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IZA DP No. 10597

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ABSTRACT

The Perfect Storm: Graduating in a Recession in a Segmented Labor Market*

This paper analyzes the effects of entry labor-market conditions on workers' career in Spain, a country well known for its highly segmented labor market and rigid labor-market institutions. In contrast with more flexible labor markets, we find that the annual earnings losses of individuals without a university degree are greater and more persistent than those of college graduates. For workers without a college degree, the effect is driven by a lower likelihood of employment. For college graduates, the negative impact on earnings is driven by both a higher probability of non-employment, and employment in jobs with fixed-term contracts. While a negative shock increases mobility of college graduates across firms and industries, there is no earnings recovery, just secondary labor-market job churning. Our results are consistent with tight regulations of the Spanish labor market such as binding minimum wages and downward wage rigidity caused by collective bargaining agreements.

JEL Classification: E32, J22, J31

Keywords: full and dynamic effect of poor labor market conditions at entry, wage rigidity, fixed-term and permanent contract

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I. Introduction

While recent research has focused on the scarring effects of unemployment in flexible labor markets, the long-term consequences of graduating during a recession when labor-market institutions are rigid, and permanent and fixed-term contracts co-exist, are less known.[1] On the one hand, institutions (minimum wage laws and collective bargaining agreements) tend to make wages rigid, potentially creating a situation where demand shocks are absorbed by employment losses. On the other, segmented labor markets tend to have a dual system of job protection, in which high-firing costs for individuals working under a permanent contract co-exist with no-firing costs for those with a fixed-term contract within the same firms and for the same type of jobs. While permanent contracts offer high levels of employment protection, accumulation of human capital, and generous benefits, fixed-term contracts impose penalties in the form of forgone experience, and higher levels of unemployment risk to those workers who hold them (Fernandez-Kranz *et al.*, 2013). If employers use fixed-term contracts as a flexible device to adjust employment in the face of adverse shocks, as opposed to screen workers to promote them into permanent contracts (Güell and Petrongolo, 2007), demand shocks could trap workers in secondary labor-market jobs. Moreover, as labor-market segmentation severely reduces mobility of workers with a permanent contract, smooth wage renegotiations based on current labor-market conditions are unlikely (Beaudry and DiNardo, 1991).

The objective of this paper is to analyze the long-term consequences of graduating during a recession in a rigid and segmented labor market. We argue that such analysis is particularly policy relevant in the current economic situation because these two features (rigid labor-market institutions and segmented labor markets) are present, with varying degrees of intensity, in many OECD economies, including many Continental European countries.

To do so, we use the Spanish Social Security records from the 2008 *Continuous Sample of Working Histories* (hereafter CSWH). We focus on male cohorts graduating from high school, vocational training, or college between 1979 and 1991, and observe their labor-market outcomes from a year after they graduated to 2008. Hence, our

[1] See Kondo (2008); Kahn (2010); Genda *et al.* (2010); Hershbein (2012); Oreopoulos *et al.* (2012); and Altonji *et al.* (*2016*) for research in North America. Several studies focus on countries outside of North America, such as Austria (Brunner and Kuhn, 2014), Flanders (Cockx and Ghirelli, *forthcoming*), Norway (Raaum and Roed, 2006), or Japan (Genda *et al.*, 2010).

longitudinal data covers a minimum of 19 years and a maximum of 29 years of work history after graduation. In addition, because we have access to contractual monthly wages, measurement error owing to recall bias or non-response is not a concern as it is with survey data. We argue that Spain is a suitable case to investigate this issue because it is probably the best example of a country that combines rigid labor-market institutions with a striking segmentation of its labor force.

We find that graduating into a time of high unemployment results in substantial and persistent annual earnings losses, which are greater and more persistent for the *least* educated. The average cumulated effect of the first ten years after entry of an eight percentage-point increase in the entry unemployment rate -- the average shift from a recession to a boom in Spain -- is a 9.6%, 12.5% and 6.4% decrease in annual earnings for high-school graduates, workers with vocational training, and college graduates, respectively. For college graduates, the negative effect persists for 5 years, and for those without a college degree, it persists for 7 years. These findings are robust to a variety of sensitivity tests and they do not appear to be driven by mobility across provinces, selective employment, or graduation decisions.

The evidence presented above shows that, in the presence of a rigid and segmented labor market, workers entering the labor market during a recession experience large and persistent earning losses, especially if they do *not* hold a college degree. These findings contrast with those found in a more flexible labor market such as the one in the US (Hershbein, 2012, and Genda *et al.,* 2010), but resemble findings by Genda *et al.* (2010) and Cockx and Ghirelli (*forthcoming*) in two other rigid labor markets: Japan and Flanders (Belgium), as explained in the discussion in Section IV.

Comparing our college graduates' results with those from the US and Canada, we find that our earning losses estimates are only slightly higher than those found by Oreopoulos *et al.* (2012) in Canada, and Altonji *et al.* (*2016*) in the US, and smaller than those found by Kahn (2010) (also in the US).[2] However, the mechanisms are quite different. In the Spanish case, *both* a higher probability of non-employment and employment in fixed-term-contract jobs drive the results (as opposed to lower wages, as

[2] Kahn (2010) finds large effects of entry conditions that are four to five times higher than those found by Oreopoulos *et al.* (2012), and Altonji *et al*. (*2016*). Kahn's estimates imply that the wages of college graduates would fall 25% the first year of entry, and 20% after five years of entry, due to an increase in the entry unemployment rate of 4 percentage points, the average increase in a typical U.S. recession. These differences may be due to the use of different datasets and estimation methods by the three authors. Our results are much more in line with those of Oreopoulos *et al.* (2012) and Altonji *et al.* (*2016*) although, as we explain in this paper, the mechanisms are quite different.

in North America). Although minimum wages are less binding among college graduates in Spain, collective bargaining agreements still drive wage determination for high-skilled workers, generating a downward wage rigidity that limits wage reductions, especially for permanent contract workers, during recessions (Font, Izquierdo, and Puente, 2014). This extremely weak wage pro-cyclicality in Spain prolongs employment losses and employment in the secondary labor market for high-educated workers, and prevents the wage adjustments observed among high-educated workers in Flanders. Indeed, Cockx and Ghirelli (*forthcoming*) find that a negative shock at labor-market entry drives down *initially* both employment and wages of high-educated workers, and, *after 5 years in the labor market*, only wages. Ten years after labor market entry, wage losses still amount to -6% in Flanders.

Another finding (that contrast from that of more flexible labor markets) is that we find no evidence that firm mobility among college graduates helps in the catch-up process. Oreopoulos *et al.* (2012) find that college graduates who entered the Canadian labor market in the midst of the recession tend to move to better jobs as their career advances, and this job mobility helps them reduce the negative wage gap from the beginning of their career.[3] While bad entry labor market is associated with an increase of the mobility of college-graduate workers across firms, industries, and provinces in Spain, this higher mobility does *not* help them catch-up; instead of moving to better jobs, workers churn across fixed-term contract jobs.[4] To put it differently, college graduates entering the labor market during a negative shock are trapped in the secondary market.

Our work also contributes to a growing literature analyzing the effects of labor-market conditions on workers careers in European countries. With the exception of Cockx and Ghirelli (*forthcoming*), most of these studies focus in a particular education group, which precludes understanding how human capital may attenuate or worsen the effects. For instance, Raaum and Roed (2006) study whether the unemployment rate during the ages of 16 to 19 affect individuals' labor-market outcomes as well as schooling choices in Norway. They find that a business-cycle slump occurring at ages 16 and 19 raises prime-age unemployment rate by as much as 1 or 2 percentage points, but has no

[3] Oreopoulos *et al.* (2012) find that "earnings adjustment process is characterized initially by increased mobility across employers and industries and improvements in the characteristics of the average employer."
[4] Unfortunately, our data does not allow us to estimate firm's average payroll or median wage, hence precluding us from directly testing the effect of entry labor-market conditions on firm quality.

effect on individuals' choice of educational attainment. Brunner and Kuhn (2014) study the careers of workers with vocational training in Austria. They find a robust negative effect of the initial unemployment rate on starting wages. Even though this effect fades away after several years, the authors estimate that entering the labor force when unemployment is high lowers the present discounted value of lifetime earnings of these workers by 15% compared to entering in average conditions. For the case of Sweden, Kwon *et al.* (2010) find similar results to the Austrian study, but for all education levels. When analyzing their results by completed education, these authors find that the negative effects on job-market entry in Sweden are similar across all education groups.

The remainder of this paper is organized as follows. The next section discusses the Spanish labor market, and Section III presents the data and empirical strategy. Section IV presents the results. Sections V and VI present the dynamic specification and mobility results, respectively, before concluding in Section VI.

II. The Spanish Labor Market

Permanent versus Fixed-Term Contracts

With unemployment over 20% in the early 1980s, the Spanish government legalized the use of fixed-term contracts for jobs lasting between 1 day and 3 years in 1984. The objective of the reform was to add flexibility and promote employment in a rigid labor market. Such flexibility came from the fact that, in contrast with permanent contracts, fixed-term contracts have much lower dismissal costs and their termination cannot be appealed in court. In particular, if a fixed-term contract worker is laid-off, he receives a severance payment of 12-day wages per year of service (with a ceiling of 36 months) as opposed to the 45-day wages per year of service paid to workers with permanent contracts (with a 42-month ceiling).[5] Moreover, if the employer waits for the fixed-term contract to expire, there is no cost to let the employee go. Even though by law, fixed-term contracts can only be used for up to a maximum of three consecutive years within the same firm, this was not strictly enforced until after 2008.

One of the most visible consequences of the 1984 reform is that, since then, the vast majority of workers in Spain are first hired under a fixed-term contract and, eventually (often after the legal-time limit of consecutive fixed-term contracts has been reached), they are promoted to a permanent one. Consequently, the conversion rate of

[5] Severance payments are lower for fixed-term- than permanent-contract layoffs not only because the amount paid per year worked is lower, but also because the average tenure is also considerably lower.

fixed-term contracts into permanent ones is low (18% in 1987) and has decreased over time (to 5% in 1996) as estimated by Güell and Petrongolo (2007).[6] All this implies that the transition from a fixed-term to a permanent contract is often a quite lengthy one. For example, Estrada *et al.* (2009) estimate that as many as 40% of fixed-term contract workers still hold such type of contract ten years after having entered the labor market.[7] To put it differently, Spanish employers use fixed-term contracts more as a flexibility device to adjust employment in the face of adverse shocks than to fill-in jobs of a temporary nature or as stepping stones towards permanent jobs. Not surprisingly, once a worker finally gets a permanent contract, he will try to maintain it at all costs, reducing his or her willingness to move to a different job. This prevents smooth wage renegotiations based on current labor market condition (Beaudry and DiNardo, 1991). Amuedo-Dorantes and Serrano-Padial (2007) estimate that the annual turnover rates among permanent contract workers are low (in the order of 10%) and most of the observed transitions are into a new permanent contract or retirement. In contrast, fixed-term contract workers' yearly turnover rates are very high (in the range of 34% to 66%) and, in this case, workers transition to a new fixed-term contract job or become unemployed.

Much evidence indicates that the labor market of fixed-term contract workers in Spain is a secondary one. For instance, fixed-term contracts impose penalties to workers in the form of forgone experience, delayed wage growth and higher odds of unemployment (Amuedo-Dorantes and Serrano-Padial, 2007). Several authors have found that the likelihood of transitioning into unemployment is considerably higher among workers with fixed-term contracts (Güell and Petrongolo, 2007; García-Ferreira and Villanueva, 2007; and Barceló and Villanueva, 2010). As such, Barceló and Villanueva (2010), estimate that for a given year the probability of entering an unemployment spell is 8 percentage points higher for workers with fixed-term contracts (10%) than those with permanent ones (2%). Amuedo-Dorantes (2000) also finds that fixed-term contract work spells in Spain are unlikely to end in permanent jobs, regardless of workers' tenure. Finally, the probability of receiving free or subsided on-the-job

[6] For example, in our sample the one-year transition probability from a temporary to a permanent contract during the first five years of potential experience is 14% for cohorts that graduated on or before 1985 and just 10% for those that graduated after 1985. However, there is substantial variation across education groups with lower educated individuals facing a lower transition probability (7% for all cohorts of high-school graduates) than more educated ones (14% for individuals with vocational training and 18% for college graduates).

[7] In our sample, 51% of individuals are either working under a fixed-term contract or out-of-work ten years after having finished their education, with the percentage being much higher for high-school graduates (57%) compared to individuals with vocational training (44%) and college graduates (36%)

training is 22% lower for workers under fixed-term contracts than for workers under permanent contracts (Dolado *et al.,* 1999); and fixed-term contract employment increases work accidents by 300% (Jimeno and Toharia, 1996).

Over the years, fixed-term contracts have contributed to employment growth and declining unemployment during economic expansions in Spain. After its inception in 1984, fixed-term employment soared, reaching a persistent one third of the Spanish labor force in the early 1990s. However, with the Great Recession and unemployment rate climbing from 8% to over 25% within five years, the share of fixed-term employment dropped to 23%, the lowest level since its inception (shown in Figure 1). During slowdowns, most employment adjustments take place via the termination of fixed-term contracts and, hence, concentrate on the young (Bentolila *et al.*, 2008). There are two reasons for this adjustment process: a rigid wage-setting process, which prevents firms from adjusting the cost of the employed workforce, and a near-zero cost of dismissing fixed-term contract workers. Since fixed-term jobs are usually of lower productivity, this vast destruction of jobs leads to the well-documented countercyclical evolution of labor productivity in Spain, according to which labor productivity and the average quality of jobs increase during recessions and decrease during expansions (Maroto and Cuadrado, 2013).

Rigid Labor-Market Institutions

A further concern with the Spanish labor market is that the lower part of the wage distribution is compressed by collective bargaining. Collective agreements, which in Spain cover about 90% of private-sector wage and salary workers, are bargained at the province/industry level, with a very low share of firm-level agreements. Collective bargaining in Spain sets "entry minimum wages" above the legal minimum wage inflating the lower part of the wage distribution and resulting in relatively high earnings for young workers and the least qualified ones. This leads to high unemployment rates for these two groups of workers (Felgueroso, 2010). Izquierdo *et al.* (2004) and Bentolila *et al.* (2010) have found that this intermediate level of decentralization provide a low association of wages and labor conditions to firms' individual performance. Moreover, Messina *et al.* (2010) find that both high inertia of wages and real downward wage rigidities in Spain are due to the strong wage indexation of wages to inflation negotiated in collective agreements.

Most recently, Font, Izquierdo, and Puente (2014) find that real wages are very weakly pro-cyclical in Spain in all stages of the business cycle. According to their estimates, an increase (decrease) of 1 percentage point in the unemployment rate is associated with a real wage decrease (increase) of between 0.24 (0.48) percentage points. This is extremely low compared to the US where wage to unemployment semi-elasticities lie above 1, or other European countries with the semi-elasticity close to 2 in the UK or above 1 in Germany, Italy or Portugal (Devereux and Hart, 2006; and Pissarides, 2009). Most interestingly for our paper, Font, Izquierdo, and Puente (2015) find that downward wage rigidities are important in the Spanish context as wage cyclicality is much lower in recessions than in expansions. In particular, they find that the level of the unemployment rate appears to be relevant only in expansionary periods. According to these authors, these important asymmetries in wage formation are likely the result of having both a highly segmented labor market and sectorial levels of negotiations. To put it differently, the Spanish collective bargaining system overly protects the incumbents or insiders, without worrying about job access for those unemployed, leading to *both* high levels of unemployment and wage growth during the first stages of economic recovery, therefore delaying the decrease in unemployment (Font, Izquierdo, and Puente, 2014).

Traditional Society

Most young individuals in Spain study and later live near the parental household. Whether this is the result of tradition or out of economic necessity, the fact is that family ties in Spain are very strong, which often implies very low geographical mobility during the life of an individual (Jimeno and Bentolila, 1998). For example, according to a recent report by Eurostat, Spanish young men do not leave the parental household until they are 30 years old, on average, compared to 20 years old in Sweden, Denmark or Finland (Eurostat, 2015). Even more striking, according to research by the Spanish Council for Youth in 2013, 93% of individuals aged 16 to 24 years old lived with their parents.[8] This includes college students who often choose the college that is nearest to their parents' home. The lack of affordable housing and job opportunities for youth is another reason why so many young individuals delay leaving their parents' nest and choose a college near their parents' home. According to the 2015 Eurostat study, only 10% of Spanish 20-

[8] Available at www.cje.org

to 24-year old young men worked while studying compared to more than 50% in Switzerland, Germany, the Netherlands and the Nordic countries.

III. Data and Empirical Specification

We use data from two different sources: social security data from the 2008 Continuous Sample of Working Histories (hereafter CSWH), and survey data from the 1980 to 2008 Spanish Labor Force Survey.

The 2008 Continuous Sample of Working Histories

The 2008 CSWH is a 4% non-stratified random sample of *all* individuals who were either working in 2008, and hence, contributing to the Social Security, or receiving Social Security payments, which includes unemployment benefits, disability, survivor pension, and parental leave.[9] As long as the individual receives unemployment benefits (or some other Social Security transfer), he or she is in the CSWH. In Spain, there are two types of unemployment benefits: Unemployment Insurance (UI) and Unemployment Assistance (UA). To be entitled to UI benefits one has to become involuntarily unemployed and have worked for at least 12 months over the 72-month period prior to unemployment. UI benefits last for a period of at least four months extendable in two-monthly periods up to a maximum of two years, depending on the worker's employment record.[10] Once UI benefits expire, workers are entitled to UA. UA is a non-contributory benefit targeted to those who no longer qualify for UI benefits due to the duration of unemployment or lack of contributions. To determine UA payments, the beneficiary's per capita family income is set to 75% of the Statutory Minimum Wage. The fact that the adult male labor-market participation rate is high and that the system of unemployment benefits in Spain is quite generous implies that our sample will suffer little from attrition. Furthermore, as explained in the *Sample Selection* sub-section below, to minimize attrition because workers may drop out of the labor force, we focus on male workers *and* use pre-Great Recession data.

The 2008 CSWH gives information of the *complete* work history of individuals sampled in 2008 back to when they first entered the labor force. More specifically, the

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[9] The random sample is selected by Social Security and shared with researchers upon request.

[10] A worker with 12 to 18 months of employment within the last 6 years is entitled to 4 months of UI benefits. If the worker has worked for a period ranging between 19 and 24 months within the last 6 years, he is entitled to 6 months of UI benefits, and so on. This implies that the UI benefit entitlement in Spain is about 30% of the months employed during the last 6 years with a maximum of 24 months.

2008 CSWH provides detailed information on: (1) socio-demographic characteristics of the worker (such as sex, nationality, province of residence at the time of labor-market entry); (2) the worker's career information (such as the dates the employment spell started and ended, monthly earnings, hours worked, type of contract, and occupation); and (3) employer's information (such as industry, public versus private sector, the number of workers in the firm, and the location).[11] Using information in the CSWH, we can calculate experience and tenure. We construct annual earnings by averaging out monthly earnings for months 1, 4, 7 and 10 for each year.[12] Annual earnings are top coded at €42,000 euros (in 2008 dollars).[13] Annual and monthly earnings are deflated using the 2008 Spanish CPI.

Matching the Social Security records with the 2008 Spanish Municipal Registry of Inhabitants, we are able to retrieve the individual's education level. We conduct our analysis separately for the following three groups according to their completed education level: (1) high-school graduates; (2) individuals with more than a high-school degree but less than college; and (3) college graduates. The second group comprises individuals with technical degrees below the college level (in Spanish, *formación profesional*) or with associate degrees (in Spanish, *diplomaturas*). These technical degrees focus on teaching a profession, such as cook, electrician, nurse, or plumber, and frequently include some internship in the field of study. Hence, they resemble vocational training such as in Germany or Switzerland, although they represent less employer commitment in the training component. Associate degrees are three-year long and have a more practical orientation than the five-year university degrees (in Spanish, *licenciaturas*) of individuals in the third group. Thereafter, we call this group "workers with vocational training".

Since we do not directly observe the year of graduation (only their highest educational degree), we impute it using information on the date of birth, the highest educational level completed, and the most common graduate age for each degree, as reported by the Spanish National Statistics Institute (*INE*). The most common graduation age is 18 for a high-school degree, 20 to 22 for degrees above high school but less than

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[11] Hours worked are usual weekly hours as reported by the employer in the job contract that the employer and the employee sign. This information is reported to the Social Security and included as a variable in the CSWH dataset.
[12] As explained below, this is done for computational efficiency given the size of the original sample.
[13] Top coding affects 5% of individuals in our sample (0.6% of high-school graduates, and 13.6% of college graduates). We found very limited effects of entry labor-market conditions on the probability of being top-coded by highest education attainment (results available from authors upon request).

college, and 23 for a college degree. This imputation technique is common in this literature when graduation year is not available (see Altonji *et al.*, *2016*).[14]

Because we use predicted year of graduation (based on year of birth and typical degree duration) instead of actual age of labor market entry, bias due to choice of entry is less of a concern in our analysis. However, measurement error may be an issue if it is correlated with the business cycle. While a priori, there is no reason why this would be the case, if it were, it is likely that individuals graduating during a recession may be less eager to graduate on time than those graduating during an expansion.[15] In this case, our annual earnings estimates would capture the full effect of graduating during a recession since we would count the extra time in school as non-work. In the case of wages conditional on working, however, our estimates for the initial years would be a *lower bound* since those graduating during bad times would delay their entry in the labor market and enter when the job opportunities have improved. Moreover, it is important to highlight that measurement error will lead to imprecise matching to the true unemployment rate at graduation, and hence lead to attenuation bias in the results.

A different but related issue is whether individuals expand their studies and get a higher degree because of finishing their first degree during a recession. Since the analysis is by highest education completed, this may affect our estimates only if the unobserved component or ability of those who act in this manner is different (higher or lower) to that of those who finish a given degree independently of the economic situation. Nonetheless, Raaum and Roed (2006) for Norway, and Oreopoulos *et al.* (2012) for Canada, find no evidence that individuals expand their studies during recessions.[16] Unfortunately, such analysis is not possible with our data, as we do not observe the year of graduation. Using the Spanish Labor Force Survey data, we estimated the effect of the business cycle on high-school completion rates and enrollment rates in vocational training and college, using a specification that follows Hershbein (2012). Estimates in Appendix Table A.1, show that, once we control for province and year fixed effects, there is no effect of labor-

[14] Using three alternative datasets, we have explored whether the imputation of the year of graduation is reasonable. With the Survey of Educational Transitions and Employability, we estimate that 89% of individuals with at most a high-school degree graduated by age 19. With the Spanish Labor Force Survey, the estimate is 86%. With Survey of Educational Transitions and Employability, we estimate that 92% of individuals with at most vocational training graduated by age 22. Finally, using Statistical Report of University Education, we estimate that 40% of those with a college degree graduated on time (age 23) and 70% by age 25.

[15] Raaum and Roed (2006) do not find evidence that unfavorable entry conditions cause students to delay graduation among 16- to 19-year olds in Norway.

[16] Kahn (2010) finds that while the national unemployment rate at time of college graduation is positively correlated with educational attainment, the state unemployment rate is not.

market entry conditions on high-school completion or enrollment in vocational training and college (as shown in column 3).

The Spanish Labor Force Survey

Using data from the Spanish Labor Force Survey, we measure province unemployment rates for each year of graduation and in each province of initial employment, U_{cp0}. We follow Oreopoulos *et al.* (2012) and use the province unemployment rate as the measure of economic conditions. Our results are robust to using state or national unemployment rates, as explained in the *Robustness Section* below. However, we prefer the more disaggregated measure, as there are 50 provinces in Spain, compared to only 17 states, hence adding useful geographic variation to supplement the time variation.[17]

In the CSWH, we do *not* observe college location, but instead the province of residence once the individual first joints the labor market. Hence, to define our economic conditions at labor-market entry, we match economic conditions to the province of the first job. Note that this is a concern only if individuals move to a province different from the one of first labor-market entry in response to adverse economic conditions. As explained earlier, migration within Spain is traditionally low. For instance, in our dataset only 5% of our sample migrates to another province during the first five years of potential experience. This percentage increases to 14% and 21% during the first 10 and 15 years of potential experience, respectively. In the *Results Section*, we first show that our findings are robust to using state level unemployment rates (as opposed to province level unemployment rates). In addition, we also show that our results are robust to *only* keeping individuals who never leave their original province. Finally, we estimate the effects of entry labor-market conditions on mobility in Section VI.

Sample Selection

We focus our analysis on individuals entering the labor market between 1980 and 1992, implying that in 2008 they are between 36 and 52 years old. The reason our youngest cohort is the one entering the labor market in 1992 is because the highest education variable was last updated in 1996. Hence, we want to prevent miscoding of the completed highest education level. We restrict our analysis to male wage and salary workers. The

[17] Kahn (2010) uses both an annual average of national monthly unemployment rates and the state unemployment rate, and Altonji *et al*. (*2016*) use census-division unemployment rate in the year of college graduation.

reason we focus on male workers is that female labor force participation has traditionally been low in Spain and drops after women's first birth.[18] As most 36- to 52-year old males are working or receiving UI (or other) Social Security benefits in Spain in 2008, attrition because the individual has dropped out of the labor force is very unlikely in our sample. Nonetheless, we decided to work with the 2008 CSWH wave as opposed to some more recent waves as the Great Recession started to shred jobs in Spain beginning the last quarter of 2008. It is important to highlight that 2008 was an excellent year in terms of employment in Spain as the Spanish economic activity had been growing at more than 3% annually since the year 2000. More specifically the unemployment rate in 2008 was below 9%, a record low for Spanish standards.[19] Finally, we also excluded immigrants from our analysis because they represented less than 1% of the Spanish labor force prior to the turn of the century.[20] Using the Spanish Labor Force Survey, we estimate that immigrant males represent 0.15 and 0.23% of the entering cohorts with at most a high-school degree and more than a high-school degree, respectively.[21]

For each individual in our dataset, we have monthly information on their work history since the year after "imputed" graduation and until 2008, covering a minimum of 16 years and a maximum of 28 years of work history after graduation.[22] Because the resulting dataset would have been huge, to reduce sample size and increase computationally efficiency, we transform the monthly to quarterly data by keeping only the last month of each quarter. This leaves us with a dataset comprising 4,878,043 quarterly individual level observations, 2,152,300 (or 44%) of which are high-school

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[18] Using data from the first half of the 1990s, Gutierrez-Domenech (2005), estimates that the proportion of women in Spain with paid work falls from 43% to 33% after their first birth and remains around 35% ten years after they gave birth.

[19] Using the Labor Force Survey, we explored whether attrition due to labor market inactivity (defined as not working and not looking for a job) was an issue in our sample of males aged 36- to 52-year olds in 2008 by education level. We found that the average inactivity rate for males within this age range is very low (in the order of 4% or lower). Furthermore, inactivity rates of men who entered the labor market during bad times are only slightly higher than inactivity rates of men who joined the labor market in good times, with the difference between the two groups being statistically significant only for individuals with more than a high-school degree (3.5% versus 2.6%).

[20] Even if we wanted to include immigrants, the CSWH lacks information on their year of arrival to Spain, their accumulated work experience after having completed their studies and prior to their arrival to Spain, or their work experience as undocumented after arrival to Spain. Moreover, we have no concise information on the correspondence between immigrants' reported education and the degree they completed in their country of origin.

[21] Another reason to exclude immigrants and women from the analysis is to limit sensitivity of our results to external factors such as discrimination (for immigrants) and childbearing and discrimination (for women).

[22] Others have comparable size and observation periods. Raaum and Roed (2006) observe 19 to 22 years of data. Kwon *et al.* (2010) observe about 20 years of data. Oreopoulos *et al.* (2012) have information covering the first 17 years of labor market experiences. Brunner and Kuhn (2014) observe 22 years of data.

graduates, 1,905,192 (or 39%) of which have vocational training, and 820,551 (or 17%) have a college degree.[23]

The period 1980 to 1992 includes a period of a deep recession (between 1984 and 1987) followed by an economic expansion (between 1988 and 1992) as shown in Figure 1. During the early 1980s recession, the unemployment rate soared from 11% in 1980 to 22% in 1985, and then decreased to below 16% in 1992. In the context of our analysis, we exploit variation in entry conditions across both time and province. Interestingly, we observe greater dispersion in the across province variation with the unemployment rate being as low as 2% in *Lleida* in 1980, and as high as 37% in *Cádiz* in 1991. Of all the unemployment-rates variation, 67% comes from across provinces and 33% across time. Hence, individuals graduating between 1980 to 1992 experienced very different labor market conditions at the time of entry.

Empirical Specification

Our objective is to estimate the impacts of labor-market entry conditions on subsequent labor-market outcomes. Identification in this analysis comes from exploiting the variation in unemployment rates at the province-year level in Spain for the period 1980 to 1992 and across 50 provinces.

Following Oreopoulos *et al.* (2012), we collapse the quarterly data at the level of education, province of initial employment, graduation cohort, potential experience and year. We then construct two different collapsed datasets to estimate the effect of labor-market entry conditions on two different groups of outcomes. First, to estimate the impact of labor-market entry conditions on employment and annual earnings, we assign zeroes to the left-hand side variable each time we observe the individual not working. Second, we construct another dataset with only those individuals with wages greater than €0 euros, and estimate the impact of labor-market entry conditions on monthly earnings, hours worked, and the probability of having a permanent contract (conditional on working). The first collapsed dataset has 43,859 observations or cells and the second one has 42,816 observations.[24]

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[23] Studies with administrative data contrast with those using the Panel Study of Income Dynamics (Devereux, 2002) and the National Longitudinal Studies of Youth (Gardecki and Neumark, 1998 and Kahn, 2010), which have considerably smaller sample sizes.

[24] The difference between the two samples is due to some cells having zero observations.

Since we have originally 4,878,043 quarterly individual-level observations, we can estimate average outcomes in each of those cells with precision.[25] We work with the collapsed datasets made of the cell means weighted by the corresponding cell sizes. For each highest education level, Appendix Table A.2 presents the number of individual-quarter observations by initial province unemployment rate for each graduation year, U_{cp0}, and potential experience. Appendix Table A.2 shows that sample sizes at all levels of unemployment and potential experience for each educational level are substantial.

Table 1 reports summary statistics of the key variables, calculated by assigning equal weight to each cohort-potential-experience-year cell. Panel A summarizes variables for the whole sample (including individuals with zero earnings). Both the likelihood of working and average annual earnings increase with education. The average probability of employment is 66% for high-school graduates, 69% for workers with vocational training, and 71% for college graduates. Looking now at average annual earnings, Table 1 shows that high-school graduates earn an average of €11,473.92 per year, workers with vocational training earn €15,544.68 per year, and college graduates earn €20,825.64 per year. The average graduation year is 1985 and the average year of an earning's observation is 1997.

In the sample with positive earnings (shown in Panel B in Table 1), average hours worked are 40 hours per week regardless of the education level, and the rate of part-time work is extremely low (between 1% and 2%). This result is not new. As explained by Fernández-Kranz and Rodríguez-Planas (2011), the incidence of part-time work among males in Spain is one of the lowest in OECD countries. We do observe some variation in the likelihood of working under a permanent contract by highest education level. Indeed, the probability of working under a permanent contract increases from 59% to 73% to 79% for high-school graduates, workers with vocational training, and college graduates, respectively.

To assess the impact of initial conditions on labor market outcomes, we begin estimating equation (1) separately by education level.

[25] As explained by Oreopoulos *et al.* (2012) in footnote 3, the small samples sizes used in studies using survey data, "do not allow controlling for cohort, state, and year effects in a flexible way, controlling for persistent correlated labor market conditions, or studying other career outcomes than wages with a sufficient degree of precision." Because of our large sample size, we are able to do all these three types of analyses.

$$Y_{cpt} = \beta_1 + \beta_2 U_{cp0} + \beta_3 U_{cp0} PE_{cpt} + \beta_4 U_{cp0} PE_{cpt}^2 + \beta_5 PE_{cpt} + \beta_6 PE_{cpt}^2 + \phi p + \eta t + gc + \epsilon_{cpt} \tag{1}$$

where Y_{cpt} is the cell mean of the labor-market outcome of interest measured at the level of graduation cohort (c), initial province of employment (p), and calendar year (t) (weighted by the corresponding cell sizes). U_{cp0} stands for the province unemployment rate when the individual's employment history begins. We standardize U_{cp0} and, therefore, the coefficients of interest show the effect of one standard deviation of the province unemployment rate the year after graduation.[26] PE_{cpt} is the cell mean of potential years of experience calculated as calendar year minus (year of graduation+1) at the level of graduation cohort (c), initial province of residence (p), and calendar year (t). Besides these potential experience controls, all models include province of residence when employment history begins (ϕ_p), current calendar year (η_t), and imputed-graduation-cohort (g_c) fixed effects. Given the presence of potential experience, year, cohort and province fixed effects, the coefficients of interest β_2, β_3 and β_4 measure deviations from the average experience profile that are due to graduating in a bad year (high U_{cp0}) or in a good year (low U_{cp0}). Hence, our estimation results show not only the average effect of initial conditions but also the persistence of those effects throughout the experience profile.[27] To account for group-specific error components, we cluster standard errors at the cohort-province level.

We call equation (1) the *full-effects* specification. In this specification, we allow the dependent variable (Y_{cpt}) to be affected by the initial unemployment rate (U_{cp0}) *and* by the sequence of unemployment rates correlated with U_{cp0}. Individuals graduating in a bad year will face not only a high rate of unemployment the year of graduation but also a particular sequence of unemployment rates the years that follow graduation which will be different from the sequence faced by individuals graduating in a good year. As we do not control for the contemporaneous rate of unemployment in equation (1), the coefficients β_2, β_3 and β_4 will capture the effect of the initial unemployment rate and also the effect of the successive rates of unemployment that are correlated with U_{cp0}. To

[26] We normalize U_{cp0} using the sample period mean, 0.16, and dividing it by the sample period standard deviation, 0.07.

[27] As graduation cohort, calendar time, and potential experience are collinear with each other, identification is only possible if one makes additional restriction on cohort effects. Following Oreopoulos *et al*. (2012), we dropped one additional cohort effect from the regression.

control for the contemporaneous rate of unemployment (U_{cpt}), we also estimate the following equation:

$$Y_{cpt} = \beta_1 + \beta_2 U_{cp0} + \beta_3 U_{cp0} PE_{cpt} + \beta_4 U_{cp}\ PE_{cpt}^2 + \beta_5 PE_{cpt} + \beta_6 PE_{cpt}^2 + \beta_7 U_{cpt} + \beta_8 U_{cpt} PE_{cpt} + \beta_9 U_{cpt} PE_{cpt}^2 + \phi p + \eta t + gc + \epsilon_{cpt} \tag{2}$$

In equation (2), we control for the contemporaneous rate of unemployment (U_{cpt}) as well as for the different impact the contemporaneous unemployment rate (U_{cpt}) has at each level of potential experience through the interaction terms $U_{cpt}PE_{cpt}$ and $U_{cpy}PE^2_{cpt}$. We call equation (2) the *dynamic specification.* Now, the coefficients of interest β_2, β_3 and β_4 measure deviations from the average experience profile that are due to graduating in a bad year or in a good year *net* of the effect of the future sequence of unemployment rates (that are correlated with the initial conditions) on the experience profile. This equation is similar in spirit to the dynamic specification in Oreopoulos *et al.* (2012).

IV. Full Effects of Graduating During a Recession

Baseline Data

Figure 2 shows the general experience profiles in annual earnings for our baseline Spanish data by highest education level. For high-school graduates and individuals with vocational training, we observe sharp and sizeable differences in starting earnings across graduation cohorts, with those entering between 1981 and 1987 (for high-school graduates) and 1983 and 1986 (for those with vocational training) having lower annual earnings. For college graduates, we also observe fluctuations of starting earnings across graduation cohorts, albeit smoother. Interestingly, Figure 2 shows a clear pattern of convergence for all education groups, suggesting that initial differences in starting conditions tend to fade over time and become negligible for all groups around 7 years after entry. In what follows, we will analyze the mechanisms that explain these earnings gaps and the convergence patterns for each education group.

Effects of Labor-Market Entry on Annual Earnings

Table 2 shows estimates of the effects of labor-market entry on annual earnings by highest education level. More specifically, we use the estimates from equation (1) to calculate the effects of an eight percentage-points increase in unemployment rate, which corresponds to the average shift from a boom to a recession in our sample, on annual earnings. Panel A shows the average effect of an 8 percentage-points increase in the province unemployment rate at labor-market entry at different years of potential experience (at 1, 3, 5, 7, and 10 years of potential experience), and Panel B shows the cumulated effect over the first five and the first ten years of potential experience in the labor market.

Focusing first on the average *full* effect of an eight percentage-point increase on earnings over a ten-year period (Panel B), we observe that the penalty is *lower* for workers with a college degree. College graduates receive, on average, 6.4% (only statistically significant at the 10% level) lower earnings if they graduated during a recession instead of a boom (shown in column 3 in Panel B). However, the effect increases by half or doubles (-9.6% and -12.5%) and is statistically significantly different from zero at the 1% level for workers with a high-school degree or with vocational training, respectively (shown in columns 1 and 2 in Panel B).

Panel A in Table 2 shows the shifts due to experience profiles. They show that the negative shock is also more persistent for individuals without a college degree. The effect of the negative shock at labor-market entry is a decrease in earnings of 25.1% and 24.9% during the first year in the labor market for high-school graduates and workers with vocational training, respectively (shown in columns 1 and 2 in panel A). The effect of the shock decreases to 17.4% and 18.5% at experience years 3, and further to 10.3% and 12.8% at experience years 5, respectively. For both groups, the negative effect of the shock on earnings fades away at experience years 10. For college graduates, the full effect of an increase of 8 percentage points in the unemployment rate fades away within 7 years (shown in column 3 in panel A). College graduates experience a 13% decrease in earnings during the first year in the labor market, a 9.3% decrease at experience years 3, and a 6.3% decrease at experience years 5. All of these effects are statistically significantly different from zero at the 1%, 5%, and 10% level, respectively. After 7 years, the effect becomes smaller and is no longer statistically significantly different from zero.

Robustness Checks

Our main results hold to a battery of robustness checks as shown in Table 3. Each panel shows the results for each highest completed education level. Column 1 in Table 3 shows our preferred estimates (also shown in Table 2). Columns 2 and 3 re-estimate equation

(1) using state and national unemployment rates, respectively. While differences between our preferred estimates and those estimated with state unemployment rates are small, we lose precision when we estimate the model with national unemployment rates for vocational-training and college graduates (but not for high-school graduates). This is not surprising given that "*national estimates may be more affected by measurement error problems due to aggregating across local labor market shocks*" (Oreopoulos *et al.*, 2012). Nonetheless, comparing state and national unemployment rate estimates to the province ones, we continue to find a greater penalty for non-college graduates than college graduates. The major difference is that the national unemployment specification displays less persistence than the specification with province or state unemployment rates for workers with a vocational degree (panel B) and college graduates (panel C). As the specification with province unemployment rates is likely to measure entry labor-market conditions more accurately, we prefer this specification, which is also the preferred specification in the literature (Oreopoulos *et al.*, 2012).

As explained earlier, regional mobility is lower in Spain than in the United States. For instance, the majority of males in our sample do not move out of their provinces of labor-market entry. By highest education level, only 4%, 5%, and 9% of high-school graduates, individuals with vocational training and college graduates, respectively, have moved out of their provinces of labor-market entry 5 years after they entered the labor market. After 10 years, 12%, 15%, and 21% have moved out of their provinces of labor-market entry, and after 15 years, 19%, 22% and 27% have moved out of their provinces of labor-market entry, respectively. Nonetheless, to address potential concerns that province of residence when graduating may inaccurately capture local labor-market conditions if there are economic differences between residence and college locations, we re-estimate our preferred specification using *only* those individuals who never switched provinces. Estimates for non-movers (shown in column 4 in Table 3) are very similar to our main results. This finding resembles that of Oreopoulos *et al.* (2012).

Given the high degree of regulation in the Spanish labor market, one may wonder how robust are these results to using the non-employment rate instead of the unemployment rate, as the former may well be a better measure of labor-market conditions. Column 5 in Table 3 presents estimates using province-level non-employment rate at labor-market entry. Overall, our main results do not change much, which is not that surprising as the correlation between the two rates is high (in the order of 70%), and both indicators move quite closely over the business cycle. While the results are practically identical for non-college graduates, the negative effects of graduating during a recession are slightly higher and more persistent for college graduates.

We already discussed and showed (in Data Section III and Appendix Table A.1) that postponing high-school completion or enrolling in vocational training or college as a response to the high-school graduation unemployment rate is unlikely in Spain. Nonetheless, if students postponed graduation due to poor labor-market conditions, the relevant unemployment rate would not be that of entry, but that of one or two years earlier. Column 6 in Table 3 explores whether our results are robust to using the unemployment rate two years prior to expected graduation. Results are similar to those in our baseline specification, corroborating that endogenous timing of degree completion is not a concern.

We have followed Altonji *et al.* (2016) and Kahn (2010), and used a second-order polynomial in experience. To address concerns on the sensitivity of the results to an alternative specification for experience, we have re-estimated our main specification using two-year buckets instead (results shown in Appendix Figure A.1). Overall, the main results remain, although the second-order polynomial overestimates slightly the loss during the first couple of years after entry and underestimates them after the tenth year of potential experience. As our paper focuses on the first ten years of potential experience, the latter difference is less relevant.

Effects of Labor-Market Entry on Other Outcomes

Employment probability. Table 4 shows estimates of the effects of an eight percentage-point increase in the entry unemployment rate on the likelihood of working by highest education level. For non-college graduates, Table 4 reveals that graduating during a recession has large and persistent effects on employment. More specifically, the negative effect on employment is quite large during the first year in the market (-0.251 with a standard error of 0.026, and -0.231 with a standard error of 0.032 for high-school and vocational-training graduates, respectively); and remains close to a statistically significant and negative -0.10 for both types of graduates in the fifth year in the market. While the effects are small and no longer statistically significant by the 7^{th} (10^{th}) year of potential experience for high-school graduates (workers with vocational training), the average cumulated impact over the first 10 years is a 10% decline in the likelihood of employment for either group.

In contrast, the average cumulated 10-year entry effect on college graduates' employment is half the size than that observed for non-graduates, and not statistically significant. While entering the labor market during a bad shock reduces college graduates' likelihood of employment by 11% during the first year in the market, this negative effect dies off within the first 5 years of experience. We proceed to explore whether labor-market entry conditions affect which segment of the labor market individuals have access to.

Likelihood of Having a Permanent Contract. Columns 2, 5, and 8 explore the effects of graduating during a recession on the likelihood of having a permanent contract by years of potential experience. Interestingly, we observe a large and persistent negative effect especially for college graduates and, to a lower extent, for workers with vocational training. Among college graduates, the cumulated 10-year effect is a 9% drop in the likelihood of working under a permanent contract. This negative effect is half the size for workers with vocational training, and inexistent for workers with a high-school degree.

At labor-market entry, we do not observe a negative effect of graduating in a recession on the likelihood of having a permanent contract for high-school graduates nor workers with vocational training. In fact, it is not until the 7^{th} (for high-school graduates) and 5^{th} year (for vocational-training graduates) of potential experience that the negative effect of entering during a recession kicks in. For college graduates, we observe a negative effect (albeit not statistically significant) from the 1^{st} year on. These results suggest that the detrimental effects of graduating in a recession in a segmented and rigid labor market mediate more through the type of contract for those workers with greater human capital and, a priori, easier access to the primary segment of the labor market. In contrast, lower educated workers who graduate during a recession suffer mainly an employment loss as the recession shreds many low-skilled fixed-term contract jobs.

Monthly Wages. We turn now to the effects of graduating during a recession on monthly wages. As explained earlier, these estimates (shown in Table 4 columns 3, 6, and 9) *only* use employed individuals.[28] They give guidance on whether differences in monthly wages conditional on working drive any of the observed negative and persistent effects

[28] We also analyzed the effects on hours worked, but found no effects of labor-market entry conditions on this outcome.

of entering the labor market during a recession. These results need to be taken with caution because of sample selection as labor-market entry conditions affect employment. If only the most able workers find jobs when entering the labor market in a recession, our estimates conditional on employment would underestimate the effect of early economic conditions on monthly wages, especially during the first years in the labor market.[29]

Not surprisingly given the wage rigidity discussed in Section II, we only find modest negative impacts of entry labor-market conditions on monthly wages for workers with vocational training (and small but not statistically significant effects for college graduates). Hence, employment drives most of the effects of entry-labor market conditions found in Table 2 (as seen in Table 4, columns 1, 4, and 7).[30] These results are consistent with similar findings from studies analyzing rigid labor markets in Norway (Raaum and Roed, 2006), Japan (Genda *et al*., 2010), and Germany (Stevens, 2008).

For high-school graduates, we observe a small, but statistically significant, and *positive* effect of unemployment on monthly wages (columns 3 in Table 4). This is consistent with the idea that during recessions there is a substantial shredding of low-skilled jobs with "bad" jobs disappearing first and only "good" jobs remaining. Therefore, those observed working do so at higher than average earnings. Appendix Table A.3 gives further supporting evidence of this. In column 4, we control for job characteristics that proxy for the quality of the job (namely, industry controls, blue- versus white-collar indicator and firm-size controls), and, as a result, the positive ten-years average wage effect decreases one fourth for high-school graduates. We also observe this pattern among college graduates. In this case, controlling for job characteristics increases the average ten-year penalty from -0.016 (standard error of 0.012) to a statistically significant -0.024 (standard error of 0.009).

Discussion

The larger losses for high-school graduates compared to college graduates contrast sharply with findings by Hershbein (2012), and Genda *et al.* (2010) for the US. Hershbein (2012) finds no effects of labor-market entry conditions on employment or weeks worked

[29] Note that, even though we find that the unemployment rate affects participation, sample selection is not an issue in our estimates on annual earnings in Table 2 as all workers are included in the sample regardless of whether they work or not (and hence have 0 earnings).

[30] For example, looking at the ten-year average effect for college graduates, and considering that only 51% of college graduates in our sample work during the first years of potential experience, the small negative wage effect (-0.016) in Table 3 represents only 12% of the overall income effect found in Table 2, with the remaining 88% explained by differences in the probability of working for different cohorts.

of high-school graduates in the US, and some persistent negative effects on hours worked. He estimates that a severe recession (raising the unemployment rate by 3 to 4 percentage points) reduces high-school graduates' wages by 7% over the first four years (statistically significant at the 5-percent level). The negative wage effect fades out within 6 years, and is considerably smaller than that found for college graduates in North America. In this case, the initial earnings losses for college graduates range between 2.5% and 6% for a 4 percentage-point increase in the unemployment rate. Although the effect eventually fades away, the earnings reductions add up to a loss of 5% to 18% of cumulated earnings over the first 10 years--see Oeropolous *et al.* (2012) for Canada, and Kahn (2010) and Altonji *et al.* (*2016*) for the US.

Hershbein (2012), and Genda *et al.* (2010) argue that, for male high-school graduate in the US, the labor market operates more like a spot market—in which workers suffer little effects of graduating during a recession because of the lack of specific investments and job ladders, and the flatter experience profiles for these workers. In contrast, we find that, in Spain, wage rigidity among the less educated creates a situation where demand shocks translate mainly into employment losses. In a typical recession, job losses concentrate in the secondary segment of the labor market, and only few high-quality jobs in the primary segment of the labor market remain, in which the "best" low-educated workers are hired. It is important to highlight that those less educated workers who find employment during a recession are much more likely to do so in a permanent contract job. The rest of high-school graduates entering the labor market become non-employed. Our finding of positive selection in the case of low educated workers is consistent with results from Brunner and Kuhn (2014) on workers with vocational training in Austria.

Interestingly, Genda *et al.* (2010) and Cockx and Ghirelli (*forthcoming*) also find larger and more persistent earnings losses from a negative shock at labor-market entry for high-school graduates than for college graduates in Japan and Flanders, respectively.

In the case of Japan, in addition to the two-tier structure within a firm (with regular and irregular workers), the Japanese labor law requires that high schools lead the matching process between graduating seniors and prospective employers. These authors find that a one-percentage point increase in the unemployment rate at entry reduces the likelihood of working by 3 to 4 percentage points over 12 years, leading to a 5% to 7% decrease in earnings. Further exploration leads the authors to conclude that this negative effect on earnings is due to a continuous decline in the probability of full-time, regular employment. For the more educated Japanese workers, a drop in regular employment does not drive the earnings losses.

In Flanders, Cockx and Ghirelli (*forthcoming*) find that a typical recession reduces earnings of low-educated workers (via hours worked full-time) by 4.5% up to 12 years after graduation (with no effect on hourly wages, and only a small short-lived effect on salary employment). The authors argue that binding high minimum wages and short-time work compensation in low-skilled jobs explain the lack of results on hourly wages.[31] Instead, they find that the effect shows through a reduction in hours worked full-time. As these authors measure the employment rate annually, they do not rule out the possibility that bad entry labor-market conditions may increase the unemployment rate (see their footnote 35). The weak employment protection legislation in Flanders among blue-collar jobs also explains that short-time work compensation is at works here as opposed to work in the secondary labor market.

For college graduates, our results on annual earnings are slightly higher than those found by Oreopoulos *et al.* (2012) in Canada, and Altonji *et al.* (*2016*) in the US, and smaller than those found by Kahn (2010) (also in the US). Oreopoulos *et al.* (2012) find that a rise in unemployment rate by 5-percentage points (the average increase in a typical Canadian recession) implies an initial loss in earnings of 9% that halves within 5 years, and fades to 0% by 10 years. The cumulated ten-year effect amounts to 5%.[32] Instead, Kahn's estimates imply that the wages of college graduates would fall 25% the first year of entry and 20% after five years of entry, due to an increase in the entry unemployment rate of 4-percentage points (the average increase in a typical U.S. recession).

However, the mechanisms differ. Entering the labor market during a recession in Spain brings a sharp and persistent drop of jobs in the primary segment of labor market, but some hiring continues to occur in the secondary labor market. In contrast with more flexible labor markets, most of the adjustment occurs through employment and permanent employment (as opposed to weekly earnings). Thus, wage rigidity also mediates on the effects of initial labor market conditions on college graduates in Spain. Below, we further

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[31] According to Cahuc (2014), "short-time work compensation schemes provide additional funds so that employees can reduce their hours of work without a proportional reduction in their take-home pay. The employees earn less than they do when in full-time employment, but more than they would receive in unemployment benefits."

[32] To obtain these estimates from column 3, Table 1 in Oreopoulos *et al.* (2012) one needs to multiply the estimates by 5, which is the average increase in a typical recession.

explore the mechanisms leading to persistent effects of initial labor market conditions across the different educational groups.

V. Full versus Dynamic Effect

How much of the persistence observed in Spain is explained by the unemployment rate at labor-market entry versus the subsequent sequence of unemployment rates that follow a negative shock (and are likely to differ from those experienced by individuals graduating in a good year)? To address this, we follow Oreopoulos *et al.* (2012), and, estimate equation 2 where we control for conditions in the labor market that take place overtime and affect workers after labor-market entry. Results are shown in Figure 3. Figure 3 also shows the full-effects specification (from Table 2 and equation 1) for comparison purposes.

By comparing the full versus the dynamic effects in Figure 3, we find that controlling for the full-sequence of unemployment rates reduces the penalty during the first two to three years but increases it afterwards. This result is consistent across the three education groups and suggests that part of the penalty of entering the labor market during a recession is simply due to the less favorable sequence of unemployment rates that follow. However, the result also suggests that unlucky individuals benefit less from the subsequent periods of economic recovery, as younger cohorts take advantage of the new and better economic environment, the gap with older, unlucky, cohorts grows.

Figure 3 also shows that the penalty persists for workers with vocational training and college graduates even after 10 years of potential experience indicating that the negative entry effect accumulates over time and persists even when the economic recession gives way to a new expansion. These findings are consistent with evidence on the persistent effects of entry labor market on earnings on more experienced workers -- see Beaudry and DiNardo (1991), McDonald and Worswick (1999), Grant (2003), and Schmieder and von Wachter (2010). They contrast with those of Oreopoulos *et al.* (2012) for Canada, who find evidence supporting "*the greater importance economic conditions have at the beginning of one's labor market career relative to their effect after an individual has begun his career*".

VI. Effects on Mobility

Our results indicate that in Spain the negative effect of bad-entry conditions is not only large across all education groups, but also very persistent even at high levels of experience. What explains this persistence? Is it the lack of mobility or a lack of improvement when workers move? What type of job transitions do workers experience when they graduate under different conditions? In this section, we explore these issues further.

Figure 4 shows whether entering the labor market during a recession is associated with higher mobility across firms, industries, and provinces, or with a lower probability of working under a permanent contract. Figure 4 portrays a quite different story for college graduates versus non-graduates. Focusing first on college graduates, Panel C shows that entering the labor market during a recession increases mobility across provinces and this effect steadily increases over time. Panels A and B show that entering the labor market during a recession also increases mobility across firms and industries, respectively. Interestingly, this effect remains constant around 10% and 5% even after 10 years of potential experience.

Oreopoulos *et al.* (2012) find that college graduates who entered the Canadian labor market in the midst of the recession tend to move to better jobs as their career advances, and this job mobility helps them reduce the negative wage gap from the beginning of their career. Is the mobility observed among Spanish college graduates also driven by workers searching for better paying jobs, as is the case in Canada, or is it driven by job churning due to the precariousness of fixed-term contracts? Panel D in Figure 4 shows that this mobility does not come with a higher likelihood of entering the primary segment of the labor market. Perhaps more concerning, this effect not only persists through the first 10 years of potential experience, but also worsens over time.

To further explore this, Appendix Table A.4 presents our baseline specification controlling for the probability of firm changes (columns 2, 4, and 6) for each education group. If mobility were the result of finding new, better paying jobs, controlling for mobility would increase the earnings losses of college graduates, since those who graduate during recessions are more mobile. Instead, if mobility is involuntary and the result of having a precarious job, controlling for it would explain part of the earnings gap due to bad entry conditions. Controlling for firm mobility has little explanatory power and, if anything, it reduces the penalty. Hence, the results in the Appendix table indicate that workers mobility due to bad-entry conditions does not lead to higher incomes but instead is more likely the result of involuntary churning.

To sum up, in the case of college graduates, the story is not as a much about non-employment but about the lack of jobs in the primary segment of the labor market. During a recession, college graduates enter the labor market under fixed-term contracts, which leads to more mobility. The high province, industry and firm mobility is associated with churning in and out of fixed-term contract jobs. To put it differently, college graduates entering the labor market during a negative shock are trapped in the secondary market.

Moving now to workers without a college degree, we observe in Panel C in Figure 4 that negative entry labor-market conditions *decrease* province mobility at labor-market entry. Nonetheless, this negative effect converges to zero within 3 years for workers with vocational training and 6 years for high-school graduates. Thereafter, the effect on provincial mobility is positive. Brunner and Kuhn (2014) also find evidence that entering the labor market during a recession has an initial negative impact on vocational-training workers' mobility across firms and industries in Austria.

Labor-market entry also affects firm and industry mobility of non-college graduates (as shown in Panels A and B, Figure 4). Although average firm and industry mobility are higher the lower the education level (probably due to the higher incidence of fixed-term contracts among those groups of workers), a negative shock at labor-market entry decreases firm and industry mobility during the first five years for workers with vocational training and during the first ten years of potential experience for high-school graduates.

Panel D shows that for non-college graduates, entering during a recession increases the odds of a permanent contract during the first couple of years of potential experience, and decreases these odds, thereafter. This initial positive effect of graduating during a recession for this group is consistent with the fact that, for low-skilled workers, only a few good jobs are offered, as there is no hiring in the secondary labor market. The bottom line is that as non-college individuals fail to find jobs when they graduate in the midst of a recession, only those of higher quality manage to find a permanent contract, and hence they stay with their employers. The lower incidence of fixed-term contracts implies less churning (note that this is due to a composition effect by which only the "best" non-graduates find "good" jobs if they graduate during a recession). Over time, as economic conditions improve and new fixed-term contract jobs are offered to the low skilled, employment in the secondary segment of labor market increases, and so does mobility.

VII. Conclusion

Using Social Security data merged with Labor Force Survey data, we analyze the effects of entry labor-market conditions on the career of Spanish male workers in a context of high wage rigidity and segmented labor markets. Most interestingly, we do the analysis by highest educational attainment. We find that labor market institutions mediate differently across the three educational levels. In particular, we find that human capital is a double edge sword as college graduates who enter during a recession suffer a relatively small employment loss but end up trapped in the secondary labor market. Interestingly, the extremely weak wage pro-cyclicality in Spain prevents the adjustment of wages among the highly educated workers. The lack of short-time work compensation among low-skill jobs in Spain limits the adjustment via hours for the less educated workers. These results contrast with findings from other European labor markets with stringent regulations, such as Flanders. This paper contributes to an emerging literature on the effects of labor-market entry on workers' careers by highlighting the relevance of labor-market institutions and individuals' human capital.

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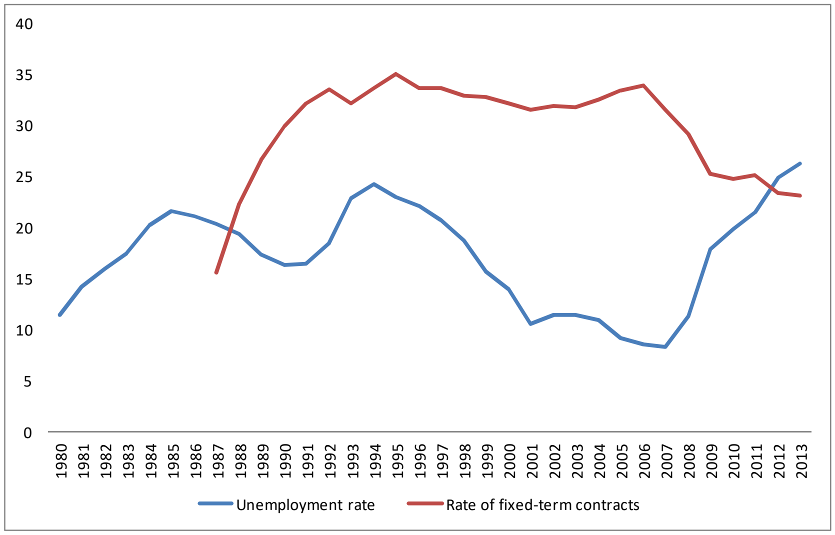
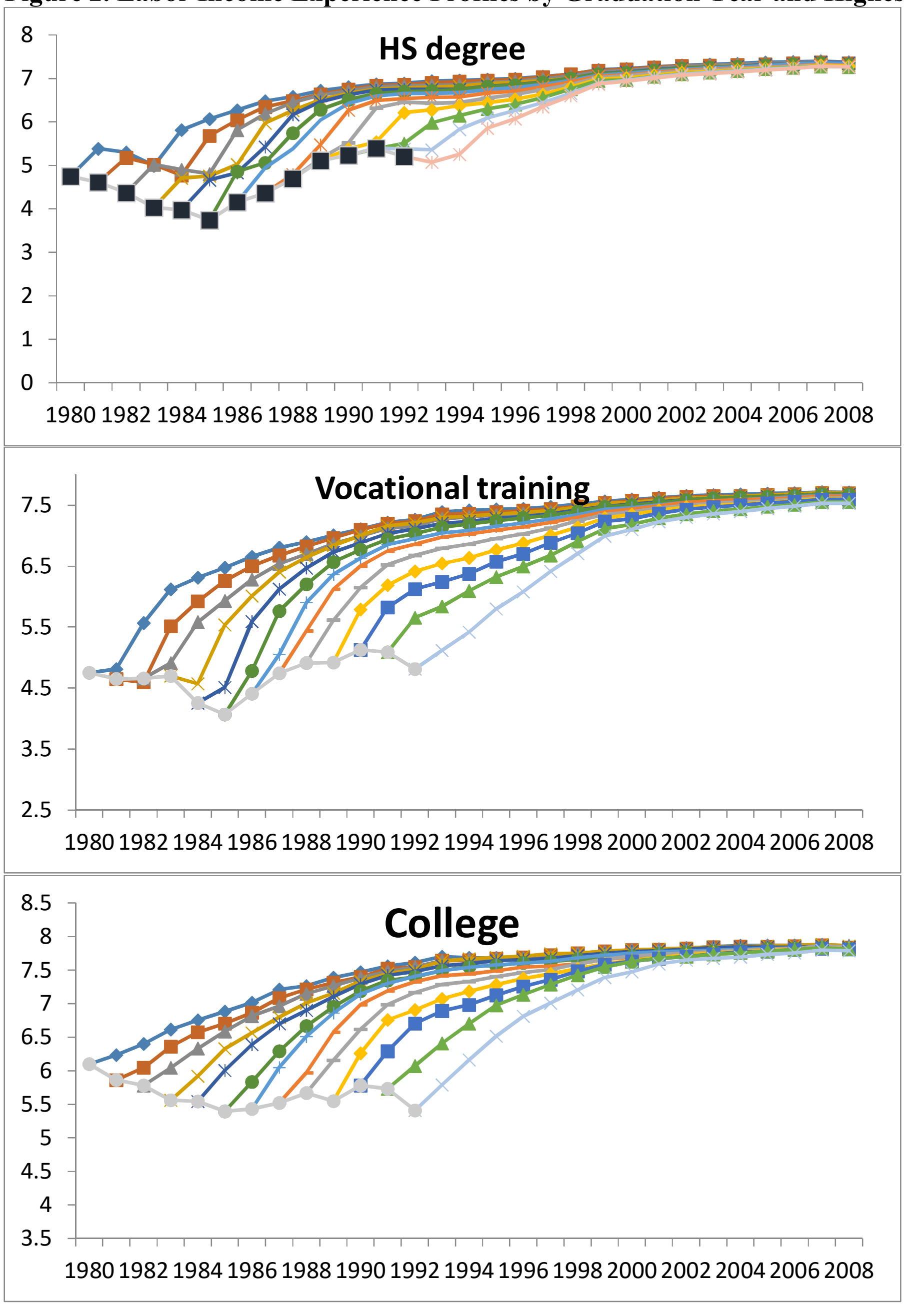
Figure 1. Rate of Unemployment and of Employment under Fixed-term Contract in Spain.



Source : O.E.C.D.

Figure 2. Labor Income Experience Profiles by Graduation Year and Highest Education Level



Note: The figure shows the log of average monthly earnings for a particular experience-year and for each entry-year cohort. The reason for the seemingly low values at initial levels of experience is due to the fact that many young individuals remain non-employed with zero income.

Table 1. Descriptive Statistics

Cohorts graduating between 1980 and 1992

	High School Degree				Vocational Training				College Graduates			
	Mean (1)	SD (2)	Min (3)	Max (4)	Mean (5)	SD (6)	Min (7)	Max (8)	Mean (9)	SD (10)	Min (11)	Max (12)
Collapsed dataset with zeroes for non-workers												
Entry unemployment rate	0.15	0.06	0.02	0.37	0.15	0.06	0.02	0.37	0.16	0.06	0.02	0.37
Potential Experience	12.30	7.13	1	29	12.30	7.13	1	29	12.25	7.10	1	29
Entry Year	1985.39	3.68	1980	1992	1985.39	3.68	1980	1992	1985.49	3.69	1980	1992
Year	1996.69	7.13	1980	2008	1996.69	7.13	1980	2008	1996.75	7.10	1980	2008
Employed	0.66	0.26	0	1	0.69	0.28	0	1	0.71	0.28	0	1
Monthly Income (in 2008€)	956.16	512.91	0	3,139.99	1,295.39	698.46	0	3,158.07	1,735.47	879.57	0	3,508.25
Nobs	14,879				14,905				14,075			
Collapsed dataset using observations of those who work only												
Month wage(2008€)	1,357.33	329.08	585.84	3,680.17	1,742.92	443.01	577.57	3,206.68	2,310.70	571.27	535.63	3,786.67
Hours worked	39.76	0.55	26.66	40	39.64	1.04	23.36	40	39.48	1.84	8	40
Permanent contract	0.59	0.24	0	1	0.73	0.23	0	1	0.79	0.24	0	1
White collar	0.02	0.04	0	0.66	0.17	0.15	0	1	0.68	0.28	0	1
Part-time	0.01	0.02	0	0.67	0.01	0.04	0	1	0.02	0.07	0	1
# of observations	14,699				14,605				13,512			

Notes: Equal Weighting across Graduation Year-Potential Experience-Province-Year Cells.

Table 2. Effect of an 8-ppt Rise in Unemployment Rate at Labor Market Entry on Annual Earnings, by Potential Experience
Cohorts Graduating Between 1980 and 1992
(Full-Effects Specification)

	High School Degree	**Vocational Training**	**College Graduates**
	(1)	(2)	(3)
Panel A. Effect at potential experience at year			
1	-0.251***	-0.249***	-0.130***
	(0.026)	(0.030)	(0.053)
3	-0.174***	-0.185***	-0.093**
	(0.022)	(0.024)	(0.043)
5	-0.103***	-0.128***	-0.063*
	(0.018)	(0.020)	(0.036)
7	-0.040***	-0.081***	-0.040
	(0.017)	(0.018)	(0.032)
10	0.031*	-0.031	-0.022
	(0.018)	(0.019)	(0.032)
Panel B. Average effect for potential experience years			
1-5 years	-.176***	-.187***	-.095**
	(.021)	(.024)	(.043)
1-10 years	-0.096***	-0.125***	-0.064*
	(0.018)	(0.019)	(0.035)
# of observations	14,879	14,905	14,075
R^2	0.32	0.36	0.48

Notes: Standard errors (in parentheses) are clustered at the graduation year-initial province. All models control for graduation cohort fixed effects, year fixed effects, initial-province fixed effects and a quadratic in experience.
***, **, * indicate significance at the 1%, 5%, 10% level.

Table 3. Sensitivity Analysis, Effect of an 8-ppt Rise in Unemployment Rate at Labor Market Entry on Annual Earnings, by Potential Experience (Full-Effects Specification)
Cohorts Graduating Between 1980 and 1992

	Province Uc	State Uc	National Uc	Non-movers	Non-employment rate province level	Uc 2 years before grad. year
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: High School Graduates						
1	-0.251***	-0.243***	-0.298***	-0.254***	-0.257***	-0.273***
	(0.026)	(0.022)	(0.066)	(0.026)	(0.019)	(0.025)
5	-0.103***	-0.106***	-0.121***	-0.105***	-0.103***	-0.104***
	(0.018)	(0.014)	(0.040)	(0.019)	(0.013)	(0.019)
10	0.031*	0.014	0.037**	0.029*	0.040***	0.043***
	(0.018)	(0.016)	(0.015)	(0.018)	(0.012)	(0.019)
1-10 years	-0.096***	-0.101***	-0.113***	-0.098***	-0.094***	-0.097***
	(0.018)	(0.014)	(0.036)	(0.018)	(0.013)	(0.018)
# of observations	14,879	5,083	299	14,879	14,879	14,879
R^2	0.32	0.92	0.96	0.87	0.87	0.87
Panel B: Vocational Training						
1	-0.249***	-0.285***	-0.236***	-0.252***	-0.303***	-0.228***
	(0.030)	(0.028)	(0.078)	(0.030)	(0.022)	(0.031)
5	-0.128***	-0.158***	-0.059	-0.131***	-0.141***	-0.094***
	(0.020)	(0.018)	(0.047)	(0.021)	(0.015)	(0.024)
10	-0.031	-0.053***	0.088***	-0.033*	0.009	-0.002
	(0.019)	(0.018)	(0.009)	(0.019)	(0.014)	(0.024)
1-10 years	-0.125***	-0.154***	-0.053	-0.128***	-0.132***	-0.094***
	(0.019)	(0.017)	(0.042)	(0.020)	(0.015)	(0.023)
# of observations	14,905	5,083	299	14,905	14,905	14,905
R^2	0.36	0.91	0.95	0.83	0.85	0.84
Panel C: College Graduates						
1	-0.130***	-0.162***	-0.071	-0.135***	-0.259***	-0.176***
	(0.053)	(0.045)	(0.088)	(0.053)	(0.032)	(0.049)
5	-0.063*	-0.073***	0.010	-0.067*	-0.119***	-0.070*
	(0.036)	(0.030)	(0.036)	(0.037)	(0.024)	(0.038)
10	-0.022	-0.011	0.061*	-0.026	0.003	-0.014
	(0.032)	(0.028)	(0.033)	(0.032)	(0.024)	(0.037)
1-10 years	-0.064*	-0.073***	0.008	-0.069**	-0.113***	-0.075***
	(0.035)	(0.029)	(0.034)	(0.036)	(0.024)	(0.037)
# of observations	14,075	5,076	299	14,075	14,075	14,075
R^2	0.48	0.85	0.94	0.64	0.64	0.64

Notes: Standard errors (in parentheses) are clustered at the graduation year-initial province level in columns (1), (4) and (5), at the graduation year-initial state level in column (2), and at the graduation year level in column (3). Models in columns (1), (4), and (5) control for graduation cohort fixed effects, year fixed effects, initial-province (or state in column 2) fixed effects and a quadratic in experience. In column (3), all models control for year fixed effects, a quadratic in experience and linear and quadratic cohorts trends.
***, **, * indicate significance at the 1%, 5%, 10% level.

Table 4. Effect of an 8-ppt Rise in Unemployment Rate at Labor Market Entry on Various Outcomes, by Potential Experience
Cohorts Graduating Between 1980 and 1992
(Full-Effects Specification)

	High School Degree			**Vocational Training**			**College Graduates**		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Probability of Working	Prob. of Permanent Contract¥	Monthly Wages	Probability of Working	Prob. of Permanent Contract¥	Monthly Wages	Probability of Working	Prob. of Permanent Contract¥	Monthly Wages
1	-0.251***	0.071	0.033***	-0.231***	0.021	-0.009	-0.107**	-0.071	-0.022
	(0.026)	(0.064)	(0.010)	(0.032)	(0.049)	(0.011)	(0.054)	(0.051)	(0.014)
3	-0.177***	0.019	0.029***	-0.163***	-0.021	-0.015	-0.074*	-0.084**	-0.018
	(0.022)	(0.046)	(0.009)	(0.025)	(0.036)	(0.010)	(0.043)	(0.040)	(0.013)
5	-0.108***	-0.019	0.025***	-0.102***	-0.053**	-0.020**	-0.047	-0.093**	-0.015
	(0.018)	(0.033)	(0.008)	(0.020)	(0.027)	(0.009)	(0.035)	(0.033)	(0.012)
7	-0.046***	-0.045*	0.021***	-0.050***	-0.076***	-0.024***	-0.027	-0.098***	-0.013
	(0.015)	(0.026)	(0.008)	(0.017)	(0.021)	(0.008)	(0.030)	(0.029)	(0.011)
10	0.025	-0.063***	0.014*	0.003	-0.095***	-0.030***	-0.010	-0.100***	-0.013
	(0.016)	(0.022)	(0.008)	(0.018)	(0.018)	(0.008)	(0.029)	(0.027)	(0.011)
1-5 years	-0.178***	0.022	0.029***	-0.164***	-0.018	-0.015	-0.076*	-0.083**	-0.018
	(0.022)	(0.047)	(0.009)	(0.025)	(0.037)	(0.010)	(0.043)	(0.041)	(0.013)
1-10 years	-0.100***	-0.014	0.024***	-0.098***	-0.051*	-0.020**	-0.048	-0.091***	-0.016
	(0.017)	(0.033)	(0.008)	(0.019)	(0.027)	(0.009)	(0.034)	(0.033)	(0.012)
# of observations	14,879	8,100	14,699	14,905	8,058	14,605	14,075	7,953	13,512
R^2	0.77	0.61	0.06	0.71	0.53	0.31	0.44	0.25	0.09

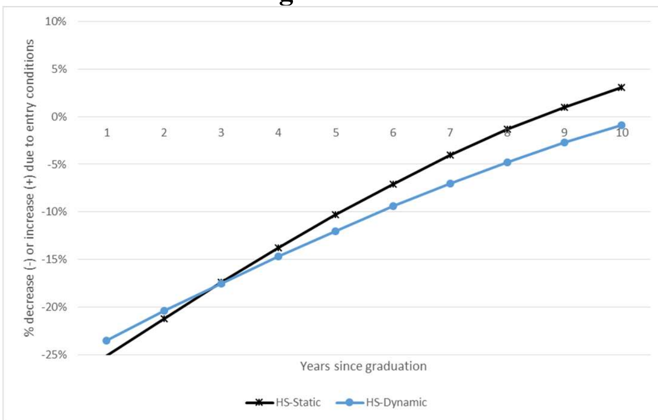
Notes: Standard errors (in parentheses) are clustered at the graduation year-initial province. All models control for graduation cohort fixed effects, year fixed effects, initial-province fixed effects and a quadratic in experience. Estimates for hours worked and monthly wages estimated only for those individuals who are working.

¥ Individuals who graduated after 1984, the year in which a law legalizing fixed-term contracts was passed in Spain.

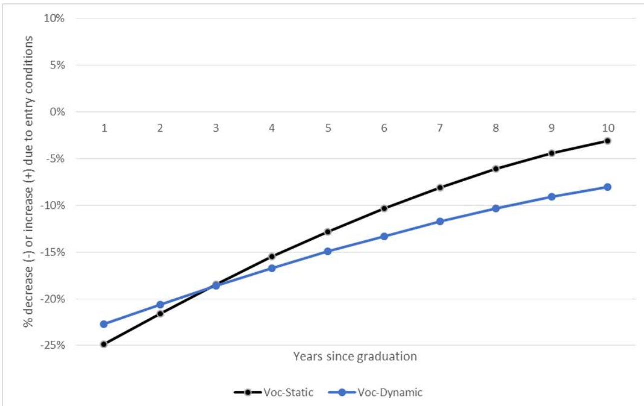
***, **, * indicate significance at the 1%, 5%, 10% level.

Figure 3. Persistent Effects of the Province Unemployment Rates the Year of Labor Market Entry on Annual Earnings
Effect of a 8-ppt Rise in Unemployment Rate at Labor Market Entry
Cohorts Graduating Between 1980 and 1992

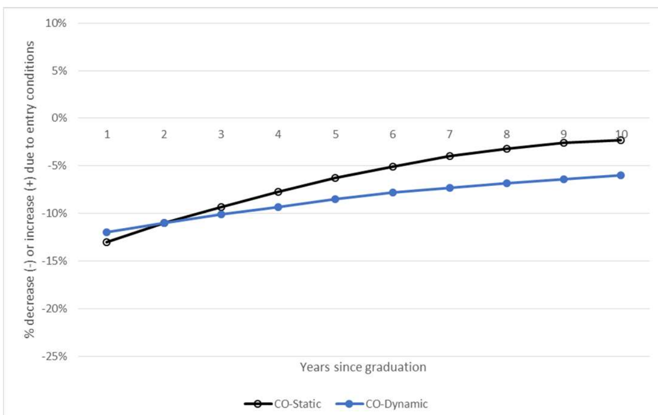
Panel A: High-School Graduates



Panel B: Vocational Training

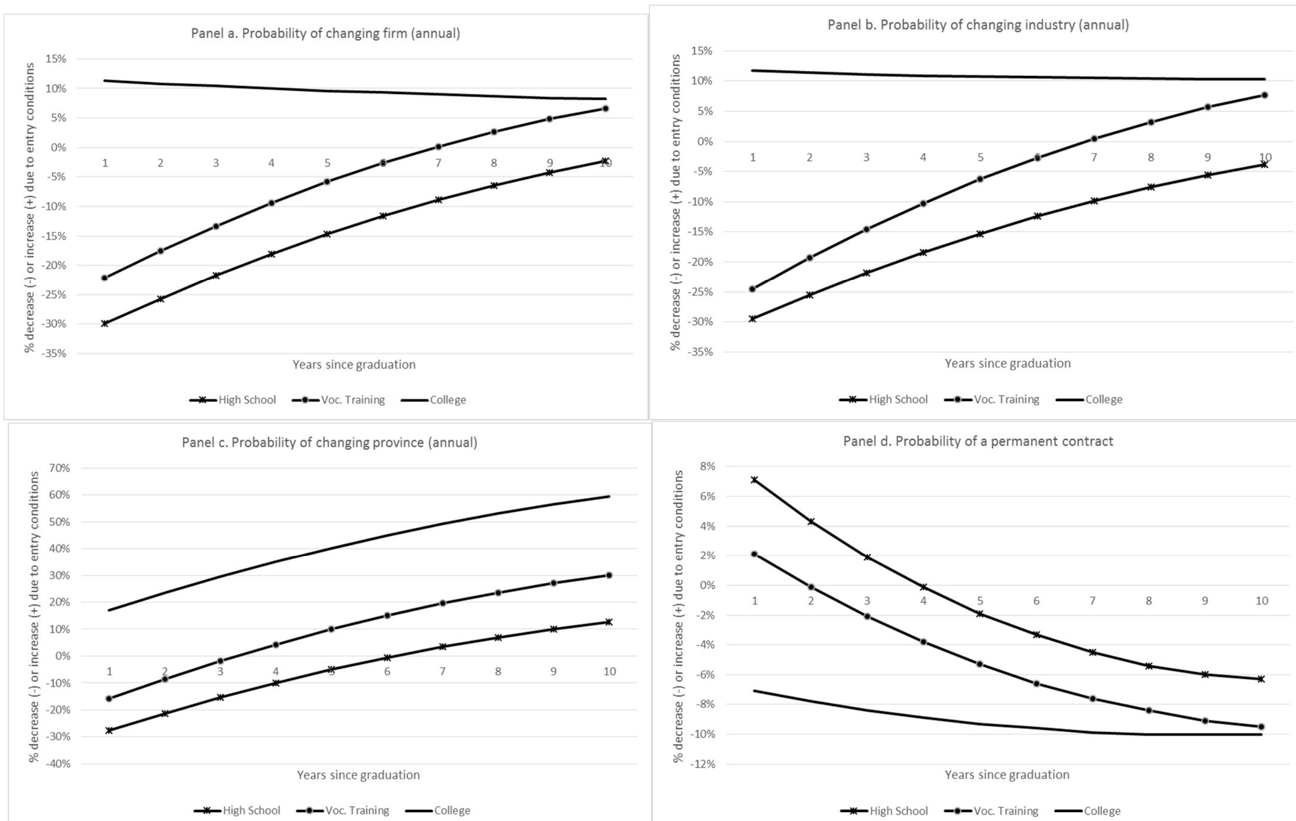


Panel C: College Graduates



Note: the figures show the effect of an eight percentage points increase in the unemployment rate at the time of entry on annual earnings under two specifications: the full-effects specification (static) and the specification that controls for the sequence of contemporaneous unemployment rates (dynamic).

Figure 4: Persistent Effects of the Province Unemployment Rates the Year of Labor Market Entry on the: (1) Probability of Changing Firms (annual change); (2) Probability of Changing Industry (annual change), (3) Probability of Changing Province (annual change); and (4) Fraction with Permanent Contract



Note: the figures display the percentage increase or decrease in the specific probability due to an eight percentage points increase in the unemployment rate at the time of entry. Full-effects specification.

APPENDIX
(Not for Publication)

Appendix Table A.1
The Effect of the Business Cycle on Males' High School Completion Rates, and College and Vocational Training Enrollment.
Spanish Labor Force Survey 1999-2015

The effect of an 8 percentage-point increase of the unemployment rate			
	(1)	(2)	(3)
Panel A. High-school completion rate			
High-school completion rate of individuals aged 16-19	-0.024*** (0.004)	0.019*** (0.001)	0.005 (0.007)
Number of (year-province) observations	750	750	750
Number of individual observations	254,674	254,674	254,674
R^2	0.23	0.57	0.59
Year Fixed Effects	X		X
Province Fixed Effects		X	X
Panel B. College and vocational enrollment rate			
College and Vocational Training Enrollment Rate of individuals 19 years old or less that finished high school when they were 18.	-0.008 (0.006)	0.023*** (0.004)	-0.001 (0.012)
Number of (year-province) observations	800	800	800
Number of individual observations	29,703	29,703	29,703
R^2	0.21	0.14	0.30
Year Fixed Effects	X		X
Province Fixed Effects		X	X

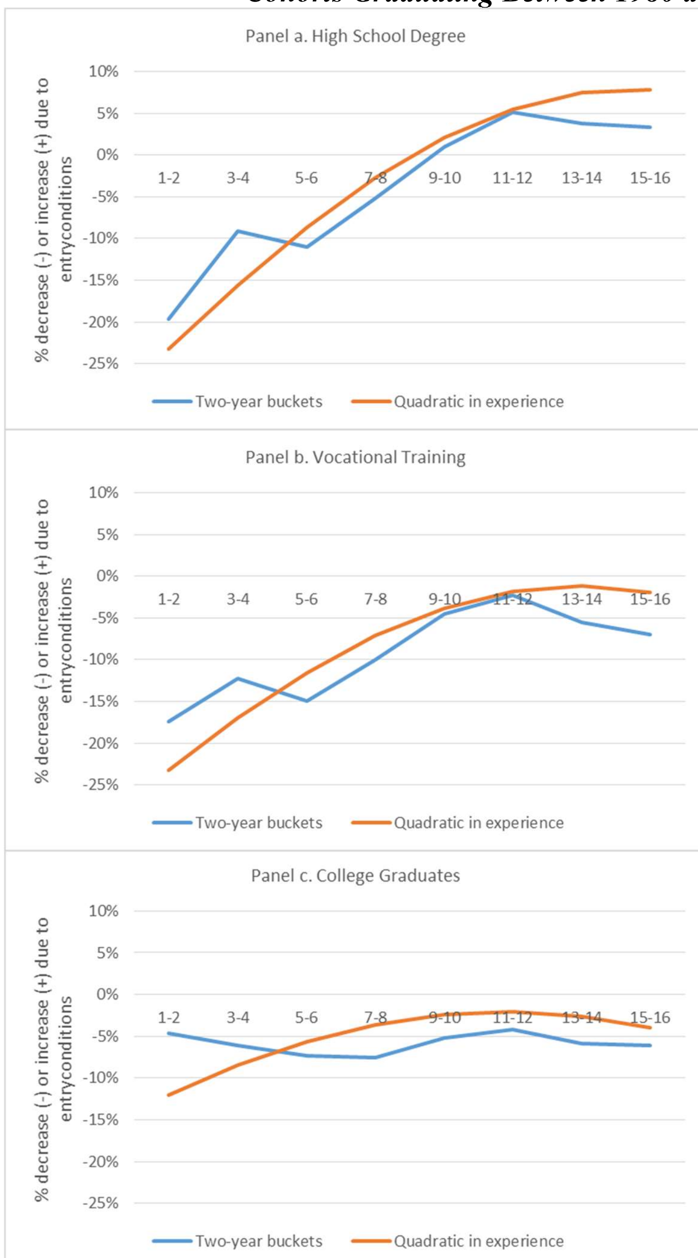
Note: Data collapsed at the province-year level. For the high-school completion rate, the dependent variable is the percent of individuals aged 16 to 19 who report having completed high school. The average graduation rate in our sample and for individuals aged 16 to 19 is 21%. This variable is regressed against the three-year average of the rate of unemployment at the province level. For the college enrollment rate, the dependent variable is the percent of individuals aged 16 to 19 who report having completed high school at the age of 18 and report studying in college or in vocational training. The average enrollment rate of individuals 19 years old or less and having finished high school at the age of 18 is 82.95%. This variable is regressed against the two-year average of the rate of unemployment at the province level. All regressions are weighted by the number of observations in each province-year cell. Robust standard errors clustered at the province level. *** Significant at the 1%; ** Significant at the 5%; * Significant at the 10%.

Appendix Table A.2.
Sample Sizes (individual-quarter) before Collapsing Data: Initial Unemployment Rate and Potential Experience Groups

	Potential Experience				
	(1-3)	**(4-6)**	**(7-9)**	**(10-12)**	**TOTAL**
		Panel A - High School Degree			
Entry unemployment rate (Uc)					
<11%	36,037	35,516	35,261	35,543	**142,357**
11-15%	55,792	55,362	55,324	55,603	**222,081**
16-19%	86,307	85,263	85,211	85,935	**342,716**
20-23%	38,771	38,278	38,413	38,768	**154,230**
>23%	72,925	72,405	72,686	72,943	**290,959**
TOTAL	**289,832**	**286,824**	**286,895**	**288,792**	**1,152,343**
		Panel B – Vocational Training			
Entry unemployment rate (Uc)					
<11%	28,953	28,681	28,662	28,778	**115,074**
12-15%	53,211	52,940	52,867	53,025	**212,043**
16-19%	79,336	78,737	78,816	79,323	**316,212**
20-23%	35,125	34,840	34,993	35,164	**140,122**
>23%	61,211	60,866	60,916	61,148	**244,141**
TOTAL	**257,836**	**256,064**	**256,254**	**257,438**	**1,027,592**
		Panel C – College Degree			
Entry unemployment rate (Uc)					
<11%	11,709	11,626	11,602	11,656	**46,593**
12-15%	27,500	27,407	27,399	27,487	**109,793**
16-19%	34,486	34,135	34,122	34,358	**137,101**
20-23%	15,187	15,112	15,213	15,290	**60,802**
>23%	22,011	21,912	21,945	21,973	**87,841**
TOTAL	**110,893**	**110,192**	**110,281**	**110,764**	**442,130**

Notes: Standard errors (in parentheses) are clustered at the graduation year-initial province. In Panel A, the coefficient effects may be interpreted as log points per a one standard deviation (SD) increase of the initial unemployment rate. They show the effect of a one SD increase of the initial unemployment rate including the effect of the sequence of unemployment rates that the individual goes through during his work life. All models control for graduation cohort fixed effects, year fixed effects, initial-province fixed effects and a quadratic in experience. ***, **, * indicate significance at the 1%, 5%, 10% level.

Figure A.1. Annual Earnings as a Function of Initial Conditions
Second-Order Polynomial versus Two-Year buckets Experience
(Full-Effects Specification)
Effect of a 8-ppt Rise in Unemployment Rate at Labor Market Entry
Cohorts Graduating Between 1980 and 1992



Notes: The figures show the coefficients from regressing annual earnings on the rate of unemployment at the time of entry. In the two-year buckets specification, the rate of unemployment at the time of entry is interacted with two-year buckets experience dummies, whereas in the quadratic in experience specification the entry rate of unemployment is interacted with a quadratic polynomial in experience. All models control for graduation cohort fixed effects, year fixed effects, initial-province fixed effects and a quadratic in experience (or experience dummies).

Appendix Table A.3. Effect of an 8-ppt Rise in Unemployment Rate at Labor Market Entry on Monthly Wages as a Function of Initial Conditions
***After* Controlling for Firm Characteristics**
(Full-Effects Specification)
Cohorts graduating between 1980 and 1992
(Individuals working at *t*)

	High School Degree				Vocational Training and Some College				College Graduates			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	Effect of a 8-ppt Rise in Unemployment Rate at Labor Market Entry											
1	0.033***	0.031***	0.032***	0.026***	-0.009	0.002	-0.006	-0.005	-0.022	-0.024*	-0.032***	-0.029***
	(0.010)	(0.010)	(0.010)	(0.009)	(0.011)	(0.011)	(0.010)	(0.010)	(0.014)	(0.014)	(0.011)	(0.011)
5	0.025***	0.022***	0.023***	0.019***	-0.020**	-0.009	-0.014*	-0.011	-0.015	-0.018	-0.024***	-0.024***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.008)	(0.007)	(0.012)	(0.012)	(0.009)	(0.009)
10	0.014*	0.011	0.012	0.009	-0.030***	-0.020***	-0.023***	-0.019***	-0.013	-0.017	-0.021**	-0.021**
	(0.008)	(0.008)	(0.008)	(0.007)	(0.008)	(0.008)	(0.007)	(0.007)	(0.011)	(0.011)	(0.009)	(0.009)
1-10 years	0.024***	0.021***	0.022***	0.018**	-0.020**	-.010	-0.015*	-0.012	-0.016	-0.019	-0.025***	-0.024***
	(0.008)	(0.008)	(0.008)	(0.007)	(0.009)	(.008)	(0.008)	(0.007)	(0.012)	(0.011)	(0.009)	(0.009)
Controls:												
Industry dummies		X	X	X		X	X	X		X	X	X
White- vs. blue-collar			X	X			X	X			X	X
Firm size				X				X				X
# of observations	14,699	14,699	14,699	14,699	14,605	14,605	14,605	14,605	13,512	13,512	13,512	13,512
R^2	0.89	0.90	0.90	0.91	0.87	0.88	0.89	0.90	0.73	0.73	0.77	0.78

Notes: Standard errors (in parentheses) are clustered at the graduation year-initial province. All models control for graduation cohort fixed effects, year fixed effects, initial-province fixed effects and a quadratic in experience.
***, **, * indicate significance at the 1%, 5%, 10% level.

Appendix Table A.4. Annual Earnings (with 0s for non-workers) as a Function of Initial Conditions Controlling for the Probability of FIRM CHANGES

Cohorts graduating between 1980 and 1992

	High School Degree		**Vocational Training and Some College**		**College Graduates**	
	No HIST.	HIST. of FIRM CHNG.	No HIST.	HIST. of FIRM CHNG.	No HIST.	HIST. of FIRM CHNG.
	(1)	(2)	(3)	(4)	(5)	(6)
			Effect of a 8-ppt Rise in Graduation Unemployment Rate			
1	-0.251***	-0.258***	-0.249***	-0.251***	-0.130***	-0.128***
	(0.026)	(0.025)	(0.030)	(0.030)	(0.053)	(0.053)
3	-0.174***	-0.185***	-0.185***	-0.189***	-0.093**	-0.091**
	(0.022)	(0.021)	(0.024)	(0.025)	(0.043)	(0.043)
5	-0.103***	-0.116***	-0.128***	-0.133***	-0.063*	-0.062*
	(0.018)	(0.019)	(0.020)	(0.020)	(0.036)	(0.036)
7	-0.040***	-0.056***	-0.081***	-0.087***	-0.040	-0.039
	(0.017)	(0.017)	(0.018)	(0.018)	(0.032)	(0.033)
10	0.031*	0.013	-0.031	-0.038**	-0.022	-0.022
	(0.018)	(0.018)	(0.019)	(0.019)	(0.032)	(0.033)
1-5 years	-0.176***	-0.186***	-0.187***	-0.191***	-0.095**	-0.093**
	(0.021)	(0.022)	(0.024)	(0.024)	(0.043)	(0.043)
1-10 years	-0.096***	-0.109***	-0.125***	-0.130***	-0.064*	-0.063*
	(0.018)	(0.018)	(0.019)	(0.020)	(0.035)	(0.035)

Notes: Standard errors (in parentheses) are clustered at the graduation year-initial province. In Panel A, the coefficient effects may be interpreted as log points per a one standard deviation (SD) increase of the initial unemployment rate. They show the effect of a one SD increase of the initial unemployment rate including the effect of the sequence of unemployment rates that the individual goes through during his work life. All models control for graduation cohort fixed effects, year fixed effects, initial-province fixed effects and a quadratic in experience. ***, **, * indicate significance at the 1%, 5%, 10% level.