Wage growth and minimum wages: cointegration results for Chile

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1. Introduction

The role of minimum wages for labor market performance has been extensively discussed in the literature 4. The central concern, however, seems to have been the effect of statutory minimum wages on employment/unemployment. The link between minimum and nominal wages has been given less attention, although the effect of minimum wages on other labor market variables usually goes through its (implicit or explicit) effect on nominal wages. In Latinamerica, were average inflation is high, it has been natural to ask for the role of the minimum wage as a initiating or propagating inflation factor, and much more research has been devoted to the effect of minimum wages on wage inflation and the wage structure 5.

This paper analyses the effect of statutory minimum wages on the evolution of aggregate nominal wages in Chile for the quarterly sample period 1979.3 - 1997.2. The aim is to identify a long run wage equation in a multivariate context and test whether minimum wages have any consistent long or short run effect on nominal wage determination. In this context, several hypothesis concerning wage behavior, present in the literature, are also tested.

To fulfill this purpose, Hendry’s (1995) general to particular framework to reduce linear dynamic econometric systems and Johansen’s (1988) methodology for estimating cointegrating vectors, in the context of a vector autoregressive system (VAR) are used. This methodology enables us to identify consistently the relevant short- and long-run parameters of a system of variables that determine the wage behavior.

The principal results of the paper are: i) the identification of a cointegration space for nominal wages, consumer prices, labor productivity, the rate of unemployment and minimum wages; ii) the identification of a wage equation within this cointegration space where wages and prices are nearly homogenous of degree one, productivity has a less than complete impact on wages, and the unemployment rate can be excluded from the long run relation; iii) the results indicate that minimum wages have no permanent (long run) effect on the aggregate wages. iv) wages showed a great desequilibrium in the early 80’s. However, as a consequence of a relatively rapid adjustment speed to desequilibria, the des-equilibrium period should have been short. v) Finally, even if minimum

1 We acknowledge gratefully comments of D. Hendry on a previous version of this paper. Of course the usual disclaimer of responsibility is in order.

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wages do not seem to affect average wages in the long run, they seem to have a significant role in the short run.

The outline of the remainder of the paper is the following. First, a basic approach to nominal wage determination, that clarifies the central hypotheses to be tested, is presented. Then, the data sources and stylized facts are discussed. Following, the fundamentals of the econometrics used is laid down. Afterwards, the empirical analysis and results are presented. Finally, the principal conclusions of the study are summarized.

2. The basic wage equation

A general view on nominal wage determination is that wages grow according to expected inflation, labor productivity growth and the excess demand situation in the labor market. Moreover, in some analyses, the influence of different institutional factors has been stressed. There exists a great deal of literature on the subject.

In formal terms, a basic linear model that includes these considerations is:

\[ \Delta w = \alpha \Delta p^e + \gamma \Delta q + \delta f(u) + \theta g(s) \]

where \( \Delta w \) is the growth rate of nominal wages, \( \Delta p^e \) is the expected inflation rate, \( \Delta q \) is the labor productivity growth rate, \( u \) is a variable that measures excess demand in the labor market (for instance the unemployment rate), \( s \) is a catch-all variable for other factors (for instance institutional factors), \( f(.) \) and \( g(.) \) represent functions with positive first derivatives and \( \alpha, \gamma, \delta, \theta \) are parameters.

No standard theoretical frame for this type of relation emerges from the literature, and probably no unique frame does exist. Intuitively, this relation is compatible with a competitive labor market model where, in the long run, nominal wages adjust to expected inflation and exogenous shifts in the marginal productivity of labor, and where in the short run, this relation is distorted by the excess demand situation in the labor market and institutional factors. Likewise, a formally equivalent relation has been derived from a bargaining model where the target wage depends on productivity shifts, the probability of getting a new job, and the real minimum wage (see Bazén and Martin, 1991).

The literature has raised and tested different hypothesis concerning wage formation. First, in the long run, expected inflation should equal actual inflation \( (p^e = p) \), and nominal wages should be homogenous of degree one in prices, i.e., \( \alpha=1 \). In the case that several commodity price indices are included in wage determination, the


7 Some studies are based on uni-equational models as ec.(1), while other include a wage equation in a simultaneous equation model.

8 The empirical literature for the Chilean case does not show coincident results. A short review of this literature can be found in Dresdner & Letelier (1995).
Second, real wages, in the long run, should change according to labor productivity growth. That is, the coefficient of the productivity variable should also show an unitary value ($\gamma = 1$). Third, with the «Phillips curve» approach in background, the parameter associated to the labor demand situation should be negatively significant ($\delta < 0$). As this variable usually is seen as a short run determinant of wages, short and long run situations should be distinguished in this case. Usually, labor excess demand has been measured by the open unemployment rate or its rate of change. Lastly, the influence and persistency in time of different institutional factors on nominal wage determination have been treated in the literature. One specific factor of great concern has been the role of increases in statutory minimum wages on average wages. The standard presumption is that increases in minimum wages should tend to increase nominal average wages. Two routes are usually pointed out for this effect. First, the increase in the minimum wage will rise the wages for those lying at or slightly above the old minimum wage level. Second, workers obtaining wages well above the minimum wage can want to compensate for this, trying to maintain the wage differential constant, or the rate of change in the minimum wage can work as a signal (a base wage change) for wage bargains.

In the present paper, we investigate these hypotheses on wage formation by estimating a cointegration space from a set of labour market variables. By applying linear restrictions on the parameters of the estimated cointegrated vectors, it is possible to test the hypotheses previously discussed. The estimation and testing procedure is applied to the aggregate wages series. The expected results are the identification of a long run relation that «explains» wage development. Moreover, contrasting this relation with actual wages can identify periods of short run disequilibrium. Finally, the specific role of minimum wages, both in the short and long run, is analysed.

3. Data sources and stylized facts

The data source for nominal wages, consumer prices, employment and unemployment rates is the Chilean National Institute of Statistics, NIS (Instituto Nacional de Estadísticas). However, quarterly employment for the period 1979.3 - 1985.2 were not computed by the NIS and therefore were obtained from Jadresic (1986) (excl. the special employment programs PEM & POJH). This serie was linked for the remaining period with the NIS serie. Average productivity is measured as the ratio of production and employment at the aggregate level. Production and minimum wage statistics were collected from the Chilean Central Bank.

If one adds to this hypothesis the unitary price homogeneity hypothesis, one obtains a wage growth path where the labor income share of total income is constant ($\Delta w - \Delta p - \Delta q = 0$).

As it is well known, aggregate figures can hide many sectoral differences that tend to blur or bias the results. Therefore, a natural extension to this work is to repeat the estimation at a more disaggregated level. However, in this paper we are interested in the aggregate effects.
(Banco Central de Chile). Finally, the social security tax rate was obtained for the period 1979.3-1985.4 from Jadresic(1989), and onwards from the Social Security Office (Superintendencia de Seguridad Social). Since only annual figures are available it was assumed that the annual tax rate was valid for all four quarters of the corresponding year.

The period considered in our analysis ranges from the third quarter of 1979 up to the second quarter of 1997. The lower bound responds to the fact that only since the second term of 1979 collective bargaining was reintroduced. This represents a fundamental institutional change that affects the wage determination mechanism.

The pattern over time of the variables considered for the whole economy can be regarded in Figure 1. In Figure 1a, the graphics on the log of average labor productivity (Q) and (the log of) the real wage (W-Pc), suggests the existence of a common pattern between these two variables at the aggregate level. Actual series for productivity register a clear increase at the beginning of the sample period. Average productivity increased 10.3% between 1979.3 and 1982.1. This tendency switch in 1982, showing and fall in productivity of 10.7% in the subsample period 1982.1-1987.4. In the final period productivity recovered, increasing in 44% between 1987.4 and 1997.2. This behaviour is compatible with the significant expansion of the Chilean economy in the second half of the 70s, the well known colapse in 1982 and the later recovery from 1986 and onwards. Real wages tend to show a similar development to productivity growth and over the whole period the two variables grew practically at the same rate. However, within subperiods the development was not totally parallel, showing the wage development a tendency to vary more than productivity. For instance, the switch in the real wage trend in 1982 corresponds well with the change in legislation that de-indexed wages from past inflation and the upsurge in inflation due to, among other factors, devaluation of the exchange rate.

In figure 1b, real average wages are plotted against the real minimum wage. The data indicates that minimum wages grew well below average wages in the first years of the sample up to 1982.1. Thereafter, minimum wages were reduced (in real terms), although falling proportionately less than average wages. Finally, after 1987, they recovered but at a slower pace than average wages. Since nominal minimum wages are a policy parameter, this development seems to show a change in the (implicit or explicit) policy rule used by the authorities to determine minimum wages. However, this policy rule can be affected by the evolution of economic variables such as the unemployment rate, consumer prices, productivity and nominal wages themselves, and at the time be a determinant of other economic variables.

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11 However, due to lags the estimation period is restricted to 1980.3-1997.2.

12 Preliminary results indicated that the stability of the estimated parameters was drastically affected when the series were extended backwards in time.

13 All the series have been normalized for expository purposes.
such as average wages, prices and the unemployment rate. In this sense, it is possible for minimum wages to be co-integrated with other variables and/or to be an important determinant of the short run dynamics of these variables. Over the whole period real average wages increased in 40%, while real minimum wages augmented in only 4%. This fact can “a priori" de-emphasize the potential role of minimum wages on average wages. Obviously, more formal tests are required to settle this issue.

To consider (tentatively) the role of the (log of the) unemployment rate (U), this variable has been graphed together with the net real wage (W-Pc-Q), i.e., the (log of the) real wage deflated by average productivity (see Figure 1c). Although the tendency for the initial years is not quite clear, it seems that from the 1982 crisis and onwards, the unemployment rate showed a negative relation to the net real wage. The unemployment rate peaked when the net real wage attained its lowest level and then it depicted a clear negative trend as real wages recovered. The

FIGURE 1
STYLIZED LABOR MARKET FACTS. CHILE: 1979.3 - 1997.3

NOTES: W-Pc = Log of Real Average Wages.
Q = Log of Average Labor Productivity.
W-Pc-Q = Log of the Real Wage Net of Productivity.
Wmin-Pc = Log of the Real Minimum Wage.
U = Log of the Unemployment Rate.

SOURCES: See Appendix.
pattern of the variables considered suggests that there is a possibility for them to be cointegrated. It is worth mentioning, that contrary to what usually might be expected, the unemployment rate does not show a stationary pattern during the sample period. The unemployment rate registered a procyclical pattern over the expansion period before 1982, and a countercyclical behavior from then onwards. This stylized fact could be a sign of the severe desequilibrium period that the Chilean labour market went through by the late 70s and early 80s and, in this sense, the observed co-movement does not necessarily indicate a long run relation. The graphic analysis is, as it is wellknown, far from conclusive as to whether the unemployment rate should be part of the long run wage equation or not.

4. Econometric theory

A vector autoregressive (VAR) model can be represented in its error correction form:

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_i \Delta Z_{t-i} + \Pi Z_{t-k} + \epsilon_t$$

where $Z$ is a column vector that contains the values for the n-relevant variables in the analysis, $e$ is a iid (nX1) random term column vector, with zero mean and constant variance $(0, \Sigma^2)$ and $D$ is a difference operator.

The short and long run effects can be distinguished in equation (2). The first ones are reflected in the parameters of the $\Gamma_i$ matrices. To recover the long run parameters, however, matrix $P$ is of interest. According to Granger’s Representation Theorem (Engle & Granger, 1987), if the rank ($R$) of this matrix is $R = n$, where $n$ is the number of variables in the model, the $Z_t$ process is stationary. This implies that all variables are I(0), and that $n$ cointegrated vectors exist. In this case, the economy has a static point solution. On the other hand, if $R = r$, where $r < n$, there exists a representation $\Pi$, such that $\Pi = \alpha \beta'$, where both $\alpha$ and $\beta$ are matrices with (nxr) dimension. $\beta$ is called the cointegration matrix and has the property that $\beta'Z_t \sim I(0)$, while $Z_t \sim I(1)$. That is, there exist one or more cointegrated vectors, that reflect (statistical) long run relations, but fewer than n, so the economy shows a dynamic solution.

$\alpha$ is known as the adjustment coefficient matrix, and is generally interpreted as the parameters that establish the way in which the model returns to its long run equilibrium once it has been shocked.

If there exists more than one cointegrated vector, $r > 1$, then the estimation will give a cointegration space, since all linear combinations of the cointegrated vectors will also be cointegrated. In this case, an unique identification of individual vectors will not exist, unless one can use exogenous information for identification (Johansen & Juselius, 1992). Therefore, the test of structural hypotheses is a crucial part of the analysis. Specifically, what one
is interested to test is whether hypotheses with economic content lie within the estimated cointegration space. In the present case, we are interested to test whether the evidence rejects the existence of a long run economic relation corresponding to (1).

5. Model estimation and results

5.1 Long run dynamics

The first step in our analysis involves the estimation of a basic Vector Autoregressive Model (VAR Model). The relevant set of variables is grouped in a so called vector Z (see eq.(2)), where \( Z' = (W, Pc, W_{\text{min}}, U, Q) \) and all variables are in logs. Here \( W \) stands for the level of nominal wages, \( Pc \) is the level of consumer prices, \( W_{\text{min}} \) denotes the level of statutory minimum nominal wages, \( U \) is the unemployment rate, and \( Q \) is the level of average labor productivity. However, we should recall that we are primarily interested in estimating a wage equation as in eq.(1). A linearized stochastic version of this equation in levels is:

\[
\beta_1 W_t - \beta_1 Pc_t - \beta_2 W_{\text{min},t} - \beta_3 Q_t - \beta_4 U_t = \mu_t
\]

where \( \mu_t \) is an error term.

Regarding the level of integration of the series, the results obtained from different unit root tests\(^{14}\) indicated that the level of productivity (\( Q \)), the unemployment rate (\( U \)) and the level of minimum wages (\( W_{\text{min}} \)) become stationary when they were differentiated once. That is to say, they probably are integrated of order one. As far as nominal wages (\( W \)) and the consumer price index (\( Pc \)) are concerned, their behavior is more compatible with an I(2) process. However, since for instance a linear combination between nominal wages and the level of consumer price index (\( W-P \)) could be integrated of order one, and this combination in turn could cointegrate with the other I(1) variables, these tests are far from conclusive from the proper way to model the vector autoregressive model (VAR).

To model this system we use a General to Particular strategy, in the sense of Hendry (1995). In principle, one can search for identifying different economic structural relations from this set of variables. However, the aim in this paper is focused primarily to identify a nominal wage equation, and the selection of variables included in the VAR was done with this purpose in mind. Therefore, other economic relations (price equation, unemployment rate equation, etc.) might not be well specified in this system.

With regards to the non stochastic component of the VAR, this was determined by a test based on the so called Pantula Principle, developed by Johansen (1992c). This test basically implies that the rank order of models, with different deterministic components, are compared and the most restrictive alternative, not rejected by the usual cointegration tests, is selected. The test results indicated that the preferred model had no trend, and it

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\(^{14}\) Augmented Dickey-Fuller tests with and without constant, trend and seasonals were performed with different lag lengths. The tests tended to show consistently the results reported in the main text so that it seemed unnecessary to use other, more sophisticated tests.
included an intercept as an unrestricted component of the VAR (it corresponds to model 2 in Table 1). In addition to the five variables in Z, three intervention dummies were also included as unrestricted regressors. The first one (Dpol) captures the change in payroll taxes during the period. These taxes showed a dramatic reduction, specially between 1980 and 1982, as a consequence of changes in the labor legislation. Thereafter the tax rate remained relatively stable\textsuperscript{15}. This should obviously affect the equilibrium wage level in the economy. The other two intervention dummies (D82 and D84), correct for the abrupt changes in price formation related to the economic crisis that detonated in 1982 and that showed an upsurge in 1984 respectively\textsuperscript{16,17}.

The reduction procedure began by determining the number of lags of the VAR. Providing that gaussian behaved errors where found, the reduction process was not rejected until a 4 lag VAR was estimated.

The application of Johansen’s test on this system provides the unrestricted cointegration vectors. Within the context of Johansen’s procedure, two basic tests are usually reported. They are the test on the trace of the stochastic matrix (trace-test) and the test on the maximum number of eigenvalues ($\lambda$ max-test). These are presented with correction for degrees of freedom in Table 1. Both tests do not reject the null of one cointegration vector ($p=1$). However, if one looks at the results not corrected for degrees of freedom (not presented here), there can be some doubt if there are one or two cointegration vectors. For that purpose we checked the plots of the two first vectors (corrected for the short run dynamics and deterministic factors) and of the corresponding recursive eigenvalues (see Figure 2), it can be seen that the second vector does nor seem very stationary and its corresponding recursive eigenvalue is not very stable either. With all these evidence we are inclined to accept the existence of one cointegration relation.

The obtained eigenvectors ($\beta$s) and adjustment coefficients ($\alpha$s) are shown in Table 2. The first vector is standardized by the level (log) of nominal wages ($W$), and it is interpreted as a wage equation. This wage equation is expected to depend on prices and labour productivity, and could also depend on minimum wages and unemployment. Of special interest in this context is

\begin{itemize}
  \item \textsuperscript{15} Since we do not have reliable information about the precise magnitude of the average tax rate changes, it seems preferable to take account of these effects by allocating approximative orders of magnitude in the dummy variable to represent the tax rate in the relevant years, instead of using the available information to calculate the net real wage level.
  \item \textsuperscript{16} D82 takes the value of 1 the quarters 1982.3 and 1982.4 and zero otherwise, while D84 takes the value of 1 the quarters 1984.3 and 1984.4 and zero otherwise.
  \item \textsuperscript{17} For different models, the standard deviations of the regressions were importantly reduced and the characteristics of the disturbances improved, when these dummies were introduced. This tends to indicate that this specification is correct. Dpol showed out to be especially relevant for the wage and unemployment equations, while D82 and D84 improved basically the price equation in the VAR. The standard deviations with (WD) and without dummies (WO) are:
\end{itemize}

\begin{table}
\begin{tabular}{|c|c|c|c|c|}
\hline
Equation & $W$ & $P$ & $W_{\text{min}}$ & $Q$ & $U$ \\
\hline
WD & 0.012 & 0.013 & 0.036 & 0.023 & 0.061 \\
WO & 0.015 & 0.016 & 0.036 & 0.023 & 0.070 \\
\hline
\end{tabular}
\end{table}
whether minimum wages have any significant effect on the long run wage development.

If we focus on the first vector (the cointegration vector), we will observe that, with the exception of the unemployment rate, all the signs of the standardized estimated $\beta$ coefficients are in line with the theory (see eq.(3)). That is to say, changes in consumer prices, minimum wages and labor productivity are all expected to affect positively the nominal wage level. On the other side, changes in the unemployment rate are expected to affect negatively the nominal wage level. Moreover, the magnitude of the $\beta$ coefficients for the minimum wage and unemployment rate variables are very small, while the coefficient for the productivity variable is well below unity, which would be required for constant labor income share (see footnote 8).

We proceed to test structural hypotheses on the wage equation. Finally, two alternative sets of restrictions imposed on the cointegration vector are presented. In the first set (I), we restrict to zero the coefficients of the minimum wage

<table>
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<tr>
<th>$\lambda_{\text{max}}$ test</th>
<th>H0: $p = 0$</th>
<th>H0: $p \leq 1$</th>
<th>H0: $p \leq 2$</th>
<th>H0: $p \leq 3$</th>
<th>H0: $p \leq 4$</th>
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<tbody>
<tr>
<td>Model 1</td>
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Trace test

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<th>H0: $p \leq 3$</th>
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<td>Model 3</td>
<td>125</td>
<td>74.8</td>
<td>41.9</td>
<td>16.2</td>
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Model 1: Intercept restricted to long run model
Model 2: Intercept in short run model
Model 3: Intercept in short run model and trend restricted to long run model.
The results presented correspond to the ones corrected for degrees of freedom.

* An asterisk indicates the most restrictive alternative not rejected by the test statistic. For this purpose the table should be read row-wise.

$p = \text{Rank of the matrix } ( = \alpha \beta^\prime)$. 
and unemployment rate variables in the cointegration vector, while leaving the coefficients of the consumer price and productivity variables free. In the second set of restrictions (II), in addition to the abovementioned restrictions, we add the hypothesis of price homogeneity and complete effect of productivity growth on nominal wages.

Thereafter, we attempted to identify weakly exogenous variables by testing the significance of the α-coefficients estimated by means of Johansen’s maximum likelihood procedure, conditioned on the previous structural test results. Given that every error correction equation will include only one error correction term, the relevant null hypothesis involves that the corresponding α will not be significantly different from zero. We tested each α coefficient separately and then in joint hypotheses, subject to the previous obtained structural results.

The results of all these tests are presented in Table 3. The first (composi-

\[18\] This because we have only one cointegration vector.
cy determined by the authorities has not have, at least for the sample period considered, long run effects on the average nominal wage level.

On the other hand, the wage equation does not include the level of unemployment either as a long run determinant. This result is much in line with previous results for the wage equation in the Chilean economy, that point out that general labor market conditions doesn’t seem to affect the nominal wage development19.

The second (composite) structural hypothesis tested additionally whether the data rejected that the coefficients for the consumer price level (unitary price homogeneity) and the productivity level were equal to one. The results reject strongly this hypothesis20. If one regards the point estimates for these variables, the results indicate that the price coefficient should be slig-

<table>
<thead>
<tr>
<th>W</th>
<th>P</th>
<th>Wmin</th>
<th>Q</th>
<th>U</th>
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<table>
<thead>
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<th>W</th>
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<th>Q</th>
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<td>0.022</td>
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<td>-0.008</td>
<td>-0.009</td>
<td>0.154</td>
<td>-0.124</td>
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<tr>
<td>0.001</td>
<td>-0.003</td>
<td>0.004</td>
<td>0.001</td>
<td>0.0004</td>
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</table>

19 For references see those given in footnote N°5.
20 We also tested whether only the price coefficient was different from one and this was rejected at the 5% level.
ghtly above one, while the productivity coefficient should well below unity. The result is interesting but one can only speculate about it.

That the price coefficient is above unity would indicate that nominal wages respond more than fully to consumer price changes in the long run. This result is difficult to accept as long run behaviour in a theoretical sense. However, this pattern is compatible with a lagged inflation wage readjustment scheme, in the presence of a declining inflation rate. This scenario was prevalent in Chile for most of the sample period.

The result for the productivity coefficient on the other hand does conflict with the notion of a constant labor share in value added. However, the empirical evidence for the Chilean economy during the sample period tends to confirm a decline in the labor share, and one explanation for this fact could be that labor productivity increases were only partially transferred to nominal wages.

### TABLE 3

**STRUCTURAL HYPOTHESES ON THE STANDARDIZED \( \beta \) AND WEAK EXOGENEITY TESTS**

\[ Z' = (W \ P \ C \ W_{min} \ Q \ U) \]

#### JOINT TESTS FOR \( \alpha \) AND \( \beta \) VECTORS

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<th>( \alpha )</th>
<th>( \chi^2 )</th>
<th>P Value</th>
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</thead>
<tbody>
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<td>( 1^*0^*0 )</td>
<td>( * * 0 0 * )</td>
<td>7.2</td>
<td>(0.12)</td>
</tr>
<tr>
<td>( 1-10-10 )</td>
<td>( * * 0 0 * )</td>
<td>68.6</td>
<td>(0.00)**</td>
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#### WEAK EXOGENEITY TESTS FOR INDIVIDUAL \( \alpha \) SUBJECT TO \( \beta \) (Ho I)

<table>
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<th>( \alpha )</th>
<th>( \chi^2 )</th>
<th>P Value</th>
</tr>
</thead>
<tbody>
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<td>( W )</td>
<td>( \alpha 1 = (0 + + + +) )</td>
<td>28.0</td>
<td>(0.00)**</td>
</tr>
<tr>
<td>( P )</td>
<td>( \alpha 1 = (+ 0 + + +) )</td>
<td>8.2</td>
<td>(0.04)*</td>
</tr>
<tr>
<td>( W_{min} )</td>
<td>( \alpha 1 = (+ + 0 + +) )</td>
<td>6.1</td>
<td>(0.11)</td>
</tr>
<tr>
<td>( Q )</td>
<td>( \alpha 1 = (+ + 0 + +) )</td>
<td>7.2</td>
<td>(0.06)</td>
</tr>
<tr>
<td>( U )</td>
<td>( \alpha 1 = (+ + + + 0) )</td>
<td>25.9</td>
<td>(0.00)**</td>
</tr>
</tbody>
</table>

#### Restricted Cointegration Vector with Ho I

\[ \beta (1 + 0 + 0) \]

<table>
<thead>
<tr>
<th>( W )</th>
<th>( P )</th>
<th>( W_{min} )</th>
<th>( Q )</th>
<th>( U )</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \beta 1 )</td>
<td>1.00</td>
<td>-1.074</td>
<td>0.00</td>
<td>-0.649 0</td>
</tr>
</tbody>
</table>

Notes: + implies that the coefficient value is free; * significant at the 95% level; ** significant at the 99% level
From the results of the weak exogeneity tests it can be concluded that both the minimum wage (Wmin) and the labor productivity (Q) variables are weakly exogenous. The weak exogeneity results are interesting, although they could be expected. They point out that neither minimum wages nor productivity were contemporaneously affected by nominal wage changes during the period.

Given these results, it is possible to obtain a restricted cointegration vector that represents the long run wage behaviour (see Table 3). This relation depends only on consumer prices and labor productivity. This will be our long run wage equation.

Before proceeding to the short run analysis, one might wonder whether the period chosen, characterized by great and rapid changes in the Chilean economy, is apt to estimate long run relations. In other words, if the economy suffered great imbalances during the sample period perhaps the model will not be able to capture the “structural long run” behaviour of the economy. One way of judging whether the results show this long run behaviour could be to reestimate the model for a reduced sample period, that excludes the most dramatic swings in the variables during the period, and compare with the results for the full sample. This procedure was implemented, excluding the period 1980-1983.4, and the results indicate that similar estimated coefficients can be obtained for this reduced sample once the same steps developed above are repeated.

**Figure 3** shows the behavior of the cointegration vector, not corrected for short run dynamics nor deterministic factors, over the period considered. If we focus on the wage standardized vector (Clw), the effect of the crisis of the early 80s, which was preceded by the economic boom of the late 70s can be identified. This is caught as a positive desalignment between nominal wages and the long run conditioning factors for the cointegration vector. The ongoing correction of the desequilibrium gap is probably the result of the desindexation of the economy, and the dramatic effect of the devaluation on the short run rate of inflation.

### 5.2 Short run dynamics

Now, we turn to the short run analysis. This involves the estimation of an Error Correction Model (ECM), similar to eq.(2). We know now that all terms

21 Note that the notion of long run implicit in this comment is that of steady state situations where the economic relations are stable, different from the one present in the cointegration literature.

22 We owe this point to C.W.J. Granger, who kindly commented a former version of this paper.

23 The results obtained when the estimations were made excluding 1980.1 1983.4 are the following:

Long Run Equilibrium Wages for the Whole Economy $W_{LR} = 1.08 P_c + 0.70 Q + 0.05U$.

24 Using the restricted coefficients of the obtained cointegration vector in Table 3, we can estimate a series for the long run equilibrium level of nominal wages ($W_{LR}$). Then, $Cl_w = W - W_{LR}$ where $W = Aggregate nominal wages$.

25 The only difference lies in that we include deterministic variables, discussed in the text, both in the levels and rates of change components of the cointegration analysis.
in the equations are of the same order of integration. Thus, least squares can be applied. We develop the following reduction procedure: First, we estimated the VAR in its error correction form (VECM), including one error correction term (the cointegrating relation). Since we are primarily interested in modelling the wage equation, we used the information on the weakly exogenous variables to reduce the system to a three equations system: the wage, consumer price and unemployment equations. Using the system FIML procedure in PcFiml a parsimonious VECM was obtained through the elimination of non significant variables, subject to the weakly exogenous variables and the condition that the equation disturbances were well behaved. We used as short run deterministic variables Dpol, D82, D84, seasonal dummies and a constant term. Various diagnostic tests indicated that the residuals had the desired properties and the reduction procedure was supported by an F-test\(^{26}\). In the final results, we detected that lagged unemployment variables did not enter any of the three equations. Thus, since the unemployment variable is not included in the error correction term, nor in the short run dynamics of the system, it seems that it is affected by the other variables, but does not affect contemporaneously or lagged any of the other two variables. Therefore, we estimated a two equation model, leaving outside the unemployment equation\(^{27}\). The estimated coefficients, as expected, were exactly the same. This allowed us to reduce the system further. The final results with diagnostic tests are presented in Table 4. Recursive test statistics, composed of equation one

\(^{26}\) The test result was \(F(30,121)= 1.1983\) (0.2439).

\(^{27}\) We included the contemporaneous unemployment rate as an unrestricted regressor in each equation and tested whether its elimination was not rejected by an F-test, which was the case.
step residuals and system one step F-tests and break-point F-tests are presented in Figure 4.

The system diagnostic tests show that the residuals do behave well and it does not seem to be any model specification problems. The model is quite parsimonious with only two other variables explaining the short run dynamics of the wage equation and one variable the short run dynamics of the price equation. Moreover, the stability test results depicted in Figure 4, show that there does not seem to be important breaks in the parameters.

The first aspect to be considered refers to the adjustment of wages to the long run equilibrium. The impact of changes in the level of wages on their rate of change (\(\frac{\partial dW}{\partial W}\)) amounts to -0.29\(^{28}\). This implies that slightly less than 30% of the same disequilibrium is adjusted within one quarter. Within one year approximately 74% of the total disequilibrium should have been adjusted. This result indicates that friction is not excessively important for the aggregate wage level, and that for instance the imbalance in the labor market that took place during the 1982 crisis should have been basically adjusted within a period slightly over one year.

When the short run dynamics of the ECM for the wage equation is analyzed, it can be seen that besides its own lagged values, nominal wages are affected by minimum wages and labor productivity. The impact of minimum wages on average wages appears to be positively significant. The rate of change in minimum wages affects contemporaneously and with two quarters lag the wage inflation rate. The total multiplicator effect is 2.5. Thus, although according to previous results this variable does not seem to affect aggregate wages in the long run, it does affect them in the short run. Heuristically, one can think of the adjustment process in the following manner. Minimum wage increases tend to raise wages in the short run. This however creates a gap between the actual wage level and the long run wage level. This gap tends to decrease the nominal wage level to its long run position, and the effect of minimum wages on the long run average wage is null, ceteris paribus.

Regarding the short run effects of other variables on wage adjustment, the results show that changes in labor productivity affect the rate of change in aggregate wages. Assume that labor productivity increases. First, there is a contemporaneous increase, but then, with a two period lag, there is a decline in short run wages. Actually the second effect is stronger implying an overall negative multiplier. This tendency should be reversed in the long run, since labor productivity increases wages, according to our long run equation.

Finally, the price equation tells that the short run rate of inflation will react positively to the one lagged value of the inflation rate itself and negatively to the lagged value of labor productivity. These results correspond to what one intuitively would expect.

\(^{28}\) This value is close to the one originally obtained from Johansens procedure (-0.23), which is an additional indication of the adequacy of the reduction procedure.
6. Conclusions

1) The application of Johansen’s methodology to estimate cointegration relations and Hendry’s general to particular approach to obtain parsimonious estimations from ECMs, allowed us to identify a cointegration vector for an aggregate labor data set.

2) Imposing restrictions on this space it was possible to identify a long run relation, that (heuristically) can be interpreted as an aggregate wage equation. The wage level depends, with (almost) unitary elasticity, on consumer prices and with less than unitary elasticity on labor productivity. Moreover, the results indicate that aggregate wages do not depend on the unemployment rate.

3) The tests do not reject the exclusion of minimum wages from the cointegrating relation, indicating that minimum wages might not affect average wages in the long run.

4) Tests on weak exogeneity were applied to both data sets. The results showed that labor productivity and minimum wages are weakly exogenous.

5) An analysis of the long run behaviour of wages compared with their actual performance showed that a severe desequilibrium situation arised in the early 80’s, that generated a positive gap between actual and long run wages. However, this imbalance should have been corrected within the first half of the eighties. This is because the results indicate that the adjustment speed of wages to desequilibria.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coeff.</th>
<th>t-value</th>
<th>t-prob</th>
<th>Variable</th>
<th>Coeff.</th>
<th>t-value</th>
<th>t-prob</th>
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<tbody>
<tr>
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<td>-0.244</td>
<td>-2.619</td>
<td>0.0113</td>
<td>DP_1</td>
<td>0.350</td>
<td>3.029</td>
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<td>0.156</td>
<td>4.314</td>
<td>0.0001</td>
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<td>-0.174</td>
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<td>3.133</td>
<td>0.0027</td>
<td>Cl_1</td>
<td>-0.154</td>
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<td>1.890</td>
<td>0.0638</td>
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<td>0.542</td>
<td>3.107</td>
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<td>Cl_1</td>
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<td>0.0000</td>
<td>D82</td>
<td>0.067</td>
<td>5.706</td>
<td>0.0000</td>
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<tr>
<td>Constant</td>
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<td>9.088</td>
<td>0.0000</td>
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<td>0.029</td>
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<td>0.7120</td>
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\[\sigma = 1.17\%\]
\[\sigma = 1.41\%\]

Multivariate tests: \(F_{AR}(20,92) = 0.98(0.49); \chi^2_N(4) = 3.44(0.49); F_{HET}(42,119) = 0.92(0.61)\)

System tests: \(F_{AR}\) is a vector autocorrelation test; \(F_{HET}\) is a vector heteroscedasticity test; \(\chi^2_N\) is a vector normality test.
is relatively high. A little less than 30% of the disequilibrium dissipates in one quarter.

6) The short run cointegration analysis shows that the rate of change in nominal wages registers a significant inertia with regard to lagged values of minimum wages variations, implying that although minimum wages do not affect average wages in the long run, they do have an impact on its rate of change in the short run.
7. References


