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CYCLICALITY OF REAL WAGES AND ADJUSTMENT COSTS

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Abstract

This paper tackles the issue of the procyclicality of the real wage. We present a dynamic relationship between real wages and employment consistent with the long-run stationary equilibrium using a cointegrated VAR model. We find that wages are anticyclical and that a negative relationship between real wages and employment is necessary to achieve an economically identifiable stationary long run solution. The contentiousness of the topic does not appear so important once we recall some measurement issues and economic features of the Italian labour market.

Introduction

Recently, several reviews of the evidence on real wage cyclicalities have stressed the difficulty of drawing any firm conclusions. The existence of a host of measurement and methodological issues (e.g. Abraham and Haltiwanger 1995) and the difficulty expressed by the theories to select among alternative movements of real wages over the cycle (e.g. Rotemberg and Woodford 1991; Brandolini 1995), make it hard to support any

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definite conclusion.¹

The rejection of a priori belief may always be questioned emphasizing different sources of disturbances.

In particular, we emphasize that the empirical evidence described in the literature, is inferred from highly aggregated data. Thus, it could be not appropriate to use it to shed light on the underlying microeconomic parameters (see, for instance, Hamermesh 1993).

This paper provides some empirical evidence on the countercyclicality of real wages using aggregate data for the Italian manufacturing sector. We recognize that it would be incorrect to associate this hypothesis with a particular approach to macroeconomics, nevertheless, we show this result for a competitive model of labour demand with a simultaneous specification of labour and real wages, an appropriate consideration of the time-series properties of the data, and a short-run structure specified in a way consistent with a cointegrated long-run structure.²

The organization of this paper is as follows. In the next section the estimates of the structural model are reported. The economic identification is drawn in Section 2 while Section 3 deals with the economic interpretation of the achieved results.

¹ See, for instance, Bils (1987), Barsky and Solon (1989), Nickell and Symons (1990), Solon, Barsky and Parker (1994) among others.

² The theoretical analysis of dynamic models of labour demand with turnover costs is discussed in Nickell (1986) and Bertola (1990). See also Hamermesh (1993).

The observation of a negative correlation of aggregate output and the real wage over the business cycle and the existence of a long-run aggregate labour demand relation are empirical matters. Although these results are consistent with the theory, we are aware that "no aggregate relation of any sort exist", because "the existence of an aggregate labor-demand curve is not demonstrable or refutable" (Hamermesh 1993).

Some concluding remarks are outlined in the final section. Unit root tests and cointegration analysis results are gathered in the Appendix.

1. A Dynamic Econometric System

Perfectly competitive models with a standard neoclassical production technology imply that output and employment fluctuations should be associated with countercyclical movements in the real wage. An aggregate demand shock (given technology and capital stock) increases output and employment reducing real wage. These aspects, however (as pointed out by the authors quoted above), arise in neoclassical as well as in neo-keynesian models. The effect now described may be obstructed by high costs of adjustment on the labour demand side. According to Neftci (1978), in this context simple regression analyses may be misleading. If these costs are relevant, the long-run relations and the short-run dynamic adjustments should receive more attention in empirical analysis.

In a neoclassical model, both employment and real wages are endogenous variables. In the presence of remarkable adjustment costs, factors affecting labour demand may change over the cycle without influencing employment. These aspects are investigated specifying and estimating a multivariate cointegration model for labour demand in Italy.

The following structural model has been formulated and estimated using a modelling sequence developed by Johansen (1988; 1991) (see also Johansen and Juselius 1990) and a growing body of literature on cointegration and dynamic econometrics (see Lütkepohl 1991 and Juselius 1993 for a methodological discussion). It consists in specifying a VAR model in $I(1)$ space, carrying out the relative cointegration analysis. Parsimony is subsequently achieved by removing the insignificant regressors in

the VAR model in $I(0)$ space. The process ends up with a structural model (estimated in FIML) imposing specific restrictions on each equation.³

-The Data

The analysis has been performed on the following data vector $x'=[n_t, y_t, (w-p)_t, h_t, T_t]$, where w_t is the wage for employee in the Italian manufacturing sector, n_t stands for manufacturing employment, p_t is the price of output, y_t is the real output (value added) in the sector. The vector includes, as proxies, per capita worked hours (h_t) and labour turnover rate (T_t) in the "large firms" (firms with more than 500 employees).⁴

The VAR is augmented by three shift dummy variables (D1 D2 and D3), respectively, to account for the short-run effects concerning the second oil shock, to cope with some anomalous data relative to the turnover variable, and to account for the effect of the industry restructuring process during the 1980s.

The data are quarterly seasonally adjusted over 1976:1-1995:1. All the variables are in logs. With the exception of the output price (which is $I(2)$), the series appear to be $I(1)$. However some ambiguity arises over the labour turnover variable T_t which

³ Inference procedures for the cointegration have been developed by Johansen on the basis of the maximum-likelihood analysis of the VAR error-correction representation. See Johansen (1995).

⁴ The labour turnover variable T_t may proxy employment adjustment costs: a higher labour turnover indicates low adjustment costs and *vice versa*.

could be $I(0)$ (see Appendix 1)⁵. Finally, the trend (t) has been forced to lie in the cointegration space. Thus, we expand the VECM to include the fact that there are linear trends in the levels of the data. This should allow the non-stationary relationships in the model to drift. Since there are no quadratic trends in the levels of the data we use in the model, there will be no time trend in the short-run "structure" (e.g. Harris 1995; Hendry 1995a).

Since we assume a competitive model of labour demand, the output price variable is presumed to be exogenous and to have only a short-run effect. Thus, we replace the $I(2)$ variable with an $I(0)$ alternative through differencing.

The usual tests (absence of serial correlation in the residual and significance of the parameter estimates) on the initial lag length suggest that a VAR with 2 lags on all stochastic variables suffices.⁶

⁵ The results based on the ADF test may be questionable, since the "appropriate lag-length" is unknown and a "comprehensive simulation study has not been done on the significance levels" (e.g. Hatanaka 1996). Thus, too many lags may lead to some loss of efficiency, while too few lags may seriously distort the test. The practical rule for determining the "appropriate lag-length" is that it should be relatively small in order to save degrees of freedom, but large enough to allow for the existence of autocorrelation. However, ignoring a sufficiently high lag order would have a serious effect upon the performance of the test. To consider the underlying properties of the processes that generate time series variables, we have followed this procedure of defining the underfitting and overfitting of lag orders.

⁶ In order to keep the length of this paper within reasonable limits, the full output results (VAR models, Parsimonious VAR models, tests and charts on parameter constancy and other statistics) is not reported. Further results may be provided by the authors upon request. Some result is discussed in more detail in Chiarini (1996).

-The multivariate cointegration model

Consider the following conditional vector autoregressive model as a statistical model for the data-generating process:

$$x_t = \sum_{i=1}^k A_i x_{t-i} + \xi + \phi D_t + \varepsilon_t, \quad \varepsilon_t = \text{NIID}(0, \Omega) \quad (1)$$

where ξ is a vector of constants and D_t is a vector of deterministic variables. We may rewrite (1) in the error-correction form, taking into account that $k=2$ (see Hendry 1995a, Johansen 1995, among others):

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \Pi x_{t-1} + \xi + \phi D_t + \varepsilon_t, \quad (2)$$

where $\Gamma_1 = -A_2$; $\Pi = -(I - A_1 - A_2) = \alpha\beta'$ is reduced rank and α and β are $n \times r$ matrices (α is the speed of adjustment to disequilibrium and β is a matrix of long-run coefficients). (2) defines a multivariate cointegration model, and the rank r determines the number of linearly-independent stationary relations between the levels of the variables.

-Testing for long-run restriction and weak exogeneity

The process of formulation and testing of hypotheses on the coefficients α and β , provides an unrestricted cointegration space spanned by three stationary eigenvectors β_1 . To identify long-run relations we formulate restrictions on the individual relations (economic questions has been formulated as linear restrictions on all of the β vectors).⁷ A general formulation of overidentifying restrictions on the individual equations is given by:

$$\beta = (H_1 \phi_1, \dots, H_r \phi_r) \quad (3)$$

⁷ On the basis of this procedure, Wickens (1996) provides a critical analysis of the practical usefulness of cointegration analysis.

where the matrices $(n \times s_i)$ H_i , $i=1, \dots, r$, express linear economic hypotheses to be tested against the data and ϕ_i is an $(s_i \times 1)$ vector of parameters to be estimated in the i th cointegration relation. We may impose the same restrictions in H_i by specifying an $(n \times k_i)$ matrix R_i , such that $R_i' \beta = 0$. Identification may, therefore, be defined with respect to the linear restrictions specified in R_i and H_i (Johansen and Juselius 1994; Johansen 1995). For $r=3$, hypotheses about β can be formulated as $\beta = (H_1 \phi_1, H_2 \phi_2, H_3 \phi_3)$, such that R_1 , R_2 and R_3 identify the vectors in β if

$$\text{rank}(R_i' H_j) \geq 1, \quad i, j=1, 2, 3; \quad i \neq j \quad \text{and} \quad \text{rank}(R_i' (H_j, H_m)) \geq 2$$

for different i, j, m . The test procedure is based on the likelihood ratio test statistic. Similar tests have been performed on the rows of α , corresponding to the test for weak exogeneity. Imposing $r-1$ "exactly identifying" restrictions and a normalizing coefficient on each cointegration vector we obtain the following restricted cointegration space:

$$\begin{aligned} (w-p) &= (y-n) + 0.84h - 0.0003t \\ h &= 0.775y - (w-p) + 0.003t \\ n &= y - h - 0.804(w-p) + 0.158T - 0.004t \end{aligned}$$

Three restrictions were imposed on the first and second vector and two on the third vector. The restricted cointegrating space, described by the last three equations, has been accepted by the data: the LR test for these hypotheses, distributed as a $\chi^2(9)$ under the null, gives a value of 2.808 [0.9714]. The graphs in Figure 1 and 2, based on the recursive estimates of the underlying VAR model $I(1)$, highlight the model costancy over the sample in terms of the 1-step residuals with $\pm 2\sigma_1$ and individual equation break-point Chow F-test, indicating a well defined cointegration analysis (see Appendix 2).

Cointegrated systems must be interpreted cautiously. The coefficients are not elasticities, and all the other dynamic relations between the variables which are specified in a VAR should be taken into account (see, Lutkepohl 1991; 1994 and Johansen 1995). Thus, although these coefficients seem represent an economically meaningful structure, they are only indicative. The first cointegration relation can be interpreted as a wage relation. The wage share, defined as $[(w-p)-(y-n)]$, cointegrates with the worked hours variable. The second and third cointegration relations seem to represent labour demand relations, respectively, for working hours and employment. The hypothesis that employment cointegrates with the labour turnover variable cannot be rejected by the data. As labour turnover may proxy turnover costs, this result implies that long-run employment may be somewhat higher if turnover costs (due, for instance, to job security provisions) reduce.

Johansen and Juselius (1994), point out that cointegrating vectors may contain evidence about two behavioral relations, reflecting at least two types of agents with disparate goals (e.g. demanders versus suppliers). The above long-run structure seems to be consistent with this interpretation, where unions and firms may interact in such a way that equilibrium is restored once it has been violated.

The association between economic theory and cointegration has been questioned in some literature (for instance, Harvey 1997; Pesaran 1997; Wickens 1996, among others). Here we justify this association. According to Hendry (1995b), "The parameters of an economic structure may include those of agent's decision rules, but there is no presumption that these must be derived by representative agent optimization". It must be emphasized the extensive gap between an abstract theoretical framework of an (inter-temporal) optimizing agent and an empirical model of aggregate behaviour. In such circumstances, where the available data do not measure the theoretical solutions directly, we cannot

apply rigid procedures: theory needs to be used flexibly to drive the empirical analysis. Here, the Haavelmo's distinction between agents' intentions and actual realizations in the form of aggregate variables may be strongly advocated (see Spanos 1989 and Juselius 1993). In this context, the Johansen' approach, which provides a framework (the cointegration space) wherein an adequate statistical model and an estimable form of the theory can be used to delineate the issues involved, is particularly relevant.

-The VECM Model

Using first differences of the variables we reformulate the model with the error-correction terms included. Furthermore, to eliminate non-significant regressors, we impose specific restrictions on each equation and then re-estimate a simultaneous model using the FIML procedure (see, for instance, Hendry 1995a).

FIML estimates : 1976:1-1995:1 (standard errors in parentheses)

$$\Delta y_t = 0.112\Delta y_{t-1} + 0.951\Delta n_{t-1} - 0.150ecm2_{t-1} + 0.104ecm3_{t-1} + \\ (.095) \quad (.283) \quad (.039) \quad (.041) \\ - 0.0124D1 - 0.035D2 + 2.15Const. \\ (.005) \quad (.008) \quad (.401)$$

$$\Delta n_t = 0.539\Delta n_{t-1} - 0.0316\Delta h_{t-1} + 0.087\Delta y_{t-1} + 0.0384\Delta^2 p_t + \\ (.098) \quad (.021) \quad (.035) \quad (.022) \\ - 0.0058D2 - 0.0025D3 \\ (.0027) \quad (.0009)$$

$$\Delta(w-p)_t = - 0.159\Delta(w-p)_{t-1} + 0.253\Delta h_{t-1} - 0.31\Delta y_{t-1} - 0.222\Delta^2 p_t \\ (.081) \quad (.09) \quad (.112) \quad (.083) \\ - 0.088ecm1_{t-1} - 0.379ecm2_{t-1} - 0.0244D1 - 0.013D2 + \\ (.038) \quad (.081) \quad (.0067) \quad (.0089) \\ - 0.0495D3 - 2.54Const. \\ (.0085) \quad (.556)$$

$$\Delta h_t = 0.186\Delta h_{t-1} - 0.113\Delta(w-p)_{t-1} + 1.66\Delta n_{t-1} - 0.013\Delta T_{t-1} + \\ (.09) \quad (.091) \quad (.377) \quad (.006) \\ - 0.38ecm2_{t-1} + 0.295ecm3_{t-1} - 0.0288D1 - 0.0317D3 \\ (.058) \quad (.044) \quad (.009) \quad (.008)$$

$$\Delta T_t = .33\Delta T_{t-1} + 12.32\Delta n_{t-1} + 5.63\Delta^2 p_t + 9.26ecm1_{t-1} + \\ (.11) \quad (4.92) \quad (1.40) \quad (.966) \\ - 7.65ecm2_{t-1} - 4.044ecm3_{t-1} - 0.344D2 - 0.347D3 + \\ (1.32) \quad (.772) \quad (.144) \quad (.126) \\ - 59.65Const. \\ (10.31)$$

$$ecm1_t = \Delta(w-p)_t - \Delta y_t + \Delta n_t - 0.84\Delta h_t + 0.00032Const. + ecm1_{t-1}$$

$$ecm2_t = \Delta h_t - 0.775\Delta y_t + \Delta(w-p)_t - 0.00317Const. + ecm2_{t-1}$$

$$ecm3_t = \Delta n_t - \Delta y_t + \Delta h_t + 0.804\Delta(w-p)_t - 0.158\Delta T_t + 0.004Const. \\ + ecm3_{t-1}$$

$$F_{AR}^o(125, 206) = 1.2517[0.0798]; \quad \chi_N^2(10) = 8.758[0.555]; \\ F_{HET}^o(315, 437) = 0.935[0.735].$$

The multivariate tests (F_{AR} : F-test for the hypothesis of no serial correlation; F_{HET} : no heteroscedasticity; χ_N^2 : chi-square test for normality) suggest that the model has approximately white, normally distributed errors. Notice that the constant coefficients for the differenced variables $ecm1_t$ are their long-run trends.

2. Economic Identification

Noteworthy features of the structural model and the underlying cointegration analysis are:

i)- With regard to the identification of the long-run structure, we find that restricting the real wage to zero in all of the cointegrating vectors is rejected by the data: $\chi^2(3) =$

28.338[0.000]. Moreover, the following restriction is also rejected:

$$(n-y+h-aT+bt) \in Sp(\beta) \quad \chi^2(2)=26.413[0.000]$$

where a, b are estimated parameters. The long-run equilibria require a negative relationship between real wages and employment.

ii)- The restricted cointegration relations were tested for lying in the cointegration space jointly with testing for employment being long-run weakly exogenous for the stationary relations β_1 ($\chi^2(15)=18.56[0.234]$). This result is consistent with a rigid labour market with high adjustment costs. However, it contrasts with the results provided by other analyses (e.g. Bertola 1990). The estimates have been performed over a period characterized by important employment protection legislation.⁸ The model shows that changes of employment still react (although with modest effects) to lagged changes of working hours and output.

iii)- Firms, keeping employment constant, operate with hours adjustments. Firms cannot adjust to shocks by laying off and hiring workers, but they can adjust the hours worked by each worker.

This result induces us to recognize that a correct analysis should model both employment and per capita worked hours as endogenous variables.

iv)- The impulse response analysis, generated by unanticipated unit shocks (Figures 3-6), shows the relative importance of the short-run adjustment to the long-run relations.

⁸ The institutional rules devoted to employment protection singled Italy out as the industrial country with the highest degree of employment rigidity. However, theoretical models and empirical evidence do not suggest a clear conclusion on the relationship between costs of adjustment and employment. See, for instance, Bertola (1990), Emerson (1988), Grubb and Wells (1994), Snower (1990) and the literature quoted therein.

v)- A short-run change in the output variable generates a countercyclical behaviour of real wages. This result is confirmed by the responses of real wage growth to a one-time unit impulse in Δy_t (see Figures 3-6). The responses return to zero after few periods. However, a one-time impulse in Δy_t yields a lasting effect on y_t and $(w-p)_t$. This is due to the nonstationary features of the data. Output and real wages may settle at a different equilibrium values. Figure 6 shows that an increase in Δn_t leads to countercyclical behaviour: a negative relationship between the product wage and output (and employment).

These results appear to be robust to the outcome created by considering the labour turnover variable to be $I(0)$. The dominant features of the estimated alternative model are:⁹

- i) A negative coefficient of the output in the real wage equation, $-0.275(0.109)$ (standard errors in parenthesis).
- ii) A short-run effect of the real wage in the output equation: $-0.192(0.078)$.
- iii) Employment is not longer weakly exogenous for the long-run parameters, but, as expected, its adjustment to deviations from disequilibria is rather sluggish: the estimate speed of adjustment parameters is $-0.033(0.009)$ and $0.043(0.0127)$. Again, this result is consistent with a labour market with a high degree of employment rigidity (high job security provisions).

⁹ Including stationary variables in VAR models will increase the cointegration rank. In this case, the stationary variable T_t forms a cointegration relation by itself. Thus, the practical implication of assuming $T_t \neq I(0)$ in our cointegration analysis is that one of the three cointegration relations β_1 contains only the labour turnover variable. The full system is provided upon request.

3. Economic Interpretation

We argue that the evidence of a countercyclical real wage is a complex outcome of various elements.¹⁰

Firstly (*measurement effect*), employment data used to specify the labour demand and construct the gross earnings per employee is provided by the National Account in *standardized labour units*. These measure the number of full-time workers, taking into account the informal economy (irregular workers, second job, work done by illegal immigrants). This measure does not include workers receiving CIG (Wage Supplementation Fund) provisions, that is, the mechanism to ensure the cyclical adjustment of labour input with very reduced adjustment costs. This measurement feature, *ceteris paribus*, tends to create a countercyclical response of earnings per employee to demand shocks.

A further measurement effect concerns the price deflator used to deflate wages. Wages deflated by a consumer price index have been found (e.g. Abraham and Haltiwanger 1995) more procyclical than wages deflated by an output price. As stressed by Nickell and Symons 1990, if the relationship between employment and the real wage is viewed on the (labour) demand side, then the correct price deflator must be an output (value-added) price.

Secondly (*cost adjustment effect*), we have stressed that high adjustment costs, due to job protection, generate a sluggish adjustment process of employment.

In this context, payments to labour in recession may fall

¹⁰ Mocan and Baytas (1991; 1993), using standard vector autoregressive models and structural time-series models (where the trend of the real wage rate is estimated through Kalman filter), show that the cyclicalities of real wages depends upon whether the business cycles are driven by demand or supply shocks. In particular, they show that a supply shock generates procyclical real wages whereas a demand shock produces countercyclicalities.

short of the wage determined by firm's profit-maximizing conditions, whereas payments in booms rise above this level because employment does not adjust. The presence of adjustment costs for changes in the labour input could make the real wage more procyclical (although the latter might move in the same direction during both, upturns and downturns). However, if employment does not adjust, the empirical evidence of a short-run negative relationship between real wages and employment may disappear. In this context, changes in output and overtime hours would detect the cyclicity of real wages.

Per capita worked hours play a central role (e.g. Bils 1987). Employment, measured in standard labour units, does not account for changes of per capita worked hours which tends to represent the overtime effect in the wage equation (the working hours variable shows a negative coefficient in the employment equation, although a one-time impulse in Δh_t has no significant short-run effect on labour demand). If hires and fires are viewed as costly by firms, it will be optimal for them to bear some costs of changing hours: overtime hours move over the cycle and raise wages (see Figures 3,5 and 6). However, this effect turns out to be smoothed by the rise of the procyclical cost of marginal hours. In booms, overtime hours lengthen the period over which capital input is utilized. This implies that the marginal product of overtime hours declines, reducing real wages. Of course this result is related to the state of the capacity utilization.

Lastly (*economic effect*), for most of the sample considered the process of wage formation was characterized by the automatic indexation of wages to prices. However, during the 1980s the degree of indexation was substantially reduced (the elasticity to CPI fell from about 0.7 to a value less than 0.5). A further important feature is the predeterminate nature of nominal wages. Nominal contracting provisions (the minimum duration of the contracts over the period considered was 36 months) weaken the

relationship between wages and current demand condition (some evidence for the Canadian economy are reported in Card 1990).

Imperfectly indexed wages¹¹ and staggered nominal contracts implied by a centralized structure of collective bargaining (although characterized by a low degree of coordination between different bargaining levels) may also generate countercyclical real wages, particularly if price (both, consumption and output prices) responsiveness to demand shock is high.

4. Conclusions

We have shown, for manufacturing industry quarterly data that: a) real wages are anticyclical; b) a negative relationship between real wages and employment in the long-run is required by the data; c) labour input is costly to adjust: employment does not react to disequilibrium errors (it is found weakly exogenous for the long-run parameters); d) employment adjustment costs permit hours adjustment and reduce the amount of labour turnover. All these results are unambiguously supported by the macroeconomic data.

We use recent contributions in cointegration analysis which provide an appropriate framework for the joint analysis of short- and long-run behaviour. Of course our findings may be sensitive to the employment measurement used and to the choice of deflator (as, for instance, in Geary and Kennan 1982, and Abraham and Haltiwanger 1995), however, staggered contracts, wage indexation schemes and employment protection regulations should provide a valid interpretation of the topic.

¹¹ The government began the process of dismantling the wage indexation in 1983; in 1984 it decreed a ceiling to wage indexation (for only one year) and in 1985 achieved a more significant reform of the wage indexation mechanism, reducing the frequency of adjustments from quarterly to half yearly and lowering the degree of indexation. The Incomes Policy Agreement in July 1993 established the definitive cancellation of the automatic indexation.

Appendix 1

Unit Root Tests: ADF(4)

	C		C	C+t		C+t
n	0.415	Δn	-4.022**	-1.617	Δn	-4.433**
y	-1.740	Δy	-5.511**	-2.688	Δy	-5.749**
(w-p)	-2.180	$\Delta(w-p)$	-5.089**	0.497	$\Delta(w-p)$	-5.544**
h	-1.663	Δh	-4.378**	-1.482	Δh	-4.498**
T	-2.146	ΔT	-3.409**	-1.916	ΔT	-4.834**
p	-1.887	Δp	-2.235	-1.232	Δp	-2.955

C=ADF with Constant; C+t=ADF with Constant and Trend.

Critical values: 5% = *; 1% = **.

Correlogram of the variables (lag length: 11)

	1	2	3	4	5	6	7	8	9	10	11
n	.996	.988	.976	.962	.947	.931	.914	.894	.874	.845	.823
y	.993	.981	.966	.955	.947	.941	.937	.934	.936	.936	.933
(w-p)	.993	.991	.990	.988	.985	.984	.984	.984	.984	.982	.980
h	.799	.624	.615	.625	.585	.331	.304	.327	.310	.250	.214
T	.491	.430	.426	.789	.339	.236	.233	.528	.123	.050	.084
p	.999	.999	.999	.999	.998	.998	.997	.997	.996	.995	.994

Appendix 2

Testing for cointegration in the five-equation system yields the results reported in Table 1.

Table 1
Cointegration Analysis

r	1	2	3	4	5
μ	.565	.515	.330	.143	.062
Max	64.17**	55.79**	30.79**	11.91	4.96
Tr	167.60**	103.51**	47.67*	16.88	4.96

Stationary eigenvectors (β_i) Nonstationary eigenvectors (v_i)

	β_1	β_2	β_3	v_4	v_5
n	1.000	0.097	-23.040	-93.320	-0.981
y	-1.461	-0.772	1.000	-15.390	1.017
(w-p)	1.405	1.012	-11.420	7.991	1.000
h	-0.843	1.000	-20.930	-14.630	-0.267
T	0.061	0.065	6.643	1.000	-0.008
t	0.003	-0.003	-0.187	-0.379	-0.016

Weights α_i

	α_1	α_2	α_3	α_4	α_5
n	0.002	0.033	-0.026	-0.000	0.015
y	-0.000	0.115	0.123	-0.002	-0.054
(w-p)	-0.082	-0.407	-0.091	-0.000	-0.069
h	-0.002	-0.441	0.217	-0.002	0.066
T	-0.217	0.269	-1.196	-0.023	0.410

Max= Maximum Eigenvalue test; Tr= Trace test; μ = eigenvalues.

This table determines the rank of Π as 3 (see equation 2). However the variables in D_t affect the distribution of test

statistics, thus, the critical values provided by the cointegration analysis (e.g. Johansen and Juselius 1990; Osterwald-Lenum 1992) are only indicative. Following Juselius (1995), additional information may be obtained looking at the eigenvalues of the companion matrix A of the model (1). This matrix has $n+k=10$ roots (inside the unit circle). The moduli of the two largest roots are 0.892 and 0.892, supporting $(n-r)=2$ common stochastic trends and three stationary relations.

Identification restrictions have been placed on the unrestricted cointegration vectors β_1 (reported in Table 1), defining the ECMS in the FIML model of Section 1. The restrictions to identify the three separate cointegrating vectors have been tested jointly with testing for n being long-run weakly exogenous for all the stationary relations (β), y and h being long-run weakly exogenous for the first cointegrating relation (β_1) and w being long-run weakly exogenous for the third stationary relation (β_3).

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Figure 1: System Recursive Evaluation Statistics

1-Step Residuals with ± 2 *S.E.

Figure 2: Individual Equation Break-Point Chow F-Test (1% Crit.)

Figure 3: Impulse Response (Shocked variable Δy)

Figure 4: Impulse Response (Shocked variable $\Delta(w-p)$)

Figure 5: Impulse Response (Shocked variable Δh)

Figure 6: Impulse Response (Shocked variable Δn)

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