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Abstract: We analyze the effect of a parametric reform of the fully-funded pension regime in Colombia on the intensive margin of the labor supply. We take advantage of a threshold defined by law in order to identify the causal effect using a regression discontinuity design. We find that a pension system that increases retirement age and the minimum weeks during which workers must contribute to claim pension benefits causes an increase of around 2 hours on the number of weekly worked hours; this corresponds to 4% of the average number of weekly worked hours or around 14% of a standard deviation of weekly worked hours. The effect is robust to different specifications, polynomial orders and sample sizes.

JEL codes: D91, J26

Keywords: Labor supply, Regression discontinuity, pension system reform, Colombia

1 Introduction

In the last decades, several countries undertook structural and parametric reforms in their pension systems. These reforms were particularly important in Latin American countries. The first country in the region that made these reforms was Chile, at the beginning of the 1980s. Some other countries followed the Chilean example; Bolivia, Mexico, Salvador and Dominican Republic substituted a fully-funded public system for a pay-as-you-go private system. Other countries, like Peru and Colombia, adopted a model in which both the public system and the private system coexist. The adoption of these structural and parametric reforms implied changes in replacement rates, minimum requirements to obtain pension benefits and contribution rates. It is well known that, economic theory predicts that many of these changes may imply changes in expected life-time income, labor choices and savings.

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These effects should be taken into account in the policy design of pension systems (Barr and Diamond 2008).

The literature has shown, that in developed countries, the design of pension systems affects labor, consumption and saving choices (e.g. Attanasio et al., 2003; Battistin et al. 2009; Friedberg, 2000). However, the analysis for Latin American countries has concentrated on topics related to the pension passive of the public systems, coverage rates and informality in the labor market (Calderon et al., 2011; Corbo et al., 2003; Mesa-Lago, 2002, 2004). To the best of our knowledge, there are no papers that consider the effect of how pension systems affect formal labor market in LA.

The aim of this paper is to evaluate to what extent the conditions to claim pension benefits in a fully-funded system affect the intensive labor supply in Colombia; for this purpose we take advantage of one of the changes that the regime had during the last decades. Since 1993 the Colombian pension system moved from a fully-funded public system to a system in which a fully-funded regime and a pay-as-you-go regime coexist and compete. The initial design of 1993 had several flaws that challenged the sustainability of the public system, so in the years following 1993 several additional partial reforms were performed. The initial reform was introduced with Law 100 of 1993; the main reform in this law was to allow the existence of a pay-as-you-go private pension system. However, the private system did not substituted completely the public system so that both systems coexist now and workers can choose to which system the want to belong. The law also established a Transition Regime (TR) for those individuals that where relatively old and near to retirement. People in the TR that decided to stay in the public regime could claim benefits with the rules that applied before the reform. In 2005, a new reform was undertaken (Legislative Act 01 of 2005) with parametric changes of the fully-funded system. Under the reform individuals that had contributed to the system less than 750 weeks in January 2005 should contribute more weeks in order to claim pension benefits; for these individuals retirement age also increased. We take advantage of the contributed-weeks threshold established by law to identify the effect of the pension reform on the intensive margin of labor supply measured as worked hours per week. We undertake a parametric sharp Regression Discontinuity Design in order to do so.

Our findings show that treated individuals work more. The average effect is around two hours per week. This effect represents 14.8% of the standard deviation and 4.2% of the average of the worked hours per week. The average effect is concentrated on relatively old employees and is independent of education level, household size and number of children. It is consistent with the identifying assumption, namely, the local continuity assumption. Moreover, results are robust to different specifications, excluding observations far away from threshold and using polynomials of different orders.

The rest of the paper is divided in six sections. Section 2 makes a literature review. Section 3 describes the main aspects of the 2005 pension reform. Section 4 presents the empirical strategy and section 5 the data. Section 6 presents the main results and section 7 concludes.

2 Literature Review

Most of the available empirical evidence about the economic effects of pension systems is for developed countries (Bodor et al., 2008). One of the concerns of the literature is to what extent the design of pension systems affect labor supply. Evidence from Europe shows that pension systems are associated with reductions in both the extensive margin (Gruber and Wise, 1999) and the intensive margin of labor supply (Borsch-Supan, 1998), specially for people near retirement.

Evidence from US is mixed. On the one hand, some contributions show that increasing pension benefits cause reductions in labor supply. Hurd and Boskin (1984) show that non anticipated increases in social security benefits reduces labor participation in 8.2% among old workers; and Burtless (1986) finds that increases in pension benefits increases the probability of retirement in about 2% among 62-65 years old people. On the other hand, Blinder et al. (1981) finds that increases in benefits might increases labor supply of people near retirement. They argue that the negative incentive to work found in the literature, might be avoided with a good actuarial adjustment in which pension benefits increase with additional hours of work.

Krueger and Pischke (1992), and Van der Klaauw and Wolpin (2008) try to understand the mixed evidence. They argue that the effect of pension systems on labor supply depends on income level, age, expectations and the available ways to smooth consumption across life cycle. Interestingly, Van der Klaauw and Wolpin (2008) find that pension reforms affect more low-income individuals.

Danzer (2010) estimates the effect of an increase in the minimal pension on retirement and labor supply choices of people near retirement. It increases retirement rates in almost 17%. The effect is larger for individuals with low education.

Evidence for middle and low income countries is very scarce and Colombia is not an exception. Calderón and Marinescu (2011) look at the affect on informality. The paper finds that both, increasing contributions of independent workers and the unification of the health and pension contribution systems increases informality.

3 The 2005 reform to the Colombian Pension System

The Colombian pension system was created in 1945 with the creation of the Social Security Institute (ISS from its acronym in Spanish), the public pension institute. Until 1993 the ISS managed pension contributions and benefits for workers from the private sector and most workers from the public sector.[1] This public system was a typical fully-funded system in which the pension payments done in a given year are funded by the workers that contribute to the system in that year. As usual contributions and benefits were tied to wage levels. As it is well know this type of system might experience strong funding problems resulting from demographic changes.

The system worked more or less in the same way until the introduction of Law 100 of 1993. This law established a system with two parallel regimes. A fully-funded public regime and a pay-as-you-go private regime. The first is very similar to that introduced in 1945 with come changes in the parameters needed to claim pension benefits (minimum number of weeks in which the worker contributed to the system and retirement age). The second is a system in which the pension benefits of an individual are based on his/her private saving account. Contributors might change from one regime to the other under some conditions.

The 1993 law also established a Transition Regime (TR) for those that had contributed to the pension system for more than 15 years by April 1994 or that were older than 40 (men) or 35 (women) years old. People in the TR could claim retirement benefits under the same conditions that existed before the 1993 reform.

Although there was a reform in 2003, it was only until 2005 (Legislative Act 01 of 2005) that the TR was modified. The 2005 reform changed the conditions to belong to the TR: individuals with less than 750 contributed weeks by January 2005 would not belong any more to the TR and can only claim pension benefits under the rules of the new public system. The 2005 reform also introduced instantaneous increases in the retirement age and the minimum of contributed weeks required to have access to pension benefits as well as reductions in the pension amounts for those individuals who do not belong to the TR.

4 Empirical Strategy

The aim of this paper is to estimate the effect of the 2005 pension reform in Colombia on the intensive margin of labor supply measured as the number of weekly worked hours. We take advantage of the fact that the reform introduces an exogenous threshold to define the

1. Public servants like teachers, army and oil-company workers had independent institutions managing their pension benefits and contributions.

population it affects. This allows us to use a sharp regression discontinuity design (RDD) in order to identify the causal effect of the reform on the intensive margin of the labor supply. The basic idea of the identification strategy is that the probability of being treated changes discontinuously at a given threshold so that individuals that are just to the left or to the right of the threshold are comparable since they are very similar in all other respects. The identifying assumption of this strategy is the so called local continuity assumption. Any difference between the treatment and the control group at the threshold may credibly be attributed to the pension reform if all other covariates, observable and non observable, are continuous at the threshold.

In our case, under the reform, individuals that have contributed for less than 750 weeks by 2005 have to contribute to the system for more time and can retire later than those that contributed for more than 750 weeks. This means that the probability to face longer contribution time and and a higher retirement age is discontinuous with respect to the numbers of weeks contributed to the system. Individuals cannot play around the discontinuity rule since the threshold is established based on the number of contributed weeks in the moment the Law was promulgated (January 2005). Therefore the threshold is exogenous.

Our outcome, y_i, is the number of weekly worked hours. All our population belonged to the Transition Regime (TR) in the moment in which they answer the survey. The assignment variable is the number of contributed weeks by January 2005, Sem normalized by subtracting the threshold to the contributed weeks CW, that is, $Sem = CW - 750$. The treatment, CR_i, is based on the contributed weeks and takes value one for for all individuals with less than 750 contributed weeks by January 2005, and value zero, otherwise. That is,

$$CR_i = \mathbb{1}(CW < 750)$$

Therefore, the treatment CR_i is a discontinuous function of the assignment variable Sem for which the following condition holds

$$Pr[CR = 1|\overline{Sem}^-] \neq Pr[CR = 1|\overline{Sem}^+]$$

where $\overline{Sem}^-$ and $\overline{Sem}^+$ are all individuals that are marginally bellow and above the threshold, respectively.

Under the assumption that individuals just at the left of the threshold are very similar to those just at the right of the threshold in all observable and non observable variables, we expect that they work the same number of weekly hours in absence of the treatment. This allows us to argue that individuals around the threshold have the same probability of being hired. Any difference in the number of worked hours at the threshold might be caused by the pension reform. Formally, the causal effect can be written as,

$$E[\theta|\overline{Sem}^{-}] \equiv E[y_0|\overline{Sem}^{-}] - E[y_0|\overline{Sem}^{+}]$$

The causal effect is identified at the threshold defined by the law. In order to identify this causal effect we estimate a parametric model as follows,

$$y_i = \alpha_0 + f(Sem_i; CR_i) + \theta CR_i + \sigma_t + \mu_m + \eta_i \tag{1}$$

where the polynomial $f(Sem_i; CR_i)$ tries to reproduce how the outcome variable, y_i, is distributed across the assignment variable Sem_i. η_i is the error term. σ_t and μ_m are year and municipality fixed effects, respectively. We report specifications with and without these fixed effects. Coefficient θ is our coefficient of interest. Under the identifying assumption discussed above it identifies the causal effect of the pension reform on the weekly worked hours. The specification implemented to calculate the main results uses a quadratic polynomial,

$$f(Sem_i; CR_i) = \alpha_1 Sem_i + \alpha_2 {Sem_i}^2 + \beta_1(Sem_i * CR_i) + \beta_2({Sem_i}^2 * CR_i) \tag{2}$$

This polynomial includes interactions between the assignment and the treatment variable in order to capture model differences at each side of the threshold.

Following Lee and Card(2008), to have proper robust results, we cluster errors by contributed weeks. Our identification strategy allow for the treatment and control groups to be different on some (or all) covariates, on average. What matters for the covariates is that they should not be different at the threshold. In order to test this assumption we have a complete battery of covariates that might be correlated with our outcome. We introduce each covariate, one by one, in Equation (1). Then, the specification becomes

$$y_i = \alpha + f(Sem_i; CR_i) + \theta CR_i + \gamma X_i + \varepsilon_i \tag{3}$$

where X_i is a covariate. In our battery of covariates we have individual variables (age, gender, years of education, marital status, whether is employee, whether is public servant); household variables (household size, number of children); and market variables (informality and market size). If the identifying assumption holds our θ estimates should remain significant.

We perform several additional robustness checks. First, we test whether our results are driven by observations far away from the threshold (outliers). To do so, we estimate the specification in Equation (1) using three windows around the threshold, namely ±600, ±500,

±400 weeks around the threshold. This restricts the sample around the threshold at the cost of reducing estimation efficiency. Second, we test whether our results are robust to polynomials of different degrees. We test for linear and cubic polynomials with interactions. Third, we test whether our results are driven by outliers in the outcome. We perform the same analysis without individuals within the 5% smallest and 5% largest outcome. Fourth, we perform a falsification test. We estimate the same specification on individuals that belong to the private pension system. Estimates of the coefficient of interest on this sample should be null.

Finally, we look at whether there exists heterogenous effects across sensible groups. To do so we obtain estimates adding an interaction term to Equation (1) between the treatment and a dummy co-variate of interest. To this end we estimate the following specification

$$y_i = \alpha + f(Sem_i; CR_i) + \theta CR_i + \gamma CR_i * Z + \beta Z + \epsilon_i \tag{4}$$

where Z is a dummy variable that takes value one if the individual has some characteristic, and zero, otherwise. Coefficient γ measures the effect above and beyond the average effect captured by θ. We estimate heterogeneous effects by gender ($man = 1$), age ($old = 1$), years of education, marital status ($single = 1$), household size ($large = 1$), number of children ($large = 1$), employee, civil servant, informality at municipality level and market size. We expect older and employee individuals to be more sensitive to the reform.[2]

5 Data

We use the Colombian Household Survey (GEIH from its acronym in Spanish) from the National Statistics Office (DANE from its acronym in Spanish) for years 2007 and 2008. This survey gathers data on individual characteristics (age, gender, birth date, marital status and years of education.) and household characteristics (household size and number of children). For occupied individuals, it gathers data on pension regime, contributed weeks, job type, hours of work, among others.

The sample is conformed by individuals that are affiliated to the Public Pension System (ISS from its acronym in Spanish) and that used to belong to the Transition Regime (TR) before 2005. We use birth date in order to calculate the individual age at the moment Law 100 began to rule (April 1994). The sample has 9901 observations. 27.44% of individuals left the TR due to the 2005 reform (treatment group).

In Table 1 we report the descriptive statistics of the variables used in the empirical analysis. As a measure of the intensive margin of labor supply we use the number of hours

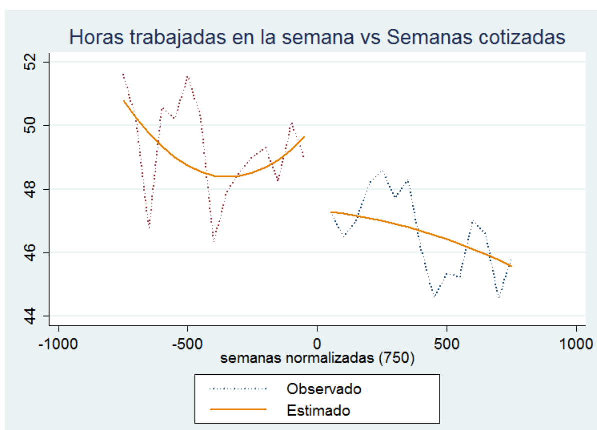
2. In addition, there is some evidence for Colombia showing that the education level might be positively correlated with savings. See Guataquí, et.al(2009) and Arango and Posada(2003).

worked per week.[3] On average individuals work almost 47 hours per week.[4] On average treated individuals offer around 2.5 hours more than non treated individuals. As expected, treatment and control groups are very different regarding the assignment variable, namely, the number of contributed weeks. On average for the whole sample individuals have contributed during 957 weeks by the moment the 2005 reform begin to rule. Treated individuals have contributed 451 weeks on average and non treated individuals have done so for 1150 weeks.

As covariates we use variables that might affect individual work choices (Meghir y Phillips, 2008). We report the number of observations, mean, standard deviation, minimum and maximum. For each variable we report descriptive statistics for the whole sample, treatment group and non treated group. Individuals in the sample are 54 years old on average, so they are near to the retirement age. 46% of those individuals are men. Household size is about 4 persons with 3 children on average. In addition, 80% of the sample are employees (wage earners).[5] Independent workers represent around 20% of the sample. 43.1% of individuals work with the government.

In Figure 1 we can see both observed and fitted data (with a quadratic polynomial) in the outcome - assignment Cartesian plane. The number of worked hours per week seems to decrease with the number of contributed weeks. It seems that there is a discontinuity of worked hours at the threshold.

Figure 1: Weekly worked hours and contributed weeks



Source: DANE-GEIH. Authors' calculations. Fitted data uses quadratic polynomial.

3. We use the question that ask individuals to report the number of hours worked in the week before the survey was undertaken.

4. We excluded the 1% of observation at each tail of the distribution.

5. Employees include workers in the private sector, the public sector and domestic employees.

Table 1: Descriptive Statistics

		N	Mean	Sd	Min	Max
Weekly Work Hours		9901	47.1	13.5	10	98
	Treated	2815	48.9	15	10	98
	Non Treated	7086	46.5	12.9	10	98
Contributed Weeks		9901	965.7	386.5	0	1500
	Treated	2815	451.1	193.8	0	700
	Non Treated	7086	1170.1	216.3	800	1500
Age		9901	54.3	4.8	41	95
	Treated	2815	55.4	5.1	46	95
	Non Treated	7086	53.9	4.6	41	85
Men		9901	0.5	0.5	0	1
	Treated	2815	0.4	0.5	0	1
	Non Treated	7086	0.5	0.5	0	1
Education		9901	6.5	3.1	0	15
	Treated	2815	6.2	3.2	0	15
	Non Treated	7086	6.6	3.1	0	15
HH size		9901	3.8	1.9	1	20
	Treated	2815	3.9	2	1	18
	Non Treated	7086	3.8	1.8	1	20
Children		9901	3.4	1.5	1	14
	Treated	2815	3.4	1.6	1	12
	Non Treated	7086	3.4	1.5	1	14
Single		9901	0.1	0.3	0	1
	Treated	2815	0.1	0.3	0	1
	Non Treated	7086	0.1	0.3	0	1
Employee		9868	0.8	0.4	0	1
	Treated	2796	0.7	0.5	0	1
	Non Treated	7072	0.9	0.4	0	1
Civil servant		7656	0.4	0.5	0	1
	Treated	1720	0.3	0.5	0	1
	Non Treated	5936	0.5	0.5	0	1
Market size (% occupies)		9901	2.9	1.4	0	4.7
	Treated	2815	2.9	1.4	0	4.7
	Non Treated	7086	2.9	1.4	0	4.7

Source: GEIH 2007 and 2008. Sample: 9901 affiliated to ISS that belonged to TR before 2005. Treated: Individuals with less than 750 contributed weeks by January 2005. Non Treated: Individuals with 750 contributed weeks or more by January 2005.

6 Results

Table 2 reports estimates of θ in Equation 1. Column (1) presents estimates with no fixed effects. Column (2) includes municipality fixed effects. Column (3) has year fixed effects. Column (4) has both municipality and year fixed effects. Results show that the pension reform seems to have increased labor supply in 2.1 hours. The result is robust (and very stable) across different specifications. Treated individuals can retire later and have to contribute during more time to retire. This makes treated individuals increase labor supply in the intensive margin. The effect is about 14% of a standard deviation and 4.3% of the number of worked hours per week, on average. The effect is thus economically important, which reflects the key role of the pension system in the labor market.

Table 2: Main Results

	(1)	(2)	(3)	(4)
CR	2.128**	2.122**	2.129**	2.127**
	(0.815)	(0.797)	(0.774)	(0.775)
Municipality FE		✓		✓
Year FE			✓	✓
Observations	9,901	9,901	9,901	9,901

Clustered standard errors by the number of contributed weeks in parenthesis.
*** p<0.01, ** p<0.05, * p<0.1

Let us now test whether the identifying assumption holds. We introduce the covariates, one by one. Results are reported in Table 3. We can see there that estimates of the coefficient of interest are robust to the introduction of covariates, and is positive and significant in all specifications. This shows that other covariates are continuous at the threshold and that the discontinuity in the outcome can be attributed to the pension reform.

We report the results of heterogeneous effects in Table 4. Recall that the variables are dummies. To transform continuous variables into dummies we use the criteria explained in the note. All criteria are equal or around the median of the respective variable. All interactions show non significant coefficients except for age and employee. Indeed, when we introduce these last two interactions the average effect becomes non significant. This means that the pension reform only has an effect on old employees. This population has a shorter time length to achieve the minimum contributed weeks needed to obtain pension benefits so they are more likely to increase the intensive margin of labor supply.[6] Interestingly, the effect is the same independently of education level, household size or number of children.

6. In table A.1 in the Appendix, we report the main results restricting the sample to employees. Results for employees are very parsimonious with results in Table 2. Average effect for employees is a bit smaller than two hours.

Table 3: Testing local continuity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
CR	2.452***	2.333***	2.106**	2.147**	2.121**	2.116**	2.160**	2.139**	2.419***
	(0.803)	(0.821)	(0.814)	(0.835)	(0.815)	(0.819)	(0.835)	(0.819)	(0.804)
Individual characteristics									
Age	0.208***								
	(0.0440)								
Gender (men)		5.910***							
		(0.379)							
Years of education			-0.111**						
			(0.0537)						
Marital status (single)				-0.933**					
				(0.370)					
HH characteristics									
HH size					0.246***				
					(0.0774)				
Children						0.289***			
						(0.0994)			
Market characteristics									
Informality							-0.124***		
							(0.0130)		
Market size								0.332***	
								(0.110)	
Job type									-3.312***
									(0.606)
Observations	9,901	9,901	9,901	9,901	9,901	9,901	9,901	9,901	9,901

Clustered standard errors by the number of contributed weeks in parenthesis. All regressions include both municipality and year fixed effects. *** p<0.01, ** p<0.05, * p<0.1

Table 4: Heterogeneous Effects

variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
CR	1.835*	2.105**	1.279	2.071**	2.305**	2.264**	2.547**	-0.189
	(0.985)	(0.817)	(1.063)	(0.824)	(0.871)	(0.933)	(1.051)	(1.340)
CR * Gender	1.157							
	(0.768)							
CR * Marital Status		0.315						
		(0.716)						
CR * Age			1.728**					
			(0.768)					
CR * Education				0.0397				
				(0.766)				
CR * HH size					-0.324			
					(0.746)			
CR * Children						-0.198		
						(0.711)		
CR * Market size							-0.829	
							(0.803)	
CR * Employee								3.409***
								(1.062)
Observaciones	9,901	9,901	9,901	9,901	9,901	9,901	9,901	9,901

Clustered standard errors by the number of contributed weeks in parenthesis. All regressions include both municipality and year fixed effects. *Gender* takes value one for men. *Marital status* takes value one for single. *Age* takes value one if 54 years old or more. *Education* takes value one if 5 years of education or more. *Household size* takes value one if 4 persons in HH or more. Children takes value one if 3 children in HH or more. *Market size* takes value one if proportion of occupies at the municipality level is above the median. *Employee* takes value one for employees in private sector, public sector and domestic service. *** $p<0.01$, ** $p<0.05$, * $p<0.1$

We make three robustness checks. First, Table 5, Panel A, reports the sensibility of our results to outliers. To do so we estimate the same regression in Equation (1) using windows of different sizes around the threshold. A limit of this test is that the smaller the window, the smaller the efficiency and coefficients can become non significant. Results for windows of ±600, ±500 and ±400 around the threshold are reported in columns (1), (2) and (3), respectively. All estimates are positive and significant. Reduction of efficiency with smallest window makes estimates to be significant at 90% level.

Table 5: Robustness checks

		(1)	(2)	(3)
		(-600, 600)	**(-500, 500)**	**(-400, 400)**
Panel A: Sensibility to outliers	CR	2.551**	4.131***	2.357*
		(0.964)	(1.178)	(1.180)
	Observations	8,197	6,634	5,446
		(1)	**(2)**	**(3)**
		Linear	Quadratic	Cubic
Panel B: Polynomial order	CR	1.149*	2.127**	2.907**
		(0.566)	(0.775)	(1.252)
	Observations	9,901	9,901	9,901
		(1)	**(2)**	**(3)**
		Benchmark	No Outliers	Falsification Test
Panel C: Outcome outliers and Falsification test	CR	2.127**	2.659***	-0.105
		(0.775)	(0.637)	(0.784)
	Observations	9,901	9,099	9,484

Clustered standard errors by the number of contributed weeks in parenthesis. All regressions include both municipality and year fixed effects. *** p<0.01, ** p<0.05, * p<0.1

Second, Table 5, Panel B, reports robustness checks to polynomial order. We estimated different specifications with interactions for linear, quadratic and cubic polynomials. Estimates are reported in columns (1), (2) and (3), respectively. All estimates are positive and significant. The linear model fails to capture some non linearities that are particularly important at the left of the threshold (See Figure 1).

Third, Table 5, Panel C, reports robustness checks to outcome outliers and a falsification test. Column (1) shows benchmark results. Column (2) reports results taking out 5% smallest and 5% largest outcomes. Column (3) reports results of making the same empirical exercise on individuals in the private pension system. Our main results are not driven by outliers in the outcome and individuals in the private pension system have no discontinuities at the threshold as expected.

On top of verifying that the local continuity assumption holds, results are robust to different specifications, are not driven by outliers in the assignment variable, nor depend on the degree of assignment variable polynomials. Besides, results are not driven by outliers in the outcome and the falsification test shows null results as expected. Therefore, we can credibly attribute the outcome discontinuity to the pension reform and argue that it causes an average increase of 2 hours in the number of worked hours per week.

7 Final Remarks

The Colombian pension system moved from a fully-funded public system to a system where two regimes coexist and compete, namely, a public fully-funded system and a pay-as-you-go system. In the process that began with Law 100 of 1993, a Transition Regime (TR) was introduced allowing individuals that where relatively old and near retirement to keep the conditions of the old fully-funded system. The 2005 reform (by Acto Legislativo 01 of 2005) increased retirement age and the number of weeks a worker has to contribute to the system in order to claim pension benefits. It also established a threshold of 750 contribution weeks, so that only individuals bellow the threshold were affected by the reform. We take advantage of that threshold to implement a parametric sharp RDD in order to identify the effect of the parametric pension reform on the intensive margin of labor supply measured as worked hours per week.

Our findings show that two-three years after the reform, those individuals under tougher pension conditions work more. The average effect is around two hours per week. This effect is economically important since it represents around 14.8% of the standard deviation and 4.2% of the average of the worked hours per week. The average effect is concentrated on relatively old employees and is independent of education level, household size and number of children. It is consistent with the local continuity assumption. Moreover, results are robust to different specifications, excluding observations far away from threshold, using polynomials of different orders, excluding outcome outliers and a falsification test.

Some policy recommendations are in order. Changing the parameters of fully funded defined benefits pension programs is sometimes seen as a reform that help to make these programs more sustainable. Our results show that this type of reform also constitutes incentives for labor supply which can also be seen as a desirable effect of these type of reforms. However, it may also be the case that such a reform reduces labor market participation (the extensive margin). If this is so the exact reform has to carefully balance both effects.

References

[1] Attanasio, O.P. and Rohwedder, S. (2003). “Pension wealth and household saving: Evidence from pension reforms in the United Kingdom”. American Economic Review. 93. 1499-1521.

[2] Barr, N.A. and Diamond, P.A. (2008). “Reforming pensions: Principles and policy choices”. Oxford University Press, USA.

[3] Battistin, E.; Brugiavini, A.; Rettore, E. and Weber, G. (2009). “The retirement consumption puzzle: Evidence from a regression discontinuity approach”. The American Economic Review. 99. 2209-2226.

[4] Blinder, A.S., Gordon, R.H. and Wise, D.E.(1981). “Reconsidering the work disincentive effects of social security”. National Bureau of Economic Research Cambridge, Mass., USA.

[5] Bodor, A., Robalino, D. and Rutkowski, M. (2008). “How Mandatory Pensions Affect Labor Supply Decisions and Human Capital Accumulation? Options to Bridge the Gap between Economic Theory and Policy Analysis”. University Library of Munich, Germany. MPRA Paper 12046.

[6] Borsch-Supan, A. (1998) “Incentive effects of social security on labor force participation: evidence in Germany and across Europe”. National Bureau of Economic Research.

[7] Burtless, G. (1986). “Social security, unanticipated benefit increases, and the timing of retirement”. The Review of Economic Studies. 53. 781-805.

[8] Calderón, V. and Marinescu, I. (2011) “The Impact of Colombia’s Pension and Health Insurance Systems on Informality”. Inter-American Development Bank.

[9] Corbo, V. and Schmidt-Hebbel, K. (2003). “Efectos macroeconómicos de la reforma de pensiones en Chile”. Resultados y desafíos de las reformas de pensiones.

[10] Danzer, A. M. (2010). “Retirement Responses to a Generous Pension Reform: Evidence from a Natural Experiment in Eastern Europe”. Institute for the Study of Labor (IZA). IZA Discussion Papers 4726.

[11] Friedberg, L. (2000).“The labor supply effects of the social security earnings test”. Review of Economics and Statistics. 82(1). 48-63. MIT Press.

[12] Gruber, J. and Wise, D. (1999) “”Introduction to ”Social Security and Retirement around the World”. National Bureau of Economic Research. 1-35.

[13] Hahn, J., Todd, P. and Van der Klaauw, W.(2001). "Identification and estimation of treatment effects with a regression-discontinuity design". Econometrica. 69(1). 201-209.

[14] Hurd, M.D. and Boskin, M.J. (1984). "The effect of social security on retirement in the early 1970s". Quarterly Journal of Economics. MIT Press. 99(4). 767-790.

[15] Imbens, G. and Kalyanaraman, K. (2009). "Optimal bandwidth choice for the regression discontinuity estimator". National Bureau of Economic Research.

[16] Imbens, G.W. and Lemieux, T. (2008). "Regression discontinuity designs: A guide to practice". Journal of Econometrics. 142(2). 615-635.

[17] Krueger, A.B. and Pischke, J.S.(1992). "The effect of social security on labor supply: A cohort analysis of the notch generation". Journal of Labor Economics. University of Chicago Press. 10(4). 412-437.

[18] Lee, D.S. and Card, D. (2008). "Regression discontinuity inference with specification error". Journal of Econometrics. 142(2). 655-674.

[19] Mesa Lago, C. (2002). "La reforma estructural de las pensiones de seguridad social en América Latina: modelos, características, resultados y lecciones". Economía y Sociedad. 19.

[20] Mesa-Lago, C.(2004). " Evaluación de un cuarto de siglo de reformas estructurales de pensiones en América Latina". Revista de la CEPAL. 84. 59-82.

[21] Van der Klaauw, W. and Wolpin, K.I. (2008) "Social security and the retirement and savings behavior of low-income households". Journal of Econometrics. 145. 21-42.

A Results for Employees

Table A.1: Main Results for Employees

	(1)	(2)	(3)	(4)
CR	1.949**	1.942**	1.953**	1.950**
	(0.821)	(0.782)	(0.828)	(0.811)
Municipality FE		✓		✓
Year FE			✓	✓
Observations	7,946	7,946	7,946	7,946

Clustered standard errors by the number of contributed weeks in parenthesis.
*** p<0.01, ** p<0.05, * p<0.1