Unemployment rate and the real wage behavior: a neoclassical hint for the Colombian labor market adjustment

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Summary

The Colombian urban unemployment rate grew dramatically over the last six years. At the same time the real wage also had a sharp increase. The empirical evidence supports the hypothesis that an exogenous increase in the real wage was a cause of the unemployment growth. The long-run elasticity suggests that one percent increase of the real wage index increases unemployment rate something between 0.7 and 1.0 percent. Therefore it seems necessary that real wage comes back to its equilibrium path for the reduction of the unemployment rate to the natural level.

JEL classification: E24, J30.

Key words: unemployment rate, real wage, cointegration, long-run elasticities.

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1. Introduction

Over the last six years the urban unemployment rate in Colombia has grown dramatically. From a minimum of 7.6% of the labor force in September 1994 , it grew up to a maximum of 20% of the labor force in June 2000. So, understanding the evolution of the labor market in this period is not a point of simple curiosity. In this work we will propose a theoretical interpretation of the rise of unemployment rate and offer some empirical evidence using a standard cointegration approach. The sample, quarterly dated, corresponds to 1984:1 - 2000:2, a period for which information is available and consistent.

The work evolves as follows. The next section shows the main facts over the sample period. Section three outlines a simple neoclassical-type model. Section four discusses the results we obtain by using a standard cointegration approach. Section five provides some conclusions.

2. The facts

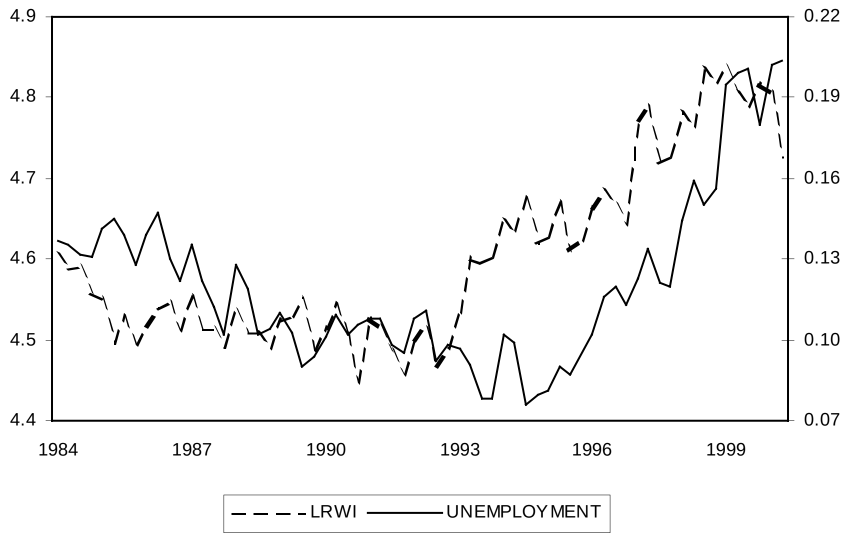
Figure 1 shows two series: the log of the real wage index and the urban unemployment rate (proportion of the labor aged people already working or looking for a job). The former computed from the *Encuesta Nacional de Hogares* (National Housing Survey) and deflated by CPI. The later corresponds to the unemployment rate of the seven major cities[1] of the country (also provided by official statistics).

As Figure 1 shows, there was a period characterized first by a sharp unemployment rate increase: 1994:4 – 2000:2, and second by a (anomalous) strong wage growth.

The hypothesis we maintain is that the increase of the unemployment rate was a reaction to an exogenous growth of the real wage out of the equilibrium path. The behavior of the real wage can be justified on, at least, a couple of points. First, by an unexpected reduction in the inflation rate, maybe due to the combination of an adaptive expectations phenomenon and lack of credibility of the monetary policy[2]. Second, by the existence of a minimum wage, which leads other nominal wage settings in the country, whose level is established on political rather than factual grounds. The theoretical model used to explain our hypothesis is presented next.

[1] These cities account for about a 75% of the total population of the country.
[2] Since 1992, the Colombian Central Bank is leading a reduction inflation program.

Figure 1. Unemployment rate and real wage index (in logs)



3. A model of neoclassical flavor

The model included in this section is neoclassical in essence except by the assumption that the real wage does not clear the market; instead it is assumed exogenous (see Arango and Posada, 2001). The model consists of three basic equations (variables in logs):

$$N_t^d = -\phi(w_t - \theta) + \varepsilon_t^d \;; \qquad \phi > 0 \qquad (1)$$

$$N_t^s = \pi w_t + \varepsilon_t^s \;; \qquad \pi > 0 \qquad (2)$$

where N_t^d stands for labor demand in period *t*; N_t^s for labor supply; w_t for the real wage, θ is a parameter related to a constant multifactor productivity and ε_t^d, ε_t^s represent demand and supply labor shocks both assumed to be $i.i.d.$, zero-mean and constant variance. Defining the unemployment rate, u_t, as[3]:

[3] More precisely, u_t is not equal to the observed unemployment rate but to the log [1/(1-observed unemployment rate)].

$$u_t = N_t^s - N_t^d \tag{3}$$

and after replacement of (1) and (2) in (3) we have:

$$u_t = -\phi\theta + (\phi + \pi)w_t + \varepsilon_t^s - \varepsilon_t^d \tag{4}$$

At least two interesting implications of the model, *vis á vis* the performance of the Colombian labor market (see Table 1 below), arise from expression (4). First, if the real wage (as measured by the log of the real wage index, *lrwi*) is exogenous and there is evidence of it having a unit root, then the observed unemployment rate should also have a unit root[4]. Second, if this joint property is satisfied, then it could be the case that the two variables (or their *proxies*) are cointegrated. Next we show some empirical evidence on this.

4. Results

Since ours is a bivariate approach, we follow Johansen´s (1995) to test for the null hypothesis of *I*(1) processes in a multivariate context[5]. Being this the case, the observed unemployment rate and *lrwi* are *I*(1), equation (4) can conveniently be rewritten as:

$$u_t + \phi\theta - (\phi + \pi)w_t = \varepsilon_t^s - \varepsilon_t^d \tag{4a}$$

so as to represent a long-run relationship between u_t and the real wage[6]. These results as well as those of the cointegrating testing procedure are included in Table 1. Notice, however, that the cointegration testing is done by using the definition of unemployment obtained from the

[4] Arango and Posada (2001) give an economic explanation for the persistence of the observed unemployment rate.
[5] This is motivated on the fact that using the covariates introduces a larger power to this approach (Balmaseda, *et al*. 2000).
[6] In a univariate context , Arango and Posada (2001) show that the null hypothesis of a *I*(1) process for the observed unemployment rate can be rejected for the sub-sample 1984:4-1994:4.

model, that is u_t = log of 1/(1-observed unemployment rate) and log of the real wage index, *lrwi*, as the proxy of the real wage.

Table 1. Cointegration and other tests
1984:1 – 2000:2

Eigenvalue	L-max	Trace	$H_0 : r$	Critical values	
				L-max	Trace
0.3425	24.32 (18.42)	30.04 (22.75)	0	10.29	17.79
0.0939	5.72 (4.33)	5.72 (4.33)	1	7.50	7.50
R	**DGF**	χ^2	***lrwi***	***U***	**Constant**
Exclusion					
1	1	3.84	17.75	5.67	17.35
Stationarity					
1	2	5.99	22.99	23.75	
Weak-exogeneity					
1	1	3.84	$10e^{-5}$	18.13	

Note: These results correspond to a system in which dummies to account for seasonal effects and 8 lags were included. The model was selected on the grounds of normality and no autocorrelation. Cheung and Lai (1993) correction for sample size in parenthesis.

Thus, we have evidence that a stable (Hansen and Johansen, 1993)[7] cointegration relationship between u_t and *lrwi* does exist, the latter variable being weak-exogenous. Table 2 and Figure 2 content the impulse response generated by the *VEC* model while Table 3 presents the results on variance decomposition of the forecast error of the unemployment rate growth. Accordingly, in the long run shocks to the real wage index (*lrwi*) explain most of the variance of the forecast error of the unemployment growth.

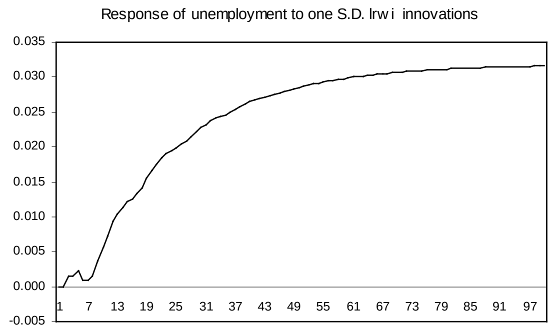
Since there is evidence in favor of a cointegrating relationship between u_t and *lrwi*, an error correction representation is feasible. This model can be written as:

$$\Delta u_t = c_u + \alpha_u \, ect_{t-1} + \sum_{i=1}^{k} \gamma_i^u \Delta u_{t-i} + \sum_{i=1}^{k} \lambda_i^u \Delta w_{t-i} + \varepsilon_t^u \qquad (5)$$

$$\Delta w_t = c_w + \alpha_w \, ect_{t-1} + \sum_{i=1}^{k} \gamma_i^w \Delta u_{t-i} + \sum_{i=1}^{k} \lambda_i^w \Delta w_{t-i} + \varepsilon_t^w \qquad (6)$$

[7] This result is not shown but is available for the reader upon request.

Figure 2. Impulse response functions



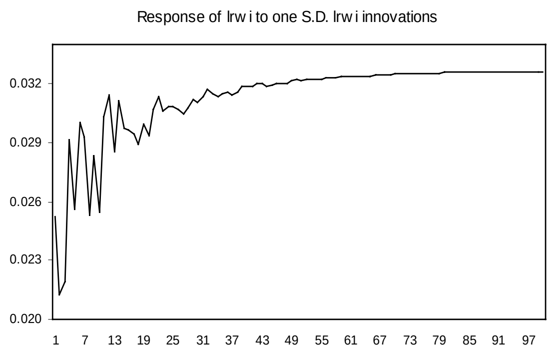


Table 2. Value of responses

	Innovation to Lwri	
Periods ahead	***u***	***w***
1	0.000	0.0252
2	-4.76E-05	0.0212
4	0.0015	0.0291
8	0.0014	0.0253
12	0.0093	0.0314
24	0.0194	0.0308
48	0.0282	0.0320
96	0.0315	0.0326
∞	0.0318	0.0326

Table 3. Variance decomposition

	Unemployment	
Periods ahead	**Unemployment**	**Lrwi**
1	100.000	0.000
2	99.996	0.003
4	94.870	5.129
8	87.448	12.551
12	34.016	65.983
24	4.506	95.493
48	2.957	97.042

where $ect_{t-1} = c + u_{t-1} + \beta w_{t-1}$ (see equation 4a). On the assumption that there is a unique long-run relationship between the variables of the system, the evidence of weak-exogeneity of *lrwi*, means that the stochastic process of *lrwi* does not content any relevant information for the estimation of the long run parameters, assuming these are the parameters of interest; consequently, the system can be reduced to a single equation[8]. The results of the model for u_t [9], after imposing weak-exogenity of *lrwi*, are shown in Table 4.

In summary, a stable cointegration relationship between u_t and *lrwi* does exist; the latter is weak-exogenous; and, in the long-run, increases of 1% in *lrwi* produce increases in u_t of 0.97% (see Tables 1 - 4, and Figure 2)[10].

[8] In a model represented by $\boldsymbol{x}_t = \mathbf{f}(\boldsymbol{z}_t)$, $\boldsymbol{z}_t$ is considered weakly exogenous when the joint distribution of $\boldsymbol{w}_t = (\boldsymbol{x}_t, \boldsymbol{z}_t)$, conditional on the past, can be factorized as the conditional distribution of $\boldsymbol{x}_t$ given $\boldsymbol{z}_t$ times the marginal distribution of $\boldsymbol{z}_t$. In this case, the parameters of the conditional and marginal distributions are not subject to cross-restrictions and the parameters of interest can be uniquely determined from the parameters of the conditional model (Mills, 1993).

[9] Consistently with the weak-exogeneity of *lrwi*, in equation (6) the error correction term happened to be not significant.

[10] However, a word of caution is necessary. Given the definition of u_t as the log [1/(1-observed unemployment rate)], this number is not easy to interpret. In order to obtain a number more readable (as an elasticity) we also made an cointegration exercise between the log (1+observed unemployment rate) and *lrwi*, keeping the same form

Table 4. Error correction model for the unemployment rate
(weak-exogeneity of *lwri* imposed)

Error correction	**Constant**	4.315 (1.890)	***u***	1.000	***Lrwi***	-0.974 (-2.129)	**Alpha**	-0.144 (-5.406)
Deterministic (Seasonal)	**March**	0.013 (1.638)	**June**	0.021 (1.755)	**September**			0.0013 (0.153)
Short-term	**Lag**	**1**	**2**	**3**	**4**	**5**	**6**	**7**
D(lu)		-0.336 (-2.986)	-0.089 (-0.776)	-0.091 (-0.669)	-0.155 (-1.134)	-0.257 (-1.943)	-0.284 (-1.899)	0.158 (-1.067)
D(lrwi)		-0.142 (-2.828)	-0.077 (-1.566)	-0.101 (-2.201)	-0.088 (-1.924)	-0.177 (-3.734)	-0.156 (-3.475)	-0.122 (-2.924)
R-squared: 0.731				**Schwarz SC**: -5.518				
Adjusted R-squared: 0.586				**Akaike AIC**: -6.264				
Sum squared residuals: 0.003				**Mean dependent**: 0.001				
S.E. equation: 0.009				**S.D. dependent**: 0.014				
Log likelihood: 202.665				**F-statistic**: 5.028				

Note: t-values of the long-run parameter estimates are computed following Bardsen (1990).

5. Conclusions

The Colombian urban unemployment rate grew dramatically over the last six years. The real wage also had a sharp increase in the same period according to the real wage index for the National Housing Survey. Moreover, the long-run elasticity suggests that one percent increase of the real wage index increases unemployment rate something between 0.7 and 1 percent.

The empirical evidence supports this hypothesis: an exogenous (or a disequilibrium) increase in the real wage was the cause, or one of the most important, of the unemployment growth. Therefore it seems necessary that real wage comes back to its equilibrium path for the reduction of the unemployment rate to a natural level.

of the cointegration space, the number of lags and the seasonal dummies. Under this view the elasticity is not 0.97% but 0.68%. Given our analysis of data we believe that the elasticity is closer to 0.68% than to 0.97%.

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