Long-Term Education and Labor Market Consequences of the Student Protests in 1970s and the Subsequent Political Turmoil in Turkey*

Ahmet Ozturk †
Ministry of Development

Semih Tumen ‡
TED University and *IZA*

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Abstract

1970s have witnessed widespread and highly-politicized student protests in Turkey. Small protests turned into violent street clashes, the death toll exceeded 5,000, and a military coup came in—which resulted in mass arrests. Universities were at the center of the turmoil and violence. Many people hesitated to continue education and, if continued, had long interruptions in education. We present a comprehensive empirical analysis of the long-term labor market consequences of this political turmoil on cohorts directly 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 and the subsequent military coup. Second, we estimate a counterfactual wage distribution for the exposed cohorts using semi-parametric methods and, then, check whether the turmoil affected their occupational structure in the long term. Finally, we use the unexpected decline in educational attainment as an instrumental variable to estimate the returns to education parameter. We find that the decline in educational attainment due to the political turmoil negatively affected the productive human capital capacity of the country by shifting the affected population from high-income occupations toward low-income ones. Accordingly, wages were compressed toward the minimum wage. Our IV estimates suggest that returns to an additional year of schooling in Turkey range between 11.6–12.8 percent for men.

JEL codes: D74; J21; J31; I26.

Keywords: Student protests; military coup; labor markets; IV; returns to education; occupational change.

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†ahmet.ozturk@kalkinma.gov.tr. Necatibey Cad. No:110-A, 06100 Yucetepe, Ankara, Turkey.

‡semihtumen@gmail.com. Department of Economics, TED University, Ziya Gokalp Cad., No.48, 06420 Kolej, Ankara, Turkey.

1 Introduction

Turkey experienced violent student protests in late 1970s and a military coup in 1980, which significantly eroded post-secondary educational attainment for the exposed cohorts. From 1978 to 1980, an average of 20 youths were killed each day on Turkey's streets and university campuses. After the coup, students were regularly snatched up in mass arrests. In this paper, we characterize the long-term labor market consequences of these dramatic events—which we hereafter refer to as the "1978–1982 turmoil." We provide a detailed documentation of the impact of the turmoil on post-secondary education. Then, we estimate a counterfactual wage distribution (as if the turmoil did not take place) and illustrate the change in occupational structure for the exposed cohorts. Finally, we use the unexpected decline in educational attainment during this episode as an instrumental variable (IV) to estimate the causal effect of education on earnings.

In 1960s and early 1970s, student movements concurrently grew in much of the world and became a global phenomenon.[1] In the United States and Europe, these movements had declined significantly by mid-1970s [Barker (2012)], which was the time they just started to escalate in Turkey [Ahmad (1993)]. The late 1970s had been one of the darkest periods in Turkey's modern history, culminated in the 1980 military coup [Ahmad (1993)]. During the 1978–1982 turmoil, student protests between youth groups on the political right and left turned extremely violent in Turkey [Zurcher (2004)].

The intensive violence adversely affected post-secondary educational attainment in Turkey in several ways. First, new enrollments declined in the 1978–1979 school year, largely due to the closure of teacher-training institutes as a result of their links to student violence. Second, graduation rates declined following massive student dropouts related to security concerns. Finally, mass student arrests in the wake of the 1980 coup kept many from completing their education.

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

Using the 2005 Turkish Household Labor Force Survey (HLFS), we document that the 1978–1982 turmoil led to a 6.6–7 percentage point decline in the probability of completing post-secondary education, a 0.22–0.28 year decline in average schooling, and a 2.6–3.5 percent decrease in wages. What is more, this educational decline led to a shift from high-income to low-income jobs/occupations. We also find that male wage earners of age 40–45 in 2005—15–20 years old in 1980; born from 1960 to 1965—had been the most severely affected group during this period.

Because birth year is random and unrelated to innate ability and other relevant characteristics (at least within a sensibly narrow birth-year interval), it seems reasonable to assert that the sole reason for wage decline in this age group, after standardizing the experience across the labor market participants, is the decrease in post-secondary education due to the turmoil. Along this line of reasoning, we use birth year as an IV in the wage equation. Our IV estimates suggest that the returns to education in Turkey range between 11.6–12.8 percent for men. These numbers are somewhat above the corresponding ordinary least square (OLS) estimates. We conclude that the IV estimates may be closely approximating the average causal effect of an additional year of schooling in post-secondary education for several reasons. First, the instrument only affects post-secondary education. Second, the individuals whose schooling attainment is changed by the instrument (i.e., the compliant sub-population) are at least 31 percent of individuals in our sample. Third, those individuals affected by the 1978–1982 turmoil are not marginal individuals who were basically indifferent between going to college or not, because those affected were the dropouts in post-secondary education or would have at least a college degree if these events had never happened.

This paper makes three important contributions. First, we argue how a violent political turmoil can erode long-term human capital productivity of a country by examining its impact on wage distribution and occupational structure. Our findings clearly indicate that such political/social turmoils have large and persistent effects on wage and occupational distributions in a society. Second, we develop a new instrument that affects post-secondary education. To our knowledge, this is the first study constructing an instrument from the global student protests in the 1960s and 1970s to estimate returns to education. The final contribution of the paper is about the size of the returns to schooling in a developing country context. The two-stage least squares (2SLS) estimates of this study are close to most estimates found for the developed countries, but smaller than the typical estimates reported for developing economies.[2] Duflo (2001) and Aydemir and Kirdar (2017) also reach similar results in an IV framework. Thus, we argue that developing countries may not be experiencing higher returns to education than the developed countries.

The plan of the paper is as follows. Section 2 discusses the related literature. Section 3 provides detailed background information about student protests in late 1970s and the subsequent military coup in Turkey. Section 4 describes the data. Section 5 explains the empirical setup and identification strategy. Section 6 presents the findings and provides an in-depth discussion of the main implications. Section 7 concludes.

2 Related Literature

We organize our discussion of the related literature under three related but distinct topics. First, we discuss the literature on the timing of school enrollment and graduation on labor market outcomes—in the sense that when one graduates and/or enrolls in school matter for future labor market success. Second, we argue how our paper can be related to the papers in the literature investigating the impact of major historical events on long-term human capital formation and growth potential. Finally, we link our paper to the literature on IV estimates of returns to schooling. We discuss the implications of our returns to schooling estimates in light of the existing evidence on Turkey and other developing countries—touching upon the differences between the estimates for developed and developing countries.

The impact of the timing of school graduation on labor market outcomes is being discussed in a large and growing literature [see, e.g., Kahn (2010), Oreopoulos, von Wachter, and Heisz (2012), Johnson (2013), and Liu, Salvanes, and Sorensen (2016)]. The luck component inter-

[2] See, e.g., Behrman (1999) and Psacharopoulos and Patrinos (2004).

acts with the economic/political/social conditions evolving over time and shapes the long-term labor market consequences of a group of individuals. There are also other cases based on abrupt changes—such as war, civil conflict, natural disasters, etc.—affecting long-term labor market prospects of the vulnerable ones. Studies including Kondylis (2010), Shemyakina (2011), Swee (2015), and Tumen (2016) address this issue along various dimensions, contexts, and settings.[3] Moreover, political instability and turmoil are shown to have not only labor market consequences, but other long-term implications for the macroeconomic performance of countries.[4] While our paper can be classified among these papers, it is the first paper in the literature focusing on the long-term labor market effects of student protests, which were widespread in a large set of countries during 1960s and 1970s. The Turkish protests generated relatively more severe and dramatic outcomes than those in other countries because of the high level of violence and the military coup following the protests. Hence, the turmoil affected directly the post-secondary school attainment, which magnifies the impact of conflict on long-term labor market outcomes and makes the political turmoil a potentially viable IV.

Our paper can also be related to the literature investigating the impact of major historical events on long-term economic outcomes. For example, there is a growing body of literature on the long-term consequences of Holocaust.[5] Similarly, the long-term effects of the World War II and the Vietnam War have also been studied and documented extensively.[6] The endogenous growth literature confirms that major historical events shaping the dynamics of human capital formation and accumulation affect long-term growth prospects of countries.[7] Our paper also suggests that the violent students protests in late 1970s, which are followed by a military coup in 1980, significantly reduced post-secondary school attainment in Turkey and, thus, inhibited growth potential in Turkey.

We exploit the sharp decline in post-secondary school attainment due to the turmoil in an IV

[3] See Blattman and Miguel (2010) for a comprehensive review of the relevant literature.

[4] See, e.g., Alesina, Ozler, Roubini, and Swagel (1996), Collier (1999), Abadie and Gardeazabal (2003), and Bohlken and Sergenti (2010).

[5] See, e.g., Waldinger (2010, 2012, 2016), Acemoglu, Hassan, and Robinson (2011), Grosfeld, Rodnyansky, and Zhuravskaya (2013), Akbulut-Yuksel and Yuksel (2015), and Pascali (2016).

[6] See, for example, Davis and Weinstein (2002), Brakman, Garretsen, and Schramm (2004), Ichino and Winter-Ebmer (2004), Miguel and Roland (2011), and Akbulut-Yuksel (2014).

[7] See, e.g., Klenow and Rodriguez-Clare (2005).

setting to estimate returns to education. There is a rather small set of commonly used IVs in the literature to estimate returns to schooling in various settings. For example, institutional features of schooling systems have often been used to obtain IV estimates of returns to education. The landmark study by Angrist and Krueger (1991) uses the interaction between compulsory school attendance rules and the quarter of birth as an instrument. Their 2SLS estimates yield around 7.5 percent returns to an additional year of schooling for men, which is close to their OLS estimates—suggesting that OLS may have little bias.[8] Acemoglu and Angrist (2000) use the same idea to estimate human-capital externalities.[9] Another highly referenced article exploiting institutional features of education system is Oreopoulos (2006). He uses an indicator about whether a cohort faced a school leaving age of 15 at age 14 in the United Kingdom (UK). Oreopoulos (2006) finds that the 2SLS returns to schooling for men range from 7–10 percent in the UK.[10] Other instruments used by researchers in estimating the returns to schooling include: change in the duration of secondary education; tuition at 2–year versus 4–year colleges; distance to nearest high school or college; and living in a university town.[11] In addition to the features of the school system, family background such as parents' and twin's education are also used as instruments in studies of returns to education.[12] We contribute 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 education.

Even though a large body of literature investigates the endogeneity of education in the estimation of the returns to schooling in developed countries, relatively few studies deal with

[8] Yet this landmark study has faced criticism. Bound, Jaeger, and Baker (1995) indicates that the IV estimates in Angrist and Krueger (1991) paper may suffer from finite-sample bias and may be inconsistent because several of their models include weakly correlated instruments. Bound and Jaeger (1996) also criticize the same study. They point out that the quarter of birth may be correlated with some unobserved variables such as family background. See Torun and Tumen (2017) for an example in which the season of birth may not be a good IV.

[9] Other recent examples exploiting these laws to estimate causal effects of education on various outcomes include Bell, Costa, and Machin (2016) and Sansani (2015).

[10] In contrast, Devereux and Hart (2010) re-analyze the same data set and find smaller 2SLS returns to schooling (of about 4–7 percent) for men in the same context.

[11] See Card (1999, 2001) for a comprehensive list of relevant references. The instruments based on the geographic location of individuals of college- or high school-going age cannot be valid if the choice of going to school and the location decision are correlated [Heckman, Lochner, and Todd (2006)]. Families may choose to locate in areas based on proximity to schools. Average tuition may also be invalid because Carneiro and Heckman (2002) show that the average college quality is correlated with the average tuition in the US.

[12] For these instruments, it is crucial to presume that potential wages in college and high school regions are independent of family characteristics, but many studies show that these are among the main determinants of ability [Heckman, Lochner, and Todd (2006)]. Thus, these instruments are controversial unless the ability is somehow included in the regressions.

this issue in developing countries. The seminal paper by Duflo (2001) is one. Duflo (2001) investigates a dramatic change in education policy that launched a major primary school construction program to target children who had not previously been enrolled in Indonesia. Duflo (2001) reports moderate economic returns to education, ranging between 6.8–10.6 percent for men. These estimates are close to other estimates found for developed countries. Fang, Eggleston, Rizzo, Rozelle, and Zeckhauser (2012) is another study in a developing country context. They construct an IV by exploiting the China Compulsory Education Law of 1986. Their 2SLS estimates for the returns to schooling for men are about 51 percent—more than five times the corresponding OLS estimate (9 percent). This finding could be controversial due to the large difference between the OLS and 2SLS results. Torun and Tumen (2016) use compulsory military service exemption (which reduces incentives to stay in college for the purpose of deferring military service) in Turkey as an IV and also report high returns to higher education for men.

With a few exceptions, research on evaluating the impact of education on earnings using an IV strategy (or other quasi-experimental designs) remains limited in Turkey. Torun (2015) and Aydemir and Kirdar (2017) exploit Turkey's compulsory schooling law of 1997 in their studies. The law introduced a continuous uninterrupted eight-year education in the same school building. Both studies use an indicator of whether birth cohorts are affected by the policy as an instrument. Torun (2015) and Aydemir and Kirdar (2017) find low returns to schooling estimates, about 2–3 percent for men, mainly because the 1997 law changes schooling distribution at the elementary school level (grades 6 through 8). Recent studies in Turkey [Cesur and Mocan (2017); Gulesci and Meyersson (2013); Torun (2015); Aydemir and Kirdar (2017)] have used the compulsory schooling law of 1997 as an instrument to investigate the causal relationship between education and economic or social outcomes. Individuals who are induced to change their behavior because of the 1997 law are still young even in the most recent data sets—of age 18 and 26. Some of these individuals may still be in college or some of them may have only recently completed their high school or college education. In our analysis, those individuals who are affected fall within the range of core age group in the labor market—40 to 45 years old. Hence, the instrument that we propose can bring a new perspective in exploring the causal relationship between education and earnings—and also other socio-economic outcomes often explored in the literature such as health, crime, religiosity, and voting preferences. Moreover, since the students protests were common in many countries in 1960s and 1970s, a similar IV approach may be implemented also for other countries depending on the context.

3 Some background: Student protests and the subsequent coup

3.1 Emergence of civil conflict in Turkey from 1960 to 1980

The army has always played an outsized role in Turkish politics, ousting elected governments nearly every decade from 1960 to 1980. The 1960 coup marked the beginning of a new phase in Turkey. Ahmad (1993, 2003) and Zurcher (2004) emphasizes that junior officers carried out this intervention against higher officials and it was Turkey's only successful military coup from outside of the army's hierarchical structure. After the military intervention, a new constitution was prepared before the free election 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; workers had the right to strike. Turkey's new freedom enabled something unprecedented: ideological politics.

Left-wing politics started to emerge, especially on university campuses. Trade unionists founded the Workers' Party of Turkey. Zurcher (2004) argues that it forced the other parties to define themselves in ideological terms. In contrast, the right was alarmed by this leftist presence, and began to organize against it. Accordingly, Turkey's nationalist movement started to grow rapidly in 1969, with the creation of the Nationalist Movement Party [Erken (2014b)].

With a push from the global events of 1968, Turkey's left became more extremist in the hopes of igniting a revolution. But the left's violence was soon met and surpassed by violence from the extremist right [Zurcher (2004)]. This violence created the political instability that laid the groundwork for the coup. In March 1971, the army forced the elected government to step down and changed the constitution. Ahmad (1993, 2003) emphasize that they amended the constitution to strengthen the state against civil society; gained control of the universities to curb radicalism; and pacified trade unions after the dissolution of the Workers' Party. The left soon rallied around the Republican People's Party, which had shifted left in the mid-1960s.

In 1973, the Republican People's Party won parliamentary elections and formed a coalition government. Right-wing parties criticized the government program that sought to heal the wounds left by the military regime. The formation of the coalition coincided with an uptick in right-wing extremist violence. According to Ahmad (1993, 2003), the aim of rightist violence was to decrease the left's potential by eroding support and causing chaos to create a climate for military intervention. Radical leftists responded with acts of violence to further increase instability. After coalition formation, political violence became a regular feature of daily Turkish life, escalating and becoming more intense in the late 1970s.

Figure (1) presents the total terrorist attacks used as a proxy for the civil conflict in Turkey from 1970 to 1985—data source is the Global Terrorism Database (GTD). According to the code book of this database, "the GTD 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." The data shows that attacks declined after the 1971 intervention, but they increased again after 1974 and were most intense during the turmoil leading up to the 1980 coup.
On 5 April 1977, the two main parties agreed on an early election, sparking more intense political violence. The street terror peaked on May Day (May 1st) 1977, four weeks before the election. The Confederation of Revolutionary Workers' Union organized a huge rally in Istanbul. Shots fired into the crowd killed 36 people and injured hundreds.

The 1977 election did not produce a strong and stable government because no party won a majority. As a result, Turkey experienced one of its darkest periods in terms of political instability and societal chaos. By July 1978, the government started to use the army to secure the country. Despite the increasing use of force, the violence continued until the slaughter reached 20 victims a day in the late 1970s [Ahmad (1993); Kaya (1981)]. From 1978 to 1980, 5,241 people were killed and 14,152 people wounded due to the political violence [Kaya (1981)].

The army took control in September 1980 and ruled until the general election 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. After a year, the number was 122,600. 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 The effect of the 1978–82 turmoil on post-secondary education

Student protests in Turkey increased with a push from the global events of 1968. But Turkey's protests soon mutated into violence, and the incidence of these acts increased in the late 1970s. University students were divided into two opposed groups, "rightists" and "leftists," and built their identities in opposition to each other [Neyzi (2001)]. Educated youth saw themselves as the moving force of society and their main mission was to modernize society [Neyzi (2001); Zurcher (2004)]. Youth violence played a key role in creating the political instability that led to military interventions both in 1971 and in 1980 [Ahmad (2003)].

In the early 1970s, extreme leftist students emulating Latin American left-wing radicals robbed banks, kidnapped American soldiers and other prominent corporate figures [Ahmad (1993, 2003)]. From the military intervention of 1971 through 1973, the student activism went into a period of silence because the activists were either under arrest or executed [Erken (2014a)]. Figure (1) shows that terrorist attacks almost disappeared in this period. Student protests started to escalate again after 1974, along with terrorist attacks.

The intense violence seen during the late 1970s adversely affected post-secondary educational attainment in Turkey through several channels. First, new enrollments in post-secondary education declined in the 1978–79 school year, largely due to the closure of teacher-training institutes driven by their links to student violence [see Figure (2)]. Based on official (TURKSTAT) data, the decline in the number of students enrolled is 37,715.

Civil conflicts and student movements caused deep polarization in Turkey's higher education institutions. Although 11 new universities were established between 1970 and 1980 in major cities, university boards did not increase enrollment capacity enough to meet the demand for higher education. In this period, the higher education system was decentralized and there was no governing authority for higher education institutions. According to Dogramaci (1989), the lack of coordination among higher education institutions made it impossible to address national priorities. The government built academies, vocational schools, and teacher-training institutes that were affiliated with certain ministries. Between 1973 and 1977, new enrollments rose by 42,570—15 percent of this rise occurred in universities, while 85 percent occurred in institutions affiliated with ministries.

In 1973, Turkey increased compulsory education from five to eight years, and thus urgently needed elementary school teachers. As a result, 76 percent of the enrollment increase mentioned above occurred in teacher training institutions. 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.

One prominent newspaper covered this issue closely: "administrators, teachers, and students in 35 teacher training institutions wanted to re-open the educational institutes. . . . in the joint statement, they argued that closure of educational institutions after attacks submit to the fascists" [Cumhuriyet (1978)]. These institutions were directly affiliated with the Ministry of Education and they were not protected by the constitution like the universities. According to Tekeli (2010), the student protests began to turn violent after 1968 and this violence escalated in the late 1970s. Authorities were thus focused on how to prevent student protests. Due to mainly from closing of these institutions, enrollment declined by 37,715—eroding about 90 percent of the enrollment increase of 1973–77.

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

The second channel that adversely affected educational attainment was that graduation rates declined following massive student dropouts related to security concerns. From 1978 to 1980, an average of 20 young people were killed per day; so many students canceled their registration in higher education institutions [Kaya (1981)]. Some students were unable to finish education because they were injured or disabled during the student violence. 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 1979–80 academic year [Kaptan (1986)]. Faculty offices and student dormitories were often turned into weapon warehouses and arsenals [Kaya (1981); Kaptan (1986)], because the law on autonomy gave allowed universities considerable immunity from police oversight [Gunter (1989)].

Finally, 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 2005 Turkish Household Labor Force Survey (HLFS) in this study.[13] The Turkish Statistical Institute (TURKSTAT) has provided the HLFS micro data in accordance with Eurostat's requirements since 2004. We focus on the 2005 wave, because the 2005 wave is the only reliable wave (as 2004 was a transition year) and after 2005 the impact of the 1997 compulsory schooling reform will start to affect educational attainment within the core working-age group (34 to 51 years old—which is our main sample. This group includes both the treatment and control groups for this study, which are explained in detail below.[14] Observe that the average wage for males is an increasing function of age [see Figure (3)]. But after age 51, it sharply falls and tends to be quite volatile. In addition, the number of wage observations declines significantly after age 51 (a similar pattern is also observed in the 2004 data). This is another reason why the 2005 HLFS is the only reliable wave for this study. The data set provides age, highest level of education completed, labor 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 payment like bonus payments and premiums), and main tasks/duties of individuals in workplace.

The data set does not have direct information on experience, so we use potential experience, as proposed by Mincer (1974): $exp = A - S - B$, where A is current age, S is the years of schooling, and B is the age at the beginning of schooling. In Turkey, age seven was approximately the beginning of schooling before 1980. Moreover, the data set does not directly reports the years of schooling; instead, it only has highest level of education successfully completed. However, Turkish Demographic and Health Surveys (TDHS) have the information on both graduation and years of schooling. Thus, we estimate 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 illiterates, 1.68 years for literates with no degrees, 5.09 years for primary school graduates, 8.34 years for elementary school graduates, 11.09 years for

[13] HLFS is a nationally-representative micro-level dataset compiled and published by TURKSTAT. It is used to calculate official employment statistics in Turkey.

[14] Note that the results of the regressions do not significantly change when we use the 2004 wave.

high school graduates, and 14.63 years for post-secondary school graduates. Based on this information, we use 0 years for illiterate people, 2 years for those who are literates with no degrees, 5 years for primary school graduates, 8 years for elementary school graduates, 11 years for high-school graduates, and 15 years for post-secondary graduates. In the TDHS estimates, we prefer to restrict the sample of individuals aged 37 to 54 years because our sample of the 2005 HLFS are aged 34 to 51 years.[15]

The data set includes monthly wages—the average wage is 588 Turkish liras in the 2005 wave. Card (1999) indicates 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, in Turkey there is a negative correlation between schooling and the number of hours worked [presented in 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 also -0.3. Therefore, in this paper, we choose hourly wages as the measure of income. We compute 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.

In all regressions, we standardize log hourly wages at 26 years of potential experience because our treatment and comparison groups have different experience levels. 26 years is the mean of potential experience of male wage earners for the age group 34–51. We estimate a log hourly wage equation separately for each educational status defined in the survey data for this group. These are no degree, primary (five-year), elementary (eight-year), high school, and post-secondary education graduates. 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. Altonji, Bharadwaj, and Lange (2012) use a similar strategy to standardize the potential experience. Lemieux (2006) also proposes the use of a quartic function in potential experience instead of a quadratic in the Mincer equation.

[15] A similar strategy is implemented by Aydemir and Kirdar (2017).

Table (2) provides descriptive statistics for individuals of age 34 to 51. For this group of individuals, 63 percent have a primary or elementary school diploma, 14 percent have a high school diploma, 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 Strategy

5.1 Visual Evidence for Instrument Validity

Effect on post-secondary educational attainment. Post-secondary education enrollment increased dramatically during the second half of the 20th century all over the world [Psacharopoulos (1991)]. Turkey is no exception. However, in Turkey, new enrollments and graduation rates in post-secondary education substantially declined between 1978 and 1982 due to the political turmoil. To find the trend in post-secondary educational attainment for Turkey using the 2005 wave of the LFS (plotting over age will suffice), 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 a 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 are included for the age range 30 to 55. As the vector of covariates we use NUTS2-level region-of-residence dummies, an urban/rural dummy, and gender. We drop the age dummy 55 in the regressions; hence, the coefficients of the age dummies should be read relative to age 55. This is a natural thing to do, because individuals of age 55 in the 2005 survey were 28 years old in 1978 and they were not exposed to violence and coup when they were at school. Therefore, each coefficient β_c can be interpreted as an estimate of the probability of completing post-secondary education for the corresponding age relative to age 55. In a developing country, it is normally expected that the coefficient of age dummies increases as age declines over the age spectrum in the absence of a negative shock on post-secondary education.

That means there should be an upward trend in post-secondary educational attainment.

Figure (4) plots β_c for the whole sample. Each dot on the solid line is the coefficient of the probability of completing post-secondary education, while the broken lines indicate the 95-percent confidence interval. The figure indicates that the coefficients increase as we move from age 54 to age 47. After age 47, there is a sharp drop, which levels off between age 45 and 40, and, then, begins to increase again.

Figure (5) and Figure (6) display the separately estimated coefficients of age dummies for male and female samples, respectively. Although the trends for completing post-secondary education look similar for men and women, the effects of the 1978–82 turmoil are more severe for men. These figures indicate that individuals of age 40–45 were the ones who are affected the worst from the turmoil. These individuals were born from 1960 to 1965, and were about 13 to 18 years old in 1978. Thus, we choose the individuals of age 40–45 in 2005 as our treatment group throughout the analysis.

The 40–45 age group has 6 age categories. We construct a control group that also has 6 age categories—those of age aged 46–51. Figure (4) shows that the probability of completing post-secondary education for the treatment group is clearly less than that for the control group. To further justify the treatment group, we run the same regressions for a narrower age spectrum by dropping age 48 (the middle age in the control group). Figures (7)–(9) show coefficients of age-dummies for whole, male, and female sample, respectively. These figures demonstrate that the probability of completing post-secondary education for individuals of age 40–45 relative to those of age 48 is negative (in a statistically significant way) for the entire sample and for males.[16] However, for females, statistical significance is less obvious. This confirms that the protests significantly affected the educational attainment of males. Similar results are also obtained if 50, 49, and 47 are taken as the reference age in the regressions.

Figure (2) already indicated that first-year enrollments in higher education declined signifi-

[16] The significance levels are 5 percent except that age 44 is at 10 percent in the whole sample; ages 44 and 41 are at 10 percent for males.

cantly for the first time in 1978 and remained low until 1982. This decline of enrollment in higher education institutions would probably affect 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 from Figure (4) to Figure (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 1978–82 turmoil.

Effects on different levels of educational attainment. To further justify the selection of the treatment group, 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. We use the same vector of covariates as in Equation (5.1). 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 that β is positive in the absence of a negative shock on education.

Table (3) presents three sets of estimates from Equation (5.2). Column [1] indicates the result for post-secondary educational attainment, column [2] for high school attainment (individual i has only high school diploma), and column [3] for elementary/primary school attainment (individual i has only elementary/primary school diploma). Column [1] shows that the probability of completing post-secondary education declined 1.5 percentage points for age group 40–45. In contrast, the probability of graduation from elementary/primary school and high school increased 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 or primary school. This suggests that those individuals affected from the protests would normally have gone to or completed a post-secondary education, but did not do so due to the 1978–82 turmoil. So, the number of high school graduates increased more than its ordinary trend, which suggests that potential college graduates have stayed as high school graduates as a consequence of violent student protests.

We also compare the age groups 46–51 and 52–57. We find that 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 46–51/52–57 gaps is 1.6 percentage points—approximately the percentage point decline in completing post-secondary education for 40–45 age group. This tells us that the main group affected from the student protests were those who would have normally continue post-secondary education in the absence of these events. Figures (10) and (11) confirm this finding.

Effects on years of schooling. We run a similar regression to visualize the trend for average years of schooling. The coefficients of age dummies are plotted in Figure (12). Clearly, the trend is flatter for age group 40–45 relative to the age groups 34–39 and 46–51. We conclude from Figures (4) – (12) that the sole reason for the break in the trend is the decline in post-secondary educational attainment due to the turmoil.

Effects on the probability of wage employment. The analysis above shows that the 1978–82 turmoil substantially reduced the post-secondary educational attainment for age group 40–45. Now we check whether it has any effect on the probability of wage employment. In the regression, the dependent variable is a dummy taking 1 if an individual is wage employed (regular employee and casual employee) and 0 otherwise (employer, self employed, and unpaid family worker). The sample includes all individuals who are employed. Figure (13) plots the coefficients of age dummies. The trend of probability of wage employment is smooth over the entire range of ages. The estimated coefficients lie almost on a slightly concave curve. Thus, the 1978–82 turmoil had no effect on wage employment. Similarly, the turmoil did not affect labor force participation, employment, and labor informality [see Section 6.3].

Although the probability of being a wage-earner did not change, the school attainment among wage-earners of different age groups should have changed as a result 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 males of age 40–45 are the ones whose post-secondary educational attainment have been affected the worst. Therefore, we can conclude that the group most affected by the 1978–82 turmoil is male wage earners of age 40–45 (in 2005 survey). Thus, we restrict our sample to male wage earners of age 34–51 for our main estimations. The age groups 46–51 and 34–39 are selected for comparison purposes.

Effects of wages. We assess whether the decrease in post-secondary education for men of age 40–45 is reflected on earnings. We run a regression [similar to Equation (5.1)] in which the dependent variable is log hourly wage standardized to 26 years of experience—following Altonji, Bharadwaj, and Lange (2012). The coefficient of ages (relative to age 51) are plotted in Figure (14) and it is clear that log hourly wages increase from age 50 to age 47 and begins to decline for younger cohorts, similar to the trend in post-secondary educational attainment in Figure (4).

On the other hand, Figures (4) – (12) show that the only reason for the slowdown in years of schooling for ages 40–45 is the decline in post-secondary educational attainment. Based on this finding, we run the same regression as mentioned above by using the data restricted to men with at least high school education. The coefficients of ages are plotted in Figure (15) and it is clear that the log hourly wages of ages 40–45 are statistically negative relative to age 51. Since the instrument exploiting the student protests affected only post-secondary education, Figure (15) confirms that the decrease in post-secondary education leads to a decrease in earnings for men aged 40–45.

5.2 Instrument validity for the IV estimation

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 wage, s_i is a measure of schooling, $\boldsymbol{X}_i$ is a set of other observables variables assumed to affect labor income, and ϵ_i is an error term assumed to be independent of the explanatory variables [Griliches (1977)]. However, in this setting, the causal effect of education on labor income may not be consistently estimated, because of the standard omitted variables problem—also known as the ability bias. A possible solution to this problem is to use an IV, which requires at least one observable covariate that affects labor income only through schooling.

Above we have indicated that the post-secondary educational attainment significantly declined for individuals aged 40–45 (approximately born between 1960 and 1965 or being the age of 13 to 18 years old in 1978) compared to individuals aged 46–51 due to the student protests in the late 1970s and the subsequent coup. It would normally be expected that the probability of completing post-secondary education has an upward trend across age cohorts as the age declines (i.e., over time). Therefore, we use a 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 schooling.

In a heterogeneous-outcome framework, the IV method has the potential to estimate 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, Imbens, and Rubin (1996); Card (2001)]. There are two key conditions [see Imbens and Angrist (1994)]. The first one is the existence of a valid instrument. Because an individual's year of birth is randomly assigned and probably unrelated to individuals' innate ability, motivation, or family characteristics, it seems reasonable to assert that the wage decline for age group 40–45 is due to decline in their post-secondary education—after standardizing the experience in the labor market. Thus, potential outcomes in the labor market are independent of the instrument and exclusion restriction assumption is satisfied. We show in the previous sub-section that completing post-secondary education is a nontrivial function of z_i. The second condition is monotonicity. This assumption ensures that the instrument affects the post-secondary education in a monotonic way [Imbens and Angrist (1994); Angrist, Imbens, and Rubin (1996)]. We also document in detail that the 1978–82 turmoil negatively affected "all sub-samples of the population" such as males, females, wage earners, and non-wage earners, which suggests that the monotonicity condition is also satisfied. Based on these assumptions, the IV estimates using z_i is the average treatment effects for those who did not continue post-secondary education due to the 1978–82 turmoil, but who would have normally had a post-secondary degree. In our IV estimates, we restrict our sample to male wage earners of age 40–51.

5.3 First-stage and reduced-form estimates for male wage earners

In this subsection, we report the first-stage and reduced form estimates for the returns to schooling. We run three different regressions based on the following equation:

$$s_i = \alpha + \beta z_i + \boldsymbol{X}_i'\boldsymbol{\Pi} + \epsilon_i, \tag{5.4}$$

where z_i is a dummy variable taking 1 if the individual i is of age 40–45 and 0 if 46–51, $\boldsymbol{X}_i$ is a vector of covariates, and ϵ_i is an error term.

In the first regression, the dependent variable (s_i) is a binary variable indicating whether the individual has completed post-secondary education or not. In the second and third regressions, we use years of schooling and log hourly wages as the dependent variables, respectively. In all regressions, we focus on male wage earners of age 40–51 (in 2005). We define male in the 40–45 sample as the treatment (or affected) group and those in the 46–51 sample as the control (or comparison) group. The results are presented in Table (5). Columns [1] to [3] indicate that the probability of finishing post-secondary education is 6.6–7 percentage point lower for the treatment group. Similarly, years of schooling is also lower by 0.22–0.28 year—see columns [4]–[6]. The last three columns document the effect of 1978–1982 turmoil on wages. The corresponding estimates suggest that wages of men in 40–45 sample is 2.6–3.5 percent lower than wages of men in the 46–51 sample. Yet, the effect on wages turn only marginally significant when we include both 26 NUTS2-level region-of-residence and urban/rural dummies as control variables in column [9].

The post-secondary educational attainment rate is 21.1 percent for male wage earners in the 46–51 sample. This rate is approximately 6.6–7 percentage points above the figures for the 40–45 sample due to the 1978–1982 turmoil. In other words, the post-secondary educational attainment of the ones in the treatment group is changed by the IV and the rate of change is around least 31 percent (6.6/21.1), which is substantial. We assume in this calculation that the post-secondary attainment rate for the 40–45 sample would have at least remained the same as that in the 46–51 sample, if the 1978–1982 turmoil had not happened.

6 Results and Discussion

This section presents (*i*) the impact of the 1978–1982 turmoil on the wage distribution and occupational structure of the affected population and (*ii*) the estimates for returns to schooling using the turmoil as an IV. The first subsection provides the results of IV estimations. The second subsection presents the effects on the wage distribution using semi-parametric counterfactual density estimation methods. The third subsection explores the effects of the turmoil on the occupational structure. The last subsection performs robustness checks.

Table (6) reports summary statistics for the treatment (40–45) and control (34–39 and 46–51) age groups. The treatment group of clearly has a lower average log hourly wage, fewer average years of schooling, and a lower average post-secondary educational attainment than the control groups. Moreover, this group has a higher average high school graduation rates compared to the previous (34–39) and subsequent (46–51) age groups. This suggests that the decline in school attainment for the 40–45 sample comes entirely from the decline in post-secondary education due to the turmoil.

6.1 Estimating returns to education

6.1.1 Estimating returns to education for an additional year in school

The 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 1978–1982 turmoil and (*ii*) the turmoil affects wages (as of 2005) only through its impact on post-secondary educational attainment. Based on these assumptions, we use the political turmoil as an IV to estimate the causal effect of additional years of schooling on wages. We have already shown that the instrument z_i —taking 1 if individual i is of age 40–45 and 0 if 46–51—is highly likely a valid one. The first-stage and reduced form of this IV specification are already presented in Table (5). The results suggest that the instrument has satisfactory explanatory power at the first stage.

Estimates for the returns to an additional year in school are presented in Table (7). In all regressions, the dependent variable is the log hourly wage standardized to experience (at 26 years) and the sample is male wage earners of age 40–51. The first row provides the OLS estimates of Equation (5.3). Column [1] indicates that the estimated return to schooling is 11.2 percent and is not affected by introducing region of residence and urban/rural dummies as control variables—also see columns [2] and [3].

The second row of Table (7) reports the 2SLS estimates. In column [1], there are no control variables and the point estimate (12.7 percent) 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. To capture this possibility, regional-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.

In a heterogeneous-outcome framework, the returns to education estimates—ranging between 11.6 and 12.8 percent—are the average causal effects of an additional year in school for those affected from the 1978–1982 turmoil. They would have completed a post-secondary degree, if these events had not occurred. We have already documented that the 40–45 sample has fewer average years of schooling and a lower average post-secondary graduation rate, but a higher average high school graduation rate compared to the previous (34–39) and subsequent (46–51) age groups. Figures (4) to (12) provide more evidence to support this claim. Therefore, our instrument only affects post-secondary education, which strengthens the LATE interpretation.

The individuals whose schooling attainment is changed by the instrument are at least 31 percent of individuals having post-secondary education in the male sample of wage earners aged 40–45. Average treatment effect on the treated is a weighted average of effects on "always-takers" and "compliers" [Angrist and Pischke (2008)]. In addition, those individuals affected from the 1978–1982 turmoil are not marginal individuals who are indifferent between going to university or not. Those affected were the dropouts in post-secondary education or would have gone to universities if these events had never happened. Carneiro (2003) shows that the return to education for the average student in college is systematically above the return to education for marginal individual in the US. Therefore, our estimates can be interpreted as a close approximation to the average causal effects of an additional year of schooling in post-secondary education.

Our estimates for returns to post-secondary schooling are similar to those reported for developed countries. Belzil and Hansen (2002) use a structural dynamic programming model to estimate marginal returns to schooling in the US. They find that log wage regression is convex in schooling and the estimated marginal returns are less than 1 percent per year until grade 11, 3.7 percent in grade 12, 6 percent in grade 13, and range from 10.8–12.7 percent between grades 14 and 16. Psacharopoulos and Patrinos (2004) also point out that average returns to higher education is 18.2 for Asian (non-OECD) countries, 18.8 percent for Europe/Middle East/North Africa (non-OECD) countries, and 11.6 percent for the OECD countries.

The convexity of the log wage regression function implies that marginal returns are increasing in the level of schooling. The instrument that we use only affects post-secondary education. Since different instruments may define different "effects" in a heterogeneous-outcome framework [Heckman, Lochner, and Todd (2006)], the findings of the current paper would be consistent with low returns to elementary school grades of Torun (2015) and Aydemir and Kirdar (2017) in light of the evidence presented by Belzil and Hansen (2002). Thus, the log wage regression may also be convex in schooling in Turkey.

We repeat the same regressions with the log monthly wage as the dependent variable instead of the log hourly wage. The results are presented in Table (8). Because there is a negative correlation between hours worked and average years of schooling in Turkey, the returns to schooling for an additional year for the log monthly wage are approximately 2–3 percentage points lower than the regressions with log hourly wage as the dependent variable. The 2SLS estimates in both regressions are slightly above the corresponding OLS estimates.

6.1.2 Estimating returns to education for post-secondary degree

In this subsection, we estimate the effects of getting a post-secondary degree on wages among males of age 40–51. The survey data includes information on the highest completed level of schooling, but there is only one level for post-secondary schooling. We define a binary variable d_i taking 1 if i has post-secondary education and 0 if high school.

We calculate the average years of schooling for the post-secondary education as four years based on the 2008 TDHS. This suggests that typical post-secondary education a four-year college in Turkey. Therefore, we estimate the effect of obtaining a four-year college degree relative to a high school diploma. The results are presented in Table (9). In all regressions, the dependent variable is the log hourly wage standardized to experience (26 years). The first row documents the OLS estimates. Column [1] shows that the returns to a four-year college degree are 50 percent greater than those for high school education 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 row of Table (9) presents the 2SLS estimates. In column [1], there is no control variable and the point estimate (58 percent) is slightly above the OLS estimate. Including region-of-residence and urban/rural dummies as control variables (columns [2] and

[3]) do not change the results significantly. The first-stage F-statistics for all regressions are close to 50, confirming the strength of the instrument.

6.2 Counterfactual density estimation

Next we implement a semi-parametric procedure to analyze the impact of 1978–1982 turmoil on the wage distribution of men of age 40–45. We implement the density re-weighting procedure developed by DiNardo, Fortin, and Lemieux (1996) to estimate a counterfactual density of labor income. With this procedure, we apply weighted kernel methods based on the assumption that the post-secondary attainment rate for men of age 40–45 remained the same as that for men of age 46–51. This assumption roughly resembles the counterfactual experiment that the 1978–1982 did not hit the 40–45 age group and, thus, the wage distribution is not changed. We basically re-weight the 40–45 sample to have the same distribution of post-secondary education as the sample of 46–51. We then compare how labor income is distributed in the re-weighted 40–45 sample in the actual distribution of the same sample. This comparison provides a clear visual representation of how the change in post-secondary attainment due to the 1978–1982 turmoil affected the density of wages in the treatment group.

The actual and counterfactual density estimates are obtained by the kernel density estimator. 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:

$$\hat{f}_h(w) = \sum_{i=1}^{n} \frac{\theta_i}{qh} K\left(\frac{w - W_i}{h}\right), \tag{6.1}$$

where $q = \sum_{i=1}^{n} \theta_i$, h is bandwidth, and $K(\cdot)$ is the kernel function. We choose analytic weights in the estimation because weights are rescaled so that $\sum_{i=1}^{n} \theta_i = n$.[17] In our data, the weights are the HLFS sampling weights.

We prefer to stick with the notations and notions introduced in the original paper by DiNardo, Fortin, and Lemieux (1996). In the estimation procedure, each individual observation belongs

[17] This ensures that the software version (STATA) of kernel density estimation is compatible with the estimator proposed by DiNardo, Fortin, and Lemieux (1996).

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 one 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 $F(d|z_d)$ at time z_d as follows;

$$\begin{aligned} f_z(w) &= \int_{d\in\Omega_d} dF(w, d|z_{w,d} = z) \\ &= \int_{d\in\Omega_d} f(w|d, z_w = z)dF(d|z_d = z) \\ &= f(w; z_w = z, z_d = z) \end{aligned} \tag{6.2}$$

where Ω_d is the domain of the individual attributes. To be compatible with the previous subsections, z is a binary variable taking 1 for the 40–45 sample and 0 for the 46–51 sample. Thus, the expression of $f(w; z_w = 1, z_d = 1)$ 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 had remained as in individuals aged 46–51 without changing the wage schedule observed for the 40–45 sample.[18]

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

$$\begin{aligned} f(w; z_w = 1, z_d = 0) &= \int_{d\in\Omega_d} f(w|d, z_w = 1)dF(d|z_d = 0) \\ &= \int_{d\in\Omega_d} f(w|d, z_w = 1)\psi_d(d)dF(d|z_d = 1), \end{aligned} \tag{6.3}$$

where the re-weighting function $\psi_d(d) = dF(d|z_d = 0)/dF(d|z_d = 1)$. As seen in Equation (6.3), 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 be-

[18] Note that the general equilibrium effects are ignored.

tween the cohorts in the distribution of d can account for any difference between the cohorts in the marginal distribution of vector of unobserved skills [Altonji, Bharadwaj, and Lange (2012)].

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

$$\hat{f}(w; z_w = 1, z_d = 0) = \sum_{i \in I_1} \frac{\theta_i}{qh} \widehat{\psi}_d(d) K\left(\frac{w - W_i}{h}\right), \tag{6.4}$$

where I_1 is the set of indices of individuals aged 40–45. Our goal is to estimate the effects of the decline in post-secondary educational attainment. Thus, the individual-level attribute in this study is a binary variable taking 1 if the individual finishes post-secondary education and 0 otherwise. In that case, the difference between the actual density and the counterfactual density indicates the effect of a decline in post-secondary attainment on the distribution of wages in the treatment group.

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

$$\psi_d(d) = \frac{\mathbb{P}[z_d = 0|d]}{\mathbb{P}[z_d = 1|d]} \cdot \frac{\mathbb{P}[z_d = 1]}{\mathbb{P}[z_d = 0]}. \tag{6.5}$$

A probit or logit model can estimate the probability of being in period z given the individual attributes d. $\mathbb{P}[z_d = 1]$ is equal to the weighted number of observations in the 40–45 sample divided by the weighted number of observations in 40–45 and 46–51 samples. We apply the probit model, as in DiNardo, Fortin, and Lemieux (1996), to estimate the re-weighting function and plot the weighted kernel density estimates of the counterfactual [dotted line in Figure (16)] and the actual [solid line in Figure (16)] densities. We choose wages as log hourly wages standardized to 26 years of experience for men—as in the previous subsections. Both lines are superimposed in Figure (16).[19]

[19] The STATA optimal bandwidth and Gaussian kernel function are chosen, but the results are not sensitive to the choice of bandwidth and alternative kernel functions.

The vertical line indicates the log minimum wage in 2005. It is computed 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 around the minimum wage and the second peak appearing around 1.45 log wage value—where the mean of this sample is approximately 1.01 in log terms.

The difference between actual (solid line) and counterfactual density (dotted line) 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 sample. Strikingly, the decline in post-secondary education pushed these individuals from the high-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 1978–82 turmoil had not occurred.

6.3 A shift in occupations

In this subsection, we explore the impact of 1978–82 turmoil on the occupational structure. Before this analysis, we will address the following question: does the turmoil affect other labor market outcomes such as labor force participation, employment, and labor informality? In order to answer this question, we run three regressions based on the following simple model:

$$s_i = \alpha + \sum_{c=30}^{54} \beta_c d_{ic} + \boldsymbol{X}_i' \boldsymbol{\Pi} + \epsilon_i, \tag{6.6}$$

where s_i is a binary variable for labor market status, d_{ic} is a 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 HLFS data used in this analysis is restricted to men of age 34–51. In all regressions, we use 26 NUTS2-level region-of-residence dummies and urban/rural dummy as the vector of covariates. Individuals of age 51 in 2005 serve as the reference group—i.e., the omitted age dummy.

The first regression is for labor force participation. Each coefficient β_c can be interpreted as an estimate of the probability of being in the labor force for the corresponding age relative to age 51. The coefficients of ages between 34 and 50 are plotted in Figure (17), which shows that the trend for labor force participation is smooth over the 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 (being registered with the social security institution in the current job), respectively. The results are reported in Figures (18) and (19). 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 affect on employment outcomes.

We show, however, that the turmoil significantly 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 compute the mean of log wage for each occupation for men and then we separately find the percentage of individuals in each occupation for three age groups (34–39, 40–45, and 46–51). We sort the occupations based on average log wage values and, accordingly, we construct five main 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 first group is corporate managers and professionals and their mean log wage values are over 1.5—approximately corresponding to the second peak in distribution of wages in Figure (16). 12 percent of those in the 40–45 sample are within this category. This ratio is less than the other two age groups—being nearly 7 percentage points lower than the 46–51 sample. This difference is compatible with the difference between actual and counterfactual density estimations in the previous sub-section. Thus, it can be concluded that individuals of age 40–45 have less attractive jobs, on average, in terms of pay.

The second occupation group consists of technicians, associate professionals, and clerks. Clearly, these occupations have less education requirements than the first group. The percentage of this group in the 40–45 sample 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 the sub-major divisions whose mean log wage values between the minimum wage and 1.5 [the second peak in Figure (16)], the fraction of individuals in groups 34–39, 40–45, and 46–51 becomes 68.7, 72.5, and 66.3, respectively. This suggests that the 40–45 sample 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.

6.4 Robustness checks for missing data

The wage data contains missing values. About 2 percent of male individuals of age 40–51 did not declare 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 schooling dummies, 27 sub-major divisions of occupations, 26 NUTS2-level region-of-residence dummies, and an urban/rural dummy. We find that some variables are statistically significant, and thus the missing values are not random.

Although the fraction of missing data in our sample is very low, we re-estimate sampling weights to adjust for the missing wage values following the procedure introduced by Altonji, Bharadwaj, and Lange (2012) to check whether the results are robust to missing data for a set of the previous estimations. The comparison is presented in Table (11). The results suggest that our estimates are robust to adjusting sample weights for missing data on wages.

7 Conclusion

Between 1978 and 1980, Turkey experienced violent student protests. Almost 20 youths were killed daily. Universities were at the center of the violence. This violence ultimately led to a military intervention in 1980, which came with additional arrests and suppression lasting several years. The 1978–1982 turmoil adversely affected post-secondary educational attainment in Turkey. In this study, we exploit the exogenous drop in the number of graduates and new admissions due to this turmoil (*i*) to document its long-term impact on wages and occupation structure and (*ii*) to estimate the causal impact of education on earnings using the turmoil as an IV.

We find that the group most profoundly affected by the 1978–82 turmoil is male wage earners of age 40–45 (in 2005), who were 15–20 years old in 1980. These events led to a 6.6–7 percentage points decline in the probability of completing post-secondary education, a 0.22–0.28 decline in average years of schooling, and a 2.6–3.5 percent drop in wages for the affected group. We also find that the decline in post-secondary educational attainment led to a permanent shift in occupations from high-pay to low-pay ones. These findings suggest that violent political turmoils can permanently erode a country's human capital and, therefore, can adversely affect long-term growth prospects.

Using the turmoil as an IV (which generates an exogenous drop in post-secondary educational attainment and wages), we find that the returns to education range between 11.6–12.8 percent in Turkey. This finding suggests that the rate of return to education in a typical developing economy is more or less similar to the estimates reported for developed economies. This is the first paper using student protests (which were widespread throughout the world in 1960s and 1970s) as an IV to estimate returns to education.

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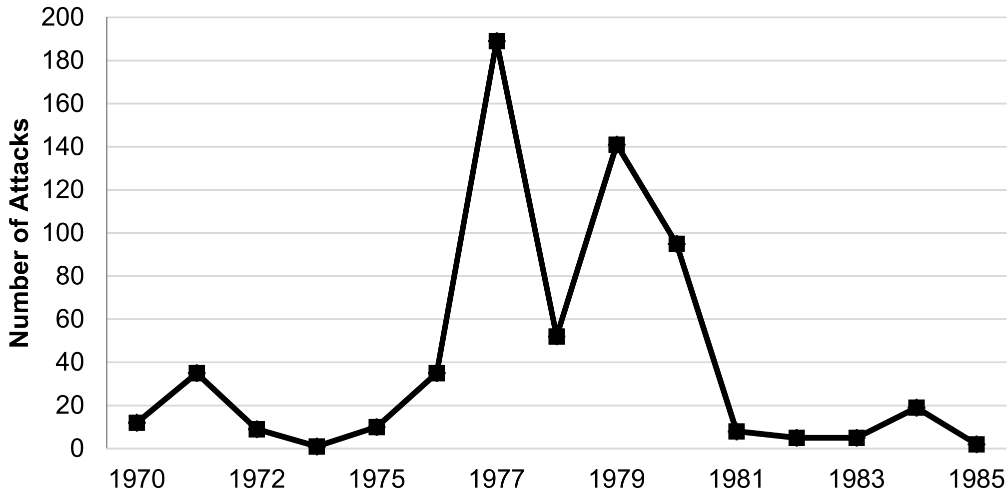

Figure 1: **Number of Terrorist Attacks in Turkey (1970–1985).** *Source:* National Consortium for the Study of Terrorism and Responses to Terrorism (START), (2016). Global Terrorism Database [GTD from 1970 to 1991]. Retrieved from https://www.start.umd.edu/Gtd

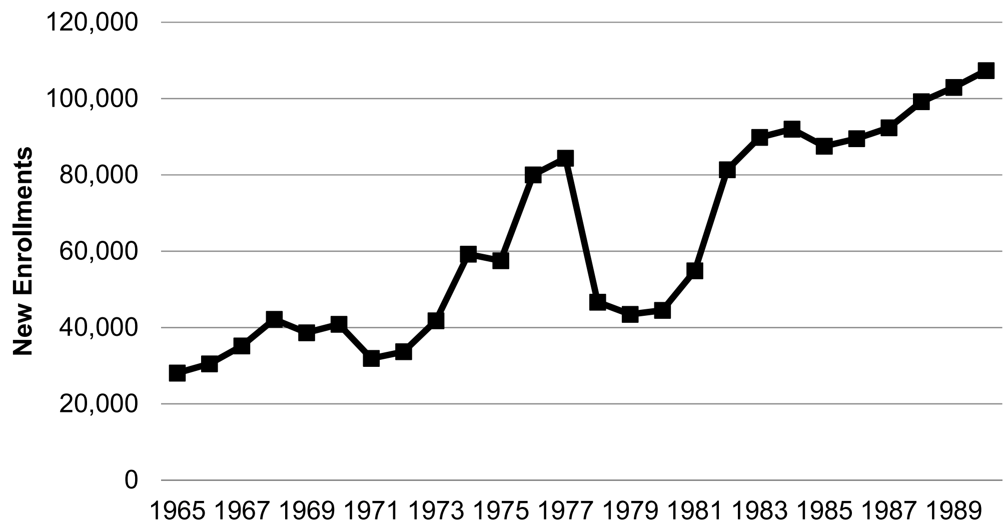

Figure 2: **New Enrollments in All Higher-Education Institutions in Turkey (1965–1990).** *Source:* Authors' calculations based on National Education Statistics (TURKSTAT) and Academic Year Higher Education Statistics (OSYM).

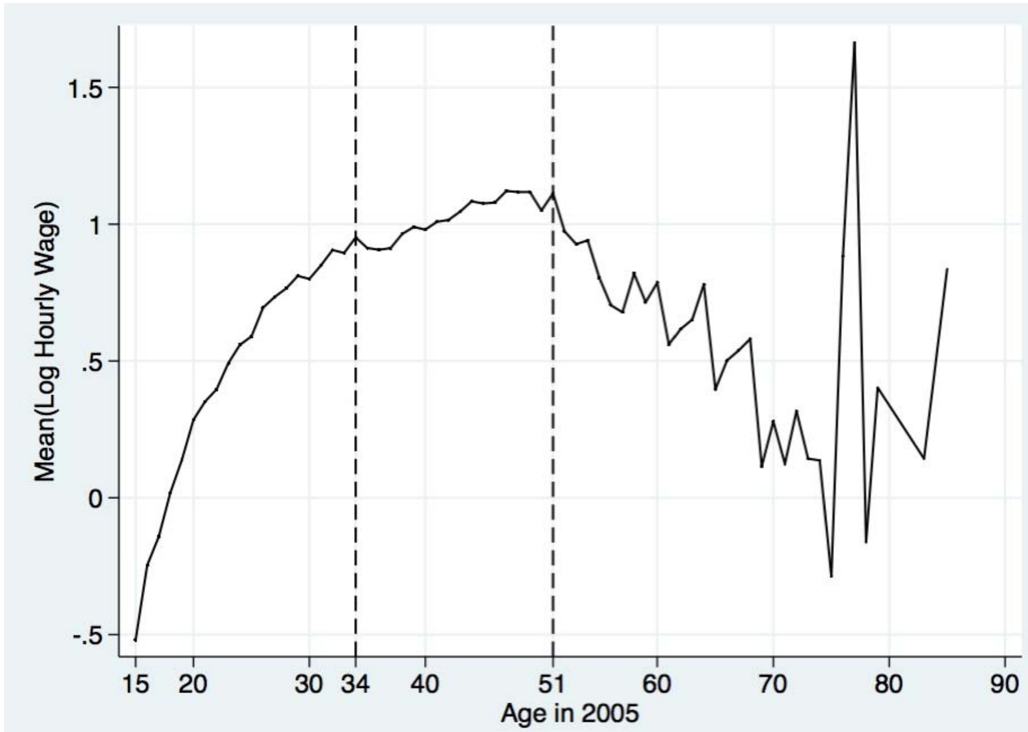

Figure 3: **Average Log Hourly Wage for Men by Age.** *Source:* Authors' calculations based on the 2005 Turkish Household Labor Force Survey.

Figure 4: **Coefficients of Age Dummies – Probability of Completing Post-Secondary Education.** The specification includes 26 NUTS2 region of residence, urban/rural, and gender dummies. Age 55 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

Figure 5: **Coefficients of Age Dummies – Probability of Completing Post-Secondary Education for Men.** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Age 55 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

Figure 6: **Coefficients of Age Dummies – Probability of Completing Post-Secondary Education for Women.** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Age 55 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

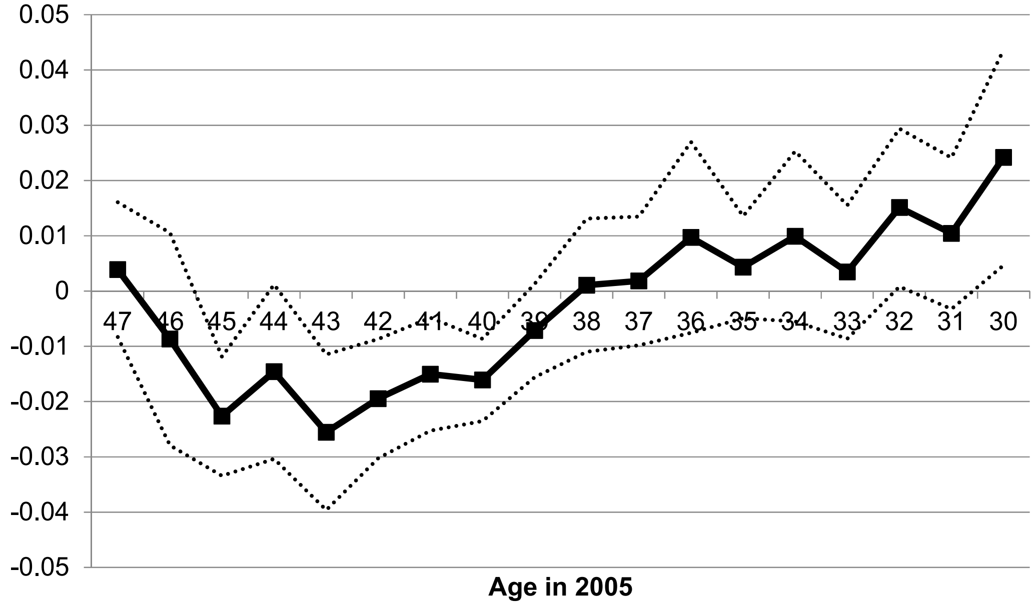

Figure 7: **Coefficients of Age Dummies – Probability of Completing Post-Secondary Education (age 48 as reference group).** The specification includes 26 NUTS2 region of residence, urban/rural and gender dummies. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

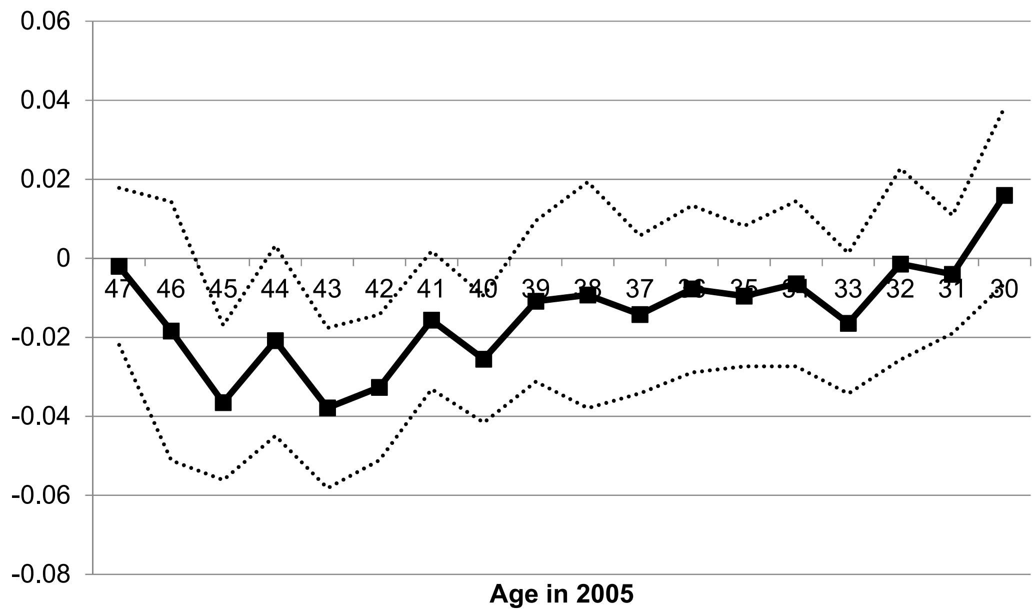

Figure 8: **Coefficients of Age Dummies in Estimating the Probability of Completing Post-Secondary Education for Men (age 48 as reference group).** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

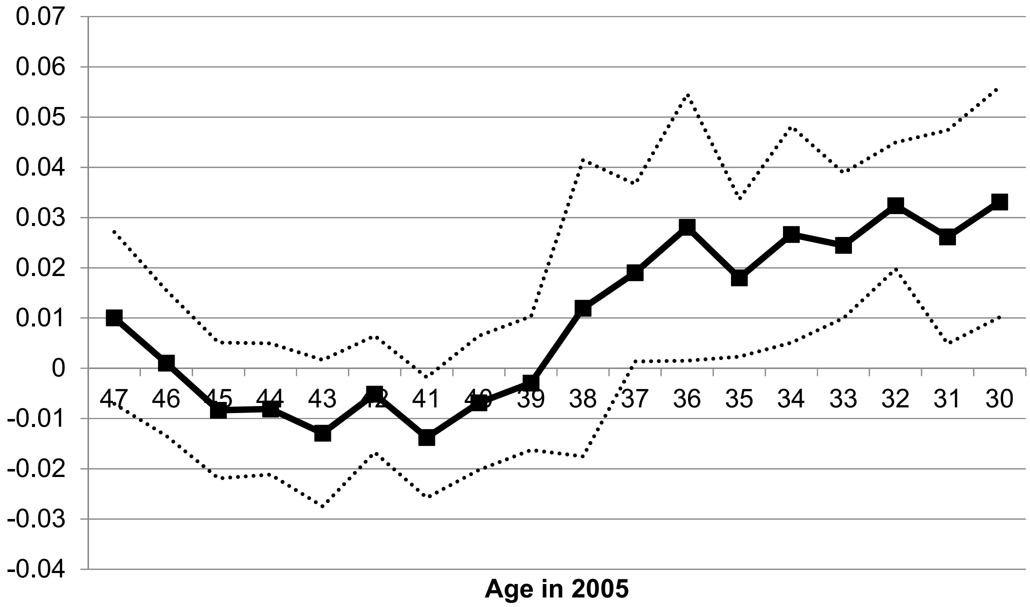

Figure 9: **Coefficients of Age Dummies in Estimating the Probability of Completing Post-Secondary Education for Women (age 48 as reference group).** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

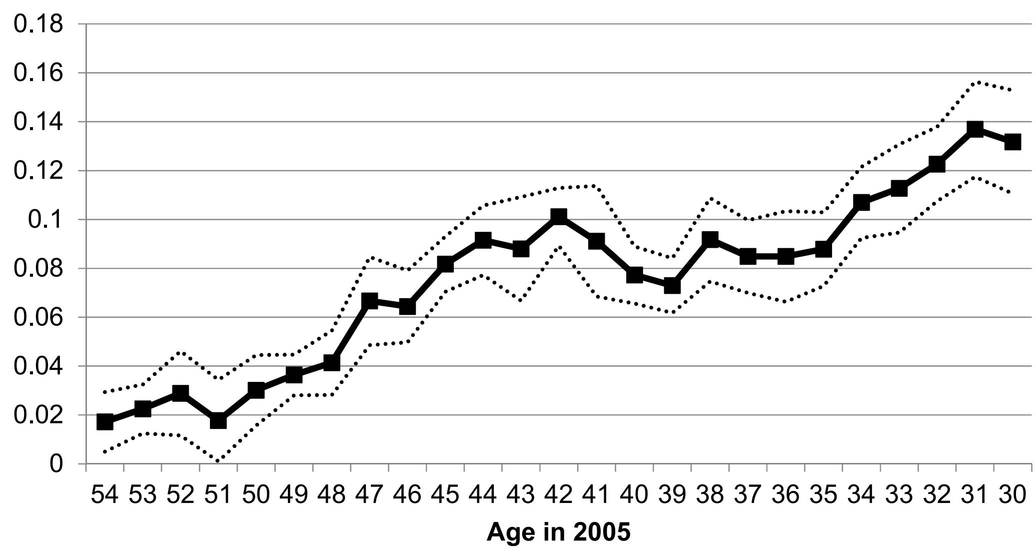

Figure 10: **Coefficients of Age Dummies in the Estimating the Probability of Graduation From High School.** The specification includes 26 NUTS2 region of residence, urban/rural and gender dummies. Age 55 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

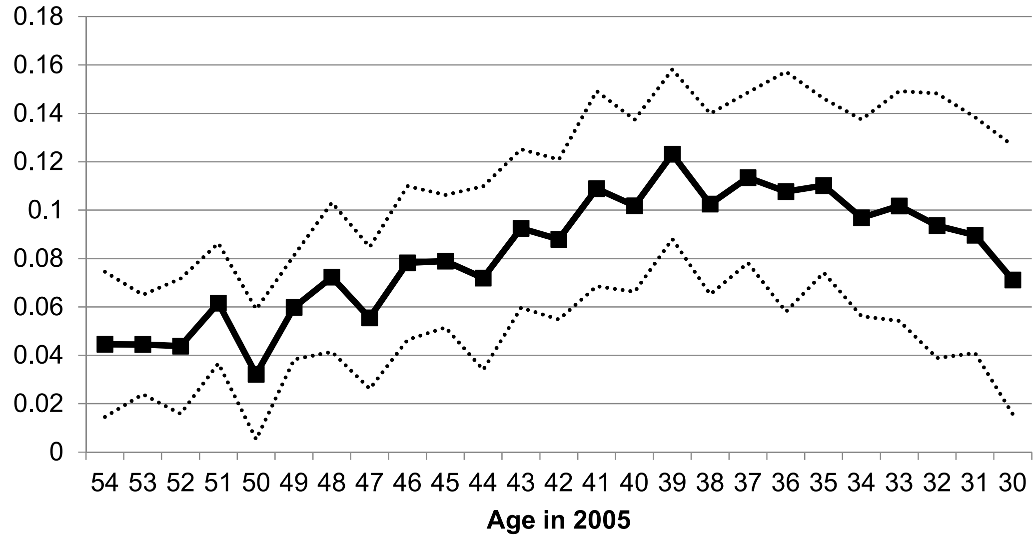

Figure 11: **Coefficients of Age Dummies in the Estimating the Probability of Graduation from Elementary/Primary School.** The specification includes 26 NUTS2 regions of residence, urban/rural and gender dummies. Age 55 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

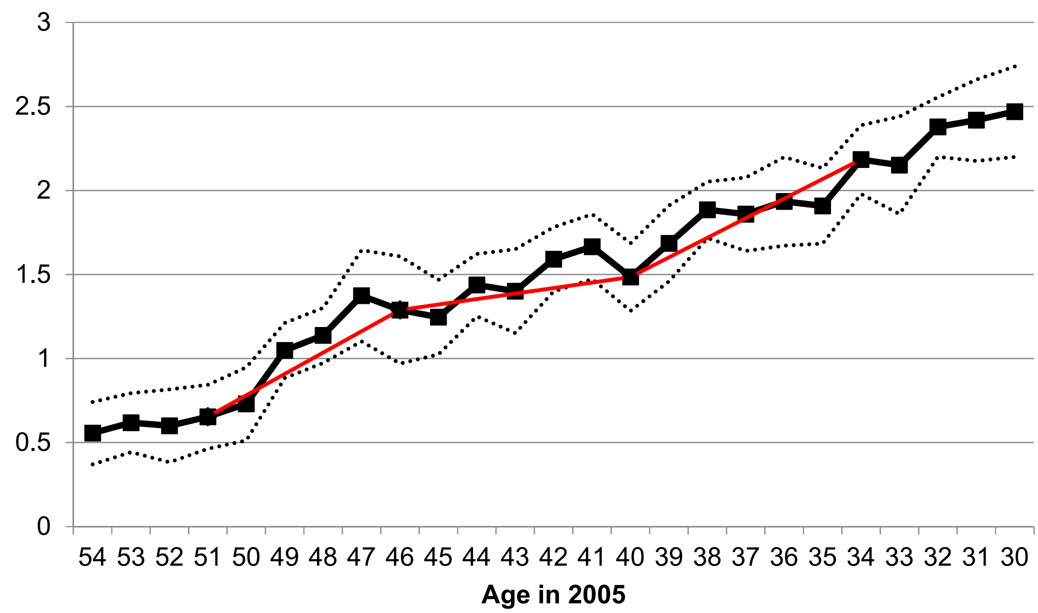

Figure 12: **Coefficients of Age Dummies in the Estimating Years of Schooling.** The specification includes 26 NUTS2 region of residence, urban/rural and gender dummies. Age 55 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

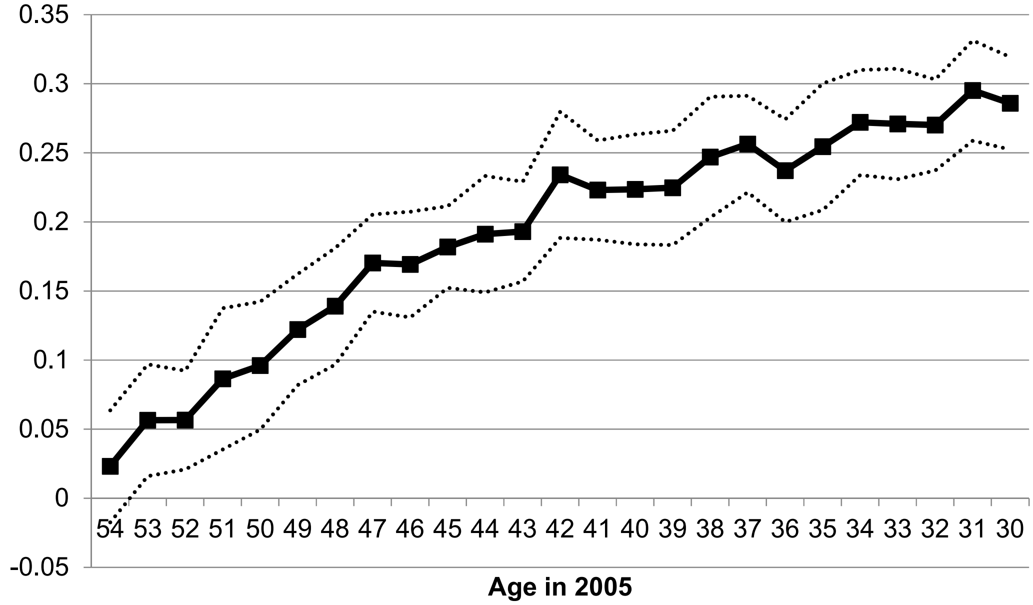

Figure 13: **Coefficients of Age Dummies in the Estimating the Probability of Wage Employment.** The specification includes 26 NUTS2 region of residence, urban/rural and gender dummies. Age 55 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

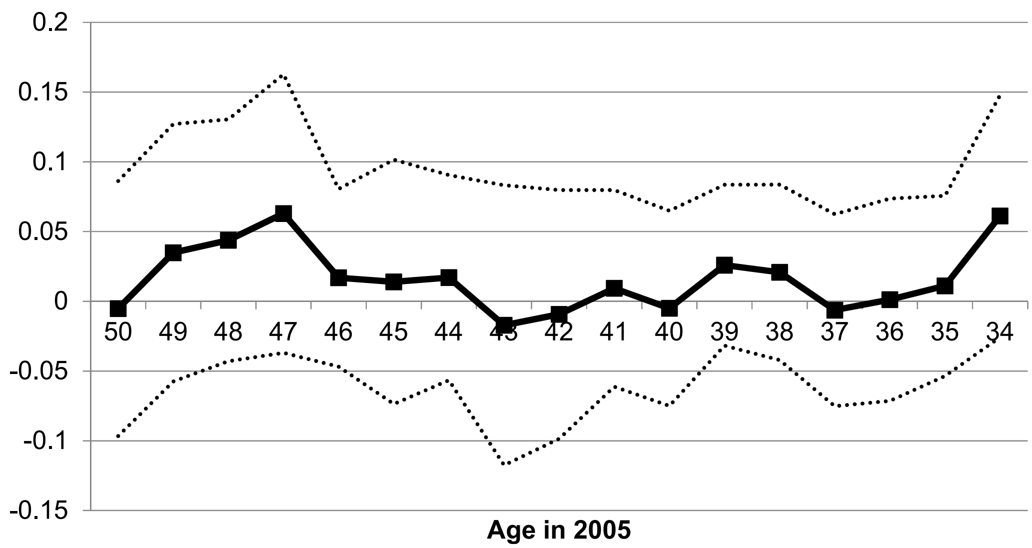

Figure 14: **Coefficients of Age Dummies in the Estimating Log Hourly Wage for Men.**The specification includes 26 NUTS2 region of residence and urban/rural dummies. Age 51 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

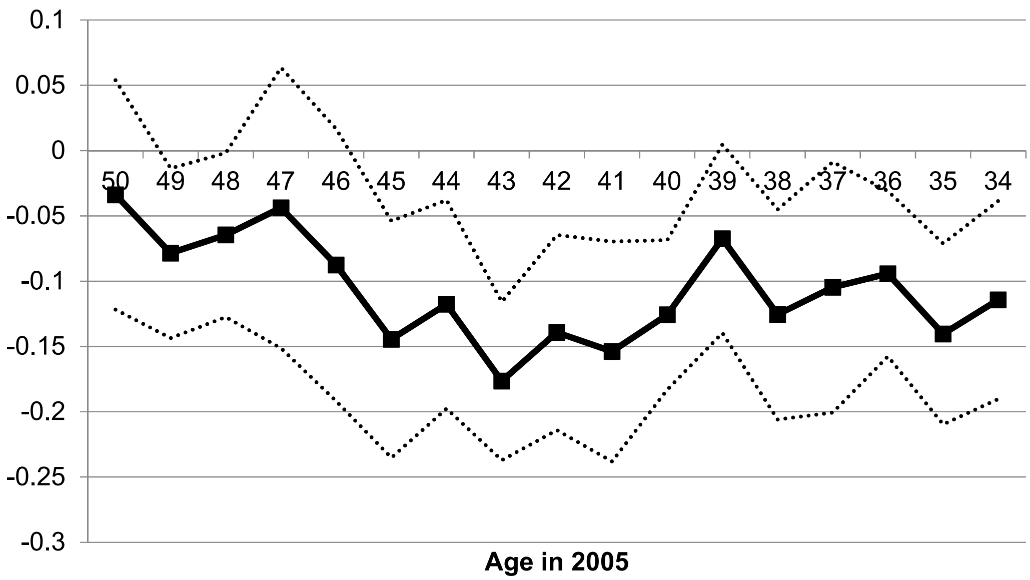

Figure 15: **Coefficients of Age Dummies in the Estimating Log Hourly Wage for Men with at least High School Education.** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Age 51 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

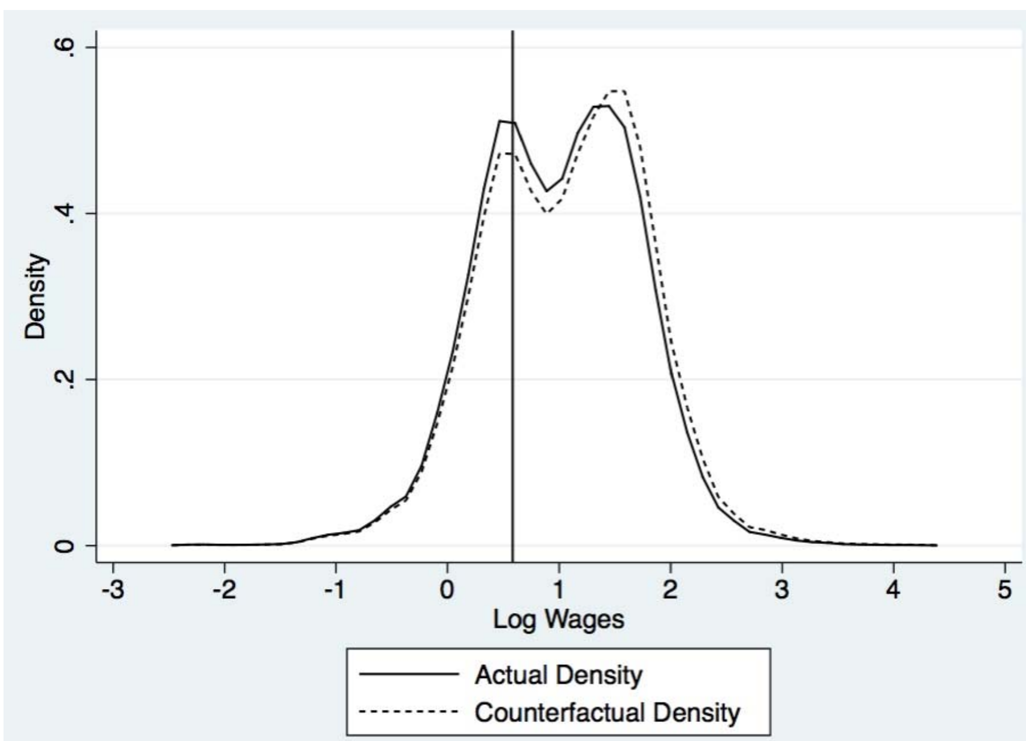

Figure 16: **The Actual and Counterfactual Density of Log Wages for Male Individuals Aged 40–45.** The sample includes male wage earners aged 40–51. Observations are weighted using the sampling weights so that the results are nationally representative.

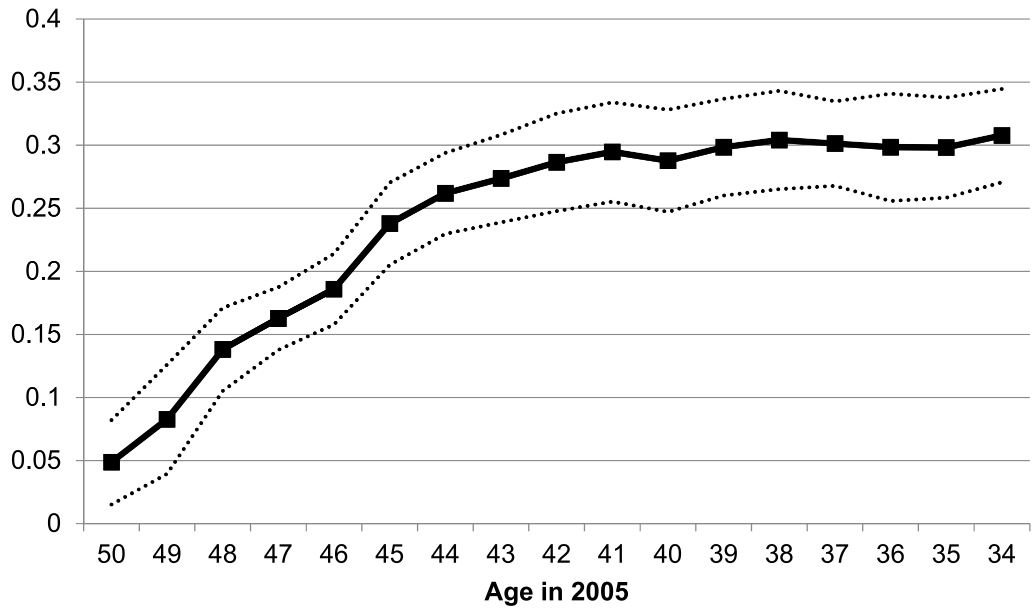

Figure 17: **Coefficients of the Age Dummies in the Estimating the Probability of Being in the Labor Force.** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Age 51 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

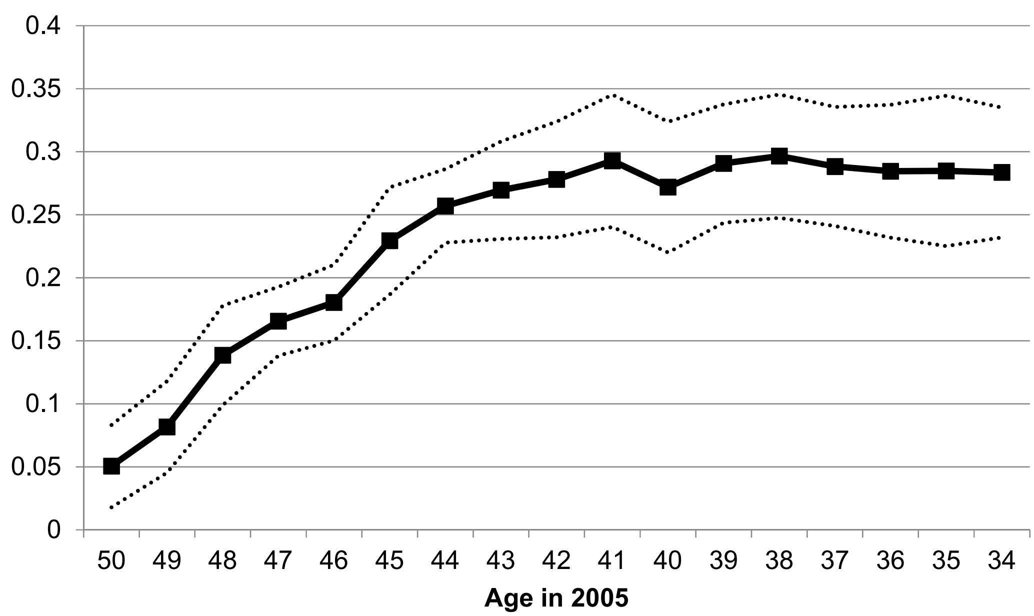

Figure 18: **Coefficients of Age Dummies in the Estimating the Probability of Being Employed.** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Age 51 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

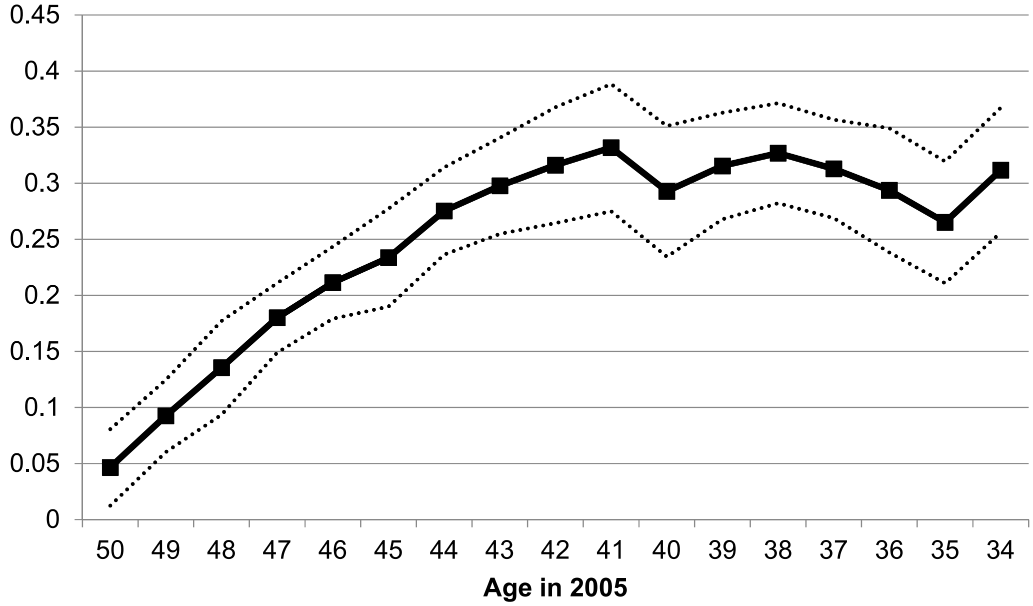

Figure 19: **Coefficients of Age Dummies in the Estimating the Probability of Being Formal Employed.** The specification includes 26 NUTS2 region of residence and urban/rural dummies. Age 51 is the reference group. Observations are weighted using the sampling weights so that the results are nationally representative. Broken lines indicate the 95-percent confidence interval based on standard errors clustered at region level (NUTS2).

Average Hours Worked in the Main Job 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

Table 1: **Source:** Authors' calculations based on the 2005 HLFS. Observations are weighted using the sampling weights so that the results are nationally representative.

Descriptive Statistics for Individuals Aged 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

Table 2: Observations are weighted using the sampling weights so that the results are nationally representative.

Estimations of the Probability Difference of Graduating between Aged 40–45 and Aged 46–51

	Dependent Variable		
	Post-secondary	**High school**	**Primary/Elementary**
	Degree==1	Degree==1	Degree==1
	Otherwise==0	Otherwise==0	Otherwise==0
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

Table 3: 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 the (NUTS2) region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

Effect of the 1978–82 Upheaval 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

Table 4: The specification includes 26 NUTS2 region of residence and urban/rural dummies. It also includes gender dummy for the total sample estimations. Observations are weighted using the sampling weights so that the results are nationally representative. Standard errors, clustered at the (NUTS2) region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

Effect of the Student Protests in the late 1970s and the Subsequent Coup in 1980 on the Probability of Completing Post-secondary Education, Years of Schooling, and Wage

	Dependent Variable								
	Post-secondary degree			**Years of schooling**			**Log hourly wage**		
	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]
Instrument (z_i)	-0.0696***	-0.0690***	-0.0664***	-0.2765***	-0.2575***	-0.2233**	-0.0351**	-0.0329*	-0.0259
	(0.0117)	(0.0119)	(0.0114)	(0.0914)	(0.0915)	(0.0881)	(0.0162)	(0.0164)	(0.0162)
Region of residence	No	Yes	Yes	No	Yes	Yes	No	Yes	Yes
Urban/rural status	No	No	Yes	No	No	Yes	No	No	Yes
# of observations	15,827	15,827	15,827	15,827	15,827	15,827	15,827	15,827	15,827

Table 5: 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 the (NUTS2) region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

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	8.193	0.142	0.243
Age 46–51	5,722	1.037	8.475	0.211	0.195

Table 6: 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.

OLS and 2SLS Estimates of the Returns to Education

	Dependent Variable: Log hourly wage		
	[1]	[2]	[3]
OLS	0.1123***	0.1125***	0.1110***
	(0.0045)	(0.0041)	(0.0039)
2SLS	0.1271**	0.1278**	0.1161*
	(0.0537)	(0.0592)	(0.0673)
Region of residence	No	Yes	Yes
Urban/rural status	No	No	Yes
F (excluded instrument)	9.2	7.9	6.4
# of Observations	15,827	15,827	15,827

Table 7: 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 the (NUTS2) region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

OLS and 2SLS Estimates of the Returns to Education

	Dependent Variable: Log monthly wage		
	[1]	[2]	[3]
OLS	0.0829***	0.0837***	0.0820***
	(0.0028)	(0.0023)	(0.0021)
2SLS	0.0981**	0.1036**	0.0912
	(0.0447)	(0.0051)	(0.0578)
Region of residence	No	Yes	Yes
Urban/rural status	No	No	Yes
F (excluded instrument)	9.2	7.9	6.4
# of Observations	15,827	15,827	15,827

Table 8: 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 the (NUTS2) region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

OLS and 2SLS Estimates of the Returns to College

	Dependent Variable: Log hourly wage		
	[1]	[2]	[3]
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)
Region of residence	No	Yes	Yes
Urban/rural status	No	No	Yes
F (excluded instrument)	49.7	51.5	52.8
# of Observations	6,309	6,309	6,309

Table 9: 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 the (NUTS2) region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.

Classification of Occupations and Their Percentages in the 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

Table 10: Log (hourly) wage values are the mean of individuals in the corresponding occupation.

OLS and 2SLS Estimates of the Returns to Education with Adjusting Missing Values

	Dependent Variable: Log hourly wage	
	Additional years of schooling	Post-secondary education
Panel A: With Missing Values		
OLS	0.1110***	0.5025***
	(0.0039)	(0.0311)
2SLS	0.1161*	0.5716***
	(0.0673)	(0.0935)
F (excluded instrument)	6.4	52.8
# of Observations	15,827	6,309
Panel B: Adjusted for Missing Values		
OLS	0.1112***	0.5037***
	(0.0039)	(0.0310)
2SLS	0.1119	0.5744***
	(0.0709)	(0.0935)
F (excluded instrument)	6.0	52.7
# of Observations	15,827	6,309
Region of residence	Yes	Yes
Urban/rural status	Yes	Yes

Table 11: 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 the (NUTS2) region level, are reported in parentheses. *** $p < 0.01$; ** $p < 0.05$; * $p < 0.1$.