The Employment Effect of Increasing Disability Benefits: A Regression Discontinuity Approach

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Abstract:

We analyze the employment effect of a law that provides for a 36 percent increase in the generosity of disability insurance (DI) in Spain for claimants who are, as a result of their lack of skills and of the labour market conditions they face, deemed unlikely to find a job. The selection process for treatment is therefore conditional on having a low probability of employment making evaluation of its effect intrinsically difficult. We exploit the fact that the benefit increase is only available to individuals aged 55 or older to estimate its impact using a regression discontinuity approach. Our first results indicate a large drop in employment for disabled individuals that receive the increase in the benefit. Testing for the linearity of covariates around the eligibility age threshold reveals that age at which individuals start claiming DI is not continuous: the benefit increase appears to accelerate the entry rate of individuals aged 55 or over. We obtain new estimates excluding this group of claimants (who enter the disability rolls at age 55 or after) and estimate the policy to decrease by 8 percent employment probability. We conclude that the observed DI generosity elasticity of .22 on employment is mostly due to income effects since benefit receipt is not work contingent.

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1 Introduction

Disabled individuals have incomes which are on average almost 15 percent lower than the rest of the population in developed Western economies and only 70 percent of the mean in the United States (OECD 2009). This is despite a very substantial documented increase in disability insurance (DI) availability and generosity in recent decades (Autor and Duggan, 2006; OECD, 2003). The cost of expanding this protection program has so far been outweighed by the sustained economic growth of the past twenty years[3]. With the fallouts of the financial crisis on government spending limitations, the burden of DI on the public purse will certainly come under renewed criticism. The recurring principal argument for reform of the disability benefit system has however not been its cost but rather its potential perverse incentive on the labour market participation (LMP) of certain groups of individuals.

The relationship between DI availability and generosity and LMP is an intrinsically difficult question to answer. The main criterion for eligibility is broadly always defined as having a physical or mental impairment that prevents a person from engaging in *substantial gainful activity*. This shows how the selection process into DI is strongly based on an individual having low LMP probability, making the claim and work decisions highly endogenous. Any evaluation of a disability benefit program must therefore carefully deal with this endogeneity issue in order not to over-estimate its impact on the labour market behaviour of recipients. There is now an influential literature exploring this relationship using various methodological approaches to the problem and we review it in the next section of this paper. The almost universal consensus is that DI has a very negative effect on the attachment to the labour market of eligible claimants. The remaining debate seems to be mostly about the size of this effect.

Autor and Duggan (2007, 2008[4]) are among the few that have recently focused on better understanding the mechanisms behind the behavioral response of DI claimants. Their main argument is that it could be due not only to the usually put forward distortionary substitution effect on incentives but also to a non-distortionary income effect. The latter interpretation would

[3] Despite increases in the number of claimants and of the average generosity of these benefits, the good economic performance of the economy has meant that the cost of DI has remained stable since the early 1990s in OECD countries at around 1.3 percent of GDP over this period

[4] We would like to thank these authors for making their 2008 unpublished report to the Social Security Administration available to us.

imply that the observed reduction in labour supply is not a deadweight loss and it is providing the right amount of transfer income in order for disabled individuals not to have to work above their substantial gainful activity level. These authors have attempted to empirically measure the importance of the income effect channel on LMP by using one of the few such DI programs in the USA which is not provided elusively on a work-contingent basis (Agent Orange). Their findings suggest a large income effect on near elderly males but are only tentative because of data limitations. We propose here to investigate this issue by exploiting certain unique features of the DI system in Spain.

The Spanish insurance system for disabled individuals is first characterized by a low and relatively stable recipiency rate by international standards[5]. Those who are eligible then receive monthly transfers which are fixed at a certain proportion of their wage level prior to the disability (i.e. it is a contributory insurance scheme[6]) and is secure until moving to retirement pension at age 65. This proportion is 55 percent for partial disability and 100 percent for total disability. Crucially the benefit amount is not contingent on not having employment income and has even built in taxation mechanisms to make LMP advantageous while on DI. Another interesting feature of the Spanish system is that certain claimants of partial disability benefits are eligible to a 36 percent increase when they turn 55 years old. This is granted to DI recipients with lower skill level and exposed to local labour market conditions which are deemed to make it difficult for them to find employment. The 75 percent replacement rate is granted to just under two thirds of partial disability claimants over the age threshold and close to none before that. This particularity enables us to investigate the impact of this large increase in DI generosity on the LMP of near elderly individuals using a regression discontinuity approach.

We use a large representative sample of the Spanish population receiving disability benefits for which we have monthly administrative data on work and benefit history between 1996 and 2007. We focus our attention on partial disability recipients who are aged between 51 and 58 and are able to identify the individuals who are treated with the DI increase. Because of selection on low LMP probability, naïve OLS estimates of the treatment effect logically generate huge policy impacts (-17 percent) even after controlling for observable characteristics. Augmenting the

[5] Only 4 percent of the population aged 20 to 64 receives disability benefits in Spain compared to an OECD average of 6 percent, which is the same number than in the United States.

[6] There is also a non-contributory disability benefits system but it is comparatively smaller in size (205.319 people received non-contributory disability pensions in Spain in 2007 as opposed to 868.026 that receive contributory disability pensions). We do not include the group of non-contributory pensioners in our analysis.

model with the inclusion of individual fixed effects yields a policy impact almost three times smaller. Our regression discontinuity (RD) approach first considers different age windows around the 55 years threshold. These results suggest that the increase in DI generosity is at least responsible for a 3 percentage points decrease in LMP. However, when we run some tests about the linearity of covariates around the eligibility age threshold, we can see that age at which individuals start claiming DI is not continuous. This result can be interpreted as evidence that the benefit increase accelerates the entry rate into the disability rolls of individuals aged 55 or over. Therefore, we proceed by estimating a new RD model but with a sample restricted to individuals who start receiving DI benefits before reaching age 55. Now results suggest that the policy prompts a 1 percentage point decrease in employment probability, which is translated into an 8 percent reduction in employment probability (as the average LMP for this sample is 12.5 percent).

We believe that these results are very much in line with the findings from previous research on this subject (specially results from Gruber, 2000) and claim that, considering the features of the Spanish system in terms of the benefit not being contingent on working status, the observed impacts of DI on LMP may be mainly due to an income effect, in line with was is argued in Autor and Duggan (2007, 2008).

The rest of the paper is structured as follows. Section 2 gives an overview of the related literature on the impact of DI on LMP. Section 3 discusses the disability benefit system in Spain and the increase in DI generosity program. Section 4 describes the data and gives some descriptive statistics. Section 5 presents the methodology. Section 6 reports and discusses the results and the final section concludes.

2 Related Literature

Much of the literature on work disincentives of permanent disability benefits is based on the analysis of the Social Security Disability Insurance (SSDI) program in the USA.

Labor force participation rates for older males in the USA have been falling during the last three decades and an extensive body of research has emerged that tries to link this evolution with the growth of the disability insurance program. The argument is based on the high implicit marginal tax rate on earnings above a threshold ($940/month in 2008) that is part of the eligibility criteria

in order to be accepted to the disability program[7]. It is widely accepted that the current design of the program creates disincentives to work for disabled individuals but there is still disagreement on the magnitude of these effects and their contribution to the decrease in labor force participation of older Americans.

The main problem encountered when trying to estimate the size of the disincentives to work resulting from the disability insurance system is the endogeneity of the receipt of disability benefits in a labor force participation equation. Therefore, in order to get unbiased results, researchers have tried to use an exogenous variation in the level of benefits or to rely on econometric techniques that can account for this endogeneity.

In this line of research, Bound (1989) compares a sample of rejected and accepted DI beneficiaries and estimates that the counterfactual labor force participation rate of disability recipients would have been 30 percentage points higher if they wouldn't have received the benefits. The validity of his estimates relies on the assumption that both groups are relatively similar in observed and unobserved characteristics although he recognizes that rejected applicants are usually somewhat healthier than accepted applicants so that he defends that his estimates represent an "upper bound" of the potential labor force participation of DI receivers.

Using a similar approach, a more recent study by Chen and Van Der Klaauw (2008) focuses on the impact of the receipt of disability benefits for marginal applicants for which access to the disability system is only decided in a second stage of the eligibility process and on the basis of vocational and age factors. Individuals that reach this second stage have been found to fulfill the medical impairment required in the first stage of the eligibility process and case workers use a grid of individual characteristics of the disabled worker such as age, education and past work experience in order to determine "the individual's residual functional capacity to perform work-related physical and mental activities". For age, the cutoff points are set at 45, 50 and 55 years old and the authors use the discontinuity in DI award rates at these points to estimate the effect of benefit receipt on labor supply. Their findings suggest that the receipt of disability benefits reduce labor force participation by 6-12 percentage points and labor supply by

[7] Applicants to the disability insurance system in the USA need to demonstrate that they did not work during the five months prior to the application and once they start receiving the benefits they cannot gain more than the threshold defined by the substantial gainful activity (SGA which was $940 per month in 2008). However, since 1960 disability recipients have access to a Trial Work Period in which, for 5 years, they can accumulate up to nine months of work above the SGA. If they earn more than the SGA for more than nine months, benefits are terminated (Maestas and Yin, 2008).

16-20 hours per month. However, the question remains to whether this result can be extrapolated to the whole population of disability benefit recipients as calculations are only done with individuals that reach the second stage of the benefits adjudication process who are also probably the ones with a relatively better health among the pool of disability recipients.

Maestas and Yin (2008) take a difference approach and exploit the fact that, once disabled individuals reach the Full Retirement Age (FRA), disability benefits are automatically converted to old-age pension. Since the year 2000, this change abolishes the working restriction embedded in the US disability system so that the tax on earnings disappears while the benefit amount remains the same. They calculate changes in labor force participation as DI beneficiaries move to old-age benefits and using a difference-in-difference approach get an estimate of 10.4 percentage points rise in labor force participation at full retirement age among former DI beneficiaries relative to non-DI participants. However, this number is calculated using a sub-population and it cannot be generalized to the whole population in the disability rolls as labor force participation naturally decreases with age. For that reason, the authors conclude that this number represents a "lower bound" estimate on the potential work capacity of individuals in the disability system (Maestas and Yin, 2008).

Apart from evidence for the United States, there are two more important studies that focus on Canadian data. The study by Gruber makes use of an increase of 36% in the level of disability benefits in all Canadian regions except Quebec introduced in 1987 in order to estimate an elasticity of labor force non-participation of 0.28-0.36 with respect to disability insurance benefits. Or course, his identification strategy relies on the exogenous variation in benefit amounts between the two regions, which assumes no correlation between the change in benefits and other differential labor market trends in the two regions. Another study by Hyatt (1996) focuses on Canadian data (Ontario) and uses a Heckman selection model to estimate the effect of permanent disability benefits on post-injury wages accounting for the self-selection of individuals into the employment option. Hyatt proves that the generosity of worker's permanent disability benefits is negatively associated with post-injury employment.

In an attempt to summarize the results and to raise the attention of both governments and the general public opinion about the distortions introduced by the relatively generous disability systems, the OECD calculated the disincentive effects to work using the "disability benefits net

replacement rates[8]" for average earners which were estimated to be 39% in Australia, between 43-60% in the UK, 74% in Luxembourg and 64%-84% in the case of partial disability holders in Spain (121% in the case of full disability). Additionally, in a companion report published in 2003 also by the OECD, a "benefit generosity indicator" was constructed and proved to have a positive correlation with both beneficiary rates and disability benefit inflows (OECD, 2003). Therefore, the literature in the European context has, for some time, accepted that disability systems in Europe are not only used for persons with health incapacities but also as an alternative pathway into early retirement or as uncovered unemployment, particularly for the case of older individuals. As a result, countries have focused on the introduction of reforms to activate disabled individuals and to substitute the disincentives by incentives to take up a job for disability benefit recipients. Consequently, most of the latest literature relative to the European disability systems consists of policy evaluations of these reforms[9].

For the particular case of Spain, a number of papers have tried to identify the extent to which disability benefits have been used as an alternative exit from the labor market for disabled individuals: Blanco (2000) uses a competing risk model to identify the characteristics that affect older individuals when leaving employment to enter early retirement or disability benefits while Jimenez et al. (2009) focus on the effects of the business cycle on older workers going into disability benefits, unemployment or inactivity. Malo (2007) and Vall-Castello (2010) focus on the factors that influence the low employment rates of disabled individuals and the later paper also finds positive but small effects of an employment promotion policy introduced in 2004 to foster the integration of disabled women in the labor market (an increase in the deductions to the Social Security contributions offered to employers that hire a disabled women). However, none of the literature in Spain has tried to estimate the disincentives to work provided by the level of disability benefits awarded.

[8] Net replacement rates compare the income situation when moving from paid work to receiving disability benefits without working (OECD, 2007a).

[9] Examples of these evaluations are Humer et al. (2007) and Lalive et al. (2009) for Austria, Duell et al. (2009) for Norway, Stafford et al. (2006) and Corden et al. (2001) for the U.K., reports from the Department of Employment and Workplace Relations (2005) and from the Department of Family and Community Services (2003) for Australia, Van Ours (2006) for The Netherlands or Hartmann (2006) for Luxembourg, among others.

3 The Disability Insurance System in Spain

3.1 Types of disability pensions

In Spain, permanent contributive disability insurance is defined as the economic benefits that aim at compensating the individual for losing a certain amount of wage or professional earnings when the person is affected by a **reduction or a complete loss** of his/her working ability in a way that is assumed to be permanent due to the effects of a pathologic or traumatic process derived from an illness or an accident[10].

In order to capture the different situations in which a person can be after suffering from a disabling condition, the Spanish Social Security administration differentiates between three main degrees of disability[11] that depend on the amount of working capacity that has been lost:

(i) Partial Disability: Individuals suffering from the kind of impairment that disables the individual to develop all or the fundamental tasks of his/her usual job or professional activity but the individual is still capable of developing a different job or professional activity.

(ii) Total Disability: Individuals suffering from the kind of impairment that disables the individual for the development of any kind of job or professional activity.

(ii) Severe Disability: Individuals who are in a situation in which, as a result of anatomic or functional loses, the individual needs the assistance of a third person to develop the most essential acts of live such as eating, moving, etc…

Therefore, the degree of disability in which the individual is classified is set in relation to the working capacity lost as the goal of the benefits is to compensate for the reduction on wages or professional earnings caused by the disabling condition. The procedure of applying for permanent disability benefits can be initiated by the provincial office of the National Institute of Social Security (NISS), by the institutions that collaborate in the process (such as hospitals), or by the individual himself (in which case, more documentation is required). The evaluation is done by the Disabilities Evaluation Team which analyses the medical report prepared by doctors in the provincial office of the NISS and the report explaining the professional background of the person. The Disabilities Evaluation Team writes a report that will be evaluated by the directors of the provincial office of the NISS in order to take a decision on the type of disability pension

[10] Own translation of the definition of permanent disability given by the Spanish Social Security administration at www.seg-social.es.

[11] These three levels of disability represented 99.6 percent of permanent contributory DI claimants in 2007. A remaining 0.4 percent of claimants received a lump sum payment for every minor disability.

granted (if any), the benefit level and the date of the next planned revision to check whether there has been an improvement or an aggravation of the condition. Similarly, any revision of the disability degree or pension amount is done by the provincial office of the NISS. A graphical description of the process in provided in Figure A1 of the appendix.

3.2 Elegibility and Pension Amount

As we are dealing with contributory pensions, there are a number of rules in terms of eligibility requirements that the individual has to fulfill in order to qualify for the pension. These requirements as well as the pension amount vary depending on the source of the disability, which can be an ordinary illness, a work-unrelated accident or a work-related accident or professional illness. In the case of an ordinary illness, eligibility to the pension requires having contributed 1/3 of the time between turning 20 years old and the appearance of the disabling condition, with a minimum of 5 years of contributions[12]. There is no contributive requirement when the disability is caused by an accident, whether or not work-related, or a professional illness.

The total amount of the pension is obtained from multiplying a percentage, which varies depending on the type of pension, to the regulatory base, which depends on the source of the disability and on previous salaries. The regulatory base of the pension in the case of an ordinary illness is computed as the average wage of the last 8 years of work. For a work-unrelated accident, is the average annual wage over a period of 24 consecutive months chosen by the person within the last 7 years of work. For work-related accident or professional illness, it is the average wage in the last year of work.

The percentage applied to the regulatory base is 55% under the Partial Disability regime, 100% under the Total Disability scheme and 100% plus another 50% to cover for the person that is taking care of the disabled in the case of Severe Disability. However, for partial disability pensioners, this 55% can be increased by 36% for individuals who are older than 55 years old and "as a result of their lack of education/preparation and the social and labor market conditions of the region where they live, it is assumed that it will be difficult for them to find a job". Therefore, individuals in the Partial Disability scheme that comply with these requirements can receive a 75% of the regulatory base instead of a 55%.. All permanent disability pensions are automatically converted to old-age pensions once the individual turns 65 years old.

[12] If the individual is younger than 31 when becoming disabled, the requirement is to have contributed for 1/3 of the time between turning 16 years old and the appearance of the disabling condition with no minimum of years required.

In terms of income taxes, Partial Disability benefits are taxable under the general income tax rules and Absolute disability pensions are always exempted from income taxes. Furthermore, if the individual works while receiving the pension, there is a reduction in the employment income used to calculate the income tax of 2.800 euros/year if their degree of disability is low (between 33% and 65%) or 6.200 if the disability level is higher (more than 65%) or if the disabled has reduced mobility. All this legislation can be translated in the following equations for recipients of Partial Disability benefits:

Partial Disability without a job:

$$y_i\,(p,a,) = \begin{cases} 0{,}55\,p_i - r \times (0{,}55\,p_i) & if\ a < 55 \mid a \geq 55\ \&\ t = 0 \\ 0{,}75\,p_i - r \times (0{,}75\,p_i) & if\ a \geq 55\ \&\ t = 1 \end{cases}$$

Where p_i is the regulatory base of the pension amount that depends on the contributions in the last job (salary) and on the source of the disability[13], t is an indicator function for treatment and a is age.

Partial Disability with a job:

$$y_i\,(d,w,p) = \alpha p_i - c \times d_i + w_i - r \times [\alpha p_i + (w_i - k(d_i)]$$

With

$$\mathrm{k(d_i)} = \begin{cases} 2.800\ if\ 33 \leq d_i \leq 65 \\ 6.200\ if\ d_i > 65 \end{cases} \qquad \begin{cases} \alpha = 0.55 \text{ if } a < 55 \mid a \geq 55\ \&\ \mathrm{t} = 0 \\ \alpha = 0.75 \text{ if } a \geq 55\ \&\ \mathrm{t} = 1 \end{cases}$$

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Where d_i is the degree of disability and $c \times d_i$ is the cost of working for a disabled individual which is a linear function of his/her degree of disability.

For this paper, our interest lies on the evaluation of the disincentives to work caused by the increase in the pension amount at age 55. For this reason, we are restricting our analysis to disabled individuals in the Partial Disability scheme where the law applies.

[13] We take it as a value; we don't calculate it as everybody in the sample receives always a pension and it would be very difficult to calculate because we don't know the source of the disability.

4 Data and Descriptive Statistics

4.1 The Data

The study will use the Continuous Sample of Working Lives ("Muestra Continua de Vidas Laborales", MCVL) which is a microeconomic data set based on administrative records provided by the Spanish Social Security Administration. It contains a random sample of 4% of all the individuals who, at some point during 2007, had contributed towards the social security system (either by working or being in an unemployment scheme) or had received a contributory DI. The random sample selected contains over one million individuals.

It contains information on the entire monthly employment and benefit history of the workers, including the exact duration of employment, unemployment and disability benefit spells, and for each spell, several variables that describe the characteristics of the job or the unemployment/disability benefits. There is also some information on personal characteristics such as age, gender, nationality and level of education. We also capture the economic business cycle by compiling quarterly unemployment rates at the province level (there are 52 provinces in Spain).

We select in our sample all individuals who are aged 51-58 and are receiving a Partial disability pension at some point between 1996 and 2007. The final sample for our analysis contains 623,228 monthly observations for 14,692 individuals.

4.2 *Descriptive Statistics*

About a third of the claimants in our sample are women (4,708) which is relatively large but follows the recent trends in the increase in the female to mail ratio among disability benefit recipients. The education level of these DI claimants is low with only about 30 percent having completed at least secondary education. This is not surprising since we are focusing on older workers which have on average less schooling but also because disability is likely to be correlated to education, as it is likely to generate differences in health investment. Almost 18 percent of the individuals in our sample (2,592) have worked at some point while they were claiming Partial Disability benefits. This is relatively high considering that this is only about half the employment rate of non disabled workers of the same age group in Spain over this period.

Our methodological approach (described below) relies on the discontinuity in the 20 percent rate of DI increase (DII) at age 55 which is a criterion for selection into the program. We show in Figure 1 that the proportion of individuals treated does dramatically jump after individuals reach

this age threshold. There also appears to be a very small number of individuals who receive the DII before being 55 which may reflect some level of administrative error or cheating by disability case workers. Still the difference in treatment is very large, about 60 percent, and does warrant evaluation using a *fuzzy* regression discontinuity design.

We check that the treatment discontinuity is not a result of differences in covariates around the age threshold graphically. Figure 2 shows that women constantly represent approximately one third of our sample of claimants throughout. In Figure 3 we must first note that older individuals are on average less educated. But the graph is reassuring and shows that this criterion for eligibility to the benefit increase appears to be very smooth around the age threshold. The pattern for age started claiming DI in Figure 4 is inverse with older benefit recipients having on average entered DI older. The distribution of average age started claiming DI still appears relatively around the eligibility threshold but we will check this more systematically in our statistical analysis.

Finally we plot the changes in our outcome of interest, LMP, by age in Figure 3. Again we see that the detachment from the labour market increases from age 52 onwards as our sampled individuals become older[14]. We also note what appears to be a small discontinuity around the threshold which could be resulting from the change in treatment probability. The significance of this gap is however difficult to judge graphically and will be confirmed or otherwise in our statistical model estimates.

5 Methodology

5.1 *OLS and Fixed Effect Models*

For individual *i* in month *t,* a simple statistical model relating LMP (our outcome of interest) to DII participation (the policy treatment) can be written as:

$$LMP_{it} = \alpha + \beta DII_{it} + \delta(Age) + u_{it} \qquad (1.1)$$

where α is an intercept, *LMP* is a dummy variable for working or not, *DII* a dummy variable for receiving the disability benefit increase and *u* is an error term. We control for the age in months of the claimants in quadratic form, $\delta(Age)$, since we have observed in Figure 3 that is negatively

[14] The marked drop in LMP at age 52 onward is still relatively puzzling to us since there are no disability insurance entitlement changes at this age. It could be due to the 'natural' effect of age on LMP as claimants become older and we therefore carefully control for this phenomenon in our statistical analysis.

correlated to LMP. If assignment to DII treatment was random, then β would be an unbiased estimator of the impact of *DII* on *LMP*.

However, the definition of having a low LMP probability for DII selection makes it non-random and so a regression estimate from equation (1.1) will be biased – overestimating the decreases in employment rates due to the programme. One possible means to deal with this is to augment (1.1) by adding observable characteristics of claimants to amend the equation as:

$$LMP_{it} = \alpha + \beta DII_{it} + \delta(Age) + \gamma_k X_{ki} + \eta UR_{it} + u_{it} \qquad (1.2)$$

Where *k* individual characteristic are included in the vector of control variables, *X,* and *UR* is the unemployment rate the claimant is exposed to in the area where he resides in month *t.* The Ordinary Least Squares (OLS) estimate of β is then the relationship between *LMP* and *DII* holding constant the X's and controlling for labour market conditions[15].

Since we have panel data information on the treatment and employment status of disability benefit claimants we can include individual fixed effects, α_i, as follows:

$$LMP_{it} = \alpha_i + \beta DII_{it} + \delta(Age) + \eta UR_{it} + u_{it} \qquad (1.3)$$

The individual characteristics, *Xs*, from (1.2) are now dropped since they do not vary over time. Our interpretation of the estimated β is that it should now be free of observed *and* unobserved claimant characteristics. Still, if selection into *DII* and *LMP* is dependent on factors not included in equation (1.3) the selection problem remains. Since the rule for DII selection is that claimants must be over 55 years old, we consider another methodology which should be better in addressing the discussed selection problem: Regression Discontinuity.

5.2 *Regression Discontinuity Models*

Regression Discontinuity (RD) design has had a long history in statistics, but has recently gained prominence among economists for its potential for dealing with the problem of unobservable characteristics alongside its conceptual simplicity[16]. This method can only be applied when there exists a cut-off point of an assignment variable *Z* above and below which there is a strong difference in treatment probability. As we clearly illustrated in Figure 1, this is the case for DII treatment depending on age of the claimant (*Z)* due to the 55 years minimum selection rule.

[15] We use the unemployment rate to control for labour market characteristics rather than month and year dummies since these would be highly correlated with the age (in months) control that we include in the model.
[16] For a clear and detailed discussion on the RD methodology, see for example Imbens and Lemieux (2008).

A widely researched and very intuitive example of RD occurs for the 50 percent cut-off rule for winning or losing an election. The argument is that different units (areas, firms) which have had very close votes around the cut-off are likely to be very similar in observed and unobserved characteristics. Still they will have opposite outcomes whether they were above or below the assignment cut-off, making it very simple to compare the difference in impact of selection or not. In this case, an unbiased treatment effect on outcome, here *LMP*, with subscripts + and – indicating proximity to either side of the threshold can be written as:

$$\beta = LMP^{+} - LMP^{-}$$

It is extremely simple to estimate β here since being above the cut-off guarantees treatment and we only have to compare the means of the outcome around that point. This is called a s*harp* RD as the probability of treatment, or inclusion into a program, jumps from 0 to 1 on either side of the cut-off.

In the case of DII treatment, as in many other programs, the change in the probability of treatment around the assignment variable threshold is not so sharp but does greatly increase. This type of set up is called a *fuzzy* RD and it is still possible to exploit the discontinuity to identify a treatment effect[17]. In this case however the difference in outcomes around the cut-off will be a function of the difference in the jump in the proportion treated around this point. Mathematically, using average LMP, the mean proportion of claimants receiving DII, and the subscript + and – as before, we can write $LMP^{+} - LMP^{-} = \beta(LMP^{+} - LMP^{-})$. This can be re-written as the RD estimator:

$$\beta = \frac{LMP^{+} - LMP^{-}}{DII^{+} - DII^{-}} \tag{2.1}$$

If it is the case that claimants just below and just above the age cut-off do have similar characteristics (observable and unobservable) then the estimator in equation (2.1) can legitimately be used to estimate the causal impact of DII on LMP. This is because it simply compares the difference in employment rates of individuals which have been randomly assigned around an assignment threshold and which should consequently have similar characteristics. Of

[17] In our case we are actually facing what has been referred to as a '*partially fuzzy*' by Battistin and Rettore (2008) or a '*simple special case*' by Blundell and Costa Dias (2009) version of the RD methodology. This is because treatment is only available but not mandatory on one side of the threshold. Both of these papers highlight the advantage of this approach relative to the standard *fuzzy* RDD.

course since not all claimants over 55 receive the benefit increase, this must be scaled by the difference in the jump in the proportion of individuals that are treated around this point.

We can estimate β using different + and - windows in terms of age on each side of the threshold. As the age difference of individuals around the threshold becomes smaller we expect the RD estimate of DII treatment to decrease since claimants on either side will be ever more similar in observed (and unobserved) characteristics. A more robust test of this similarity is to include individual characteristics of our claimants as controls when obtaining RD estimates of DII on LMP in the smallest age window available. If these characteristics are randomly distributed around the age threshold, we expect that including controls will not significantly change the RD estimates. But if they become significantly different with the inclusion of certain individual trait of DI claimants, it will suggest that we must consider how this affects the validity of our RD estimates.

6 Results

The results from the OLS, columns (1) to (3), and fixed effect estimates, columns (4) and (5), of DII on LMP are reported in Table 1. The simple OLS estimate including a quadratic term in age is very large and only slightly smaller when controlling for individual characteristics (gender, education, age started claiming disability). Also taking into account changes in labour market conditions by including provincial unemployment rates does not change the estimated 17 percent impact of the policy on employment participation. This large effect was to be expected considering that selection for DII is contingent on having a low LMP probability. The fixed effect estimates are much smaller and suggest that the program only reduces the probability of working by 6 percent. We believe this to be a superior estimate of the DII impact to the OLS one since it measures the average effect of switching from non-treatment to treatment in terms of changes in LMP for the same individuals. As we explained above we may still be concerned that individual fixed effects and labour market conditions cannot account for the endogenous nature selection for treatment of a claimant and her behavioural response. For this we turn to the RD analysis.

Table 2 reports RD estimates for four different age windows (from +/- 4 years to +/- 1 year) around the 55 year threshold. The discontinuity in DII is clearly important and represents

a jump of between 50 and 60 in the proportion of treated claimants. The difference in LMP is significant and between 5.1 and 1.6 percentage points lower after the age cut-off. The RD coefficients are the ratios of these differences and they are all statistically significant[18]. However it is clear that the impact of DII on LMP becomes much smaller as the age window around the threshold is reduced as the individuals in our sample become ever more similar. A straightforward RD methodology therefore suggests that an increase in DI generosity reduces work probability by, at least, 3 percentage points.

In Table 3 we report RD estimates which include DI claimant individual characteristics for the 55 years old +/-1 year sample. The second and third columns of results show that the inclusion of gender and secondary education completion dummies does not significantly change the estimated impact of the DI increase on LMP which remains roughly -.03. However, the inclusion of the age at which individuals started claiming disability benefits generates an RD estimate three times smaller. This suggests that this covariate is not linearly distributed around the eligibility age threshold although this was not graphically obvious from Figure 4. We may, for example, worry that the benefit generosity increase incites a number of individuals to start claiming DI after turning 55 and have low LMP probabilities (i.e. the new higher replacement rate is enough for them to use DI as a form of early retirement).

Figure 6 plots the proportion of claimants who entered DI at different ages in our sample. We do find here that there appears to be a substantial acceleration in the proportion of new entries of individuals aged 55 or older into DI. This cannot be entirely explained by the natural increase of probability to claim DI with age. The jump in entry rates between ages 54 to 55 is substantial and we must therefore conclude to an important entry effect into DI of increasing the generosity of benefits available. To obtain an impact of DII on LMP net of entry effects, we restrict our sample to individuals who started claiming disability benefits before age 55. We report our RD estimates for the +/- 1 year window including individual characteristic controls for this sample in Table 4. We first note that our RD estimates are robust to the inclusion of claimant individual characteristics suggesting that observed (and unobserved) characteristics of claimants

[18] As the estimate is similar to a local IV estimate of DII on LMP instrumented by a claimant being older than the cut-off age, we are able to obtain standard errors as recommended by Hahn et al. (2001)

around the threshold are linearly distributed for this sample. We therefore conclude that the net impact of DII is to reduce by 8 percent employment probability[19].

7 Conclusions

In this paper we analyze the employment effects of a 36 percent increase in the amount of disability pensions (DII) that is granted to almost 60 percent of the individuals aged 55 or above who are receiving a partial disability pension in Spain. The DI system in Spain allows partially disabled claimants to combine the receipt of benefits with income from employment without any implicit tax on labour supply being levied (and even providing for tax incentives to work).

We exploit this discontinuity of DII provision from age 55 onwards and first use a straightforward fuzzy regression discontinuity approach to estimate the effect of treatment on LMP (which are likely to be endogenously determined). We generate RD estimates using different age windows, from +/-4 to 1 year before and after the cut-off age, which suggests that the increase in DI generosity reduces the probability of working by, at least, 3 percentage points.

However, once we model more carefully for the impact of other individual characteristics we discover an acceleration of entry into DI of claimants aged 55 and older which we believe is partly due to the increased generosity of available benefits. Once we take this phenomenon into account, we estimate that the employment of DI recipients would have been 8 percent higher had they not received the benefit increase. Since the replacement rate is in practice increased by 36 percent, this translates into an elasticity of DI generosity to LMP of approximately .22.

These results are very much in line with the literature on the employment effect of DI almost unanimously concluding to a negative causal relationship. Our results nevertheless are an important contribution for two distinct reasons. Firstly, they are among the first, together with Gruber (2000), to focus on the impact of benefit generosity rather than entitlement. Secondly, the features of the DI institutional system we study make it possible to rule out that this impact stems from a substitution effect. As benefit receipt is not work contingent, it suggests that there is an important income effect at play in the work decision of older workers as observed by Autor and Duggan (2007, 2008).

[19] The average LMP for this sample is of 12.5 percent and thus a 1 percentage point decrease corresponds to an 8 percent drop in employment probability.

This latter point is important in terms of the efficiency of policy reforms since an income effect does not imply any deadweight loss. We believe this should be seriously considered in future reforms which may use the disincentive substitution argument to cut the benefits of a group in an already relatively weak income position. This is especially true considering the findings by Bound et al. (2004) that workers on average value increased benefits somewhat above the average cost of providing them.

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Figure 1: Proportion of DI Clamants Receiving Rate Increase by Age

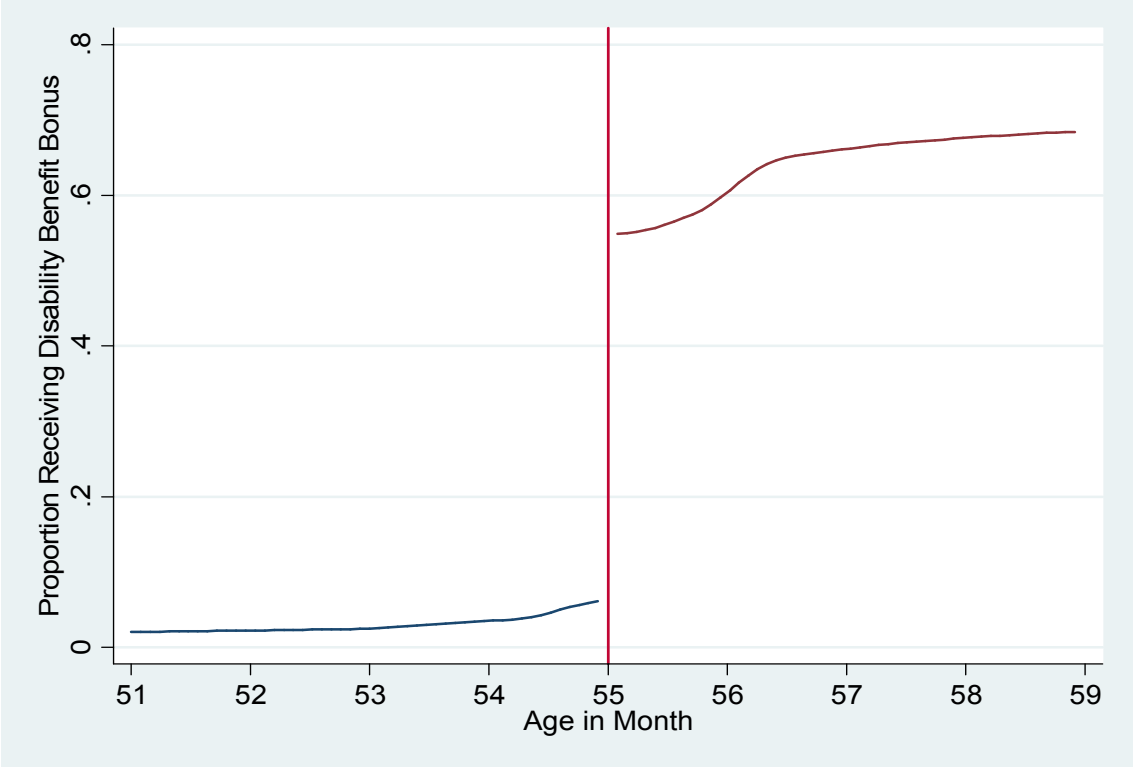

Figure 2: Proportion of DI Claimants who are Women by Age

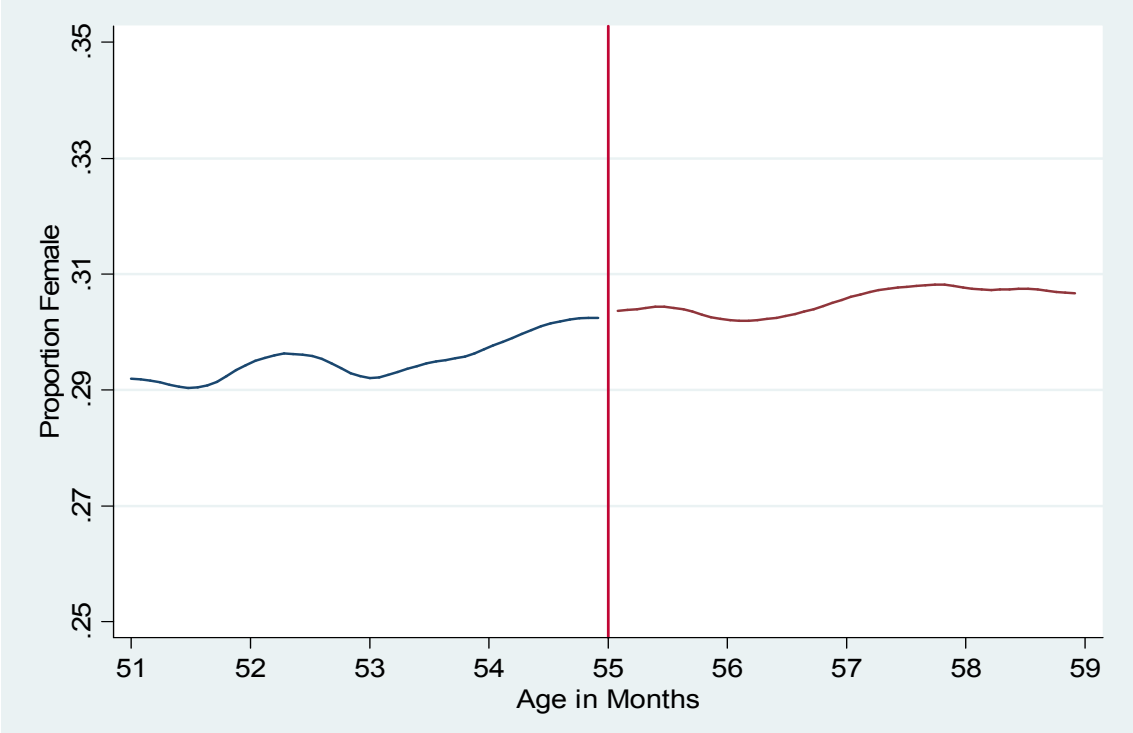

Figure 3: Proportion of DI Claimants with Secondary Education by Age

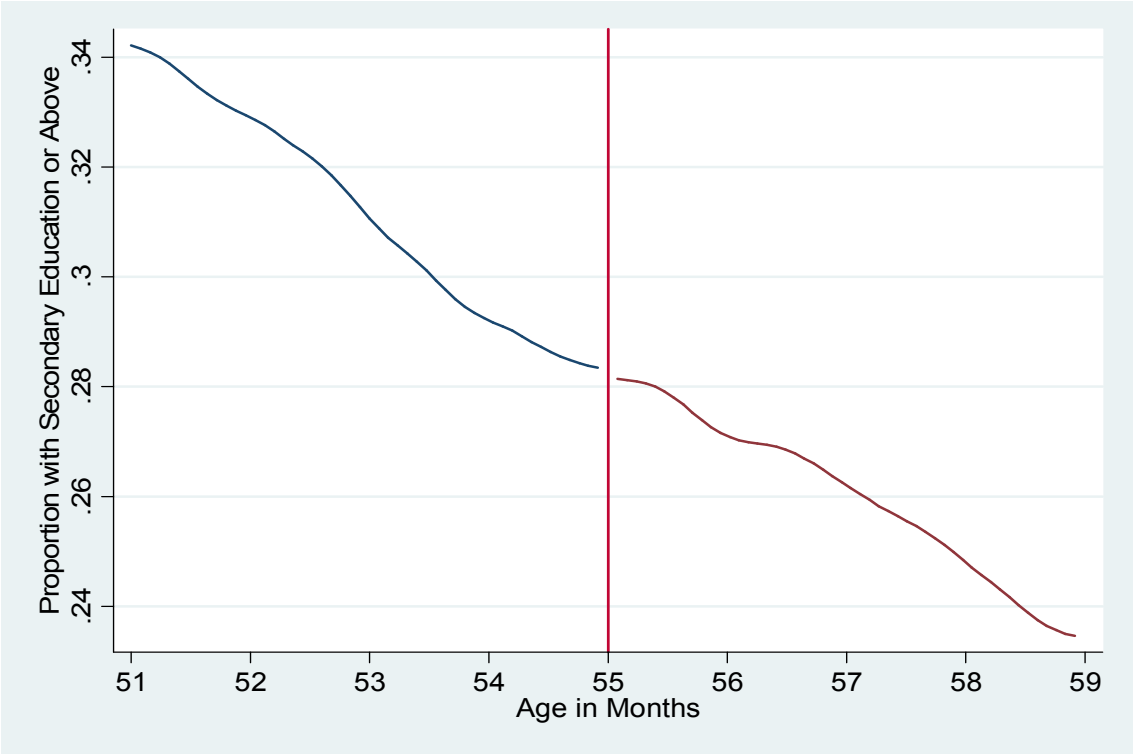

Figure 4: Age Started Claiming DI by Age

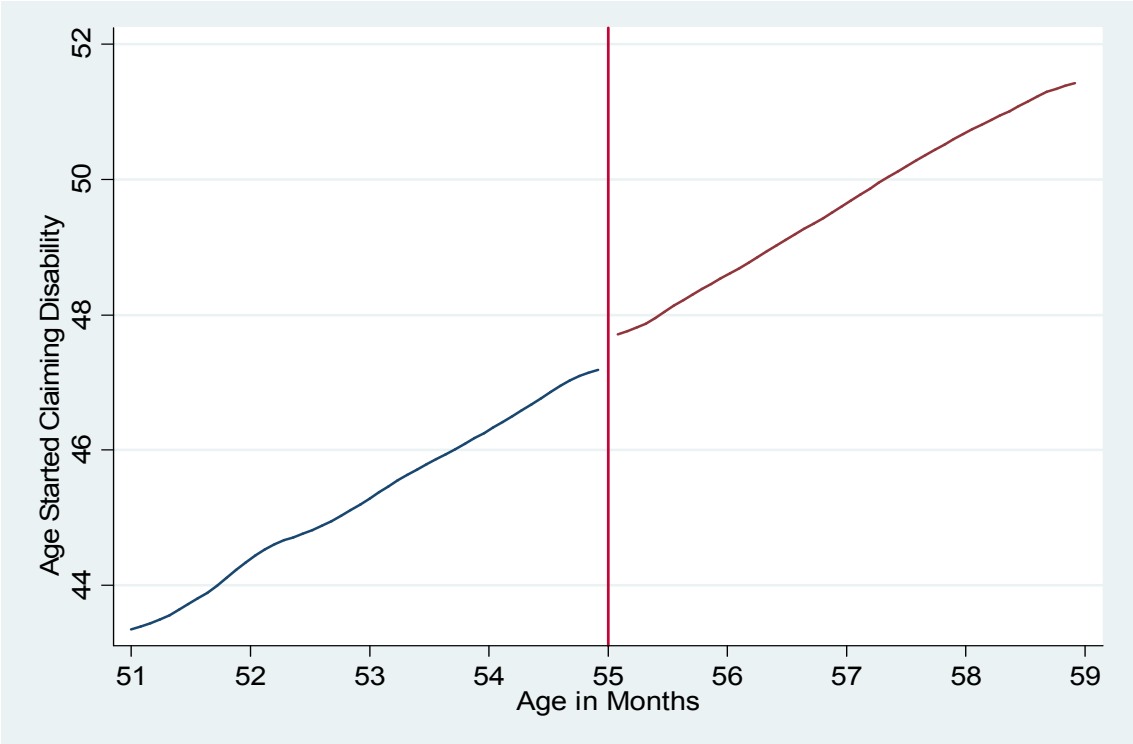

Figure 5: LMP of DI Claimants by Age

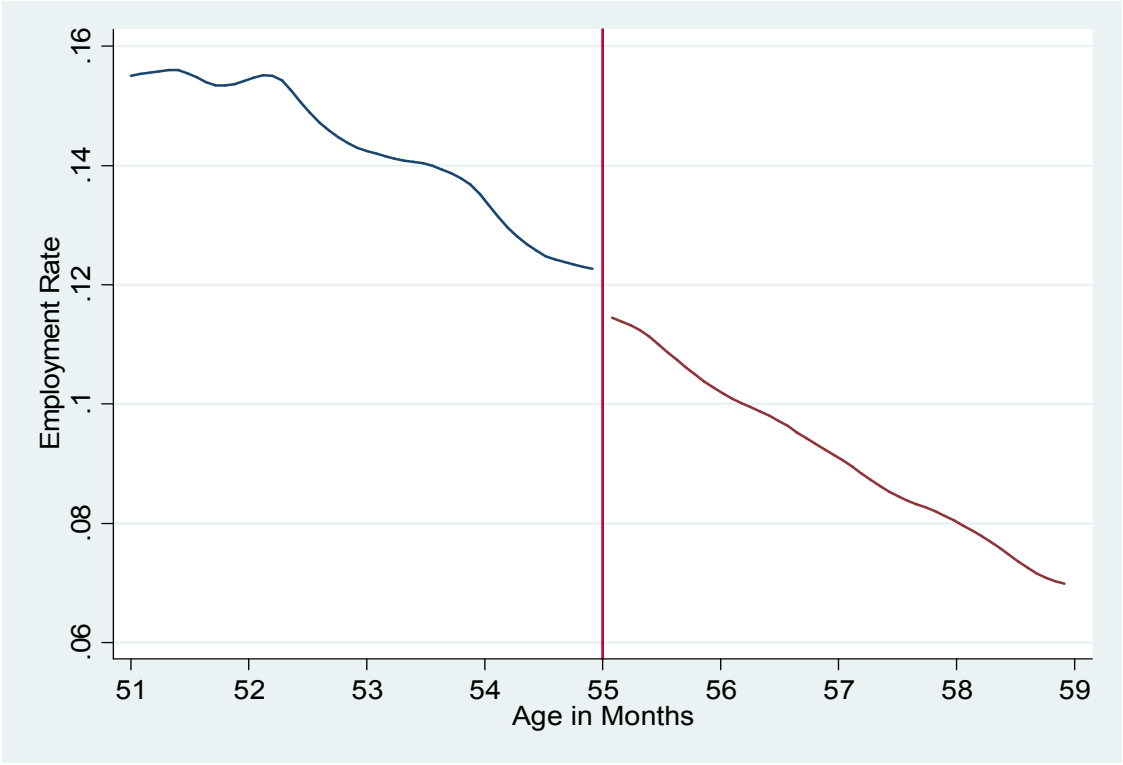

Figure 6: Distribution of Age Started Claiming DI

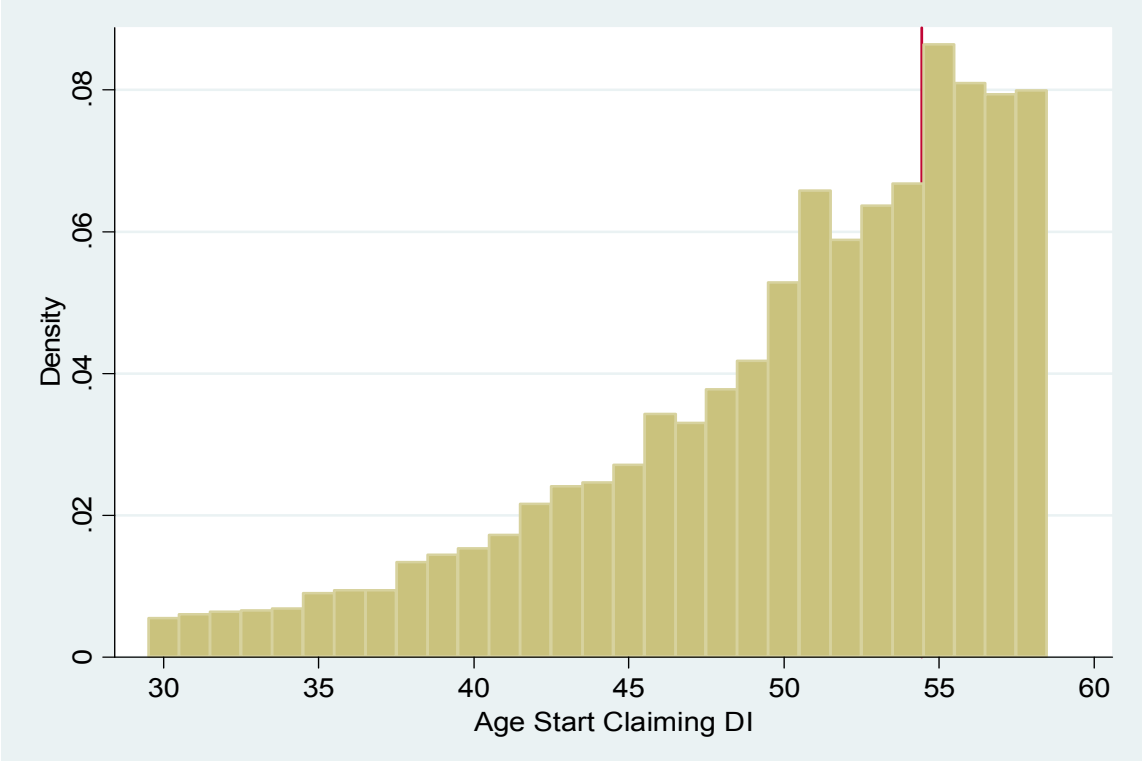

Table 1: OLS and Fixed Effect Results

	Dependent Variable: Labour Market Participation				
	OLS			FE	
	(1)	(2)	(4)	(5)	(6)
Disability Benefit Bonus (DBB)	-.187 (.001)	-.170 (.001)	-.170 (.001)	-.060 (.001)	-.060 (.001)
Age/Age^2/Age^3	Yes	Yes	Yes	Yes	Yes
Individual Controls	No	Yes	Yes	No	No
Unemployment Rate	No	No	Yes	No	Yes
Individual Fixed Effects	No	No	No	Yes	Yes
Number of Observations	623,228	623,228	623,288	623,288	623,288
Number of Individuals	14,692	14,692	14,692	14,692	14,692

Note: Age is in month; Age^2 and Age^3 are respectively the square and cube of the difference from the mean Age. The Individual Controls are: gender; secondary school completion; and age start claiming disability benefits. The Unemployment Rate is quarterly for the 51 Spanish administrative regions.

Table 2: RD Results for Different Age Windows around Eligibility Age Threshold

	Estimation on Individuals Aged 55 and +/- 4, 3, 2 and 1Year			
	All Ages	**+/- 3 Years**	**+/- 2 Years**	**+/- 1 Years**
Discontinuity of DII Treatment Around Threshold ($DII^{+} - DII^{-}$)	.612 (.001)	.595 (.001)	.567 (.001)	.517 (.002)
Difference in LMP Around Threshold ($LMP^{+} - LMP^{-}$)	-.051 (.001)	-.041 (.001)	-.030 (.001)	-.016 (.002)
Estimated Effect of DII on LMP Participation ($LMP^{+} - LMP^{-}$)/ ($DII^{+} - DII^{-}$)	*-.084* *(.001)*	*-.070* *(.002)*	*-.052* *(.002)*	*-.031* *(.003)*
Number of Observations	632,228	463,438	306,460	151,904
Number of Individuals	14,692	12,424	10,314	8,313

Table 3: RD Results +/- 1 Year Around Eligibility Age Threshold with Individual Controls

	Estimation on Individuals Aged 55 and +/- 1Year			
Discontinuity of DII Treatment Around Threshold ($DII^{+} - DII^{-}$)	.517 (.002)	.517 (.002)	.517 (.002)	.508 (.002)
Difference in LMP Around Threshold ($LMP^{+} - LMP^{-}$)	-.016 (.002)	-.016 (.002)	-.015 (.001)	-.006 (.002)
Estimated Effect of DII on LMP Participation ($LMP^{+} - LMP^{-}$)/ ($DII^{+} - DII^{-}$)	*-.031* *(.003)*	*-.030* *(.003)*	*-.029* *(.002)*	*-.011* *(.003)*
Proportion Female	No	Yes	Yes	Yes
Proportion with Secondary Education	No	No	Yes	Yes
Age Started Claiming Disability	No	No	No	Yes
Number of Observations	151,904	151,904	151,904	151,904
Number of Individuals	8,313	8,313	8,313	8,313

Table 4: RD Results +/- 1 Year Around Eligibility Age Threshold with Individual Controls for Individuals who Started Claiming before Age 55

	Estimation on Individuals Aged 55 and +/- 1Year			
Discontinuity of DII Treatment Around Threshold (DII^{+}– DII^{-})	.500 (.002)	.500 (.002)	.500 (.002)	.498 (.002)
Difference in LMP Around Threshold (LMP^{+}– LMP^{-})	-.007 (.002)	-.007 (.002)	-.007 (.002)	-.005 (.002)
Estimated Effect of DII on LMP Participation (LMP^{+}–LMP^{-})/ (DII^{+}– DII^{-})	*-.014* *(.003)*	*-.015* *(.003)*	*-.013* *(.002)*	*-.010* *(.003)*
Proportion Female	No	Yes	Yes	Yes
Proportion with Secondary Education	No	No	Yes	Yes
Age Started Claiming Disability	No	No	No	Yes
Number of Observations	138,121	138,121	138,121	138,121
Number of Individuals	6,879	6,879	6,879	6,879

Figure A1: Description of the Permanent Disability Pension Claim Process