Wage Inequality in a Developing Country:
Decrease of Minimum Wages or Increase of
Education Returns∗

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Abstract

In this paper we analyze the increasing wage inequality observed in the Uruguayan labour market during the last decade, by studying how the changes in minimum wages and returns to education have affected the wage structure. Despite of the fact that in most developed countries an important proportion of the increased wage inequality is explained by a fall of the real minimum wages, this is not the case for the Uruguayan labour market. We find that the increased wage dispersion is mostly explained by shifts of the upper tail of the wage distribution, i.e. substantial wage increments for the most skilled-educated workers. To derive these conclusions we follow a parametric and nonparametric quantile regression approach.

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1.-INTRODUCTION

Labour earnings are by far the most important component of income for individuals working in urban areas. As a consequence, the level of wage inequality generated by the labour market is fundamental for those interested in understanding poverty, internal migrations or economic incentives facing workers (Blau and Khan, 1996). While in developed countries the evolution of wages has attracted a large amount of interest recently, mainly because of the observed increase in wage dispersion, there are not too many studies analyzing the wage structure of developing countries, as pointed out by Harrison and Leamer (1997; see also MacIsaac and Rama, 1997 and references therein).

In this paper we are concerned with studying the wage structure of the Uruguayan labour market from 1986, the end of the military regime, until 1997, the last year of available data. In particular, we are interested with whether the explanations given in developed countries to the increase in wage dispersion are appropriate for the Uruguayan labour market. Although there is no consensus concerning the fundamental causes of wage dispersion in developed countries, several explanations have received much attention: decline in manufacturing employment and a shift towards educational intensive sectors; shifts in relative demand for labour favoring more-educated, skilled and flexible workers; changes in wage-setting institutions or pay norms; fall of the legal minimum wage; increased openness of the economies, are among the most common explanations (see Bound and Johnson, 1992; Murphy et al., 1992; Katz and Murphy, 1992; Juhn et al., 1993; Freeman, 1995; Wood, 1995; Meghir and Whitehouse, 1996; Topel, 1997; Fortin and Lemieux, 1997; Moane and Wallerstein, 1997, among others). Though we will focus basically on the minimum wage and returns to education, we will mention the other possible causes taking into account that there are no unique nor exclusive explanations to the increase in wage
To discuss the evolution of the Uruguayan wage structure we follow an empirical approach based on parametric and nonparametric quantile regression analysis. On the one hand, to study the possible effects of the fall of minimum wages on the increase in wage inequality we estimate the nonparametric quantiles of the wage distribution conditional on education and experience. Clearly, the analysis of quantiles is a way of characterizing the probability distribution of wages. What we observe is that the lower quantiles of this distribution have moved upwards between 1986 and 1997 despite of the fact that the real hourly minimum wage fell by nearly fifty percent during this period. This result is the opposite to the one found in developed countries, where the lower tail of the wage distribution shift downward following the fall of the minimum wages (see Fortin and Lemieux, 1997; Lee, 1998 among others).

On the other hand, returns to schooling analysis is based on the estimation of the quantile Mincer specification regression equation, following what is usual in this literature. However, the novelty of the application we present here is the fact that we test the Mincer specification for the conditional mean and quantiles using tests which are consistent against nonparametric alternatives (Zheng, 1998; Stute, 1997). The test rejects the Mincer specification for the conditional mean of the real hourly wages and for some but not all the conditional quantiles considered between 1986 to 1997. Upto our knowledge, it is the first time the Mincer equation specification is tested against nonparametric alternatives.

The paper is divided in six sections. In the next section we discuss the data used. In section three we study the changes in the wage structure following a descriptive approach. In section four we discuss the minimum wage effects on the increase in wage dispersion. In section five we analyze the returns to schooling. In section six we conclude.
2.-DATA

This study is based on data from the Household Survey of Uruguay from 1986 through 1997 (Encuesta de Hogares, Instituto Nacional de Estadística, Uruguay). The survey frame is the civilian population of Uruguay living in housing units, decomposed in a survey for the metropolitan area of Montevideo and another for the population living in cities in the rest of the country. Here, we used the survey corresponding to the Montevideo given that: first, nearly a half of the Uruguayan population lives in the metropolitan area of Montevideo, while in the rest of urban country the population is dispersed in particularly small cities\(^1\); second, two thirds of Uruguayan economic activity takes place in Montevideo; finally, nearly all university colleges are in Montevideo.

The survey contains individual data on monthly labour earnings, non labour earnings, age, sex, educational level, hours worked per week, marital status, occupation characteristics, and other relevant variables.

The sample we used here is composed of all the males older than 13 years old, given that this is the legal working age in Uruguay and including: a) those who worked during the week prior the interview; b) those whose current job were either in the private or in the public sector (excluded are persons who were self-employed, working without pay, entrepreneurs, or who had never worked), and c) those who earn a positive wage in that period of reference. The variable of interest is the real hourly wage (see DiNardo et al. 1996 for a justification of the use of real hourly wages).

\(^1\)Uruguayan total population is of 3.2 million people, with 90 percent living in cities: 45% in Montevideo; five of the biggest cities in the rest of the country, with less than one hundred thousand inhabitants, sum up to a 10% of total urban population; the rest lives dispersed in cities generally smaller than 30.000 inhabitants (Instituto Nacional de Estadística, Censo 1996 Uruguay).
3.-REAL HOURLY WAGE CHANGES 1986-1997

In this section we describe the changes in real hourly wages (rhw) for working men in Montevideo for the 1986-97 period and three subperiods, 1986-1989, 1989-1994 and 1994-1997, corresponding to the different democratic administrations that followed the military regime ended in 1985. Before going on, we point out two historical facts that could help to understand the evolution of wages during these periods. The first democratic administration after the military regime legalized labour unions and establish a centralized collective bargaining where wages were settled uniformly for all workers. During the two subsequent administrations, government intervention in the labour market diminished, wage bargaining were decentralized, workers’ participation in labour unions began to diminish and an increasing openness process began in the 1990s with the consolidation of the free trade commerce zone, Mercosur.

A first approximation to the evolution of wages is presented in Table 1, where we show the rhw rate of change at different points of wage distribution and for different education-experience groups for each of the periods detailed above. From this table we observe that all workers enjoyed an increase of their rhw over the full period, though there are important differences between these rates of change, i.e. while for those at the tenth percentile of the wage distribution the increase was of about 20 percent, for those at the ninetieth percentile it was of 70 percent. Also, the different education-experience groups present significant differences in their median rhw rate of change, where more educated- experienced received much higher rhw increases than

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2In the eighties, 4 of every 10 workers were members of labour unions. In 1997 only 1 of every 10 (Cassoni et al. 1999); Between 1990 and 1997 imports of manufactured goods multiplied nearly by 3, exports only by 1.5, while flow trades were reoriented towards Mercosur countries, i.e., in 1997, 50 percent of total Uruguayan trade was with Argentina and Brazil; The contribution to the GDP of manufacturing production fell from 26 to 20 percent, substituted by an increase in service (retail trade, restaurant etc. ) and tourism sector participation.
less educated-experienced workers. This divergence in wage increments suggests that
the Uruguayan labour market has moved towards a greater wage inequality.

Insert Table 1

This table also shows that the timing of the rhw movements was clearly different
between subperiods. Between 1986 and 1989, the centralized collective bargaining
increased wages substantially and uniformly for all workers at different points of the
wage distribution, i.e. near a 30 percent for all of them. In the subsequent adminis-
trations, the rate of change of the rhw was significantly different between workers at
distinct percentiles, i.e. an increase of 5 percent for those at the ninetieth percentile
against a 10 percent fall for those at the tenth percentile between 1994 and 1997. This
dispersion in the rate of rhw change corresponds not only to the decentralization of
collective bargaining but also to the increasing openness linked with the exchange
rate stabilization policies that the Mercosur countries were undergoing during the
nineties.

In Figure 1 we present the median, tenth and ninetieth percentile of the hourly real
wage distribution of working men for 1986-97, together with the minimum wage and
men unemployment in the metropolitan area of Montevideo.

Insert Figure 1

Regard first the dramatic fall of the legal real minimum wage during the whole period,
i.e. it fell a 55 percent in the whole period. Notice also that the tenth percentile of
the rhw distribution did not follow the drop of the minimum wages, which anticipates
the idea that the lower tail of the wage distribution did not shift towards the lower
minimum wage. The erosion of the real minimum wage was linked to the target of
diminishing public deficit, given that the increases of many social benefits, such as
pensions, annuities or poverty transfers, were indexed to the evolution of the minimum wages.

This figure clearly shows the increase in wage inequality in the Uruguayan labour market during this period, i.e. the distance or the ratio between the rhw of the least skilled workers, as measured by the tenth percentile of the wage distribution, and the most skilled workers, as measured by the ninetieth percentile, increased dramatically during the overall period. This increase in wage dispersion has been slightly larger above the median than below it, i.e. the ninetieth median ratio is slightly larger than the median-tenth percentile ratio, as observed from the bottom graph in the figure.

To sum up, the evolution of wages during the last decade suggests an increase in wage dispersion in the Uruguayan labour market. In the following sections we discuss whether this increase can be attributed to the dramatic fall of the minimum wage, i.e. by shifting downward the lower tail of the wage distribution or to the increase of the rhw of the most skilled-educated workers.

4.- THE LOWER TAIL OF THE WAGE DISTRIBUTION: DECREASE OF MINIMUM WAGE

The more basic government intervention in the labour market consists of the legislation of minimum wages. Regardless of its possible effects on employment, earnings or income, the minimum wage sets an explicit floor on the wage distribution, i.e. acting as a backstop for the bottom end of the wage distribution, it should tend to reduce wage dispersion (Fortin and Lemieux, 1997). As a consequence, movements of this bottom floor should affect the dispersion of the wage distribution. Di Nardo, Fortin and Lemieux (1996) demonstrate that the minimum wage is an important empirical feature of the observed distribution of United States wages during the eigthies. Lee (1998) argues that decreases in the minimum wage tend to increase measured wage inequality.
However, in this section we show that in the Uruguayan labour market the decrease of minimum wages did not pushed downward the lower tail of the wage distribution. Nevertheless, the effect of this minimum wage decrease could have been to limit the rate of increase of the rhw for workers in the lower percentiles of the wage distribution, i.e. the lower minimum wage left more margin in the decentralized bargaining to settle smaller wage increments for lower earning workers. In this sense, the fall of the minimum wage affected wage dispersion by constraining the wage increase of rhw at the lower tail of the distribution, while those in the upper tail could freely increase.

In Figure 2 we present the nonparametric density estimates of the log rhw together with the log of the legal minimum rhw for the years 1986 and 1997, considering different educational-experience groups and where we have explicitly undersmoothed.

Insert Figure 2

From this figure it is clear that the lower tail of the wage distribution did not collapse towards the lower minimum wage in 1997. Practically no worker was paid the minimum wage in this last year while in 1986 only the less educated workers were eligible for minimum wage jobs. In addition, the minimum wage seems no to act as a bottom floor for the wage distribution, i.e. there is no mass accumulation near the minimum wage, as it should be the case if an important proportion of workers were paid the minimum wage (see, for example Fortin and Lemieux, 1997). Finally, the most striking fact of this figure is the significant difference of wage distributions between less and more educated workers, at all levels of experience, which suggests that returns to schooling are high in the Uruguayan labour market.

To characterize the movements of the lower tail of the wage distribution we have estimated the fifth and tenth percentiles of the distribution of rhw conditional on education and experience by nonparametric methods (see Magee et al 1991; Mehra et al. 1991 or Pagan and Ullah, 1999 among other). On the one hand, quantile functions
characterizes the probability distribution of a random variable. On the other hand, not assuming any functional forms avoid the problem of misspecification.

Following Magee et al., let $q_\theta(x)$ be the $\theta$ quantile of the distribution of $Y$ given $X = x$, $F(q_\theta(x) | x) = \theta$, where $F(y | x)$ is the conditional distribution of $Y$ given $x$ evaluated at $y$. An estimate of $q_\theta(x)$ is the one that solves $\hat{F}(q_\theta(x) | x) = \theta$, were $F$ is estimated by

$$\hat{F}(y | x) = \frac{\sum_i K(u) I(Y_i < y)}{\sum_i K(u)},$$

$u = (X_i - x)/h$, $K(\cdot)$ is a kernel function and $I(\cdot)$ is the indicator function and $h$ is the bandwidth parameter. The asymptotic properties of these estimators can be found in Härdle (1984) or Robinson (1984).

To avoid the problem of course of dimensionality, we follow the idea of Philips (1983) or Katz and Murphy (1992), of dividing the sample in different cells depending on workers characteristics and taking into account the number of observations in each cell (see Delgado and Miles, 1997). Here we divided the sample in three groups in terms of years of experience: new entrants in the labour market, those upto 5 years of experience; consolidated workers, those with 6 to 15 years of experience and old workers, those with more than 16 years of experiences. For each of these groups we estimate the fifth and tenth percentiles of the log rhw distribution conditional on the level of education.

Insert Figure 3

The graphs of this figure shows the nonparametric estimates of the fifth and tenth percentiles for each year, i.e. $q_{0.05}_86$, $q_{0.05}_97$, $q_{0.10}_86$ and $q_{0.10}_97$ respectively, together with the minimum wages in both years in those cases where the range of these percentiles includes the value of the minimum wages.

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3All the estimators presented here were programmed by the authors in GAUSS and are available upon request.
Regard first that in most cases both percentiles has moved upwards between 1986 and 1997, i.e. the lower tail of the wage distribution shift to the right despite of the left movement of the minimum wage. As suggested before, the decrease in real minimum wage did not increase wage dispersion by shifting downward the wage distribution. Second, only low educated workers earned less than the minimum wage in 1986, while in 1997 mostly all of these lower earning workers were paid above the corresponding minimum wage. Third, the range of wages of both percentiles is significantly higher for more educated workers than for less educated ones, as should be expected from the discussion of Figure 2.

The fact that wage dispersion increased while the lower percentiles of the wage distribution were also increasing indicates that the rate of change of these percentiles was much smaller than the rate of change of the percentiles at the upper tail of wage distribution, i.e. most experienced-educated workers experienced much higher wage increases than least experience-educated ones (see Table 1). Clearly, the drop of the minimum wage could have slow down the rate of increase of rhw for those workers in the lower tail of the wage distribution.

5-HIGHER RETURNS TO MOST SKILLED-EDUCATED WORKERS

In Figure 4 we present the ratio of the nonparametric ninetieth, median and tenth percentiles between 1997 and 1986 for each of the subgroups considered, i.e. \( \frac{q_{90\_97}}{q_{90\_86}} \) represents the ratio between the ninetieth percentile in 1997 to the ninetieth percentile in 1986. We use this ratio as an estimator of the theoretical ratio.
between both percentiles, which can be interpreted as a percentile rate of change\textsuperscript{4}.

Insert Figure 4

Notice, first, that in general the ninetieth percentile ratio is larger than the tenth percentile ratio for mostly all levels of education, i.e. most skilled workers, as measured by the ninetieth percentile of the distribution of each subgroup, had enjoyed relatively higher wage increases than least skilled workers, as measured by the tenth percentile. On the other hand, the range of variation of these ratios is higher for more educated workers, thirteen or more years of education, than for less educated workers. That is, more educated workers experienced higher wage increases than less educated ones, i.e. the more skilled-educated workers had enjoyed higher wage increases than the least skilled-educated workers. Finally, the behaviour of these ratio is clearly different between each subgroup, i.e. for less educated workers, the three graphs present distinct evolution patterns of the estimated quantiles in terms of age. The erratic evolution of these ratios observed in the graphs for more educated workers, 13 or more years of education, could be attributed to the lack of observations in the upper tail of the education distribution, which makes the nonparametric estimator to estimate the different percentiles so as to be practically the same for higher levels of education.

Given this higher raw increase for more skilled-educated workers in what follows we present the quantile regression estimation of the Mincer equation to evaluate returns to schooling and experience. However, instead of directly estimating the Mincer equation, as is usual in the literature, we have first tested this specification using tests which are consistent against nonparametric alternatives, i.e. the power of the test statistic tends to one when we move out of the null in any direction.\textsuperscript{4}

\textsuperscript{4}Given that \(g(Y_n, X_n) = Y_n / X_n\) is a regular function, if \((Y_n, X_n)\) has a bivariate asymptotical distribution, it is possible to derive the asymptotic distribution of the ratio, so as to build confidence intervals to study if they are significantly different.
In order to test $H_0$ in the direction of nonparametric alternatives $H_1$, two alternative strategies have been developed in the literature. On the one hand, tests have been constructed based on a distance between a nonparametric fit and its parametric counterpart (see, for example, Härdle and Mammen, 1993). An advantage of these tests is that their null asymptotic distribution is a normal. A disadvantage is that size and power properties of the asymptotic test depend on the choice of the amount of smoothing. On the other hand, the estimation of the model under the alternative hypothesis and hence, the choice of the bandwidth parameter can be avoided by using tests based on weighted empirical process as suggested by Bierens (1982, 1991), Bierens and Ploberger (1996), Stute (1997) and Delgado and Dominguez (1997).

For testing quantile regression specifications, to our knowledge, the only test which has been published is the one developed by Zheng (1998), which falls in the first type of tests described above. Therefore, we have tested the Mincer specification for the conditional quantiles using Zheng (1998) and for the conditional mean using the test developed by Stute (1997).

The null hypothesis to be tested is

$$H_0 : \Pr (F (g (x, \beta_0) \mid x) = \theta) = 1$$

in the conditional quantile case or

$$H_0 : \Pr (m (x) = g (x, \beta_0)) = 1$$

in the conditional mean case, where $E (Y \mid X = x) = m (x)$, for some $\beta_0 \in B, g (x, \beta_0)$ is the Mincer parametric specification. The alternative hypothesis is the general negation of the null. The statistic developed by Zheng (1998) is given by

$$T_n = nh^{m_2/2} W_n / \hat{\sigma} \xrightarrow{d} N (0, 1)$$
where $m_2$ is the number of continuous random variables in the model,

$$ W_n = \frac{1}{n(n-1)} \sum_{i=1}^{n} \sum_{j \neq i}^{n} \frac{1}{h^{m_2}} K \left( \frac{x_i - x_j}{h} \right) \times \left( I \left( y_i \leq g \left( x_i, \hat{\beta} \right) \right) - \theta \right) \left( I \left( y_i \leq g \left( x_i, \hat{\beta} \right) \right) - \theta \right) $$

where $\hat{\beta}$ is an estimator of $\beta$ and

$$ \hat{\sigma}^2 = 2\theta^2 (1-\theta)^2 \frac{1}{n(n-1)} \sum_{i=1}^{n} \sum_{j \neq i}^{n} \frac{1}{h^{m_2}} K^2 \left( \frac{x_i - x_j}{h} \right) $$

The bandwidth $h$ is selected by minimizing the generalized cross-validation function

$$ GCV (h) = \frac{\sum_{i=1}^{n} \left( I \left( y_i \leq g \left( x_i, \hat{\beta} \right) \right) - \theta - \hat{m}_h (x_i) \right)}{n \left( 1 - tr \left( H \right) / n \right)^2} $$

where

$$ \hat{m}_h (x_i) = \frac{\sum_{i=1}^{n} K (u) \left( I \left( y_i \leq g \left( x_i, \hat{\beta} \right) \right) - \theta \right)}{\sum_{i=1}^{n} K (u)} $$

with $u = (x_i - x) / h$ and $tr \left( H \right)$ is the trace of the matrix $H = (h_{i,j})_{n \times n}$ with

$$ h_{i,j} = \frac{K \left( \frac{x_i - x_j}{h} \right)}{\sum_{i=1}^{n} K \left( \frac{x_i - x_j}{h} \right)} $$

The test developed by Stute (1997) is based on the weighted empirical process

$$ R_n (x) = \frac{1}{n} \sum \left( Y_i - g \left( x_i, \hat{\beta} \right) \right) I (x_i \leq x) $$

and the statistic adopts a Cramer-Von Misses form, $T_n^2 = \sum R_n (x)^2$. The main disadvantage of $T_n^2$ is that its asymptotic null distribution is not distribution free. However, Stute et al. (1996) have proposed to approximate the critical values by wild bootstrap$^5$.

$^5$Both tests were programmed by the authors in GAUSS. The number of rejections when simulating Zheng test, for $n=100$, was of 0.048 under the null and 0.991 under the first of the alternatives established by Zheng in his article. Simulations of Stute test are presented in Miles and Mora (1999). For computing regression quantiles the Fortran algorithm of Koenker and D’Orey (1987) was translated to GAUSS language. All of these programs are available upon request.
In Table 2 we present the p-value of the Stute and Zheng statistics when testing the Mincer specification to the Uruguayan data, i.e. the logarithm of rhw as dependent variable and a constant, education, experience and experience squared as independent variables.

The most striking feature of this table is the fact that Stute statistic rejected the null in all cases whilst the Zheng statistic did not basically reject the null. As stated before, a disadvantage of Zheng statistic is that it depends on the selection of a smoothing parameter and consequently, we could force other values of Zheng’s statistic by changing this parameter. In this application we have selected the smoothing parameter by minimizing the GCV, as stated before, which gives some objectivity to the results of the test.

Therefore, given that we have not rejected the Mincer quantile specification, we have estimated the parameters of the quantile Mincer specification regression. Following Koenker and Bassett (1978), this can be written as

\[ y_i = x_i' \beta_{\tau} + u_{\tau i} \quad \text{with} \quad Q_{\tau} (y_i \mid x_i) = x_i' \beta_{\tau} \]

for \( i = 1, \ldots, n \), where \( \beta_{\tau} \) and \( x_i \) are \( K \times 1 \) vectors and \( x_{i1} = 1 \). \( Q_{\tau} (y_i \mid x_i) \) denotes the \( \tau \)th conditional quantile of \( y \) given \( x \), that is, \( Q_{\tau} (y_i \mid x_i) = \inf \{ y_i : F_{Y \mid X} (y_i \mid x_i) \geq \tau \} \), where \( F_{Y \mid X} \) is the conditional distribution of \( Y \) given \( X \). The quantile regression model allows one to estimate the entire conditional distribution of \( y \) given \( x \). While more quantile regressions can potentially be more informative, estimation is restricted here to five quantiles: 0.1, 0.25, 0.50, 0.75 and 0.90.

If the distribution of the covariate vector \( x \) has finite support, with \( \Pr (x = x^j) = \phi_j \), for \( j = 1, \ldots, J \), then it is possible to obtain a minimum distance (MD) estimator for \( \beta_{\tau} \) (Chamberlain, 1991; Buchinsky, 1994).\(^6\)

\(^6\)We have simulated the Minimum Distance estimator of Chamberlain (1991) and the estimator
Here we use the standard model relating education, experience and earnings, based on works of Mincer (1974). The dependent variable is the logarithm of real hourly wages while the explanatory variables are a constant, education, experience and experience squared (see Mwabu and Schultz, 1996; Buchinsky, 1995, 1994; Rupert et al. 1992, among others, for similar specifications). Results of these estimations are presented in Figures 5 and Tables A.2 to A.3 in the appendix7.

Insert Figure 5

In Figure 5 we present the returns to schooling and experience at the ninetieth and tenth percentile. In first place, the estimates of education returns we obtain are relatively higher than those reported by Buchinsky (1994), while experience returns are more similar. This could help to explain why inequality has increased as a consequence of the shift of the upper tail of the distribution, i.e. it pays to be highly educated. Also, these returns present a similar U-shape pattern as those of Buchinski, i.e. returns are higher for workers in the lower or upper tail of the log rhw distribution than for those in the middle of the distribution. That is, returns of an additional year of education for least skilled workers or highly skilled workers are relatively higher than for the mean skilled type of worker, i.e. there are many ”mean” skilled workers. Finally, from Figure 5 we observe that experience-education returns remained oscillating around the same levels during the whole period.

CONCLUSIONS

In this paper we have discuss whether the increase in wage in the Uruguayan labour market is explained by the dramatic drop of the real minimum wage or by the high proposed by Koenker and D’Orey (1987), and the results were basically the same.

7The quantile regression estimation was programmed in GAUSS by the authors. Standard Deviations were estimated by bootstrap.
wage increase of the more skilled-educated workers. Using an empirical nonparametric quantile approach we conclude that this increase in wage dispersion can be basically attributed to the second of the reasons mentioned.
REFERENCES


