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The revolution is dead, long live the demolition: Education and labor market consequences of student riots

Abstract

The 1970s witnessed violent, widespread, and highly-politicized student protests in Turkey. Small protests turned into bloody street clashes, the death toll exceeded 5,000, and a military coup came in—which resulted in mass arrests. The universities were at the center of violent conflict. We study the education and labor market consequences of this political turmoil on the cohorts exposed to educational disruptions. First, we document that the number of new admissions and graduates in post-secondary education declined significantly due to the turmoil. The decline in post-secondary graduation ratio is 6.6 percentage points for the exposed individuals. Second, we estimate a counterfactual wage distribution for the exposed cohorts and check whether the turmoil affected their wage and occupation distributions. We show that the decline in educational attainment pushed the exposed population toward medium- and low-income occupations, and compressed their wages toward the minimum wage. Finally, we use the unexpected decline in post-secondary education as an instrument to estimate returns to college. We find that the college-premium is around 15 percent per year for men.

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1 Introduction

From the Arab Spring to yellow vests movement in France and from anti-Revolution protests in Iran to Hong Kong's umbrella revolution, it is common to see mass socio-political movements in the early 21st century. Factors such as resurgence of far-right political views, authoritarianism, widespread human rights violations, war/conflict, increased volume of forced displacement, technological change, and the Covid-19 outbreak have contributed to the increased risk of social uprisings that might devastate economies, reduce well-being, and ramp up inequalities. This paper looks back to an earlier major uprising—i.e., the student protests that took place in Turkey between 1978 and 1980 and eventually led to the 1980 military coup—to explore how exposure to violent conflict that led to prolonged disruptions in tertiary education affected people's educational and labor market outcomes later in life.

In the 1960s and early 1970s, student movements concurrently grew in much of the world and became a global phenomenon.[1] In the US and Europe, these movements had declined significantly by mid-1970s (Barker, 2008), which was the time they escalated in Turkey (Ahmad, 1993). Turkey experienced violent, widespread, and politically-motivated student protests in late 1970s followed by a military coup in 1980. During the peak years of the protests, clashes between youth groups affiliated with the political right and left turned extremely violent (Zurcher, 2004)—an average of 20 young individuals were killed each day on Turkey's streets and university campuses. After the 1980 coup, students were regularly snatched up in mass arrests.

In this paper, we document the education and labor market consequences of those dramatic events.[2] We start with a detailed analysis of the impact of the turmoil on post-secondary education. Then, we estimate a counterfactual wage distribution for the exposed cohorts and discuss what would happen to wage and occupation distributions if the turmoil had not taken place. Finally, we use the unexpected decline in tertiary education for the exposed cohorts as an instrument to estimate the causal effect of tertiary education on earnings.

Violent conflict adversely affected post-secondary education through three channels. First, new enrollments declined after the closure of teacher-training institutes because of their involvement in violence. Second, graduation rates fell following the massive student dropouts related to security concerns. Finally, repression in the wake of the 1980 coup kept many from completing their education. Using the Turkish Household Labor Force Surveys (HLFS), we document that male wage earners, with birth years from 1960 to 1965, were the most adversely affected ones. Specifically, the 1978-1982 turmoil led to a 6.6 percentage point decline in the probability of completing postsecondary education for the exposed cohorts. Our counterfactual analysis suggests that the decline in college attainment led to a clear shift from high-income to low-income jobs and occupations.

We use the turmoil as a plausibly exogenous event affecting post-secondary education for the exposed cohorts and implement an IV strategy to estimate the returns to college. We find that the college premium ranges between 56-58 log points for men, which corresponds to around 15 percent returns to per year of post-secondary education. We also provide convincing

1 See, e.g., Samuelson (1968), Kazuko (1968), Flacks (1970), Rothman and Lichter (1978), Koopmans (1993), and Thomas (2002) for background information on student protests in different countries and contexts during this episode.

2 Throughout the paper, we interchangeably refer to those events as the "1978-1982 turmoil," the "political turmoil," or, shortly, the "turmoil."

evidence that the decline in earnings for the exposed cohorts is solely due to the decline in their post-secondary educational attainment, not due to other confounding factors—such as long-term psychological effects, etc.—that could potentially dampen labor productivity.

The plan of the paper is as follows. Section 2 reviews the literature and discusses our main contributions. Section 3 provides detailed background information about the student protests in the 1970s and the subsequent military coup in Turkey. Section 4 describes the data. Section 5 reports detailed evidence on the decline in post-secondary education, presents IV estimates for returns to college along with several robustness checks, and documents results from a counterfactual analysis of wage distributions and occupational shift. Section 6 concludes.

2 **Literature review**

Violent conflict might affect various outcomes of the exposed individuals, and there is a large literature documenting the short- and long-term consequences of conflict exposure.[3] Specifically, violent conflict may lead to disruption in education and negatively affect long-term human capital and labor market outcomes. Studies in this literature have focused on a wide range of events from different countries/contexts and documented large negative impact of conflict exposure on schooling.[4] Almost all the papers in this literature—especially the ones focusing on less developed countries—document education consequences at primary- and/or secondary-school levels.

Although there is vast evidence documenting the impact of conflict exposure on primary and secondary education, evidence on tertiary-level effects is scant. There might be several reasons for this evidence gap. First, children are more vulnerable to conflict than adults. Second, sometimes violence explicitly target children (Valente, 2014). Third, and most importantly, the type of events studied in the literature mostly lead to educational disruptions either more generally across all education categories or exclusively for primary- and/or secondary-school levels. We contribute to this literature by providing direct evidence on violent conflict exposure exclusively targeting tertiary education. In particular, we focus on a politically-motivated and widespread student movement that was centered around Turkish higher education institutions in the 1970s and repressed by the 1980 military coup. We show that the set of events before and after the coup, as a whole, generated significant disruptions in higher education and, as a consequence, longer-term outcomes of the exposed cohorts were negatively impacted. We also find that the wage and occupation distributions were distorted for the exposed cohorts.

The closest paper to ours is Maurin and McNally (2008), who focus on the cohorts exposed to the well-known May 1968 student riots in France. Although the French student riots were very influential, they neither lasted long nor prevented students from continuing higher education due to violence. As a response to the riots, the French government lowered the passing thresholds for critical exams, which increased post-secondary educational attainment for the

3 For example, there is a growing body of literature on the long-term consequences of Holocaust—see, e.g., Waldinger (2010, 2012, 2016), Acemoglu et al. (2011), Grosfeld et al. (2013), Akbulut-Yuksel and Yuksel (2015), and Pascali (2016). Similarly, the long-term effects of the World War II and the Vietnam War have also been studied and documented in detail—see, for example, Davis and Weinstein (2002), Brakman et al. (2004), Ichino and Winter-Ebmer (2004), and Miguel and Roland (2011).

4 For some recent examples in this literature, see Akresh and de Walque (2008) for Rwanda, Leon (2012) for Peru, Verwimp and Van Bavel (2014) for Burundi, Justino et al. (2014) for Timor-Leste, Brown and Velasquez (2017) for Mexico, Bertoni et al. (2019) for Nigeria, Brueck et al. (2019) for Palestine, and Koppensteiner and Menezes (2021) for Brazil.

exposed cohorts. So, the main difference between our paper and Maurin and McNally (2008) is that the student protests generated a decline in school attainment in Turkey, while the French riots increased education for the exposed cohorts. Both papers identify certain birth cohorts as "exposed" and document exogenous changes in the educational attainment of those cohorts.

Another similarity with Maurin and McNally (2008) is that both papers use the abrupt change in tertiary education for the exposed cohorts as an exogenous source of variation to estimate the returns to college in an IV setting. While exogenous variation is relatively easier to obtain at primary or secondary education levels[5], it is rather more difficult to find exogenous events affecting tertiary education. The Turkish student protests led to severe and dramatic outcomes because of the high intensity of violence they generated. More specifically, the violent conflict abruptly and unexpectedly affected tertiary education attainment, which makes the political turmoil a reasonable source of exogenous variation. Therefore, we contribute to this literature by introducing the idea that the exogenous nature of political instability due to student revolts and the resulting decline in school enrollment can be used in an IV setting to estimate returns to higher education.[6] We report a huge drop in post-secondary education for the exposed cohorts. Using the exposed cohorts as an IV, we estimate the college premium to be around 15 percent for an additional year of postsecondary education, which is in the ball-park of typical estimates reported in the literature on returns to higher education.[7] Remarkably, our estimates are very similar to those (around 14 log points) reported by Maurin and McNally (2008).

The instrument that we propose can bring a new perspective in exploring the causal relationship between higher education and earnings—and also other socio-economic outcomes often explored in the literature such as health, crime, religiosity, and voting preferences—in Turkey. Since the student protests were common in many countries in the 1960s and 1970s, a similar IV approach might be implemented also for other countries depending on the context.

3 Background information

3.1 Emergence of civil conflict in Turkey (1960-1980)

The army played a dominant role in Turkish politics by ousting elected governments nearly once in every decade from 1960 to 1980. The 1960 coup marked the beginning of a new era in Turkey. After the military coup, a new constitution was prepared before the free elections in 1961. The new constitution was more liberal and people had more civil rights than ever before. Universities had greater autonomy, students had the freedom to organize their own associations, and workers had the right to strike. As a result, left-wing politics started to gain

5 The most common example is a change in compulsory education laws, which typically expose certain birth cohorts to extended compulsory schooling. In the second half of the 20th century, many countries introduced major educational reforms that increased the years of compulsory schooling. Numerous studies exploit these reforms to estimate returns to schooling in various countries and contexts. Breakthrough papers in this literature include Angrist and Krueger (1991), Harmon and Walker (1995), Duflo (2001), Meghir and Palme (2005), Oreopoulos (2006a,b), Pischke and von Wachter (2008), Aakvik et al. (2010), Devereux and Hart (2010), Fang et al. (2012), Stephens and Yang (2014), Sansani (2015), and Bell et al. (2016).

6 Other examples of instruments used in this literature include college-proximity (which is rather controversial), veteran rehabilitation acts, and discontinuities in college admissions and dismissals.

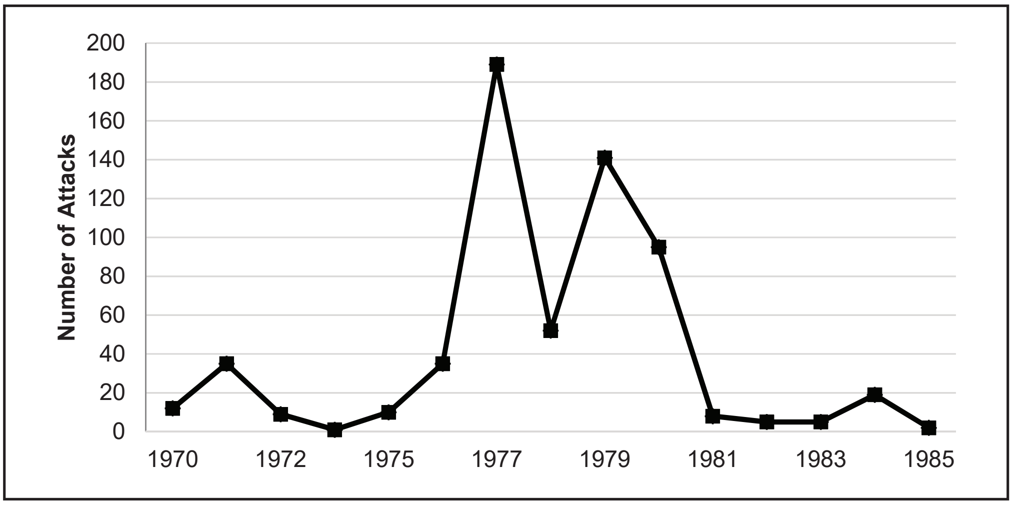
7 See, e.g., Kane and Rouse (1995), Card (1995), Lemieux and Card (2001), Belzil and Hansen (2002), Carneiro (2003), and Ost et al. (2018).

strength, especially on university campuses. Over time, with the push from the global political events of 1968, Turkey's left became more extremist in the hopes of igniting a revolution. But the left's extremism was soon met and surpassed by the right, which generated conflict and violence (Zurcher, 2004). In March 1971, the army forced the elected government to step down and changed the constitution again (*i*) to strengthen the state against civil society; (*ii*) to gain control of the universities and curb radicalism; and (*iii*) to phase out trade unions (Ahmad, 1993, 2003).

The left soon rallied around the Republican People's Party. In 1973, the Republican People's Party won the parliamentary elections and formed a coalition government with the National Salvation Party. The extremist right-wing parties criticized the government program, which sought to heal the wounds caused by the military regime. Radical leftists responded with acts of violence and political violence became a regular feature of daily life in Turkey in the 1970s. Figure 1 presents the total number of terrorist attacks used as a proxy for conflict intensity in Turkey from 1970 to 1985. The data source is the Global Terrorism Database (GTD), which defines "a terrorist attack as the threatened or actual use of illegal force and violence by a non-state actor to attain a political, economic, religious, or social goal through fear, coercion, or intimidation." Figure 1 shows that the attacks declined following the 1971 military intervention, but they increased again after 1974 and were quite intense during the turmoil leading up to the 1980 coup.

In April 1977, political parties agreed on an early election, which led to increased violence intensity. Street terror peaked on the May Day (May 1st) of 1977, four weeks before the early elections. The unions organized a huge rally in Istanbul. Shots fired into the crowd killed 36 people and injured hundreds. Additionally, the 1977 election did not produce a strong and stable government because no party won a clear majority. As a result, Turkey experienced one of its darkest periods in terms of political instability and societal chaos. In July 1978, the government started to use the army due to internal security concerns. Despite the increasing use of force, the violence continued until the slaughter reached 20 victims a day in the late 1970s

Figure 1 Number of terrorist attacks in Turkey (1970-1985).



Source: National Consortium for the Study of Terrorism and Responses to Terrorism (START), 2016. Global Terrorism Database [GTD from 1970 to 1991]. Retrieved from https://www.start.umd.edu/Gtd.

(Ahmad, 1993). From 1978 to 1980, 5,241 people were killed and 14,152 people wounded due to violent political conflict (Kaya, 1981).

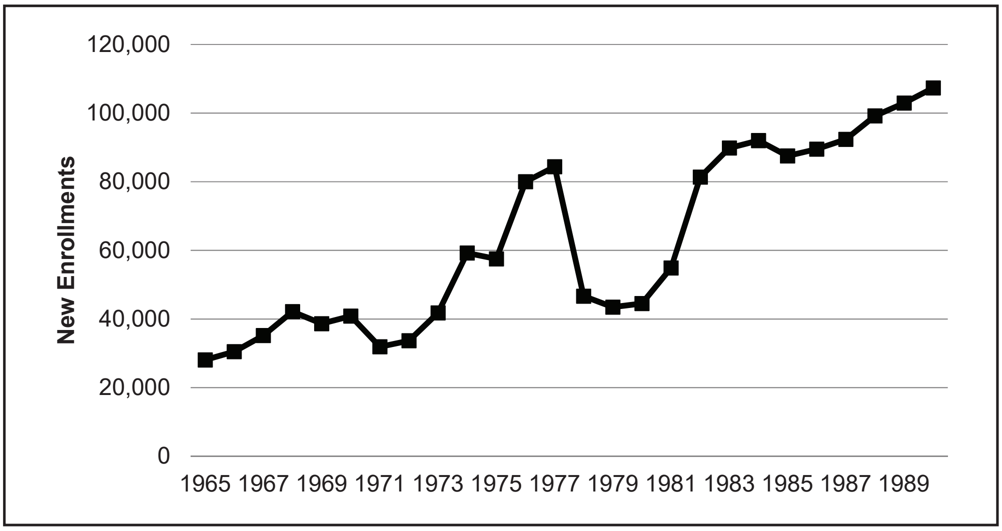
The army took control in September 1980 and ruled until the general elections of November 1983. The public welcomed the military intervention, and the army crushed almost all movements from the left and right to de-politicize urban youth (Ahmad, 1993). In the first three months after the coup, 30,000 people were arrested, which increased to 122,600 after a year. By September 1982, 80,000 people were still under arrest, with 30,000 awaiting trial (Zurcher, 2004). Meanwhile, the number of terrorist attacks declined by 90 percent after the intervention (Figure 1).

3.2 **Channels that affected tertiary education during the turmoil**

Violent conflict adversely affected post-secondary educational attainment in Turkey through several channels. **First**, new enrollments in post-secondary education declined in the 1978-1979 school year (Figure 2). Based on the Turkstat data, the decline in the number of students enrolled is 37,715. This decline was mostly related to the closure of state-controlled educational institutions, which were highly involved in the conflict and violence. Those institutions included academies, vocational schools, and teacher-training institutes that were directly affiliated with certain ministries. Yet, most violence among students was seen in those institutions (Binbasioglu, 2005; Tekeli, 2010), which led the government to close 41 institutions out of 64 in 1978. Students already enrolled in the closed institutions were allowed to complete their programs, but new students were not admitted.

After the 1980 coup, the Council of Higher Education was established as a governing board to plan, coordinate, and review the activities of higher education institutions in Turkey (Dogramaci, 1989). This central body would also determine the enrollment capacity of post-secondary education institutions. Figure 2 indicates that enrollments started to increase right after 1982, when all ministry-affiliated higher education institutions were reorganized under the university system.

Figure 2 New enrollments in all higher-education institutions in Turkey (1965-1990).



Source: Authors' calculations based on National Education Statistics compiled by TURKSTAT and Academic Year Higher Education Statistics compiled by OSYM.

The **second** channel that adversely affected educational attainment was that graduation rates declined because of massive student dropouts related to security concerns. Due to the high death toll, many students withdrew their registration in higher education institutions (Kaya, 1981). Some students were unable to finish their education because they were injured or disabled during the turmoil. In addition, some families chose not to send their children to higher education in this period due to heightened risks. Courses were often suspended or canceled during this time. For instance, classes were canceled for 116 days in Ege University and for 421 days in Istanbul University—two of the largest universities in Turkey. The School of Dentistry in Hacettepe University was completely closed during the 1979-1980 academic year (Kaptan, 1986). Faculty offices and student dormitories were often used as weapon warehouses and arsenals (Kaya, 1981; Kaptan, 1986; Gunter, 1989).

And, **third**, mass student arrests in the wake of the 1980 coup kept many from completing their education. According to a Turkish government report, by 1981, one year after the coup, 9,760 of the state's "captured terrorists" were students. Moreover, 57 percent of the state's 43,140 "captured terrorists" were of age 16-25 (and most were men). In addition to these channels, new enrollments in open education declined by 12,479 between 1977 and 1978.

4 Data

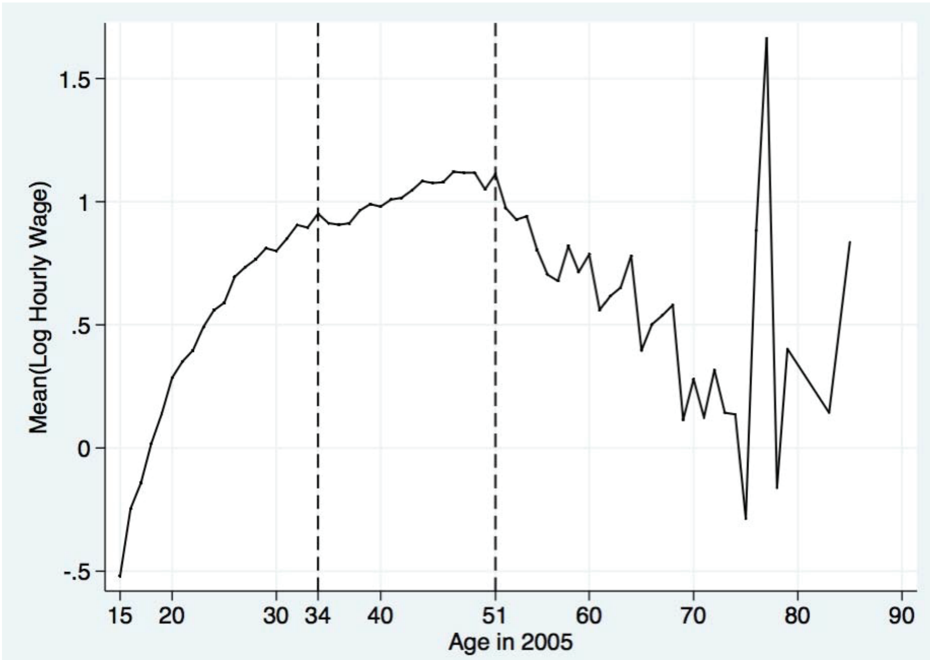
We use the Turkish Household Labor Force Surveys (HLFS) in our empirical analysis.[8] The Turkish Statistical Institute (Turkstat) has prepared and published the HLFS micro-level data sets in accordance with Eurostat's guidelines since 2004 on an annual basis. As 2004 was a transition year in terms of the general structure of the survey (Tunali, 2022), the earliest reliable wave is the 2005 wave. We aim to capture the labor market outcomes of the exposed individuals at the time when they were still within the core group in the labor market. The average wage for males is an increasing function of age; but, after age 51, it sharply falls and tends to be quite volatile (Figure 3). In addition, the number of wage observations goes down significantly after age 51, since older workers tend to leave the labor market due to early retirement, human capital depreciation, and other factors. To avoid the loss of observations, we mainly focus on the 2005 wave of the survey. But, we also use the 2004, 2006, 2007, and 2008 HLFS waves for reliability checks. The HLFS data set provides detailed information on age, education, labor market status, number of hours per week usually worked in the main job, earnings of individuals from the main job during the past month including any irregular payments like bonus and premiums, and main tasks/duties of individuals in the workplace.

The data set does not directly include a variable for actual labor market experience; thus, we use potential experience as proposed by Mincer (1974): $X = A - S - B$, where A is current age, S is years of schooling, and B is school-start age. The typical school-start age was 7 before 1980 in Turkey. Moreover, the HLFS does not directly report the years of schooling; instead, it reports the highest level of education completed.[9] However, Turkish Demographic and Health Surveys (TDHS) contain information on both graduation and years of schooling. We estimate

8 HLFS is a nationally-representative micro-level data set compiled and published by the Turkish Statistical Institute. It is used to produce the official labor market statistics in Turkey.

9 Note that, for the post-secondary degree, we only observe the category "college or above," which means that we cannot distinguish between individuals with 2-year college, 4-year college, and graduate degrees.

Figure 3 Average log hourly wages for men by age.



Source: Authors' calculations based on the 2005 Turkish HLFS.

the mean years of schooling conditional on the highest completed schooling level by using the 2008 TDHS. We find that the average years of schooling is 0.14 years for illiterate individuals, 1.68 years for literate individuals with no educational degree, 5.09 years for primary school graduates, 8.34 years for elementary school graduates, 11.09 years for high school graduates, and 14.63 years for postsecondary school graduates.[10] Based on this supporting evidence, we impute the average years of schooling for the post-secondary education as 4 years in our HLFS data sets.

The data set includes monthly wages—the average wage is 588 Turkish liras in the 2005 wave. Card (1999) emphasizes that the estimated coefficient of annual earnings could comprise the effect of schooling on hourly earnings, hours per week, and weeks per year. Also, in the US data, individuals with higher years of schooling tend to work more. In contrast, there is a negative correlation between schooling and the number of hours worked in Turkey (Table 1); as schooling increases, average hours worked in the main job fall. The pairwise correlation coefficient between hours worked and mean years of schooling is -0.3. Therefore, we choose hourly wages as the measure of labor income. We calculate hourly wages as the monthly wage in the main job divided by (52/12) and, then, by the number of hours per week usually worked in the main job.

The average potential experience of male wage earners for the age group 34-51 is 26 years. In all regressions, we standardize log hourly wages at 26 years of potential experience because our treatment and comparison groups have different experience levels and explicitly controlling for experience in our regressions may create a collinearity problem as we define the treatment and control groups based on age (or birth cohort). We estimate a log hourly wage

10 In the TDHS exercise, we restrict our sample to individuals of age 37-54. A similar strategy is implemented by Aydemir and Kirdar (2017).

Table 1 Average hours worked in the main bob by educational attainment in the sample of wage earners in Turkey

Educational attainment	# of observations	Mean
No schooling	3,305	55.3
Primary school (5 years)	26,065	55.5
Elementary school (8 years)	11,046	54.9
High school	19,498	51.8
Post-secondary degree	13,396	44.1

*Source***:** Authors' calculations based on the 2005 HLFS. Observations are weighted using the sampling weights so that the results are nationally representative.

Table 2 Descriptive statistics for individuals of age 34-51

Variables	Mean
Primary or elementary sch. grad. rate	0.63
High sch. grad. rate	0.14
Post-secondary sch. grad. rate	0.08
Years of schooling	6.36
Labor force participation	0.58
Employment rate	0.54
Sample size	115,410

Observations are weighted using the sampling weights so that the results are nationally representative.

equation separately for each education category defined in the survey data for this group. These categories are no degree, primary school (five-year), middle school (eight-year), high school, and post-secondary degrees. Following Altonji et al. (2012), we include potential experience as a quartic function and, from these regressions, we compute the predicted log hourly wage for a common experience of 26 years and add the residual.

Table 2 provides descriptive statistics for the 34-51 age group. Among this group, 63 percent have primary or middle school degree, 14 percent have high school degree, and approximately 8 percent have a post-secondary degree. In addition, the employment rate is 54 percent, while the labor force participation rate is 58 percent.

5 Empirical analysis, results, and discussion

5.1 Baseline estimates

In this section, we argue that the turmoil generated a substantial decline in post-secondary education among the exposed population—the ones, especially males, born between 1960 and 1965. The decline in higher education as a consequence of a political turmoil in a developing country is an interesting finding in its own. However, we do not stop at this point. In Section 5.2, we use this decline as an instrument based on exposed versus non-exposed cohorts to estimate the effect of post-secondary education on labor market earnings. Finally, in Section 5.3, we show that the decline in post-secondary education during the turmoil had severe negative implications on adult wage and occupational distributions.

In addition to documenting the decline in post-secondary education, this section also provides a background analysis for determining the birth cohorts exposed and non-exposed to the turmoil, e.g., the treatment versus control groups, that will be used in the IV analysis. Although we carry out our baseline empirical analyses focusing on specific treatment and control cohorts, we relax these restrictions in Section 5.2.4, where we perform several robustness checks, and show that the qualitative nature of our baseline results is not overly sensitive to inclusion or exclusion of certain birth cohorts in defining the treatment and control groups. It should also be noted that the baseline analysis is performed using the 2005 wave of the HLFS data set, but the results are robust to including more HLFS waves—as we show in Section 5.2.4.

Post-secondary education. Post-secondary school enrollment rates increased dramatically during the second half of the 20th century all over the world (Psacharopoulos, 1991). Turkey was no exception. However, in Turkey, the enrollment and graduation rates in post-secondary education substantially declined between 1978 and 1982 due to the political turmoil. To analyze the trends in post-secondary education, we use the following regression model:

$$s_i = \alpha + \sum_{c=30}^{54} \beta_c d_{ic} + \boldsymbol{X}_i' \boldsymbol{\Pi} + \epsilon_i, \tag{5.1}$$

where s_i is a binary variable indicating whether individual i has a post-secondary degree, d_{ic} is an age dummy indicating whether individual i is c years old, $\boldsymbol{X}_i$ is a vector of covariates, and ϵ_i is an error term.

The age dummies cover the 30-55 age interval. The vector of covariates include 26 NUTS2-level region-of-residence dummies, an urban/rural dummy, and a gender dummy. We omit age 55 in the regressions; therefore, each coefficient β_c can be interpreted as the probability of completing post-secondary education for the corresponding age relative to age 55. In a developing country, one would normally expect β_c to increase as age declines—in the absence of a negative shock on post-secondary education.

Figure 4 plots the estimated β_c's for the whole sample. The solid line shows the probability of completing post-secondary education, while the dashed lines indicate the 95-percent confidence interval. The substantial drop in post-secondary education between ages 40 and 45 is not typically observed in developing countries and may be hard to justify in the absence of a big negative shock affecting post-secondary educational outcomes for those cohorts. Figures 5 and 6 plot the estimated coefficients for men and women, respectively. Although the trends for completing post-secondary education look similar in both figures, the effect of the turmoil is much more pronounced for men than that for women. Based on these results, individuals of age 40-45 are marked as "exposed individuals" throughout the analysis. Those individuals were born between years 1960-1965, and were about 13 to 18 years old as of 1978.[11]

The 40-45 age group has 6 age categories. We construct a comparison group which also has 6 age categories: the 46-51 age group. We choose this older group as the comparison group for two main reasons. First, after the coup, the educational institutions had changed and many

11 Figure 2 shows that first-year enrollments in higher education declined significantly for the first time in 1978 and remained low until 1982. This decline probably affected young adults of age 17 and 18 from 1978 to 1982. Therefore, the affected group was approximately from 13 to 18 years old in 1978, which is line with the findings reported in Figures 4-6. In addition, student dropouts related to security concerns and mass student arrests after the coup also affected the educational attainment in this age group. Based on these findings, we assert that men of age 40-45 (in the 2005 survey) were the ones most severely affected by the turmoil.

Figure 4 Coefficients of age dummies – Probability of completing post-secondary education.

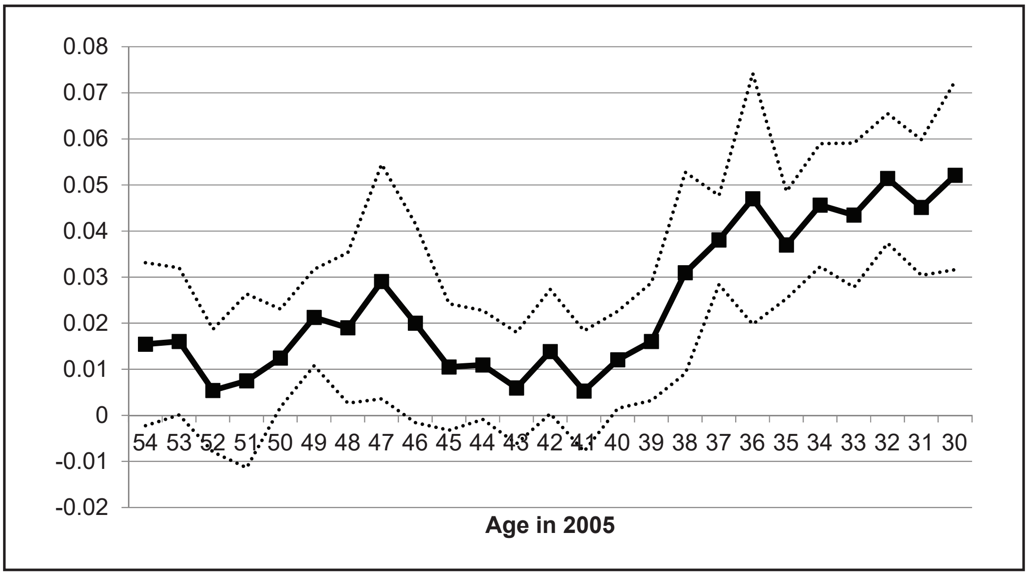
The specification includes region of residence, urban/rural, and gender dummies. Age 55 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

Figure 5 Coefficients of age dummies – Probability of completing post-secondary education (men).

The specification includes region of residence and urban/rural dummies. Age 55 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

new universities were constructed; so, the younger birth cohorts cannot be used as a relevant comparison group. Second, older cohorts were exposed to similar political conditions except violence. Given a natural upward trend in a developing-country context, one would expect a higher post-secondary educational attainment for the 40-45 age group relative to the 46-51

Figure 6 Coefficients of age dummies – Probability of completing post-secondary education (women).



The specification includes region of residence and urban/rural dummies. Age 55 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

age group. Figure 4 shows that the probability of completing post-secondary education for the exposed group is clearly less than that for the comparison group. Alternative age groups are used in Section 5.2.4 for robustness purposes.

Different levels of education. Next we compare the educational attainment of age groups 40-45 and 46-51 using the following regression model:

$$s_i = \alpha + \beta z_i + \boldsymbol{X}_i' \boldsymbol{\Pi} + \epsilon_i, \tag{5.2}$$

where s_i is a binary variable indicating whether individual i has graduated from a school (post-secondary, high school, elementary/primary school) or not, z_i is a dummy variable taking 1 if the individual i is in the age group 40-45 and 0 if 46-51, $\boldsymbol{X}_i$ is a vector of covariates, and ϵ_i is an error term.[12] The coefficient β can be interpreted as the gap between probability of graduation from school between age groups 40-45 and 46-51. One would normally expect β to be positive in the absence of a negative shock on education.

Table 3 presents three sets of estimates from Equation 5.2. Column 1 displays the result for postsecondary degree, column 2 for only high school degree, and column 3 for only elementary/primary school degree. Column 1 shows that the probability of completing post-secondary education is 1.5 percentage points lower for the 40-45 age group. In contrast, the probabilities of graduation from elementary/primary school and high school increase significantly as expected—see columns 2 and 3. However, the increase in the probability of graduation from high school is 4.5 percentage points, or about twice the increase in the probability of graduation from elementary/primary school. This suggests that those individuals affected from the protests would have normally gone to or completed a post-secondary education, but could not

12 We use the same vector of covariates as in Equation 5.1.

do so due to the 1978-1982 turmoil. So, the number of high school graduates increased more than it normally would, which implies that potential college graduates have remained as high school graduates as a consequence of the turmoil.

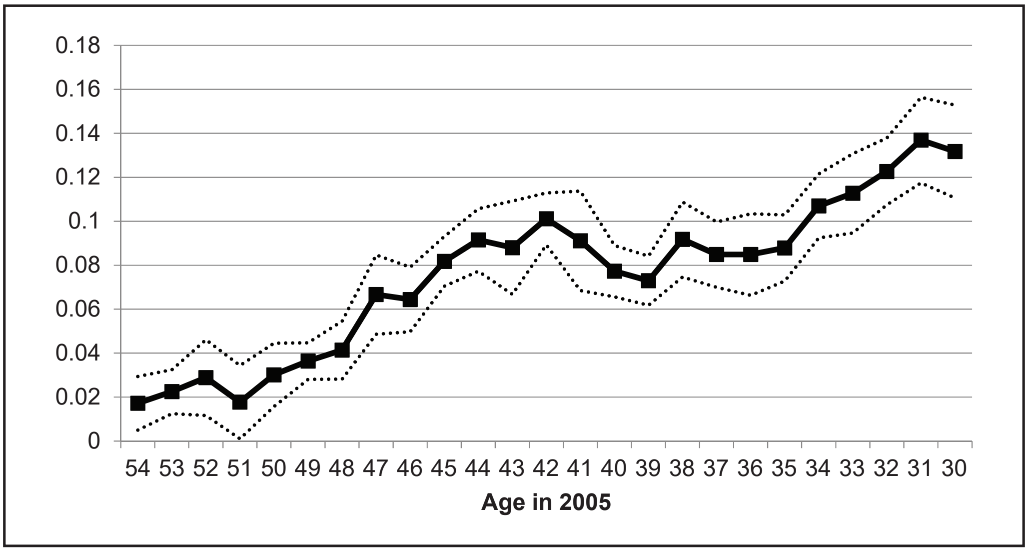
We also compare the age groups 46-51 and 52-57. The high school graduation probability is 2.9 percentage points higher for the 46-51 age group. The difference between 40-45/46-51 and 4651/52-57 differences is 1.6 percentage points—approximately equal to the percentage point decline in completing post-secondary education for the 40-45 age group. This tells us that the main group affected from the student protests were those who would have normally continued post-secondary education in the absence of the turmoil. Figures 7 and 8 confirm this point.

Table 3 Difference in the probability of graduation between the 40-45 and 46-51 age groups

	Dependent variable		
	Post-secondary	**High school**	**Primary/Elementary**
	Degree==1	**Degree==1**	**Degree==1**
	Otherwise==0	**Otherwise==0**	**Otherwise==0**
	[1]	**[2]**	**[3]**
Age 40–45	−0.0148***	0.0450***	0.0325***
	(0.0051)	(0.0027)	(0.0092)
# of obs.	74,903	74,903	74,903
R^2	0.0364	0.0375	0.0524

The specification includes 26 NUTS2 region of residence, urban/rural and gender dummies. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

Figure 7 Coefficients of age dummies – Probability of high school graduation.



The specification includes region of residence, urban/rural, and gender dummies. Age 55 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

Figure 8 Coefficients of age dummies – Probability of graduation from elementary/primary school.

The specification includes region of residence, urban/rural, and gender dummies. Age 55 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

Probability of wage employment. We check whether the turmoil had any effect on the probability of wage employment. In the regression, the dependent variable is a dummy variable taking 1 if an individual is wage employed (regular employee or casual employee) and 0 otherwise (employer, self employed, or unpaid family worker). The sample includes all employed individuals. Figure 9 plots the estimated coefficients of age dummies. The trend is plausibly smooth over the age horizon; so, the turmoil had no effect on wage employment.[13]

Although the probability of being a wage-earner did not change, the school attainment among wage-earners of different age groups might have changed as a consequence of the turmoil. To test this conjecture, we estimate the probability of post-secondary educational attainment for wage earners (column 1) and non-wage earners (column 2), separately. Table 4 reports the results. The estimates clearly suggest that wage-earner men of age 40-45 are the ones whose post-secondary educational attainment have been affected the worst. We conclude that the group most affected from the turmoil is wage-earner men of age 40-45 (in the 2005 survey, with birth years from 1960 to 1965).

5.2 **IV estimates for returns to higher education**

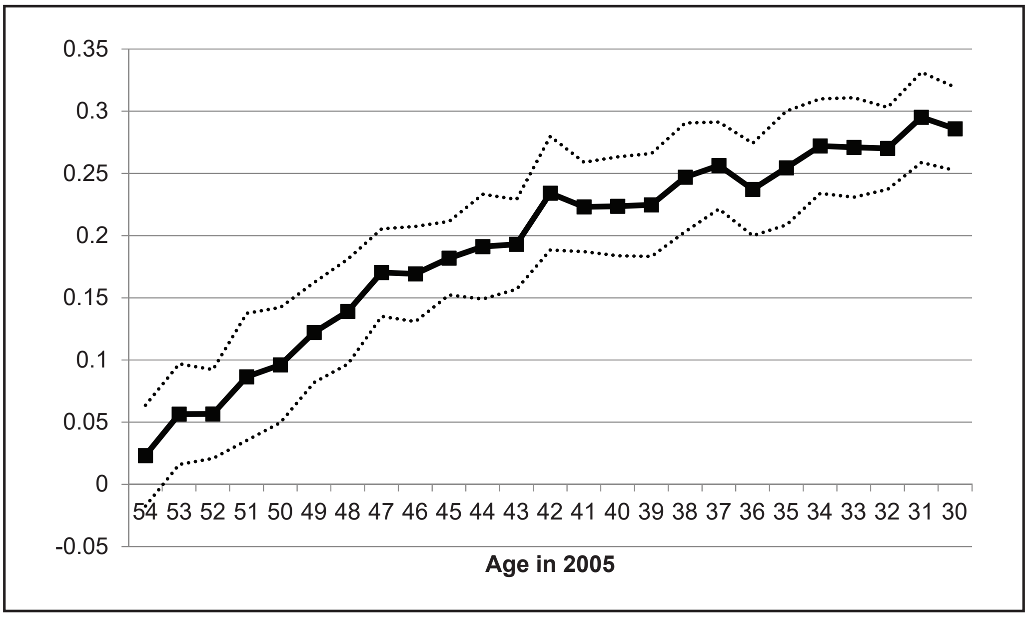
The following Mincerian setting is often used to estimate the effect of education on wages:

$$w_i = \alpha + \beta s_i + \boldsymbol{X}_i' \boldsymbol{\Pi} + \epsilon_i, \tag{5.3}$$

where w_i is a measure of labor income, s_i is a measure of schooling, $\boldsymbol{X}_i$ is a vector of observables, and ϵ_i is an error term assumed to be independent of the explanatory variables

13 We show in Section 5.3 that there is also no impact on labor force participation, formal employment, and informal employment.

Figure 9 Coefficients of age dummies – Probability of wage employment.



The specification includes region of residence, urban/rural, and gender dummies. Age 55 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

Table 4 The effect of the turmoil on the probability of completing post-secondary education

	Dependent variable: Post-secondary degree==1; Otherwise==0					
	Total		Men		Women	
	[1]	[2]	[1]	[2]	[1]	[2]
Age 40-45	−0.0587***	−0.0084**	−0.0664***	−0.0102	−0.0147	−0.0031
	(0.0075)	(0.0034)	(0.0114)	(0.0065)	(0.0247)	(0.0053)
# of obs.	18,730	18,852	15,827	12,798	2,903	6,054
R^2	0.0476	0.0570	0.0298	0.0474	0.0780	0.1077

Region of residence and urban/rural dummies are included in all regressions. A gender dummy is also included for the total sample estimations. Columns [1] and [2] in each of the three separate regressions present results for wage earners and non-wage earners, respectively. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

(Griliches, 1977). In this setting, the causal effect of schooling on labor income may not be consistently estimated because of the classical omitted variables problem—also known as the ability bias. A possible solution to this problem is to use an instrument, which requires an exclusion restriction, i.e., at least one observable covariate that affects labor income only through schooling.

In Section 5.1, we show in detail that post-secondary educational attainment declined significantly for individuals born between 1960 and 1965 due to the student protests in the late 1970s and the subsequent military coup. We also argue that we set the non-exposed cohort—the ones born between 1959 and 1954, i.e., the 46-51 age group in the 2005 HLFS wave—as the comparison group.[14] Therefore, we use the dummy variable z_i—taking 1 if the individual i belongs to age group 40-45 and 0 if s/he belongs to age group 46-51—as an IV for estimating the returns to college. In Section 5.2.4, we perform robustness checks using alternative age intervals as treatment and control groups.

In a heterogeneous-outcome framework, the IV method potentially estimates the average treatment effect (ATE) of schooling on earnings for the sub-group whose schooling attainment is changed by the instrument—i.e., the local average treatment effect (LATE) (Imbens and Angrist, 1994; Angrist et al., 1996; Card, 2001). There are two key conditions (Imbens and Angrist, 1994). The first one is the existence of a valid instrument. Because an individual's year of birth—within a reasonably narrow year of birth interval—is randomly assigned and probably unrelated to individuals' innate ability, personal characteristics, or family characteristics, it seems reasonable to assert in our case that the wage decline for the 40-45 age group relative to the 46-51 age group is due to the exposure to turmoil and the associated educational disruptions—after standardizing labor market experience and controlling for other observables. Thus, potential outcomes should be independent of the instrument and the exclusion restriction assumption should be satisfied. We show in Section 5.1 that the probability of completing post-secondary education is related to z_i in a non-trivial way. The second condition is monotonicity. This condition ensures that the instrument affects the post-secondary education in a monotonic way (Imbens and Angrist, 1994; Angrist et al., 1996). We also document in detail that the 1978-1982 turmoil negatively affected all sub-samples of the relevant population, which suggests that the monotonicity condition should also be satisfied. Based on these arguments, our IV estimates can be interpreted as the local average treatment effect for those who did not continue post-secondary education due to the 1978-1982 turmoil, but who would have normally obtained a post-secondary degree.

One potential concern about the relevance of our instrument is the possibility that the events leading to post-secondary education disruptions might be a response to declining returns to higher education. We believe that this is not the case as the student protests were not specific to Turkey and happened parallel to global political developments. In fact, it is well

Table 5 Comparisons of age groups for male wage earners

	[1]	[2]	[3]	[4]	[5]
Age 34-39	10,774	1.023	8.363	0.166	0.228
Age 40-45	10,105	1.002	8.198	0.142	0.243
Age 46-51	5,722	1.037	8.475	0.211	0.195

Columns [1] - number of observations; Column [2] - mean log hourly wage, Column [3] - mean years of schooling; Column [4] - mean post-secondary attainment rate; Column [5] - mean high school completion rate. Observations are weighted using the sampling weights so that the results are nationally representative.

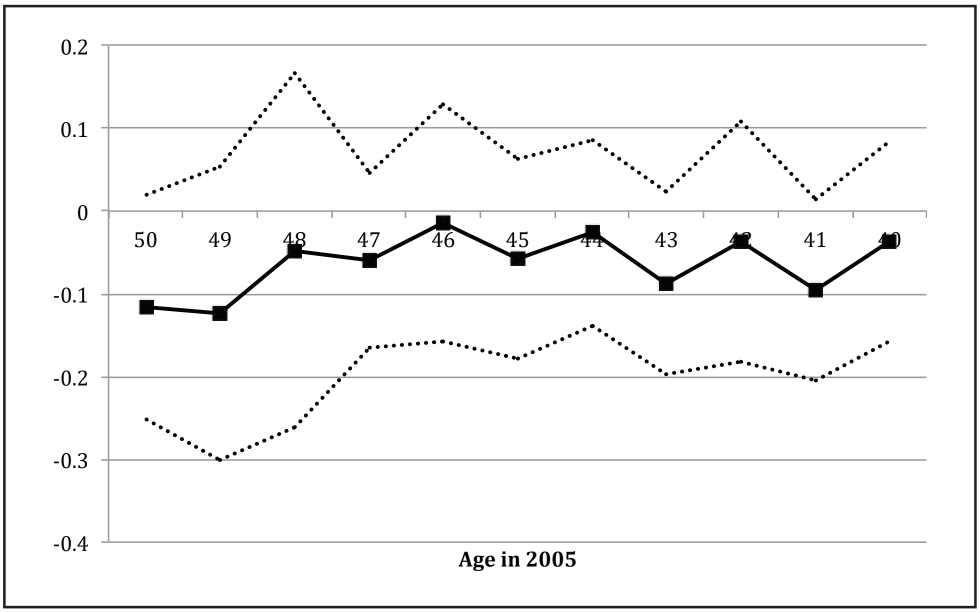
14 The 34-39 age group may not be an appropriate comparison group as they were subject to post-coup educational institutions, which enhanced post-secondary education opportunities through new university openings and reorganization of tertiary education system. However, our findings are robust to using preceding or succeeding cohorts as IV. Table 5 reports summary statistics for three age groups: 34-39, 40-45, and 46-51. The 40-45 age group clearly has lower average log hourly wage, fewer average years of schooling, and lower average post-secondary educational attainment than the younger and older groups. Moreover, this group has a higher average high school graduation rates compared to other age groups.

known that the college premium significantly increased across the world during that period (Juhn et al., 1993; Oreopoulos and Petronijevic, 2013).

5.2.1 Other effects as confounding factors? A test for instrument validity

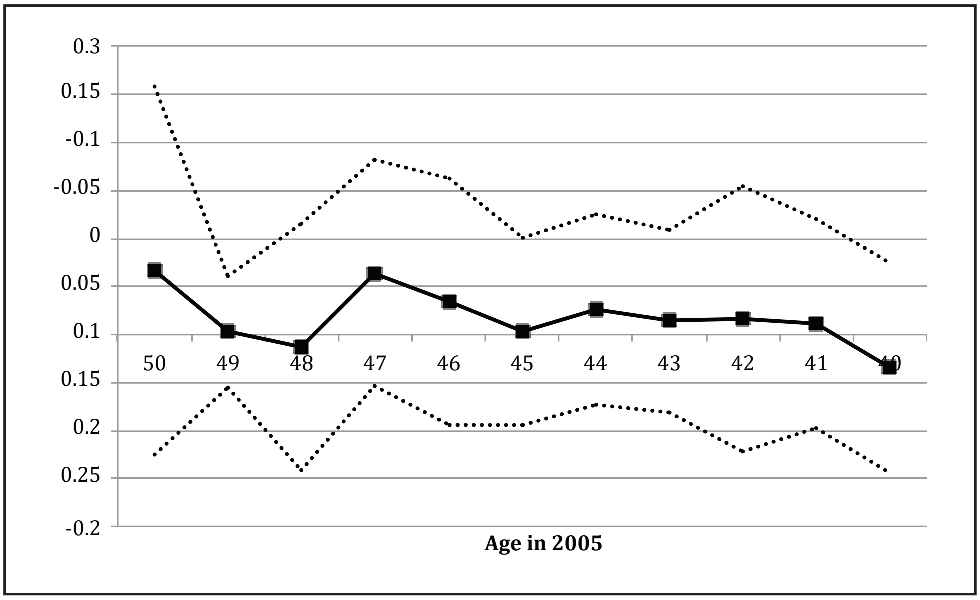
One potential threat to instrument validity is the presence of other confounding variables, such as taste differences for risk or discount factors, psychological factors, a fall in the quality of schooling for the exposed cohorts, or different entry effects into the labor market during the turmoil that would affect wages of the exposed individuals independent from schooling quantity outcomes. If confounding factors, other than schooling, influence the wages of the 40-45 age group due to the turmoil, one would observe this by looking at wages "within the same education group." The individuals affected from the turmoil are high school graduates, who could not complete their postsecondary education. To address this concern, we run separate regressions (similar to Equation 5.1) for both high school and college graduates in which the dependent variable is the log hourly wage standardized to 26 years of labor market experience following Altonji et al. (2012) and the explanatory variables are the age dummies, dummies for region of residence, and an urban/rural dummy. We omit age 51 in the regressions; therefore, the coefficients of age dummies can be interpreted as the gap between the wage of the corresponding age relative to age 51. By these regressions, we also test whether the composition of high school and college graduates changed due to the turmoil in terms of wages. In the regressions, the first sample includes male wage earners of age 40-51 with only a high school degree, the second sample includes only college graduates for male wage earners of age 40-51. The coefficients of age dummies are plotted in Figures 10 and 11, respectively.

Figure 10 Coefficients of age dummies – Log hourly wages for male wage earners having only a high school degree.



The sample includes male wage earners aged 40-51 having only high school degree. Log hourly wage is standardized to experience 26 years. The specification includes region of residence and urban/rural dummies. Age 51 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

Figure 11 Coefficients of age dummies – Log hourly wages for male wage earners having only a postsecondary degree.



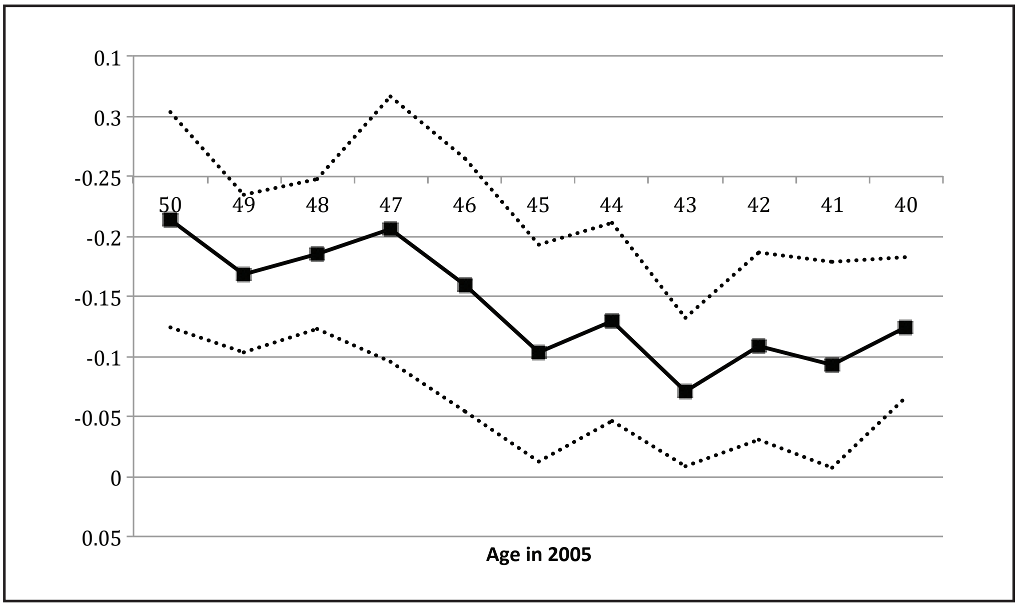
The sample includes male wage earners aged 40-51 having only post-secondary degree. Log hourly wage is standardized to experience 26 years. The specification includes region of residence and urban/rural dummies. Age 51 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

They lie almost on a straight line and none of the age coefficients are statistically significant. Therefore, the log hourly wages of high school and college graduates born between 1960-1965 (e.g., the exposed ones) are not statistically different from those of the non-exposed cohort. In other words, the wage pool of high school and college graduates did not change due to the turmoil. This suggests that the decline in earnings for the exposed cohorts is solely due to the decline in their post-secondary educational attainment, not due to other confounding factors.

To assess whether the decline in post-secondary education for men of age 40-45 is reflected on earnings, we run a third regression for the high school and college graduate men because we already show that the instrument affects only the post-secondary education. The estimated coefficients of ages are plotted in Figure 12 and, clearly, log hourly wages begin to decline for younger cohorts after age 47, similar to the trend in post-secondary educational attainment in Figure 4. The substantial drop in wages between ages 40 and 45 are statistically significant. We also show the average wage for males for only high school graduates, only college graduates, and both high school and college graduates in the same plot in Figure 13. Clearly, the average log hourly wages of the exposed and non-exposed groups follow almost a straight line for only high school and only college graduates. However, the mean wage of the exposed group is smaller than that for the comparison group. This difference is also statistically significant based on Figure 12. These findings suggest that there are likely no permanent psychological or other confounding effects of the political turmoil on wages.

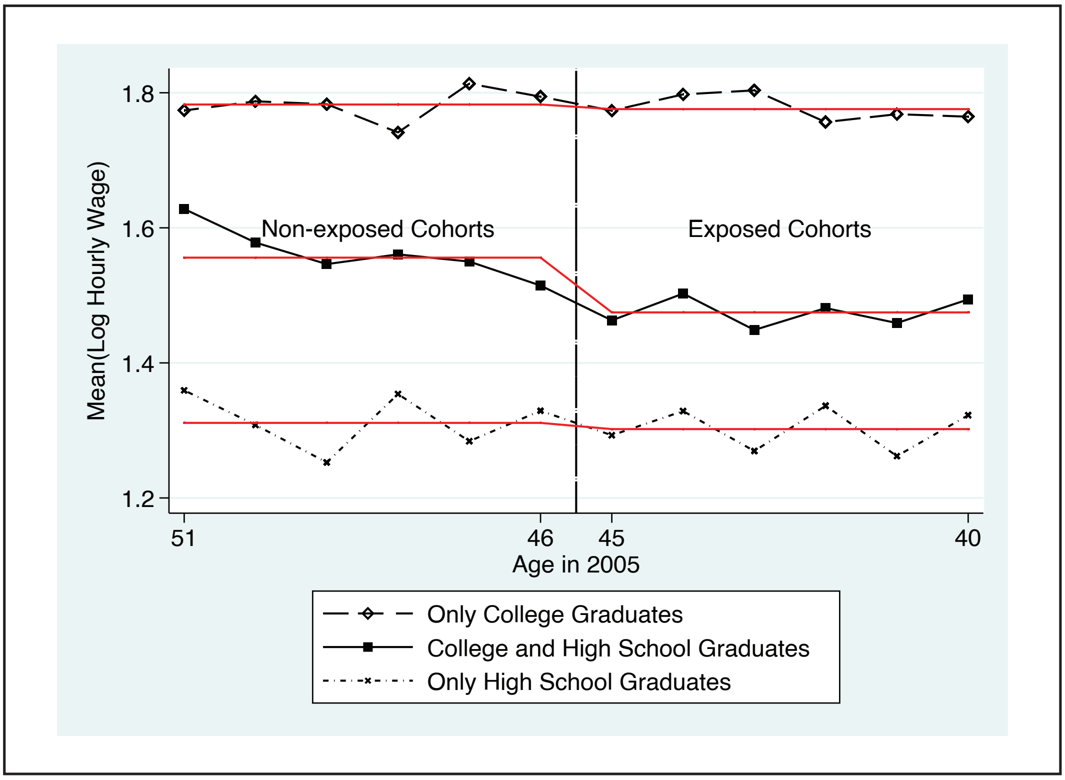
As a side note, since we analyze wage earnings measured about 25 years after the 1978-1982 turmoil, any initial/temporary effects might have vanished over time.

Figure 12 Coefficients of age dummies – Log hourly wages for men with at least a high school degree.



The sample includes male wage earners aged 40-51 with at least a high school degree. Log hourly wage is standardized to experience 26 years. The specification includes region of residence and urban/rural dummies. Age 51 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

Figure 13 Average log hourly wages for men by age.



Source: Authors' calculations based on the 2005 Turkish HLFS. The red lines indicate the average log hourly wages for exposed and non-exposed individuals for the corresponding educational level.

5.2.2 **First-stage and reduced-form estimates for male wage earners**

We run two different regressions based on the following equation:

$$y_i = \alpha + \beta z_i + X_i' \Pi + \epsilon_i, \tag{5.4}$$

where z_i is a dummy variable taking 1 if the individual i is of age 40-45 and 0 if 46-51, X_i is a vector of covariates, and ϵ_i is an error term.

In the first regression, the dependent variable, y_i, is a dummy indicating whether the individual has completed post-secondary education or high school. In the second regression, we use log hourly wages as the dependent variable. In all regressions, we focus on male wage earners of age 40-51 with at least a high school degree in the 2005 wave. The results are presented in Table 6. Columns 1 to 3 indicate that the probability of finishing post-secondary education is at least 15 percentage points lower for the treatment group relative to the control group. The last three columns document the effect of the turmoil on wages. The corresponding estimates suggest that wages of men in the 40-45 sample is 8.6-8.8 percent lower than wages of men in the 46-51 sample. All coefficients are highly statistically significant.

The main identifying assumptions are that (*i*) post-secondary educational attainment for the 40-45 sample would not be lower than that for the 46-51 sample in the absence of the political turmoil and (*ii*) the turmoil affects wages only through its impact on post-secondary educational attainment. Based on these assumptions, we use birth cohorts as IV to estimate the causal effect of getting a post-secondary degree versus a high school degree on wages among males. We redefine the variable s_i as follows: taking 1 if i has a post-secondary degree and 0 if high school. We also calculate the average years of schooling for the post-secondary education as 4 years based on the 2008 TDHS. This suggests that the typical post-secondary education is 4-year college in Turkey. Therefore, our estimates can be interpreted as the returns to a 4-year college degree relative to a high school degree. The first-stage and reduced-form of this IV estimations are presented in Table 6. The results suggest that the instrument has satisfactory explanatory power at the first stage.

5.2.3 **Estimating the returns to post-secondary education**

Estimates for the returns to a 4-year college degree are presented in Table 7. In all regressions, the dependent variable is the log hourly wage standardized to 26 years of potential experience[15] and the sample is male wage earners of age 40-51 with at least a high school degree. Panel A of Table 7 provides estimates for just-identified models, in which only one dummy variable—taking 1 for the 40-45 age group and 0 for 46-51—is used as the instrument. The first line presents OLS estimates of Equation 5.3. Column 1 indicates that the returns to a 4-year college degree are 50 log points greater than those for a high school degree and this estimate is not affected by introducing region-of-residence and urban/rural dummies as control variables—see columns 2 and 3.

The second line of Panel A reports the 2SLS estimates using only one instrument. In column 1, there are no control variables and the point estimate, 58 log points, is slightly above the OLS estimate. The protests could be more widespread across some regions due to some unobserved factors that are possibly correlated with schooling and labor market outcomes.

15 The results are almost the same when the log hourly wage is standardized to age rather than the potential experience.

Table 6 The effect of the turmoil on the probability of completing post-secondary education and wage

	Dependent variable					
	Post-secondary degree			Log hourly wage		
	[1]	[2]	[3]	[4]	[5]	[6]
Instrument (z_i)	−0.1507***	−0.1518***	−0.1502***	−0.0873***	−0.0879***	−0.0859***
	(0.0214)	(0.0211)	(0.0207)	(0.0160)	(0.0159)	(0.0154)
Region of residence	No	Yes	Yes	No	Yes	Yes
Urban/rural status	No	No	Yes	No	No	Yes
# of observations	6,309	6,309	6,309	6,309	6,309	6,309

The sample includes male wage earners aged 40-51 with at least a high school degree. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

Table 7 OLS and 2SLS estimates of the returns to college

	Dependent variable: Log hourly wage		
	[1]	[2]	[3]
Panel A: Just-identified models			
OLS	0.5022***	0.5062***	0.5025***
	(0.0320)	(0.0304)	(0.0311)
2SLS	0.5795***	0.5790***	0.5716***
	(0.0965)	(0.0935)	(0.0935)
F-statistic (first stage)	139.54	142.56	140.13
Panel B: Over-identified models			
2SLS	0.5688***	0.5659***	0.5588***
	(0.0887)	(0.0855)	(0.0841)
LIML	0.5704***	0.5676***	0.5605***
	(0.0908)	(0.0880)	(0.0865)
F-statistic (first stage)	25.01	25.51	25.15
p-value Hansen's test	0.87	0.80	0.80
Region of residence	No	Yes	Yes
Urban/rural status	No	No	Yes
# of observations	6,309	6,309	6,309

The sample includes male wage earners aged 40-51. In just-identified models, a dummy for individuals of age 40-45 is used as instrument. In over-identified models, six age dummies in the 40-45 age group are used as instruments. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

To capture this possibility, region-of-residence and urban/rural dummies are also included in the IV regressions. We find that including region-of-residence and urban/rural dummies as control variables (columns 2 and 3) do not change the results significantly.

Panel B of Table 7 shows the results of the over-identified models, in which six age dummies within the age interval 40-45 are used as instruments. The results are very similar to the IV estimations using only one instrument. The F-statistics on the excluded instruments in all regressions are much higher than the Staiger and Stock (1997) rule of thumb of 10, suggesting that the weak instrument problem may not be a serious concern. However, we follow the suggestion of Angrist and Pischke (2009) and report limited-information maximum likelihood (LIML) estimates in case of the possibility of weak instruments. Stock and Yogo (2005) and Angrist and Pischke (2009) argue that LIML is superior to 2SLS in weak-instruments cases. Our LIML estimates are almost identical to 2SLS. In addition, high p-values of Hansen's test indicate that the null hypothesis of instrument validity is not rejected at conventional significance levels.

Overall, our IV estimates suggest that the returns to 4-year college degree is around 56-58 log points for men. This corresponds to around 15 percent returns to an additional year of post-secondary education.[16]

5.2.4 Robustness checks

Fuzzy regression discontinuity analysis. We use a fuzzy regression discontinuity design (RDD) as an alternative estimation strategy to test the robustness of our baseline results. The main idea is that college attainment changes discontinuously based on age cohorts; therefore, there is scope for implementing an RDD analysis over the age horizon. A fuzzy design is more appropriate in our setting due to possible residual effects caused by the treatment. We use the dummy variable taking 1 for the 40-45 age group and 0 for the 46-51 age group as the treatment variable. Age of individuals is the running variable and we implement the fuzzy RDD procedure described by Imbens and Lemieux (2008) and Lee and Lemieux (2010). As the average wage for men falls and tends to be volatile after age 51, we set an analysis window consisting of 12 age categories. Specifically, we compare the individuals in the 40-45 age group to those in the 46-51 age group. We prefer specifications with the first and second order polynomials as the risk of overfitting increases at higher orders. The estimates are presented in Table 8. In our RDD estimates (columns 2 and 3), the log hourly wage is not standardized to experience as we already use polynomial age variables in the estimation. We include region of residence and urban/rural dummies as control variables in all regressions. Although the precision of the estimates somewhat declines relative to the IV-2SLS estimates, magnitudes of the RDD estimates are similar to our baseline findings. Overall, our findings are robust to using a fuzzy RDD design.

Table 8 Fuzzy RDD estimates of the returns to college

	Dependent variable: Log hourly wage		
	[1]	[2]	[3]
Treatment	0.5716***	0.5629**	0.6492*
	(0.0935)	(0.2425)	(0.3777)
F-statistic (discontinuity)	140.13	18.63	8.24
Region of residence	Yes	Yes	Yes
Urban/rural status	Yes	Yes	Yes
# of observations	6,309	6,309	6,309
Polynomial degree	–	linear	quadratic

The sample includes male wage earners aged 40-51. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

16 We calculate the average years of schooling for the post-secondary education as 4 years based on the 2008 TDHS.

Correcting for missing wage data. The wage data in the HLFS contains missing values. About 2 percent of male individuals of age 40-51 did not report their wages, either because they started their current job within the survey month or did not want to disclose. We define a binary variable for male wage earners taking 1 if the wage data is missing and 0 otherwise. We regress this variable on age cohorts, 6 education dummies, 27 sub-major divisions of occupations, 26 NUTS2level region-of-residence dummies, and an urban/rural dummy. We find that some variables are statistically significant, and thus the missing values may not be random.

Although the fraction of missing data in our sample is low and unlikely to affect our main results, we still re-estimate the sampling weights to adjust for missing wage values following the procedure introduced by Altonji et al. (2012) to check whether the results are robust to missing data. The comparison is presented in Table 9, which says that our estimates are robust to adjusting sample weights for missing data on wages.

Using alternative birth cohorts to define treatment and control groups. Table 10 presents the estimates for alternative definitions of treatment/control groups and compares those estimates with the baseline findings. Age 46 in our control group might also be partly affected from the turmoil. We drop this cohort and also narrow the age windows by dropping an additional cohort in each specification (41-45 versus 47-51; 42-45 versus 47-50; 43-45 versus 47-49). We include region of residence and urban/rural dummies as control variables in all regressions. Estimates for the returns to college are highly significant and similar to our baseline estimates in terms of magnitude. Overall, our findings are robust to using alternative birth cohort specifications in defining the treatment and control groups.

Table 9 OLS and 2SLS estimates of the returns to college with adjusted missing wages

	Dependent variable: Log hourly wage	
	With missing values	**Adjusted for missing values**
Panel A: Just-identified models		
OLS	0.5025***	0.5037***
	(0.0311)	(0.0310)
2SLS	0.5716***	0.5744***
	(0.0935)	(0.0935)
F-statistic (first stage)	140.13	139.44
Panel B: Over-identified models		
2SLS	0.5588***	0.5601***
	(0.0841)	(0.0837)
LIML	0.5605***	0.5618***
	(0.0865)	(0.0861)
F-statistic (first stage)	25.15	25.04
p-value Hansen's test	0.80	0.80
Region of residence	Yes	Yes
Urban/rural status	Yes	Yes
# of observations	6,309	6,309

The sample includes male wage earners aged 40-51. In just-identified models, a dummy for individuals of age 40-45 is used as instrument. In over-identified models, six age dummies in the 40-45 age group are used as instruments. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

Table 10 OLS and 2SLS estimates of the returns to college with different samples

	Dependent variable: Log hourly wage			
	40-45	**41-45**	**42-45**	**43-45**
	46-51	**47-51**	**47-50**	**47-49**
	[1]	**[2]**	**[3]**	**[4]**
Panel A: Just-identified models				
OLS	0.5025***	0.5145***	0.5152***	0.5105***
	(0.0311)	(0.0333)	(0.0367)	(0.0342)
2SLS	0.5716***	0.5148***	0.4698***	0.4855***
	(0.0935)	(0.0690)	(0.0802)	(0.1016)
F-statistic (first stage)	140.13	181.59	144.01	94.10
Panel B: Over-identified models				
2SLS	0.5588***	0.5078***	0.4774***	0.5039***
	(0.0841)	(0.0716)	(0.0827)	(0.1032)
LIML	0.5605***	0.5076***	0.4767***	0.5037***
	(0.0865)	(0.0731)	(0.0840)	(0.1053)
F-statistic (first stage)	25.15	37.43	36.65	31.97
p-value Hansen's test	0.80	0.72	0.71	0.61
Region of residence	Yes	Yes	Yes	Yes
Urban/rural status	Yes	Yes	Yes	Yes
# of observations	6,309	5,042	4,093	3,088

The sample includes male wage earners of age 40-51. In just-identified models, a dummy for individuals of age 40-45 is used as instrument. In over-identified models, six age dummies in the 40-45 age group are used as instruments. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

Including more HLFS waves. As a final robustness check, we provide estimates based on four HLFS waves (2005-2008) rather than focusing only on the 2005 wave. From 2009, Turkey has gradually changed the HLFS's sampling addresses with a new national address database. In addition, as we go beyond the 2008 wave, individuals in our treatment and control groups get older and drop out of the core group in the labor market; thus, they start exhibiting non-standard labor market behavior. We focus on the 2005-2008 waves due to these two considerations. As we discuss in the data section, the average wage for men sharply falls and tends to be quite volatile after age 51; therefore, we drop the oldest cohort when we include a new wave. For example, treatment group is the 41-46 age group and control group is the 47-51 age group in the 2006 wave; thus, age 52 in the control group is dropped. Wages for different waves are adjusted to the 2005 prices using the CPI. The estimates are presented in Table 11. We include region of residence and urban/rural dummies as control variables in all regressions. Including more observations/waves allow for much higher precision in the estimates. Moreover, the estimates are similar to our baseline findings.

Clearly, our results are also robust to the inclusion of more HLFS waves.

5.3 Counterfactual wage distribution and occupational shift

We implement the semi-parametric procedure developed by DiNardo et al. (1996) to analyze the impact of the political turmoil on the wage distribution of men born between 1960-1965,

Table 11 OLS and 2SLS estimates of the returns to college using more HLFS waves

	Dependent variable: Log hourly wage	
	2005	**2005-2008**
Panel A: Just-identified models		
OLS	0.5025***	0.5380***
	(0.0311)	(0.0305)
2SLS	0.5716***	0.5644***
	(0.0935)	(0.0801)
F-statistic (first stage)	140.13	301.03
Panel B: Over-identified models		
2SLS	0.5588***	0.5950***
	(0.0841)	(0.0883)
LIML	0.5605***	0.5957***
	(0.0865)	(0.0895)
F-statistic (first stage)	25.15	55.62
p-value Hansen's test	0.80	0.85
Region of residence	Yes	Yes
Urban/rural status	Yes	Yes
# of observations	6,309	21,717

The sample includes male wage earners of age 40-51. In just-identified models, a dummy for individuals of age 40-45 is used as instrument. In over-identified models, six age dummies in the 40-45 age group are used as instruments. As the vector of covariates, 26 NUTS2-level region-of-residence dummies and an urban/rural dummy are used. Additionally, survey-year dummies are used in the regressions using multiple HLFS waves. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at region level, are reported in parentheses.

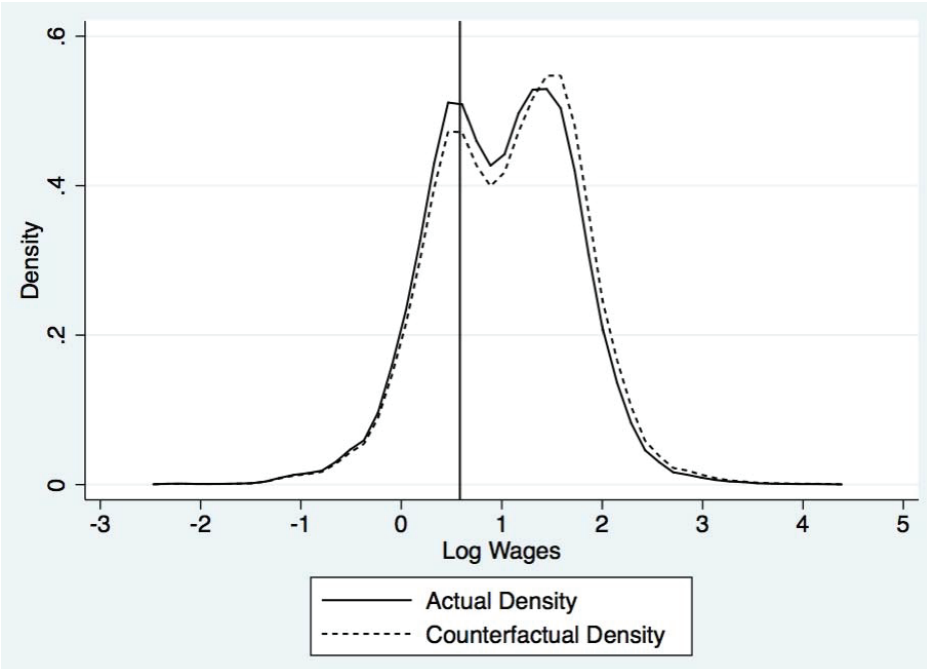
*** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

i.e., the exposed individuals (the 40-45 sample of males). To be consistent with our baseline analysis, the 2005 wave of the HLFS is used in this exercise. We re-weight the 40-45 sample to have the same distribution of post-secondary education as the 46-51 sample. We then compare how labor income is distributed in the re-weighted (counterfactual) 40-45 sample versus the actual 40-45 sample. This comparison roughly demonstrates how the decline in post-secondary educational attainment due to the turmoil affected the density of wages in the treatment group. A formal description of how we formulate and implement the density estimation are presented in Appendix A.

We run a probit model, as in DiNardo et al. (1996), to estimate the re-weighting function and plot the weighted kernel density estimates of the counterfactual (dotted line in Figure 14) and the actual (solid line in Figure 14) densities. We use log hourly wages standardized to 26 years of labor market experience for men—as in the previous sections. Both lines are superimposed in Figure 14.[17] The vertical line indicates the log minimum wage in 2005. It is calculated as the net monthly minimum wage (350 Turkish liras for 45 hours per week) divided by (52/12) and then by 45. Clearly, the minimum wage in Turkey compresses the lower tail of the density of the male wage earners. So, the distribution is twin-peaked, with the first peak settling

17 The Stata optimal bandwidth and Gaussian kernel function are chosen; but, note that the results are not sensitive to the choice of bandwidth and alternative kernel functions.

Figure 14 Actual and counterfactual density of log wages for male individuals in the 40-45 age group.



The sample includes male wage earners of age 40-51. Observations are weighted using the sampling weights so that the results are nationally representative.

around the minimum wage and the second peak appearing around 1.45 log wage value. The mean of the entire sample is approximately 1.01 in log terms.

The difference between actual and counterfactual densities represents the effect of the decline in post-secondary educational attainment—due to the political turmoil—on the distribution of wages for the 40-45 age group. Strikingly, the decline in post-secondary education pushed these individuals from the higher-income group toward the minimum-wage group. Those individuals who would have otherwise completed a post-secondary degree would have earned much more than their actual earnings if the turmoil had never occurred.

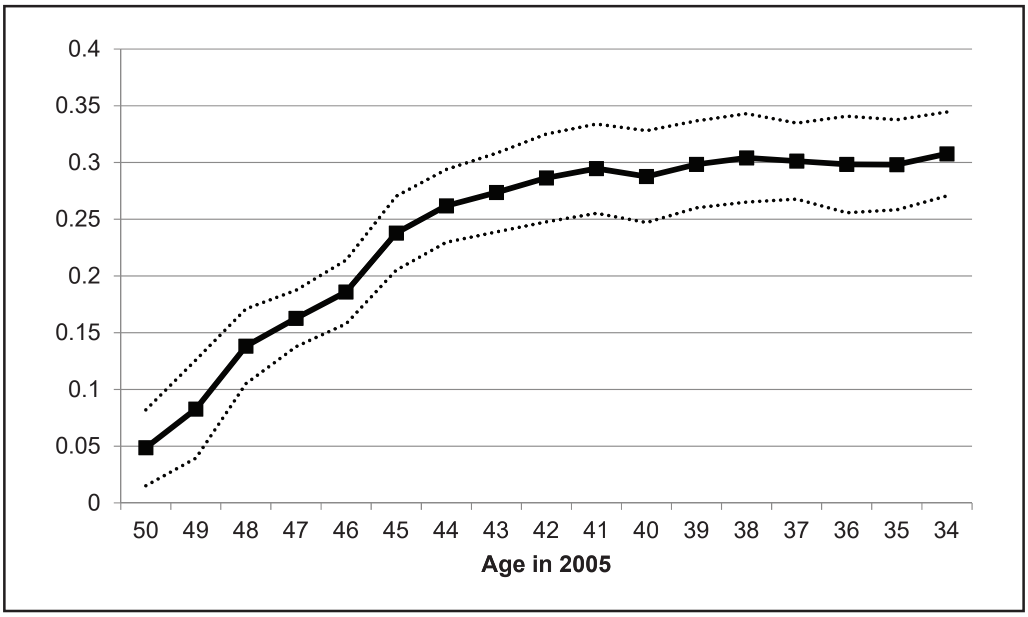
Next we explore the impact of the turmoil on the occupational structure. Before this analysis, we address the following question: does the turmoil affect other labor market outcomes such as labor force participation, employment, and labor informality? To answer this question, we run three regressions based on the following simple model:

$$s_i = \alpha + \sum_{c=34}^{50} \beta_c d_{ic} + \boldsymbol{X}_i' \boldsymbol{\Pi} + \epsilon_i, \tag{5.5}$$

where s_i is a binary variable for labor market status, d_{ic} is a dummy variable indicating whether individual i is c years old, $\boldsymbol{X}_i$ is a vector of covariates, and ϵ_i is an error term. The HLFS data used in this analysis is restricted to men of age 34-51. In all regressions, we use region-of-residence dummies and an urban/rural dummy as the vector of covariates. Age 51 is the omitted age category.

The first regression is for labor force participation. The estimated coefficients of ages are plotted in Figure 15, which shows that the trend for labor force participation is smooth over the

Figure 15 Coefficients of age dummies – Probability of being in the labor force.



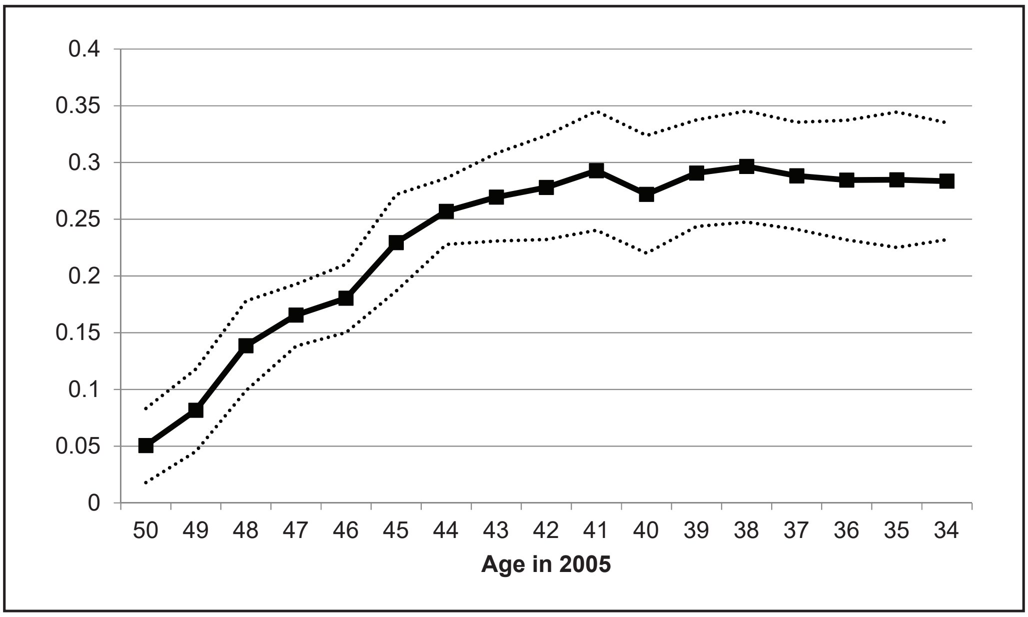
The specification includes region of residence and urban/rural dummies. Age 51 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

age horizon. Thus, the decline in post-secondary educational attainment did not affect labor force participation for men. The second and third regressions are for employment and formal employment (defined as being registered with the social security institution in the current job), respectively. The results are reported in Figures 16 and 17. Both figures point out that the decline in post-secondary educational attainment did not affect employment and formal employment. We conclude that the turmoil did not have any statistically significant effect on the main employment outcomes.

We show, however, that the turmoil largely affected the occupational structure in the labor market—as hinted by the counterfactual shift in wages. Our data set contains 27 sub-major divisions of occupations and they are classified according to International Standard Classification of Occupations (ISCO-88). We calculate the mean of log wage for each occupation for men and, then, we separately find the share of individuals in each occupation for three age groups (34-39, 40-45, and 46-51). We classify the occupations based on average log wage values and, accordingly, we construct five broad occupation groups. The first two groups can easily be defined, because they contain similar sub-major divisions. The occupations in the last three groups are in different majors. Thus, we classify them based on mean log wage values. The results are presented in Table 12.

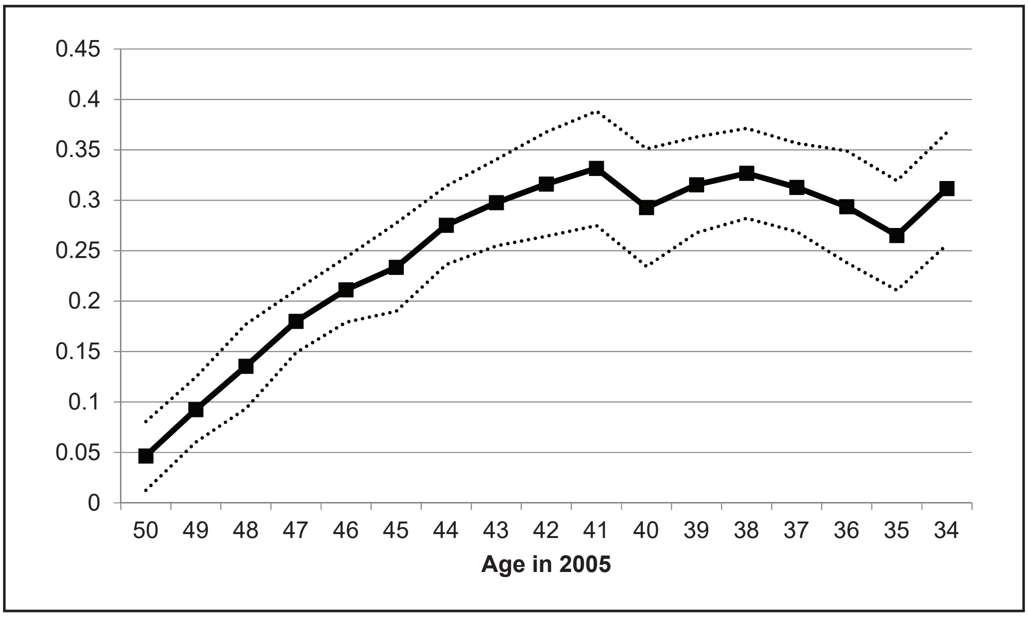
The top group is corporate managers and professionals and their mean log wage value is over 1.5—approximately corresponding to the second peak of the distribution of wages in Figure 14. 12 percent of those in the 40-45 age group are within this category. This ratio is much less than the other two age groups—being nearly 7 percentage points lower than the 46-51 age group. This difference is consistent with the difference between actual and counterfactual wage density estimations. We observe that individuals in the 40-45 age group—i.e., the exposed ones—have less attractive jobs on average.

Figure 16 Coefficients of age dummies – Probability of being employed.



The specification includes region of residence and urban/rural dummies. Age 51 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

Figure 17 Coefficients of age dummies – Probability of being formally employed.



The specification includes region of residence and urban/rural dummies. Age 51 is the omitted age category. Observations are weighted using the sampling weights so that the results are nationally representative. Dashed lines indicate the 95-percent confidence interval based on standard errors clustered at region level.

The second occupation group consists of technicians, associate professionals, and clerks. These occupations have less education requirements than the top group. The percentage of this group in the 40-45 age group is higher than the other two age groups and this result is also consistent with the counterfactual density estimation. The third and fourth groups have also confirmed the same result—as the percentages for the 40-45 sample are higher. If we combine

Table 12 Classification of occupations and their percentages in age groups

		Percentage in age group		
ISCO-88 codes	**Classification**	**34-39**	**40-45**	**46-51**
12, 21, 22, 23, 24	Corp. managers and professionals (1.53<log wage<2.07)	13.73	12.00	18.68
31, 32, 33, 34, 41, 42	Technicians, assoc. professionals & clerks (1.29<log wage<1.45)	15.12	17.39	16.66
11, 13, 51, 72, 81	Average wage earners (0.92<log wage<1.10)	20.32	20.72	18.40
71, 73, 82, 83, 91	Between min. wage & av. wage (0.69<log wage<0.87)	33.29	34.37	31.28
52, 61, 62, 74, 92, 93	Approx. less than min. wage (log wage<0.61)	17.54	15.52	14.98

Log (hourly) wages in this table are the means in the corresponding broadly-defined occupation group.

the sub-major divisions whose mean log wage values between the minimum wage and 1.5 [the second peak in Figure 14], the fractions of individuals in age groups 34-39, 40-45, and 46-51 become 68.7, 72.5, and 66.3, respectively. This suggests that the 40-45 age group have a higher fraction of low-pay occupations. Therefore, we conclude that the decline in post-secondary attainment led to a shift in occupations from high-pay to low-pay ones. These findings suggest that violent political turmoils or other large-scale turbulent events affecting large fractions of the productive population may have permanent negative consequences on countries' growth potentials.

6 Concluding remarks

Between 1978 and 1980, Turkey experienced violent student protests. Almost 20 young individuals were killed daily. Universities and other higher education institutes were at the center of the violence and conflict. This violence ultimately led to a military intervention in 1980, which came with additional arrests and suppression lasting several years. We document in detail that the political turmoil adversely affected post-secondary educational attainment in Turkey. Furthermore, we exploit the exogenous drop in the number of graduates and new admissions due to this turmoil (*i*) to estimate the causal effect of college education on earnings using birth cohorts as IV and (*ii*) to document its long-term impact on wages and occupation structure.

We find that the group most severely affected by the turmoil is male wage earners with birth years from 1960 to 1965. These events led to a substantial decline in the probability of completing postsecondary education for the exposed cohorts. The decline in post-secondary educational attainment pushed the wage distribution to the left and led to a permanent shift in occupations from high-pay to low-pay ones. These findings also suggest that violent political turmoils can permanently erode a country's human capital and, therefore, can adversely affect long-term growth prospects.

Using birth cohorts as IV, we estimate that the return to an additional year of college education is approximately 15 percent. This is one of the first papers using the widespread student protests—almost globally took place in the second half of the twentieth century—as a source of exogenous change in college attainment to estimate returns to higher education. The long duration of the turmoil and high violence intensity jointly make the Turkish case a unique example of student protests with severe adverse effects on post-secondary education and long-term labor market outcomes.

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Appendix A

Description of Counterfactual Density Estimations

The actual and counterfactual density estimates are obtained by the kernel density estimator proposed by DiNardo et al. (1996). Let $W_1,\ldots,W_n$ be a random sample of size n, with weights $\theta_1,\ldots,\theta_n$ drawn from some distribution with an unknown density f. Its kernel density estimator is defined as follows:

$$\hat{f}_h(w) = \sum_{i=1}^{n} \frac{\theta_i}{qh} K\left(\frac{w - W_i}{h}\right), \tag{A.1}$$

where $q = \Sigma_{i=1}^{n} \theta_i$, h is bandwidth, and $K(\cdot)$ is the kernel function. We choose the HLFS analytic sampling weights in the estimation, because weights are rescaled so that $\Sigma_{i=1}^{n} \theta_i = n$.[18]

We adopt the notations and notions introduced in the original paper by DiNardo et al. (1996). Each individual observation belongs to a joint distribution $F(w,d,z)$; where w is wage, d corresponds to individual-level attributes, and z is a time variable. The joint distribution of wages and individual attributes at a point in time is the conditional distribution $F(w,d|z)$. In that case, the density of wages at a point in time, $f_z(w)$, can be defined as the integral of the density of wages conditional on individual attributes and a time z_w, $f(w|d,z_w)$, over the distribution of the individual attributes at time z_d, $F(d|z_d)$, as follows;

$$\begin{aligned} f_z(w) = \int_{d\in\Omega_d} dF(w,d|z_{w,d} = z) &= \int_{d\in\Omega_d} f(w|d,z_w = z)\,dF(d|z_d = z) \\ &= f(w;\, z_w = z,\, z_d = z), \end{aligned} \tag{A.2}$$

where Ω_d is the domain of individual-level attributes. To be consistent with our notation, z is a binary variable taking 1 for the 40-45 sample and 0 for the 46-51 sample. Thus, the expression $f(w;z_w = 1,z_d = 1)$ represents the actual density of wages in the 40-45 sample, whereas $f(w;z_w = 1,z_d = 0)$ represents the counterfactual density of wages in the 40-45 sample, if the characteristics of these workers are the same as the 46-51 sample without changing the wage schedule observed for the 40-45 sample.[19]

Under the assumption that conditional density $f(w|d,z_w = 1)$ does not depend on the distribution of attributes, the counterfactual density $f(w;z_w = 1,z_d = 0)$ can be written as

$$\begin{aligned} f(w;\, z_w = 1,\, z_d = 0) &= \int_{d\in\Omega_d} f(w|d,\, z_w = 1)\,dF(d|z_d = 0) \\ &= \int_{d\in\Omega_d} f(w|d,\, z_w = 1)\,\psi_d(d)\,dF(d|z_d = 1), \end{aligned} \tag{A.3}$$

where the re-weighting function is $\psi_d(d) = dF(d|z_d = 0)/dF(d|z_d = 1)$. As in Equation A.3, the counterfactual density is obtained by re-weighting the actual density. The conditional density of wages may depend on the distribution of attributes due to non-random selection. Therefore, we assume that the distribution of the unobserved attributes conditional on the observed attribute d is the same for the two groups, which means that the difference between the cohorts in

18 This ensures that the software version (Stata) of kernel density estimation is compatible with the estimator proposed by DiNardo et al. (1996).

19 The general equilibrium effects are ignored.

the distribution of d can account for any difference between the cohorts in the marginal distribution of the vector of unobserved skills (Altonji et al., 2012).

After estimating $\hat{\psi}_d(d)$, the counterfactual density is estimated by the weighted kernel method as follows:

$$\hat{f}(w;\ z_w = 1,\ z_d = 0) = \sum_{i \in I_1} \frac{\theta_i}{qh} \hat{\psi}_d(d) K\left(\frac{w - W_i}{h}\right), \tag{A.4}$$

where I_1 is the set of indices for individuals of age 40-45. We want to estimate the effects of the decline in post-secondary educational attainment; so, we are interested in whether the individual finishes post-secondary education or not. The difference between the actual and counterfactual densities indicates the effect of a decline in post-secondary educational attainment on the distribution of wages within the treatment group.

The re-weighting function $\psi_d(d) = dF(d|z_d = 0)/dF(d|z_d = 1)$, by applying the Bayes' rule, can be rewritten as follows:

$$\psi_d(d) = \frac{\mathbb{P}\left[z_d = 0|d\right]}{\mathbb{P}\left[z_d = 1|d\right]} \cdot \frac{\mathbb{P}\left[z_d = 1\right]}{\mathbb{P}\left[z_d = 0\right]}. \tag{A.5}$$