Modelling the Causal Relationship Between Prices and Wages in Greece: A Multivariate Empirical Approach

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Abstract

This paper attempts to investigate the causal relationship between wages and prices in Greece, using quarterly data for the period 1976:1 to 1993:2. We employ multivariate cointegration techniques in conjunction with variance decomposition and impulse response analysis. The results provide evidence of a significant causal effect running from wages to prices, thus supporting the Post–Keynesians' "mark-up pricing" view.

Περίληψη

Η μελέτη αυτή επιχειρεί να διερευνήσει την αιτιώδη σχέση μεταξύ μισθών και τιμών χρησιμοποιώντας τριμηνιαία στοιχεία για την περίοδο 1976:1 έως 1993:2. Στην εμπειρική ανάλυση χρησιμοποιήθηκε η τεχνική της πολυμεταβλητής συνολοκλήρωσης (multivariate cointegration) σε συνδυασμό με τη μέθοδο διάσπασης της διακύμανσης (variance decomposition) και την πρόβλεψη της αντίδρασης των μεταβλητών του συστήματος σε τεχνητές εξωτερικές διαταραχές (impulse response analysis). Τα αποτελέσματα παρέχουν ενδείξεις για την ύπαρξη μιας σημαντικής αιτιώδους επίδρασης από τους μισθούς προς τις τιμές, στηρίζοντας έτσι τη θέση των Μετα–Κεϋνσιανών στο σχετικό θέμα ("mark-up pricing" view).

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1. Introduction

There is a widely held view that wage growth and inflation are causally related. The nature of the relationship between these two economic indices has long been the subject of theoretical debates concerning the question of their relative exogeneity.

The main theories that have been put forward and tested can be summarised as bellow: 1) The expectations-augmented Phillips-curve theory supports the existence of a two-way causal relationship between wage growth and inflation; 2) the original wage-type Phillips-curve model argues that it is inflation that causes wages to increase; 3) Post-Keynesians' "mark-up pricing" view supports that wages cause inflation; 4) Neoclassicists consider that for a given level of productivity inflation causes nominal wage growth to respond so as to preserve the real wage; 5) Monetarists deny the existence of any causal relationship between wage growth and inflation.

The literature concerning the empirical analysis of the relationship between wage inflation and price inflation reports rather mixed findings. Thus, for example, Mehra (1977) and Fosu and Huq (1988) suggest a bi-directional causality between wage growth and inflation; Barth and Bennett (1975) and Mehra (1991) find one way causal effects running from prices to wages; Shannon and Wallace (1986) present evidence supporting unidirectional causal effects from wages to prices. Last, Gordon (1988) and Darrat (1994) report lack of any causal relationship between the two variables. For Greece, Psaradakis (1991) using cointegration finds causal effects running from both directions.

In this paper, we investigate empirically the nature of the relationship between wages and prices in Greece, using quarterly data for the period 1976:1 to 1993:2. We employ multivariate cointegration modelling in conjunction with variance decomposition and impulse response analysis. The results suggest the existence of a causal effect running from wage inflation to price inflation, thus supporting the Post-Keynesian "mark-up pricing" view, for the case of Greece.

The rest of this paper is organised as follows: Section 2, discusses theoretical and methodological issues. In section 3, the empirical results are reported and discussed. Concluding remarks are given in Section 4. The relevant Tables and Figures of the study, are presented in the Appendix.

2. Theoretical and Methodological Issues

2.1 Stationarity

A common fact in the empirical analysis is that many macroeconomic series are characterised by nonstationarities, implying that the classical t and F-tests are not

appropriate (Fuller, 1976), since they may lead to invalid results. Thus, as it is required in standard econometric analysis, the series under consideration are examined, first, for unit root nonstationarity employing the methods developed by Fuller (1976) and Dickey and Fuller (1981).

Testing for a unit root in a time series, requires the computation of one of the three OLS regressions presented below.

$$y(t) = \alpha y(t-1) + u_1(t) \tag{1}$$

$$y(t) = \mu^* + \alpha^* y(t-1) + u_2(t) \tag{2}$$

$$y(t) = \tilde{\mu} + \tilde{\beta}T + \tilde{\alpha}y(t-1) + u_3(t) \tag{3}$$

In model (1) the adjusted t-Statistic $Z(t_\alpha)$ is used to test the null hypothesis of a unit root, i.e. H_0: $\alpha = 1$ against the stationary alternative $\alpha < 1$. If the equation includes a constant (model 2), we calculate the $Z(t_{\alpha^*})$ and the joint statistic $Z(\Phi_1)$ to test the null H_0: $(\alpha^*, \mu^*) = (1, 0)$ against the alternative $(\alpha^*, \mu^*) \neq (1, 0)$. Further, if the equation includes a constant and a trend (model 3), $Z(t_{\tilde{\alpha}})$ and the joint statistics $Z(\Phi_2)$ and $Z(\Phi_3)$ are used. In particular, $Z(\Phi_2)$ tests the null $H_0: (\tilde{\alpha}, \tilde{\beta}, \tilde{\mu}) = (1, 0, 0)$ while $Z(\Phi_3)$ tests the null $H_0: (\tilde{\alpha}, \tilde{\beta}, \tilde{\mu}) = (1, 0, \mu)$.

Thus, if we establish I(1) properties, we can proceed with testing for cointegration among the examined variables.

In case where the variables are found to be stationary, the estimation of a classical vector autoregressive (VAR) system specification is the appropriate methodology to examine the information content of the variables concerned.

2.2 Cointegration

The long-run relationship between a number of series can be looked at from the viewpoint of cointegration (Engle and Granger, 1987). Let x(t) be a vector of n-component time series each integrated of order one. Then x(t) is said to be cointegrated CI(1, 0), if there exists a vector φ such that

$$s(t) = \varphi' x(t) \tag{4}$$

is I(0). Stationarity of s(t) implies that the n variables of x(t) do not drift away from one another over the long-run, obeying thus an equilibrium relationship. If φ exists, it will not be unique, unless x(t) has only two elements. The Engle and Granger (1987) approach can deal with the possibility of only one linear combination of variables that is stationary. Recent advances in cointegration theory (Johansen and Juselius, 1990) have developed a maximum likelihood (ML) testing procedure on the number of cointegrating vectors which also allows inferences on parameter restrictions. The ML method uses a vector autoregressive (VAR) model

$$\Delta x(t) = \sum_{i=1}^{q-1} \Pi_i \Delta x(t-i) + \Pi_q x(t-q) + \mu + v(t) \tag{5}$$

where $x(t)$ is a $n \times 1$ vector of variables, Π_q is a $n \times n$ matrix of rank $r \leq n$, μ is a $n \times 1$ vector of constant terms and $v(t)$ is a $n \times 1$ vector of residuals. The testing procedure involves the hypothesis $H_2{:}\alpha\beta'$, where α and β are $n \times r$ matrices of loadings and eigenvectors respectively, that there are r cointegrating vectors $\beta_1, \beta_j^i, \ldots, \beta_r$ which provide r stationary linear combinations $\beta' x(t-q)$. The likelihood ratio (LR) statistic for testing the above hypothesis.

$$-2 \ln Q = -T \cdot \sum_{i=r+1}^{T} \ln\left(1-\hat{\lambda}_i\right) \tag{6}$$

is a test that there are at most r cointegrating vectors versus the general alternative (trace), where λ_i corresponds to the $n-r$ smaller eigenvalues. The $n \times r$ matrix of cointegrating vectors β can be obtained as the r, n-element eigenvectors corresponding to λ_i.

The LR test statistic for testing r against $r+1$ cointegrating vectors is given by

$$-2 \ln (Q{:}r \mid r+1) = -T \cdot \ln (1-\hat{\lambda}_{r+1}) \tag{7}$$

The above tests (6) and (7) are used to determine the significant eigenvalues and the corresponding number of eigenvectors.

2.3 Variance decomposition and impulse response analysis

The next step of this study considers variance decomposition analysis. This technique involves the transformation of the system into its moving average representation and then to obtain a vector of orthogonal innovations estimated from the data. Furthermore, the analysis traces the dynamics of an innovation in any of the involved variables over time to account for the total amount of system variation attributable to each innovation.

More specifically, according to the Wold decomposition theorem, any finite linearly regular covariance stationary process $y(t)$, $m \times 1$, has a moving average representation

$$y(t) = \sum_{s=0}^{\infty} \Phi(s)\, u(t-s) \tag{8}$$

with $Var[u(t)] = \Sigma$.

Even although $u(t)$ is serially uncorrelated by construction, the components of $u(t)$ may be contemporaneously correlated, so an orthogonalizing transformation to $u(t)$ is done so that (8) can be rewritten as[1]

$$y(t) = \sum_{s=0}^{\infty} \Phi(s) P^{-1} P u(t-s) = \sum_{s=0}^{\infty} \Theta(s)\, w(t-s) \tag{9}$$

where $\Theta(s) = \Phi(s) P^{-1}$, $w(t-w) = Pu(t-s)$ and $Var[w(t)] = Var[Pu(t)] = I$

When P is taken to be lower triangular matrix, the coefficients of $\Theta(s)$ represent "responses to shocks or innovations" in particular variables. We can also allocate the variance of each element in y to sources in elements of w, since w is serially and contemporaneously uncorrelated. The orthogonalization provides

$$\sum_{s=0}^{T} \Theta^2 (s)_{ij} \tag{10}$$

which is the components of error variance in the $T+1$ step ahead forecast of y_i which is accounted for by innovations in y_j.

Finally, the study uses impulse response analysis to trace the observable responses of the system to shocks in the variables involved. This technique provides both a directional and quantitative measure of the system's response to each variable.

3. Empirical Analysis

3.1 Unit-root stationarity and cointegration results

The empirical analysis employs quarterly data[2] for the consumer price index (CP), the wage rate (WR), the nominal effective exchange rate (EX), the volume index of production in industry (Y), the productivity of labour adjusted per hour of work (PR) and the high-powered money (HM). All variables are expressed in logs and cover the period from 1976, 1st quarter to 1993, 2nd quarter.

Since cointegration requires nonstationary time series of the same order of integration, we firstly tested for unit-roots in the employed series using the augmented Dickey-Fuller (ADF) stationarity test. The results are presented in Table 1 and suggest

1 The representation (9) is obtained by decomposing Σ^{-1} as $\Sigma_T^{-1} = P'P$.

2 Data are collected from the Statistical Bulletin of Bank of Greece (various issues) and OECD: Main Economic Indicators (various issues). All empirical analysis employs MICROFIT 386 and RATS version 3.10 computer packages.

that all series are nonstationary in levels whereas they exhibit stationarity in first-differences.

Having established that the variables concerned are all I(1), we tested for common trends among the investigated economic series, by means of maximum likelihood multivariate cointegration technique developed by Johansen (1988 and 1989). The first step requires the postulation of a VAR model to obtain the long-run relationships among the examined series. The employed VAR system includes a dummy variable (DUMP) to account for the oil crisis effects and three seasonal dummies (S1, S2, S3). In order to specify the optimal number of lagged terms in the VAR, we adopted a strategy based on Sims (1980) likelihood ratio (LR) test. Thus, we calculated LR values for lag lengths varying from 1 to 6; the results indicated that a 3-lag VAR system was the appropriate specification. The examined lag-structures were also checked for serial correlation and heteroscedasticity.

The next step is to test for the existence of cointegrating vectors using LR tests based on the Maximal Eigenvalue and the Trace of the stochastic matrix Π. The results, reported in Table 2, suggest the existence of three cointegrating vectors at the 95% confidence level. Further examination of the signs and the graphs of the residuals in conjunction with DF unit root tests indicated the rejection of the first two cointegrating vectors.

Normalising the accepted third cointegrating vector on the price index, yields the following cointegrating equation

$$CP(t) = -2.092 + 0.396WR(t) + 0.467HM(t) + 1.053Y(t) - 0.278EX(t) - 2.479PR(t)$$

Having established the existence of joint integration properties among the investigated economic series we proceed with exploring the dynamic features of the variables and the directions of the detected causal relationships among them by means of variance decomposition and impulse response analysis.

3.2 Variance decomposition and impulse response analysis

In this section, first, we present the results derived from variance decomposition analysis. It should be noted that since the examined economic series exhibit cointegration properties, variance decomposition and impulse response analysis use the series in levels (Lutkepohl, 1991).

Table 3 presents the percentage of the variance of the T-quarter ahead forecast error of the variables that is attributable to each of the shocks for T = 1, 4, 8 and 20. We report only these four time horizons, since according to Blanchard and Watson (1986), they could be interpreted as the short-run (1 quarter ahead), the medium-run (4 or 8 quarters ahead) and the long-run (20 quarters ahead). The reported results refer

only to the prices equation and the wages equation and suggest that prices are explained by the course of wages, productivity and output. In particular, the effect of the wage rate on prices becomes significant at the 8th quarter being capable of explaining thereafter about 20% of the prices behaviour. Productivity causes significantly the price level after the 12th quarter, being capable to explain about 20% of the variance of prices while at the end of the 20th quarter it captures the 30%. Last, the results concerning the output variable show that it is the most significant determinant of prices being capable to explain a 20% of the variance at the 6th quarter, a 30% at the 8th quarter and about 35% by the end of the 20th quarter forecast horizon.

As it concerns the wage rate equation, according to the results in Table 3, the wages behaviour was found to be explained only by productivity and to a less degree by output. The price level explains only a 4% of the wage rate variance during the whole forecast period.

Next, we present the results from the impulse response analysis. Figures 1 and 2 plot impulse response functions-the 1 to 20 quarter response of the price level to wage rate shocks and of the wage rate to price level shocks. The numerical values calculated from the impulse analysis are reported in Table 4.

The overall evidence, emerging from both variance decomposition and impulse response analysis, is clearly in favour of the "mark-up pricing" view, that is, wage rate growth causes inflation to increase.

4. Concluding Remarks

This paper has investigated the causal relationship between wages and prices in Greece, using quarterly data for the period 1976:1 to 1993:2. We employed multivariate cointegration techniques in conjunction with variance decomposition and impulse response analysis. The results provided evidence of a significant causal effect running from wages to prices, thus supporting the Post-Keynesians' "mark-up pricing" view.

References

Barth, J. R. and Bennett, J. T., (1975). "Cost-Push Versus Demand-Pull Inflation: Some Empirical Evidence". Journal of Money, Credit and Banking, August 1975:391–397.

Blanchard, O. J. and Watson, M. W., (1986). "Are Business Cycles all Alike:", in Gordon R. J. (ed.) The American Business Cycles, NBER.

Darrat, A. F., (1994). "Wage Growth and the Inflationary Process: A Reexamination". Southern Economic Journal, 61:181–190.

Dickey, D. A. and Fuller, W. A., (1981). "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root". Econometrica, 49:1057–1072.

Engle, R. F. and Granger, C. W. J., (1987). "Cointegration and Error Correction: Representation, Estimation and Testing". Econometrica, 55:251–276.

Fosu, A. K. and Huq, M. S., (1988). "Price Inflation and Wage Inflation: A Causal-Effect Relationship". Economic Letters, 27:35–40.

Fuller, W. A., (1976). Introduction to Statistical Time Series. Wiley, New York.

Gordon, R. T., (1988). "The Role of Wages in the Inflationary Process". American Economic Review, May 1988, (Papers and Proceedings): 276–283.

Granger, C. W. J., (1986). "Developments in the Study of Cointegrated Economic Variables". Oxford Bulletin of Economics and Statistics, 48:213–228.

Johansen, S., (1988). "Statistical Analysis of Cointegration Vectors". Journal of Economic Dynamics and Control, 12:231–254.

Johansen, S., (1989). "Likelihood Based Inference on Cointegration". Lecture Notes for a Course on Cointegration, Seminario Estivo di Econometria, Centro Studi Sorelle Clarke, Bagni di Lucca, Italy.

Johansen, S. and Juselius, K., (1990). "Maximum Likelihood Estimation and Inference on Cointegration with Applications to the Demand for Money". Oxford Bulletin of Economics and Statistics, 52:169–210.

Katrakilidis, K. P., (1992). "The Long-run Relationship Between Price Inflation and Wage Inflation in Greece". Memorial volume, School of Forestry, University of Thessaloniki.

Lutkepohl, H., (1991). Introduction to Multiple Time Series Analysis. Springer-Verlag, New York.

Mehra, Y. P., (1977). "Money, Wages, Prices and Causality". Journal of Political Economy, December 1977:1227–1244.

Mehra, Y. P., (1991). "Wage Growth and the Inflationary Process: An Empirical Note". American Economic Review, September 1991:931–937.

Psaradakis, G. Z., (1991). "Modelling the Structure of Wage and Price Determination in Greece". Discussion Paper, No 91/2, University of Southampton.

Shannon, R. and Wallace, M. S., (1986). "Wage and Inflation: An Investigation into Causality". Journal of Post-Keynesian Economics, Winter 1986:182–191.

Sims, C., (1980). "Macroeconomics and Reality". Econometrica, 48:1–48.

Appendix

Table 1: Augmented Dickey–Fuller (ADF) Unit Root Tests

Variables	Levels (with trend)	First Differences (with trend)
CP	−1.264(2)	−5.295(2)
WR	−0.329(2)	−5.585(2)
HM	−0.673(2)	−7.378(2)
EX	−1.970(2)	−4.896(2)
PR	−2.264(2)	−14.551(2)
Y	−2.716(2)	−12.512(2)

Notes:

i) The ADF statistics were calculated with two lags (indicating in the parentheses) to ensure that the residuals were "white noise".

ii) The critical value from Fuller (1976), for the respective degrees of freedom and the 5% level of significance, is −3.477 for the trended case.

iii) CP, WR, HM, EX, PR, Y are logs of the consumer price index, wage rate, high-powered money, productivity of labour adjusted per hour and volume index of production in industry, respectively, over the period 1976:1 to 1993:2.

Table 2: Johansen Maximum Likelihood Tests for Cointegration

List of variables included in the cointegrating vector:
CP WR HM EX PR Y Intercept
List of additional I(0) variables included in the VAR:
S1 S2 S3 DUMP.
66 observations from 1977Q1 to 1993Q2. Maximum lag in VAR = 3.

Cointegration LR Test Based on Maximal Eigenvalue of the Stochastic Matrix.

Null	Alternative	Statistic	95% Crit. Value	90% Crit. Value
r = 0	r = 1	81.5197	40.3030	37.4480
r ≤ 1	r = 2	29.1011	34.4000	31.6640
r ≤ 2	r = 3	22.4420	28.1380	25.5590
r ≤ 3	r = 4	14.5513	22.0020	19.7660
r ≤ 4	r = 5	12.4387	15.6720	13.7520
r ≤ 5	r = 6	4.2298	9.2430	7.5250

Table 2: (continued)

Cointegration LR Test Based on Trace of the Stochastic Matrix

Null	Alternative	Statistic	95% Crit. Value	90% Crit. Value
r = 0	r ≥ 1	164.2827	102.1390	97.1780
r ≤ 1	r ≥ 2	82.7630	76.0690	71.8620
r ≤ 2	r ≥ 3	53.6619	53.1160	49.6180
r ≤ 3	r ≥ 4	31.2199	34.9100	32.0030
r ≤ 4	r ≥ 5	16.6685	19.9640	17.8520
r ≤ 5	r = 6	4.2298	9.2430	7.5250

Estimated Cointegrated Vectors (Normalized in Brackets)

	Vector 1	Vector 2	Vector 3
CP	-0.9735	-0.2321	-2.2935
	(-1.0000)	(-1.0000)	(-1.0000)
WR	1.4079	1.0177	0.9120
	(1.4462)	(4.5124)	(0.3968)
HM	-0.3909	-1.8845	1.0748
	(-0.4015)	(-8.1166)	(0.4676)
EX	0.4209	-1.7685	-0.6402
	(0.4324)	(-7.6168)	(-0.2785)
PR	-0.4408	5.3262	-5.6984
	(-0.4528)	(22.9397)	(-2.4792)
Y	2.1372	-6.7841	2.4205
	(2.1954)	(-29.2191)	(1.0531)
Intercept	-9.5244	44.8193	-4.8098
	(-9.7834)	(193.0350)	(-2.0926)

Table 3: Variance Decompositions

Variance Decomposition of CP:

	Percentage of variance of error due to innovations in					
	CP	WR	HM	EX	PR	Y
Forecast horizon						
1	100.00	0.000	0.000	0.000	0.000	0.000
4	81.653	3.638	1.611	1.982	1.907	9.206
8	30.661	18.065	2.649	5.411	13.102	30.110
20	8.391	20.389	1.327	3.152	31.555	35.183

Variance Decomposition of WR:

	Percentage of variance of error due to innovations in					
	CP	WR	HM	EX	PR	Y
Forecast horizon						
1	0.003	99.996	0.000	0.000	0.000	0.000
4	3.905	67.456	8.734	4.652	6.618	8.632
8	4.000	45.326	9.737	6.532	23.100	11.302
20	4.413	26.706	6.376	2.899	44.520	15.082

Table 4: Impulse Responses to One Standard Deviation Innovations (mult. 10^2)

Response of CP:

	CP	WR	HM	EX	PR	Y
Forecast horizon						
1	0.8320	0.0000	0.0000	0.0000	0.0000	0.0000
4	0.4296	0.2223	-0.1893	0.1889	0.1808	0.3199
8	0.0181	0.6599	-0.1315	0.3658	0.5480	0.8763
20	-0.382	0.7104	0.2557	-0.0959	1.0631	0.8662

Response of WR:

	CP	WR	HM	EX	PR	Y
Forecast horizon						
1	-0.0111	1.8159	0.0000	0.0000	0.0000	0.0000
4	-0.3826	1.0751	0.6151	0.6572	0.7099	0.7845
8	-0.2770	0.8792	0.6020	0.4786	1.3941	0.7479
20	-0.6460	0.8858	0.5000	-0.2377	1.5143	0.9104

Response of CP to One S.D. Innovations

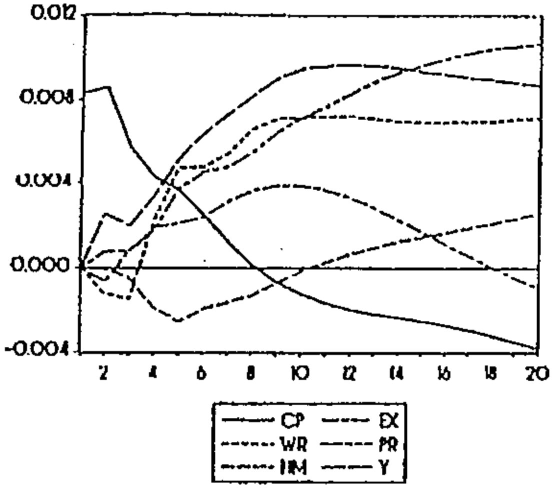

Figure 1

Response of WR to One S.D. Innovations

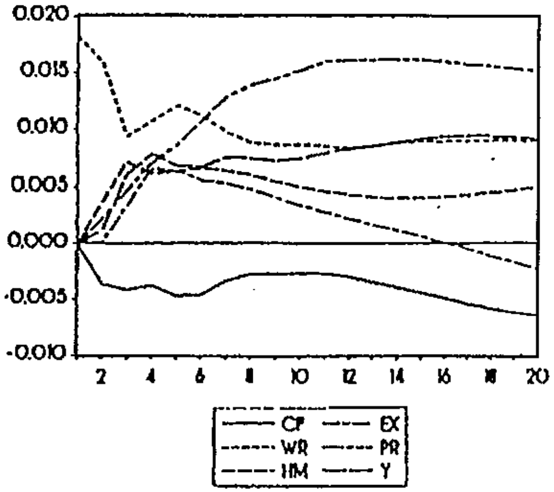

Figure 2