy in a statistical analysis of regressing logarithm of earnings onto higher education dummy and a non-linear function of potential experience.

HLFS collects data on monthly income and weekly hours of work. By using this information, we measure earnings by hourly wages. The idea behind such a measure is as follows: The variations in the monthly income may originate from the differences in both the hours of work and the hourly wages. Since the hours worked in a given month could differ across educational groups, we use hourly wages as the measure of earnings. We compute the nominal hourly wages by dividing monthly income by monthly hours of work ($4.3 \times$ weekly hours of work). Further, we adjust the nominal figures by the annual regional (NUTS-2 level) consumer price index.⁵⁶ We further winsorize hourly wages of the top and bottom 1% of the wage distribution for each wave to exclude outlier observations. We use the log of hourly wages as our main dependent variable.

To account for the effect of experience in our estimates, we add age and age-squared variables in our regressions and capture the effect of potential experience (data on labor market experience is not available in the HLFs). The years of tenure with the current employer is introduced as an additional control for potential experience. Hence, we control for both firm tenure and the second-order polynomial of age (age and age-squared) to account for potential experience.

We further introduce occupation, industry, and sector dummies using the categorical information available in HLFs on occupation (ISCO-88 until 2013, ISCO-08 in 2014 and later⁷), industry (Nace Rev.2), and sector (available only in

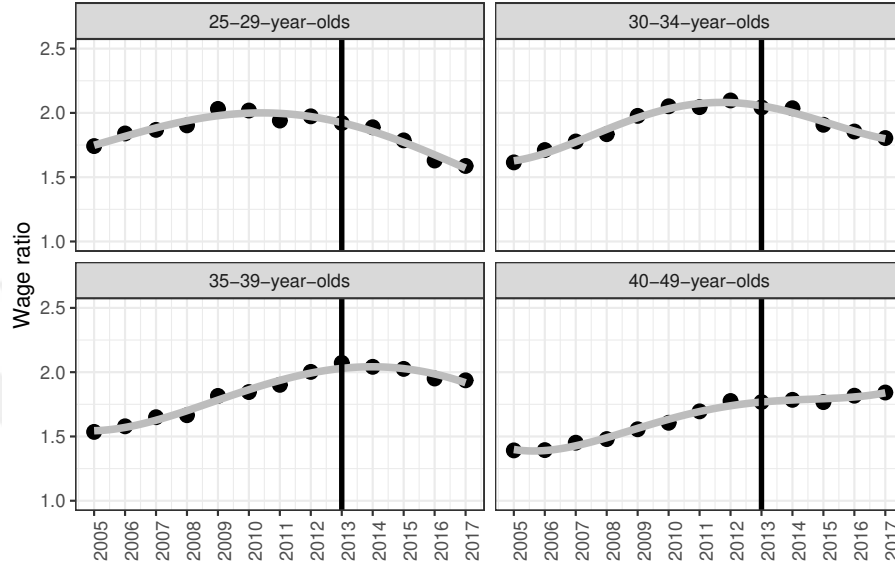
⁵⁶We use data on regional consumer price indices, published by the Central Bank of Turkey (2020) to deflate the earnings to account for the price variations across regions in addition to across years.

⁶Moretti (2013) points out that the nominal wage differentials by educational attainment significantly exceed the real wage inequality due to the geographical concentration of skilled workers into high-cost locations.

⁷We convert ISCO-08 classification to ISCO-88 and use ISCO-88 classification to construct

2009 and later waves) in some specifications to control for the wage differentials across occupations and workplace characteristics.

Figure 3.4: Evolution of college wage premium



Source: Authors' estimations using HLFS 2005-2017.

Notes: The wage ratio for each year gives the ratio of average hourly wages for university graduates to high school graduates for the age group given in each panel. The sample includes full-time paid employees with at least a high school degree observed in HLFS 2005-2017. The gray curves are the local linear fits to capture the time trend in the wage differentials.

Figure 3.4 displays the evolution of the college wage premium over time captured by the ratio of average hourly wages for university graduates to high school graduates. The lines show the local linear fits to capture the time trend of the wage premium. Our figure documents an increasing trend in the first decade of the twenty-first century in the college wage premium of all age-groups of 25-49. Many developed and developing countries experienced a similar increase, mostly associated with the skill biased technological change. However, around 2010s we observe a decline in the wage advantage enjoyed by college graduates. Moreover, the decline is steeper for younger adults in the labor market. To associate the recent decline in the college wage premium, especially for young adults, with the reform that increased access to higher education, we use data on both supply and demand-side variations in higher education, as described next.

Table 3.3 additionally provides information on the wage dispersion separately for occupation dummies

high school and college graduates observed before and after 2012. Consistent with our observations in Figure 3.4, the ratios of both mean and median wages of university graduates to high school graduates fall for younger age groups after 2012, parallel to the increase in their supply. On the other hand, the same wage ratios remain flat or slightly increases for older age groups. The last two columns show the q90/q10 and q75/q25 ratios separately for high school and college graduates from pre and post-2012 waves. We observe that the spread of the real wages remains relatively the same or slightly widens for college graduates, while it significantly narrows down for high school graduates in post-2012 for all age groups.

Table 3.3: Distributional evolution of wages by age group and educational attainment

Age group	Educational attainment	Time period	Mean	Median	q90/q10	q75/q25
25-29-year-olds	High school	Pre-expansion	2.041	1.771	3.028	1.717
		Post-expansion	2.397	2.173	2.562	1.584
	University	Pre-expansion	3.922	3.641	4.415	2.390
		Post-expansion	4.318	3.857	4.159	2.422
30-34-year-olds	High school	Pre-expansion	2.529	2.124	3.649	2.013
		Post-expansion	2.682	2.370	2.827	1.675
	University	Pre-expansion	4.748	4.641	3.640	1.908
		Post-expansion	5.254	5.119	3.854	2.118
35-39-year-olds	High school	Pre-expansion	3.052	2.693	3.929	2.295
		Post-expansion	2.981	2.582	3.197	1.821
	University	Pre-expansion	5.305	5.190	2.909	1.679
		Post-expansion	5.905	6.495	3.339	
40-49-year-olds	High school	Pre-expansion	3.622	3.544	3.931	2.091
		Post-expansion	3.663	3.261	3.644	2.288
	University	Pre-expansion	5.550	5.341	2.477	1.576
		Post-expansion	6.567	6.448	2.706	1.571

Notes: Authors' calculations using 2005-2017 HLFS.

To construct our key variable, the measure of exposure to expansion policy in a given region and year, we collect data on the number of available slots in four-year higher education programs and the population of 18-year-olds. We define the measure of exposure as the increase in available slots divided by the 18-year-old population where the former captures the supply-side variations and the latter represent the demand changes in higher education. We use the online almanacs of the OSYM to obtain the data on the available slots. Each year OSYM publishes data on the number of available slots in all programs (majors) in all universities. We collected and digitalized the data for the 2005-2012 period. The location identifiers (Province (NUTS-3), and NUTS-2 regions) of each university are also known. We aggregate the available slots at the NUTS-2 level to merge with our

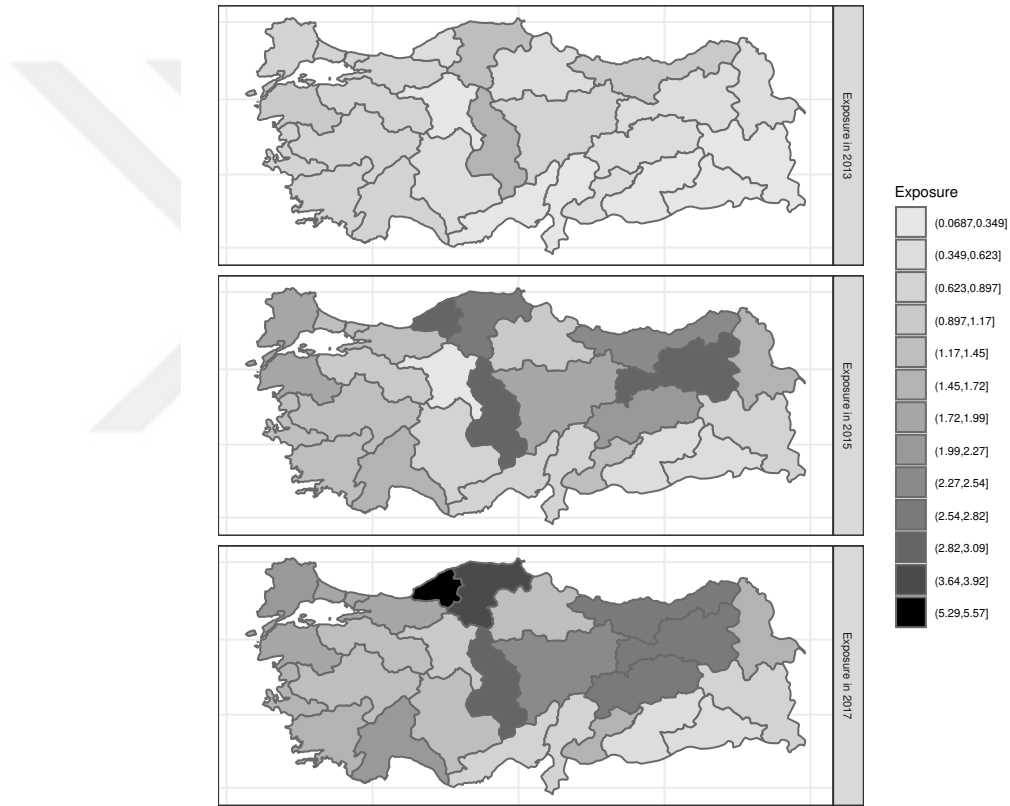
HLFS data and for each year we subtract the base number of slots in 2005 (before the expansion policy) to define the increase in the stock of educational opportunities with the intervention.

We obtain the population data of 18-year-olds at the province level from the Address Based Population Registry System of Turkstat, available since 2007. We aggregate the 18-year-old population information in 2007-2012, at the NUTS-2 level to merge with our HLFS data. The number of 18-year-olds represents the size of the population that benefits from the higher education opportunities in a particular region since an individual is typically 18-year-old when he graduates from high school. We use the number of 18-year-olds to normalize the available slots in order to control for any scale differences across regions.

Our final variable to measure the level of exposure to the reform is the increase in the available slots (compared to the base year 2005) per 18-year-old population. Since the reform was initiated in 2006, exposure is zero in the earlier years, as done in the earlier studies in the literature (Duflo 2001, Berlinski et al. 2009). We refer to these years as the pre-expansion years in the rest of the paper.

The individuals enrolled in higher education programs are expected to participate in the labor market five years later than their enrollment (Valero & Van Reenen 2019). Hence, we measure the level of exposure of an individual who participates in a local labor market by the increase in available slots relative to the number of slots in 2005 (before expansion), normalized by the 18-year-old population in the same region, five years before the year that the individual is observed in the HLFS. In other words, we lag our exposure measure by five years to evaluate its impact on the labor market when the students who benefit from the exposure actually enter the labor market. This approach, similar to Fortin (2006), allows us to exclude the concerns on the migration of skilled labor to the region to benefit from higher returns to education. Our exposure measure is calculated using data on the number of slots from 2006-2012, associated with the labor markets in 2011-2017. As discussed above, in the definition of *exposure*, the increase in slots is normalized by the population of 18-year-olds to control for the scale differences across regions. Thus, for example, for an individual observed in the 2015 survey living in Konya region, exposure is measured as the increase in available slots from 2005 (the year that the expansion policy was initiated) to 2010 in Konya region, divided by the number of 18-year-olds in 2010 in Konya region.

Figure 3.5: The spatial distribution of exposure across years.



Source: OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: Authors' calculations. Exposure for the year t is measured by the increase in available slots per 18-year-olds in the year $t - 5$.

Figure 3.5 depicts a detailed demonstration of the evolution of the expansion's spatial distribution over time in 2013, 2015, and 2017 (5 years lagged) and provides evidence on the idiosyncratic nature of the expansion policy. Figure shows that exposure has significantly varied over time, as well as across regions. Some regions started with a much higher exposure than others in the early years of the expansion. There are several regions with low exposure at the beginning of the expansion, but reached the highest exposure in the later years of the expansion, and vice versa. In other words, the expansion did not proceed at a uniform pace across regions.

We use the 2006-2017 waves of HLFS to estimate our main regressions. We also make use of an earlier wave of the HLFS (2005) to show that the intervention to higher education is exogenous to pre-existing trends in the labor market.

3.5. Empirical strategy

Our main aim is to identify how the expansion policy affects the additional wage enjoyed by college graduates compared to high school graduates in the local labor markets. The local labor market is determined by the region of residence of the individual in the year an individual is observed with a wage. In other words, year and region jointly determine the local labor market that an individual participates in and also the level of exposure to the expansion policy.

We identify the causal impact of the reform by exploiting the spatial and temporal variations in the exposure level. Our empirical approach is similar to Card (1992), Duflo (2001), Berlinski et al. (2009), and Müller & Wrohlich (2020).

We introduce a generalized regression framework of the identification with the following equation:

$$\log(w_{irt}) = \beta_1 h_{irt} + \beta_2 s_{rt} + \beta_3 h_{irt} \times s_{rt} + \alpha_0 + X_{irt} \alpha_1 + \gamma_r + \mu_t + \epsilon_{irt} \quad (3.1)$$

where w_{irt} represents our dependent variable, wage rate of individual i living in region r and observed in year t , and s_{rt} represents our explanatory variable of interest, exposure to treatment in region r and year t , defined as the increase in the

higher education slots in region r and year $t-5$ compared to the pre-expansion year 2005. The variable h_{irt} is a binary indicator whether the individual is a college graduate or not. The interaction of h_{irt} and s_{rt} aims to account for possible differential effect of the expansion on high school versus college graduates. X_{irt} represents individual-level controls introduced in a classical Mincer Model. The inclusion of region effects, γ_r control for time-invariant unobserved regional differences which could impact our dependent variable. The year fixed effects, μ_t capture any time-related differences observed in the dependent variable at the country-level. By the inclusion of two-way fixed effects, our parameter of interest β_3 captures the correlation between the exposure to reform and within region variations over time in the differential wage of college graduates. Therefore, our estimates for β_3 can be interpreted as the causal effect of the expansion policy on the returns to a college education.

The vector of individual controls consists of potential experience, captured by age, age squared and firm tenure. In some specifications we additionally control the occupation and workplace characteristics to check the robustness of our results.

We estimate the equation (4.2) separately for the age groups 25-29, 30-34, 35-39, and 40-49. The error term ϵ_{irt} captures the impact of individual-specific idiosyncrasies that influence the dependent variable. We calculate standard errors in three ways, following different strands in the literature. First, we cluster errors at the region level by following Abadie et al. (2017), that suggest the standard errors should be clustered at the level that treatment is clustered. In our study, the expansion policy (i.e., our treatment variable) is determined at the region level. Hence, we first estimate the standard errors by allowing within-region correlation. Second, we cluster errors at the year of birth level following Bertrand et al. (2004a), who suggest that clustering should be performed to capture time-trend related autocorrelation in the sample. Since consecutive waves of the HLFS include individuals born in the same year who were probably observed in the labor market in the same year, their wages may be correlated, leading to serial autocorrelation in the data. To account for that, we allow the errors to be correlated among those born in the same year. Finally, we combine these two approaches and cluster standard errors at the region $\times$ year of birth level. In the rest of the paper, we only report the standard errors clustered at the region level since it gives the largest standard errors.

3.5.1. Sample

We use micro-data from the 2005-2017 waves of HLFS to construct our sample. The year 2006 is considered to be the transition years into the expansion policy. Hence, we exclude five years posterior to this transition year, i.e., year 2011, from the sample to reduce the fuzziness in the treatment status. Therefore, our sample includes individuals observed in the six waves (2005-2010) that belong to the pre-expansion period and six waves (2012-2017) in the post-expansion period. We avoid extending the time window around the transition years beyond six years with the concern of maintaining exogeneity of the reform to demand and also to exclude the impact of longer-term macroeconomic fluctuations⁸ on the labor market. We trim down the time window and construct narrower alternative samples with 4 and 3 years around the policy transition years, to control the robustness of our results by excluding possible impacts of the political events⁹ and other policy interventions occurred in 2016 on the labor market¹⁰

We further restrict the sample to those working full-time, i.e., working at least 30 hours a week by following the earlier literature (Aydemir & Kirdar 2017*a*, Kawaguchi & Mori 2016), with a wage, since the wage information is not available for employers, or self-employed.

Next, we restrict our sample to those who completed at least secondary education (graduated from high school) for two reasons: First, secondary education graduates are the ones who have the minimum qualification to attend higher education and so they are the primary target of the expansion policy. Second, we aim to compare college graduates with high school graduates instead of with all individuals who may have many different educational qualifications Goldin & Katz (2007).

We further restrict our sample to males aged 25-49. First, females are excluded

⁸Turkey went into recession at the end of 2018 according to official statistics published by Turkstat (2019).

⁹Due to the coup attempt in 2016, a state of emergency was declared in the country. Many people with the suspicion of being related to the parties organizing the failing coup attempt were dismissed from their jobs with a series of decree laws.

¹⁰At the end of 2015, the governing party has increased the minimum wage by around 30% as a part of their general election promises. The result of such a large increase in the minimum wage on the wage differentials is not very clear since most high school graduates are paid with minimum wage according to statistics derived from 2015 HLFS.

to avoid bias in the sample due to the self-selection of women into employment.¹¹ Individuals who are younger than 25 are excluded to avoid selection due to those who did not complete their educational attainment, those who are 50 and above are excluded to avoid selecting those who retire early, especially in the early years of our sample.

Finally, we divide our sample into several subsamples by the age-group since, first, individuals from distinct age profiles may not be close substitutes of each other in the labor market (Card & Lemieux 2001), and second, the channels at work might differ across our subsamples. The experiences of older and younger workers may exhibit large differences, especially as the age-gap increases, that could lead to heterogeneous impacts of the expansion policy on different age-groups. Moreover, as explained in the conceptual framework, younger workers are affected by the reform via two possible channels (productivity and labor market equilibrium channel), whereas older workers are affected only through the higher number of younger graduates in the labor market (labor market equilibrium channel). Although we separately analyze several age groups, our primary sample of interest is 25-29-year-old men in this study. We examine the results for 25-29-year-old women to complement our results for young men. Similar to our sample of interest, we explore the impact on the remaining age to provide evidence on the full picture of how the expansion alters the labor market outcomes of college graduates in the labor market.

3.6. Identification

In this section, we first verify the validity of our identification strategy by providing evidence on each assumption embedded in the interpretation of the estimates from equation (4.2) as the causal effects of the reform. Then, we conduct a regression analysis to show the impact of the expansion policy on the share of college graduates.

¹¹Female labor force participation rate in Turkey, at 28.9 percent, is the lowest among OECD countries and the 22nd lowest in the world (World Bank 2018)

3.6.1. Exogeneity of expansion

In estimating the impact of an expansion policy on labor market outcomes, researchers face the problem of identifying the effects of a compensatory intervention or an intervention that targets the areas that are already improving their labor market outcomes. To address this concern, we investigate whether the expansion is correlated with the average education and labor market outcomes of individuals in ages 25-64 with at least secondary education attainment, i.e., our sample of interest at the regional level prior to the expansion. First, we calculate the higher education attainment among those with at least high school education in 2005 Turkish HLFS across regions. Next, we introduce mean employment, unemployment, informal employment, and average hourly wages in the same sample. Finally, we introduce the gross domestic product (GDP) per capita in 2005, using the data published by Turkstat, across regions. We show in Table A.6 that the increase in available slots per 18-year-olds (exposure) in 2013, 2015, and 2017 is not correlated with any of these variables observed in the pre-expansion year 2005¹².

Table 3.4: The exogeneity of expansion policy

<i>Dependent variable:</i>	Exposure in 2012	Exposure in 2013	Exposure in 2014	Exposure in 2015	Exposure in 2016	Exposure in 2017
Higher education attainment of aged 25-34	1.227 (1.876)	5.424** (2.172)	5.365 (3.519)	3.473 (5.725)	3.663 (6.422)	6.909 (7.947)
Higher education attainment of aged 25-44	1.074 (2.148)	6.431** (2.452)	6.841 (3.965)	4.967 (6.502)	5.278 (7.296)	9.422 (9.004)
Higher education attainment of aged 25-64	0.766 (2.196)	4.809* (2.66)	4.098 (4.204)	0.832 (6.707)	0.638 (7.519)	4.590 (9.341)
Number of universities per 18-year-old	-3.885 (2.23)	2.929 (2.99)	6.205 (4.431)	7.581 (7.046)	11.064 (7.764)	9.535 (9.905)
Number of public universities per 18-year-old	-2.337 (1.514)	0.63 (2.04)	1.209 (3.08)	-1.485 (4.827)	0.252 (5.421)	0.821 (6.766)
Number of available slots per 18-year-old	-1.933*** (0.467)	0.285 (0.786)	1.416 (1.156)	1.199 (1.848)	1.507 (2.067)	1.393 (2.593)
Number of available public slots per 18-year-old	-1.687*** (0.436)	0.148 (0.715)	0.991 (1.063)	0.438 (1.692)	0.715 (1.893)	0.785 (2.365)
Gross domestic product per capita	3.051 (15.426)	20.637 (19.437)	10.098 (29.991)	-28.27 (46.698)	-21.062 (52.564)	-5.280 (65.831)
Unemployment among working age population	-0.121 (1.434)	-2.19 (1.792)	-4.436 (2.641)	-7.267* (4.111)	-7.883 (4.627)	-8.410 (5.870)
Unemployment among aged 25-34	0.028 (1.354)	-1.721 (1.708)	-3.222 (2.553)	-4.61 (4.018)	-5.32 (4.496)	-5.323 (5.671)
Unemployment among at least high school graduates	0.215 (1.512)	-0.505 (1.946)	-1.673 (2.925)	-2.689 (4.577)	-2.532 (5.141)	-1.907 (6.439)
Unemployment among at least college graduates	-0.206 (1.936)	-1.134 (2.484)	-2.421 (3.739)	-3.788 (5.852)	-2.603 (6.595)	-1.028 (8.258)
Labor productivity in industry sector	-19.135 (12.555)	5.056 (16.907)	12.402 (25.478)	-15.907 (39.943)	-21.933 (44.694)	-24.397 (55.854)
Labor productivity in all sectors	4.81 (7.455)	13.824 (9.266)	8.251 (14.544)	-7.666 (22.864)	-10.017 (25.605)	-5.415 (32.048)

Notes: A constant is included as a control variable. The standard errors are given in parentheses.*p<0.1 **p<0.05 ***p<0.01.

¹²We also experiment with the same explanatory variables calculated for ages 25-34, 25-39, and 25-49 and confirm that the results are similar to the ones presented in Table A.6

3.6.2. The expansion policy on the share of college graduates

The increase in the share of college graduates observed in the recent years is associated with the large-scale expansion policy. In Turkey, the share of individuals in ages 25-29 with a higher education degree was 15 percent in 2008, varying substantially across regions from 8 to 24 percent. After the expansion, the percentage of college graduates among 25-29-year-olds rose to 25 in 2013 and 36 in 2017. When we compare year 2010, five years into the expansion, with the final year of our sample, 2017, we observe that the increase in the share of college graduates among 25-29-year-olds is around 17-24 percentage points in all regions.

In our empirical strategy, we assume that the slots introduced in higher education shaped the local labor market five years later. This assumption would not be valid if, first, the additional slots did not translate into college graduates, or second, if the increase in the share of college graduates were not a result of the reform. Here, we check the validity of this assumption by identifying the impact of the expansion on the share of college graduates.

We conduct a regression analysis by exploiting the variations in the level of exposure across regions and years using the following equation:

$$c_{rt} = \theta_1 s_{rt} + \theta_2 x_{rt} + \gamma_r + \mu_t + \nu_{rt} \quad (3.2)$$

where c_{rt} is the share of college graduates in region r and year t , s_{rt} is the exposure, the increase in available slots per 18-year-old five years before t , and x_{rt} is the vector of control variables that includes a constant and the number of 18-year-olds. The inclusion of region fixed effects allows us to control for time-invariant regional differences correlated with the share of college graduates. Year fixed effects capture any time-related differences observed in the dependent variable at the country-level. Hence the coefficient θ_1 of the exposure variable captures the correlation between the share of college graduates and the exposure to the reform, which varies by region and year; and can be interpreted as the causal impact of the expansion policy.

Table 3.5 present the estimation results of the equation (3.2) where the dependent variable is the share of college graduates among 25-29-year-olds in

column (1), 30-34-year-olds in column (2), 35-39-year-olds in column (3), and 40-49-year-olds in column (4). Our results suggest that a one-unit increase in the exposure significantly increases the share of college graduates by around 12 percentage points among 25-29-year-olds.

Table 3.5: The effect of expansion policy on the share of college graduates

<i>Dependent var: Share of college grad. among</i>	<i>22-24-year-olds</i>	<i>25-29-year-olds</i>	<i>30-34-year-olds</i>	<i>35-39-year-olds</i>	<i>40-49-year-olds</i>
Exposure	0.009* (0.004)	0.012*** (0.003)	0.005 (0.003)	0.004 (0.003)	0.003 (0.002)
Observations	234	234	234	234	234
<i>Control for:</i>					
18-year-old population	✓	✓	✓	✓	✓
Region effects	✓	✓	✓	✓	✓
Year effects	✓	✓	✓	✓	✓

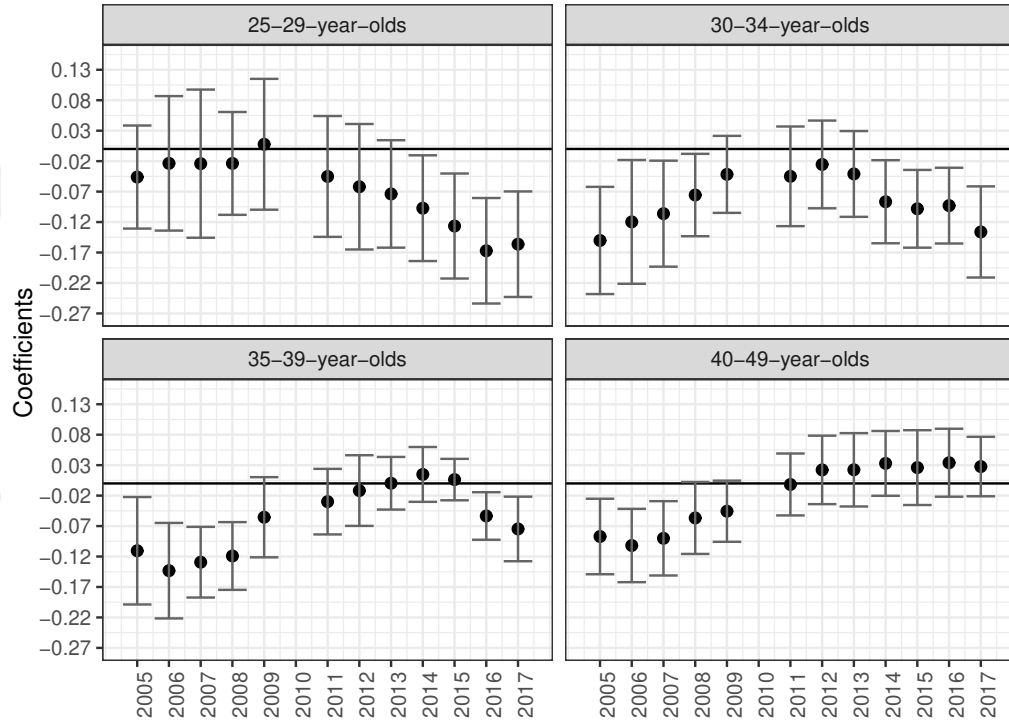
Notes: The standard errors clustered at region level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

Our findings indicate that older age groups are not likely to have benefited from the expansion. The increase in the share of college graduates in these groups, although estimated to be positive, is very small and not statistically significant. Note that the slight increase in the share of college graduates in the older age-groups in higher exposure regions might have resulted from the migration of high-skilled workers to the region to benefit from job creation in the region due to the establishment of new universities. To conclude, our results suggest that underlying regional trends in attaining a college degree that are systematically correlated with the expansion policy are not driving our results about the share of college graduates.

3.6.3. Parallel trend analysis

Another assumption that we make in equation 4.2 is that our dependent variable has a common time trend across regions. To test the validity of this assumption, following the earlier literature, we divide 26 NUTS-2 regions into two groups by ranking them by their level of exposure in 2017 (the increase in available slots per 18-years-old from 2005 to 2012). We call the regions with an above-median exposure *high exposure* regions and others *low exposure* regions. After introducing a binary variable *high exposure* indicating whether the region of residence of an

Figure 3.6: Event study analysis of the expansion policy



Source: HLFS 2005-2017, OSYM Almanacs and TurkStat Adress Based Population Registry System.

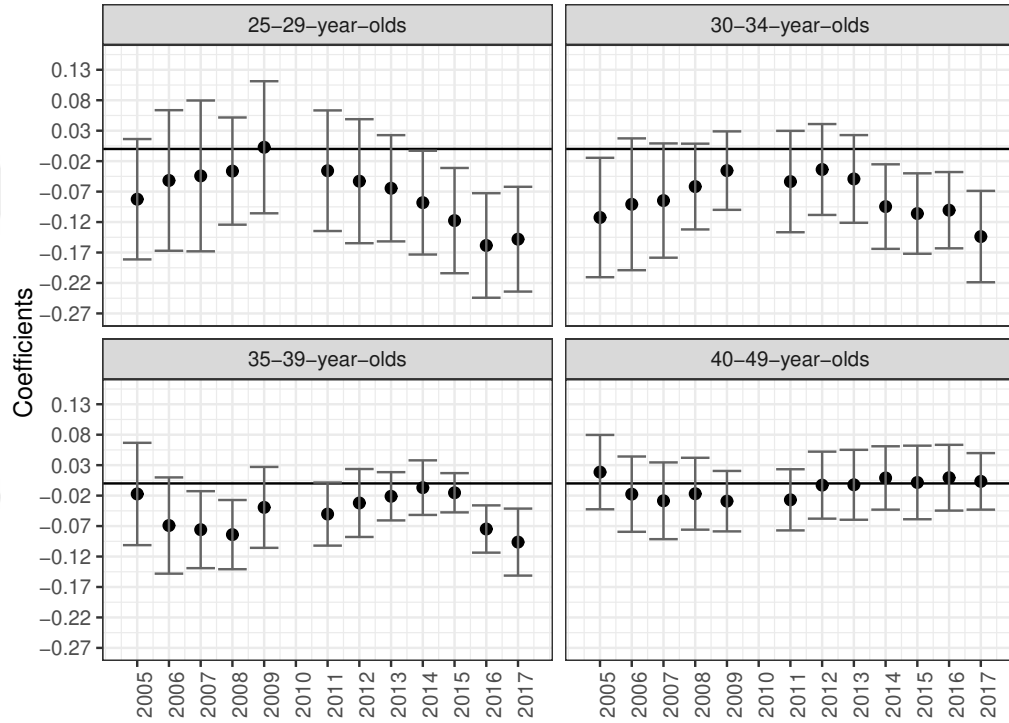
Notes: The dependent variable is the log of hourly wages. All regressions include controls for a constant, age and age squared, firm tenure, education dummy, binary high exposure variable, region fixed effects, and year fixed effects. The sample restricted to full-time paid employees of the given age group with at least a high school degree observed in the sample of 2005-2017. Each point represents the coefficient estimate for the interaction of the given year dummy with the time-invariant binary high exposure variable, and the error bars extending from each point represent the 95% confidence intervals based on the standard errors clustered at region level. The omitted year is 2010, the last year before the treatment.

individual belongs to high exposure regions, we estimate our main equation by including the high exposure dummy instead of our continuous exposure variable, and by interacting all year dummies with this high exposure dummy. After introducing a binary *high exposure* variable indicating whether the region of residence of an individual belongs to high exposure regions, we conduct an event study analysis as follows: We include the high exposure dummy and its interaction with the education dummy and full set of year dummies in our main equation instead of a continuous exposure variable that varies across regions and years. The base year is determined to be the year right before the new graduates enter the labor market, i.e., the year 2010, and omit the year 2010 dummy in our estimations. We keep the individual controls the same as in equation (4.2).

Figure 3.6 presents the coefficient estimates for the interaction of high exposure dummy with education and year dummies, capturing the differential change in the additional wages enjoyed by college graduates compared to their high school peers living in treated regions compared to the same additional wages from control regions. The coefficient estimates for the pre-expansion years are close to zero and statistically insignificant for those aged 25-29, confirming the parallel trend assumption. We find that point estimates for all post-expansion years are negative, becoming statistically significant in 2014 and later.

Our estimates for the older age groups, however, show a violation of the parallel trend assumption with increasing, and statistically different than zero, wage differentials in treated regions compared to the control ones in pre-expansion years. The eyeballing of the event study figures for older age groups indicates a linear trend in wage differentials, and therefore, we include a linear trend variable for the pre-expansion years to absorb this differential trend between treatment and control regions. We present the same coefficient estimates after the pre-expansion de-trending in Figure 3.7. The inclusion of the linear pre-trend leads to around zero pre-trend differentials in treated regions compared to control regions, and therefore, we additionally include a linear pre-trend to our control variables in equation (4.2).

Figure 3.7: Event study analysis of the expansion policy with pre-trend control



Source: HLFS 2005-2017, OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: The dependent variable is the log of hourly wages. All regressions include controls for a constant, age and age squared, firm tenure, linear pre-trend, education dummy, binary high exposure variable, region fixed effects, and year fixed effects. The sample restricted to full-time paid employees of the given age group with at least a high school degree observed in the sample of 2005-2017. Each point represents the coefficient estimate for the interaction of the given year dummy with the time-invariant binary high exposure variable, and the error bars extending from each point represent the 95% confidence intervals based on the standard errors clustered at region level. The omitted year is 2010, the last year before the treatment.

3.7. Results

3.7.1. The expansion policy on college wage premium

Prior to the expansion (2005-2010), the difference between the average earnings of high school and college graduates had been increasing over time so that by 2010 university graduates in ages 25-29 were paid about 109% more in an hour than high school graduates. This wage gap decreased in the post-expansion years (2012-2017) so that in 2017, young college graduates are paid only 63% more in hourly wages than high school graduates. In order to examine the impact of the expansion policy on wage differentials, we estimate the equation (4.2) and present our results in Table 3.6 for those aged 25-29, 30-34, 35-39, and 40-49.

Table 3.6: The effect of expansion policy on hourly wages

Ages:	25-29		30-34		35-39		40-49	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Education	0.616*** (0.030)	0.620*** (0.031)	0.691*** (0.023)	0.698*** (0.024)	0.649*** (0.016)	0.635*** (0.016)	0.528*** (0.015)	0.534*** (0.016)
Education×exposure	-0.092*** (0.018)	-0.079*** (0.022)	-0.069*** (0.013)	-0.057** (0.021)	-0.027** (0.010)	0.008 (0.009)	-0.022** (0.010)	-0.031** (0.013)
Observations	60,202	41,409	62,768	43,158	52,612	35,422	75,325	50,171
Mean exposure in the post expansion years	1.000	0.739	1.028	0.746	1.074	0.773	1.071	0.768
Sample:	2005-2017	2007-2015	2005-2017	2007-2015	2005-2017	2007-2015	2005-2017	2007-2015

Notes: The dependent variable is the log of hourly wages. All regressions include controls for a constant, age and age squared, firm tenure, linear pre-trend, education dummy, continuous exposure variable, region fixed effects, and year fixed effects. The sample restricted to full-time paid employees with at least a high school degree observed in the sample of years given each column. The observations from year 2011 are excluded from all regressions. The standard errors clustered at region 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 coefficient of the education dummy captures the mean college wage premium in the years where the exposure is zero, i.e., the pre-expansion years. Our baseline results given in the column (1) and (2) for the younger age group of 25-29 suggests that, after accounting for potential experience, time trend, region and year fixed effects, in the pre-expansion period, college graduate men earn about 61-62% more than their peers with at most high school degree, whereas this rate is 69% for aged 30-34, 64% for aged 35-39, and 53% for men aged 40-49.

The coefficient estimates of the interaction term *education×exposure* show the causal impact of the reform on college wage premium. The estimates are negative and strongly significant for men aged 25-29 in column (1), where we present our baseline results for the younger age group showed to be the compliers of the expansion policy. Specifically, the wage premium received by 25-29-year-olds

decreased by 9% with a one-unit increase in the exposure. Since the mean exposure in the post expansion years was 1.00 on average, the reform resulted in 9% fall in the college wage premium of 25-29-year-old men observed in post-expansion years. The exposure to the expansion varied between 0.37 and 5.56 across 26 NUTS-2 regions in the final year of our sample, 2017. Hence, our findings indicate that the wage advantage of 25-29-year-old college graduates fell by around 3-57% due to government's expanding educational opportunities. Narrowing down the sample to four years around the policy cutoff only slightly changes the estimates for our parameters of interest and keep the level of statistical significance the same.

Our estimates on the interaction term on the sample of older age groups show that the adverse effects of increasing the supply of skilled workers on wages are statistically significant but relatively small in magnitude. By accounting for the mean exposure levels in the post expansion years for each age group, we find that the expansion policy reduces the returns to university education by 7% for men aged 30-34, 3% for those aged 35-39, and 2% for those aged 40-49.

The older employees whose results are presented in columns (3)-(8) are expected to have started their higher education in 2007 or earlier, and therefore mostly not expected to be exposed to the expansion policy. Our findings on the statistically significant impact of the exposure on the share of college graduates aged 30 and older, presented in Table 3.6 provide suggestive evidence that the labor market equilibrium channel is at work, and, therefore, increasing the supply of skilled workers slight diminishes the college wage premium of older workers not exposed to the policy.

Our findings on the higher negative impact of the expanding educational opportunities on the labor market advantage of younger college graduates compared to their high school graduate peers of the same age suggest that supply increase may not be the only driving mechanism for the fall in the college wage premium and the lower productivity of young college graduates might be another mechanism of lower premium for that age group. On the other hand, the channel of the supply increase is expected to lower the wages of all college graduates to varying degrees. However, since older and younger workers are not close substitutes to each other due to their different experience profiles, we might expect to observe a higher impact on the younger college graduates' relative advantage in the labor market compared to the older workers.

Hence, though our findings presented in this section confirms the existence of the labor market equilibrium channel, does not entirely confirm or rule out the hypothesis that expanding the educational opportunities with a sudden and not well-prepared policy leads to lower productivity for those benefited from these additional opportunities.

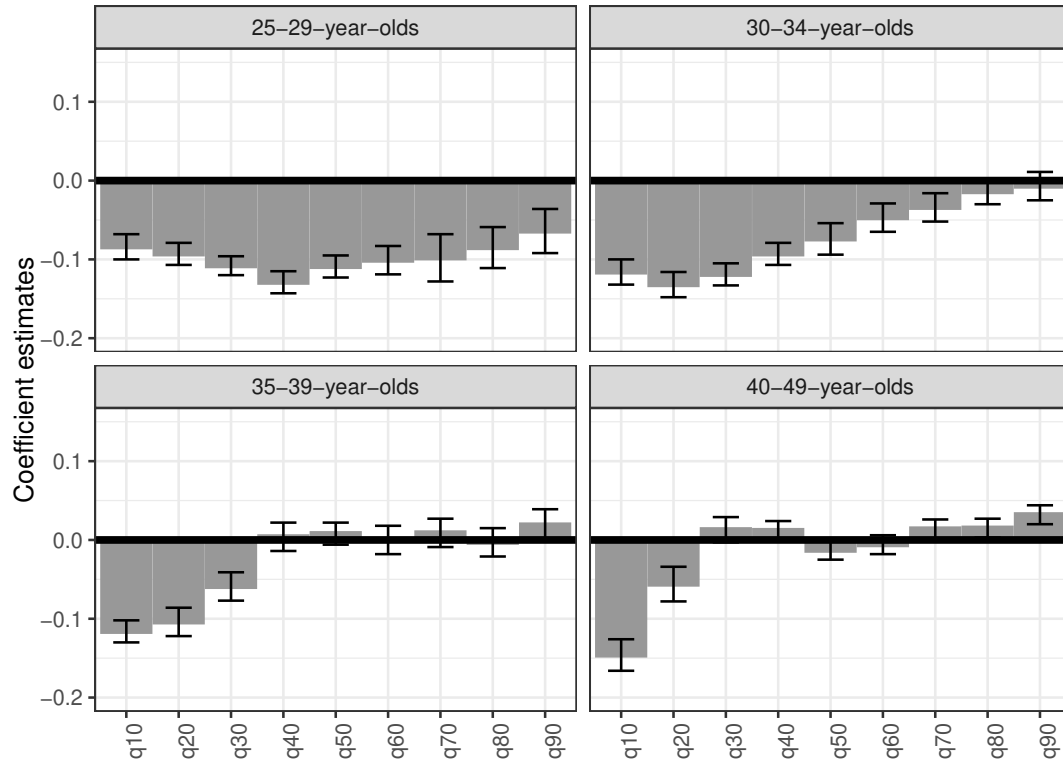
3.7.2. The effect of the expansion policy by quantile regression

In our conceptual framework, we conjecture that the fall in the college wage premium might result from either the increase in the supply of high-skilled labor beyond demand or the fall in the ability structure/value-added in higher education, or both. The lower productivity channel mostly affects the wages of those who attained education due to the additional educational opportunities. However, the effect of the expansion on the earnings of those who attained education in more prestigious universities is expected to be limited. Our sample consists of workers who attained education in well-established prestigious universities as well as those who studied in newly established universities with fewer educational resources. For this reason, we would expect a higher wage dispersion among younger college graduates if it is mainly the productivity channel that is driving our results (in Table 3.6). Here, we use quantile regression analysis to investigate whether the impact of the reform significantly varies along with the wage/ability distribution.

We estimate the impact of the expansion policy on the wages of workers at different quantiles, by similarly controlling for potential experience, captured by age, age squared, and firm tenure variables, time trend, in addition to region and year fixed effects. Figure 3.8 gives the coefficient estimates for the interaction term separately for each age group. Our findings indicate that the wage premium decreased significantly in all quantiles for the youngest age group exposed to the expansion policy. Hence, men in the upper tail of the wage distribution faced a lower wage advantage of their educational attainment in the post-expansion period. However, the decline is slightly smaller than those in the lower tail. However, our highest and lowest quantiles estimates are not statistically different, supporting the supply increase of skilled workers as the primary contributor to our results in Table 3.6.

Under the assumption that well-established universities do not suffer from

Figure 3.8: Distributional effects of expansion policy by unconditional quantile regression analysis



Source: HLFS 2005-2017, OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: The dependent variable is the log of hourly wages. The points shows the coefficient estimate for the interaction of exposure with the binary education variable. All regressions include controls for a constant, age and age squared, firm tenure, linear pre-trend, education dummy, continuous exposure variable, region fixed effects, and year fixed effects. The sample restricted to full-time paid employees with at least a high school degree of aged given in each panel. The sample includes observations from 2005-2010 and 2012-2017 HLFS. The standard errors are bootstrapped and error bars shows the 95% confidence intervals of the coefficients.

quality of students and teaching faculty, if lower productivity is the only driving channel for the fall in the college wage premium, then we would not observe any fall in the highest quantiles, which consists of higher ability individuals who are more likely to attain their education in these universities. Hence, our finding that the exposure reduces the relative wage advantage of college graduate men in all quantiles of the wage distribution leads to two possible explanations: First, the lower productivity is the result of the expansion in all types of higher education institutions, i.e., there exists a fall in the ability structure of college entrants and/or quality of education even in the most prestigious universities. Second, the increase in the supply of high skilled workers beyond demand is the primary mechanism at work in reducing college wage premium. Even though we show that there is not any observable change, on average, in the teaching faculty in already existing universities (see Figure 3.2), it is not possible for us to detect whether the ability distribution of college students in these universities changed over time. (The university entrance exam results are not comparable across years since the score calculation rules and the difficulty of the questions may change from year to year.)

Our estimates on the youngest age group not exposed to the expansion policy (30-34) similarly show the adverse effects of expanding the supply of skilled workers on the relative wage advantage of college graduate men compared to their high school graduate peers at almost all quantiles. These estimates are insignificant only on the upper tail of the wage distribution. Similarly, we find statistically significant negative impacts for the older age groups only on the lower tail of the wage distribution, and the effect of the expansion becomes insignificant on the upper tail. A more striking feature deduced from our findings is that the magnitude of the negative impact we reported for the older groups in the lower tail is very close to our estimates on the impact on the youngest age group, suggesting that the compliers of the policy are only close substitutes to the workers in the lower tail of the wage distribution for the older age groups.

3.8. Robustness Checks

Here, we conduct three key robustness checks using different samples and specifications. In the first one, we explore the impact of the expansion policy by changing our definition of exposure. Next, we vary the time window around the policy cutoff, and then restrict the full sample to sub-samples by excluding the

areas where the labor market might differ from the rest of the country. Finally, in the third one, we control for occupation and industry specifics of jobs to show whether the college wage premium can be partially explained by the segregation of workers by their level of education.

3.8.1. Alternative measure of exposure

Our measure of exposure relies on the increase in the available slots in both public and private universities. Here, to eliminate the concern that the available slots in private universities could be more demand-driven than those in public universities, we introduce an alternative measure of exposure, defined as the available slots in public universities per 18-year-olds, called *public exposure*. We replicate our regression analysis by replacing the exposure variable with public exposure. Our results, given in the even columns of Table 3.7, show slightly smaller coefficient estimates (in magnitude) than our baseline results given in odd columns of Table 3.7, but with the same patterns of statistical significance.

Table 3.7: Robustness check with public exposure

Ages:	25-29		30-34		35-39		40-49	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Education×exposure	-0.092*** (0.018)		-0.069*** (0.013)		-0.027** (0.010)		-0.022** (0.010)	
Education×Public exposure		-0.082*** (0.018)		-0.062*** (0.015)		-0.026** (0.011)		-0.026** (0.011)
Observations	60,202	60,202	62,768	62,768	52,612	52,612	75,325	75,325

Notes: The dependent variable is the log of hourly wages. All regressions include controls for a constant, age and age squared, firm tenure, linear pre-trend, education dummy, continuous exposure (public exposure) variable in odd columns (even columns), region fixed effects, and year fixed effects. The sample restricted to full-time paid employees with at least a high school degree observed in the waves 2005-2017. The observations from year 2011 are excluded from all regressions. The standard errors clustered at region 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.

3.8.2. Alternative sample specifications

In Table 3.6 that presents our baseline results, we provide the estimates on the main sample that includes six years around the policy cutoff, and also on the narrower sample trimmed down the time window that includes four years around the policy year 2011 to exclude the possible impact of macroeconomic conditions and other policy interventions observed in the years that are of interest to us.

One possible concern one might consider is the effect of the coup attempt in 2016 on the labor market. Following the failing attempt, a state of emergency was declared in the country. Many people with the suspicion of being related to the parties organizing the coup attempt were dismissed from their jobs with a series of decree laws. To the best of the authors' knowledge, no study in the literature identifies the impact of the coup attempt on the labor market, if any. However, to eliminate the concern that it might shift the structure of our sample of interest, men, and women employed with an observable wage, we focus on the years prior to 2016, as an alternative sample specification.

On the other hand, one possible policy intervention that could deteriorate our results is the minimum wage hike in 2016. Turkey experienced a 33% increase in the minimum wage in January 2016, following the promise of the governing party in general election campaigns in 2015. Turkey has one of the highest minimum wage-to-average wage ratios among OECD countries (Işık et al. 2020) making minimum wage an important determinant in the wage differentials. According to the findings of a recent study by Işık et al. (2020), there is no significant impact of the minimum wage hike on employment. However, a substantial positive impact on wages is identified that is also amplified with either education for the working-age population and age for the sample of all education levels. Even though their findings come short on answering whether the minimum wage jump increases or decreases the college wage premium of our sample of different age groups, it certainly signals some impact on the wage differentials. Hence, we aim to control whether our results remain the same in the sample of years prior to the government's minimum wage intervention beyond the current inflation rate.

Here, we focus on these two samples of interest, and further investigate whether our results are robust to restricting these sample geographically.

First, we consider the possible concern that our baseline results on the effect of the expansion policy on college wage premium would be different in the metropolitan areas than in the rest of the country. Between 2005 and 2010, referred to as pre-expansion years in our sample, among 25-29-year-olds with secondary education attainment, the share of college graduates was 50% in the three largest cities, namely Ankara, Istanbul, and Izmir, while it was 43% in the rest of the country. In addition to that, in the three largest cities, college graduates earned 98% more on average than high school graduates, while they only earned

80% more in the other cities. Hence, one could argue that the effects of the increase in the available slots would be stronger in the smaller cities due to lower college wage premium observed prior to the expansion. We exclude the three largest cities, restrict our sample to *smaller cities*, and examine the impact of the expansion in this sample.

Another concern about our baseline analysis may be that Turkey was the largest recipient of Syrian migrants displaced due to the Syrian civil war, hosting the highest number of refugees in the world since 2015. The number of registered Syrians rose to 2.5 million or 3.2% of the native population in 2015. The earlier literature examines the impact of this massive influx of Syrian migrants on the labor market outcomes of the natives by exploiting the variations in the share of migrants among the native population across NUTS-2 regions (Ceritoglu et al. 2017, Del Carpio & Wagner 2015, Aksu et al. 2018*b*). Aksu et al. (2018*a*), the study using the most recent datasets shows that the total employment of women decreased with the arrival of Syrian refugees. Hence, one could argue that this could affect our estimation results, especially for women, by changing the structure of our sample of full-time paid employees. On the other hand, due to the need for social assistance in the regions with higher share of Syrian migrants, the demand for high skilled workers might increase leading to a higher college wage premium in those regions. To avoid these concerns, we exclude the five regions with the highest share of Syrian refugees over time; in other words the regions that were identified as treated regions in Aksu et al. (2018*a*)¹³, and replicate our estimations for this sample, called *low refugee cities*.

Table 3.8 present our baseline results on the full sample in the first panel, on the smaller cities sample in the second panel, and finally on low refugee cities in the third panel.

First, we point out that the estimated effect of the reform on the college premium in men's wages is very similar in the samples including different time windows around the policy cutoff. The similar coefficients for men observed in years 2005-2017 and 2007-2015 suggest that the fall in the college wage premium of younger men cannot only be explained by the minimum wage increase. Therefore,

¹³By Aksu et al. (2018*a*), the regions with the highest share of Syrian refugees ranked by their share are Gaziantep (TRC1), Hatay (TR63), Sanliurfa (TRC2), Adana (TR62), and Mardin (TRC3).

Table 3.8: Robustness check with alternative sample specifications

Ages:	25-29		30-34		35-39		40-49	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Full sample	-0.092*** (0.018)	-0.079*** (0.022)	-0.069*** (0.013)	-0.057** (0.021)	-0.027** (0.010)	0.008 (0.009)	-0.022** (0.010)	-0.031** (0.013)
Observations	60,202	41,409	62,768	43,158	52,612	35,422	75,325	50,171
Smaller cities	-0.089*** (0.021)	-0.076*** (0.026)	-0.070*** (0.015)	-0.053** (0.023)	-0.033*** (0.010)	0.006 (0.010)	-0.019 (0.012)	-0.031** (0.013)
Observations	42,841	29,242	45,241	30,855	39,001	25,961	56,592	50,171
Low refugee cities	-0.082*** (0.017)	-0.068*** (0.022)	-0.063*** (0.013)	-0.051** (0.021)	-0.020* (0.011)	0.018* (0.010)	-0.018 (0.010)	-0.031** (0.013)
Observations	52,223	35,983	54,715	37,626	45,858	30,896	65,744	50,171
Sample:	2005-2017	2007-2015	2005-2017	2007-2015	2005-2017	2007-2015	2005-2017	2007-2015

Notes: The dependent variable is the log of hourly wages. All regressions include controls for a constant, age and age squared, firm tenure, linear pre-trend, education dummy, continuous exposure variable, region fixed effects, and year fixed effects. The sample restricted to full-time paid employees with at least a high school degree observed in the waves 2005-2017. The observations from year 2011 are excluded from all regressions. The standard errors clustered at region level are given in round parentheses with the patterns of p-values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$.

we observe a reduced wage premium even in the absence of the minimum wage hike. The larger decline in the premium in the wider time windows could be due to either the increasing effect of the expansion over the years or the effect of the minimum wage increase experienced in the later years.

In the second panel, where we present our results for the sample of smaller cities, we observe that the causal impact is statistically significant for younger men, with more mixed results for the older workers. Our main findings on the causal impact of the reform on college wage premium of aged 25-29 and 30-34 are robust to restricting the sample to smaller cities.

Finally, in the third panel, where we check the robustness of our results to restricting the sample to low refugee regions, we observe that the adverse effects are lower in the low refugee cities compared to our baseline results on the full sample.

3.8.3. Controlling for job and workplace characteristics

Skill biased technological change is described as the shift in the production technology that favors skilled over unskilled labor by increasing its relative productivity. This demand shift could lead to the polarization of college graduates into some occupations and industries. Here, we investigate whether our results change once we control for job and workplace characteristics, which could be closely related to the labor demand variations, by including first the occupation dummies in the second panel, next, the industry dummies in the third panel, and finally,

both for occupation and industry dummies in the fourth panel of Table 3.9. We also present our baseline analysis in the first panel for easy reference. We observe that the coefficient estimates are relatively smaller than baseline analysis once we include the additional controls, which suggests that part of the wage differentials between college and high school graduates could result from the segregation of paid employees into occupations and industries by their level of education. The change induced by controlling for industries is more sizable. Notwithstanding, the fall in the college wage premium due to the reform is statistically significant for all age groups after controlling for job and workplace characteristics.

Table 3.9: Robustness check with occupation and workplace controls

Ages:	25-29		30-34		35-39		40-49	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Baseline results	-0.092*** (0.018)	-0.079*** (0.022)	-0.069*** (0.013)	-0.057** (0.021)	-0.027** (0.010)	0.008 (0.009)	-0.022** (0.010)	-0.031** (0.013)
Observations	60,202	41,409	62,768	43,158	52,612	35,422	75,325	50,171
Occupation controls	-0.073*** (0.015)	-0.061*** (0.019)	-0.048*** (0.011)	-0.037* (0.018)	-0.041*** (0.008)	-0.014 (0.008)	-0.023** (0.009)	-0.029** (0.012)
Observations	60,202	41,409	62,768	43,158	52,612	35,422	75,325	50,171
Industry controls	-0.067*** (0.012)	-0.056*** (0.017)	-0.058*** (0.010)	-0.045*** (0.016)	-0.034*** (0.007)	-0.008 (0.009)	-0.029*** (0.008)	-0.037*** (0.011)
Observations	60,202	41,409	62,768	43,158	52,612	35,422	75,325	50,171
Occupation and industry controls	-0.058*** (0.011)	-0.048*** (0.015)	-0.043*** (0.009)	-0.029* (0.015)	-0.043*** (0.006)	-0.020** (0.008)	-0.027*** (0.008)	-0.031*** (0.011)
Observations	60,202	41,409	62,768	43,158	52,612	35,422	75,325	50,171
Sample:	2005-2017	2007-2015	2005-2017	2007-2015	2005-2017	2007-2015	2005-2017	2007-2015

Notes: The dependent variable is the log of hourly wages. All regressions include controls for a constant, age and age squared, firm tenure, linear pre-trend, education dummy, continuous exposure variable, region fixed effects, and year fixed effects. The sample restricted to full-time paid employees with at least a high school degree observed in the waves 2005-2017. The observations from year 2011 are excluded from all regressions. The standard errors clustered at region level are given in round parentheses with the patterns of p-values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$.

Many studies in the literature point out the labor market institutions, such as the unionization, firing, and minimum wage regulations, as important contributors to the educational wage differentials (Lee 1999, Koeniger et al. 2007). The effect of the reform might differ for the workers employed in the private and public sector since the payment arrangements are more centralized in the public sector and the labor market institutions are more effective than the private sector, whose equilibrium dynamics are closer to the free market (Grimshaw 2000).

The information on the public or private sector employment of individuals is available in HLFS starting from 2009. Hence, we exclude the years before 2009 and re-estimate the impact of the reform with public and private sector controls to capture the variations in the educational wage differentials due to the sector of employment. Table 3.10 gives the replication of baseline analysis for the restricted sample in odd columns, by controlling public sector and private sector employment

dummies in columns even columns. The causal impact of the reform is negative and significant for men in all columns, except for the sixth one, which suggests the additional controls does not change the qualitative power of our results.

Table 3.10: Robustness check with sector controls

Ages:	25-29		30-34		35-39		40-49	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Education $\times$ exposure	-0.078*** (0.019)	-0.055*** (0.014)	-0.063*** (0.015)	-0.050*** (0.014)	-0.031*** (0.009)	-0.001 (0.008)	-0.029** (0.011)	-0.038*** (0.012)
Observations	40,528	31,269	45,699	34,083	39,785	28,790	54,221	39,553
Control for: Sector dummies	$\times$	$\checkmark$	$\times$	$\checkmark$	$\times$	$\checkmark$	$\times$	$\checkmark$

Notes: The dependent variable is the log of hourly wages. All regressions include controls for a constant, age and age squared, firm tenure, linear pre-trend, education dummy, continuous exposure variable, region fixed effects, and year fixed effects. The sample restricted to full-time paid employees with at least a high school degree observed in the waves 2009-2017. The observations from year 2011 are excluded from all regressions. The standard errors clustered at region level are given in round parentheses with the patterns of p-values denoted by * for $p < 0.1$, ** for $p < 0.05$, and *** for $p < 0.01$.

3.9. Conclusion

Many countries experienced inflation in the number of college graduates at varying times in different countries, which has led to dramatic changes in the labor market structure all around the world. While an increasing wage dispersion between college and high school graduates documented in the US (Bound & Johnson 1995, Katz & Murphy 1992), flatter educational wage differentials are observed in European countries as the supply of college graduates expanded gradually.

Earlier studies established the relation of wage differentials with demand increase -due to skill-biased technological change- and labor market institutions in the age of rapid increase of college graduates. However, the impact of supply increase on the college wage premium, the additional wage enjoyed by college graduates compared to their high school graduate peers got very little attention, partly because establishing a causal link between supply and wage differentials is hard to investigate owing to data limitations. The higher education expansion in Turkey provides an excellent case to identify the role of the supply increase in the educational wage differentials.

The country experienced a politically driven higher education expansion occurred by the establishment of new universities all around the country, and the increase in educational opportunities in all higher education institutions. It has experienced a fall in the college wage premium of young adults parallel to the

entering of young college graduates exposed to the expansion to the labor market. The aim of this paper is to address this recent fall in the college wage premium. In doing so, we investigate the role of the recent higher education expansion policy, exogenous to existing labor market conditions and macroeconomic conditions, on the fall in the college wage premium of young adults.

Our findings indicate a significant fall in both college wage premium of 25-29-year-old men, exposed to the expansion, by 9%. We conjecture that this fall could be attributed to either the fall in the productivity of college graduates who attained their education in the post-expansion period, or the sudden jump in the number of college graduates, or both.

We further investigate the causal effect of the expansion on the wage premiums of the older age groups in order to determine if the rise in the supply of skilled worker is a mechanism driving our results for the younger workers. These employees were not expected to have lower productivity in the post expansion years since their human capital accumulation occurs prior to the establishment of new universities and expanding the educational opportunities in the existing universities. Because younger and older graduates are not close substitutes for one another, given their varied experience profiles in the labor market, we only anticipate labor market equilibrium impacts for these cohorts, and even then, these benefits are probably going to be minimal. Our findings confirm the labor market equilibrium channel by showing a negative, statistically significant, but decreasing effect with age.

To explore the impact through lower productivity channel, we examine the additional wage that college graduates earn compared to their high school graduate peers at different quantiles of the wage distribution. We find that college wage premium decreases similarly along the entire wage distribution for younger age groups, while the negative impact vanishes for older age groups on the upper tail of the wage distribution. The relatively homogeneous effect of the policy on younger workers, compliers of the policy, indicates either the fall in the productivity of recent graduates is not a contributor to our findings or that all post-expansion graduates are subject to the same productivity fall. Hence, our investigation of the productivity channel is not conclusive.

CHAPTER 4

GENDER GAP IN INTERGENERATIONAL EDUCATIONAL PERSISTENCE: CAN COMPULSORY SCHOOLING REDUCE IT?

4.1. 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 has 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 towards sons and daughters can also

¹Hertz 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 fifty years.

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 eight 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 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 (Rosenzweig & Wolpin 1994, Behrman et al. 1994, Currie & Moretti 2003, 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 offspring's education. Studies in this literature are all from advanced economies with high levels of income equality and intergenerational educational mobility (Meghir & Palme 2005, Aakvik et al. 2010, Betthäuser 2017). 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 (Spohr 2003, Fang et al. 2012, Tsai et al. 2009, Dayioglu et al. 2016). Several studies identified differential preferences of parents for their sons and daughters as a source of women's lower educational attainment in Turkey (Rankin & Aytac 2006, Caner et al. 2016), 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 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: Section 2 and 3 introduce the compulsory schooling reform and the conceptual framework. Section 4 and 5 explain the data and empirical strategy. Section 6 presents and discusses the results and robustness checks. Finally, Section 7 states the concluding remarks.

4.2. 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 five years of primary education, followed by voluntary three years of lower secondary and three 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

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

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 five 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 five years of primary education to eight years by incorporating junior high school into compulsory education and delayed starting a vocational school including religious vocational schools by three 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 eight 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 five 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 three years (Aydemir & Kirdar 2017b).

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 2017b, 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 percent, while prior to compulsory schooling law, it was about 15 percent (Dulger 2004, Dayioglu et al. 2016).

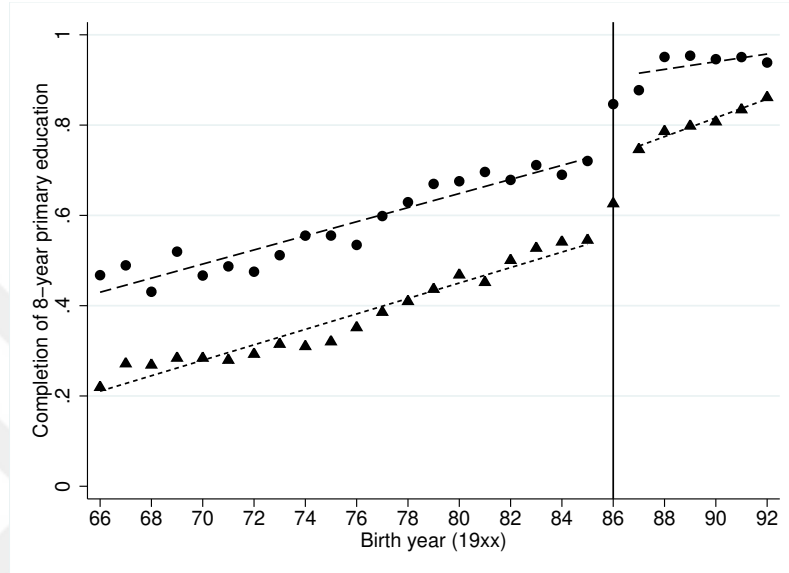
The compulsory education law came into force in August 1997, just before the 1997-1998 school year. Students who had not completed their 5-year education on this date and had not been awarded the primary school diploma are required to complete 8-years of schooling to earn a primary education diploma. 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 three 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).

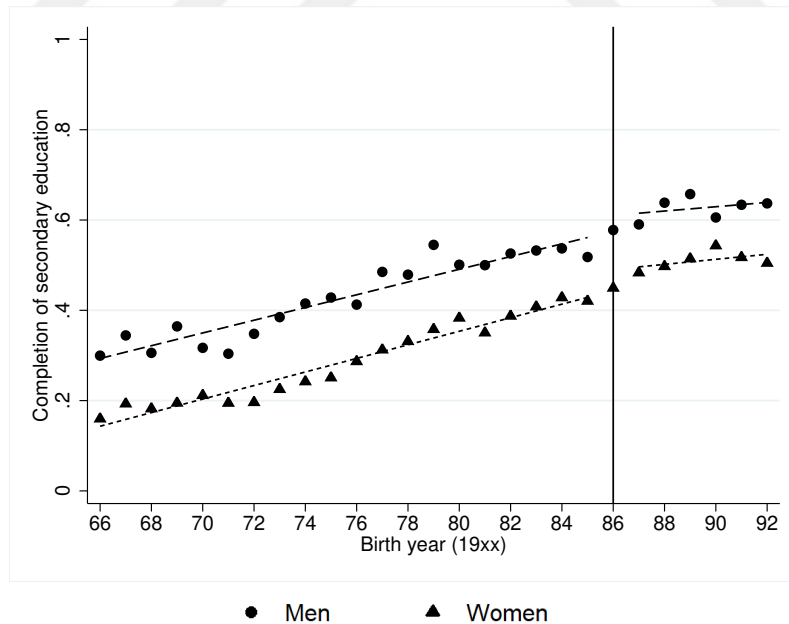
We further present the change in school attainment with the nationally representative micro-level data from Turkish Adult Education Survey. Panel (a) of Figure 4.1 displays the share of individuals who completed at least 8-year primary education by birth year separately for men and women. Separate linear 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

Figure 4.1: Mean education completion rates by birth year

(a) 8-year primary education



(b) Secondary education



● Men ▲ Women

Source: Adult Education Survey, 2012 and 2016.

Notes: 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.

trends for pre and post-policy cohorts, confirming their partial exposure. The increasing trends are similar for both genders 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 Figure 4.1, 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.

4.3. 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 modelled in two theories in the literature. The human capital theory suggests that education is an investment activity raising future earnings by increasing worker productivity (Mincer 1974, Becker 1975). 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, 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.

Table 4.1: Educational completion percentages of offspring by paternal 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 educ	92%	100%	97%	100%	88%	99%
Completing 8-year educ	48%	94%	60%	96%	39%	92%
Completing secondary educ	35%	88%	43%	89%	28%	87%
Completing higher educ	14%	61%	17%	62%	12%	60%
Observations	9,894	1,750	4,185	809	5,709	941

Notes: 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.

We present completion rates of five years of compulsory education, eight years of education and secondary education for pre-policy cohorts born in 1976-1985 according to their parental education levels in Table 4.1. While those with educated parents almost perfectly comply with five years of compulsory education,

³Jensen (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.

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 eight 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 eight 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 are 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 & 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 travelling 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 & 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 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.

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

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 4.1). Among those not exposed to the extension of compulsory schooling (individuals born in 1976-1985), the completion rate of five 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 eight 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 4.1.

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 noncompliance⁵, it increased the cost of not attending the new compulsory level of education. Hence, these penalties can be interpreted as the negative cost of

⁵Parents 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 (Dayıoglu et al. 2016).

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 & 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 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 non-compliance 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 three 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 five 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 five years of education in the absence of the policy to complete at least eight 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 eight years of education even before the enactment of the law (with 96 percent for men and 92 percent 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 eight 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 three 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.

4.4. 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 this 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.

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

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

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 4.2 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 4.2, 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. 50 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

Table 4.2: 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)
<i>Panel A: Education outcomes</i>				
Completed 5-year education	0.908 (0.289)	0.961 (0.194)	0.868 (0.338)	0.093 (0.005)
Completed 8-year education	0.710 (0.454)	0.809 (0.393)	0.636 (0.481)	0.173 (0.008)
Completed secondary education	0.498 (0.5)	0.569 (0.495)	0.446 (0.497)	0.123 (0.009)
<i>Panel B: Parental education outcomes</i>				
Father's (secondary) education	0.174 (0.379)	0.190 (0.393)	0.161 (0.368)	0.029 (0.007)
Father's higher education	0.066 (0.248)	0.074 (0.263)	0.060 (0.237)	0.014 (0.004)
Mother's (secondary) education	0.076 (0.266)	0.084 (0.277)	0.071 (0.257)	0.013 (0.005)
Mother's higher education	0.024 (0.154)	0.027 (0.163)	0.022 (0.146)	0.005 (0.003)
Either parent has at least secondary education	0.188 (0.391)	0.206 (0.405)	0.175 (0.38)	0.031 (0.007)
Both parents have at least secondary education	0.062 (0.241)	0.068 (0.252)	0.058 (0.233)	0.010 (0.004)
Either parent has at least higher education	0.072 (0.259)	0.080 (0.272)	0.066 (0.249)	0.014 (0.004)
Both parents have at least higher education	0.018 (0.133)	0.021 (0.145)	0.015 (0.123)	0.006 (0.002)
Observations	13,807	5,896	7,911	

Notes: 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.

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.

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' make it the primary determinant of both parents' educational attainment.

An unusual feature observed in Table 4.2 is that parental education is systematically lower for daughters than sons. Though the available data does 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.

⁷Altindag (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.

4.5. 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 eight years of schooling, while those born in 1986 or earlier could drop out after five 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:

$$y_i = \beta d_i + f(t_i) + \alpha X_i + \epsilon_i \quad (4.1)$$

$$\forall t_i \in (-h, +h)$$

$$d_i = \begin{cases} 1 & t_i > 0 \\ 0 & t_i \leq 0 \end{cases}$$

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

⁸Several 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 & Kirdar (2017b).

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 six years around the cutoff year 1986.

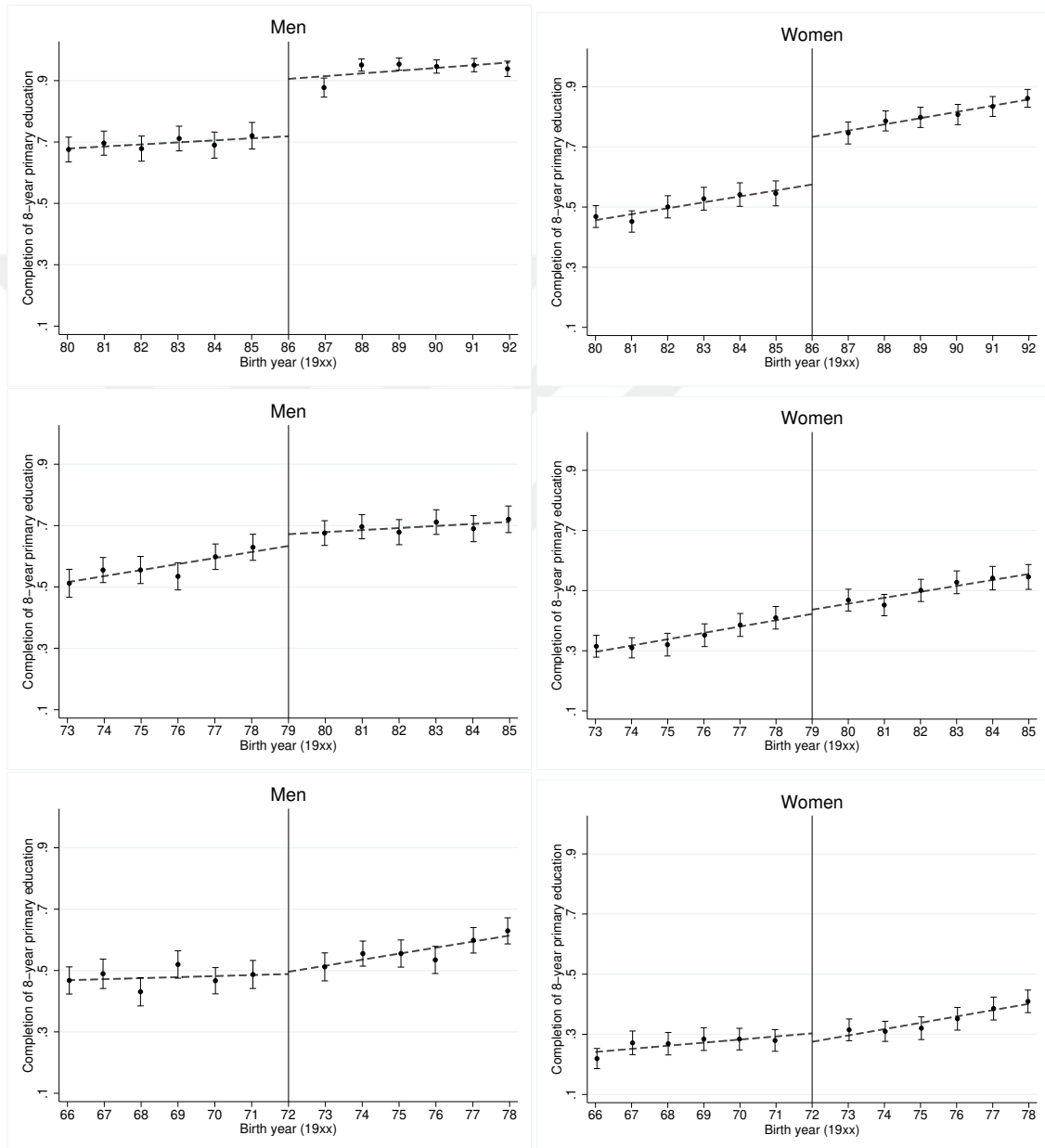
The outcome variables in this study are two binary educational attainment variables: completion of 8-year primary education (new compulsory level) and completion of secondary education (high school). The former and the latter are employed to capture the direct and spillover effects of the reform, respectively. 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*b*).

We visualize our identification strategy, in the first panel of Figure 4.2, by showing the mean completion of 8-year primary education rates separately for men and women born six 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 Figure 4.2, 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,

Figure 4.2: Mean education completion rates by birth year in treatment and placebo samples



Source: Adult Education Survey, 2012 and 2016.

Notes: 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.

and our results are given in the 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 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 the equation (4.1) 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:

$$y_i = \beta_1 p_i + \beta_2 d_i + \beta_3 p_i d_i + \alpha X_i + \epsilon_i \quad (4.2)$$

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 equation (4.2) 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:

$$y_i = \gamma_0 w_i + \gamma_1 p_i + \gamma_2 w_i p_i + \gamma_3 d_i + \gamma_4 w_i d_i + \gamma_5 p_i d_i + \gamma_6 w_i p_i d_i + \theta_1 X_i + \theta_2 w_i X_i + \nu_i \quad (4.3)$$

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 4.2.

In equation (4.3), 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 equation (4.3) 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 4.1). 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* $\times$ *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. (2004b), 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 the Appendix.

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

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.

4.6. Results

4.6.1. The impact of policy on intergenerational educational persistence

We first estimate equation (4.2) separately for men and women with alternative parental education variables as our key explanatory variable and present our results in Table 4.3. 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 shows 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 $\text{parental education} \times \text{policy}$ 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.

Table 4.3: The effect of compulsory schooling on the intergenerational education transmission

<i>Dependent var: Completion of 8-year primary education</i>						
	Men			Women		
Father's education	0.323*** (0.008) [0.0000]			0.494*** (0.020) [0.0000]		
Father's educ.×Policy	-0.249*** (0.015) [0.0000]			-0.306*** (0.025) [0.0000]		
Mother's education		0.220*** (0.050) [0.0010]			0.413*** (0.038) [0.0000]	
Mother's educ.×Policy		-0.149** (0.052) [0.0000]			-0.217*** (0.039) [0.0000]	
Parental education			0.314*** (0.011) [0.0250]			0.458*** (0.019) [0.0000]
Parental educ.×Policy			-0.239*** (0.016) [0.0000]			-0.284*** (0.024) [0.0000]
Observations	5,401	5,401	5,401	7,277	7,277	7,277
<i>Dependent var: Completion of secondary education</i>						
	Men			Women		
Father's education	0.452*** (0.016) [0.0150]			0.548*** (0.022) [0.0130]		
Father's educ.×Policy	-0.089*** (0.020) [0.0030]			-0.088*** (0.028) [0.0060]		
Mother's education		0.452*** (0.021) [0.0150]			0.590*** (0.020) [0.0180]	
Mother's educ.×Policy		-0.093*** (0.023) [0.0020]			-0.098*** (0.022) [0.000]	
Parental education			0.450*** (0.014) [0.0220]			0.550*** (0.023) [0.0140]
Parental educ.×Policy			-0.084*** (0.020) [0.0030]			-0.075** (0.028) [0.0180]
Observations	5,401	5,401	5,401	7,277	7,277	7,277
Birth cohorts	1980-1992	1980-1992	1980-1992	1980-1992	1980-1992	1980-1992

Notes: 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.

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.

4.6.2. 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 equation (4.3) 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. We present our results in Table 4.4 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* $\times$ *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

Table 4.4: The effect of compulsory schooling on the gender gap in intergenerational education transmission

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

Notes: 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.

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, $woman \times policy \times parental\ education$) are 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 eight 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 & 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 4.4, 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 4.4 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 4.1). 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 eight 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.

4.6.3. 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 equation (4.3) 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 4.5 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

Table 4.5: The effect of compulsory schooling on the gender gap in intergenerational education transmission: Heterogeneity analysis by the number of educated parent

<i>Dependent var:</i>	<i>Completion of 8-year primary educ</i>				<i>Completion of secondary educ</i>			
		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 educ.	0.322*** (0.009) [0.0220]	0.496*** (0.019)			0.450*** (0.014) [0.0280]	0.550*** (0.023)		
Parental educ× Woman	0.174*** (0.018) [0.0000]				0.099*** (0.023) [0.0160]			
Parental educ× Policy	-0.246*** (0.016) [0.0000]	-0.300*** (0.027)			-0.084*** (0.020) [0.0010]	-0.075** (0.028)		
Parental educ× Woman×Policy	-0.054** (0.022) [0.0460]				0.009 (0.024) [0.7117]			
One parent is educ.			0.304*** (0.012) [0.0250]	0.459*** (0.019)			0.409*** (0.015) [0.0100]	0.489*** (0.028)
Two parents are educ.			0.358*** (0.004) [0.0000]	0.563*** (0.019)			0.534*** (0.016) [0.0100]	0.660*** (0.019)
One parent is educ.× Woman			0.154*** (0.020) [0.0000]				0.080** (0.031) [0.0360]	
Two parents are educ× Woman			0.205*** (0.020) [0.0000]				0.125*** (0.017) [0.0220]	
One parent is educ× Policy			-0.228*** (0.017) [0.0000]	-0.278*** (0.025)			-0.077*** (0.025) [0.0260]	-0.052 (0.035)
Two parents are educ× Policy			-0.282*** (0.017) [0.0000]	-0.333*** (0.025)			-0.104*** (0.020) [0.0000]	-0.096*** (0.023)
One parent is educ.× Woman×Policy			-0.049* (0.023) [0.0871]				0.024 (0.035) [0.5205]	
Two parents are educ.× Woman×Policy			-0.051* (0.026) [0.0881]				0.009 (0.027) [0.7548]	
Observations	12,678	12,678	12,678	12,678	12,678	12,678	12,678	12,678
Birth cohorts	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92

Notes: 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.

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 equation (4.3) with these two variables and add a full set of interactions of these variables with policy and woman dummies. Table 4.6 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 4.6: The effect of compulsory schooling on the gender gap in intergenerational education transmission: Heterogeneity analysis by the level of parental education

<i>Dependent var:</i>	<i>Completion of 8-year primary education</i>				<i>Completion of secondary education</i>			
		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 educ	0.322*** (0.009) [0.0220]	0.496*** (0.019)			0.450*** (0.014) [0.0280]	0.550*** (0.023)		
Parental educ× Woman	0.174*** (0.018) [0.0000]				0.099*** (0.023) [0.0160]			
Parental educ× Policy	-0.246*** (0.016) [0.0000]	-0.300*** (0.027)			-0.084*** (0.020) [0.0010]	-0.075** (0.028)		
Parental educ× Woman× Policy	-0.054* (0.022) [0.0460]				0.009 (0.024) [0.7117]			
Secondary educ			0.314*** (0.011) [0.0380]	0.459*** (0.019)			0.415*** (0.022) [0.0090]	0.487*** (0.027)
Tertiary educ			0.334*** (0.009) [0.0240]	0.560*** (0.019)			0.505*** (0.016) [0.0060]	0.657*** (0.021)
Secondary educ× Woman			0.144*** (0.023) [0.0000]				0.072** (0.032) [0.00851]	
Tertiary educ× Woman			0.226*** (0.014) [0.0000]				0.152*** (0.017) [0.0310]	
Secondary educ× Policy			-0.239*** (0.016) [0.0000]	-0.284*** (0.024)			-0.084** (0.028) [0.0220]	-0.059* (0.033)
Tertiary educ× Policy			-0.257*** (0.018) [0.0000]	-0.330*** (0.025)			-0.088*** (0.020) [0.0010]	-0.105*** (0.025)
Secondary educ× Woman× Policy			-0.045 (0.028) [0.1672]				0.024 (0.033) [0.5385]	
Tertiary educ× Woman× Policy			-0.073*** (0.018) [0.0040]				-0.017 (0.024) [0.5035]	
Observations	12,678	12,678	12,678	12,678	12,678	12,678	12,678	12,678
Birth cohorts	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92

Notes: 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 eight years of education than men without an educated parent. Similarly, women whose either parent completed secondary or tertiary education are more likely to finish eight 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.

4.6.4. 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 B.1 shows the West versus East classification on a Turkey map. Next, we estimate equation (4.3), first on the pooled sample of West and East (in other words, by restricting the sample to AES, 2012), and then separately for two samples. Table 4.7 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).

Table 4.7: Regional differences on the effect of compulsory schooling on the gender gap in intergenerational education transmission

<i>Dependent var:</i>	<i>Completion of 8-year primary education</i>			<i>Completion of secondary education</i>		
	All (1)	West (2)	East (3)	All (4)	West (5)	East (6)
Woman	-0.141*** (0.022) [0.0230]	-0.150*** (0.031) [0.0541]	-0.130 (0.087) [0.3153]	-0.054** (0.019) [0.0180]	-0.040 (0.028) [0.1862]	-0.211** (0.078) [0.0470]
Policy	0.272*** (0.037) [0.0000]	0.268*** (0.033) [0.0000]	0.287*** (0.067) [0.0350]	0.073** (0.027) [0.0460]	0.085** (0.034) [0.0120]	0.031 (0.041) [0.6266]
Woman×Policy	-0.046 (0.030) [0.1992]	0.010 (0.034) [0.8048]	-0.194** (0.076) [0.0060]	-0.035 (0.029) [0.3674]	-0.029 (0.036) [0.4945]	-0.029 (0.057) [0.6316]
Parental education	0.314*** (0.023) [0.0000]	0.298*** (0.024) [0.0000]	0.386*** (0.041) [0.0010]	0.450*** (0.018) [0.0030]	0.428*** (0.018) [0.0070]	0.550*** (0.060) [0.0090]
Parental education×Woman	0.172*** (0.034) [0.0100]	0.183*** (0.037) [0.0050]	0.142* (0.073) [0.1502]	0.084** (0.033) [0.0631]	0.111*** (0.029) [0.0430]	-0.028 (0.089) [0.7768]
Parental education×Policy	-0.281*** (0.027) [0.0000]	-0.268*** (0.027) [0.0000]	-0.335*** (0.051) [0.0000]	-0.131*** (0.020) [0.0000]	-0.113*** (0.024) [0.0000]	-0.209** (0.067) [0.0150]
Parental educ×Policy×Woman	-0.065* (0.032) [0.0841]	-0.102** (0.037) [0.0100]	0.018 (0.078) [0.8168]	0.043 (0.036) [0.2953]	-0.002 (0.036) [0.9650]	0.193* (0.098) [0.1051]
Observations	7,756	5,883	1,873	7,756	5,883	1,873
Birth cohorts	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92

Notes: 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.

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

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 eight 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 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 leads 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 percent 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 Section A.1 of the 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.

4.6.5. 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 B.2 in the 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 B.2, 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 B.3 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 is 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 B.4 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.

4.7. 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.

CHAPTER 5

CONCLUSION

In this thesis, we examine the effects of two education policies on education and labor market outcomes in Turkey. First, we explore the causal effects of the drastic expansion in Turkish higher education, which occurred in 2006-2008 with the establishment of new universities and increasing the educational opportunities both in the new and established universities, on the gender gap in tertiary education attainment to the detriment of women by exploiting the region and cohort variations in the exposure intensity to the expansion policy using Turkish Household Labor Force Surveys (HLFS).

The results show that overall, the expansion policy increased the higher education attainment of both men and women but did not reduce the gender gap in attainment above and beyond the time trend. Furthermore, we find the intriguing result that the increase in available slots from newly established universities benefited men more than women, thereby having an increasing effect on the gender gap. To examine this perplexing result, we examine fields of study offered in these new universities and existing ones. While the increase in available slots is similar in old (existing) and new universities, the highest growth in available slots occurred in social sciences and engineering. We find that the expansion in social sciences benefited men and women almost evenly; therefore, it did not affect the gender gap. However, the expansion in engineering in newly established universities benefited men more than it did women, as it is a major traditionally preferred by men.

The literature on the effects of expansion policies on the socio-economic gap in higher education generally finds that expansion policies do not decrease the socio-economic gap (Blanden & Machin 2004, Reimer & Pollak 2010). In a similar

vein, we show that the expansion policy did not reduce the gender gap in higher education in Turkey. Hence, our results support the argument that the expansion policies as they have been designed may not have been the best tools to address inequalities across different socio-economic and demographic groups. Our findings on how the allocation of the additional slots to different fields of study has affected men versus women inform us that an expansion policy with a specific target to reduce inequality needs to consider how the preferences of compliers of the policy may vary across different socio-economic and demographic groups. As such, our findings can inform future expansionary policies in parts of the world where women are still under-represented in tertiary education.

In the second essay, we investigate the causal impact of the same expansion policy in higher education on college wage premium in Turkey. The expansion increases the share of college graduates significantly and decreases the college wage premium for young men exposed to the policy. This finding can be due to the labor market equilibrium effect of the expansion as the supply of college graduates increased while demand remained constant. It can also be due to lower productivity of new graduates as the expansion happened over a short period, at a large scale when resources in terms of faculty and infrastructure were limited and the sorting power of the university entrance exam according to ability has decreased.

To explore whether the channel of the supply increase in skilled workers is at work, we further explore the causal impact of the expansion on the wage premiums of the older age groups, who were not subject to the policy when they were most likely to make their college entry decision. These workers were not exposed to quality issues related to newly established universities or lower selectivity of the university admission process that younger cohorts were exposed to. Therefore, we expect only labor market equilibrium effects for these cohorts; moreover, such effects are likely to be limited since younger and older graduates are not close substitutes due to their different experience profiles in the labor market. Our findings indicate a negative, statistically significant, but decreasing effect with age, providing evidence on the labor market equilibrium channel. This result provides causal evidence that the increase in the supply of skilled workers reduces the wage advantages enjoyed by these workers. To disentangle the impact through lower productivity channel from labor market equilibrium effects, we examine the additional wage that college graduates earn compared to their high school graduate peers at different quantiles of the wage distribution. We find that college wage

premium decreases similarly along the entire wage distribution for younger age groups, while the negative impact vanishes for older age groups on the upper tail of the wage distribution.

Finally, in the third essay, we investigate the impact of extending mandatory years of schooling from 5 to 8 years in 1997 on the gender gap in intergenerational educational persistence. Prior to the reform, there was a gender gap in the association of parents' educational attainment with their offspring's. Daughters' educational attainment is more dependent on their parents' educational background. By leveraging the policy intervention extending compulsory years of schooling by three years, 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.

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 (Duflo 2012).

Bibliography

- Aakvik, A., Salvanes, K. G. & Vaage, K. (2010), ‘Measuring heterogeneity in the returns to education using an education reform’, *European Economic Review* **54**(4), 483–500.
- Abadie, A., Athey, S., Imbens, G. & Wooldridge, J. (2017), ‘When should you adjust standard errors for clustering?’, *National Bureau of Economic Research* .
- Aksu, E., Erzan, R. & Kirdar, M. (2018*a*), The impact of mass migration of syrians on the turkish labor market, Technical report, Working Paper.
- Aksu, E., Erzan, R. & Kirdar, M. (2018*b*), The impact of mass migration of syrians on the turkish labor market, Technical report, Working Paper.
- Akyol, P. & Okten, C. (2019), ‘The role of culture on female labor supply: Evidence from turkey’, *IZA Discussion Paper* .
- Altindag, O. (2016), ‘Son preference, fertility decline, and the nonmissing girls of turkey’, *Demography* **53**(2), 541–566.
- Angrist, J. D. & Keueger, A. B. (1991), ‘Does compulsory school attendance affect schooling and earnings?’, *The Quarterly Journal of Economics* **106**(4), 979–1014.
- Arap, K. S. (2010), ‘Türkiye yeni üniversitelerine kavuşurken: Türkiye’de yeni üniversiteler ve kuruluş gerekçeleri’, *Ankara Üniversitesi SBF Dergisi* **65**(3), 001–029.
- Attanasio, O. P. & Kaufmann, K. M. (2009), ‘School choices, subjective expectations, and credit constraints’, *NBER Working Paper* **15087**.
- Aydemir, A. & Kirdar, M. (2017*a*), ‘Low wage returns to schooling in a developing country: Evidence from a major policy reform in turkey’, *Oxford Bulletin of Economics and Statistics* **79**(6), 1046–1086.

- Aydemir, A. & Kirdar, M. (2017*b*), ‘Low wage returns to schooling in a developing country: Evidence from a major policy reform in turkey’, *Oxford Bulletin of Economics and Statistics* **79**(6), 1046–1086.
- Aydemir, A. & Yazici, H. (2019), ‘Intergenerational education mobility and the level of development’, *European Economic Review* **116**, 160–185.
- Barro, R. J. & Lee, J. W. (2013), ‘A new data set of educational attainment in the world, 1950–2010’, *Journal of development economics* **104**, 184–198.
- Becker, G. S. (1975), Investment in human capital: effects on earnings, in ‘Human Capital: A Theoretical and Empirical Analysis, with Special Reference to Education, Second Edition’, NBER, pp. 13–44.
- Behrman, J. R., Rosenzweig, M. R. & Taubman, P. (1994), ‘Endowments and the allocation of schooling in the family and in the marriage market: the twins experiment’, *Journal of political economy* **102**(6), 1131–1174.
- Berlinski, S., Galiani, S. & Gertler, P. (2009), ‘The effect of pre-primary education on primary school performance’, *Journal of public Economics* **93**(1-2), 219–234.
- Bertrand, M., Duflo, E. & Mullainathan, S. (2004*a*), ‘How much should we trust differences-in-differences estimates?’, *The Quarterly journal of economics* **119**(1), 249–275.
- Bertrand, M., Duflo, E. & Mullainathan, S. (2004*b*), ‘How much should we trust differences-in-differences estimates?’, *The Quarterly journal of economics* **119**(1), 249–275.
- Betthäuser, B. (2017), ‘Fostering equality of opportunity? compulsory schooling reform and social mobility in germany’, *European Sociological Review* **33**(5), 633–644.
- Björklund, A. & Salvanes, K. G. (2011), Education and family background: Mechanisms and policies, in ‘Handbook of the Economics of Education’, Vol. 3, Elsevier, pp. 201–247.
- Black, S. E., Devereux, P. J. & Salvanes, K. G. (2005), ‘Why the apple doesn’t fall far: Understanding intergenerational transmission of human capital’, *American economic review* **95**(1), 437–449.

- Blanden, J. & Machin, S. (2004), ‘Educational inequality and the expansion of uk higher education’, *Scottish Journal of Political Economy* **51**(2), 230–249.
- Bound, J. & Johnson, G. (1995), ‘What are the causes of rising wage inequality in the united states?’, *Economic Policy Review* **1**(1).
- Bratti, M., Checchi, D. & De Blasio, G. (2008), ‘Does the expansion of higher education increase the equality of educational opportunities? evidence from italy’, *Labour* **22**, 53–88.
- Callaway, B., Goodman-Bacon, A. & Sant’Anna, P. H. (2021), ‘Difference-in-differences with a continuous treatment’, *arXiv preprint arXiv:2107.02637*.
- Callaway, B. & Sant’Anna, P. H. (2021), ‘Difference-in-differences with multiple time periods’, *Journal of Econometrics* **225**(2), 200–230.
- Calonico, S., Cattaneo, M. D., Farrell, M. H. & Titiunik, R. (2017), ‘rdrrobust: Software for regression-discontinuity designs’, *The Stata Journal* **17**(2), 372–404.
- Cameron, A. C. & Miller, D. L. (2015), ‘A practitioner’s guide to cluster-robust inference’, *Journal of human resources* **50**(2), 317–372.
- Caner, A., Guven, C., Okten, C. & Sakalli, S. O. (2016), ‘Gender roles and the education gender gap in turkey’, *Social Indicators Research* **129**(3), 1231–1254.
- Caner, A. & Okten, C. (2010), ‘Risk and career choice: Evidence from turkey’, *Economics of Education Review* **29**(6), 1060–1075.
- Card, D. (1992), ‘Using regional variation in wages to measure the effects of the federal minimum wage’, *ILR Review* **46**(1), 22–37.
- Card, D. (1995), ‘Using geographic variation in college proximity to estimate the return to schooling, aspects of labour market behaviour: essays in honour of john vanderkamp. ed’, *Unpublished manuscript*.
- Card, D. (1999), ‘The causal effect of education on earnings’, *Handbook of labor economics* **3**, 1801–1863.
- Card, D. & DiNardo, J. E. (2002), ‘Skill-biased technological change and rising wage inequality: Some problems and puzzles’, *Journal of labor economics* **20**(4), 733–783.

- Card, D. & Lemieux, T. (2001), ‘Can falling supply explain the rising return to college for younger men? a cohort-based analysis’, *The Quarterly Journal of Economics* **116**(2), 705–746.
- Ceritoglu, E., Yunculer, H. B. G., Torun, H. & Tumen, S. (2017), ‘The impact of syrian refugees on natives’ labor market outcomes in turkey: evidence from a quasi-experimental design’, *IZA Journal of Labor Policy* **6**(1), 5.
- Chevalier, A. (2004), ‘Parental education and child’s education: A natural experiment’, *IZA discussion paper* .
- COHE (2004), *Türk Yükseköğreniminin bugünkü durumu*.
- COHE (2007), *Türkiye’nin Yükseköğrenim Stratejisi*.
- Crivellaro, E. (2014), College wage premium over time: trends in europe in the last 15 years, university of venice, Technical report, Department of Economics Working Paper 03.
- Currie, J. & Moretti, E. (2003), ‘Mother’s education and the intergenerational transmission of human capital: Evidence from college openings’, *The Quarterly journal of economics* **118**(4), 1495–1532.
- Dayıoğlu, M. & Kırdar, M. G. (2010), ‘Determinants of and trends in labor force participation of women in turkey’.
- Dayioglu, M., Kırdar, M. & Koc, I. (2016), ‘Does longer compulsory education equalize schooling by gender and rural/urban residence?’, *World Bank Economic Review* **30**(3).
- De Chaisemartin, C. & d’Haultfoeuille, X. (2020), ‘Two-way fixed effects estimators with heterogeneous treatment effects’, *American Economic Review* **110**(9), 2964–96.
- De Chaisemartin, C. & D’Haultfoeuille, X. (2022), Two-way fixed effects and differences-in-differences with heterogeneous treatment effects: A survey, Technical report, National Bureau of Economic Research.
- Del Carpio, X. V. & Wagner, M. (2015), *The impact of Syrians refugees on the Turkish labor market*, The World Bank.

- Delaney, J. & Devereux, P. J. (2019a), It's not just for boys! understanding gender differences in stem. CEPR Discussion Paper No. DP13558.
- Delaney, J. M. & Devereux, P. J. (2019b), 'Understanding gender differences in stem: Evidence from college applications', *Economics of Education Review* **72**, 219–238.
- Demirel, M. & Okten, C. (2020), 'Gender gap in intergenerational persistence in education', *Unpublished manuscript*.
- Devereux, P. J. & Fan, W. (2011a), 'Earnings returns to the british education expansion', *Economics of Education Review* **30**(6), 1153–1166.
- Devereux, P. J. & Fan, W. (2011b), 'Earnings returns to the british education expansion', *Economics of Education Review* **30**(6), 1153–1166.
- Devereux, P. J. & Hart, R. A. (2010), 'Forced to be rich? returns to compulsory schooling in britain', *The Economic Journal* **120**(549), 1345–1364.
- Di Pietro, G. & Cutillo, A. (2008), 'Degree flexibility and university drop-out: The italian experience', *Economics of Education Review* **27**(5), 546–555.
- Dinçer, M. A., Kaushal, N. & Grossman, M. (2014), 'Women's education: Harbinger of another spring? evidence from a natural experiment in turkey', *World Development* **64**, 243–258.
- Dougherty, C. (2005), 'Why are the returns to schooling higher for women than for men?', *Journal of Human Resources* **40**(4), 969–988.
- Duflo, E. (2001), 'Schooling and labor market consequences of school construction in indonesia: Evidence from an unusual policy experiment', *American economic review* **91**(4), 795–813.
- Duflo, E. (2012), 'Women empowerment and economic development', *Journal of Economic literature* **50**(4), 1051–79.
- Dulger, I. (2004), Turkey: Rapid coverage for compulsory education—the 1997 basic education program, in 'Scaling Up Poverty Reduction: A Global Learning Process and Conference, Shanghai', pp. 25–27.
- Duta, A. (2016), Iannelli. cristina, jacob. marita and emer smyth (2019).” the institutional embeddedness of gendered choices of field of study, in 'Rediscovering Inequalities: AQMeN International Conference, Edinburgh'.

- Ehrlich, I., Cook, A. & Yin, Y. (2018), ‘What accounts for the us ascendancy to economic superpower by the early twentieth century? the morrill act–human capital hypothesis’, *Journal of Human Capital* **12**(2), 233–281.
- Emerson, P. M. & Souza, A. P. (2007), ‘Child labor, school attendance, and intrahousehold gender bias in brazil’, *The World Bank Economic Review* **21**(2), 301–316.
- Emran, M. S. & Shilpi, F. (2015), ‘Gender, geography, and generations: Intergenerational educational mobility in post-reform india’, *World Development* **72**, 362–380.
- Eren, O. et al. (2018), Evolution of the university wage premium in turkey: 2004-2015, Technical report.
- Erten, B. & Keskin, P. (2018), ‘For better or for worse?: Education and the prevalence of domestic violence in turkey’, *American Economic Journal: Applied Economics* **10**(1), 64–105.
- Fang, H., Eggleston, K. N., Rizzo, J. A., Rozelle, S. & Zeckhauser, R. J. (2012), The returns to education in china: Evidence from the 1986 compulsory education law, Technical report, National Bureau of Economic Research.
- Fortin, N. M. (2006), ‘Higher-education policies and the college wage premium: Cross-state evidence from the 1990s’, *American Economic Review* **96**(4), 959–987.
- Frenette, M. (2004), ‘Access to college and university: Does distance to school matter?’, *Canadian Public Policy/Analyse de Politiques* pp. 427–443.
- Glick, P. & Sahn, D. E. (2000), ‘Schooling of girls and boys in a west african country: the effects of parental education, income, and household structure’, *Economics of education review* **19**(1), 63–87.
- Goldin, C. D. & Katz, L. F. (2009), *The race between education and technology*, harvard university press.
- Goldin, C. & Katz, L. F. (2007), Long-run changes in the us wage structure: Narrowing, widening, polarizing, Technical report, National Bureau of Economic Research.

- Goodman-Bacon, A. (2021), ‘Difference-in-differences with variation in treatment timing’, *Journal of Econometrics* **225**(2), 254–277.
- Grimshaw, D. (2000), ‘Public sector employment, wage inequality and the gender pay ratio in the uk’, *International Review of Applied Economics* **14**(4), 427–448.
- Gulesci, S., Meyersson, E. & Trommlerová, S. (2020), ‘The effect of compulsory schooling expansion on mothers’ attitudes toward domestic violence in turkey’, *The World Bank Economic Review* **34**(2), 464–484.
- Güneş, P. M. (2015), ‘The role of maternal education in child health: Evidence from a compulsory schooling law’, *Economics of Education Review* **47**, 1–16.
- Harmon, C. & Walker, I. (1995), ‘Estimates of the economic return to schooling for the united kingdom’, *The American Economic Review* **85**(5), 1278–1286.
- Hertz, T., Jayasundera, T., Piraino, P., Selcuk, S., Smith, N. & Verashchagina, A. (2008), ‘The inheritance of educational inequality: International comparisons and fifty-year trends’, *The BE Journal of Economic Analysis & Policy* **7**(2).
- Hungerford, T. & Solon, G. (1987), ‘Sheepskin effects in the returns to education’, *The review of economics and statistics* pp. 175–177.
- Işık, E., Orhangazi, Ö. & Tekgüç, H. (2020), ‘Heterogeneous effects of minimum wage on labor market outcomes: A case study from turkey’, *IZA Journal of Labor Policy* **10**(1).
- Jaeger, D. A. & Page, M. E. (1996), ‘Degrees matter: New evidence on sheepskin effects in the returns to education’, *The review of economics and statistics* pp. 733–740.
- Jensen, R. (2010), ‘The (perceived) returns to education and the demand for schooling’, *The Quarterly Journal of Economics* **125**(2), 515–548.
- Kara, M. & Okten, C. (2019), Test anxiety and gender gap: Evidence from a high stakes university entrance exam in turkey. Working paper.
- Katz, L. F. & Murphy, K. M. (1992), ‘Changes in relative wages, 1963–1987: supply and demand factors’, *The quarterly journal of economics* **107**(1), 35–78.
- Kaufmann, K. M. (2014), ‘Understanding the income gradient in college attendance in mexico: The role of heterogeneity in expected returns’, *Quantitative Economics* **5**(3), 583–630.

- Kawaguchi, D. & Mori, Y. (2016), 'Why has wage inequality evolved so differently between japan and the us? the role of the supply of college-educated workers', *Economics of Education Review* **52**, 29–50.
- Kingdon, G. G. (2002), 'The gender gap in educational attainment in india: How much can be explained?', *Journal of Development Studies* **39**(2), 25–53.
- Kingdon, G. G. (2005), 'Where has all the bias gone? detecting gender bias in the intrahousehold allocation of educational expenditure', *Economic Development and Cultural Change* **53**(2), 409–451.
- Koeniger, W., Leonardi, M. & Nunziata, L. (2007), 'Labor market institutions and wage inequality', *ILR review* **60**(3), 340–356.
- KONDA (2020), Barometer survey - violence against women. KONDA, Istanbul.
- Krishna, K., Lychagin, S. & Frisancho, V. (2018), 'Retaking in high stakes exams: Is less more?', *International Economic Review* **59**(2), 449–477.
- Kyui, N. (2016), 'Expansion of higher education, employment and wages: Evidence from the russian transition', *Labour Economics* **39**, 68–87.
- Lee, D. S. (1999), 'Wage inequality in the united states during the 1980s: Rising dispersion or falling minimum wage?', *The Quarterly Journal of Economics* **114**(3), 977–1023.
- Li, S., Whalley, J. & Xing, C. (2014), 'China's higher education expansion and unemployment of college graduates', *China Economic Review* **30**, 567–582.
- Lochner, L. & Moretti, E. (2004), 'The effect of education on crime: Evidence from prison inmates, arrests, and self-reports', *American economic review* **94**(1), 155–189.
- Lucas, R. E. (1989), 'On the mechanics of economic development', *Journal of Monetary Economics* .
- Mankiw, N. G., Romer, D. & Weil, D. N. (1992), 'A contribution to the empirics of economic growth', *The quarterly journal of economics* **107**(2), 407–437.
- Meghir, C. & Palme, M. (2005), 'Educational reform, ability, and family background', *American Economic Review* **95**(1), 414–424.

- Mincer, J. (1974), ‘Schooling, experience, and earnings. human behavior & social institutions no. 2.’.
- MONE (2007), ‘Formal and informal education statistics’, *Ministry of National Education* .
- Moretti, E. (2013), ‘Real wage inequality’, *American Economic Journal: Applied Economics* **5**(1), 65–103.
- Müller, K.-U. & Wrohlich, K. (2020), ‘Does subsidized care for toddlers increase maternal labor supply? evidence from a large-scale expansion of early childcare’, *Labour Economics* **62**, 101776.
- Nguyen, T. (2008), ‘Information, role models and perceived returns to education: Experimental evidence from madagascar’, *Unpublished manuscript* .
- OECD (2017), *Education at a glance 2017: OECD indicators*, OECD.
- OECD (2020a), Population with tertiary education. accessed: July 2020.
- OECD (2020b), ‘Population with tertiary education’.
URL: <https://data.oecd.org/eduatt/population-with-tertiary-education.htm>
- Oppedisano, V. (2011), ‘The (adverse) effects of expanding higher education: Evidence from italy’, *Economics of Education Review* **30**(5), 997–1008.
- Orazem, P. F. & King, E. M. (2007), ‘Schooling in developing countries: The roles of supply, demand and government policy’, *Handbook of development economics* **4**, 3475–3559.
- Oreopoulos, P. (2006), ‘Estimating average and local average treatment effects of education when compulsory schooling laws really matter’, *American Economic Review* **96**(1), 152–175.
- Oreopoulos, P. (2007), ‘Do dropouts drop out too soon? wealth, health and happiness from compulsory schooling’, *Journal of public Economics* **91**(11-12), 2213–2229.
- Oreopoulos, P., Page, M. E. & Stevens, A. H. (2006), ‘The intergenerational effects of compulsory schooling’, *Journal of Labor Economics* **24**(4), 729–760.

- Pal, S. (2004), ‘How much of the gender difference in child school enrolment can be explained? evidence from rural india’, *Bulletin of Economic Research* **56**(2), 133–158.
- Pischke, J.-S. & Von Wachter, T. (2008), ‘Zero returns to compulsory schooling in germany: Evidence and interpretation’, *The Review of Economics and Statistics* **90**(3), 592–598.
- Polat, S. (2017), ‘The expansion of higher education in turkey: Access, equality and regional returns to education’, *Structural Change and Economic Dynamics* **43**, 1–14.
- Rankin, B. H. & Aytaç, I. A. (2006), ‘Gender inequality in schooling: The case of turkey’, *Sociology of Education* **79**(1), 25–43.
- Reimer, D. & Pollak, R. (2010), ‘Educational expansion and its consequences for vertical and horizontal inequalities in access to higher education in west germany’, *European Sociological Review* **26**(4), 415–430.
- Roodman, D. (2020), ‘Boottest: Stata module to provide fast execution of the wild bootstrap with null imposed’.
- Rosenzweig, M. R. & Wolpin, K. I. (1986), ‘Evaluating the effects of optimally distributed public programs: Child health and family planning interventions’, *The American Economic Review* **76**(3), 470–482.
- Rosenzweig, M. R. & Wolpin, K. I. (1994), ‘Are there increasing returns to the intergenerational production of human capital? maternal schooling and child intellectual achievement’, *Journal of Human Resources* pp. 670–693.
- Roth, J., Sant’Anna, P. H., Bilinski, A. & Poe, J. (2022), ‘What’s trending in difference-in-differences? a synthesis of the recent econometrics literature’, *arXiv preprint arXiv:2201.01194* .
- Sacerdote, B. (2002), ‘The nature and nurture of economic outcomes’, *American Economic Review* **92**(2), 344–348.
- Sacerdote, B. (2004), ‘What happens when we randomly assign children to families?’, *National Bureau of Economic Research* .
- Saygin, P. O. (2016), ‘Gender differences in preferences for taking risk in college applications’, *Economics of Education Review* **52**, 120–133.

- Saygin, P. O. (2020), ‘Gender bias in standardized tests: evidence from a centralized college admissions system’, *Empirical Economics* **59**(2), 1037–1065.
- Schofer, E. & Meyer, J. W. (2005), ‘The worldwide expansion of higher education in the twentieth century’, *American sociological review* **70**(6), 898–920.
- Schultz, T. P. (2002), ‘Why governments should invest more to educate girls’, *World development* **30**(2), 207–225.
- Spence, M. (1978), Job market signaling, in ‘Uncertainty in economics’, Elsevier, pp. 281–306.
- Spiess, C. K. & Wrohlich, K. (2010), ‘Does distance determine who attends a university in germany?’, *Economics of Education Review* **29**(3), 470–479.
- Spohr, C. A. (2003), ‘Formal schooling and workforce participation in a rapidly developing economy: Evidence from “compulsory” junior high school in taiwan’, *Journal of development Economics* **70**(2), 291–327.
- Stephens Jr, M. & Yang, D.-Y. (2014), ‘Compulsory education and the benefits of schooling’, *American Economic Review* **104**(6), 1777–92.
- Sun, L. & Abraham, S. (2021), ‘Estimating dynamic treatment effects in event studies with heterogeneous treatment effects’, *Journal of Econometrics* **225**(2), 175–199.
- Taber, C. R. (2001), ‘The rising college premium in the eighties: Return to college or return to unobserved ability?’, *The Review of Economic Studies* **68**(3), 665–691.
- Tansel, A. (2015), ‘Intergenerational educational mobility in turkey’, *Unpublished Manuscript*.
- Trostel, P., Walker, I. & Woolley, P. (2002), ‘Estimates of the economic return to schooling for 28 countries’, *Labour economics* **9**(1), 1–16.
- Tsai, W.-J., Liu, J.-T., Chou, S.-Y. & Thornton, R. (2009), ‘Does educational expansion encourage female workforce participation? a study of the 1968 reform in taiwan’, *Economics of Education Review* **28**(6), 750–758.
- Turkstat (2012), ‘Adult education survey’, *Turkish Statistical Institute*.

- Turner, S. E. & Bowen, W. G. (1999), ‘Choice of major: The changing (unchanging) gender gap’, *ILR Review* **52**(2), 289–313.
- Urbina, D. R. (2018), ‘Intergenerational educational mobility during expansion reform: evidence from mexico’, *Population research and policy review* **37**(3), 367–417.
- Valero, A. & Van Reenen, J. (2019), ‘The economic impact of universities: Evidence from across the globe’, *Economics of Education Review* **68**, 53–67.
- Walker, I. & Zhu, Y. (2008), ‘The college wage premium and the expansion of higher education in the uk’, *Scandinavian Journal of Economics* **110**(4), 695–709.
- World Bank, W. (2018), ‘World development indicators- female labor force participation rate’.
URL: <https://datacatalog.worldbank.org/dataset/world-development-indicators>
- Xing, C., Yang, P. & Li, Z. (2018), ‘The medium-run effect of china’s higher education expansion on the unemployment of college graduates’, *China Economic Review* **51**, 181–193.
- Yilmaz, B. (2014), The effects of the university reform in Turkey on narrowing gender gap in higher education, PhD thesis, Bilkent University.

APPENDICES

A. ADDITIONAL RESULTS FOR CHAPTER 2

A.1. Returns to education in Turkey

Here, we estimate the gender differential employment and wage returns to higher education in Turkey. Our sample includes individuals in ages 25 or older, observed in 2004-2013 HLFS (the earlier waves of HLFS are not compatible with later waves and therefore excluded from this analysis), to capture the returns to education when the individuals in our main analyses make their higher education decisions. We further restricted the sample to at least high school graduates to estimate the returns to a higher education degree compared to the highest level of education before the higher education in the labor market. First, we regress a binary employment variable on a binary higher education degree variable and its interaction with woman dummy. We additionally control for age and age squared to capture potential experience, and region and survey year effects. We present the estimates for ages 25-34, 25-44, and 25-64 in the first three columns of Table A.1. Similarly, we regress the log of real hourly wages on educational attainment and its interaction with woman dummy for the sample of full-time wage employees with the same control variables, in addition to firm tenure. We report our coefficient estimates, which capture wage returns for the same age groups, in the last three columns. Similar to what many studies found for the US (many of them cited in Dougherty (2005)) and documented for many European countries (Trostel et al. 2002); labor market returns to education are generally higher for woman in Turkey for the 10-year period that we focus in our regression analysis. Our coefficient estimates for the linear combination of woman and higher education dummies given in the last row of the table show the relative labor market advantage of college

graduate women compared to college graduate men indicating significant employment disadvantage of college graduate women. When we focus on wage returns to education, we do not find statistically different wage levels for men than women with college education.

Table A.1: Returns to higher education by gender

<i>Dependent variable:</i>	<i>Employment</i>			<i>Log(Hourly wage)</i>		
	(1)	(2)	(3)	(4)	(5)	(6)
Woman	-0.584*** (0.012)	-0.593*** (0.011)	-0.554*** (0.011)	0.007 (0.022)	-0.042* (0.023)	-0.060*** (0.021)
Higher education	-0.010* (0.005)	0.014*** (0.004)	0.043*** (0.003)	0.617*** (0.027)	0.607*** (0.019)	0.583*** (0.014)
Woman×Higher education	0.391*** (0.012)	0.398*** (0.013)	0.328*** (0.011)	0.013 (0.024)	0.054** (0.026)	0.075*** (0.024)
Woman + Woman×Higher ed.	-0.193*** (0.007)	-0.194*** (0.006)	-0.226*** (0.006)	0.019 (0.014)	0.013 (0.012)	0.015 (0.012)
Observations	264,760	439,467	615,541	142,762	240,052	296,012
Age window	25-34	25-44	25-64	25-34	25-44	25-64

Notes: The sample for each regression analysis includes the age window given below each column observed in HLFS 2004-2013. Higher education is a binary variable equal to 1 if the individual has graduated from a higher education program. Employment is a binary variable that is 1 if an individual self-reports to be employed. Hourly wages are similarly self-reported real wages. A constant, age and age squared variables, and region, and survey year fixed effects are included as control variables. Columns (4)-(6) additionally controls for firm tenure. The standard errors clustered at region level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

A.2. Expansion policy on higher education enrollment: Public vs. private universities

Here, we investigate how the additional supply in public higher education institutions affects the enrollment in public and private universities and look for evidence of the substitution effects of the expansion policy. We do not observe in the micro data whether students are enrolled in public or private institutions; we only observe that they are enrolled in tertiary education. Typically, the public universities are already established in a region before private universities enter the market. Indeed, before the expansion, only four regions (Istanbul, Ankara, Izmir, and Adana-Mersin region) had at least one private university and all five provinces in those regions had at least one public university before the expansion (Table A.6 in the Appendix provides more information). Moreover, the share of private universities in total available slots was low both before and after the expansion (11% in 2005, 13% in 2009, and 17% in 2013). Also, there are strict rules that regulate the transfers of students between universities.

By using the data on new enrollment numbers of all four-year higher education

programs (available at the NUTS-3 level), we investigate how the enrollment into private higher education institutions evolves with the expansion. We exploit the exogenous variation in our measure of intensity across regions and years with the following equation:

$$\ln(1 + enrollment_{rt}) = \beta s_{rt} + \alpha X_{rt} + \gamma_r + \mu_t + \epsilon_{rt} \quad (5.1)$$

where $enrollment_{rt}$ is the number of students enrolled in higher education institutions in a geographical region r and year t , and s_{rt} is the intensity in region r and year t . We control for time-varying regional controls X_{rt} that includes the population of 18-year-olds, region effects γ_r and year effects μ_t . Data on enrollment starts from 2000 since the earlier years are not published on the Measuring, Selection, and Placement Center's website. Hence, our largest sample includes 6 years before the policy cutoff in 2006 and 7 years after the cutoff.

Table A.2: The expansion policy on public and private enrollment at the NUTS-3 level

Dependent variable:	$\ln(1+Public\ enrollment)$			$\ln(1+Private\ enrollment)$		
	(1)	(2)	(3)	(4)	(5)	(6)
Intensity	5.610*** (1.699)	7.238*** (1.874)	9.394*** (1.888)	-1.036* (0.539)	-1.035* (0.535)	-0.868 (0.548)
Observations	1,053	972	810	1,053	972	810
Sample	-6/+7 years	± 6 years	± 5 years	-6/+7 years	± 6 years	± 5 years

Notes: Intensity in a given year and region is defined as the increase in the available slots in public universities in that year relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region, where regions are given at the NUTS-3 level. A constant, number of 18-year-olds, region (NUTS-3), and year fixed effects are included as control variables. The standard errors clustered at region (NUTS-3) level are given in parentheses. * $p < 0.1$ ** $p < 0.05$ *** $p < 0.01$.

Table A.3: Higher education statistics across cohorts

Cohort	Higher education graduation	Higher education attainment
2000	0.218	0.233
2001	0.229	0.244
2002	0.243	0.261
2003	0.244	0.266
2004	0.275	0.300
2005	0.278	0.302
2006	0.297	0.324
2007	0.307	0.343
2008	0.327	0.371
2009	0.327	0.379
2010	0.326	0.402
2011	0.291	0.410
2012	0.249	0.418
2013	0.176	0.406

Source: HLFS, 2016-2017.

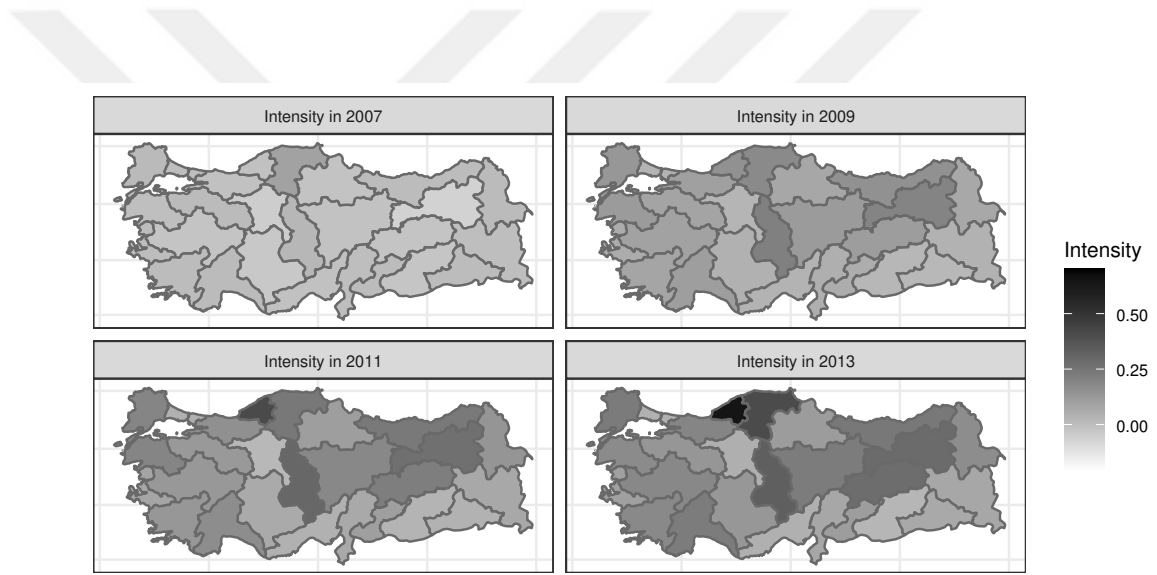
Notes: Authors' calculations. For cohorts given in each row, higher education attainment gives the share of individuals who graduated from or enrolled in a higher education program. Higher education graduation gives the share of individuals graduated from a higher education program.

Table A.4: Programs by field of study

Major	Programs
Social Sciences	Economics, business and administration, anthropology, philosophy, sociology, psychology.
Engineering	All engineering programs, architecture, construction and building, computing, manufacturing and processing.
Science	Biology, molecular biology and genetics, physics, chemistry, mathematics, statistics.
Teaching	All education programs except for teaching in higher education.
Health	Medicine, nursing, dentistry, pharmacy, veterinary.
Law	Law
Services	Social services, personal services, transportation services, security services.
Other	Arts, journalism, agriculture, forestry and fishery.

Notes: Authors' definitions.

Figure A.1: The spatial distribution of intensity across years.



Source: OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: Authors' calculations. Authors' calculations. Intensity in a given year and region is defined as the increase in the available slots in public universities in that year relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region.

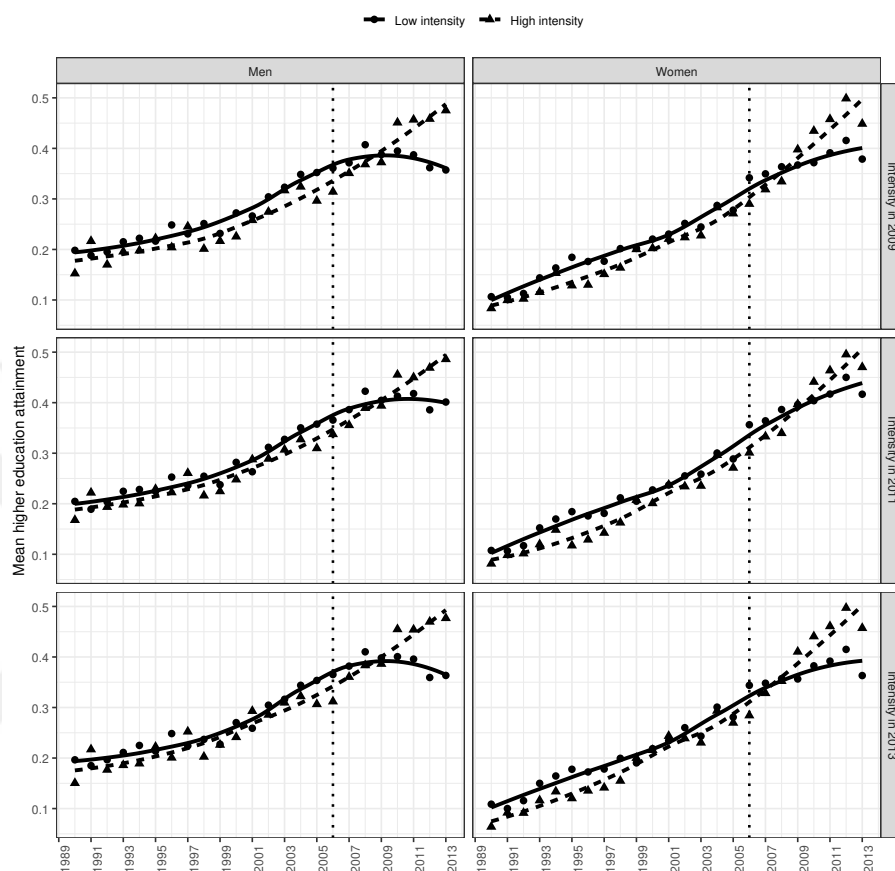
Table A.5: Descriptive statistics

	Men						Women					
	Pre-expansion			Post-expansion			Pre-expansion			Post-expansion		
	Attained	Not attained		Attained	Not attained		Attained	Not attained		Attained	Not attained	
Age [years]	32.741	32.375		24.524	24.445		32.696	32.370		24.628	24.355	
Resides in West regions [0-1]	0.371	0.407		0.332	0.385		0.349	0.438		0.305	0.397	
Resides in Central regions [0-1]	0.180	0.204		0.172	0.181		0.177	0.203		0.166	0.200	
Resides in South regions [0-1]	0.108	0.096		0.104	0.104		0.117	0.113		0.102	0.104	
Resides in North regions [0-1]	0.113	0.106		0.101	0.117		0.114	0.103		0.101	0.108	
Resides in East regions [0-1]	0.227	0.187		0.290	0.212		0.243	0.143		0.325	0.190	
Intensity	0	0		0.093	0.102		0	0		0.090	0.102	
Old intensity	0	0		0.049	0.054		0	0		0.046	0.056	
New intensity	0	0		0.045	0.048		0	0		0.044	0.046	
Pop. of 18-year-olds [in 1,000,000]	0	0		0.061	0.067		0	0		0.060	0.069	
Observations	13,212	34,266		16,827	26,452		12,214	39,400		17,189	26,762	

Source: HILFS, OSYM Almanacs and Turkstat Education Statistics.

Notes: Pre-expansion: Averages calculated for 1999-2005 cohorts. Post-expansion: Averages calculated for 2007-2013 cohorts. Not attained and attained: Averages for those whose higher education attainment is 0 and 1, respectively, where higher education indicates whether the individual has graduated from or is enrolled in a higher education program. The regions of residence presented here are West (NUTS-1 regions 1 to 4), Central (NUTS-1 regions 5 and 7), South (NUTS-1 region 6), North (NUTS-1 regions 8 and 9) and East (NUTS-1 regions 10 to 12). This is also the classification used in the Demographic and Health Surveys of Turkey.

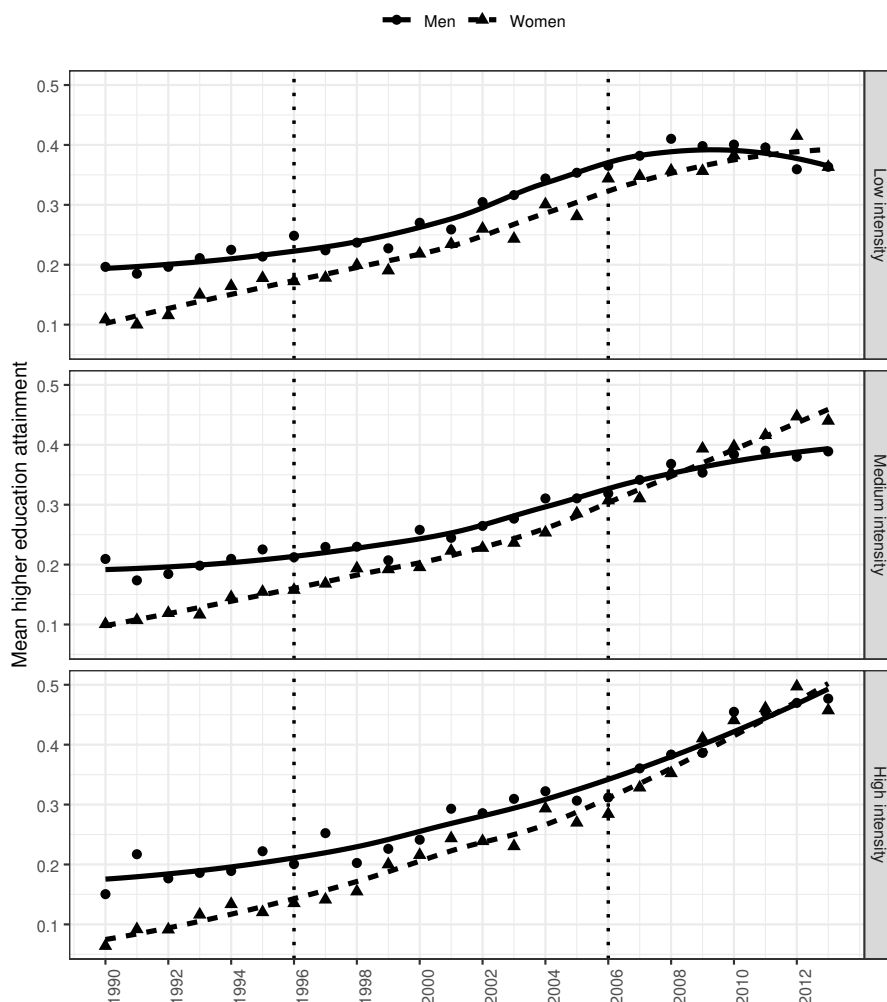
Figure A.2: The higher education attainment across cohorts: High, Low and medium intensity regions in years 2009, 2011 and 2013.



Source: HLFS, 2016-2017, OSYM Almanacs and TurkStat Address Based Population Registry System.

Notes: Authors' calculations. Higher education attainment is defined as a binary variable equal to 1 if the individual is a graduate or is currently enrolled in a higher education program. High and low intensity regions are the top and bottom quartiles of intensity in 2009, 2011, or 2013 (increase in available slots per 18-year-old population relative to slots in 2005). The regions in the two quartiles in the middle are medium intensity regions. Attainment rates are the cohort-level averages for men and women.

Figure A.3: The gender gap in higher education attainment across cohorts: High, Low and medium intensity regions.



Source: HLFS, 2016-2017, OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: Authors' calculations. Higher education attainment is defined as a binary variable equal to 1 if the individual is a graduate or is currently enrolled in a higher education program. High and low intensity regions are the top and bottom quartiles of intensity in 2013 (increase in available slots per 18-year-old population in 2013 relative to slots in 2005). The regions in the two quartiles in the middle are medium intensity regions. Attainment rates are the cohort-level averages for men and women.

Table A.6: The exogeneity of expansion policy intensity

<i>Dependent variable:</i>							
	Intensity in 2007	Intensity in 2008	Intensity in 2009	Intensity in 2010	Intensity in 2011	Intensity in 2012	Intensity in 2013
University exam takers per 18-year-old in 2005	-0.0030 (0.0051)	0.0098 (0.0066)	0.0177* (0.0100)	0.0258 (0.0159)	0.0318* (0.0177)	0.0397* (0.0217)	0.0410 (0.0253)
Higher education attainment rate of ages 25-34 in 2005	0.0682 (0.1894)	0.4168* (0.2373)	0.3369 (0.3838)	0.1664 (0.6141)	0.1306 (0.6888)	0.3058 (0.8480)	0.2337 (0.9744)
Higher education attainment rate of ages 25-44 in 2005	0.0427 (0.2165)	0.4988* (0.2689)	0.4594 (0.4347)	0.2931 (0.6990)	0.2594 (0.7846)	0.5062 (0.9644)	0.4395 (1.1092)
Higher education attainment rate of ages 25-64 in 2005	0.0032 (0.2209)	0.3221 (0.2857)	0.1622 (0.4522)	-0.1473 (0.7147)	-0.2374 (0.8003)	-0.0108 (0.9889)	-0.1262 (1.1344)
Gross domestic product (GDP) per capita in 2005	-0.4276 (1.5469)	0.2159 (2.0555)	-1.8833 (3.1568)	-5.7647 (4.8765)	-5.5246 (5.5088)	-5.2828 (6.8512)	-6.6754 (7.8400)
Number of public universities per 18-year-old in 2005	-0.3966* (0.2233)	0.2958 (0.3094)	0.6372 (0.4699)	0.8683 (0.7484)	1.2200 (0.8253)	1.1064 (1.0391)	1.1043 (1.1990)
Number of universities per 18-year-old in 2005	-0.3013* (0.1469)	-0.0927 (0.2105)	-0.1200 (0.3260)	-0.3508 (0.5107)	-0.2223 (0.5763)	-0.2526 (0.7111)	-0.4424 (0.8130)
Available public slots per 18-year-old in 2005	-0.1988*** (0.0461)	0.0234 (0.0813)	0.1359 (0.1229)	0.1273 (0.1971)	0.1468 (0.2208)	0.1250 (0.2736)	0.1082 (0.3145)
Available slots per 18-year-old in 2005	-0.1799*** (0.0420)	-0.0056 (0.0740)	0.0686 (0.1136)	0.0248 (0.1805)	0.0377 (0.2023)	0.0245 (0.2496)	-0.0104 (0.2865)
Industry labor productivity in 2005	-2.2923* (1.2339)	-0.2820 (1.7501)	-0.0439 (2.7085)	-2.6708 (4.2376)	-3.6261 (4.7316)	-4.3996 (5.8383)	-5.7998 (6.6731)
Total labor productivity in 2005	0.1337 (0.7541)	0.4796 (0.9966)	-0.6084 (1.5439)	-2.3417 (2.3960)	-2.7880 (2.6789)	-2.8176 (3.3286)	-3.5088 (3.8089)

Source: OSYM Almanacs and Turkstat Education Statistics.

Notes: Intensity in a given year and region is defined as the increase in the available slots in public universities in that year relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. Each cell presents the coefficient estimate from a regression analysis where the dependent and independent variables are given in the associated column and row, respectively. The number of observations is 26 (the number of NUTS-2 regions) for each regression analysis. Each specification controls for a constant in addition to the independent variable given in the row. Higher education rate in a region is calculated as the share of 25-34-year-olds (to rely on the increasing trend in educational attainment) with higher education attainment among 25-34-year-olds. Real production and total employment in the industry sector (non-agriculture and non-service sectors) are used for the industry labor productivity calculations. Real total production and total employment are used for total labor productivity. Robust standard errors are given in the parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table A.7: The orthogonality of migration probability to expansion policy intensity

	<i>Dependent variable: Migrated since age 16</i>					
	Men			Women		
	(1)	(2)	(3)	(4)	(5)	(6)
Intensity	0.148 (0.107)	0.145 (0.093)	0.104 (0.101)	0.070 (0.114)	0.102 (0.115)	0.134 (0.115)
Observations	90,757	77,050	63,469	95,565	81,266	67,393
Cohort window	±7 cohorts	±6 cohorts	±5 cohorts	±7 cohorts	±6 cohorts	±5 cohorts

Notes: Migrated since age 16 is a binary variable that is 1 if the individual has migrated since age 16. Intensity in a given year and region is defined as the increase in the available slots in public universities in that year relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. A constant, number of 18-year-olds, region, cohort and survey year fixed effects are included as control variables. The standard errors clustered at region level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

Table A.8: The effect of compulsory schooling law on higher education attainment

	<i>Dependent variable: Higher education attainment</i>			
	Men		Women	
	(1)	(2)	(3)	(4)
Compulsory schooling policy	-0.024** (0.011)	-0.020 (0.012)	-0.003 (0.008)	-0.010 (0.013)
Observations	51,148	38,149	55,131	41,344
Cohorts	2001-2008	2002-2007	2001-2008	2002-2007

Notes: Compulsory schooling policy is a binary variable that is 1 if the individual is exposed to the 1997 compulsory schooling law, i.e., born in 1987 or later. Higher education attainment is defined as a binary variable equal to 1 if the individual is a graduate or is currently enrolled in a higher education program. A constant, time trends before and after the compulsory schooling policy, region and survey year fixed effects are included as control variables. The standard errors clustered at region×cohort level are given in parentheses. *p<0.1 **p<0.05 ***p<0.01.

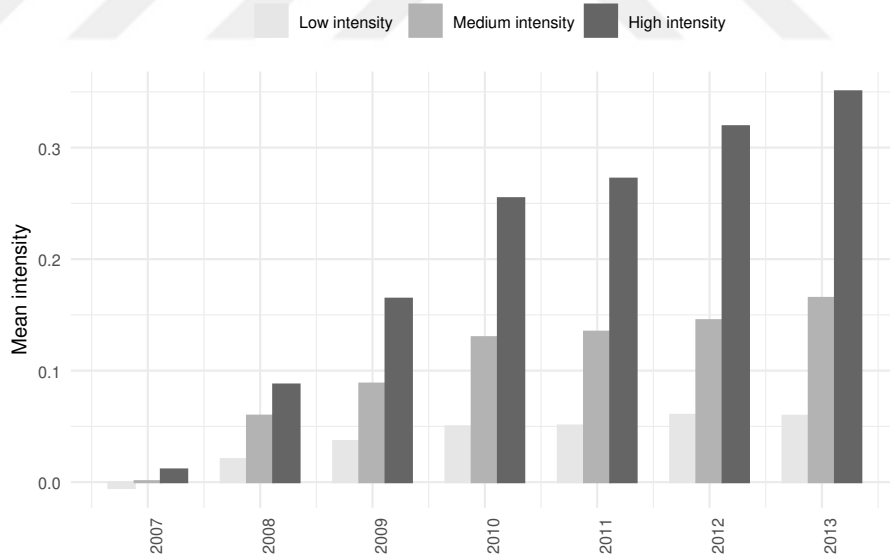
Table A.9: Orthogonality of the expansion policy to regional conservativeness

<i>Dependent variable:</i>							
	Intensity in 2007	Intensity in 2008	Intensity in 2009	Intensity in 2010	Intensity in 2011	Intensity in 2012	Intensity in 2013
The share of women wearing headscarf	-0.0310 (0.0367)	-0.0216 (0.0492)	0.0691 (0.0751)	0.1624 (0.1159)	0.1350 (0.1322)	0.1141 (0.1650)	0.1155 (0.1897)
The share of conservative individuals	-0.0368 (0.0463)	-0.0130 (0.0622)	0.1087 (0.0938)	0.2271 (0.1448)	0.1857 (0.1661)	0.1470 (0.2080)	0.1546 (0.2390)
The share of religious individuals	-0.0486 (0.0875)	-0.0211 (0.1168)	0.0523 (0.1804)	0.1495 (0.2835)	0.1587 (0.3179)	0.1459 (0.3931)	0.1401 (0.4514)

Source: KONDA Barometer.

Notes: Intensity in a given year and region is defined as the increase in the available slots in public universities in that year relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. Each cell presents the coefficient estimate from a regression analysis where the dependent and independent variables are given in the associated column and row, respectively. The number of observations is 26 (the number of NUTS-2 regions) for each regression analysis. Each specification controls for a constant in addition to the independent variable given in the row. The independent variables are constructed from KONDA Barometer survey examining the perspectives on gender attitudes, conducted in 2018 and representative at the NUTS-1 level. Robust standard errors are given in the parentheses. *** p<0.01, ** p<0.05, * p<0.1

Figure A.4: Variations in the mean of intensities across years: High, medium and low intensity regions.



Source: OSYM Almanacs and TurkStat Adress Based Population Registry System.

Notes: Mean intensity in years 2007-2013. Intensity in a given year, in a given region, is defined as the increase in the available slots in public universities in that year relative to the number of slots in 2005 (before expansion), divided by the 18-year-old population in that region. High intensity and low intensity regions are defined as the ones whose intensity in 2013 is in the top and bottom quartiles, respectively. The remaining regions whose intensity is in the mid quartiles are determined to be medium intensity regions.

B. ADDITIONAL RESULTS FOR CHAPTER 4

B.1. 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 B.1 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.

Figure B.1: Geographical classification: West vs. East

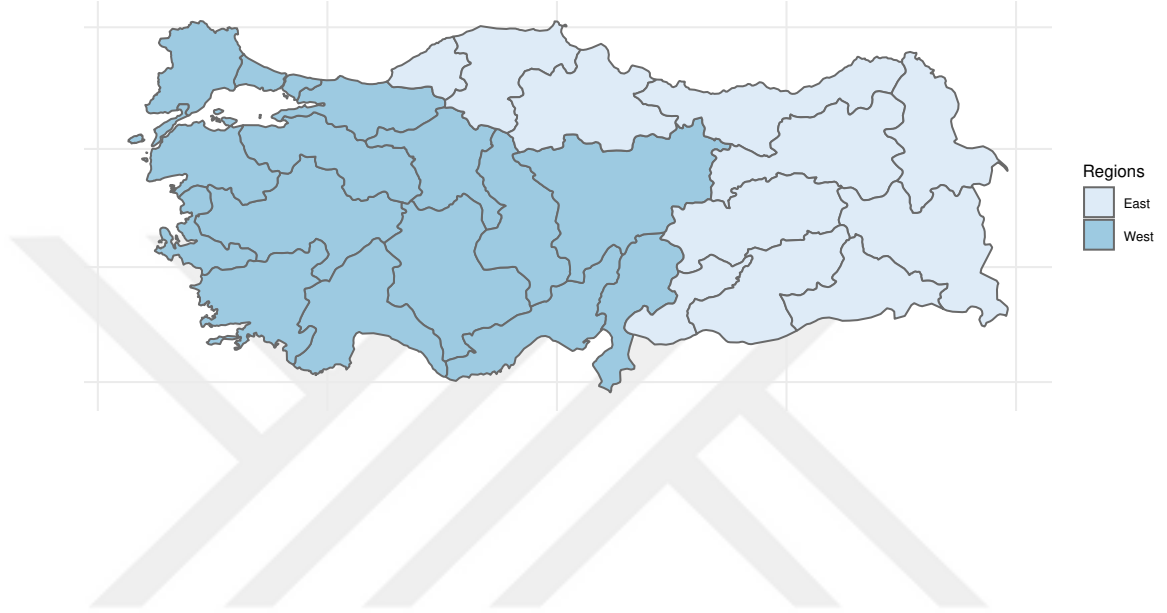


Table B.1: The gender view index at the NUTS-1 level

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

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

Table B.2: Robustness checks: Alternative bandwidths and time trend controls

<i>Dependent var:</i>	<i>Completion of 8-year primary education</i>			<i>Completion of secondary education</i>		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Controls for linear time trend						
Woman	-0.162*** (0.023)	-0.173*** (0.016)	-0.194*** (0.020)	-0.099*** (0.015)	-0.093*** (0.010)	-0.106*** (0.016)
Policy	0.248*** (0.036)	0.244*** (0.028)	0.222*** (0.029)	0.101*** (0.030)	0.162*** (0.044)	0.119*** (0.038)
Woman×Policy	-0.042 (0.032)	-0.032 (0.024)	-0.002 (0.025)	-0.044** (0.018)	-0.099*** (0.029)	-0.064** (0.028)
Parental education	0.322*** (0.009)	0.331*** (0.009)	0.349*** (0.016)	0.450*** (0.014)	0.450*** (0.011)	0.457*** (0.010)
Parental education×Woman	0.174*** (0.018)	0.182*** (0.016)	0.177*** (0.015)	0.099*** (0.023)	0.115*** (0.021)	0.120*** (0.018)
Parental education×Policy	-0.246*** (0.016)	-0.267*** (0.015)	-0.289*** (0.019)	-0.084*** (0.020)	-0.095*** (0.025)	-0.109*** (0.024)
Parental educ×Woman×Policy	-0.054** (0.022)	-0.062*** (0.019)	-0.060*** (0.017)	0.009 (0.024)	-0.009 (0.027)	-0.019 (0.026)
Observations	12,678	17,012	19,948	12,678	17,012	19,948
Panel A: Controls for second-order polynomial of time trend						
Woman	-0.182*** (0.031)	-0.156*** (0.037)	-0.134*** (0.035)	-0.073*** (0.015)	-0.080*** (0.016)	-0.062*** (0.019)
Policy	0.155*** (0.031)	0.236*** (0.044)	0.274*** (0.046)	0.068* (0.037)	0.044* (0.024)	0.166** (0.058)
Woman×Policy	0.044 (0.032)	-0.027 (0.043)	-0.074 (0.047)	-0.088** (0.038)	-0.015 (0.028)	-0.104** (0.044)
Parental education	0.322*** (0.009)	0.331*** (0.009)	0.350*** (0.016)	0.450*** (0.014)	0.450*** (0.011)	0.458*** (0.010)
Parental education×Woman	0.173*** (0.018)	0.182*** (0.016)	0.177*** (0.015)	0.100*** (0.023)	0.115*** (0.021)	0.120*** (0.017)
Parental education×Policy	-0.246*** (0.016)	-0.267*** (0.016)	-0.290*** (0.020)	-0.084*** (0.020)	-0.095*** (0.025)	-0.109*** (0.025)
Parental educ×Woman×Policy	-0.055** (0.023)	-0.062*** (0.018)	-0.060*** (0.017)	0.009 (0.024)	-0.009 (0.027)	-0.019 (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

Notes: 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 B.3: Robustness check with cohort 1986 is included in the sample

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

Notes: 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 B.4: Robustness check with sample restricted to 2012 and 2016 waves

<i>Dependent var:</i>	<i>Completion of 8-year primary education</i>			<i>Completion of secondary education</i>		
	AES 2012-16 (1)	AES 2012 (2)	AES 2016 (3)	AES 2012-16 (4)	AES 2012 (5)	AES 2016 (6)
Woman	-0.162*** (0.023)	-0.151*** (0.017)	-0.138*** (0.040)	-0.099*** (0.015)	-0.075*** (0.016)	-0.092* (0.050)
Policy	0.248*** (0.036)	0.276*** (0.036)	0.201*** (0.051)	0.101*** (0.030)	0.080** (0.027)	0.127** (0.044)
Woman×Policy	-0.042 (0.032)	-0.051* (0.025)	-0.032 (0.051)	-0.044** (0.018)	-0.043 (0.028)	-0.034 (0.062)
Parental education	0.322*** (0.009)	0.329*** (0.019)	0.311*** (0.017)	0.450*** (0.014)	0.464*** (0.018)	0.429*** (0.012)
Parental education×Woman	0.174*** (0.018)	0.184*** (0.029)	0.155*** (0.028)	0.099*** (0.023)	0.096*** (0.028)	0.109*** (0.022)
Parental education×Policy	-0.246*** (0.016)	-0.268*** (0.024)	-0.216*** (0.024)	-0.084*** (0.020)	-0.120*** (0.021)	-0.039 (0.028)
Parental educ×Policy×Woman	-0.054** (0.022)	-0.053 (0.030)	-0.046 (0.038)	0.009 (0.024)	0.050 (0.034)	-0.047 (0.032)
Observations	12,678	7,756	4,922	12,678	7,756	4,922
Birth cohorts	1980-92	1980-92	1980-92	1980-92	1980-92	1980-92

Notes: 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 $p < 0.01$.