The implications of a rise in the minimum wage on the Mexican labour market.

Jorge Bouchot
University of Birmingham

Abstract

This paper performs a comprehensive evaluation of the implications of a rise in the minimum wage on the Mexican labour market. The study estimates the impact on real wages, the distribution of earnings, employment, and informal employment, using as a natural experiment the partial harmonization of regional minimum wages in 2012. Correcting for sample selection bias, there is no evidence of adverse effects on the labour market; the estimates suggest positive effects on real wages, employment, and occupation in the formal sector. The model, which is robust to different specifications and to the control group choice, supports the existence of monopsonistic labour markets in Mexico. In addition, by the use of unconditional quantile regressions, the distributional wage effects suggest a small improvement on earnings for the targeted lowest income workers, although spillover effects on the upper percentiles of the distribution widen the wage dispersion.

JEL Classification: J23, J32

Keywords: minimum wage, employment, informal sector, earnings distribution, selection bias, unconditional quantile regression

1 Introduction

Until November 2012, Mexico was divided into three minimum wage zones: A, B, and C. This paper exploits a natural experiment where the minimum wage in Zone B was unexpectedly increased by 2.9 percent to bring it into line with that in Zone A. I evaluate, in a comprehensive way, the labour market consequences of this intervention, analysing the effect on real wages, the earnings distribution, the overall level of employment, and occupation in the informal[1] labour market.

[1]Throughout the paper the terms informal workers, informal sector or informal labour market are used indistinctly for all those workers whose labour relation is not recognized by the employer.

By the implementation of difference in differences procedures and correcting for sample selection bias, the paper provides evidence that the minimum wage increase had no adverse effects on the labour market. On average, real wages increased, the level of employment did not diminish — actually evidence is found that employment rose for the eldest in the labour force — and participation in the informal sector decreased. Nevertheless, examining the impact on the earnings distribution by unconditional quantile regressions, the results show an important minimum wage *lighthouse effect*; that is, the policy intervention does not necessarily affect the targeted workers, and there are considerable spillover effects on the upper side of the distribution, which increases the wage dispersion.

The Mexican labour market has some particular features related to the institutional minimum wage setting and to the workforce composition, which are worth highlighting because of their relevance for identifying and interpretating of the coefficient estimates presented below.

First, as a consequence of recurrent episodes of macroeconomic crises in the 80's and 90's, the real minimum wage value decreased by around 70% in the last 30 years. Therefore, the current minimum wage does not accomplish its primary goal of satisfying the most basic requirements of the workers.[2]

According to data from the Organisation for Economic Co-operation and Development (OECD), the annual real minimum wage in Mexico corresponded to $1,750 PPP US dollars in 2013, the lowest value among the OECD country members. Moreover, comparing the real minimum wage variation among Latin American countries for the period 2000-2013, statistics from the United Nations Economic Commission for Latin America and the Caribbean (ECLAC) reveal that the real minimum wage in Mexico has among the lowest in the region, only above that in Venezuela.[3]

Second, with respect to the institutional minimum wage setting there are two elements to consider. On the one hand, the main challenge to evaluating the consequences of a minimum wage rise in Mexico is the lack of significant statistical variation in the minimum wage. During the last twenty years, the real minimum wage has kept almost constant and, even though there were different wage zones, the annual nominal increases have been kept proportional between them.[4] So, in the ab-

[2]Taking as a reference the poverty lines from the National Council for Evaluation of Social Development Policy, workers earning the minimum wage do not reach the 'welfare line', which is the monetary value of food and services as clothing, housing and transport. Considering a worker with one economic dependent, minimum wage is not enough for both individuals to reach the extreme poverty line ('minimum welfare line') corresponding to the monetary cost of the basic caloric intake.

[3]Based on own calculations with data retrieved from OECD and ECLAC websites: `https://stats.oecd.org/Index.aspx?DataSetCode=RMW`, `http://estadisticas.cepal.org/cepalstat/web_cepalstat/estadisticasIndicadores.asp?idioma=i`

[4]Minimum wage literature in United States has longly exploited the state variation of minimum

sence of structural changes and regional differentiation, the intervention in November 2012 — although it only increases minimum wage in only 2.9% in nominal terms — represents the best opportunity to identify the impact of a minimum wage change on the labour market.

On the other hand, minimum wages in Mexico also work as a reference price in the labour market. Remunerations in the formal and informal sector are tied to multiples of these minimum wages (Fairris et al., 2008). In addition, until 2016 many government-set prices were tied to multiples of the nominal minimum wage. These included, for instance, pensions, labour bonuses or benefits, social security fees, income brackets for income tax rates, fines, among others.[5] Therefore, changes to the minimum wage are likely to affect not only the lowest paid workers. The findings suggest there are sizable effects on the top deciles of the earnings distribution.

And third, regarding the composition of the Mexican labour force, the paper puts special emphasis on the high level of participation in the informal sector. Informal employment Latin America has been one of the main obstacles to quantifying the impact of minimum wage policies. This study draws on data provided by the National Survey on Employment and Occupation (ENOE) to incorporate the informal sector into the evaluation, which represents almost 60% of the labour force in Mexico. The estimates show that the policy intervention also affects remunerations in the informal sector and, more importantly, it generates a reduction in the informal employment.

In addition, the sociodemographic characteristics of Mexican workers on the minimum wage are different from the characteristics observed in developed countries. For instance, the minimum wage literature in the United States has largely focused on its impact for the youngest segment of the workforce, while in Mexico 69% of workers on or below the minimum wage are aged 30 or older. Similarly, minimum wage workers in developed countries do not tend to live in poverty and they are not normally economically responsible for their households. In contrast, in the Mexican case, 42% of minimum wage workers represent the main source of household income, and more than 96% of them have no access to social security services.[6] Under this framework, minimum wage legislation has important social implications for its effects on this vulnerable segment of the population.

All these features of the Mexican labour market are incorporated into the research

---

wage levels (see Card (1992b), Card and Krueger (1994), and Dube et al. (2010) for the most influential papers.) In contrast, this regional variation is not observed in Mexico, where minimum wage is set by a single institution, the National Commission on Minimum Wage (CONASAMI).

[5]In January 2016 Mexican Congress passed a reform called 'de-indexation of the minimum wage', in which minimum wage is not tied anymore to any price.
http://www.dof.gob.mx/nota_detalle.php?codigo=5423663&fecha=27/01/2016

[6]Own calculations with data from ENOE, corresponding to the third quarter of 2012.

design to evaluate the effects of the 2012 minimum wage harmonization. The impact is estimated for the whole population, and it is also disaggregated into age thresholds, income quantiles, by different types of labour status, and by informality condition.

The econometric models rely on two empirical tools. On the one hand, the mean effects on real earnings, employment and informality are tested and corrected for sample selection bias using the procedure introduced by Heckman (1979).[7] On the other hand, in order to estimate the wage effects across the whole earnings distribution, as well as the spillover effects, uncondotional quantile regressions are implemented. The novel uncondotional quantile regression procedure, developed by Firpo et al. (2009), offers the analytical advantage of estimating marginal treatment effects at different points of the distribution.

Thus, using quarterly data from the ENOE for 2012Q1-2013Q4, pooled difference in differences models suggest that minimum wage harmonization in 2012 had on average a positive and statistically significant impact on real wages. However, this positive impact is, paradoxically, at their weakest for the lowest percentiles of the distribution. On the employment side, there is no statistical evidence that the Zone B minimum wage increase affected the probability of being active in the labour market, but the probability of being employed increased. The results also suggest that the intervention caused a switch from informal to formal employment, specially for the eldest in the labour force.

The results are robust to different econometric specifications, particularly to the control group used, and to the exact specification of the difference in differences variable. Our analyses provide solid empirical evidence on positive employment effects (and also on formal employment) of the minimum wage rise, albeit the institutional setting framework led to spillover effects that increased the wage dispersion.

The rest of the paper is organised as follows: Section 2 presents a review of the minimum wage literature. The data and some descriptive statistics are presented in Section 3, as well as the description of the minimum wage intervention used as a natural experiment. The details of the econometric specifications, the process of sample selection bias correction, and the unconditional quantile regression are specified in Section 4. The resulting estimates are presented in Section 5, describing the evaluation of the impact on real wages, its distribution, employment and informal employment, also including models of the dynamics of the effects. Finally, Section 6 concludes the document.

[7] Campos et al. (2015) used the same source of variation, the minimum wage harmonization of 2012, finding no evidence of negative effects on income and employment, while some positive effects on the formal sector are estimated. Nevertheless, their evaluation restricted the sample to the segment of the labour force actually working, generating the possibility of biased estimates.

2 Literature review

In the economics literature there has been a long and controversial debate on the effects of changes in minimum wages on the labour market. Until now, a consensus of the impact of the increase in minimum wages on employment has not been reached. Even though the discussion has focused mainly on the employment effects, there is also a vast literature on the distributional impacts of minimum wages arguing the existence of wage spillover effects.

At the theoretical level, the standard competitive model (in its simplest version) assumes that labour force is homogeneous in terms of workers' skills. Given that output depends exclusively on the quantities of inputs used, not on the workers productivity, wage rate is the same for all workers. Firms optimize profits given the wage determined by the market. In consequence, there must be an inverse relationship between wages and employment.

There are several extensions to the basic model incorporating some kind of heterogeneity, for example, the degree of compliance of the minimum wage regulations (Welch, 1974), the effect on the unemployed sector (Mincer, 1976; Gramlich et al., 1976), and different levels of skills among the labour force (Brown et al., 1982). In all of these versions, the model unambiguously anticipates employment reductions as a result of minimum wages increases, at least for the directly affected workers.

Card and Krueger (1995) asserted that although the standard economic theory is quite powerful in explaining labour markets, it is incomplete for specific contexts where firms are not necessarily price takers. In fact, it seems logical to assume that firms have some kind of power on wage setting. Thus, models of monopsonistic labour markets have been widely used as the theoretical approach to support findings of non-negative effects of minimum wage increases on employment. The basic idea behind monopsony is the following: given that the wage of equilibrium may be higher than the marginal cost, increases to the minimum wage do not necessarily reduce employment.

In dynamic monopsonic models (Manning, 2004; Flinn, 2006; Hirsch et al., 2015), the source of the discretionary wage setting by the firms can be originated by several kinds of frictions in the labour market. For instance, hiring costs, spatial location of the firms, bargaining search, among others. Under this framework, minimum wage increases lead to a reduction of the turnover costs for the workers with lowest wages, so employment level is not unquestionably lowered. Given that our analysis does not aim at investigating what is the source of the discretionary setting, the term 'monopsony' is used indistinctly from 'dynamic mononopsony'.

The empirical minimum wage literature is extensive, specially for the US. Between 1961 and 1981, seventeen nominal increases to the federal minimum wage were recorded. Exploiting this variation, time-series studies achieved an apparent consensus: an increase of 10 percent in the federal minimum wage reduces employment by between 1 and 3 percent (Brown et al., 1982).

But in the decade of the 90's, as a result of differentiated labour policies at the state level in the US, a new wave of cross-section studies started to challenge the previous findings, claiming that the effect on the employment level was not negative and could be even positive. Thus, Card (1992a,b) and Card and Krueger (1994, 1995) represented the turning point in the literature. Afterwards, the discussion on the changes in the minimum wage policy in the United Kingdom brought new cross-sectional evaluations that supported the findings by Card and Krueger (Machin and Manning, 1994; Dickens et al., 1999; Stewart, 2002).

However, there is still the opposite position that asserts that, those cross-sectional and panel-data analyses fail to capture the whole long-term effect on employment, arguing also that those results are only applicable for some specific sectors. To mention just some of the most relevant studies, Brown et al. (1995), Neumark (2001) and Sabia (2009) found negative and significant effects on employment.

In a new wave of studies, Dube et al. (2010) (which in the words of Schmitt et al. (2013) is probably the 'most important and influential paper in minimum wages' during the last decade) developed a different way to construct control groups. They identified all the contiguous counties in the US with different minimum wage settings at the state level from 1990 to 2006, estimating no adverse effects on employment.

With respect to the distributional effects of the minimum wage, Lee (1999) analysed the impact of the real minimum wage erosion in the US during the decade of the 80's on the increasing inequality. By the use of state variation in the relative level of minimum wages for the period 1979-1991, Lee found that a considerable part of the wage gap between the 10th and 50th percentiles was almost totally explained by the fall of the real value of minimum wages. The effect was estimated stronger and more convincing for the female wage distribution.

In a reassessment of the significance of the minimum wage on earnings inequality in the US, Autor et al. (2016) enlarged the analyses by Lee (1999) in two senses. They extended the period of analysis to 2012, and tested the minimum wage spillover effects for workers earning above the minimum, concluding that the significant impacts for higher percentiles estimated in previous literature were upward biased by errors of measurement. Once that the bias is purged by instrumenting effective minimum wage with legislated minimum wage, their results suggest that minimum wage effectively affects the wage dispersion, but only in the lower tail of the distribution: up to the 25h percentile for woman and up to the 10th percentile for men.

For the French labour market, Aeberhardt et al. (2015) evaluated the minimum wage changes between 2003 and 2005 — generated by the working time reduction law — on the earnings distribution. Using also unconditional quantile regressions, they found that the minimum wage increase had impacts over a large part of the distribution: up to the seventh decile for men and up to the fifth decile for women.

Regarding the empirical analyses implemented in Mexico and Latin America, it is necessary to separate them from the minimum wage literature in developed countries; the particular features of Latin American labour markets make it even more difficult to obtain generalised results of the effects of minimum wages. Neumark and Wascher (2006) argued that the main complications to implement these evaluations are the high presence of informal labour markets and the lack of full enforcement of the minimum wage regulations.

There is also a high degree of heterogeneity in the estimates. For instance, for the Brazilian case Lemos (2009) found no adverse effects on employment, while for the case of Colombia, both Bell (1997) and Maloney and Mendez (2004) estimated strongly negative impacts. Moreover, there are cases of mixed results within the same labour market; Montenegro and Pages (2004) concluded that minimum wages reforms in Chile reduced the employment level for youth and unskilled workers, but generated a positive impact on employment among women.

For the specific case of Mexico, there are three relevant studies. In the first of them, Bell (1997) used both, time series and panel data models (data from 1972 to 1990 and 1985 to 1990, respectively), finding no effect on employment for manufacturing firms. Bell argued the results are explained by the fact that minimum wage was not an effective wage, at least on the analysed data.

For a similar period, and using the large reductions in real minimum wage levels as a source of variation, Feliciano (1998) found employment increases for women workers between 15 and 64 years of age, but adverse effects for men between 55 and 64 years old. Feliciano asserted that unlike Bell (1997), she found significant effects on employment because her analyses explored the impact on the whole labour market, and not only for the manufacturing sector.

In a more recent paper, using also the regional variation observed in November 2012, Campos et al. (2015) estimated cross-sectional and panel-data regressions concluding that there is no empirical evidence of adverse effects generated by the minimum wage increase. Nevertheless, there are important methodological issues to be analysed, especially to consider the presence of sample selection bias. In addition, there are important implications which need to be studied more deeply, like the differences in the wage impact of the minimum wage reform on the formal and informal labour market, as well as the effect on the distribution of earnings.

Related to the distributional effects, Bosch and Manacorda (2010) demonstrated that real minimum wage reduction in Mexico explained most of the the growth of earnings inequality for the period 1989-2001, finding also evidence that minimum wage can affect earnings up to the sixth decile of the earnings distribution, but they failed to find a significant effect on informal workers. Nevertheless, the main caveat of their analysis is that in absence of a structural change in the minimum wage for that period, they instrumented the called 'effective minimum wages' using social security data. Thus, they took the erosion of the real minimum wage as exogenous, which is not necessarily true. Moreover, the database used was restricted to urban areas, impeding to observe a significant segment of the informal labour market.

3 Data and descriptive statistics

3.1 The data

The dataset for the empirical analysis is obtained from the ENOE, which constitutes the official source on the dynamics of the Mexican labour market, including quarterly information on participation rates, unemployment, and informal labour market participation. It is produced quarterly by the National Institute of Statistics and Geography (INEGI), sing a sample size fixed at 120,060 dwellings.

The period relevant for the purposes of our study is from 2012Q1 to 2013Q4, although some robustness exercises presented at the end of Section 5 use data from 2011Q1 to 2014Q4 aiming at including two years before and after the minimum wage reform of November 2012.

The sample includes individuals between 12 and 97 years old.[8] Even though the legal minimum age for performing working activities was 14 years old,[9] ENOE collects labour information on individuals from 12 years old. By definition, workers aged younger than 14 belonged to the informal sector, so for the objectives of our analysis it is necessary to include them in the sample. Thus, the total number of observations for the eight quarters analysed is 2,458,053.

Given that our models evaluate the effects of minimum wage harmonization on wages, its distribution, employment, and informality, different variables are used as a dependent variables. For earnings, all the specifications use the logarithm of real hourly wages,[10] while for the labour status variables, the model estimates the effect on active population and employed workers.

[8]Observations with age equal to 98 or 99 denote non-specified age.

[9]In June 2015, the minimum compulsory age for working increased from 14 to 15 years old. http://www.diputados.gob.mx/LeyesBiblio/ref/lft/LFT_ref27_12jun15.pdf

[10]Appendix C presents alternative estimates on real monthly wages.

Table 1
Sample descriptive statistics. 2012Q1-2013Q4

	Pre-treatment period			**Post-treatment period**			**National**
	Zone A	**Zone B**	**Zone C**	**Zone A**	**Zone B**	**Zone C**	
Dependent Variables							
Wages							
Hourly real wage							
Mean	34.45	34.72	29.62	33.72	35.64	29.23	30.52
Std. Deviation	45.99	44.19	65.04	95.94	45.61	43.47	56.95
Observations	53,599	42,188	348,085	62,598	47,951	406,747	961,168
Monthly real wage							
Mean	5,545.35	5,572.03	4,755.33	5,446.71	5,668.47	4,734.89	4,917.16
Std. Deviation	5,369.97	5,463.97	5,448.21	7,514.70	5,170.70	5,431.49	5,592.36
Observations	53,599	42,188	348,085	62,598	47,951	406,747	961,168
Labour Status							
Active labour market population among these aged 12 to 97							
Mean	0.5827	0.5864	0.5704	0.5821	0.5821	0.5683	0.5723
Std. Deviation	0.4931	0.4925	0.4950	0.4932	0.4932	0.4953	0.4947
Observations	135,589	107,607	879,955	163,630	126,651	1,044,621	2,458,053
Employed population among these aged 12 to 97							
Mean	0.9399	0.9436	0.9534	0.9431	0.9460	0.9525	0.9507
Std. Deviation	0.2377	0.2306	0.2108	0.2316	0.2261	0.2128	0.2165
Observations	79,014	63,097	501,951	95,253	73,723	593,703	1,406,741
Informal Sector							
Informal workers among employed population							
Mean	0.5005	0.4498	0.5849	0.4835	0.4319	0.5772	0.5561
Std. Deviation	0.5000	0.4975	0.4927	0.4997	0.4953	0.4940	0.4968
Observations	74,263	59,540	478,544	89,836	69,740	565,486	1,337,409
Waged informal among formal workers							
Mean	0.3735	0.3179	0.4292	0.3577	0.2992	0.4224	0.4052
Std. Deviation	0.4837	0.4657	0.4950	0.4793	0.4579	0.4939	0.4909
Observations	59,208	48,028	348,022	72,238	56,535	413,991	998,022
Self-employed informal among formal workers							
Mean	0.2518	0.2268	0.3314	0.2427	0.2185	0.3260	0.3063
Std. Deviation	0.4340	0.4188	0.4707	0.4287	0.4133	0.4688	0.4610
Observations	49,572	42,368	297,108	61,262	50,697	354,788	855,795
Non-waged informal among formal workers							
Mean	0.0649	0.0549	0.1390	0.0556	0.0509	0.1303	0.1151
Std. Deviation	0.2464	0.2278	0.3459	0.2292	0.2199	0.3366	0.3192
Observations	39,667	34,664	230,724	49,128	41,744	274,931	670,858
Sociodemographic Controls							
Head of the Household							
Mean	0.3511	0.3402	0.3375	0.3504	0.3417	0.3383	0.3398
Std. Deviation	0.4773	0.4738	0.4729	0.4771	0.4743	0.4731	0.4736
Observations	135,589	107,607	879,955	163,630	126,651	1,044,621	2,458,053
Female							
Mean	0.5155	0.5154	0.5248	0.5201	0.5163	0.5249	0.5232
Std. Deviation	0.4998	0.4998	0.4994	0.4996	0.4997	0.4994	0.4995
Observations	135,589	107,607	879,955	163,630	126,651	1,044,621	2,458,053
Age							
Mean	37.52	37.84	37.23	37.64	38.12	37.27	37.36
Std. Deviation	17.87	18.14	18.32	17.86	18.18	18.26	18.23
Observations	135,567	107,501	879,645	163,602	126,481	1,044,201	2,456,997
Rural							
Mean	0.0820	0.0625	0.2039	0.0624	0.0458	0.1937	0.1691
Std. Deviation	0.2744	0.2421	0.4029	0.2420	0.2090	0.3952	0.3748
Observations	135,589	107,607	879,955	163,630	126,651	1,044,621	2,458,053
School Level							
Mean	2.787	2.836	2.659	2.828	2.859	2.687	2.707
Std. Deviation	1.027	0.987	1.058	1.015	0.979	1.053	1.046
Observations	135,500	107,554	879,523	163,536	126,578	1,044,070	2,456,761

Note: sample restricted to individuals aged between 12 and 97. For wage variables, observations with non-reported values for hourly earnings are excluded. For sociodemographic controls, observations with non-specified responses are also omitted.
Appendix A describes the variables construction procedure.

Following the International Labour Organisation (ILO) parameters, ENOE classifies individuals into active and inactive labour market population. Active population refers to individuals who, during the week of interview were working (employed), or in the previous two weeks developed activities to find a job (unemployed). Inactive individuals are those neither working, nor looking for a job.

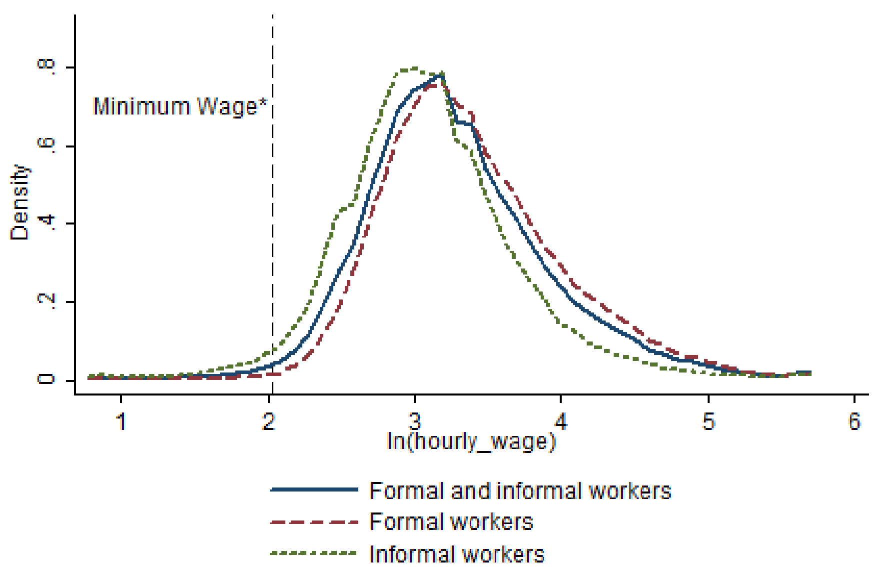
Employed workers are also classified into formal and informal workers. In the same way, following the ILO conceptual framework, the key feature to distinguish between formal and informal workers is the lack of recognition by the employer of the labour relation between them, which generates a vulnerable condition for the worker. ENOE uses the Hussman's matrix to identify and classify informal workers into waged, unwaged, and self-employed. Separate effects for each category are estimated in the next section, as well as the overall effect on informal workers.

Table 1 shows the sample descriptive statistics for the variables included in the econometric models. The sample is disaggregated by wage zones and periods, taking as a reference the policy intervention.

In terms of mean wages — in contrast to Zones A and C — treated Zone B exhibits an increase in 2013 with respect to the pre-treatment period, but for the labour status variables, the pattern is not as clear as in wages. In this case, Zone B had on average a reduction in the active labour market population and a small increase in the total employed population. For the informal sector, the number of informal workers decreased for all the wage zones.

Figure 1

Wage distribution density estimates by formality condition, Zone B
(pooled sample, 2012-2013)



* Average minimum wage weighted by active labour market population by zones (considering an eight hours working day).

With respect to the earnings distribution, Figure 1 plots the kernel density estimates of the real log-wages by formality condition for treated Zone B.[11] As expected, informal workers exhibit lower wages among the distribution. The dotted vertical line expresses the minimum wage considering a working day of 8 hours, showing that for both formal and informal sectors, there are workers with earnings below the minimum wage value.

3.2 The intervention

On November 26, 2012 the Council of Representatives of CONASAMI agreed to change the configuration of the minimum wage zones, incorporating Zone B into Zone A, while the minimum wage Zone C remained the same. The classification of the minimum wage zones depends mainly on the level of economic development of each municipality. Thus, the main argument for this change in the legislation was that since 1988 the municipalities that belonged to Zone B had experienced a development process that had led them to possess similar economic conditions to those observed in Zone A.[12] In addition, this resolution represented an important step towards reaching a unique minimum wage level in Mexico.[13]

The minimum wage increase observed in Zone B is used as the source of identification to measure the impact on wages, employment and informality. Using the terminology by Angrist and Pischke (2008), these observational data (not generated by a randomized trial) are going to be utilized to 'approximate a real experiment'. A fundamental characteristic of the intervention is that it was implemented with no anticipation from the labour market. One day after the announcement, on 27 November 2012, the legislation came into force. Therefore, it is not necessary to model any anticipated responses.

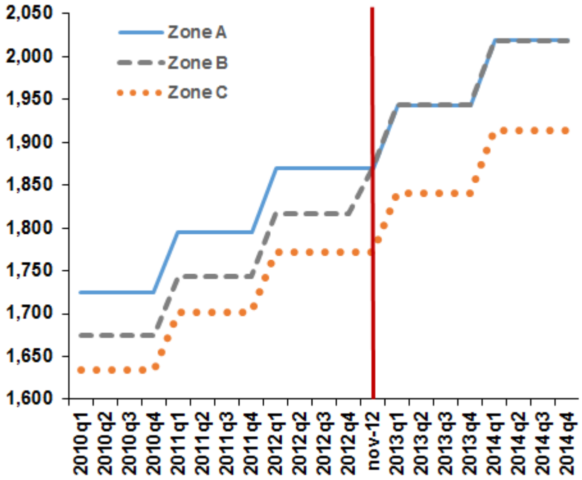
To illustrate the minimum wage change generated by the policy intervention, Figure 2 describes the monthly minimum wages for the three different zones for the period 2001-2014. In monetary terms, the monthly increase in the former Zone B was \$53.00 MXN, which represented a nominal increase of 2.9%.

[11] In line with the sample used in Section 5.2, and with previous literature on the distributional effects of minimum wage (Lee, 1999; Autor et al., 2016), self-employed workers are not included in the analysis and the data *winsorizes* the extreme 0.2 percentiles of the wage distribution by assigning the 0.02 and 99.8 percentile value to the respective quantiles. Epanechnikov kernel using the optimal cross validation bandwith: $h = 0.9 \min(\sigma, IQR/1.349)n^{-1/5}$, where σ is the standard deviation of the log hourly real wages and IQR denotes the interquartile range.

[12] Source: Official Journal of the Federation. November 26 2012.
`http://dof.gob.mx/nota_detalle.php?codigo=5279120&fecha=26/11/2012`

[13] On 20 March 2015 the Ministry of Labour announced a new legislation to have a unique national minimum wage. The increase for Zone B was divided into two parts; a half of the increment entered into force in April and the rest in October 2015. Source: Ministry of Labour and Social Welfare.

Figure 2
Nominal monthly minimum wage by zones
(current Mexican pesos, 2010-2014)



Source: INEGI, and own calculations with data from ENOE.

With respect to the demographic characteristics of the wage zones, according to the ENOE and with data for 2012Q3, 55 municipalities with approximately 11.54 million inhabitants lived in Zone B (9.7% of the total population in Mexico). In particular for the active labour market population, Panel (a) of Figure 3 shows that 5.3 million workers (10.2%) were potentially affected by this policy.

In addition, a considerable proportion of the population of wage zone B (71.1%) was concentrated in the second and third biggest cities in Mexico. Approximately 4.4 million of inhabitants lived in the Metropolitan Zone of Guadalajara and 3.5 million of inhabitants lived in the Metropolitan Zone of Monterrey. These two metropolitan areas accounts for 13 of the 55 municipalities in Zone B.[14] The rest of Zone B includes municipalities in the states of Sonora, Jalisco and Veracruz (see Panel (b) of Figure 3 for the geographical distribution).[15]

Finally, regarding the workers on or below the minimum wage, for 2013Q3 there were 6.79 million of workers in Mexico with this level of remuneration, which represented 15.3% of the employed population. Regarding the distribution across wage zones, 6.1% of this workers were employed in Zone B, while, 13.9% and 80.0% worked in Zone A and Zone C, respectively. The fact that Zone C had a bigger proportion of the minimum wage workers is explained by the own design of the wage zones; this zone covers those municipalities with lower levels of economic development, including all the rural municipalities.

[14]Data for the population in each metropolitan zone were obtained from the Mexican National Council on Population. `http://www.conapo.gob.mx/es/CONAPO/Zonas_metropolitanas_2010`

[15]The details for the identification of the wage zones across the municipalities in ENOE are described in Appendix A.

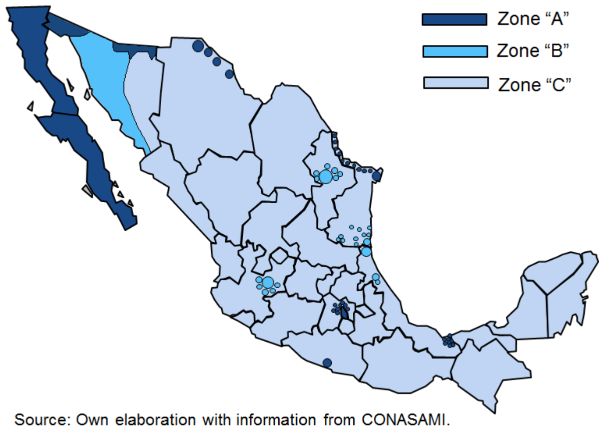
Figure 3
Minimum wage zones valid until 2012

Zone	Population	%
Zone A	11,811,875	22.75
Zone B	5,300,422	10.21
Zone C	34,814,753	67.05
TOTAL	51,927,050	

Source: Own calculations with data from ENOE.

Population values by the use of expansion factors.

(a) Active Labour Market Population by zones (3Q 2012)



(b) Geographical Distribution

Pre-treatment trends: graphical inspection

The legislation described above left Zones A and C unaffected, so they constitute the natural control group. But, given that the intervention was not performed in a randomized way, a common concern in this kind of observational studies is the validity of the control group.

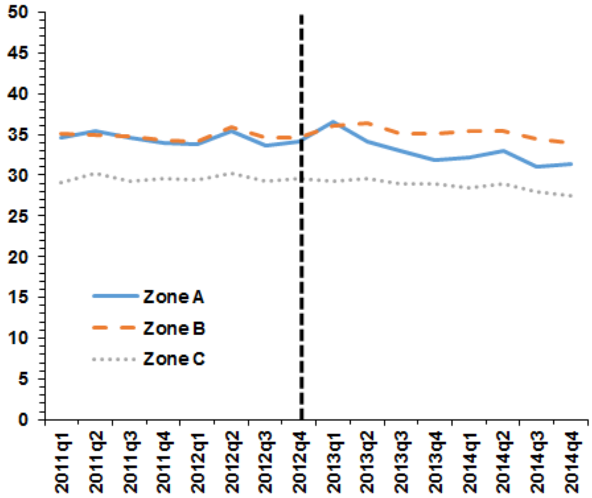
By construction minimum wages zones are different. So, under these circumstances in which control and treatment were not randomized and they are not identical in their observable characteristics; the key identifying assumption is that, in absence of the intervention, control and treatment groups follow the same trend in the outcome variables.

In order to verify the validity of treatment and control groups, it is implemented a graphical inspection of the trend by the outcome variables before and after the policy intervention.

Figure 4 shows the hourly real mean wage by zones. Before the intervention (dotted vertical line denotes the date of the minimum wage legislation) the trends in all three zones are practically equal and parallel, that is, the difference between the mean wage in Zone B with respect to zones A and C remains basically constant. This suggests that the treatment and control groups are valid. In addition, during the post-treatment period the gap between Zone B and the other two zones increased, which suggests that the intervention could have improved the relative mean wage for Zone B.

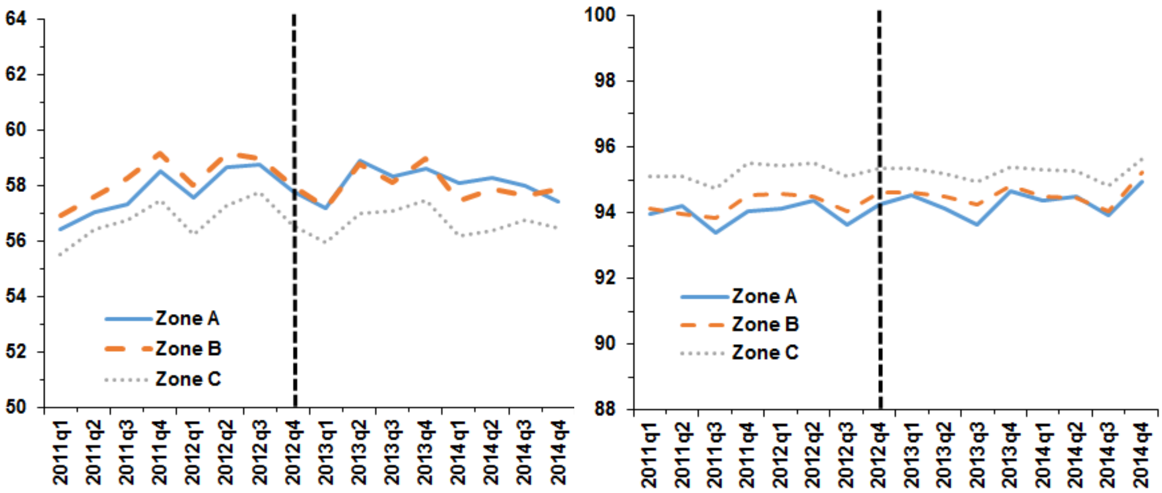
For the analyses of labour status, Panel (a) of Figure 5 shows the active labour market rate by wage zones, that is, the proportion of individuals with labour activities or looking for a job with respect to the total population. Panel (b) describes the employment rate, which corresponds to the percentage of working population relative to active labour market population. In both cases the pre-treatment trend is similar among zones, but there is not a clear pattern after the intervention, so the treatment effect is not distinguishable.

Figure 4
Mean real hourly wage by zones
(Mexican pesos of 1F December 2010)



Source: Own calculations with data from ENOE, INEGI.

Figure 5
Participation rates by wage zone
(%)



Source: Own calculations with data from ENOE, INEGI.

(a) Active labour market rate

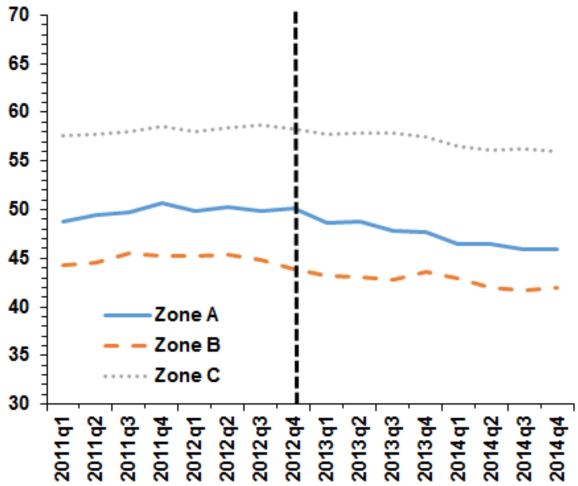
Source: Own calculations with data from ENOE, INEGI.

(b) Employment rate

Finally, Figure 6 shows the trends on the informal labour market participation, constructed as the percentage of individuals working under informal conditions with respect to the total working population. The difference observed in the informality rates across the zones is practically constant before the intervention. In contrast, after the treatment Zones B and C experimented a higher decrease in comparison to Zone A, but it is not possible determine the magnitude, nor the sign of the impact.

Hence, treatment (Zone B) and control groups (Zones A and C) show similar trends in the outcome variables before the intervention, which means that it is possible to implement difference in differences procedures to estimate the treatment effect.[16] So far it is not possible to determine if the effects — if there are any — are positive or negative.

Figure 6
Proportion of informal workers by minimum wage zone
(%)



Source: Own calculations with data from ENOE, INEGI.

4 Identification strategy

Given that the survey allows us to observe the labour market before and after the policy intervention, and based on the assumption that all three zones exhibit the same trends in the pre-treatment period, it is possible to implement difference in differences estimates. This section describes the difference in differences econometric specifications, as well as the main empirical techniques used in the paper: sample selection bias correction and unconditional quantile regressions.

[16]Section 5.5 presents some falsification tests to validate in a more formal way control and treatment groups.

Regarding the econometric specifications, the control group represents not only the counterfactual of the treated observations, strictly speaking it is the reference group to measure the relative difference generated by the intervention. In other words, for invalid control groups, the choice of the reference group can alter significantly the magnitude of the estimated impact. Then, using the fact that there are two untreated zones, and with the aim at verifying the robustness of the model, for all the subsequent estimations two basic specifications are used. In the first of these specifications only individuals in Zone C are included in the control group (observations from Zone A are not dropped from the analyses, instead of that, a dummy variable for this zone is included as a regressor). For the second specification, the control group is integrated by individuals from zones A and C.

Although zones A and B can have more similarities in terms of economic development, Zone C is the most preferred control group by its size. In the sample, 80% of the observations correspond to this zone.

Thus, in the first specification, (1a), we have the traditional difference in differences equation including a shift effect ($Period2$) and an identifier for the treated zone ($ZoneB$). For the second specification, (1b), is included the dummy variable $ZoneA$ and also — for completeness — its interaction with the shift effect, $Period2$. Then, the pooled OLS specifications are the following:

$$\begin{aligned} Y_i = {} & \beta_0 + \delta_1 ZoneB_i * Period2_i + \delta_2 Period2_i + \delta_3 TrendB \\ & + \delta_4 TrendA\&C + \beta_1 ZoneB_i + \sum_{k=2}^{k} \beta_k X_{ki} + e_i \end{aligned} \tag{1a}$$

$$\begin{aligned} Y_i = {} & \beta_0 + \delta_1 ZoneB_i * Period2_i + \delta_2 ZoneA_i * Period2_i + \delta_3 Period2_i + \delta_4 TrendB \\ & + \delta_5 TrendA + \delta_6 TrendC + \beta_1 ZoneB_i + \beta_2 ZoneA_i + \sum_{k=3}^{k} \beta_k X_{ki} + e_i \end{aligned} \tag{1b}$$

δ_1 is the parameter of interest. In equation (1a), it corresponds to the estimated impact of the intervention on treated individuals with respect to individuals in zones A and C, while in specification (1b), δ_1 is the estimated effect with respect to individuals Zone C. It is important to emphasize that δ_2 in equation (1b) is not a treatment effect; the purpose is not trying to measure the impact on the untreated Zone A. It is included to test the robustness of our parameter of interest, and for completeness to avoid losing all the observations from Zone A when control group is only Zone C.

Regarding the dependent variable, Y_i, it changes across the models depending on the impact under evaluation. Thus, in subsections 5.1 and 5.2 the dependent variable is the logarithm of real *hourly wage*. In Subsection 5.3 the response variable is the employment status, and two different binary variables are used: *active labour market* and *employed*. Finally in Subsection 5.4, the impact on different categories of informal workers is estimated: *informal*, *informal waged*, *informal self-employed*, and *informal non-waged*.

For the set of independent variables, $Period2_i$ is an indicator variable that identifies the post-treatment period. $ZoneB_i$ and and $ZoneA_i$ are dummy variables that recognize the observations that belong to Zone B or A, respectively. In addition, the inclusion of time effects (separate linear trends by zones) allows that the identifying variation in the minimum wage comes from the intervention on Zone B in November 2012.[17]

X_{ki} corresponds to the set of sociodemographic variables at the individual level including gender, age, squared age, schooling level, and an indicator of rural residence. Interactions of schooling level with rural residence and gender are also included.[18] e_i is the error term.

In addition to the full sample specification, some models are also estimated restricting the sample to the following age thresholds: individuals aged between 12 and 29, between 30 and 49 , and finally, individuals aged equal or older than 50.

Given that the models on wage rates, employment and informality systematically exclude inactive labour market population, all the models are estimated in two versions: by pooled OLS (not reported), and correcting for sample selection bias. A general description of the implementation of the latter procedure is presented in the following subsection, while the unconditional quantile regression for the model of the effects on the earnings distribution is described in Subsection 4.2.

4.1 Sample selection bias correction

To test and correct for selection bias in the models for the mean effect on wages, employment and informality, the two stage procedure developed by Heckman (1979) is used. The first step estimates the selection equation by a probit model (in our case it corresponds to the probability of participating actively in the labour market), and the fitted values for the inverse Mills ratio need to be calculated ($\hat{\lambda}$). Then, to obtain unbiased estimators in the second stage, $\hat{\lambda}$ is included as a regressor in the selected sample. It is enough to implement a standard t test on the $\hat{\lambda}$ coefficient to test the null hypotheses of no selection bias.

[17]Similarly, Autor et al. (2016) include time effects in the specifications to evaluate the minimum wage impact on inequality in the United States.

[18]The models of the effect on wages and its distribution include also state employment rate (defined as the percentage of employed workers over the active population) as a regressor to control for the macroeconomic trend at the state level.

For the binary dependant variable models on employment and informality, the Heckman two-stage procedure is not feasible, but Van de Ven and Van Praag (1981) developed a maximum likelihood probit estimator to correct sample selection models. The procedure is similar, but the first stage is estimated by fitted maximum likelihood estimators, and to test the selection bias it is necessary to estimate the correlation between the fitted errors of the selection equation and the equation of interest (ρ)[19].

Wooldridge (2010) states that technically, the set of regressors for the two stages do not have to be different. However, collinearity between $\hat{\lambda}$ and the set of explanatory variables can lead to large standard errors of the parameters. To avoid this, the binary variable $Head$ (equal to one if the individual is the head of the household, and zero otherwise) is used as an exclusion restriction in the selection equation.

As any instrument, it needs to satisfy two conditions. *Exogeneity* in the equation of interest — specifications (1a) and (1b) — that is, the covariance between $Head_i$ and the error term e_i has to be zero. And *relevance* in the reduced form, which in our case means that the parameter on $Head_i$ in the selection equation must be statistically different from zero.

The fact that an agent is the head of a household affects directly the decision to participate in the labour market because she is the main responsible to satisfy needs of the household (*relevance* condition). But, it does not affect any of the dependent variables: it does not determine the remuneration obtained (wage level), nor the labour status once they are active in the labour market (employed/unemployed), nor the formality conditions under the worker develops their labour activities.

Nevertheless, special attention has to be devoted to the estimates on hourly wage. The wage rate should not be affected by the variable $Head$, but it can have an influence on the decision on the number of hours worked. In that case, Heckman correction procedure would not be invalid, but the standard errors could be overestimated.

Independently of the control group used, the reduced form for the selection equation on labour active market are expressed by the following specification:

$$Y_i = \theta_0 + \theta_1 Head_i + \theta_2 Period2_i + \theta_3 Trend + \sum_{k=4}^{k} \theta_k X_{ki} + r_i \tag{2}$$

Notice that unlike equation (1), this specification does not include the variables $ZoneB$, $ZoneA$, $ZoneB * Period2$ and $ZoneA * Period2$. The purpose is not to estimate a treatment effect. In addition, they are not relevant in the model. Testing the individual and jointly significance of these variables all of them resulted not statistically different from zero (see Table 5 for the parameters on $ZoneB{*}Period2$).[20]

[19] It is estimated the transformation athan(ρ): athan(ρ) $= 1/2 \cdot log[(1+\rho)/(1-\rho)]$.

[20] Testing the joint significance of parameters of $ZoneB * Period2$, and $ZoneB$ on dependent

4.2 Unconditional quantile regression

The main motivation for implementing quantile regressions is to estimate the effect of the intervention on the distribution of earnings. That is, beyond answering what is the effect on the conditional mean, it allows to obtain differential impacts along the distribution. Given the nature of the minimum wage policy purposes — related to ensure a minimum level of earnings for the segment of the labour force with the lowest level of income — and the potential spillover effects, it is fundamental to evaluate the impact of the minimum wage increase at different points of the wage distribution.

Conditional quantile regression (CQR), developed by Koenker and Bassett (1978), became a useful empirical tool to characterize the full distribution of a certain outcome conditioned on a set of covariates. Nevertheless, the interpretation of the estimated parameters in a evaluation program setting is complicated; the coefficients do not translate to the relevant policy questions that are linked to the covariates, they do not summarize the causal effect of a treatment (Borah and Basu, 2013; Frolich et al., 2010). This subsection describes the method recently proposed by Firpo et al. (2009), which allows to evaluate how a marginal change in one variable — in this case, the minimum wage variation in Zone B — affects the entire wage distribution, keeping the distribution of the rest of covariates constant.

Formally, the aim is to estimate the effect of the minimum wage rise, denoted as mw, on the τth quantile of the marginal earnings unconditional distribution, $F_Y(y)$. In the OLS regression framework, the coefficient δ is interpreted as the impact on the conditional mean: $\delta = d\mu(mw)/dmw = E(Y|mw = 1) - E(Y|mw = 0)$.[21] In contrast, the coefficient δ_τ^{CQR} from conditional quantile regression is generally different from the partial effect: $\delta_\tau^{CQR} = F_Y^{-1}(\tau|D = 1) - F_Y^{-1}(\tau|D = 0) \neq dq_\tau(mw)/dmw$.[22]

The reason of the inequality is simple, conditional and unconditional distributions are not necessarily the same. For example, the set of workers at the 5th percentile of the unconditional earnings distribution of Y may not be the same as the workers at the 5th percentile of the conditional distribution of $Y|mw$ (Borah and Basu, 2013).

Firpo et al. (2009) showed that the *unconditional quantile partial effect* can be obtained by running an OLS regression of the recentered influence function (RIF) of the unconditional quantile on the explanatory variables.

variable *Active Labour Market*, for equation (1a) we cannot reject the null hypotheses that these variables are jointly insignificant (with a p-value of 0.91). For equation (1b), the null hypotheses that the parameters of $ZoneB * Period2$, $ZoneB$, $ZoneA * Period2$, and $ZoneB$ are jointly equal to zero cannot be rejected (p-value of 0.91).

[21] $mw = 1$ denotes treatment, that is, it identifies those workers performing labour activities in some municipality of Zone B after 26 November 2012; $mw = 0$ indicates absence of treatment.

[22] The coefficients of a covariate on a specific quantile outcome are the same by conditional or unconditional quantile regressions if there are no additional covariates influencing the data generating process, or if the effect is constant across levels of other covariates (Borah and Basu, 2013).

The influence function, $\text{IF}(Y; v; F_Y)$, of a distributional statistic $v(F_Y)$ represents the influence of a single observation on that distributional statistic, for instance variance, quantiles, or Gini coefficient.For the τth quantile, the influence function corresponds to: $\text{IF}(Y; q_\tau; F_Y) = \tau - \mathbb{1}\{Y \leq q_\tau\})/f_Y(q_\tau)$. While the recentered influence function is obtained just adding back the statistic $v(F_Y)$ to the influence function, in this case the τth quantile: $\text{RIF}(Y; q_\tau; F_Y) = q_\tau + IF(Y; q_\tau; F_Y)$.

Firpo, Fortin and Lemieux demonstrated that the average derivative of the conditional expectation of the RIF, $E[\text{RIF}(Y; q_\tau; F_Y)|X] = m_v(X)$, corresponds to the marginal effect on the unconditional quantile of a small location shift in the distribution of covariates, holding everything else constant. Therefore, the RIF regression model can be viewed as an unconditional quantile regression.

Hence, to implement the unconditional quantile regression the first step is to estimate the dependent variable $\text{RIF}(Y; q_\tau; F_Y) = q_\tau + \tau - \mathbb{1}\{Y \leq q_\tau\})/f_Y(q_\tau)$. So, it is necessary to compute each of its components: the sample quantile q_τ, a dummy variable $\mathbb{1}\{Y \leq q_\tau\}$ indicating wheter the outcome variable is below q_τ, and the density $f_Y(q_\tau)$ at the point q_τ by Kernel procedures, or other non-parametric methods. Finally, this new dependent variable is regressed on the set of covariates.

Subsection 5.2 presents the estimates of the impact of the minimum wage increase on the unconditional earnings distribution.

5 Results

5.1 The effect on real wages

Although the central research question in the minimum wage literature has been if minimum wage changes have an impact on employment, the first step to evaluate the minimum wage intervention of 2012 is to analyse if the policy change is truly affecting real earnings. In addition, the model also explores the effects on uncovered sectors; this paper uses the official data from ENOE to analyse the impact on the informal labour market, in which by definition, there is no full enforcement of the legal labour market framework.

The econometric specifications correspond to equation (1),[23] but the variable state

[23]It is important to remark that for these models, all those individuals with non-valid response for wage are omitted from the analyses. Around one fourth of the employed individuals in the sample (343,889 observations) did not specify or did not answer correctly their wages. Although these observations could have been included in the first stage of the Heckman sample correction procedure (estimating the probability of being active in the labour market, jointly to the probability of reporting valid earnings), this model would have a lack of economic interpretation. In addition it would be necessary to assume homogeneous distribution of wages across the non-respondents.

employment rate is also included in the set of covariates to control for the economic trend at the macro level. With respect to potential endogeneity problems by the inclusion of this variable, Appendix B presents some analytical arguments, as well as the Durbin-Hu-Hausman test to demonstrate the exogeneity of this covariate in the wage equation.

To obtain unbiased estimators, sample selection techniques need to be implemented. In all the wage regressions presented, as tables 2 and 3 show, the Inverse Mills ratio ($\hat{\lambda}$) results statistically significant. So, the null hypotheses of no selection bias is rejected, which implies that pooled OLS models are biased. With regard to the first stage, the parameter of exclusion restriction, $Head$, in the selection equation for the wage specification is positive and highly significant in all the cases (not reported).

Given that the sample selection procedure is explicitly estimated in two stages, robust or clustered standard errors are not feasible. So, standard errors are obtained by the two-step variance estimator derived by Heckman (1979), who indeed states that in presence of sample selection bias, the usual procedure to estimate standard errors can lead to an overestimation of significance levels.

Thus, correcting for sample selection bias, Table 2 shows that independently of the specification used, the impact for the full age threshold is positive and statistically significant. The minimum wage rise in Zone B increased real hourly wages, on average, by around 3.6%. The effect is economically significant, particularly if one considers the decreasing pattern in the purchasing power of wages (Figure 4). The legislation, with a nominal increase of only 2.9%, is breaking the negative trend in real earnings, which is actually observed in the other two zones.

Regarding previous findings in the literature for the minimum wage harmonization of 2012, Campos et al. (2015) is the only reference. For similar pooled OLS specifications, but without correcting for sample selection bias, they estimated an impact in the range of 1.67% and 1.98%. This confirms that the lack of sample selection bias procedures underestimates the evaluation of the minimum wage raise.[24]

Table 2 also presents the analyses of the subsamples by age thresholds. The effect is present only for subsamples of individuals younger than 50 years old; for the youngest sector of the workforce the impact is a little bit lower than 4%, while for workers aged between 30 and 49 the effect is estimated between 4.7% and 4.9%. In contrast to the traditional minimum wage literature, our analysis suggests that the wage effect is not present only on young workers. This is explained by the age distribution of the minimum wage workforce: 69% of workers on or below the minimum wage are aged older than 29.

[24] They also implemented panel data models estimating stronger effects (although not robust to different time-spans): up to 3.5% for hourly wages.

Table 2

The impact on hourly wages.
Heckman second stage for sample selection bias.

Dependent variable:	ln(*hourly_wage*)			
Equation:	(1a)		(1b)	
Full age threshold: 12 $\leq$ Age $\leq$ 97				
ZoneB*Period2	0.0359***	(0.00942)	0.0363***	(0.00948)
$\hat{\lambda}$ (IMR)	-0.1588***	(0.00541)	-0.1580***	(0.00541)
Total observations	2,112,508		2,112,508	
Uncensored observations	960,550		960,550	
Age threshold: 12 $\leq$ Age $\leq$ 29				
ZoneB*Period2	0.0399***	(0.01418)	0.0384***	(0.01426)
$\hat{\lambda}$ (IMR)	-0.2616***	(0.00961)	-0.2564***	(0.00958)
Observations	886,481		886,481	
Uncensored observations	309,008		309,008	
Age threshold: 30 $\leq$ Age $\leq$ 49				
ZoneB*Period2	0.0473***	(0.01374)	0.0492***	(0.01383)
$\hat{\lambda}$ (IMR)	-0.1553***	(0.00841)	-0.1554***	(0.00840)
Total observations	687,799		687,799	
Uncensored observations	450,695		450,695	
Age threshold: 50 $\leq$ Age $\leq$ 97				
ZoneB*Period2	-0.0075	(0.02411)	-0.0075	(0.02426)
$\hat{\lambda}$ (IMR)	-0.1033***	(0.01455)	-0.1052***	(0.01454)
Total observations	538,228		538,228	
Uncensored observations	337,381		337,381	

Note: the covariates included are state employment rate, gender, age, squared age, rural, schooling level, and interactions of schooling level with rural and gender.
Observations with non-reported wages are excluded from the analysis.
Standard errors in parentheses, by two-step variance estimator Heckman (1979).
Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Our analyses focus the main estimates on hourly wage because it addresses also the impact on employment. That is, it absorbs the minimum wage effect on the number of hours worked. In contrast, regressions on monthly wage only indicate the impact on the earnings that workers are taking home given that they kept their jobs. For completeness, monthly real wages estimates are presented in Appendix C. Table C.1 shows that the effect is estimated statistically significant only for the full sample, between 1.6% and 1.9%. Nevertheless, there is no evidence of negative effects on real monthly wage in any of the models estimated.[25]

Table 3 shows the analysis of the wage effects by formality condition. First, there is a statistically significant impact on wages in the formal sector, which means that the legislation is actually in force; the estimated effect on hourly wages for the full age threshold is 3.2%. Unlike previous estimates on the sample pooling formal and

[25] Campos et al. (2015) found no evidence of significant effects on monthly wages by pooled OLS regressions. By panel date models they estimated an increase on monthly wages by 7.7%, but their results are not robust to different time-spans.

informal workers, the effect is stronger on workers younger than 30 years old (between 4.6% and 4.7%), although the impact is still present on middle age workers (3.5%).

Table 3
The impact on hourly wages by formality condition.
Heckman second stage for sample selection bias.

	Formal workers				Informal workers			
Dependent variable:	ln(*hourly_wage*)				ln(*hourly_wage*)			
Equation:	(1a)		(1b)		(1a)		(1b)	
Full age threshold: 12 ≤ Age ≤ 97								
ZoneB*Period2	0.0319***	(0.01137)	0.0316***	(0.01149)	0.0316**	(0.01458)	0.0309**	(0.01464)
$\hat{\lambda}$ (IMR)	-0.1121***	(0.00560)	-0.1124***	(0.00560)	-0.1482***	(0.00685)	-0.1474***	(0.00684)
Total observations	1,567,202		1,567,202		1,664,940		1,664,940	
Uncensored Observations	428,783		428,783		531,767		531,767	
Age threshold: 12 ≤ Age ≤ 29								
ZoneB*Period2	0.0470***	(0.01673)	0.0456***	(0.01690)	0.0281	(0.02240)	0.0263	(0.02247)
$\hat{\lambda}$ (IMR)	-0.1763***	(0.00912)	-0.1750***	(0.00912)	-0.2056***	(0.01232)	-0.1990***	(0.01226)
Total observations	704,294		704,294		751,993		751,993	
Uncensored observations	130,639		130,639		178,369		178,369	
Age threshold: 30 ≤ Age ≤ 49								
ZoneB*Period2	0.0351**	(0.01620)	0.0347**	(0.01636)	0.0576**	(0.02241)	0.0599***	(0.02250)
$\hat{\lambda}$ (IMR)	-0.1148***	(0.00867)	-0.1153***	(0.00867)	-0.1182***	(0.01003)	-0.1187***	(0.01001)
Total observations	452,579		452,579		455,842		455,842	
Uncensored observations	221,419		221,419		229,276		229,276	
Age threshold: 50 ≤ Age ≤ 97								
ZoneB*Period2	-0.0033	(0.03185)	-0.0011	(0.03213)	-0.0167	(0.03358)	-0.0203	(0.03374)
$\hat{\lambda}$ (IMR)	-0.0805***	(0.01930)	-0.0826***	(0.01929)	-0.0841***	(0.01828)	-0.0866***	(0.01826)
Total observations	410,329		410,329		457,105		457,105	
Uncensored observations	76,725		76,725		124,122		124,122	

Note: the covariates included are state employment rate, gender, age, squared age, schooling level, rural, and interactions of schooling level with rural and gender. Observations with non-reported wages are excluded from the analysis.
Standard errors in parentheses, by two-step variance estimator Heckman (1979). Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Second, for the informal sector, there is also evidence of a positive impact on real hourly wages by around 3.1%. This confirms that the intervention is having spillover effects to this specific 'uncoverded' sector. It is relevant to point out that the effect comes from the impact on workers aged between 30 and 49, which is estimated between 5.8% and 6.0%. Actually, this is the highest estimated effect on earnings regressions, and the fact that it is taking place in the informal labour market implies that the minimum wage legislation is affecting substantially the wage incentives in this sector. A possible explanation is the importance of relative remunerations[26] on wage setting: given that the minimum wage rise is making more attractive the formal sector, in response, informal employers also increase the wage rates offered in order to retain their workers — even in higher magnitudes.

[26]See Grossman (1983) for a detailed analysis on the relevance of relative wages on wage setting.

5.2 The impact on earnings distribution

The previous subsection showed that the rise of the minimum wages in Zone B truly had an impact on real wages, but the analysis does not provide information on how the intervention impacted earnings across different points of its distribution, especially on the targeted poorest segment of the labour force.

Quantile regression takes special relevance if we consider the context of the Mexican labour market. First, as described in the introductory section, the loss of purchasing power of the minimum wage has generated the belief that there are no workers earning the minimum wage (Calva and Picard, 2007). Then, if this is the case, the significant mean conditional effects estimated by Heckman pooled OLS regressions would be explained as a consequence of the 'numeraire' or 'lighthouse' effect of the minimum wage, in which minimum wage is taken as a reference rate for wage setting (Bosch and Manacorda, 2010; Castellanos et al., 2004; Kaplan and Pérez-Arce, 2006). Thus, changes to the minimum wage would have an impact on the conditional mean of real wages, but not necessarily on the targeted population: lowest income workers.

In this regard, the role of reference price of the minimum wage on prices has been widely discussed in Mexico not only by academic researchers, but also by policy designers.[27] Recognizing the likely price consequences of the lighthouse effect of the minimum wage, in January 2016 Mexican Congress passed a reform called 'de-indexation of the minimum wage', in which minimum wage is not tied any longer to any price, including labour benefits, tax rates, social security fees, among others.

Second, the size of the informal labour market in Mexico makes the practical implementation of minimum wage policies in the labour market difficult. Although minimum wage regulation is not accomplished in the informal sector, our estimates also suggest that there is an important effect on the informal workers earnings, nonetheless the impact is only present in workers aged between 30 and 49. But once again, the conditional mean regressions do not reveal what is the effect on the unconditional earnings distribution of informal workers.

Hence, this subsection implements the innovative and advantageous technique developed by Firpo et al. (2009) that allows to estimate directly the marginal treatment effects along the the wage distribution.[28] The objective is twofold: on the one hand to evaluate if there is actually an impact on the lowest segment of the earnings distribution, and on the other, to verify the existence of spillover or lighthouse effects.

[27]See, for example a proposal for minimum wage reform in 2014 by the Government of Mexico City (Gob.Distrito-Federal, 2014)

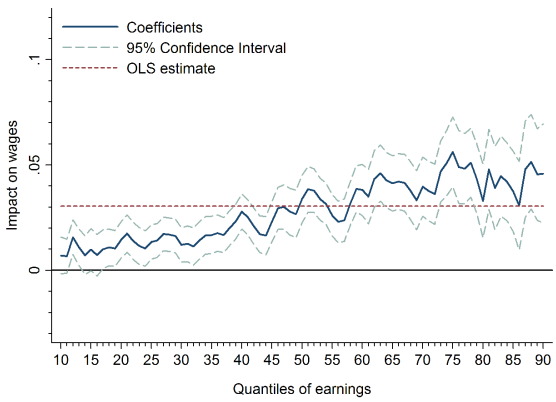
[28]For the unconditional quantile regression analysis sample section bias correction is not implemented. The purpose itself of this procedure is to estimate the effect on the earnings distribution, so there is no reason to to consider into the analysis inactive labour market population, which by definition does not perform labour activities and in consequence, does not perceive earnings.

In terms of the sample used, there are two important differences with respect to the sample used in the rest of the regressions. Following previous research on wage distribution effects by minimum wage changes, self-employed workers independently of their formality condition are excluded (Lee, 1999; Autor et al., 2016). In addition, in line with the procedure by Autor et al. (2016), to reduce the influence of the outliers the data *winsorizes* the extreme 0.2 percentiles of the wage distribution by assigning the 0.02 and 99.8 percentile value to the respective quantiles. Nevertheless, these modifications do not alter significantly the estimates obtained, nor the conclusions of the analysis.

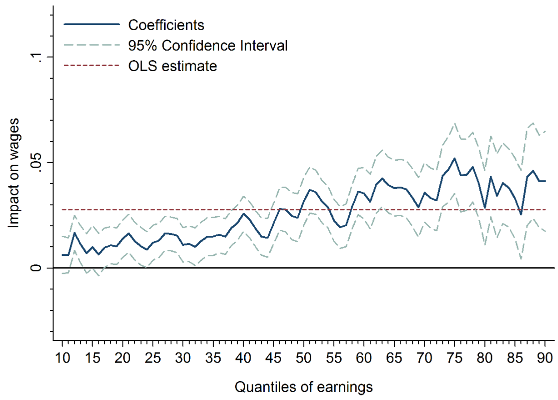
Figure 7 shows the RIF regression coefficients for the pooled sample including formal and informal waged workers, as well as its 95% confidence interval. The first aspect to highlight is the fact that the intervention has a positive, although weak effect on the lowest deciles of the distribution. For percentiles 10th, 11th, 14th, 15th, and 16th the effect is not statistically significant, at least at the 5% level. For the rest of the percentiles below the 20th percentile, the effect on earnings is significant with a magnitude of around 1%. Even though the impact is small, there exists evidence of significant and positive effects at the bottom of the distribution.

Figure 7

Unconditional quantile regressions on earnings distribution.
Pooled sample, formal and informal workers



(a) Equation (1a)



(b) Equation (1b)

A common argument against minimum wage increases in Mexico, is that the policy target group performs the labour activities out of the formal labour market, and minimum wages are not enforced in the informal sector, so that minimum wage policies cannot affect the lowest income workers. Our results demonstrate that there is a positive, although small, impact on the lowest quantiles of the earnings distribution.

The size of the coefficients becomes higher as we move to the right of the distribution; the strongest effect is observed at the 75th percentile, with an estimated impact of 5.6%, while for the 90th percentile the effect is 4.6% (see Table 4 for the estimates on the key percentiles of the distribution). This implies that the wage dispersion in the Mexican labour market is increasing by the minimum wage intervention.

The magnitude of the estimated impacts reveals the relevance of the institutional wage setting. The fact that the intervention took place before the 'de-indexation reform' of 2016 can explain the size of the effect beyond the 20th percentile, in which minimum wages were an important determinant of wage setting in the whole labour market, not only for the lowest earnings sector.

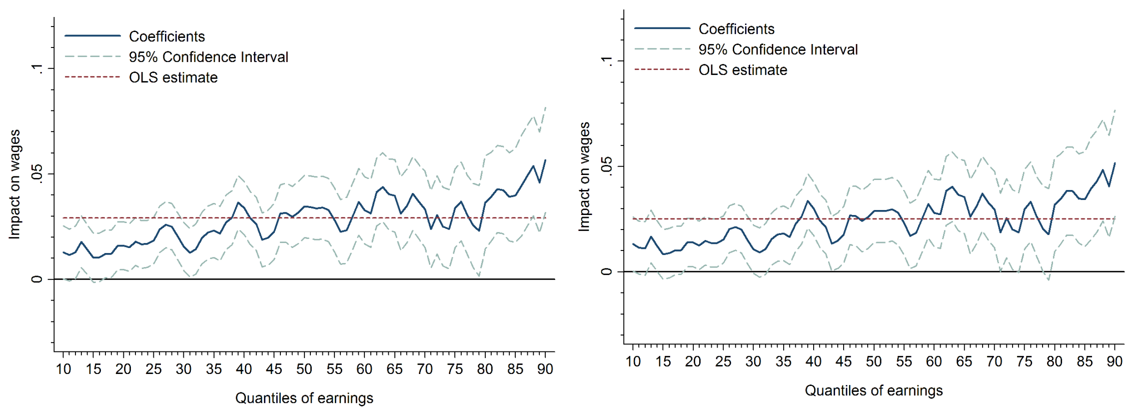
Unfortunately, it is not possible to construct a counterfactual to estimate the distributional effects of the minimum wage reform in presence of the 'de-indexation reform' to distinguish between the 'pure' minimum wage effect on the lowest deciles, and the spillover effects at the top of the distribution.

Nonetheless, our results corroborate the findings in previous literature in two ways. They confirm that minimum wage affects several occupational wages, not only the lower end of the distribution (Grossman, 1983). And also that minimum wage in Mexico has a role of a reference rate for wage and price setting (Fairris et al., 2008; Bosch and Manacorda, 2010; Castellanos et al., 2004; Kaplan and Pérez-Arce, 2006).

For the formal workers earnings distribution, Figure 8, and Panel (b) of Table 4 shows that the impact along the distribution is basically the same with respect to the pooled sample: minimum wage effects are stronger for the top of the distribution. But, the impact for the lowest percentiles of the distribution are slightly higher with respect to the pooled sample: for the bottom decile the estimated effect is 1.3%, while for the 25th percentile is 1.8%. It is worth highlighting that these estimates confirm that minimum wages are truly in force in the formal labour market.

Figure 8

Unconditional quantile regressions on formal workers earnings distribution



(a) Equation (1a) **(b)** Equation (1b)

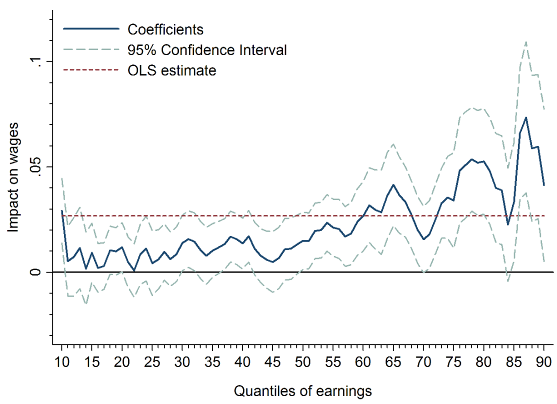
With respect to the earnings distribution of the informal labour market, the pattern of the effect is different. Figure 8 shows isolated significant effects at the 10th, 30th, and 40th percentiles, but the RIF regression shows that the intervention increased real wages on the informal sector for workers with earnings above the median.[29]

There is no a clear explanation for these results, but a likely reason for the positive and strong impact on the first decile of the distribution is that there could be a pressure from informal workers with the lowest earnings to obtain a higher remuneration, as occurred under formality conditions. On this issue, Maloney and Mendez (2004) state that although minimum wage is not enforced by law in this sector, it appears to be a benchmark for 'fair' remuneration. For the effect on the highest quantiles, the incentives are different; the results can confirm the hypotheses that an increase of wages in formal labour market affect the opportunity cost for high waged informal workers of remaining employed under informal conditions. In response, informal employers increase wages for these workers in order to retain them.

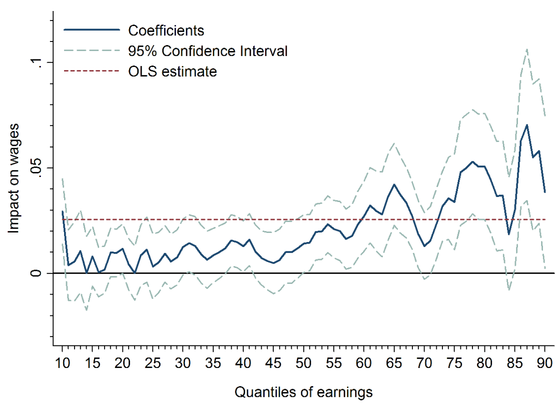
Thus, the unconditional quantile regression analysis confirms that the minimum wage rise actually improved real wages for the targeted workers, even if the increase is only 2.9% in nominal terms. Simultaneously, it corroborates the presence of the lighthouse effect of minimum wages in Mexico at the moment of the intervention, suggesting that the institutional setting of the minimum wage as a reference price can have negative repercussions on the labour market, specifically on wage dispersion.

Figure 9

Unconditional quantile regressions on informal workers earnings distribution



(a) Equation (1a)



(b) Equation (1b)

[29]Table C.3 in Appendix reports the unconditional quantile regression parameters for the effect on the distribution of monthly wages. The conclusions are not different: weak effects at the bottom of the distribution, and evidence of a higher impact on the top deciles of the distribution

Table 4

The impact on the hourly earnings distribution. Full age threshold: $12 \leq \text{Age} \leq 97$

Dependent variable:	ln(*hourly_wage*)											
	Pooled OLS		q10		q25		q50		q75		q90	
(a) Pooled sample, formal and informal workers												
Equation (1a)												
ZoneB*Period2	0.0305***	(0.00439)	0.0070	(0.00443)	0.0135***	(0.00417)	0.0339***	(0.00573)	0.0562***	(0.00844)	0.0459***	(0.01200)
Equation (1b)												
ZoneB*Period2	0.0277***	(0.00443)	0.0062	(0.00451)	0.0121***	(0.00423)	0.0317***	(0.00577)	0.0521***	(0.00849)	0.0412***	(0.01207)
Observations	767,006		767,006		767,006		767,006		767,006		767,006	
(b) Formal workers												
Equation (1a)												
ZoneB*Period2	0.0291***	(0.00543)	0.0128**	(0.00643)	0.0184***	(0.00564)	0.0346***	(0.00758)	0.0338***	(0.00955)	0.0566***	(0.01272)
Equation (1b)												
ZoneB*Period2	0.0251***	(0.00550)	0.0132**	(0.00655)	0.0154***	(0.00571)	0.0289***	(0.00765)	0.0297***	(0.00965)	0.0516***	(0.01285)
Observations	405,217		405,217		405,217		405,217		405,217		405,217	
(c) Informal workers												
Equation (1a)												
ZoneB*Period2	0.0268***	(0.00728)	0.0292***	(0.00781)	0.0043	(0.00781)	0.0150**	(0.00694)	0.0341***	(0.01155)	0.0413**	(0.01842)
Equation (1b)												
ZoneB*Period2	0.0256***	(0.00731)	0.0294***	(0.00789)	0.0032	(0.00788)	0.0141**	(0.00698)	0.0339***	(0.01159)	0.0386**	(0.01846)
Observations	361,789		361,789		361,789		361,789		361,789		361,789	

Note: the covariates included are state employment rate, gender, age, squared age, schooling level, rural, and interactions of schooling level with rural and gender.
Observations with non-reported wages are excluded from the analysis.
Bootstrapped standard errors in parentheses, 100 repetitions. Statistical significance: *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

5.3 The effect on employment status

In a very general framework, a higher wage rate makes more attractive the labour market in terms of remunerations, so it increases the labour supply. But, the effect on labour demand is ambiguous, it is not clear if employers can absorb the increasing labour supply given the higher labour costs generated by the minimum wage legislation; employers can reduce the level of employment (standard competitive labour model), or to hire more workers (monopsonistic labour model). Thus, the purpose of this this subsection is to estimate the impact of the minimum wage raise on the probability of being active on the labour market (with respect to being inactive), and the probability of being employed (with respect to being unemployed).

Second and third columns in Table 5 show that there is no effect of the minimum wage increase in Zone B on the probability of being active in the labour market. All the parameters of interest, independently of the age threshold are not statistically different from zero. This result makes sense: if individuals are initially out of the labour market, an increment of 3% in the nominal minimum wage seems to be an insufficient incentive to change their status. It is important to highlight that the model on labour market active population does not need selection bias correction; all individuals in the sample are considered in this unrestricted regression.

For the case of effect on the employment rate, pooled OLS models restrict the sample to employed and unemployed individuals excluding inactive population. Therefore, as in Subsection 5.1 sample selection procedures are implemented. According to the fourth and fifth columns in Table 5, the impact is positive and statistically significant independently of the zone used as a control group. For specification (1a) the marginal effect at the mean (MEM) is estimated at 0.24%, while for specification (1b) the effect is 0.19%.[30]

This result is highly significant: the minimum wage raise in Zone B did not reduce the employment level. In contrast, the employment rate increased by around 0.2%. Although it is still necessary to check the effects on the informal sector, these estimates suggest that the Mexican labour market exhibits characteristics of monopsonistic competition: labour demand is increasing.

Regarding the subsample groups by age, for the youngest and middle age workers, the estimated effects are positive but insignificant. So, changes in minimum wages does not seem to affect the probability of being employed for workers younger than 50 years old. In contrast, for the oldest segment of the labour force the effect is stronger, between 0.48% and 0.43%, depending on the specification used.

[30]The athan(ρ) parameter to test the null hypotheses of no selection bias, is negative and statistically significant for all the regressions, which implies that the pooled OLS parameters are overestimated (coefficients of the probit regression are not reported).

Table 5
MEM's probit for the impact on labour status.

Dependent variable:	*Labour market active*		*Employed*	
Specification:	(1a)	(1b)	(1a)	(1b)
Full age threshold: 12 ≤ Age ≤ 97				
ZoneB*Period2	-0.0010	-0.0066	0.0019***	0.0024***
	(0.00258)	(0.00271)	(0.00068)	(0.00073)
Total observations	2,455,814	2,455,814	2,455,814	2,455,814
Uncensored observations	—	—	1,405,478	1,405,478
Age threshold: 12 ≤ Age ≤ 29				
ZoneB*Period2	0.0032	0.0031	0.0009	0.0014
	(0.00577)	(0.00566)	(0.00169)	(0.00171)
Total observations	995,920	995,920	995,920	995,920
Uncensored observations	—	—	464,598	464,598
Age threshold: 30 ≤ Age ≤ 49				
ZoneB*Period2	-0.0054	-0.0046	0.0019	0.0025
	(0.00505)	(0.00510)	(0.00153)	(0.00153)
Total observations	833,501	833,501	833,501	833,501
Uncensored observations	—	—	464,598	464,598
Age threshold: 50 ≤ Age ≤ 97				
ZoneB*Period2	-0.0007	-0.0088	0.0043***	0.0048***
	(0.00499)	(0.00517)	(0.00096)	(0.00106)
Total observations	626,393	626,393	626,393	626,393
Uncensored observations	—	—	305,081	305,081

Note: the covariates included are gender, age, squared age, rural, schooling level, and interactions of schooling level with rural and gender. The model on active labour market population, in consistency with equation (2), also includes *Head* as a regressor. Clustered standard errors at the state level in parentheses.
Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Campos et al. (2015) evaluated the impact on three different labour status: 'workers' (which in our case would correspond to the variable *employed*), 'unemployed individuals', and 'workers out of the labour force' (corresponding to 'inactive' individuals). By pooled and panel data models, they do not find statistical evidence of the effect of the minimum wage legislation on any of the dependent variables used.[31]

Hence, our analysis indicates that after implementing the sample selection bias procedure, there is no statistical evidence of adverse effects on employment generated by the minimum wage increment, and, for the oldest sector of the working population, the effect is positive and strongly significant: the employment rate is augmenting by almost half a percentage point.

[31] However, their estimates do not look consistent. The effect on the probability of being working, as well as the probability of being unemployed exhibit negative coefficients (although insignificant), while one should expect symmetric coefficients with opposite signs. By the number of observations included in the models, the problem could be that the dummy variables are not correctly restricted. For example, for the 'unemployed' dummy variable, it seems that it takes the value of zero for both, 'employed' and 'out of the labour force' individuals, which has a lack of interpretation. Appendix A explains in a detailed way how the dependent variables are defined in our sample.

5.4 The effect on informality

The evaluation of the effect of the minimum wage legislation on the levels of employment in the Mexican labour market is incomplete if the analysis does not incorporate the estimates on the informal sector. The previous subsection estimated the impact on the probability of being working by 0.2%, but the models do not differentiate between individuals performing their labour activities under formal or informal conditions. So, it is possible that the increase on the employment rate was generated by a higher level of occupation in the informal sector.

As a remark, sample selection correction for the impact on informality takes the usual concept of active labour market participation, but restricts the sample to the informality category under examination. For example, in the model for waged informal workers, the first stage excludes from the analysis those informal workers classified as self-employed or non-waged.

For the full age threshold analysis, Table 6 shows that the estimated effect on the entire classification of informal workers is negative, but not statistically significant. Yet, for the category of waged informal workers, which is expected to be the most affected segment, the intervention is reducing the probability of participating in the informal sector by around 0.9%.[32]

Comparing these estimates with the findings in Campos et al. (2015), they found no significant effects on informality by pooled probit regressions. By panel data models, they estimated that the probability of remaining as a formal worker increases by 5.3%, while the probability of remain as a waged informal worker diminishes by up to 9.8%. But, the low number of observations included in their sample (only those workers that exhibited transition in the formal/informal status are included) generates skepticism about the magnitude of the parameters. For this reason our analyses implement only pooled analysis, aiming to keep all the observations in the sample.

Regarding the age threshold subsamples, there is no effect on the youngest sector of the employed population, but important effects emerges for older workers. For workers between 30 and 49 years old, there is evidence of significant marginal effects on the whole classification of informal workers, but also on the waged informal and self-employed informal segments.

For the eldest workers of the labour force the estimated effects are stronger: for informal workers in general, the impact is between -1.3% and -1.4%, while for waged informal workers the effect is estimated at the range of -1.4% and -1.5%. For self-

[32]The presence of sample selection bias is confirmed, with negative and statistically significant coefficients on the athan(ρ) regressor (not reported), which suggests overestimation with respect to pooled OLS models.

Table 6
MEM's probit for the impact on informality.
Second stage for sample selection bias correction.

Dependent variable:	*informal*		*waged_informal*		*self_emp_informal*		*non_waged_informal*	
Specification:	(1a)	(1b)	(1a)	(1b)	(1a)	(1b)	(1a)	(1b)
Full age threshold: 12 ≤ Age ≤ 97								
ZoneB*Period2	-0.0075	-0.0067	-0.0094**	-0.0088*	-0.0053	-0.0045	0.0008	0.0009
	(0.00533)	(0.00552)	(0.00445)	(0.00486)	(0.00438)	(0.00440)	(0.00167)	(0.00168)
Total observations	2,455,814	2,455,814	2,455,814	2,455,814	2,455,814	2,455,814	2,455,814	2,455,814
Uncensored observations	1,336,180	1,336,180	997,073	997,073	855,037	855,037	670,300	670,300
Age threshold: 12 ≤ Age ≤ 29								
ZoneB*Period2	0.0065	0.0054	0.0013	-0.0002	0.0158	0.0150	0.0038	0.0039
	(0.00805)	(0.00839)	(0.01022)	(0.01058)	(0.00990)	(0.01010)	(0.00624)	(0.00647)
Total observations	995,920	995,920	995,920	995,920	995,920	995,920	995,920	995,920
Uncensored observations	426,114	426,114	344,527	344,527	204,813	204,813	214,954	214,954
Age threshold: 30 ≤ Age ≤ 49								
ZoneB*Period2	-0.0153***	-0.0132***	-0.0144***	-0.0125***	-0.0097**	-0.0083**	0.0007	0.0010
	(0.00456)	(0.00496)	(0.00341)	(0.00406)	(0.00418)	(0.00413)	(0.00124)	(0.00131)
Total observations	833,501	833,501	833,501	833,501	833,501	833,501	833,501	833,501
Uncensored observations	612,879	612,879	471,838	471,838	429,611	429,611	326,846	326,846
Age threshold: 50 ≤ Age ≤ 97								
ZoneB*Period2	-0.0139**	-0.0127**	-0.0158***	-0.0143***	-0.0123*	-0.0118*	-0.0003	-0.0002
	(0.00554)	(0.00518)	(0.00356)	(0.00322)	(0.00675)	(0.00660)	(0.00156)	(0.00153)
Total observations	626,393	626,393	626,393	626,393	626,393	626,393	626,393	626,393
Uncensored	297,187	612,879	180,708	180,708	220,613	220,613	128,500	128,500

Note: the covariates included are gender, age, squared age, schooling level, rural, and interactions of schooling level with rural and gender. Clustered standard errors at the state level in parentheses. Statistical significance: *** p<0.01, ** p<0.05, * p<0.1

employed workers the impact is also higher, around -1.2% in both specifications.

It is important to emphasize that none of the specifications show evidence of an increase of the informal labour market, and the estimated model demonstrates that the probability of working under formal conditions increases. This result implies that although the minimum wage intervention is not biting the informal sector in terms of a compulsory wage rate, a higher minimum wage in the formal market is changing the incentives to work under informal conditions.

On the other hand, for non-waged informal workers, there is no evidence of some effect generated by the policy intervention. This result is to be expected; this sector consists mainly on workers with labour activities within the household, whose incentives are different from the rest of the labour market.

In terms of the impact of the minimum wage rise as a public policy, the result is highly relevant, specially for the vulnerable group of the eldest workers in the labour force. According to CONEVAL, 45.9% of individuals older than 64 years old live in poverty conditions. In addition, it is very unlikely that these workers could aspire to earn a higher salary, or to transit to formality; there are not many options for an increase in their productivity by the improvement of their skills, experience or qualifications. Our results show that as a consequence of the minimum wage legislation, formal employment is increasing for workers older than 50 years old.

Moreover, given these set of results showing that employment rate is raised, and the participation in the informal sector of the labour market is diminished, it is possible to confirm the existence of monopsonistic labour markets in Mexico. As a result of the minimum wage raise, labour demand in the formal market absorbs more workers in spite of higher wages rates.

5.5 Dynamics of the impact

This subsection aims to evaluate how the estimated effect is changing over time. For this purpose, a second set of specifications are used, in which the post-treatment period variable, $Period2_i$, is decomposed into four quarterly dummy variables: 2013_$Q1$, 2013_$Q2$, 2013_$Q3$ and, 2013_$Q4$. These four time-dummy variables, as well as their respective difference in differences regressors are included in the models to capture the treatment effect in each quarter after the intervention. Following the same structure of previous analyses, specification (3a) uses as a control group Zone C, while equation (3b) includes zones A and C in the reference group.

$$Y_i = \beta_0 + \sum_{j=1}^{4} \delta_j ZoneB_i * 2013\_Q_{ji} + \sum_{j=1}^{4} \delta_{j+4} 2013\_Q_{ji} + \delta_9 TrendB + \delta_{10} TrendA\&C + \beta_1 ZoneB_i + \sum_{k=2}^{k} \beta_k X_{ki} + e_i \tag{3a}$$

$$Y_i = \beta_0 + \sum_{j=1}^{4} \delta_j ZoneB_i * 2013\_Q_{ji} + \sum_{j=1}^{4} \delta_{j+4} ZoneA_i * 2013\_Q_{ji} + \sum_{j=1}^{4} \delta_{j+8} 2013\_Qj_i + \delta_{13} TrendB + \delta_{14} TrendA + \delta_{15} TrendC + \beta_1 ZoneB_i + \beta_2 ZoneA_i + \sum_{k=3}^{k} \beta_k X_{ki} + e_i \tag{3b}$$

For the examination on the dynamics of the effect on real wages, Table 7 shows that the impact on earnings is present three quarters after the intervention. The strongest effect is observed in the second quarter of 2013, with estimated impacts in the range of 2.2% and 2.6%. Subsequently, the effect becomes lower and statistically insignificant. In addition, sample selection bias in pooled OLS models is corroborated.[33]

[33] The estimates on the dynamics of monthly earnings are presented in Table C.4. The impact persists up to the second quarter after the intervention with coefficients around 2%.

Table 7
The dynamics for the impact on hourly wages.
Second stage for Heckman correction for sample selection bias.

Dependent variable:	ln(*hourly_wage*)			
Equation:	(3a)		(3b)	
ZoneB*D2013_Q1	0.022***	(0.0078)	0.023***	(0.0077)
ZoneB*D2013_Q2	0.022***	(0.0078)	0.026***	(0.0077)
ZoneB*D2013_Q3	0.019**	(0.0079)	0.021***	(0.0079)
ZoneB*D2013_Q4	0.006	(0.0077)	0.011	(0.0077)
$\hat{\lambda}$ (IMR)	-0.161***	(0.0054)	-0.162***	(0.0054)
Total observations	2,112,508		2,112,508	
Uncensored observations	960,550		960,550	

* Observations with non-reported wages are excluded from the analysis.
Note: the covariates included are state employment rate, gender, age, squared age, rural, schooling level, and interactions of schooling level with rural and gender.
Standard errors in parentheses, by two-step variance estimator Heckman (1979).
Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Campos et al. (2015) estimated a similar regression by pooled OLS models on the dynamics of the impact on wages. According to their results, the effect on hourly wages is also present up to the second quarter, but they found no effects on monthly wages in any quarter after the minimum wage legislation.

With respect to the evaluation of dynamics of the effect on employment status, the pooled OLS regression confirms the absence of an impact on the probability of being active in the labour market; the coefficients are statistically insignificant for the four quarters after the minimum wage increase. But for the effect on employment, the impact is present in the second quarter of 2013, although the magnitude of the marginal effects are similar to those observed in Table 8, between 0.17% and 0.20% depending on the composition of the control group. Even though the effect is found statistically non-different from zero after the second quarter of 2013, it is important to highlight that the statistical evidence on positive effects on employment remains.

Finally, for the persistence of the impact on informality, we observe a similar pattern to that in the employment estimates. At least for the full age sample analyses, the effect is statistically significant only for the second quarter of 2013, and afterwards the impact becomes lower, but the impact is not reversed.

Nevertheless, the effect is statistically significant for all informal workers, whose estimated impact is -0.9%, and also for the waged informal workers, where the effect is between -1.6% and -1.7%. For self-employed informal workers, the coefficients remain negative but insignificant.

Therefore, this exercise confirms that there is no presence of adverse effects on the labour market by the minimum wage increase in Zone B. For the dynamics of the impact on real wages, the effect is observed immediately after the intervention, and it remains up to three quarters after the policy change. Whereas for the effect on employment and informality, the labour market takes two quarters to respond to the policy change, and the effect is only present in the second quarter of 2013.

Table 8

MEM's probit for the dynamics of the impact on labour status.
Second stage for sample selection bias correction.

Dependent variable:	*Labour market active*		*Employed*	
Specification:	(3a)	(3b)	(3a)	(3b)
ZoneB*D2013Q1	-0.0034	-0.0035	0.0009	0.0014
	(0.00454)	(0.00459)	(0.00215)	(0.00210)
ZoneB*D2013Q2	0.0018	0.0027	0.0017**	0.0020**
	(0.00478)	(0.00406)	(0.00085)	(0.00088)
ZoneB*D2013Q3	-0.0044	-0.0043	0.0019	0.0019
	(0.00393)	(0.00271)	(0.00155)	(0.00160)
ZoneB*D2013Q4	0.0015	0.0015	0.0019	0.0024
	(0.00659)	(0.00672)	(0.00167)	(0.00168)
Total observations	2,455,814		2,455,814	
Uncensored observations	1,405,478		1,405,478	

Note: the covariates included are gender, age, squared age, rural, schooling level, and interactions of schooling level with rural and gender.
Clustered standard errors at the state level in parentheses.
Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Table 9

MEM's for the dynamics of the impact on informality.
Second stage MEM's for sample selection bias correction.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Analysis sample:	Employed population		Formal & waged informal		Formal & self-employed informal		Formal & non-waged informal	
Dependent variable:	*informal*		*waged_informal*		*self_emp_informal*		*non_waged_informal*	
Specification:	(3a)	(3b)	(3a)	(3b)	(3a)	(3b)	(3a)	(3b)
ZoneB*D2013Q1	-0.0076	-0.0072	-0.0060	-0.0057	-0.0095	-0.0093	0.0011	0.0011
	(0.00957)	(0.00955)	(0.00772)	(0.00772)	(0.00761)	(0.00758)	(0.00262)	(0.00265)
ZoneB*D2013Q2	-0.0090***	-0.0084***	-0.0160***	-0.0161***	-0.0038	-0.0025	0.0028	0.0030
	(0.00276)	(0.00248)	(0.00311)	(0.00331)	(0.00533)	(0.00500)	(0.00187)	(0.00189)
ZoneB*D2013Q3	-0.0094	-0.0077	-0.0113	-0.0099	-0.0071	-0.0054	-0.0004	-0.0001
	(0.00943)	(0.00965)	(0.00953)	(0.00989)	(0.00680)	(0.00685)	(0.00223)	(0.00227)
ZoneB*D2013Q4	0.0013	0.0025	-0.0028	-0.0008	0.0053	0.0048	0.0003	0.0004
	(0.00698)	(0.00763)	(0.00764)	(0.00845)	(0.00625)	(0.00646)	(0.00163)	(0.00168)

Note: the covariates included are gender, age, squared age, schooling level, rural, and interactions of schooling level with rural and gender.
Clustered standard errors at the state level in parentheses.Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

5.5.1 Falsification test on wage estimates

Following the structure of the analyses on the dynamics of the effect, this section uses a falsification test to check the validity of the control groups in our estimates.

Using a specification similar to Autor (2003), this variation of the model now includes difference in differences estimators for every quarter in the sample even before the intervention. The logic behind the falsification test is that the only difference between control and treatment groups is precisely the policy intervention, so for legitimate control groups all those parameters before the intervention should be insignificant. That is, in absence of the policy reform control and treatment groups are identical, so there should be no effect.

In addition, to test the robustness of the model to the period of analysis, the sample is increased from 2011 to 2014.

$$\begin{aligned} Y_i = &\beta_0 + \sum_{j=1}^{7} \tau_j ZoneB_i * Q_{-ji} + \sum_{j=1}^{8} \delta_j ZoneB_i * Q_{ji} \\ &+ \sum_{j=1}^{7} \phi_j Q_{-ji} + \sum_{j=1}^{8} \eta_j Q_{ji} + \beta_1 ZoneB_i + \sum_{k=2}^{k} \beta_k X_{ki} + e_i \end{aligned} \tag{4a}$$

$$\begin{aligned} Y_i = &\beta_0 + \sum_{j=1}^{7} \tau_j ZoneB_i * Q_{-ji} + \sum_{j=1}^{8} \delta_j ZoneB_i * Q_{ji} + \sum_{j=1}^{7} \phi_j Q_{-ji} \\ &+ \sum_{j=1}^{8} \eta_j Q_{ji} + \sum_{j=1}^{7} \varphi_j ZoneA_i * Q_{-ji} + \sum_{j=1}^{8} \psi_j ZoneA_i * Q_{ji} \\ &+ \beta_1 ZoneB_i + \beta_2 ZoneA_i + \sum_{k=3}^{k} \beta_k X_{ki} + e_i \end{aligned} \tag{4b}$$

Where Q_j correspond to binary variables that identify the number of quarters j before and after the intervention. τ_j and φ_j are the pre-treatment parameters, corresponding to the period from the first quarter of 2011, to the third quarter of 2013. Our parameters of interest are δ_j, while ψ_j are the post-intervention parameters for Zone A. The quarter where the legislation came into force (4Q 2012) is omitted, so this period constitutes the base for comparison. With exception of the variable for the simple linear trend, which is contained in the quarterly dummy variables Q_{ji}, the set of control variables is the same to that used in Subsection 5.1, including the the state employment rate.

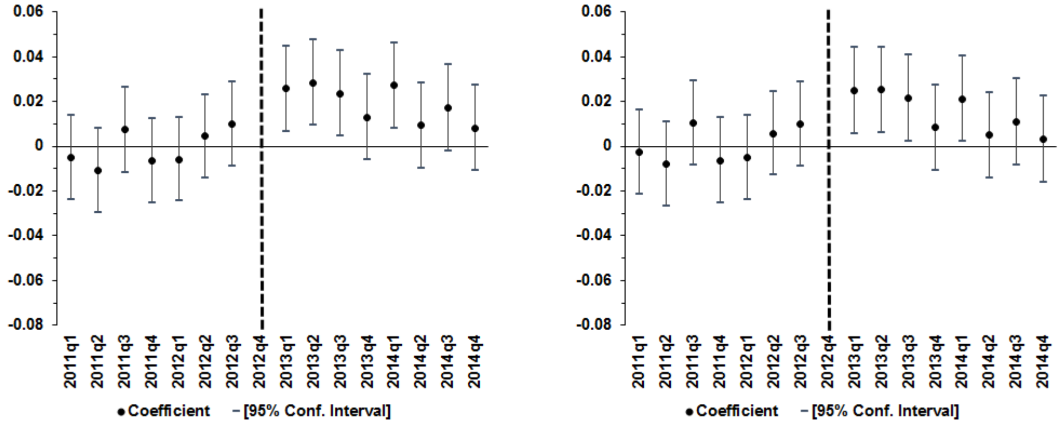
Figure 10 shows the value of the wage coefficients, as well as the 95% confidence intervals of the difference in differences estimators for the treated Zone B. For both specifications, the pre-treatment coefficients are statistically non-different from zero.

But, after the intervention, the parameters become positive and statistically significant. In line with our previous results in subsection 6, the effect is not statistically significant beyond the third quarter of 2013.

Then, this analyses confirms that the evaluation of the increase of minimum wages in Zone B using as a control groups zones A and C generates valid and successful estimators by difference in differences procedures.

Figure 10

Falsification test for the effect on hourly wages.
Difference in differences coefficients for Zone B



(a) Equation (4a) **(b)** Equation (4b)

6 Conclusions

This paper contributes to the discussion on the impact of minimum wages on the labour market providing strong empirical evidence on the existence of monopsonistic labour markets in Mexico.

Given the current context of the labour market, in which the real value of the minimum wage has decreased almost 70% in the last 30 years, more than a half of the population lives in poverty, and 60% of the labour force is employed in the informal sector, the implementation of public policies to adjust the value of the real minimum wage seems unavoidable. Using as a natural experiment the minimum wage harmonization of 2012, the study provides a formal evaluation of the labour market implications of a minimum wage rise.

Previous evaluations of the effects on minimum wages on the Mexican labour market have omitted the fact that restricting the analyses to active labour market population generates a problem of sample selection bias. Our analyses proves that there is evidence of biased estimators in the previous findings.

Robust to different specifications, difference in differences models correcting for sample selection do not find statistical evidence of adverse effects on the labour force. The results show positive effects on real wages and employment, as well as a reduction on the informal employment.

The central finding of this research is the existence of monopsonistic labour markets in Mexico, which implies that minimum wage increases (in real terms) do not necessarily mean negative effects on employment (or informality), specially for the most vulnerable workers.

Nevertheless, the analysis on the earnings distribution by unconditional quantile regressions suggests the presence of important spillover effects on the top of the distribution, which increased the wage dispersion. Previous literature suggests that the institutional wage setting in Mexico, specifically the role of the minimum wage as a reference rate, could be originating this lighthouse effect. Even though this institutional setting has been legally modified (by the 'de-indexation reform' of 2016), it is essential to consider this likely repercussions in the implementation of further policies in the labour market.

A new legislation, with stronger changes on the minimum wage is expected to occur at the end of 2016. This would represent the suitable source of variation to test the minimum wage effects on the labour market once the 'de-indexation reform' was implemented. Moreover it would be the opportunity to test the robustness of the estimates presented in this paper, but in presence of deeper variations in the minimum wage level.

Thus, the main contribution of this study is that it serves as a guideline to understand the Mexican labour market responses to changes in the minimum wage level, which is essential in the process of the design of a more intensive minimum wage legislation.

Finally, there are some analytical issues pending for future research. First, for the informal employment models, the labour status of the workers is restricted to formal and informal employment, but there are other possibilities; workers actually could transit to unemployment or even to be inactive in the labour market. To evaluate the complete set of choices faced by informal workers, multinomial logit regressions can be implemented, but sample selection correction under this methodological approach represents the main analytical challenge.

And second, the nature of the database allows for panel data analyses, but it implies to loose an important proportion of the observations. It is necessary to explore the validity of the results in order to incorporate fixed effects estimations to the set of evaluations.

References

Aeberhardt, R., Givord, P., and Marbot, C. (2015). Spillover effect of the minimum wage in france: An unconditional quantile regression approach.

Angrist, J. D. and Pischke, J.-S. (2008). *Mostly harmless econometrics: An empiricist's companion*. Princeton University Press.

Autor, D., Manning, A., and Smith, C. L. (2016). The contribution of the minimum wage to us wage inequality over three decades: a reassessment. *American Economic Journal: Applied Economics*, 8(1):58–99.

Autor, D. H. (2003). Outsourcing at will: The contribution of unjust dismissal doctrine to the growth of employment outsourcing. *Journal of Labor Economics*, 21(1):1–42.

Bell, L. A. (1997). The impact of minimum wages in Mexico and Colombia. *Journal of Labor Economics*, 15(S3):S102–S135.

Borah, B. J. and Basu, A. (2013). Highlighting differences between conditional and unconditional quantile regression approaches through an application to assess medication adherence. *Health economics*, 22(9):1052–1070.

Bosch, M. and Manacorda, M. (2010). Minimum wages and earnings inequality in urban Mexico. *American Economic Journal: Applied Economics*, pages 128–149.

Brown, C., Freeman, R. B., Hamermesh, D. S., Osterman, P., and Welch, F. R. (1995). Comments by reviewers. *Industrial and Labor Relations Review*, 48.

Brown, C., Gilroy, C., and Kohen, A. (1982). The effect of the minimum wage on employment and unemployment. *Journal of Economic Literature*, pages 487–528.

Calva, J. L. and Picard, A. A. (2007). *Empleo, ingreso y bienestar*, volume 11. UNAM.

Campos, R. M., Esquivel, G., and Santillán, A. S. (2015). El impacto del salario mínimo en los ingresos y el empleo en méxico. *Serie Estudios y Perspectivas - México No.162*.

Card, D. (1992a). Do minimum wages reduce employment? A case study of California, 1987–89. *Industrial & Labor Relations Review*, 46(1):38–54.

Card, D. (1992b). Using regional variation in wages to measure the effects of the federal minimum wage. *Industrial & Labor Relations Review*, 46(1):22–37.

Card, D. and Krueger, A. B. (1994). Minimum wages and employment: A case study of the fast-food industry in New Jersey and Pennsylvania. *The American Economic Review*, 84(4):772–793.

Card, D. and Krueger, A. B. (1995). *Myth and measurement*. Princeton University Press Princeton, NJ.

Castellanos, S. G., García-Verdú, R., and Kaplan, D. S. (2004). Nominal wage rigidities in Mexico: evidence from social security records. *Journal of Development Economics*, 75(2):507–533.

Dickens, R., Machin, S., and Manning, A. (1999). The effects of minimum wages on employment: Theory and evidence from Britain. *Journal of Labor Economics*, 17(1):1–22.

Dube, A., Lester, T. W., and Reich, M. (2010). Minimum wage effects across state borders: Estimates using contiguous counties. *The review of economics and statistics*, 92(4):945–964.

Fairris, D., Popli, G., and Zepeda, E. (2008). Minimum wages and the wage structure in Mexico. *Review of Social Economy*, 66(2):181–208.

Feliciano, Z. M. (1998). Does the minimum wage affect employment in Mexico? *Eastern Economic Journal*, pages 165–180.

Firpo, S., Fortin, N. M., and Lemieux, T. (2009). Unconditional quantile regressions. *Econometrica*, 77(3):953–973.

Flinn, C. J. (2006). Minimum wage effects on labor market outcomes under search, matching, and endogenous contact rates. *Econometrica*, 74(4):1013–1062.

Frolich, M., Melly, B., et al. (2010). Estimation of quantile treatment effects with stata. *Stata Journal*, 10(3):423–457.

Gob.Distrito-Federal (2014). Política de recuperación del salario mínimo en México y en el Distrito Federal. Propuesta para un acuerdo nacional. *http://salarioscdmx. sedecodf. gob. mx/documentos/Politica_de_recuperacion_de_Salarios_Minimos. pdf.*

Gramlich, E. M., Flanagan, R. J., and Wachter, M. L. (1976). Impact of minimum wages on other wages, employment, and family incomes. *Brookings Papers on Economic Activity*, pages 409–461.

Grossman, J. B. (1983). The impact of the minimum wage on other wages. *Journal of Human Resources*, pages 359–378.

Heckman, J. J. (1979). Sample selection bias as a specification error. *Econometrica: Journal of the econometric society*, pages 153–161.

Hirsch, B. T., Kaufman, B. E., and Zelenska, T. (2015). Minimum wage channels of adjustment. *Industrial Relations: A Journal of Economy and Society*, 54(2):199–239.

Kaplan, D. S. and Pérez-Arce, F. (2006). El efecto de los salarios mínimos en los ingresos laborales de México. *El Trimestre Econmico*, 73(289(1)):139–173.

Koenker, R. and Bassett, G. (1978). Regression quantiles. *Econometrica: journal of the Econometric Society*, pages 33–50.

Lee, D. S. (1999). Wage inequality in the united states during the 1980s: Rising dispersion or falling minimum wage? *Quarterly Journal of Economics*, pages 977–1023.

Lemos, S. (2009). Minimum wage effects in a developing country. *Labour Economics*, 16(2):224–237.

Machin, S. and Manning, A. (1994). The effects of minimum wages on wage dispersion and employment: Evidence from the UK Wages Councils. *Industrial & Labor Relations Review*, 47(2):319–329.

Maloney, W. and Mendez, J. (2004). Measuring the impact of minimum wages. evidence from Latin America. In *Law and Employment: Lessons from Latin America and the Caribbean*, pages 109–130. University of Chicago Press.

Manning, A. (2004). Monopsony in motion: Imperfect competition in labor markets. *ILR Review*, 57(2):85.

Mincer, J. (1976). Unemployment effects of minimum wages. *Journal of Political Economy*, pages S87–S104.

Montenegro, C. E. and Pages, C. (2004). Who benefits from labor market regulations? Chile, 1960–1998. In *Law and employment: Lessons from Latin America and the Caribbean*, pages 401–434. University of Chicago Press.

Mroz, T. A. (1987). The sensitivity of an empirical model of married women's hours of work to economic and statistical assumptions. *Econometrica: Journal of the Econometric Society*, pages 765–799.

Neumark, D. (2001). The employment effects of minimum wages: Evidence from a prespecified research design the employment effects of minimum wages. *Industrial Relations: A Journal of Economy and Society*, 40(1):121–144.

Neumark, D. and Wascher, W. (2006). Minimum wages and employment: A review of evidence from the new minimum wage research. Technical report, National Bureau of Economic Research.

Sabia, J. J. (2009). Identifying minimum wage effects: New evidence from monthly cps data. *Industrial Relations: A Journal of Economy and Society*, 48(2):311–328.

Schmitt, J. et al. (2013). Why does the minimum wage have no discernible effect on employment? *Center for Economic and Policy Research*, 22:1–28.

Stewart, M. B. (2002). Estimating the impact of the minimum wage using geographical wage variation. *Oxford Bulletin of Economics and Statistics*, 64(supplement):583–605.

Van de Ven, W. P. and Van Praag, B. M. (1981). The demand for deductibles in private health insurance: A probit model with sample selection. *Journal of econometrics*, 17(2):229–252.

Welch, F. (1974). Minimum wage legislation in the United States. *Economic Inquiry*, 12(3):285–318.

Wooldridge, J. M. (2010). *Econometric analysis of cross section and panel data*. MIT press.

Appendix A Variables construction

This appendix describes in a detailed way the process of variable generation procedure. The variables used are contained in the database *sociodem*, from ENOE, for all the quarters from 2011 to 2014.

Dependent variables

- **Wages**

ENOE provides two variables for earnings. On the one hand, *ingocup* reports the nominal monthly wage by worker, while ing_x_hrs contains the hourly wage.

Both variables are used for the analyses of the impact on wages, but in real terms. Nominal wages are deflated using the National Consumer Price Index (INPC), which is not contained in ENOE but it is also obtained from INEGI, and whose base period corresponds to the second fortnight of December 2010.[34]

In addition, it is relevant to mention that there is a discrepancy in the number of valid observations between monthly and hourly expressions. Hourly wages are obtained simply from the division of monthly wages over the number of hours worked (*hrsocup*); when the hours worked are not reported or wrongly answered, hourly wage is a missing value. Aiming to have a consistent sample size, observations with missing values for hourly wage are not considered in the regressions. These dropped observations represent 3.26% of the valid earnings sample. Estimates for the whole set of observations for monthly wage are not reported, but its exclusion does not affect significantly.

Thus, the logarithm of wage variables are generated by the following expressions:
$\ln(monthly\_wage_i) = \ln[(ingocup_i/INPC) * 100] \iff ing\_x\_hrs_i > 0$
$\ln(hourly\_wage_i) = \ln[(ing\_x\_hrs_i/INPC) * 100]$

- **Labour status**

To identify the labour status for each individual, the survey provides the variables $clase1$ and $clase2$. In the first of them, it is possible to diferentiate between active labour market individuals ($clase1 = 1$), from those inactive ($clase1 = 2$). $clase2$ disaggregates active and inactive population in employed ($clase2 = 1$), unemployed ($clase2 = 2$), available for working ($clase2 = 3$), and unavailable individuals

[34]It is not possible to construct price indices at the municipality level, so the national price index is employed to deflate wages. For the process of generation of the INPC, only 46 cities are included in the analyses. In consequence, many of the municipalities in ENOE are not considered for the INPC. A likely option was to generate approximations of the price indices by state or wage zone, but important biases can be produced specially in rural areas. Then, in order to avoid subjective methods to construct local price indices, national price index is used.

($clase2 = 4$). The dichotomous variables on labour status are constructed by the following way:

$labour\_market\_active_i = 1 \; \forall \; i \; clase1_i = 1$; $labour\_market\_active_i = 0$ otherwise.
$employed_i = 1 \; \forall \; i \; clase2_i = 1$; $employed_i = 0 \forall i \; clase2_i = 2$.

The definition of these variables implies that *employed* is restricted to active labour market population.

- **Informality status.**

ENOE follows the Hussman's Matrix criteria to identify formal and informal workers. We use the variable mh_col, which reports the columns of the matrix. The variable construction procedure is the following:

$informal_i = 1 \; \forall \; i \; (mh\_col_i = 1, 3, 5, 7, 9)$; $informal_i = 0$ otherwise.

$waged\_informal_i = 1 \; \forall \; i \; mh\_col_i = 1, 3$;
$waged\_informal_i = 0 \; \forall \; i \; mh\_col_i = 2, 4, 6, 8$.

$self\_employed\_informal_i = 1 \; \forall \; i \; mh\_col_i = 5, 7$;
$self\_employed\_informal\_i = 0 \; \forall \; i \; mh\_col_i = 2, 4, 6, 8$.

$non\_waged\_informal_i = 1 \; \forall \; i \; mh\_col_i = 9$;
$non\_waged\_informal_i = 0 \; \forall \; i \; mh\_col_i = 2, 4, 6, 8$.

It is important to remark that the values $mh\_col = 2, 4, 6, 8$ correspond to formal workers. So, for the analysis on the sub-categories of informality (waged, self-employed and non-waged), we are comparing them with respect to all formal workers. Transition from informality to unemployment or to inactive status is not considered.

Control variables. Minimum wage zones and post-treatment period dummy

- **Zone B and Zone A**

ENOE contains a variable to identify the minimum wage zone for each observation (*zona*). Nevertheless, using this variable is not possible to follow individuals in Zone B after the intervention; by construction these individuals belong to Zone A after 26 November 2012. In addition, there are some mistakes detected in the classification by INEGI; three municipalities are classified in Zone A, but according to CONASAMI they belong to Zone B (Apodaca, Zapopan, and Tlajomulco, all of them in the state of Jalisco), while four municipalities from Zone C are wrongly included in Zone B (Panuco and Pueblo Viejo, in the state of Veracruz, and Salinas Victoria and Juarez, in the state of Nuevo Leon).

In consequence, it was necessary to generate a variable that allows to identify minimum wage zones for all the period of analyses. The municipality classification by minimum wage zones is obtained from CONASAMI,[35] and using the codes by states and municipalities from INEGI, it is possible to identify the municipalities in ENOE database.

First, I generate a variable that allows to identify the municipality, by concatenating the state code (variable *ent*, two digits) and the municipality code (variable *mun*, three digits): $id\_mun =' ent' + `mun'$

Then, it is possible to construct the variables $ZoneB$ and $ZoneA$:

$ZoneB_i = 1 \; \forall \; i \; id\_mun_i = 14039, 14070, 14097, 14098, 14101, 14120, 19006, 19019, 19021, 19026, 19039, 19046, 19048, 26004, 26007, 26012, 26016, 26017, 26018, 26020, 26021, 26022, 26025, 26026, 26029, 26030, 26033, 26035, 26036, 26042, 26045, 26046, 26047, 26056, 26058, 26060, 26062, 26064, 26065, 26071, 26072, 28002, 28003, 28004, 28009, 28011, 28012, 28021, 28028, 28029, 28038, 28043, 30040, 30131, 28038, 28043, 30040, 30131, 30189;$
$ZoneB_i = 0$ otherwise.

$ZoneA_i = 1 \; \forall \; i \; ent_i = 02, 03, 09$ or $id\_mun_i = 08028, 08037, 08053, 12001, 15013, 15020, 15024, 15033, 15057, 15104, 15109, 15121, 26002, 26019, 26039, 26043, 26048, 26055, 26059, 26070, 30039, 30048, 30061, 30082, 30108, 30111, 30204, 30206, 28007, 28014, 28015, 28022, 28024, 28025, 28027, 28032, 28033, 28035, 28040;$
$ZoneA_i = 0$ otherwise.

- **Post-treatment period**

The policy change came into force on 27 November 2012. Given that the survey presents the information in a quarterly basis, for the fourth quarter 2012 it is necessary to differentiate individuals interviewed before and after the legislation. To do that, the variable *d_sem* is used, whose last two digits shows the number of the week when the interview took place for urban households (taking values from 01 to 13), or the month of the interview for rural households (taking values from 01 to 03).

$Period2_i = 1 \; \forall \; t \geq 2013Q1$, or
$Period2_i = 1 \; \forall \; i \; d\_sem_i \geq 09 \iff 2012Q4 = 1$ and $rural_i = 1$, or
$Period2_i = 1 \; \forall \; id\_sem_i = 03 \iff 2012Q4 = 1$ and $rural_i = 0$;
$Period2_i = 0$ otherwise.

[35] Retrieved from: `http://www.conasami.gob.mx/clasif_muni_area_geografica2.html`

Sociodemographic variables

- **State employment rate.**

  First, a constant for the number of active (*n_active*) and employed population (*n_employed*) by zones and by quarter are generated. Then, the employment rate is generated followind the occupation rate definition by INEGI:
  $EmpRate_i = (n\_employed/n\_active) * 100$

- **Head of the household.**
  $head_i = 1 \ \forall \ i \ par\_c_i = 101$; $head_i = 0$ otherwise.

  Variable *par_c* identifies the family relationship of each member of the household with respect to the head of the household.

- **Female individuals.**
  $female_i = 1 \ \forall \ i \ sex_i = 2$; $female = 0$ otherwise.

- **Age**
  $age_i = eda_i \ \forall \ i \ eda_i \leq 97$
  Observations with non-specified age are excluded. $eda = 99$ denotes non-specified age for workers older or equal to 12 years old. $eda = 98$ denotes non-specified age for workers younger than 12 years old.

- **Rural municipalities**
  $rural_i = 1 \forall i$ whose $t\_loc_i = 4$; $rural = 0$ otherwise.
  Variable *t_loc* describes the population size in the village or municipality. We follow the definition by INEGI, in which rural municipalities are those with a population lower than 2,500 inhabitants.

- **School Level**

  Primary basic school completed (from first to sixth year):
  $school\_level_i = 1 \ \forall \ i \ niv\_ins_i = 1$

  Secondary basic school completed (from seventh to sixth year):
  $school\_level_i = 2 \ \forall \ i \ niv\_ins_i = 2$;

  High school completed (from ninth to twelfth year):
  $school\_level_i = 3 \ \forall \ i \ niv\_ins_i = 3$;

  Undergraduate and Post-graduate degree:
  $school\_level_i = 4 \ \forall \ i \ niv\_ins_i = 4$;

  In all cases observations with non-specified level of education are excluded from the sample.

Appendix B Potential endogeneity of the state employment rate as a covariate

A valid concern with respect to the inclusion of employment rate as a regressor in the wage equation is its potential endogeneity with respect to the disturbance term e_i. In this case, the source of endogeneity is simultaneity because employment is the dependent variable in other structural equation estimated: the causal effect on labour status.

Nevertheless, the main argument for its inclusion is that given that the database provides information at the individual level, the state employment rate can be taken as given in the wage equation. First, the purpose is not to estimate the equilibrium wage at the national scale, nor the employment rate in equilibrium. Second, employment level by states is not the dependent variable when the effect of the Zone B minimum wage increase is estimated; it is a dummy variable indicating the individual labour status. And third, although zone dummies and separate linear trends by zones are included as regressors, it is not enough to control for the structural economic activity at a more local level. Given the size of the labour market in Mexico, it is possible to exploit the differences of the labour force participation by states (32 states). Municipalities belonging to the same wage zone can have very different characteristics in terms of development, employment, and even geographical location.

Then, employment rate at the state level is not necessarily correlated with the unexplained factors in the wage equation at the individual level. A relevant example of the inclusion of regional employment rates in the wage equation is Mroz (1987).[36] In this paper, aiming to correct previous misspecifications on female labour supply, Mroz performs an exhaustive analysis on the exogeneity assumptions of the woman's labour supply equation with respect to her wage rate. By the use of instrumental variables, he found that the use of experience variables as an instrument in the wage equation fails to be exogenous. To correct this endogeneity problem, several sets of instruments are used in the reduced form of the wage equation (family and regional background, children variables, husband's age and education, among others). County unemployment rate is always part of the set of instruments used in the baseline model, that is, the exogeneity of the regional employment rate in the wage equation is never subject to discussion.

In addition, a simple Durbin-Hu-Hausman test is implemented to verify the exogeneity of employment rate in the wage equation (Wooldridge, 2010).

The instruments used for the employment rate at the sate level are the following.

[36] With respect to minimum wage literature, Autor et al. (2016) also use sate-level employment rate as a control variable in the model to estimate the effect of minimum wage on inequality.

First, a categorical variable to control for the industrial and institutional composition of the firms, allowing to distinguish among governmental, agricultural formal and informal, and household units. Second, a variable to control for the size of the economic units within the states. And finally, to control for the geographical differences among the municipalities the minimum wage zones variable is used. In addition a dummy variable to identify the rural municipalities, and a simple linear trend variable are also included.

To test the null hypothesis that employment rate is not correlated with the error term of the wage equation, the test adds the fitted values of the residuals of the employment rate equation as a regressor in the wage equation. In spite of the number of observations, the coefficient of the residuals is statistically equal to zero (-0.0009 with a p-value of 0.785). Therefore, we cannot reject the null hypotheses of the exogeinity of employment rate in the wage equation.

Table B.1 shows the parameters of both equations of the Durbin-Hu-Hausman test. We also carried on the test including in the employment rate equation the DiD parameters, as well as the post-treatment period dummy. This specification does not alter the results of the test: employment rate is not endogenous in the wage equation.

Table B.1

Durbin-Hu-Hausman test.

Exogeneity of employment rate in the wage equation

(a) Reduced form. Employment

Dependent Variable:	*EmpRate*	
Institutional sector of the firm		
Non-financial firms	-0.0841***	(0.01172)
Formal non-agricultural	-0.0482***	(0.01201)
Governmental	0.8056***	(0.09688)
Private non-profit org	-0.0156	(0.01364)
Public non-profit org	0.1333***	(0.01224)
Household informal	-0.0447***	(0.01245)
Waged household work	0.4072*	(0.21127)
Size of the establishment		
No establishment	-0.0905***	(0.00424)
Small establishment	-0.1198***	(0.00563)
Medium establishment	-0.1680***	(0.00665)
Large establishment	-0.3531***	(0.00672)
Governmental	-0.8565***	(0.09618)
Other	-0.4715**	(0.21090)
Rural	-0.1110***	(0.00370)
Trend	-0.0234***	(0.00049)
Zone B	-0.0714***	(0.00397)
Zone C	0.6051***	(0.00371)
Observations	851,475	

Robust standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

(b) Durbin-Hu-Hausman test.

Dependent Variable:	ln(*hourly_wage*)	
Residuals EmpRate	-0.0009	(0.00313)
ZoneB*Period2	0.0369***	(0.00908)
ZoneB	0.6686	(0.41462)
Period2	0.0061**	(0.00302)
TrendB	-0.0095***	(0.00187)
TrendA&C	-0.0069***	(0.00065)
EmpRate	-0.0410***	(0.00304)
Female	-0.3117***	(0.00490)
Age	0.0451***	(0.00030)
Age^2	-0.0004***	(0.00000)
Rural	-0.1879***	(0.00671)
School_level2	0.0929***	(0.00379)
School_level3	0.1948***	(0.00359)
School_level4	0.5829***	(0.00375)
School_level2*Rural	0.0705***	(0.00840)
School_level3*Rural	0.0549***	(0.00771)
School_level4*Rural	0.0121	(0.00905)
School_level2*Female	0.0838***	(0.00591)
School_level3*Female	0.1832***	(0.00542)
School_level4*Female	0.2808***	(0.00560)
Observations	850,933	

Robust standard errors in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Appendix C The impact on real monthly wages and its distribution

Table C.1

The impact on monthly wages.

Heckman second stage for sample selection bias.

Dependent variable:	$\ln(monthly\_wage)$			
Equation:	(1a)		(1b)	
Full age threshold: 12 ≤ Age ≤ 97				
ZoneB*Period2	0.0195**	(0.00967)	0.0161*	(0.00974)
$\hat{\lambda}$ (IMR)	-0.3667***	(0.00569)	-0.3667***	(0.00569)
Total observations	2,112,508		2,112,508	
Uncensored observations	960,550		960,550	
Age threshold: 12 ≤ Age ≤ 29				
ZoneB*Period2	0.0199	(0.01497)	0.0136	(0.01508)
$\hat{\lambda}$ (IMR)	-0.4832***	(0.01073)	-0.4833***	(0.01073)
Observations	886,481		886,481	
Uncensored observations	309,008		309,008	
Age threshold: 30 ≤ Age ≤ 49				
ZoneB*Period2	0.0211	(0.01386)	0.0203	(0.01396)
$\hat{\lambda}$ (IMR)	-0.4240***	(0.00880)	-0.4239***	(0.00880)
Total observations	687,799		687,799	
Uncensored observations	450,695		450,695	
Age threshold: 50 ≤ Age ≤ 97				
ZoneB*Period2	-0.0019	(0.02443)	-0.0053	(0.02460)
$\hat{\lambda}$ (IMR)	-0.2110***	(0.01485)	-0.2111***	(0.01485)
Total observations	538,228		538,228	
Uncensored observations	337,381		337,381	

Note: the covariates included are state employment rate, gender, age, squared age, schooling level, and interactions of schooling level with rural and gender.

Observations with non-reported wages are excluded from the analysis.

Standard errors in parentheses, by two-step variance estimator Heckman (1979).

Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Table C.2

The impact on monthly wages by formality condition.
Heckman second stage for sample selection bias.

	Formal workers				**Informal workers**			
Dependent variable:	ln(*monthly_wage*)				ln(*monthly_wage*)			
Equation:	(1a)		(1b)		(1a)		(1b)	
Full age threshold: 12 ≤ Age ≤ 97								
ZoneB*Period2	0.0170*	(0.01010)	0.0161	(0.01020)	0.0213	(0.01499)	0.0166	(0.01507)
$\hat{\lambda}$ (IMR)	-0.1762***	(0.00501)	-0.1763***	(0.00501)	(0.00724)	-0.3738***	(0.00724)	-0.3738***
Total observations	1,567,202		1,567,202		1,664,940		1,664,940	
Uncensored Observations	428,783		428,783		531,767		531,767	
Age threshold: 12 ≤ Age ≤ 29								
ZoneB*Period2	0.0268*	(0.01495)	0.0241	(0.01511)	0.0225	(0.02354)	0.0159	(0.02366)
$\hat{\lambda}$ (IMR)	-0.2554***	(0.00839)	-0.2555***	(0.00839)	-0.4464***	(0.01366)	-0.4464***	(0.01366)
Total observations	704,294		704,294		751,993		751,993	
Uncensored observations	130,639		130,639		178,369		178,369	
Age threshold: 30 ≤ Age ≤ 49								
ZoneB*Period2	0.0199	(0.01432)	0.0193	(0.01446)	0.0201	(0.02263)	0.0190	(0.02275)
$\hat{\lambda}$ (IMR)	-0.1872***	(0.00773)	-0.1871***	(0.00773)	-0.4015***	(0.01051)	-0.4014***	(0.01050)
Total observations	452,579		452,579		455,842		455,842	
Uncensored observations	221,419		221,419		229,276		229,276	
Age threshold: 50 ≤ Age ≤ 97								
ZoneB*Period2	-0.0110	(0.02845)	-0.0103	(0.02871)	0.0002	(0.03370)	-0.0072	(0.03389)
$\hat{\lambda}$ (IMR)	-0.0890***	(0.01726)	-0.0894***	(0.01726)	-0.1776***	(0.01847)	-0.1774***	(0.01847)
Total observations	410,329		457,105		457,105		457,105	
Uncensored observations	76,725		76,725		124,122		124,122	

Note: the covariates included are state employment rate, gender, age, squared age, schooling level, rural, and interactions of schooling level with rural and gender. Observations with non-reported wages are excluded from the analysis.
Standard errors in parentheses, by two-step variance estimator Heckman (1979).
Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Table C.3

The impact on the monthly earnings distribution. Full age threshold: $12 \leq \text{Age} \leq 97$

Dependent variable:	ln(*hourly_wage*)											
	Pooled OLS		q10		q25		q50		q75		q90	
(a) Pooled sample, formal and informal workers												
Equation (1a)												
ZoneB*Period2	0.0305***	(0.00439)	0.0070	(0.00443)	0.0135***	(0.00417)	0.0339***	(0.00573)	0.0562***	(0.00844)	0.0459***	(0.01200)
Equation (1b)												
ZoneB*Period2	0.0277***	(0.00443)	0.0062	(0.00451)	0.0121***	(0.00423)	0.0317***	(0.00577)	0.0521***	(0.00849)	0.0412***	(0.01207)
Observations	767,006		767,006		767,006		767,006		767,006		767,006	
(b) Formal workers												
Equation (1a)												
ZoneB*Period2	0.0291***	(0.00543)	0.0128**	(0.00643)	0.0184***	(0.00564)	0.0346***	(0.00758)	0.0338***	(0.00955)	0.0566***	(0.01272)
Equation (1b)												
ZoneB*Period2	0.0251***	(0.00550)	0.0132**	(0.00655)	0.0154***	(0.00571)	0.0289***	(0.00765)	0.0297***	(0.00965)	0.0516***	(0.01285)
Observations	405,217		405,217		405,217		405,217		405,217		405,217	
(c) Informal workers												
Equation (1a)												
ZoneB*Period2	0.0268***	(0.00728)	0.0292***	(0.00781)	0.0043	(0.00781)	0.0150**	(0.00694)	0.0341***	(0.01155)	0.0413**	(0.01842)
Equation (1b)												
ZoneB*Period2	0.0256***	(0.00731)	0.0294***	(0.00789)	0.0032	(0.00788)	0.0141**	(0.00698)	0.0339***	(0.01159)	0.0386**	(0.01846)
Observations	361,789		361,789		361,789		361,789		361,789		361,789	

Note: the covariates included are state employment rate, gender, age, squared age, schooling level, rural, and interactions of schooling level with rural and gender.
Observations with non-reported wages are excluded from the analysis.
Bootstrapped standard errors in parentheses, 100 repetitions. Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.

Table C.4

The dynamics for the impact on monthly wages.
Second stage for Heckman correction for sample selection bias.

Dependent variable:	ln(*hourly_wage*)			
Equation:	(1a)		(1b)	
ZoneB*D2013_Q1	0.017**	(0.0080)	0.019**	(0.0079)
ZoneB*D2013_Q2	0.018**	(0.0080)	0.022***	(0.0079)
ZoneB*D2013_Q3	0.007	(0.0081)	0.011	(0.0081)
ZoneB*D2013_Q4	0.002	(0.0079)	0.005	(0.0079)
$\hat{\lambda}$ (IMR)	-0.368***	(0.0057)	-0.368***	(0.0057)
Total observations	2,112,508		2,112,508	
Uncensored observations	960,550		960,550	

Note: the covariates included are state employment rate, gender, age, squared age, rural, schooling level, and interactions of schooling level with rural and gender.
Observations with non-reported wages are excluded from the analysis.
Standard errors in parentheses, by two-step variance estimator Heckman (1979).
Statistical significance: *** p<0.01, ** p<0.05, * p<0.1.