Real Wages and the Lucas Critique: Can the Government Tax Policy Influence Wage Growth in Brazil?*

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Abstract

The paper uses cointegration analysis to investigate the long run behaviour and short term dynamics of aggregate real wages in Brazil over the period 1986 to 1993. First, we show that real wages, inflation, productivity, unemployment and the tax wedge are cointegrated. Second, the dynamics of this relationship are specified to obtain a satisfactory representation of the real wages, which is structurally stable over the period despite major policy changes. Thirdly, evidence is presented for the super-exogeneity of the tax wedge, an important result which shows that the path of real wages can be influenced by government policy without inducing Lucas-critique problems.

Resumo

O artigo usa técnicas de cointegração para investigar o comportamento de longo prazo e a dinâmica de curto prazo dos salários reais no Brasil entre 1986 e 1993. Em primeiro lugar, mostramos que salários reais, inflação, produtividade, desemprego e uma proxy para o custo do trabalho (o tax wedge) são cointegrados. Em segundo lugar, a dinâmica desse relacionamento é especificada de forma a obter uma representação satisfatória dos salários reais que é estruturalmente estável ao longo do período analisado, apesar de se observarem sucessivas intervenções na política oficial de salários. Finalmente, apresentamos evidência sobre a super-exogeneidade da proxy para custos do trabalho (tax wedge), um resultado

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importante que mostra que a trajetória dos salários reais pode ser influenciada pela política governamental sem que se incorra em problemas relacionados com a Crítica de Lucas.

Key Words: Cointegration analysis, Exogeneity, Wage determination .

JEL Code: C22, C52, J30 .

1. Introduction.

The study of wage determination has become a major issue in the recent literature. Most of the empirical work which has now appeared is concerned with the degree of real wage flexibility and the occurrence or not of hysteresis in the labour market. It has been claimed that countries with centralised or decentralised wage bargaining structures tend to have flexible labour markets while countries with an intermediate-centralised bargaining structure are bound to suffer from wage rigidity and hysteresis (Calmfors and Driffill, 1988). There has been considerable work done on this for the industrialised countries, but very little written about the characteristics of collective bargaining in developing countries. In the former, a recurring finding points to the existence of insider effects not only in highly unionised countries but also in non-unionised economies (see Lever, 1995, for an empirical survey).

In this paper, we present the results of a study of Brazilian earnings behaviour over the period 1986 to 1993. Our sample finishes in 1993 as the Real Plan, implemented in 1994, abolished the formal wage policy and changed the pattern of collective bargaining in the country. We estimate our earnings equations using cointegration analysis, since it allows for the separation of issues about the long run behaviour of real wages from their short term dynamics. Also we test for the stability of the estimated wage equation (e. g. , Ericsson

and Irons, 1995) in the context of a developing economy facing persistent high inflation. This allows us to draw important conclusions concerning the effects of government policy towards the structure of labour taxation on the path of real wages.

The structure of the paper is the following. Section 2 presents the theoretical framework in which we model the Brazilian wage bargaining process. Section 3 discusses the integration properties of the data and in section 4 we analyse long run determinants of aggregate real wages. In section 5 we model short run wage dynamics while section 6 examines exogeneity issues and discusses policy implications. Section 7 summarises the main findings of the paper.

2. Theoretical Framework.

We assume that the determination of the wage rate can be described by the solution to an asymmetric Nash-bargain¹ :

$$W_i = \arg \max (.) = [V_i(.)]^\mu [\Pi_i(.)]^{1-\mu} \quad (1)$$

The parameter μ in equation (1) captures the relative power of the two sides in the wage bargain and is usually proxied by a measure of union power, such as density. As μ tends to 1, we approach the monopoly union solution in which the union sets the wage unilaterally while the firm sets employment at its profit-maximising point given this wage; i. e. , the union chooses the wage that will maximise its welfare function subject to the constraint of a downward-sloping labour demand schedule. In doing so, it trades off the gains from raising real wages against the employment loss that will follow. If $\mu = 0$, the firm can choose a wage level and associated employment level to maximise its profits only constrained by having to pay workers the fall-back wage level, and for $0 < \mu < 1$, any quasi-rents

¹Further details can be found in Carneiro (1999)

originating from a monopolistic product market of monopsony in other factor markets are shared between the union and the firm (e. g., Malcomson, 1987).

The wage equation, therefore, is obtained from the first order condition for maximisation of equation (1). In the case of the right-to-manage model or the monopoly union model, this solution is additionally constrained by the marginal productivity condition. A formal solution requires the full functional specification of the production technology underlying the profit relationship $\Pi_i(\cdot)$ and the union utility function, $V_i(\cdot)$. It will generate an aggregate wage equation of the following implicit form:

$$W_i = W_i(P^e, N_i, M_i, H, \theta, \mu, X_i) \quad (2)$$

The appropriate expected prices will be captured through a consumer price index. The insider power and bargaining power parameters, θ and μ , are usually allowed to vary and are proxied through variables such as union density and/or strike intensity. Unfortunately such variables are either unavailable or very unreliable in the Brazilian case². We must assume in the present case that any movement in μ over time is proxied by movement in the other variables used. Insider power is also typically captured through the correlation with the wage of variables which capture the productive performance of the employer, such as labour productivity or corporate profitability (e. g. , Carruth and Oswald, 1989; Nickell and Wadhwani, 1990).

A further problem arises in the parameterisation of equation (2) because of the simultaneous determination of wages and employment in the model (Manning, 1993). Typically, therefore, explanatory variables which include information about the employment level,

² Union membership is available on an annual basis, but is self-reported by the individual union confederations. The data are therefore often mutually contradictory.

such as the unemployment rate or labour productivity, are either instrumented or lagged. In empirical practice, reliable interpretation of the parameters of the wage equation rests on the identification of a single long-run relationship between wages and, in our case, the unemployment rate and labour productivity and its interpretation as a reduced form wage equation³.

3. Data Description and Unit Roots Tests.

The wage data for Brazil used in the present paper is an index of average monthly nominal industrial wages for the state of São Paulo⁴, available from 1985 until 1993. Given the difficulty in obtaining, and/or the lack of, variables measuring union power/union density (with monthly frequency) and the meaninglessness of a benefit replacement ratio for Brazil, we are left with a more limited choice on the selection of our explanatory variables than is usually available for such analysis on developed economies. Therefore, we use consumer prices, productivity, unemployment and the tax wedge as explanatory variables. Precise definitions and sources of the explanatory variables are given in the Appendix.

Table 1 reports summary statistics for the data in levels, logs and differenced logs. Nominal wage and price series are indices based in 1985. Their means are so high because of the effect of accumulated rapid inflation over nine years. Average monthly wage inflation is 17.7 per cent; average monthly consumer price inflation very slightly lower at 17.5 per cent. The tax wedge variable H is defined as the ratio between real employers' labour cost (deflated by the industrial price index) and the post-tax consumption wage. On average over

³Therefore caution should be exercised in the causal interpretation of empirical estimates of such a model.

⁴In 1985, São Paulo accounted for nearly half of the industrial output of Brazil.

the sample period employers labour cost is 179% of take-home pay and this has been a constant source of concern amongst employers and employees alike. Specifically, such high non-wage labour costs create the possibility that labour turnover costs are high, which in turn will allow the exploitation of insider bargaining power [Lindbeck and Snower (1989)]. Average productivity growth is 0.5% per month; the unemployment rate, which averages 5.3%, shows a slight upward trend over the sample period.

Table 1: Descriptive Statistics

<i>levels</i>	W	P^c	Q/L	U	H
Mean (Std.Dev.)	$4.82.10^8$ ($1.80.10^9$)	$4.46.10^8$ ($1.64.10^9$)	1.138 (0.26)	5.348 (1.50)	1.788 (0.26)
<i>logs</i>	w	p^c	$q - l$	u	h
Mean (Std.Dev.)	12.33 (5.68)	12.30 (5.73)	0.109 (0.20)	1.636 (0.29)	0.577 (0.09)
<i>differenced logs</i>	Δw	Δp^c	$\Delta(q - l)$	Δu	Δh
Mean (Std.Dev.)	0.177 (0.12)	0.175 (0.11)	0.005 (0.06)	0.0003 (0.13)	0.001 (0.03)
Units of measurement	1985=100	1985=100	1985=1	%	ratio - see appendix

Table 2 reports unit root tests for log levels and monthly differenced log levels for the various data series. The Dickey-Fuller (DF) and particularly the Augmented Dickey-Fuller (ADF) test statistics suggest that nominal wages and prices are probably only integrated in the second order (i. e. , stationary after twice differencing), whereas the other variables are first order integrated. This result is consistent with other studies which investigate integration properties of price indices during hyperinflationary conditions and find evidence for higher order of integration, I(2) or even I(3) (see Taylor, 1991, and Engsted, 1993).

Table 2: Unit Root Tests

	w	p	$(q - 1)$	h	u	$(w - p)$
I(0)–Levels						
DF	16.81	17.79	0.23	-0.17	-0.12	-2.43
ADF	4.53	3.03	0.24	-0.39	-0.11	-2.75
I(1)–First Differences						
DF	-0.88	-1.36	-10.20	-6.22	-9.92	-8.85
ADF	-0.57	-1.32	-6.12	-5.37	-7.60	-9.39
I(2)–Second Differences						
DF	-11.80	-9.96				
ADF	-11.01	-7.16				

Note: Critical values for 5% and 1% significance are -1.94 and -2.59, respectively.

The result that wages and prices are of a higher order of integration than the other series suggests using the real wage. Unit root testing of $(w - p)$, reported in Table 2, confirms that it is $I(1)$. It may be surprising that real wages and Δp are $I(1)$. In the United States, for example, real wages and inflation usually appear to be stationary, depending upon the sample period and precise definitions of the variables. However, other real variables such as real money as well as inflation itself do appear to be $I(1)$ for some developed countries, such as the United Kingdom (e. g. , Johansen, 1992). In Argentina, which like Brazil experienced persistent high inflation for quite some time, real money and inflation have also been found to be $I(1)$ through the 1980s (see Kamin and Ericsson, 1993)⁵. Thus, our cointegration analysis begins with the $I(1)$ variables $w - p$, u , h , $q - l$, and Δp . The presence of Δp in the vector is to allow for the possibility of any long-run nominal illusion in the wage setting process.

⁵Cati et al. (1999) have proposed a methodology for testing for unit roots in the presence of structural breaks and concluded that the series for monthly inflation and interest rates in Brazil are $I(1)$ in the period 1974-1996.

4. Long Run Determinants of Aggregate Real Wages.

In this section we present an analysis of the long run properties of the data. However, an important caveat in our case regards the power of unit-root/cointegration tests in the presence of a relatively short span of the data (e. g. , Campbell and Perron, 1991). As argued by Hakkio and Rush (1991), though, the relevant issue is the length of the sample period relative to the length of what might be considered the long-run⁶. In Brazil, severe inflationary conditions have necessitated frequent index rebasing and lead to the complications of periodic introductions of new currencies. Consequently, the extremely truncated time horizons of bargainers brought about as a result of hyperinflation means that in economic terms a period of eight years may constitute a long run for wage bargainers⁷.

To analyse cointegration of our variables further, the Johansen (1988) procedure is applied and the results of this are reported in Table 3. The results are obtained using lags up to the 8th order in the vector autoregression⁸. Both the eigenvalue statistic and the trace test strongly reject the null of no cointegration in favour of one cointegrating relationship. The Table also reports the standardized

⁶Moreover, recent research has suggested that a long span of data is preferred, but also that for a given time span the use of high frequency data (i. e. , monthly or quarterly observations) does not reduce the power of cointegration tests (e. g. , Haug, 1995, 1996; and Hooker, 1993).

⁷Campos and Ericsson (1990) offer empirical evidence with the analysis of consumption expenditures in Venezuela with a data base of 16 annual observations that can be used to support this general view. According to these authors, what is essential to a long run analysis is the information content per observation in a given period. In Venezuela (1970 to 1985), as well as in Brazil (1985 to 1993), the high volatility of the data in question support this argument.

⁸In deciding the appropriate lag length of the vector-autoregressive (VAR) system for cointegrating purposes we first ran an unrestricted reduced form (URF) for our five 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 five equation VAR system taken together) indicated that no lags above the eight order are significant.

eigenvectors (β') and adjustment coefficients (α).

Table 3: Cointegration Analysis of Real Wages
Sample:1986(1)–1993(11)

Eigenvalues	0.440	0.221	0.116	0.049	0.004
Hypotheses	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	$r \leq 4$
λ_{max}	55.12	23.78	11.7	4.79	0.37
95% critical value	33.5	27.1	21.0	14.1	3.8
λ_{trace}	95.77	40.65	16.87	5.17	0.37
95% critical value	68.5	47.2	29.7	15.4	3.8

Standardised Eigenvectors (β')

$w - p$	u	$q - 1$	h	Δp
1.00	0.81	-2.16	0.76	0.14
34.27	1.00	-30.10	46.81	25.86
-1.34	-0.63	1.00	0.69	-0.69
0.24	-0.01	-3.25	1.00	2.09

Standardised Adjustment Coefficients (α')

0.135	0.008	0.083	0.022	-0.197
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Note: The vector autoregression includes eight lags on each variable, a constant term and monthly dummies. The λ_{max} and λ_{trace} are Johansen's maximal eigenvalue and trace statistics.

The adjustment coefficients (α) measure the feedback effect of the (lagged) disequilibrium in the cointegrating relation onto the variables in the VAR. Specifically, 0.135 is the estimated feedback coefficient for the real wage equation. This positive coefficient implies that lagged changes in the real wage induce further changes (in the

same direction) in current real wages. Its low absolute value implies slow adjustment to remaining disequilibrium, which is compatible with high frequency aggregate data.

Table 4: Estimated Long-Run Relationships
for the Period 1986(1) to 1993(11)

Dep.Variable $\Delta(w - p)$	Engle- Granger	Johansen	ADL Solved Long Run Solution
Δp	-0.27	-0.14	-0.76
$(q - 1)$	0.92	2.16	0.76
h	-0.79	-0.76	-1.15
u	-0.16	-0.81	-0.05
R^2	0.77		
DW	1.03		
DF	-5.36		
ADF	-4.64		

Note: All regressions include intercept term and seasonal dummy variables.

In the long run, therefore, real wages are negatively affected by the tax wedge, the inflation rate and unemployment. The labour productivity elasticity is similar to other examples in the literature (e. g. , Hall, 1986, for the UK; and Carruth and Schnabel, 1993, for Germany) and suggests the existence of a sizeable long run wage-productivity relationship which appears inconsistent with competitive labour markets. This is reinforced by the very low unemployment elasticity reported in columns 1 and 3 of Table 4. Such result is particularly interesting in the context of a developing country where the labour force is usually assumed to grow rapidly over time and new workers entering the labour market may over time be expected to compete away wage gains.

5. Econometric Modelling of Short Run Wage Dynamics.

The next stage is to investigate a dynamic model using information from the preceding co-integration analysis⁹. The Engle-Granger representation theorem (Engle and Granger, 1987) establishes that if a group of variables form a valid co-integrating vector then it is possible to obtain a valid error correction representation of the data which is not liable to the problems of spurious regression. Following Engle and Granger (1987), a valid dynamic specification should include information from the long-run co-integrating relationship between the variables in question (see Table 4), in addition to differenced (and therefore stationary) terms in the other explanatory variables. After some experimentation with the dynamic terms, the following preferred model is obtained:

$$\begin{aligned} \Delta(w - p)_t = & \underset{(0.02)}{-0.04} - \underset{(0.05)}{0.08}\Delta p_{t-1} - \underset{(0.04)}{0.09}\Delta u_{t-1} - \underset{(0.14)}{0.66}\Delta h_t \\ & + \underset{(0.09)}{0.29}\Delta\Delta_{12}h_t + \underset{(0.6)}{0.05}\Delta(q - l)_{t-1} \\ & - \underset{(0.04)}{0.08}\Delta(mw - p)_{t-1} + \underset{(0.04)}{0.40}\Delta\Delta_{12}(w - p)_t \\ & - \underset{(0.05)}{0.11}\Gamma_{t-1} + \sum s \end{aligned} \quad (3)$$

Coefficient standard error in parentheses

$R^2 = 0.83$ $\sigma = 2.9\%$ AR $F(6, 69) = 0.10$

$ARCH$ $F(6, 63) = 1.13$

$NORM$ $X^2(2) = 0.31$ HET $F(27, 47) = 0.85$

$RESET$ $F(1, 74) = 0.71$

$N = 1986(1) - 1993(11)$

⁹To establish appropriate orders of dynamics for modelling consideration a starting point was the estimation of an autoregressive distributed lag (ADL) with eight successive lags and seasonal dummies. No significant loss of information is found in restricting the model to six lags. The ADL model is not reported but its solved long run solution is presented in the last column of Table 4.

The term τ is the error correction term obtained from the ADL model reported in Table 4, σ is the estimated standard deviation of residuals and $\sum s$ are monthly seasonal dummy variables. The equation satisfies all test statistics at the 5% level of significance and presents coefficient estimates which accord with theoretical predictions. We have also included the lagged change in the official minimum wage ($\Delta(mw - p)_{t-1}$) as an additional explanatory variable. Its negative coefficient suggests that the rate of growth of real wages will slow down in face of previous real changes in the minimum wage; i. e. the minimum wage has a short-run restraining effect illustrating its importance in coordinating wage inflation¹⁰. However, the inclusion of $\Delta\Delta_{12}(w - p)$ shows that real wage growth will be higher if wage bargainers observe their wage accelerating over the previous year.

A growing tax wedge (Δh) will restrain wage growth in the short run, although this is mitigated by the positive relationship between real wage growth and an acceleration in the tax wedge over the previous year ($\Delta\Delta_{12}h$). Lagged changes in the price level (Δp) affect negatively current changes in the real wage as do lagged changes in unemployment (Δu); the coefficient for the latter is statistically significant but rather small in absolute terms. The short run productivity elasticity ($\Delta(q - l)$) has its expected coefficient but it is also small and not significant. The error correction term (τ) attracts the expected negative coefficient and shows that 11 per cent of any deviation from long run equilibrium is made up in a month.

¹⁰However, Carneiro and Henley (1998) show that the importance of the minimum wage as a coordinating tool has declined through the 1980s.

Figure 1

Recursive estimates of the coefficients on lagged unemployment and lagged inflation in equation (3)

Parameter constancy is evaluated through the recursive coefficients plotted in Figures 1 through 3. Figures 1 and 2 show recursive estimates for the main coefficients in (3) which appear well inside the ex ante standard errors and empirically constant over time. The coefficient of the error-correction term appears remarkably stable.

Figure 2

Recursive estimates of the coefficients on annual changes in the real wage and the error correction terms in equation

In addition, from Figure 3, none of the breakpoint Chow statistics (showing the effect on structural stability of each additional observation) are significant at the 5 percent level. Panel (a) in Figure 3 plots the one-step residuals and the corresponding equation standard errors and show that these vary little over time, supporting empirical constancy.

Figure 3

One step residuals and the corresponding calculated standard errors; and the sequence of break-point Chow statistics for equation (3)

6. Exogeneity and Policy Implications.

Weak exogeneity of the current-dated regressors in equation (3) above is required for its analysis as a single equation 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 changes¹¹. If the marginal processes of current dated variables change while the conditional model remains constant, then super exogeneity holds.

The tax wedge enters (3) contemporaneously, and so the non-constancy of its marginal processes is of particular 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) and Ericsson and Irons (1995), we use a univariate autoregressive (AR) model for Δh to evaluate its constancy over time¹².

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

$$\Delta h_t = \underset{(0.02)}{0.05} + \underset{(0.11)}{0.34} t_{-1} - \underset{(0.19)}{0.35} h_{t-2} - \underset{(0.19)}{0.14} h_{t-3} + \underset{(0.11)}{0.08} h_{t-4} + \text{seasonals} \quad (4)$$

$$\begin{aligned} N = 1986(1) - 1993(11) \quad R^2 = 0.29 \quad \sigma = 3\% \quad AR \quad F(6, 73) = 2.02 \\ ARCH \quad F(6, 67) = 0.60 \quad NORM \quad X^2(2) = 42.8 \\ HET \quad F(19, 58) = 0.84 \quad RESET \quad F(1, 78) = 0.53 \end{aligned}$$

¹¹ See the collection of papers in Ericsson and Irons (1994) and Charemza and Deadman (1997) for an illustrative and didactic explanation on how to test for exogeneity.

¹² The annual acceleration of the real wage ($\Delta \Delta_{12}(w-p)$) and the annual acceleration in the tax wedge ($\Delta \Delta_{12}h$) also enter contemporaneously in the conditional model. The marginal models for these variables, however, will not be tested for exogeneity here as they are not the parameters of interest of our model.

Figure 4 shows both the one-step residuals and the sequence of break-point Chow statistics for equation (4). 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. According to their approach, the determinants of the non-constancies should be insignificant if added to the conditional model (3). One way of accomplishing this is by including binary dummies at each structural break in the marginal model (4).

Figure 4

One step residuals and the corresponding standard errors, and the sequence of break-point Chow statistics for equation (4)

The expanded model now becomes:

$$\begin{aligned} \Delta h_t = & \frac{0.04}{(0.02)} + \frac{0.36}{(0.09)} t_{-1} - \frac{0.29}{(0.15)} h_{t-2} - \frac{0.19}{(0.15)} h_{t-3} + \frac{0.05}{(0.09)} h_{t-4} \\ & + \frac{0.09}{(0.03)} D87(1) - \frac{0.17}{(0.03)} D89(7) + \text{seasonals} \end{aligned} \quad (5)$$

$N = 1986(1) - 1993(11) \quad R^2 = 0.58 \quad \sigma = 2\% \quad AR \quad F(6, 71) = 1.05$
 $ARCH \quad F(6, 65) = 0.58 \quad NORM \quad X^2(2) = 3.98$
 $HET \quad F(19, 58) = 0.73 \quad RESET \quad F(1, 78) = 0.59$

The improvement in the diagnostic statistics is clear. The dummy variables are for the observations indicated by their names. They coincide with two of the major attempts at economic intervention in Brazil, namely the Bresser Plan of 1987 and the Summer Plan of 1989. Engle and Hendry (1993) propose the use of the determinants of non-constancies in the marginal process to test for super-exogeneity:

$$\begin{aligned} \Delta(w - p)_t = & - \frac{0.4}{(0.02)} - \frac{0.07}{(0.05)} \Delta p_{t-1} - \frac{0.09}{(0.04)} \Delta u_{t-1} - \frac{0.67}{(0.15)} \Delta h_t \\ & + \frac{0.29}{(0.09)} \Delta \Delta_{12} h_t + \frac{0.04}{(0.06)} \Delta(q - 1)_{t-1} \\ & - \frac{0.08}{(0.04)} \Delta(mw - p)_{t-1} + \frac{0.41}{(0.04)} \Delta \Delta_{12}(w - p)_t \\ & - \frac{0.12}{(0.05)} \Gamma_{t-1} + \frac{0.4}{(0.03)} D87(1) + \frac{0.02}{(0.04)} D89(7) + \text{seasonals} \end{aligned} \quad (6)$$

$R^2 = 0.84 \quad \sigma = 2.9\% \quad AR \quad F(6, 67) = 0.18$
 $ARCH \quad F(6, 61) = 0.84$
 $NORM \quad X^2(2) = 0.51 \quad HET \quad F(29, 43) = 0.95$
 $RESET \quad F(1, 72) = 1.88$
 $N = 1986(1) - 1993(11)$

The two dummy variables are insignificant confirming the super exogeneity of the tax wedge in the conditional model. As a result, equation (5) cannot be used to determine the tax wedge even when (3) is identified as a constant real wage equation. As argued by Hendry and Ericsson (1991), super exogeneity is not invariant to renormalisation and therefore the inverted real wage equation might not be constant. Estimating a tax wedge equation inverted from (3) should provide further empirical evidence of super exogeneity:

$$\begin{aligned}\Delta h_t = & -\frac{0.03}{(0.02)} - \frac{0.05\Delta p_{t-1}}{(0.03)} - \frac{0.002\Delta u_{t-1}}{(0.03)} - \frac{0.33\Delta(w-p)_t}{(0.07)} \\ & + \frac{0.42\Delta\Delta_{12}h_t}{(0.5)} + \frac{0.005\Delta(q-1)_{t-1}}{(0.04)} - \frac{0.06\Delta(mw-p)_{t-1}}{(0.03)} \\ & + \frac{0.08\Delta\Delta_{12}(w-p)_t}{(0.04)} - \frac{0.09\Gamma_{t-1}}{(0.04)} + \text{seasonals}\end{aligned}\tag{7}$$

$$R^2 = 0.67 \quad \sigma = 2.1\% \quad AR \quad F(6, 69) = 4.51$$

$$ARCH \quad F(6, 63) = 2.64$$

$$NORM \quad X^2(2) = 6.16 \quad HET \quad F(27, 47) = 1.43$$

$$RESET \quad F(1, 74) = 3.73$$

$$N = 1986(1) - 1993(11)$$

The sequence of break-point Chow tests for this equation is shown in Figure 5 and indicates the non-constancy of the inverted equation for Δh .

Figure 5
The sequence of break-point Chow statistics for equation (7)

The mis-specification of equations (4) and (7) has an important policy implication. While it is possible to obtain a real wage equation which appears constant over the sample period examined, it is not possible to invert that equation to generate a constant, well specified equation for changes in the tax wedge. As illustrated by Ericsson and Irons (1995), one important implication of super exogeneity is that error correction models such as (3) can have a forward-looking interpretation. Also, it means that policy can affect agent behaviour. This happens through the variables entering the conditional model, although not through the parameters of that model. For example, government policy might very well affect the tax wedge and so the real wage. 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, are therefore unwarranted in the present context. To summarise, an important implication of the present work is that government policy in Brazil does have the potential to influence the path of real wages: that follows because (3) has a constant parameterisation in spite of structural change within sample.

7. Conclusions.

This paper has investigated long run determinants and short term dynamics of aggregate real wages in Brazil over the period 1986 to 1993. In the long run, it has been found that real wages are positively influenced by labour productivity while only weakly affected by unemployment. In the short run, changes in the real wages respond negatively to changes in the tax wedge and the unemployment rate, although the latter effect is again not very strong. However, wage bargainers seem to be able to obtain further wage increases whenever they observe an acceleration in the annual change in the tax wedge and also if their real wage had been increasing in the previous year. This result added to the weak unemployment effect over real wages both in the long and short terms is suggestive of the presence of insider power in Brazilian wage determination.

The paper also discusses exogeneity issues and applies several tests. The conditional model for real wages is structurally stable, but combined with an empirically nonconstant univariate marginal model for the tax wedge this implies the super-exogeneity of this latter variable in the conditional model. Since super-exogeneity is not invariant to renormalisation, a structurally stable real wage equation inverted to model the tax wedge may be non-constant. This is in-

deed the case in Brazil. Therefore, the constant relationship found for real wages cannot be used to derive models for the tax wedge. This result is also consistent with the overall finding of Cati et al. (1999) in the sense that the different economic stabilisation plans implemented in Brazil did not induce structural breaks in the Brazilian economy. From this, we reach the conclusion that government policy directed towards changing the structure of labour taxation does have the potential to discipline real wage growth.

APPENDIX.

Wages (W): index of nominal average industrial wages for the state of São Paulo, published by the Federation of Industries of the State of São Paulo (FIESP).

Consumer Prices (P^c): national consumer price index, published by the Brazilian Institute of Geography and Statistics (IBGE).

Unemployment (U): open unemployment rate, published by the Brazilian Institute of Geography and Statistics (IBGE).

Productivity (Q/L): industrial production index divided by the employment index, both published by the Federation of Industries of the State of São Paulo (FIESP).

Tax Wedge (H): producer real product wage ((nominal wages + employers payroll taxes) / wholesale prices) divided by consumer real post-tax wage; wages published by FIESP, prices published by IBGE, employers and employee payroll tax rates published by the Internal Revenue Service. Data on average income tax rates were not available. Marginal income tax rates remained unchanged through the period, varying between 10 and 25% per cent. Thus an average figure of 15% was assumed for the whole time period.

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