THE IMPACT OF MINIMUM WAGE CHANGES ON EMPLOYMENT AND WAGES IN BRAZIL:
Evidence from Time Series and Longitudinal Data

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ABSTRACT

The impact of minimum wages on the labour market has been a recurrent issue in the theoretical and empirical literature [e.g., Cubitt and Heap (1999), Dickens et al. (1998), Waltman et al. (1998), and Dolado et al. (1996)]. There are two clear and divergent views on the normative impact of minimum wages on labour markets in developing economies. The advocate view holds that minimum wages redistribute resources in a welfare enhancing way, and as such have the potential to reduce poverty, enhance productivity, and foster economic growth. The distortionist view suggests that minimum wage interventions misallocate labour, waste resources through rent-seeking, impair adjustment to economic shocks, deter investment and reduce growth rates with the effect of depressing wages where most of the poor are found – i.e., in the urban informal sector and in the rural sector [Freeman (1994)]. In this paper, we present estimates on the impact of the minimum wage on the Brazilian labour markets using time series data for the period between 1982 to 1999, and longitudinal information on households for the period 1995-99. The results suggest a robust and negative impact of the minimum wage on formal employment and a sizeable impact on the distribution of wages.

Keywords: Minimum Wages, Time Series Models, Longitudinal Data Analysis.

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1. Introduction

The debate about the effects of a minimum wage in the economy has for long been characterised by intense controversy. The textbook description of a competitive labour market predicts that the introduction of a minimum wage will price workers out jobs. This traditional view, however, has recently been challenged on the grounds that, in certain circumstances, the introduction of a minimum wage might generate an employment gain [Card and Krueger (1995a)] and positive impacts on the other wages in the economy [Machin and Manning (1994); Tiffin and Dawson (1996)]. On this regard, Freeman (1996) argues that minimum wages can have redistributive effects, depending on how the minimum wage interacts with the labour relations system. If higher-paid workers have some market power, any redistributive impact of minimum wages might be offset. On the other hand, in the absence of insider power, the minimum wage might decrease wage dispersion [Freeman (1994)].

A number of recent studies has attempted to assess the impact of the minimum wage on the labour market for different countries. For Britain, for example, Bell and Wright (1996), analysed the impact of Wage Boards and Councils and found that minimum wages did not push the wages of workers in covered jobs above those in non-covered jobs, suggesting thus that there are only small wage and employment effects. Machin and Manning (1994) concluded that the decline of the level of the minimum wage relative to the average wage significantly contributed to widening wage dispersion over the 1980s in Britain. Machin and Manning (1996), on the other hand, stressed the finding that in Britain the abolition of the Wage Boards and Councils resulted in wage falls in new jobs and no employment gains (p. 672).

There are also studies that associate minimum wage increases with business failure rates. Waltman et al. (1998) and Fischer (1997), for example, analysed whether minimum wage increases drive small businesses out of the market entirely. The argument behind this issue is that since business firms are presumably employing resources at maximum efficiency, rising labour costs must be absorbed elsewhere. As demand may be inelastic for some industries, and substitution of labour by capital may require expensive capital investment, many firms may actually exit the market. The general conclusion of studies of the US economy is that minimum wage hikes do not drive businesses into bankruptcy at a higher rate than would have happened had there been no increase in minimum wages.

In the case of developing countries, there are mixed results concerning the impact of minimum wages on the labour market. Bell (1997), for example, has found substantial disemployment effects of minimum wages for the case of Colombia and no wage or employment effects in the formal sector of the labour market in Mexico. For the case of Brazil, Carneiro and Faria (1997) and Carneiro and Henley (1998) have found that the minimum wage was an important determinant of other wages during the 1980s, but that this importance in affecting wage determination was gradually eroded throughout the 1990s. Lemos (1997) showed that the minimum wage affects positively other wages for a period of five quarters, following an initial chock. Soares (1998), on the other hand, concluded that the minimum wage had a reactive behaviour during the 1990s, while Lemos (2000) estimates that increases in the value of the minimum wage tend to compress the distribution of wages with moderate adverse effects on the level of employment.

Until recently, the most influential estimates of the impact of the minimum wage on the labour market came from time series studies. Nevertheless, there are criticisms on the use of time series data to estimate the wage and employment elasticities of the minimum wage on the grounds that (i) it is difficult to choose the appropriate set of control variables and (ii) that there are possible ambiguities about how to measure the minimum wage [Card and Krueger (1995a)]. Additionally, Card and Krueger (1995b) argued that most time series studies were contaminated with publication bias since as researchers extended the sample into the 1980s, the impact of the minimum wage on employment declined and became less significant. However, Neumark and Wascher (1998) rejected the hypothesis of publication bias and in a recent paper Williams and Mills (1998) argued that previous time-series studies have essentially ignored the many advances in modelling time series.

In this paper, we use a double strategy to assess the impact of the minimum wage on the Brazilian labour market. First, we use aggregate time-series data combined with modern time-series methods to assess the employment effects of minimum wages in Brazil. We rely on the arguments of Williams and Mills (1998) and Neumark and Wascher (1998) to explore time-series techniques to investigate the impacts of minimum wages. Second, we complement the time series analysis by using longitudinal data to investigate the impact of minimum wages on employment levels and on the distribution of wages. The case of Brazil is of interest as the government has made use of enforced minimum wages with a variety of objectives. In the past, the minimum was an effective tool to control wage bargaining and currently it is used as a parameter for social benefits such as unemployment insurance and social security benefits.

The paper is structured as follows. Section 2 presents the results of cointegration analysis of the impact of minimum wages on employment, including super exogeneity tests for the minimum wage. Section 3 presents the longitudinal analysis of the impact of the minimum on the distribution of wages and on employment levels. Section 4 concludes arguing that changes in the minimum wage have sizeable impacts on the distribution of wages and can actually impact employment levels in the formal and informal sectors of the Brazilian labour market. Following a minimum wage hike, the number of formal jobs is reduced whereas the size of the informal labour force increases. This overall result is in line with most theoretical studies on the impact of minimum wage on the labour market.

2 Empirical Analysis

As argued by Williams and Mills (1998), the previous time-series literature failed to adequately address the important issues of the stationarity of the variables and the dynamic specification of the employment equations, since most of the previous literature was based on static models. These caveats may have contributed to inconsistent results. Seminal work by Granger and Newbold (1974) and Phillips (1986), and research on cointegration, casts doubt on empirical evidence based on regression analysis of time-series data given any nonstationary variables. To avoid the problem of the spurious regression and the failure to account for the appropriate dynamic specification, we first performed unit roots tests on our variables. Following the unit root tests we then used cointegration analysis to investigate the existence of any lasting long-run equilibrium amongst the variables and the short-term dynamics of the employment adjustment process.

The stationarity of the variables was tested by means of standard Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) tests, both with seasonal dummies and with and without a trend. To determine the appropriate lag length of the tests, we used the Schwarz BIC model selection criterion as recommended by Stock (1994). The results, not reported due to space limitations, have not allowed the rejection of the null hypothesis of non-stationarity for the levels of the variables. When we considered their first differences, however, we were able to reject the null of non-stationarity, confirming the findings of previous studies based on alternative methodologies as in Cati et al. (1999) and Campelo and Cribari-Neto (2000).

Although cointegration analysis represents now a mandatory procedure to deal with time series data, several Monte Carlo studies about the power of cointegration tests show that in general

these tests are not powerful [Maddala and Kim (1998), p. 233]. Another problem is how to interpret the results of cointegration analysis. If there is only one cointegration relationship, then it may be easy to interpret it as a long-run equilibrium. However, if we find the number of cointegration vectors to be greater than one, this creates problems of interpretation, and more economic theory is necessary. In addition, there is the issue of whether one should use single-equation or system estimation methods. We are conscious of these caveats and thus we follow the advice of Maddala and Kim (1998) that it is important for empirical researchers to report the results from several tests of cointegration.

2.1 The Long-Run Impact of Minimum Wages on Employment

As in most empirical studies which have attempted to assess the employment effects of minimum wage changes [e.g., Machin and Manning (1994), Bell (1997)], we have opted for estimating a simple employment equation of the following general form:

$E = f(m, Y, t)$

where E is employment, m is the minimum wage, Y is aggregate output and t represents a time trend. According to the traditional textbook view on the effects of minimum wages on the labour market, the coefficient for the variable m should attract the negative sign. On the other hand, Y, representing the effects of GDP growth on employment, should have a positive coefficient, at least for the formal sector of the economy. The time series on employment and aggregate output are monthly indices published by *the Brazilian Institute of Geography and Economics* (IBGE). The official minimum wage rate is available from the Ministry of Labour. For estimation purposes, all variables were considered in logarithmic form.

The Johansen (1988) procedure was initially applied and the results of this are reported in Tables 1 and 2. The results are obtained using lags up to the 8th order in the vector autoregression[1]. Both the eigenvalue statistic and the trace test strongly reject the null of no cointegration in favour of a single cointegrating relationship. The Tables also report the standardized eigenvectors (β') and adjustment coefficients (α). The adjustment coefficients (α) measure the feedback effect of the (lagged) disequilibrium in the cointegrating relation onto the variables in the VAR. The positive adjustment coefficient for formal sector employment (0.007) implies that lagged changes in formal sector employment induce further changes in the same direction in current employment in this sector. For the case of informal sector employment, the adjustment coefficient was negative (-0.139) indicating that lagged employment changes induce further changes in current employment in the opposite direction in the informal sector. The low absolute values imply slow adjustment to remaining disequilibrium, which is compatible with high frequency aggregate data.

[1] In deciding the appropriate lag length of the vector-autoregressive (VAR) system for cointegrating purposes we first ran an unrestricted reduced form (URF) for our six variables with a maximum lag length of 12. F-tests for the significance of retained regressors (i.e., the contribution of each lag on each variable to the six-equation VAR system taken together) indicated that no lags above the eighth order are significant.

As there was evidence of a single cointegrating vector, we decided to use the fully parametric least square estimator (FPLS), as proposed by Inder (1995)[2], which is an efficient single-equation method [Maddala and Kim (1998)]. This is a version of the fully-modified ordinary least square estimator (FM-OLS) and is similar in spirit to Phillips-Perron (1989) unit root tests in the sense that it starts with the OLS estimator and applies corrections to it to take care of the endogeneity and serial correlation problems[3]. For the sake of comparison, we also ran an autoregressive distribute lag model with the same lag length as in the Johansen analysis. The static long-run solution to the ADL model is reported in Tables 3 and 4 along with the results of the Johansen (1988) and Inder (1995) cointegration tests.

The results revealed a robust and negative impact of changes in the minimum wage in the employment levels of formal workers in the long run, with the reverse taking place for the case of informal workers. That is, increases in the value of the official minimum wage tend to decrease formal employment and increase informal employment. The elasticity of employment with respect to changes in the minimum wage ranged from – 0.001 to – 0.024 for formal sector workers and from 0.0004 to 0.003 for informal workers. Although the elasticities were very low in absolute terms they suggest important long-term trends in the employment adjustment process for both sectors. In the formal sector, changes in minimum wages tend to affect negatively employment as largely suggested by the empirical and theoretical literature. On the other hand, changes in the value of the minimum wage tend to affect positively employment in the informal sector since this is the sector where the workers who lose their jobs in the formal sector tend to find new and temporary occupations.

Another interesting result is concerned with the way in which employment levels in both sectors behave over the business cycle. Formal sector employment tends to react pro-cyclically to changes in economic activity whereas informal employment reacts anti-cyclically to output fluctuations. The intuition behind this process is straightforward since economic growth tends to create more jobs, encourage formal employment and, therefore, discourage informal employment.

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[2] This method starts with a general VAR model of I(1) variables and derives the implied single equation $y_{1t} = b'y_{2t} + A(L)Dy_{1t} + B(L)Dy_{2t} + v_t$, where A(L) and B(L) are polynomials in the lag operator L. Then, after suitable truncation of the lag polynomials A(L) and B(L), one must regress y_{1t} on y_{2t}, leads and lags of Δy_{2t}, and lags of Δy_{1t}. The leads and lags of Δy_{2t} eliminate any effect of endogeneity and the lags of Δy_{1t} capture remaining autocorrelation in the stationary component of the regression.

[3] The traditional two-step cointegration procedure proposed by Engle and Granger (1987) is similar to Augmented-Dickey Fuller (ADF) tests in the sense that it modifies the estimating equations.

2.2 Short-Run Employment Dynamics

In order to obtain a valid dynamic specification we have included information from the long-run co-integrating relationship between the variables in question (see Tables 1 to 4), in addition to differenced (and therefore stationary) terms in the other explanatory variables. The subscripts CC and SC refer to formal (workers with signed labour cards) and informal (those without signed labour cards) sectors, respectively. The term τ is the error-correction term obtained from the Johansen models reported in Tables 1 and 2, σ is the estimated standard deviation of residuals and *S Seasonals* are monthly seasonal dummy variables. We have also experimented error-correction terms formed by the residuals from the solved static-long run solution to the ADL models, which yielded very similar results and were not reported. The equations satisfy all test statistics at the 5% level of significance and present coefficient estimates that accord with theoretical predictions. The preferred models for formal and informal sector employment were the following:

$$\begin{aligned} DE_{CC\,t} = {} & 0.0009 + 0.1731\, DE_{CC\,t-11} - 0.01059\, Dm_t \\ & (0.0013) \quad (0.0544) \quad (0.0058) \\ & - 0.01202\, Dm_{t-3} + 0.08629\, DY_{t-5} - 0.09127\, t_{CC\,t-1} \\ & (0.0056) \quad (0.0363) \quad (0.0740) \\ & + S\ seasonals \end{aligned} \qquad (1)$$

$R^2 = 0.5814$ F(20,156) = 10.834 $\sigma = 0.0127$ DW = 2.10
AR 1- 7 F(7,149) = 2.599 ARCH 7 F(7,142) = 0.38582 NORM $\chi^2(2)$ = 2.499
HET F(25,130) = 0.70859 RESET F(1,155) = 1.3654
N = 1982(1) to 1999(11)

$$\begin{aligned} DE_{SC\,t} = {} & -0.0233 - 0.2461\, DE_{SC\,t-1} + 0.0176\, Dm_t + 0.0262\, Dm_{t-6} \\ & (0.0081) \quad (0.0694) \quad (0.0122) \quad (0.0121) \\ & - 0.2069\, DY_{t-5} - 0.0575\, t_{SC\,t-1} + S\ seasonals \\ & (0.0770) \quad (0.0214) \end{aligned} \qquad (2)$$

$R^2 = 0.4000$ F(18,158) = 5.8472 $\sigma = 0.0267$ DW = 2.01
AR 1- 5 F(5,153) = 1.7689 ARCH 5 F(5,148) = 1.3946 NORM $\chi^2(2)$ = 0.17743
HET F(23,134) = 0.6092 RESET F(1,157) = 0.3120
N = 1982(1) to 1999(11)

Overall, the pattern observed for the long run remained the same in the short run. In the cases of both formal and informal sectors, the equations satisfy all test statistics at the 5% level of significance and present the expected coefficient estimates. The dynamics found was very simple.

Lagged changes in aggregate output *(D Y)* have a positive impact on current changes in employment in the formal sector and a negative impact on current changes in employment in the informal sector. The impact of changes in the minimum wage on employment is also different for each sector. Current changes in the minimum wage impact positively current changes in informal employment with the opposite taking place for the case of formal employment. After some months, however, there is a change in the direction of the impact of minimum wage changes over employment for both sectors, what is indicative that some form of catchup effect might take place following minimum wage increases. The error-correction term *(t)* attracted the expected negative coefficient, but its low absolute value indicates that adjustment to deviations from long-run equilibrium is slow.[4]

Weak exogeneity of the current-dated regressors in equations (1) and (2) above is required for their analysis as single equations to be efficient [Engle et al. (1983)]. Weak exogeneity can be tested as an implication of super exogeneity, requiring constant parameters in the conditional model. Super exogeneity requires both weak exogeneity and structural invariance, so finding super exogeneity implies weak exogeneity. Demonstrating super exogeneity relies on showing that the parameters of the conditional model remain constant even though the marginal process changes[5]. If the marginal process of current dated variables changes while the conditional model remains constant, then super exogeneity holds.

The minimum wage enters the employment equations contemporaneously and the nonconstancy of its marginal process is of interest. An appealing feature of testing for super exogeneity is that only a simple marginal model needs to be nonconstant. Thus, as in Ahumada (1992), Ericsson and Irons (1995) and Carneiro and Henley (2000), we use a univariate autoregressive (AR) model for *Dm* to evaluate its constancy over time.

The following model is obtained by simplifying an AR(13) model:

$$\begin{array}{lccc} Dm_t = & 0.02039 & + 0.166\ Dm_{t-1} & + 0.103\ Dm_{t-2} \\ & (0.01452) & (0.0621) & (0.06323) \end{array}$$

[4] A finding that is consistent with the theoretical and empirical literature on labour demand, as surveyed by Hamermesh (1993).

[5] See Charemza and Deadman (1997) for an illustrative and didactic explanation on how to test for exogeneity.

$$- 0.05746\ Dm_{t-3} \underset{(0,06142)}{} + \underset{(0,06034)}{0.6119\ Dm_{t-4}} + S\ Seasonals \qquad (3)$$

$R^2 = 0.559914$ F(15,150) = 12.723 [0.0000] $\sigma = 0.127811$ DW = 1.91
AR 1- 7 F(7,143) = 0.82024 ARCH 7 F(7,136) = 1.6301 NORM $\chi^2(2) = 29.97$
HET F(69, 80) = 1.6744 RESET F(1,149) = 8.3488 N = 1982(1) to 1999(11)

Figure 1 shows both the one-step residuals and the sequence of 1-step ahead Chow statistics for equation (3). Constancy of the marginal process is rejected. Engle and Hendry (1993) propose how to use determinants of non-constancies in the marginal model to test super exogeneity. One way of accomplishing this is by including binary dummies at each structural break in the marginal model (3). The expanded model now becomes:

$$\begin{aligned} Dm_t = {} & \underset{(0.01291)}{0.00938} + \underset{(0.05533)}{0.1977\ Dm_{t-1}} + \underset{(0.05608)}{0.1469\ Dm_{t-2}} \\ & - \underset{(0.05421)}{0.02887\ Dm_{t-3}} + \underset{(0.05364)}{0.5815\ Dm_{t-4}} - \underset{(0.1171)}{0.2651\ D89(3)} \qquad (4) \\ & - \underset{(0.1203)}{0.3468\ D90(4)} + \underset{(0.1171)}{0.3448\ D91(8)} + \underset{(0.1177)}{0.6093\ D92(5)} + S\ Seasonals \end{aligned}$$

$R^2 = 0.668841$ F(19,146) = 15.52 $\sigma = 0.11238$ DW = 1.84
AR 1- 7 F(7,139) = 1.1141 ARCH 7 F(7,132) = 1.3027 NORM $\chi^2(2)$ = 15.71
HET F(23,122) = 1.92 RESET F(1,145) = 21.781
N = 1982 (1) to 1999 (11)

The dummy variables are for the observations indicated by their names and appear highly significant for the marginal model of the minimum wage. They coincide with two of the major attempts at economic intervention in Brazil, namely the Summer Plan of 1989 and the Collor Plan of 1990. According to Engle and Hendry (1993), the determinants of the non-constancies should be insignificant if added to the conditional models (1) and (2). Thus, we have also re-estimated equations (1) and (2) including the binary dummy variables which coincide with each structural break:

$$\begin{aligned} DE_{CC\,t} = {} & \underset{(0.0013)}{0.0012} + \underset{(0.0536)}{0.1549\ DE_{CC\,t-11}} - \underset{(0.0059)}{0.00847\ Dm_t} \\ & - \underset{(0.0055)}{0.00795\ Dm_{t-3}} + \underset{(0.0229)}{0.07661\ DY_{t-5}} - \underset{(0.0036)}{0.00539\ t(E)_{CC\,t-1}} \qquad (5) \\ & - \underset{(0.01303)}{0.0018\ D89(3)} - \underset{(0.01301)}{0.01199\ D90(4)} - \underset{(0.01323)}{0.0068\ D91(8)} \\ & + \underset{(0.01390)}{0.02094\ D92(5)} + S\ seasonals \end{aligned}$$

$R^2 = 0.5488$ F(13,163) = 15.253 $\sigma = 0.0128$ DW = 2.24
AR 1- 7 F(7,156) = 2.283 ARCH 7 F(7,149) = 0.4953 NORM $\chi^2(2) = 0.9619$
HET F(28,144) = 0.6588 RESET F(1,162) = 0.2258
N = 1982(1) to 1999(11)

$$DE_{SC\,t} = \underset{(0.0087)}{-0.01976} \underset{(0.0727)}{-0.25836\, DE_{SC\,t-1}} \underset{(0.0141)}{+0.0226\, Dm_t} \underset{(0.0132)}{+0.0286\, Dm_{t-6}} \underset{(0.08256)}{-0.2055\, DY_{t-5}} \underset{(0.0229)}{-0.0526\, t(E)_{SC\,t-1}} \underset{(0.0289)}{-0.02133\, D89(3)} \underset{(0.0289)}{+0.0062\, D90(4)} \underset{(0.0292)}{+0.0013\, D91(8)} \underset{(0.0307)}{+0.0318\, D92(5)} + S\ seasonals \quad (6)$$

$R^2 = 0.4102$ F(18,158) = 4.5539 $\sigma = 0.0274$ DW = 2.02
AR 1- 7 F(7,137) = 1.5417 ARCH 7 F(7,130) = 0.5108 NORM $\chi^2(2) =$ 0.0728
HET F(27,116) = 0.6074 RESET F(1,143) = 0.2439
N = 1982(1) to 1999(11)

The dummy variables are insignificant in models (5) and (6), confirming the super exogeneity of the minimum wage in their conditional models, both for formal and informal workers. As a result, equation (4) cannot be used to determine the minimum wage even when equations (1) and (2) are identified as constant employment equations.

As illustrated by Ericsson and Irons (1995), one important implication of super exogeneity is that error-correction models such as those in equations (1) and (2) can have a forward-looking interpretation. In practical terms this means that policy can affect agent behaviour. For example, government policy might very well affect the minimum wage and so employment levels of formal and informal workers. However, under super exogeneity, the precise mechanism that the government adopts for such a policy does not affect agent behaviour, except when the mechanism affects actual outcomes (op. cit., p. 35). Skepticism concerning the effectiveness of such policies, founded on the Lucas critique, is therefore unwarranted in the present context.

3. Longitudinal Analysis

3.1 The Impact on Wages

To investigate further the impact of minimum wages for the case of Brazil, we wiil use now longitudinal data from the Monthly Employment Survey of IBGE. The survey covers the main six

metropolitan regions of the country (Porto Alegre, São Paulo, Rio de Janeiro, Belo Horizonte, Salvador and Recife). The period of analysis is 1995-99, what allows one to assess the impacts of the minimum wage policy without the influence of the high inflation rates observed prior to 1994 in Brazil. The longitudinal character of the survey makes it possible to use information about a household for a period of four consecutive months. We begin the analysis first by analyzing the impact of the minimum wage on the wage distribution, and then we follow on to investigate the employment effects of the minimum wage.

To identify the impacts of the minimum wage on the other wages we monitor the value of the earnings of those workers employed during the months close to the date of the official change in the value of the minimum wage. During the period of analysis, the minimum wage changed once a year in May of each year. As in the moment of the survey (period t) the workers tend to report their earnings referring to last month's activities (period $t-1$), we will use information regarding the month of May as reported by workers in the month of June to capture the impact of the change of the minimum wage on the other wages.

Barros et al. (2000) developed a methodology based on longitudinal data to investigate the impacts of changes in the minimum wage on poverty levels. Their methodology can easily be used to investigate also the impact of the minimum on employment and wages. The authors suggest a series of steps to compare poverty levels in April of each year (one month prior to the increase in the minimum wage) with those simulated for May (the official date of change in the minimum wage). Their strategy can actually be seen as sequence of counter-factuals that allows one to measure what would be the level of poverty in May in case one considered the impact of the minimum wage only on the earnings of certain groups of workers. Thus, using this methodology we can easily assess how changes in the minimum wage impact the distribution of wages of the workers employed in the months of April and May. For this purpose we focused on the workers who earned a fixed proportion of the minimum wage in April and May.

In what follows, we compare the distribution of wages observed in April and May for the workers who composed the sample in April, May and June for each of the years of the period 1995-99. Figures 2 to 10 show these distributions, as well as a counter factual distribution. The Figures reveal 4 important features as regards the impacts of the minimum wage:

- First, the minimum wage seems to strongly impact the distribution of wages. Actually, there is an important concentration of workers earning the equivalent to one minimum wage and smaller concentrations around earnings equivalent to 2 minimum wages.

- Second, the figures show that there is a change in the distribution of wages between April and May, centered on earnings below the value of 2 minimum wages. That is, the spillover effect of the minimum wage for wages beyond the equivalent to 2 minimum wages was quite limited in the period we considered.

- Third, by comparing the wage distribution observed in May with the counter-factual distribution we observe that both of them follow a very close path. This is suggestive that changes in the minimum wage were actually responsible for changes in the wages of those workers who had earnings close to the value of the minimum wage in the months of April and May.

- Fourth, the only points in which the wage distributions diverge correspond to workers with no earnings or those with earnings in the region of the value of the minimum wage in April and May (for example, R$70 and R$100, in 1995). We can notice that the number of workers with no earnings is much higher in May than in the counter-factual distribution. On the other hand, the relationship between the two distributions is inverted in the region of the minimum wage in the months of April and May. This actually suggests that these were the workers who lost their jobs and therefore their remuneration. The Figures suggest that a lower share of these workers seems to have come from the region of two minimum wages.

3.2 The Impact on Employment Levels

To estimate the impact of the minimum wage on the level of employment we will use the method of differences-in-differences. This method allows for the investigation of the effect of an event on the level of a given variable by comparing the evolution of this variable for two distinct groups of a population [Angrist and Krueger (1999)]. These groups are called of treatment group and control group.

In implementing this methodology, first only the treatment group must be affected by the event to be analyzed. Then, both groups should be equally affected by all of the other determinants

of the variable of interest. In our case, we have to find a group of workers (treatment group) whose level of employment is affected by the minimum wage, and another group of workers (control group) whose level of employment is not affected by the minimum wage but that might be affected by other factors in the same way as the first group. The difference in the variation in the level of employment of the treatment group in relation to the variation in the level of employment of the control group can be seen as an estimate of the variation in the level of employment of the treatment group in case this group had been affected only by the increase in the minimum wage.

That is, if we call Y the percentage change in the level of employment for the treatment group, we can express it as:

$$Y = (DE^t/E^t) - (DE^c/E^c) \qquad (7)$$

where E^t and E^c denote the employment levels of the treatment and control groups, respectively, and Δ denotes the change in each variable before (April) and after (May) the increase in the minimum wage.

This methodology also allows for the analysis of the impact of changes in the value of the minimum wage on the degree of informality by simply adapting equation (7) to take account of the workers who have become employed in the informal sector (without a signed labor card) between April and May.

To identify the treatment and control groups we use the results of the wage distributions of the previous section. Thus, we will use as the treatment group the workers who earned the equivalent to one minimum wage between April and May – as affirmed earlier, there is evidence that the level of employment of these workers is affected by changes in the value of the minimum wage.

However, other factors could also have affected the level of employment of these workers (i.e., the level of economic activity). Therefore, it is necessary to identify a control group that should be formed by individuals with similar characteristics as those in the treatment group. The control group we chose will be formed by individuals who earned in April the equivalent to a value between one minimum wage in May and two times the minimum wage of April (for example, workers with salaries ranging between R$100 and R$140, in 1995). We have decided to exclude

from the sample the workers with salaries between 1.5 and 2 times the value of the minimum wage in April.

To implement equation (7), it is necessary to calculate the number of workers who have lost their jobs in both groups. This is made possible by the longitudinal character of the PME. The workers who received no income in May have been considered as those who have lost their jobs. This procedure is adopted for both workers with and without signed labor cards.

The results for the estimator in differences (Y) appear in the fifth and sixth columns of Table 5, as a proportion of the stock of individuals with positive income in the first month. Two important results arise. First, the effect of the minimum wage on the level of employment is much larger in the formal sector as compared to the informal sector. Furthermore, we can notice a declining trend of values observed for the formal sector mainly for the years of 1998 and 1999. The first column of the Table shows that this trend is not so evident for the treatment group. That is, when we take into consideration all of the factors that affect the level of employment in the treatment group, the drop in employment is always close to 10%. However, the share of the drop relative to all of the other factors but the minimum wage, reported in the third column, presents a growing trend with a strong acceleration between 1998 and 1999. Therefore, the adverse effect of the minimum wage seem to have been attenuated, especially after 1998.

The second important result is that the evolution of these trends may be influenced by the pattern and size of adjustment of the minimum wage in the period considered. Actually, the size of the adjustment allowed for the minimum wage has declined between 1995 and 1999 (in line with lower inflation rates in the period). Thus, we report in the seventh and eightieth columns of Table 1 the results in the form of elasticities. Note that the magnitude of the elasticities oscillates throughout the period, especially in the years 1995 and 1997. The first year presents the lower sensibility of changes in employment levels following an increase in the value of the minimum wage while the second year presents a higher sensibility. While an increase of 10 percent in the value of the minimum wage reduced the level of employment in the treatment group by less of 3 percent in 1995, the drop in employment levels in 1997 was of 13 percent.

The same experiment was made to assess the impact of changes in the degree of informality in the labor market following an increase in the minimum wage. Now, the treatment group is formed by the workers employed with signed labor card who received a salary equivalent to the

value of the minimum wage between April and May. Those workers who received a salary equivalent to the value of the minimum wage in May and twice the minimum wage in effect in April formed the control group. In both cases we accounted for the number of people who ceased to have a signed-labor-card job to become informal (without signed labor card).

Table 6 presents the results for the impact of the minimum wage on informality in the same fashion as Table 5. The results for the transition of workers with signed labor card to the informal sector show an upward trend that came to a halt in 1999. This evolution was mostly determined by the treatment group, as the values for the control group were very low, with the exception of those observed in 1999.

The results in the form of elasticities magnify the upward trend for the whole period, including 1999. Our estimates suggest that an increase of 10% in the value of the minimum wage would be associated with a transition of 2.2% of workers from the treatment group to an informal job in 1995, and roughly 15% in 1999. That is, within the period we have considered, increases in the value of the minimum wage have contributed to the informalization of the labor force in Brazil.

4. Conclusions

This paper has investigated the impact of minimum wages on the Brazilian labour market. The approach adopted in the paper complemented a time-series aggregate analysis with evidence generated by the use of longitudinal data. Initially, in line with recent developments concerning time-series econometric analysis, we have investigated the long-run structure as well as the short-term dynamics of employment equations in which the minimum wage enters as an explanatory variable. Then, we used a different methodology and explored the longitudinal character of the data to see whether the trends found with the aggregate data could be confirmed at a more disaggregate level if we used longitudinal data.

The strategy yielded convergent results that confirmed previous findings in the literature. Both the time-series and the longitudinal analyses revealed a robust and negative impact of changes in the minimum wage in the employment levels of formal sector workers and a sizeable impact on the distribution of wages. Our results suggest that increases in the value of the minimum wage have been responsible for the growth of informal employment in the labour market. That is, due to labour mobility, increases in the value of the minimum wage provoke disemployment effects in the formal

sector, which are then matched by increases in informal occupations. Another interesting result is that formal sector employment reacts pro-cyclically to changes in economic activity whereas informal employment reacts anti-cyclically to output fluctuations.

The paper also discusses the exogeneity of the minimum wage in the employment equation. The conditional model for employment is found to be structurally stable, but combined with an empirically nonconstant univariate marginal model for the minimum wage this implies the super-exogeneity of this latter variable in the conditional model. Since super-exogeneity is not invariant to renormalisation, a structurally stable employment equation inverted to model the minimum wage may be non-constant. This is indeed the case in Brazil. Therefore, the constant relationship found for employment cannot be used to derive models for the minimum wage. From this, we reach the conclusion that government policy directed towards changing the value of the minimum wage does have the potential to impact employment in Brazil.

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Figure 1: One-Step Residuals and Recursive Chow Statistics for Equation (9)

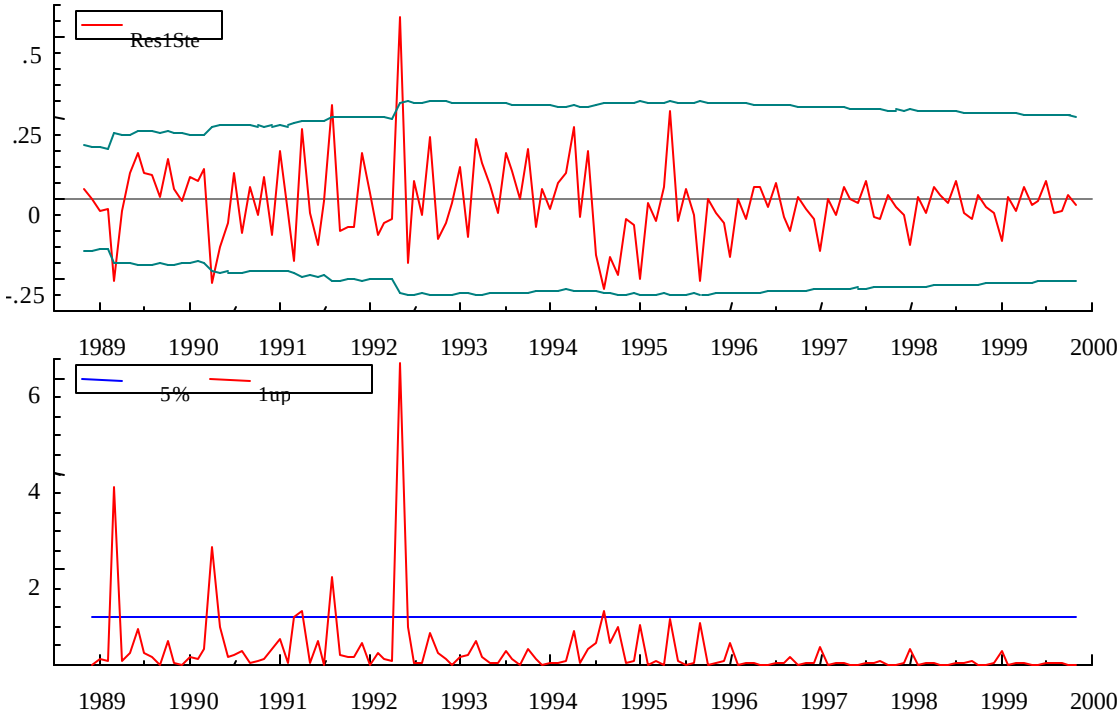

Table 1: Cointegration Analysis of Formal Sector Employment – 1986(1) – 1999(11)

Eigenvalues	0.1513	0.0737	0.0123
Hypotheses	r = 0	r ≤ 1	r ≤ 2
λMax	30.02	14.00	2.26
95% critical value	25.50	19.00	12.30
λtrace	46.29	16.27	2.26
95% critical value	42.40	25.30	12.30
Standardized Eigenvectors (β')			
Ecc	Y	mw	Trend
1.000	-3.938	0.006	0.008
-0.906	1.000	-0.019	-0.001
649.130	-251.140	1.000	0.581
Standardized Adjustment Coefficients (α')			
0.007	0.189	-0.029	NA

Note: The vector autoregression includes eight lags on each variable, a constant term and monthly dummies. The $?_{max}$ and $?_{trace}$ are Johansen's maximal eigenvalue and trace statistics.

Table 2: Cointegration Analysis of Informal Sector Employment – 1986(1) – 1999(11)

Eigenvalues	0.234	0.078	0.028
Hypotheses	r = 0	r ≤ 1	r ≤ 2
λMax	45.55	14.98	5.25
95% critical value	25.50	19.00	12.30
λtrace	64.46	18.90	4.91
95% critical value	42.40	25.30	12.30
Standardized Eigenvectors (β')			
Esc	Y	mw	Trend
1.000	1.320	-0.003	-0.004
-0.566	1.000	-0.003	-0.000
-96.44	-914.340	1.000	3.766
Standardized Adjustment Coefficients (α')			
-0.139	-0.108	-0.750	NA

Note: The vector autoregression includes eight lags on each variable, a constant term and monthly dummies. The $?_{max}$ and $?_{trace}$ are Johansen's maximal eigenvalue and trace statistics.

Table 3: Estimated Long-Run Relationships for Formal Sector Employment – 1986(1) – 1999 (11)

Dependent Variable Ecc	FPLS	ADL Solved Long Run Solution	Johansen
mw	-0.001	-0.024	-0.006
Y	0.044	2.220	3.938
Trend	N.A.	-0.003	-0.008

Notes: FPLS estimations were performed with 2 leads and 2 lags for each variable, selected by the method proposed by Inder (1995).

Table 4: Estimated Long-Run Relationships for Informal Sector Employment – 1986(1) – 1999 (11)

Dependent Variable Wsc	FPLS	ADL Solved Long Run Solution	Johansen
mw	0.002	0.0004	0.003

Y	-0.027	-0.3670	-1.320
Trend	N.A.	0.0035	0.004

Notes: FPLS estimations were performed with 2 leads and 2 lags for each variable, selected by the method proposed by Inder (1995).

Tabela 5: The Impact of Minimum Wage Changes on Employment Levels

	Treatment Group			Control Group			Estimator Y			Elasticity	
	CC	SC		CC	SC		CC	SC		CC	SC
1995	0.113	0.194		0.001	0.159		0.111	0.035		0.260	0.081
1996	0.103	0.190		0.007	0.154		0.096	0.036		0.803	0.303
1997	0.113	0.185		0.018	0.172		0.095	0.013		1.328	0.180
1998	0.091	0.190		0.011	0.162		0.080	0.028		0.966	0.338
1999	0.088	0.167		0.044	0.167		0.044	0.000		0.945	0.000

Source: Based on information from the Montly Employment Survey.

Table 6: The Impact of Minimum Wage Changes on Informality

	Treatment Group			Control Group			Estimator Y			Elasticity	
	SC	Self-Employed		SC	Self-Employed		SC	Self-Employed		SC	Self-Employed
1995	0.094	0.023		0.000	0.022		0.093	0.001		0.217	0.003
1996	0.098	0.013		0.006	0.007		0.093	0.006		0.772	0.051
1997	0.100	0.015		0.014	0.019		0.085	-0.005		1.197	-0.063
1998	0.125	0.015		0.007	0.017		0.118	-0.002		1.417	-0.026
1999	0.108	0.002		0.040	0.018		0.068	-0.016		1.474	-0.356

Source: Based on information from the Montly Employment Survey.

Legend: CC = Workers with Signed Labour Cards; SC = Workers without Signed Labour Cards.