

\\99\\

**The Relationship(s) among Wages,
Prices, Unemployment and Productivity
in Italy**

di

Marcello D'Amato e Barbara Pistoresi

Marzo 1994

Dipartimento di Economia Politica
Viale Berengario, 51
41100 Modena (Italia)

***The relationship(s) among wages, prices,
unemployment and productivity
in Italy***

Marcello D'Amato* - Barbara Pistoresi**

(J.E.L C32,J23)

Abstract: In this work we ask whether a long run Phillips Curve is supported by Italian data. This is performed via multivariate cointegration analysis as in Johansen (1988) and Johansen and Juselius (1990) using both annual and quarterly data. A clear-cut result of cointegration, describing a long run relationship among wage and price inflation, productivity growth and unemployment rate, is obtained. The presence of a trade off between nominal wage inflation and unemployment rate crucially depends on the spanning of the sample. A standard Phillips Curve with money illusion is found on annual data covering a period from 1960 to 1990. A less clear-cut result is obtained on quarterly data spanning from 1970Q1 to 1992Q2: the test for no money illusion can be just accepted on the border of 5% significance level. This implies that a trade off between *real* wage inflation and unemployment can not be excluded.

* Dipartimento di Scienze Economiche, Università' di Salerno, Via Ponte Don Melillo, Fisciano, Salerno
** Dipartimento di Economia Politica, Università' di Modena, Via Berengario 51, Modena.

1. Introduction¹

Phillips, in a well-known paper published in 1958, found a clear non-linear inverse relationship between the change in money wages and the unemployment on an estimation period from 1861 to 1957². Moreover, he found that before the World War II the data describe counterclockwise loops and lie above the fitted curve when unemployment rate is falling and inflation is rising (boom), while after the World War II the loops become clockwise. The reason he put forward was that, for a given rate of unemployment, employers would bid more vigorously for labour when unemployment is falling than it is rising.

The empirical Phillips' relationship was elaborated by Lipsey (1960). Lipsey was the first to linearise the Phillips Curve and extended it by adding further variables. Lipsey's single structural equation becomes a stochastic difference equation with the change in the nominal wage as the dependent variable on the LHS and both the levels and changes in unemployment (to explain the loops) and in consumer prices as independent variables on the RHS. Price inflation was introduced to take into account the cost of living and Lipsey found that it improved the fit. He found the negative trade-off, unemployment-inflation, and that the coefficient of price change was significantly less than one (money illusion).

Policy consequences of the Phillips-Lipsey framework was that if a stable negative relation exists and if a social welfare function, defined over the two variables inflation and unemployment, could be chosen then it would be possible to choose an optimal point on the Phillips Curve representing the optimal decision available to the policy makers.

The Phillips-Lipsey arguments are opened to two objections. One is that wages and prices should be estimated in a multivariate model. This point was recognised by Phillips (1959) himself and by Samuelson-Solow (1960). Phillips estimated two equations for 1947-1958 Australian annual data, one in the wage changes and one in the price changes and wrote: "Changes in wage rate are influenced by earlier changes in consumer prices, but changes in consumer prices are themselves largely determined by earlier changes in wage rates".

The other and more important stated by Friedman (1968) regards the size of the parameter of money illusion. In his own words "There is always a temporary trade-off between inflation and unemployment, there is no permanent trade-off. The trade-off comes not from inflation per-se, but from unanticipated inflation". This theoretical point of view has an important impact on the debate on the effectiveness of the intervention policy to reduce the unemployment. Friedman's point is simple: in the short run there exist money illusion and the possibility to expand the aggregate demand reducing the unemployment, but in the long run, when the agents adjust the adaptive expectations, the

¹We would like to thank Mark Stewart for helpful criticisms on the first draft. The authors are also grateful to Adalgiso Amendola, Mario Biagioli and Mario Forni for suggestions and comments on the final version of this work. The usual disclaimers apply.

²A variant of this relation was investigated by Fisher (1926). Two contemporary papers at the Phillips's one are: Dicks, Mireaux and Dow (1959) and Klein and Ball (1959).

aggregate demand cannot be increased, unless an acceleration of the inflationary process. Thus the long run Phillips Curve becomes vertical.

Friedman's position paved the way for the rational-expectation model of the supply side in macroeconomics. In a context of maximising agent over an infinite horizon, the Phillips Curve becomes vertical and the monetary intervention policy becomes *neutral* also in the short-run.

From an empirical point of view, the negatively sloped short-run Phillips Curve collapsed in the seventies, but during the past decade was supported by the data again. The theoretical background becomes a context of welfare maximising behaviour consistent with temporary departure from equilibrium (Ball, Mankiw and Romer 1988). It is possible if, for example, costs of renegotiating contracts are higher than the costs of disequilibrium and/or the contracts are staggered (New-Keynesian approach ³).

In this work we perform a purely statistical analysis of the Phillips Curve. We are not interested in an empirical rehabilitation of a *short run* curve. Instead, our aim is to use multivariate cointegration analysis to test if a *long run* Phillips Curve is supported by Italian data. The short run dynamics towards the long run target, represented by the estimated cointegration vector(s), is also analysed. This approach (Johansen 1988, Johansen-Juselius 1990) permits to avoid the simultaneity problem arising from usual estimation of a single equation and the problem of non standard distribution of the statistics arising from testing regressions performed on non stationary variables. This makes a first improvement on similar studies on Italian data such as the Phillips type equations estimate by Onofri and Salituro (1985) and Zenezini (1986).

Four variables are *a priori* selected for possibly entering the final specification of the model: nominal wages and prices inflation, unemployment rate and labour productivity growth. The choice of variable changes rather than levels will be made both on theoretical and statistical ground (degree of integration).

The introduction of the *inflation rate* (with a positive coefficient) into the theoretical Phillips equation represents the hypothesis that the wage-earners are sensitive to the effects of inflation on their purchasing power.

Growth of productivity may strengthen union demands and increase the willingness of firms to satisfy these demands. Labour productivity or labour productivity growth only occasionally is used in empirical work as an explanatory variable in the specification of the Phillips Curve because this variable could be considered as an alternative to the unemployment rate: if average productivity is a good substitute for marginal productivity it should be a prime determinant of labour demand. However, we consider that the marginal productivity primarily affects the demand side but we don't want to restrict a priori the supply side of the labour market to the no money illusion hypothesis because, for example, bargaining procedure, asymmetric information about marginal productivity etc. may induce rational labour supplier to take into account unemployment.

Vanderkamp (1972), for example, includes both productivity and unemployment as explanatory variables. He argues that because of labour hoarding and slow adjustment to equilibrium by firms, productivity can be used together with unemployment or vacancies to provide an improved proxy for net excess of demand. The idea is that when demand falls, output falls but workers are not fired at least initially. Productivity falls as

³ See Blanchard (1983), Fischer (1977) and Taylor (1979).

well and the opposite occurs during recovery. Hence, average productivity is a proxy for the amount of hoarded labour⁴.

Other variables appear in several studies: *profits* and *union power*. Two are the justifications to include profits in the Phillips specification: one is that a high rate of profit will be followed by a fall of the redistribution in favour of workers the other is that in period of high profitability the opportunity cost of strikes is higher for employers who will be inclined to concede more to wage demands. However, profits could be considered a proxy for productivity introducing in our case a collinearity problem (Kuh 1967). The justification for including the union power is that an increase in the militancy should bring increase in the wage rate. However, Ashenfelter and Johnson (1972) argue that it exists a causality from wages to militancy and not vice versa. In particular they find that union power is not significant in explaining wages. Moreover, as stressed by Tobin (1972) the level of union power can explain the level of wage rate but not the wage inflation. Hence, we doubt the importance of the inclusion of this variable in the selected specification.

In Section 2 multivariate cointegration approach is presented. In Section 3 univariate integration analysis to test the stationarity of nominal wage, consumer prices, unemployment rate and labour productivity is performed via Dickey Fuller tests. The sample period is 1960-1990 and the data are annual. Section 4 presents tests for stationarity performed on the estimated cointegration vectors as a more efficient alternative to the univariate integration analysis. Some linear restrictions of economic interest on the estimated cointegration vectors are finally performed and discussed. In Section 5 the same exercise is repeated on a shorter quarterly data set: 1970Q1 - 1992Q2. Section 6 presents the short run dynamics estimated on the reparameterised ECM version of the original VAR, under the hypothesis of cointegration. The presence (and direction) of the loops is tested on the coefficients of lagged change of unemployment in the short run equation. Nominal wage flexibility is tested by introducing partitioned (negative and positive) cointegration residuals in the error correction specification of the wage equation. Section 7 concludes.

Hence, identification and specification of long run equations is achieved by linear restrictions tested on long run parameters. This allows to compare the empirical relevance of different theories about the role of the selected variables in the long run equilibrium in the labour market. In particular, the main role played by unemployment rate in defining the long run target for the wage inflation is interpreted as evidence against insider-outsider models and in favour of bargaining models in the tradition of Phillips' arguments.

2. The analysis of long run relationships

In this section the basic concepts of multivariate cointegration analysis, estimation and testing of long run relationships are discussed. This is done estimating the cointegration space as in Johansen (1988) and Johansen and Juselius (1990) and then testing more specific hypotheses of economic interest within this space as presented in Johansen and Juselius (1992). The *stationarity testing in a multivariate framework* is

⁴The impact of productivity is also shown in Simler and Tella (1968) and Kuth (1967).

also performed. This is more efficient than the usual univariate one, because a multivariate model is likely to give lower residuals variance. Moreover, the multivariate stationarity testing is different from the univariate one by formulating the null of stationarity instead of non stationarity, as it is the case in the univariate test. In particular, this procedure is performed by testing the stationarity of the series both jointly and separately on the estimated cointegration vector(s). If stationarity is rejected it will be possible to analyse whether these variables are cointegrated or not. Instead, if some variables result to be stationary, it is possible to exclude them and to test for a cointegration subset. So we will also consider stationarity tests in the multivariate model as an *exclusion test* for the variable of interests. This, as well as a priori consideration on the nature of the long run relationship, will have some consequences on the specification of the final equation.

The cointegration analysis is carried out for the observational variable vector $z'_t = (\Delta w_t, \Delta p_t, u_t, \Delta q_t)$ where Δw is the first differences of logged nominal wage, Δp is the first differences of logged consumer price, u is logged unemployment rate and Δq is logged labour productivity growth. We insert the wage and price inflation separately instead of real wage dynamics, because we don't want to impose a sort of *a priori* "no-money-illusion" condition as it is currently preferred (see for example Hall(1986) and Hall(1989), Alexander (1993) and Darby and Wren-Lewis (1993)). Money illusion will be tested on the data as well as other hypothesis.⁵

The estimation period is 1960-1990 for the cointegration analysis on annual data (Source: OECD, Main Economic Indicators) while is 1970Q1- 1992Q2 for the analysis on quarterly data (Source: OECD, Main Economic Indicators)

The long run relationships are estimated jointly with the short-run dynamics (by applying the M-L procedure) thus using all the information available in the data. This is the main difference with respect to Engle and Granger (1987), Stock (1987), Phillips (1987) *two step procedure* and implies more efficient estimates.

In the two step procedure, the cointegration relation is estimated OLS by a static regression and in the second step the cointegration residuals are tested for stationarity. If stationarity is accepted, residuals are included as an Error Correction Terms in the final ECM model. The main problem of this procedure concerns the estimated long-run parameters. If the variables are cointegrated, OLS static regression yields superconsistent estimates of the long-run responses of the model, but the same cannot be said of the estimated standard errors of the regression. This is because the distribution of the OLS estimators is generally non-standard (Phillips 1988). Hence, it is not possible to test the significance of the estimated coefficients and perform long-run restrictions of economic interest on the static regression. This problem is avoided by applying the multivariate procedure.

The model to be estimated is the reparametrisation of a VAR in the ECM form

⁵For a discussion addressing the issue of expectations underlying the VAR specification refer to Juselius (1991) p.4. The intuition is that a reparameterised VAR in the ECM form requires that endogenous variables always adjust to the long run target. This implies that if, for example, in the wage equation, a price inflation coefficient less than one is found, agents find too costly fully adjusting wage inflation to price inflation. This may be due either to the impossibility of continuous recontracting wages, or to the high costs for the economy of a full indexation that is taken into account by bargaining agents.

$$(1) \Delta z_t = \mu + \sum_{i=1}^{k-1} \Gamma_i \Delta z_{t-i} + \Pi z_{t-k} + Bx_t + u_t$$

where Γ_i and Π are $n \times n$ matrices of unknown parameters, Δz_t and Δz_{t-i} are $n \times 1$ vectors of I(0) variables, while the z_{t-k} is a $n \times 1$ vector of I(1) variables. B is an $m \times s$ matrix and x_t is a $s \times 1$ vector of I(0) variables or dummies. μ is a drift parameter capturing the trend in the series, while u_t is a vector of Gaussian disturbances.

In order to estimate (1) by Johansen's (1988) procedure difference stationarity of the involved variables must be accepted by univariate and multivariate stationarity tests. Then, assuming that the vector z_t is I(1), we need to test the hypothesis of reduced rank of the Π matrix

$$H_0: \Pi = \alpha \beta'$$

where the adjustment coefficient matrix α is $(n \times r)$ and β is the $(r \times n)$ matrix of cointegration vectors. r is the rank of Π and determines the number of linearly independent stationary relations between the levels of the variables, i.e. the number of cointegration vectors in the data. If the rank is zero there is no cointegration, while if Π is of full rank, then z_t is stationary. Clearly if the rank is $r < n$, we have r cointegration vectors.

3. Univariate integration analysis using annual data

The necessary condition to perform the Johansen procedure is that all the variables in the VAR are I(1), as mentioned above. Hence, we test, via the Dickey Fuller tests, the stochastic non-stationarity of the series. Table 1 reports the results of integration tests for the basic set of logged variables: nominal wage, consumer prices, unemployment rate and labour productivity. Following Perron (1988) we initially include a trend in the testing regressions. Excluding a relevant trend, may give a bias, making rejection of the null hypothesis of non-stationarity unlikely. However if an unnecessary time trend is included a loss of power may occur. To guard against this, an inability to reject the null hypothesis is always investigated further by performing an additional test omitting the trend. The number of augmentation used in the ADF tests to avoid the autocorrelation of the residuals are reported in brackets in Table 1. In general one lag is enough given the low dynamics of annual data.

Unit root tests which incorporate structural breaks in the series are available (Perron 1989). These may be useful because often the presence of unit root is only due to the change in the slope or/and in the intercept of the time trend (Rappoport -Reichlin 1989, Perron 1993). However, we have insufficient observations here to perform these tests.

Fig. 1, 2, 3, 4 plot the changes of the logged series. From the plots we can deduce that nominal wage and price inflation are not stationary while the change in unemployment rate seems to be stationary. The stationarity of the first difference of labour productivity exhibit a negative trend. Thus, from a simple inspection of the plots, the nominal variables seem to be I(2), while the unemployment rate seem to follow an I(1) process. The labour productivity could be I(1) too, but the evidence of one (versus two) unit root is not so clear.

Fig.1

Fig.2

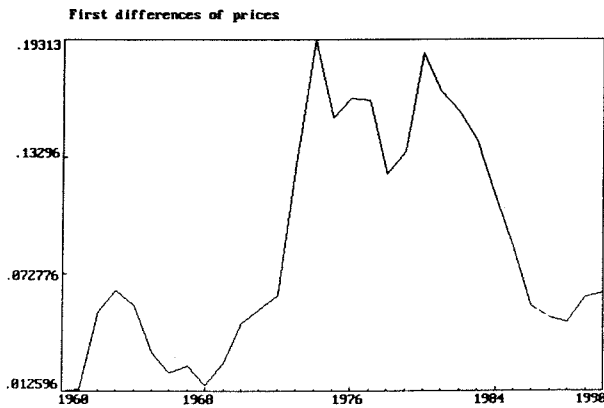
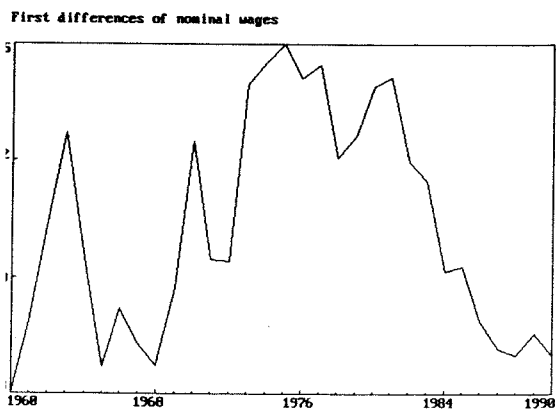


Fig.3

Fig.4

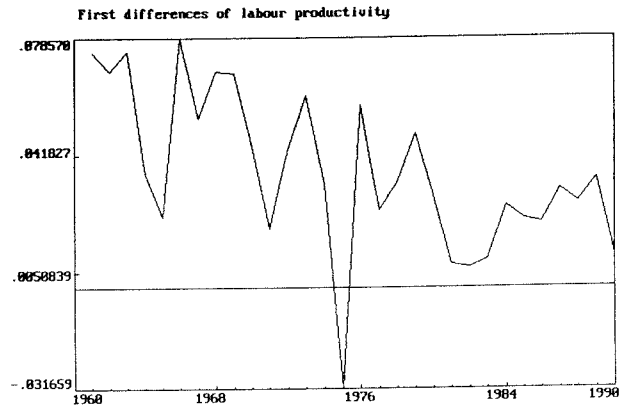
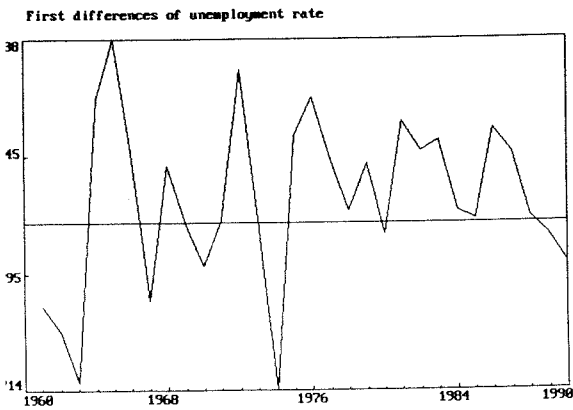


Table 1

Degree of integration: Dickey Fuller and Augmented Dikey Fuller tests

	<i>Including trend</i>				<i>No trend</i>	
	DF	ADF	Φ_3	Φ_2	DF	ADF
p	-2.31	-2.50 (1)	3.31	3.23	1.80	-1.13 (1)
w	-1.24	-1.67 (1)	2.04	2.68	0.17	-1.28 (1)
u	-3.61	-3.61 (0)	7.23		-0.12	-0.12 (0)
q	-2.04	-2.04 (0)	10.78		-4.50	-4.50 (0)
Δp	-1.14	-1.14 (0)	1.51		-1.67	-1.67 (0)
Δw	-1.87	-1.87 (0)	2.67		-2.03	-2.03
Δu	-4.27	-4.27 (0)	9.34		-4.37	-4.37 (0)
Δq	-5.21	-5.21 (0)	13.62		-3.88	-3.88 (0)
$\Delta\Delta p$	-4.62	-4.62 (0)	10.75		-4.48	-4.48 (0)
$\Delta\Delta w$	-5.12	-5.12 (0)	13.17		-5.03	-5.03 (0)

Note: The DF regression with trend is $y_t = \mu + \tau + \alpha y_{t-1}$ where t is the time trend, while the ADF(1) regression is $y_t = \mu + \tau + \alpha y_{t-1} + \Delta y_{t-1}$. The critical values from Dickey - Fuller (1976, 1981).

$\tau_\tau (H_0: \alpha = 1)$ in model with trend: -3.80 (2.5%), -3.50 (5%), -3.18 (10%)

$\Phi_3 (H_0: \alpha = 1, \tau = 0)$ in model with trend: 7.81 (2.5%), 6.73 (5%), 5.61 (10%)

$\Phi_2 (H_0: \alpha = 1, \tau = 0, \mu = 0)$ in model with trend: 5.94 (2.5%), 5.13 (5%), 4.31 (10%)

$\tau_\mu (H_0: \alpha = 1)$ in model without trend: -3.22 (2.5%), -2.92 (5%), -2.60 (10%)

Performing the unit root tests on logged nominal wage (w) and logged consumer price (p) it is not possible to reject the hypothesis that these variables are $I(2)$ ⁶. There is a clear-cut result of unit root in the levels and in the first difference of these variables, while the second differences appear to be stationary.

Logged unemployment rate (u_t) yields a less clear result. It is non stationary at 2.5% critical value, if the model with trend is considered, while the unit root in the series, in the model without trend, is out of doubt. The first difference of unemployment rate is stationary. This result is the same in Hall (1986) for UK data. A quick review of similar studies has shown similar controversial results⁷. Despite the statistical findings, on the

⁶See Hall (1986, 1989) and Alogouskufis- Smith (1991).

⁷A visual inspection of the series shows as the logged unemployment rate has the same pattern of the GDP and so of the labour productivity with a distinguished drop in 1973-74, where structural break could be identified. Hence, a problem of segmented trend rises as in the case of DGP and unemployment rate. In this case the unit root could be emerge as a spurious result of the segmentation of the time trend (an analysis of unit root in the Italian DGP is in Caselli and Marinelli 1993). Blanchard and Fisher (1988) refer to the unemployment series as stationary, but note 16 p.38 stresses as the evidence is not clear cut depending on the characteristics of the single economies and on the sample. Nymoen (1989), testing integration on Norwegian quarterly data, concludes that the null of non stationary is rejected, although

ground that a deterministic trend in the unemployment rate is difficult to conceive from a theoretical point of view, we regard unemployment as a $I(1)$ process.

Logged labour productivity (q) is non stationary in the more appropriate model with trend while the first differences of the variable appear to be stationary. Hence, univariate Dickey-Fuller test suggests a random walk with drift process. However, from the multivariate integration analysis the stationarity of the first difference of productivity is rejected. The more efficient multivariate outcome will be preferred.

These result implies that our selected specification of the Phillips type equation is in inflation of wages and prices⁸. This is not the common specification in the current literature (see note 2), the usual specification being in level.

A possible rationale for the choice of levels ($I(2)$) of nominal variables could be subset cointegration: as stressed in Hylleberg and Mizon (1990) p. 116, when the dependent variables is $I(1)$, "the regressor must include some $I(1)$ variables or a combination of variables of higher order of integration which are cointegrated (to $I(1)$). If in such a case some of the regressor are $I(0)$, these variables will affect the short run behaviour of the dependent variable only. Furthermore there can be no deterministic trend variables as a regressor in such cases" . This argument gives some rationale for the specification adopted by Hall (1986), where the final selected long run equation has an $I(2)$ wage level as a dependent variable that cointegrates with another $I(2)$ variable, prices, among regressors, all other regressors being $I(1)$. But the way in which Hall finds (subset) cointegration between nominal wage rates and prices is not the usual one: he computes the difference in logs and apply the DF and ADF to the resulting real wage variable which ends to be $I(1)$. Italian annual data accept the same test, but strongly reject the cointegration computed in the proper way, i.e. by testing subset cointegration on a multivariate VAR, between wages and prices inflation. For this reason it is difficult to accept Hall's procedure to specify the Phillips relation on the sample at hand.

4. Results from cointegration analysis using annual data

The lag length of (1) that ensures residuals approximately white noise normal is estimated before applying Johansen procedure. Table 2 shows diagnostics relating to the specification of a VAR(1). Normality Jarque-Bera tests, Godfrey's test of residual serial correlation and the tests for heteroskedasticity are derived from the analysis of the residuals of each equation of the system. These tests shows that one lag is enough to whiten the residuals in each of the equations.

not very comfortably. Instead, Alexander (1993), on quarterly UK data and Darby and Lewis (1993) on annual UK data, accept the unit root in the series of unemployment rate. Evidence of rejections of the unit root with historical data from many countries can be found in Raj (1992), in particular if the unemployment rate is considered in Mitchell (1993).

⁸This will presumably loose the dynamic properties of the short run ECM interpreted as a structural form for adjustment in wages or prices.

Table 2
Diagnostics for a VAR(1)

Equations	Godfrey's test p-value	Jarque-Bera test p-value	Heteroskedasticity p-value
<i>Wage equat.</i>	0.29	0.50	0.96
<i>Price equat.</i>	0.11	0.62	0.15
<i>Unempl. equat.</i>	0.53	0.73	0.21
<i>Prod. equat.</i>	0.16	0.95	0.012

Note: Godfrey's test of residual serial correlation distributed as an F(1,22)

Jarque-Bera's test of the normality of residuals based on skewness and kurtosis, distributed as a $\chi^2(2)$

Heteroskedasticity test (F-version) based on the regression of squared residuals on squared fitted values, distributed as an F(1, 26).

p-value: probability error of type I

In the case of a VAR(1) and no dummies, the model to be estimated becomes

$$(2) \Delta z_t = \mu + \Pi z_{t-1} + u_t$$

The tests of the hypothesis of reduced rank of Π matrix enable us to accept $r=2$. In fact, both the maximal eigenvalue and trace tests are consistent, at 95% significance level, with the presence of two cointegration vectors (Table 3). The estimated coefficients of these vectors and the associated adjustment coefficients to the long run (dis)-equilibrium are presented in Table 4. The clear-cut result of cointegration according to both the tests means that the cointegration relationships are strong and confirm that the speed of adjustment to the long run is not slow (see also the estimated adjustment coefficients in Table 4). The cointegration residuals are shown in Fig. 5.

When the dimension of the cointegration space is larger than one, a problem of direct economic interpretation may arise (Table 4). It is possible to identify some relationships, for example a wage equation or a price equation, by imposing linear restrictions on the unemployment or labour productivity equation. At first sight, the first vector may reflect a wage equation describing the negative relation of the nominal wage inflation with unemployment rate, the positive relation with price inflation and the negative relation with labour productivity growth (efficiency effect (Alexander 1993): greater efficiency of the workforce could increase unemployment and therefore decrease the wage inflation, via a reduction of wage claims). If the restriction of no-money illusion in the changes of nominal wages and prices will be accepted all these relationships can be regarded as explaining the real wage dynamics. The second cointegration vector, normalised with respect to price inflation, may be a mark-up equation in which, with wages adjusted for labour productivity, and with a possibly small, even zero, coefficient of unemployment. Whether the latter coefficient is either zero or not can be used to test the hypothesis of cyclicity of mark-up. (See. Blanchard and Fisher (1989) p.543).

The estimated matrix $\Pi = \alpha\beta'$ (Table 5), which is the matrix of long run responses (Hylleberg and Mizon (1989)), gives information regarding the long run behaviour of the system. This matrix describes the overall dynamics of the model due to long run disequilibria, under the null of cointegration.

The signs on the main diagonal are all negative indicating for all the variables that an error correction is working with different strength: the nominal wage inflation is the more flexible variable followed by the price inflation and the unemployment rate. Since the coefficient of productivity growth is -1.13, it should be tested if it is not significantly different to one. In such a case long run disequilibria are not permanent at all, i.e. the productivity growth doesn't depend at all, *in the system*, on its past values. Inflation seems to react principally to wages disequilibrium in the sense predicted by the Phillips Curve with a coefficient of 0.47. The opposite relation holds more strongly (0.64) suggesting that the wage-price spiral mechanism is at work. The low value of the parameter linking unemployment variation to the past value in the level (-0.12) suggests a high value of shocks permanence in this variable (a degree of hysteresis in the system).

Table 3

Johansen Maximum Likelihood Procedure (Trended variables, Trend in DGP) for a VAR(1)

H_0	H_1	Statistic	95% CV
<i>Cointegration LR test based on maximal eigenvalue of the stochastic matrix</i>			
$r = 0$	$r = 1$	51.09	27.96
$r \leq 1$	$r = 2$	24.92	20.96
$r \leq 2$	$r = 3$	3.87	14.06
<i>Cointegration LR test based on trace of the stochastic matrix</i>			
$r = 0$	$r \geq 1$	81.85	47.21
$r \leq 1$	$r \geq 2$	30.76	29.68
$r \leq 2$	$r \geq 3$	5.83	15.41

Note: Critical Values in Osterwald-Lenum (1992)

Table 4
 Estimated values of cointegration vectors and related adjustment coefficient

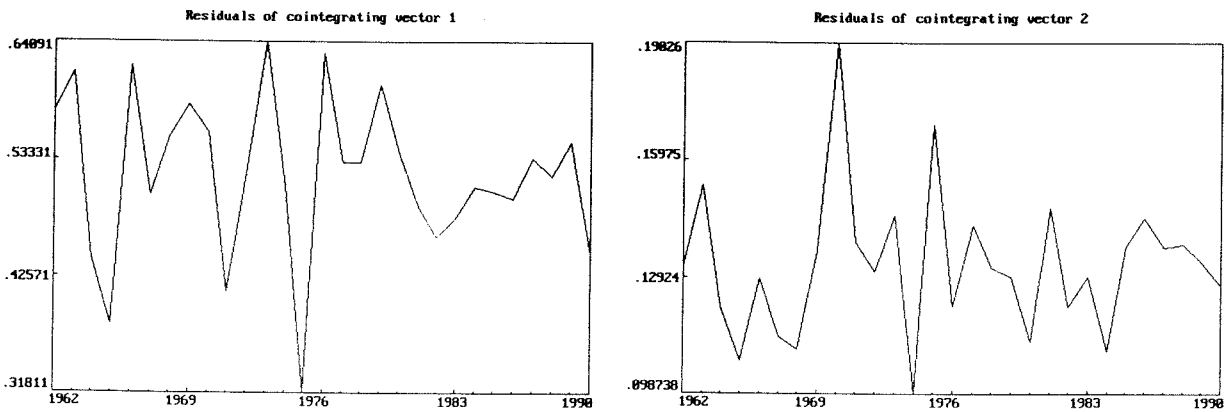
Variables	Vector 1	Vector2	Adjustment coeff. vector1	Adjustment coeff. vector2
Δw	- 1	- 1	-0.082	0.74
Δp	0.29	0.90	- 0.13	- 0.33
u	- 0.14	- 0.049	0.75	0.22
Δq	- 3.49	0.33	0.30	-0.25

Note: the estimated cointegration vectors and the relative adjustment coefficient vectors presented above are normalised with respect to the nominal wage inflation.

Table 5
 Estimated restricted ($r=2$) $\Pi = \alpha\beta'$

ECM	$\Delta w(-1)$	$\Delta p(-1)$	$u(-1)$	$\Delta q(-1)$
$\Delta\Delta w$	-0.66	0.64	-0.02	0.53
$\Delta\Delta p$	0.47	-0.34	0.03	0.35
Δu	-0.97	0.42	-0.12	-2.56
$\Delta\Delta q$	0.05	-0.13	-0.03	-1.13

Fig. 5



We perform, now, some tests (exclusion restrictions) both for stationarity and subset cointegration. As referred above test for stationarity can be performed by imposing proper linear restrictions to cointegration parameters. This procedure is more efficient than the standard univariate one.

The hypothesis to test are the following ones

Restricted cointegration vector: testing stationarity on $\hat{z}_t = (\Delta\hat{w}_t, \Delta\hat{p}_t, \hat{u}_t, \Delta\hat{q}_t)$

$H_01(0, 0, -1, 0) 2 Sp(\beta)$	LR(2) = 22.90
$H_02(0, 0, 0, -1) 2 Sp(\beta)$	LR(2) = 19.05
$H_03(-1, 0, 0, 0) 2 Sp(\beta)$	LR(2) = 19.80
$H_04(0, 1, 0, 0) 2 Sp(\beta)$	LR(2) = 21.27

The likelihood ratio tests, asymptotically distributed as a χ^2 with the appropriate degree of freedom given in brackets, indicate that the stationarity is rejected in all the cases. H1, H3 and H4 confirm the univariate integration analysis, i.e. the unemployment rate, wage inflation and price inflation are non stationary, while H2 reverses the result obtained by the univariate analysis, stating that the changes in labour productivity are non stationary. This justifies, on statistical grounds, the inclusion of the differences of labour productivity in the estimation of the VAR.

The test for the suggested interpretation of cointegration vectors as defined above will be performed by testing H5

$$H_05(-1, 1, 0, 1) 2Sp(\beta) \quad LR(2)=5.45 (pv=0.066)$$

This hypothesis on the cointegration vector interpreted as a roughly constant mark-up equation is accepted by the LR test.⁹ This means that a consistent long run price equation is: $\Delta p = \Delta w - \Delta q$. Under this restriction the estimated second long run equation of the system, interpreted as a long run Phillips curve, becomes: $\Delta w = 0.59\Delta p - 0.093u - 1.59\Delta q$. The coefficient of Δp is rather low suggesting that the widespread (see Hall (1986) and (1989), Alexander (1993), and Darby and Wren-Lewis (1993) on UK data) a priori hypothesis of no money illusion is rejected. This could be explained in terms of the accommodation of the monetary authority to wage shocks in the 60's and 70's and as evidence that full wage indexation only lasted for some time during

⁹ A possible objection would be that this hypothesis (weakly) accepts a real wage equation around the productivity growth and no long run trade off. But this contrasts with the evidence that the consequent coefficient of wages in the price equation would be larger than 1 (1.59), and furthermore price inflation would positively depend on unemployment and productivity. Another argument comes from the remarkable long run trade off estimated by the same procedure on real wage inflation, unemployment and productivity growth, under the a priori hypothesis that the endogenous variable in the system would be the real wage (no money illusion by wage setters). We argue that statistical evidence for a long run trade off under no money illusion, a fortiori, implies a long run trade off under money illusion. It has to be stressed that a constant mark-up in the long run price equation doesn't prevent a cyclical mark-up that should be tested on the ECM

70s and 80s. The negative coefficient of productivity growth may track an efficiency effect.

Solving for wage inflation in the restricted vectors and assuming an average productivity growth of 3% per year (sample mean), the rate of unemployment, corresponding to a wage inflation of 5% per year (price inflation of around 2%), is around 10%. A long run wage inflation of 8% (corresponding to a 5% price inflation), the non accelerating inflation rate of unemployment is around 9%.¹⁰

5. The analysis of long run relationships using quarterly data¹¹

In our analysis of annual data we found a negatively sloped long run Phillips Curve. We now turn to a quarterly data set spanning over a shorter and more recent period: 1970Q1 - 1992Q2. The series considered in the VAR are the same as before. Nominal wage, consumer prices, unemployment rate retain the same order of integration as before, while the productivity index appears to be I(1), see Table 6 and the plots of differences, Figg.6, 7, 8, 9, and stationarity testing in the multivariate framework.

Some doubts arise from univariate analysis of productivity levels (q): ADF with trend gives I(0) around trend, while ADF without trend gives I(1). The choice between the two options can be based on the procedure outlined in Perron(1988), and this would induce to accept the hypothesis of non stationarity (Φ_3), in particular, by rejecting Φ_2 (see the note in Table 6), the test yields a unit root process with drift as an outcome. Multivariate analysis, that has already been seen to be more efficient, will definitely confirm the I(1) nature of the process. The same argument can be made for wage inflation.

¹⁰Layard et al (1991) p. 437 prefer to simulate Nairu estimates on subsample periods rather than solving for the long run equilibrium the estimated unemployment equation. This is done under the assumption of a changing equilibrium unemployment rate, anyway if the cointegrating relations found here track a stable long run equilibrium, it is possible to solve for a long run equilibrium unemployment rate considering it as a broad measure valid over the sample period. For a similar assumption cfr. Juselius (1991) p.13.

¹¹Source: OECD. Wages and prices are expressed in index number: 1985=1.

Table 6
Degree of integration: Dickey Fuller and Augmented Dickey Fuller tests

	<i>Including trend</i>				<i>No trend</i>	
	DF	ADF	Φ_3	Φ_2	DF	ADF
w	1.46	0.42(1)	4.73	8.15	-5.10	-2.83(1)
p	2.37	-0.43(2)	2.52	3.08	-3.46	-2.25(2)
u	-2.13	-3.11(3)	5.08	3.95	-1.22	-1.39(3)
q	-2.92	-3.53(2)	6.55	7.62	-1.45	-1.26(2)
Δw	-5.58	-5.58(0)	15.61		-4.46	-4.46(0)
Δp	-3.49	-3.49(0)	6.73		-2.91	-2.91(0)
Δu	-10.78	-4.04(2)	8.20	5.46	-10.77	-4.01(2)
Δq	-7.42	-6.19(1)	19.18	12.82	-7.34	-6.16(1)
$\Delta\Delta w$	-11.28	-7.28(3)	26.61	17.74	-11.33	-7.29(3)
$\Delta\Delta p$	-12.74	-12.74(0)	81.27		-12.69	-12.69(0)

Note: The DF regression with trend is $y_t = \mu + \tau + \alpha y_{t-1}$ where t is the time trend, while the ADF(1) regression is $y_t = \mu + \tau + \alpha y_{t-1} + \Delta y_{t-1}$. The critical values from Dickey - Fuller (1976, 1981).

$\tau_\tau (H_0: \alpha = 1)$ in model with trend: -3.80 (2.5%), -3.50 (5%), -3.18 (10%)

$\Phi_3 (H_0: \alpha = 1, \tau = 0)$ in model with trend: 7.81 (2.5%), 6.73 (5%), 5.61 (10%)

$\Phi_2 (H_0: \alpha = 1, \tau = 0, \mu = 0)$ in model with trend: 5.94 (2.5%), 5.13 (5%), 4.31 (10%)

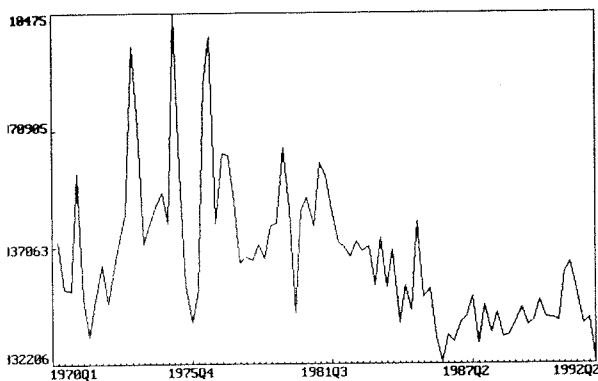
$\tau_\mu (H_0: \alpha = 1)$ in model without trend: -3.22 (2.5%), -2.92 (5%), -2.60 (10%)

(.): number of augmentation used in the ADF regression.

Fig. 6

Fig. 7

First differences of nominal wages



First differences of prices

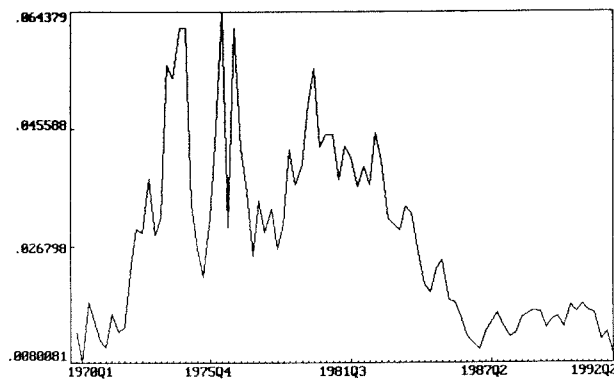


Fig. 8

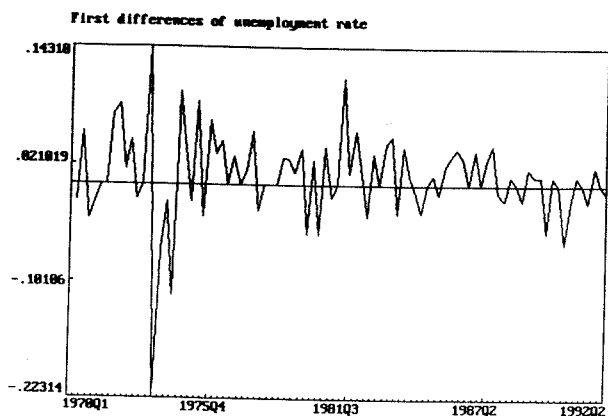
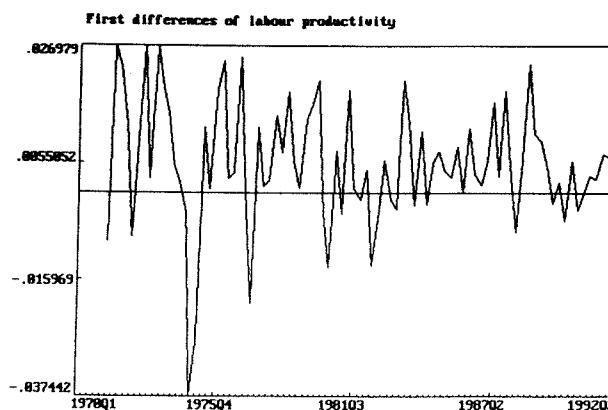


Fig. 9



The I(1) productivity index introduces the issue of the specification of the VAR. Two options are possible: either to insert the productivity level as a proxy for technological change explaining real wage dynamics or inserting the change in productivity of labour and testing its stationarity directly on the estimated cointegration vector(s). In the latter case, if stationarity is accepted, this can be regarded as an exclusion test and a cointegration subset is considered (Juselius 1991). On theoretical grounds there is no reason because the wage and price inflation should be explained by the level rather than changes in labour productivity so that the second option is preferred.

Dynamic specification on quarterly data is more complex than for annual. The lag-length of the VAR that whitens the residuals of each equation is lag four. Normality is still a problem for the wage and price equations, possibly because of the presence of some outliers (with respect to the estimated relation) in the early observations of the sample presumably due to the oil shock in early 70s. Heteroskedasticity is present in the price and unemployment equation. We got rid of these problems by including dummies in the years following the oil shock: D73 and D74 appear to be significant in the wage equation, D76 in price equation and D73:2 and D73:3 in the unemployment equation (see Table 7).

Table 7
Diagnostics for a VAR(4)

Equations	Godfrey's test p-value	Jarque-Bera test p-value	Heteroskedasticity p-value
<i>Without intervention dummies</i>			
<i>Wage equat.</i>	0.095	0.00	0.37
<i>Price equat.</i>	0.30	0.013	0.043
<i>Unempl. equat.</i>	0.11	0.094	0.00
<i>Prod. equat.</i>	0.87	0.97	0.14
<i>With intervention dummies</i>			
<i>Wage equat.</i>	0.097	0.21	0.30
<i>Price equat.</i>	0.10	0.86	0.64
<i>Unempl. equat.</i>	0.15	0.77	0.40
<i>Prod. equat.</i>	0.87	0.97	0.14

Note: Godfrey's test of residual serial correlation distributed, F-version

Jarque-Bera's test of the normality of residuals based on skewness and kurtosis, distributed as a $\chi^2(2)$

Heteroskedasticity test based on the regression of squared residuals on squared fitted values, distributed as an F

p-value: probability error of type I

Given the choice of the second option above and diagnostics on dynamic specification, the Johansen procedure is performed by testing the ECM restriction on a four dimensional VAR(4) defined for the observational variable vector: $z'_t = (\Delta w_t, \Delta p_t, u_t, \Delta q_t)$ and dummies. The model to be estimated is

$$(3) \Delta z_t = \mu + \sum_{i=1}^3 \Gamma_i \Delta z_{t-i} + \Pi z_{t-4} + Bx_t + u_t$$

where x_t is a vector of $I(0)$ variables. In this case these stationary variables are the dummies mentioned above. Both the maximal eigenvalue and trace tests allow to reject the null of no cointegrating vectors (Table 8) and two cointegrating vectors can be accepted: $\beta_1 = (-1, 1.21, -0.075, -16.28)$, $\beta_2 = (-1, 0.72, -0.029, 0.28)$.

At first sight, the coefficient for productivity growth in the first vector is not plausible and casts some doubts on the cointegration relationship found. However, if labour productivity growth, Δq_t , included in the estimated cointegrating vectors, is stationary, as expected from the univariate integration analysis, this is a sufficient condition for cointegration among the other three variables. Stationarity can be easily tested by imposing the following linear restriction: $H(0, 0, 0, 1)$. This test allows to accept the null of stationarity of labour productivity growth ($\chi^2(2)=1.67$, PValue=0.43) and, hence, the exclusion of this variable from the cointegrating set.

We further attempt to establish a cointegrating relationship for the subset of the other three variables, $y'_t = (\Delta w_t, \Delta p_t, u_t)$, with y_t in place of z_t and x_t including Δq_t in addition to the dummies. Both cointegration tests yield evidence for one cointegrating vector, see Table 8.

Table 8
Johansen Maximun Likelihood Procedure (Trended variables, Trend in DGP) for a VAR(4)

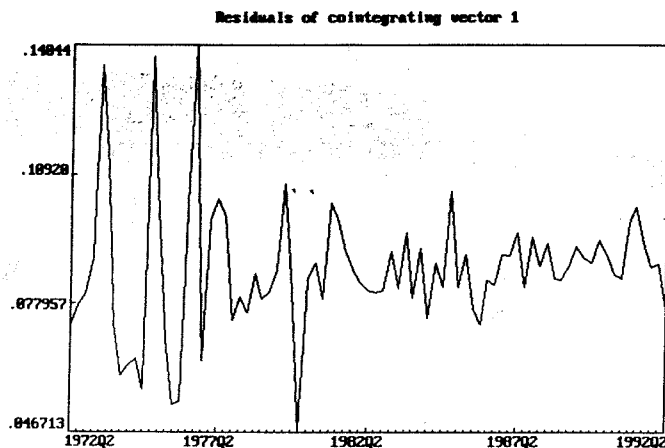
H_0	H_1	Statistic	95% CV	λ_i
<i>Cointegration LR test based on maximal eigenvalue of the stochastic matrix</i>				
$r = 0$	$r = 1$	35.49	20.96	0.35
$r \leq 1$	$r = 2$	12.67	14.06	0.14
$r \leq 2$	$r = 3$	0.42	3.76	0.05
<i>Cointegration LR test based on trace of the stochastic matrix</i>				
$r = 0$	$r \geq 1$	48.59	29.68	0.35
$r \leq 1$	$r \geq 2$	13.10	15.41	0.14
$r \leq 2$	$r \geq 3$	0.42	3.76	0.05

Note: Critical Values in Osterwald-Lenum (1992)

The correspondent cointegrating vector is $\beta' = (-1, 0.70, -0.032)$ and the vector of adjustment coefficients is $\alpha(1.10, 0.022, 1.10)$.¹² The cointegration residuals are shown in Fig. 10. By a simple inspection these residuals seem to be stationary.

¹² The estimation of long run relationship without introducing intervention dummies gives the same result of an $I(0)$ productivity growth and the following cointegration vector: $(-1, 0.72, -0.029)$ so that the introduction of somewhat arbitrary variables like dummies doesn't affect results in any relevant way.

Fig 10



The estimated long run equation, $\Delta w = 0.70\Delta p - 0.032u$, is very similar to that estimated on annual data. It shows a positive relationship between wage inflation and price inflation, the magnitude is slightly less than expected from the no money illusion hypothesis. The negative elasticity of wages to unemployment may imply, again, a long run Phillips curve type relationship. The magnitude in the adjustment coefficient indicates a very high speed of adjustment in wages and unemployment equation and almost no adjustment in the price equation: this will be interpreted as evidence supporting the idea that the relation found is a wage equation rather than a price equation. This will be discussed later, when the ECM will be estimated for short run dynamics and t-tests on α values will be performed.

The following table exhibits some linear restriction on the β vector.

Restricted cointegration vector: testing linear restriction on $\hat{z}_t = (\hat{\Delta w}_t, \hat{\Delta p}_t, \hat{u}_t)$

H_0 1(-1, 1, 0) 1Sp(β)	LR(2) = 19.82	prob value (0.000)
H_0 2(-1, 0, 0) 1Sp(β)	LR(2) = 24.81	(0.000)
H_0 3(0, 1, 0) 1Sp(β)	LR(2) = 23.76	(0.000)
H_0 4(0, 0, -1) 1Sp(β)	LR(2) = 30.40	(0.000)
H_0 5(-1, 0.70, 0) 1Sp(β)	LR(1) = 23.01	(0.000)
H_0 6(-1, 1, a) 1Sp(β)	LR(1)=4.16	(0.041)
H_0 7(-1, 1, -0.032) 1Sp(β)	LR(1)=5.98	(0.05)

H1, H2, H3, H4 reject the hypothesis of stationarity of real wage inflation, nominal wage and price inflation, and unemployment rate. H5 rejects the null of a zero coefficient of unemployment rate in the estimated cointegration vector. H1 can be also interpreted as rejecting the null of a zero coefficient of unemployment under no money illusion, so that the trade off is significant independently from money illusion.

We address now the careful issue of the presence of money illusion in the data performing some linear restrictions on the long run relationship presented above. H6 (estimated $\alpha = -0.023$) rejects (on the border) the null of no money illusion and a re-estimated unemployment coefficient under the null, while H7 accepts (on the border) the null of no money illusion in the estimated cointegration vector. The overall outcome seems to be that money illusion may be small. As a consequence of that, it is also possible to re-express the wage inflation equation in real terms so that the trade off appears to be defined between real wage inflation and unemployment rate. Hence, a consistent long run wage equation, under H7 becomes $\Delta w = \Delta p - 0.032u$. This equation describes a negative relation between real wage and unemployment. There are different rationales for this statistical outcome. We suggest that the negative relation keeps two main phenomena: one is that firms may have been rationed on the good market, as described by disequilibrium models in the case of Keynesian unemployment, so that a lower real wage makes more binding for firms the constraint on the good market and an higher real wage reduces unemployment. The second is possibly linked to a real, long lasting effect of monetary deflation of 80s: because of lackness of credibility, the only way to deflate the economy was to sustain high unemployment costs, in order to reduce wage claims and price setters expectations.

In the long run equilibrium, considering the non zero mean in the cointegration residuals, the equilibrium rate of unemployment is around 10%. This is a rather high figure computed under the hypothesis of a constant equilibrium rate of unemployment over the sample period, as described by the cointegration vector.

6. Estimating the short run dynamics

The negative relationships between nominal wage inflation and unemployment rate is explored in more details using the error correction form of the VAR(4). The OLS estimated ECM is given in Table 9. Dummies are included but not reported and t-ratio are in brackets. Diagnostics tests are satisfied in each equation (Table 10) but R^2 and F statistic of joint significance suggests that the best specified ECM equation between wage and unemployment equation is the first one (Table 9), this somewhat supports the preferred normalisation of the cointegration relationship with respect to wage inflation. The ECM term is not significant in the short run price equation suggesting that the specification is not satisfying for this variable.

The short run wage equation shows a trade off between the change in unemployment rate and acceleration of wage inflation, a positive relation to the price acceleration, with near to unity coefficients (this supports the idea for a very small measure for money illusion even in the short run). The negative coefficient of lagged change in unemployment gives evidence for counter-clockwise loops: as originally found

by Phillips for a sample period before World War II, it implies that if unemployment is decreasing in a given period (boom), employers will bid more vigorously for labour and wage inflation will accelerate.

The negative effect of productivity growth is interpretable as an efficiency effect working through the unemployment rate. The negative effect of productivity growth on unemployment can be seen in the unemployment ECM. This seriously makes a case for displacement effect as a main determinant of unemployment performance in recent years.

Table 9
Error Correction Models (1970Q1-1992Q2)#

Regressor	Dependent Variable		
	$\Delta\Delta w$	$\Delta\Delta p$	Δu
Constant	0.11 (6.02)*	0.002 (0.27)	0.12 (2.31)*
$\Delta\Delta w(-1)$	-0.96 (-8.12)*	0.069 (1.20)	-0.22 (-0.64)
$\Delta\Delta w(-2)$	-1.09 (-6.93)*	0.015 (0.20)	-1.02 (-2.38)*
$\Delta\Delta w(-3)$	-1.16 (-6.38)*	-0.057 (-0.66)	-1.14 (-2.19)*
$\Delta\Delta p(-1)$	0.97 (4.79)*	-0.25 (-2.29)*	-0.65 (-1.17)
$\Delta\Delta p(-2)$	1.16 (4.85)*	-0.22 (-1.93)*	0.99 (1.54)
$\Delta\Delta p(-3)$	0.88 (3.79)*	-0.26 (-2.31)*	1.94 (3.00)*
$\Delta u(-1)$	-0.0005 (-0.014)	-0.05 (-3.29)*	-0.09 (-0.88)
$\Delta u(-2)$	-0.10 (-2.58)*	-0.073 (-4.50)*	0.05 (0.57)
$\Delta u(-3)$	-0.04 (-1.13)	0.02 (1.23)	0.14 (1.51)
$\Delta q(-1)$	-0.31 (-2.15)*	0.18 (2.44)*	0.03 (0.074)
$\Delta q(-2)$	0.08 (0.52)	0.09 (1.21)	-0.11 (0.26)
$\Delta q(-3)$	0.06 (0.40)	0.09 (1.18)	-0.91 (-2.09)*
$\Delta q(-4)$	0.12 (0.89)	-0.10(-1.53)	-0.87(-2.19)*
ECM(-4)	-1.28 (-5.90)*	-0.04 (-0.40)	-1.25 (-2.01)*
R^2	0.70	0.69	0.65
DW	2.06	1.65	2.01
F	F(16,60) = 9.13	F(15,61) = 9.4	F(16,60) = 7.04

Note: t-statistics in the brackets

Intervention dummies are included in each estimated equation, but not reported.

* Significantly different from zero variables

Table 10
Diagnostics for the ECMs

ECMs	Godfrey's test p-value	FF- tests	Jarque-Bera- tests	Heterosk-tests
$\Delta\Delta w$	0,40 F(4,56)	0.39 F(1,59)	0.38	0.27
$\Delta\Delta p$	0.71 F(4,57)	0.75 F(1,60)	0.75	0.48
Δu	0.14 F(4,56)	0.76 F(1,59)	0.70	0.41

Note: Godfrey's test of residual serial correlation (degrees of freedom given in the table)
FFtest, Ramsey's Reset test of functional form (degrees of freedom given in the table)
Jarque-Bera test for the normality of residuals based on skewness and kurtosis, distributed as a $\chi^2(2)$
Heteroskedasticity test based on the regression of squared residuals on squared fitted values, distributed as an F(1,75)
p-value: probability error of type I

We partitioned the cointegration residuals into negative and positive to analyse the symmetry of adjustment to the long run target for each ECM above, via the associated t-ratio. The magnitude and significance of the estimated parameters of the other variables in the ECM don't change so we only reported the partitioned error correction terms. We tested for equality of the coefficients of the partitioned residuals in the wage equation ($\chi^2(1)=0.03$, pvalue=0.85), the result is a symmetric adjustment of the wage inflation to its long run equilibrium. Nominal wage is either below or above its target, because of a shock in one of the cointegrating variables, will adjust and hence this can be interpreted as no evidence for a downward rigidity of nominal wage inflation.

Table 11
Partitioned Error Correction Terms

ECMs	Dependent Variable		
	$\Delta\Delta w$	$\Delta\Delta cp$	Δu
Ecm(+)(-4)	-1.26 (-5.10)*	-0.03 (-0.31)	-0.86 (-1.27)
Ecm(-)(-4)	-1.30 (-4.96)*	-0.05 (-0.40)	-1.82 (-2.47)*

Note: t-statistics in brackets
Ecm(+) partitioned residuals: positive
Ecm(-) partitioned residuals: negative
* Significantly different from zero variables

7. Conclusions

The overall evidence from applying cointegration techniques to wage and price inflation, unemployment rate and labour productivity growth is that of a significant trade off between wages inflation and unemployment. The two examined data set give somewhat different results about money illusion. In the case of annual data spanning from 1960 to 1990, a negatively sloped long run Phillips Curve emerges, with a money illusion coefficient significantly less than one. In the case of quarterly data spanning from 1970q1 to 1992q2 there is some, but not absolutely reliable (the relevant hypothesis is accepted on the border of significance level), evidence for a rejection of the money illusion hypothesis so that the long run trade off may be expressed in real terms.

This difference can be interpreted as due to the different spanning of the data and to the different inflation expectation process in each period. Before the 70s inflation was low and relatively stable and the expectations of future inflation were low as well, after the 70s inflation became more volatile and much higher on average than before, endogenous variables and expectations were subject to frequent shocks, vanishing the stability of the trade off. In the 80s deflation was started at world-wide level and a stable trade off arose again. The scatter plots of nominal inflation variables against unemployment rate in the selected sub-sample (Fig. 11 for the wage inflation against the unemployment rate and Fig. 12 for the price inflation, see Appendix) very much resemble the original found by Phillips before the end of the 60s and after mid-70s.

Annual data comprehend two sub-periods where the Phillips mechanism is at work: the 60s characterised by clockwise loop and the 80s, characterised by long lasting deflation. The weakness of the Phillips relation in the period 1968-75, is not enough to prevent that, at statistical level, a long run relation is found with money illusion. Quarterly data set only comprehend 70s and 80s, with a large weight of the shocks of the early 70s and the higher incidence of complete wage indexation in the sample period, so that the coefficient of wages to price inflation is higher and not different from one. However, this statistical relation gives evidence for the trade off and, in the ECM formulation, the unemployment rate is a determining variable for the wage inflation target. This is consistent with wage bargaining theory and casts doubt on the empirical validity of insider-outsider models, according to which the relationship between unemployment and real wages is weak.

References

- Alexander (1993), 'The Changing Relationship between Productivity, Wages and Unemployment in the U.K', *Oxford Bulletin*, Feb., vol. 55, Iss: 1, pp. 87-102.
- G. Alogouskufis and R. Smith (1991), 'On Error Correction Models: Specifications, Interpretation, Estimation', *Journal of Economic Surveys*, vol. 5, n.1, pp. 97-128.
- O. Ashenfelter and G.E. Johnson (1969), 'Bargaining Theory, Trade Union and Industrial Strike Activity', *AER*, March, 59(1), pp. 35-49.
- L.N. Ball, N.G. Mankiw and D. Romer (1988), 'The New Keynesian Economics and the Output-Inflation Trade-Off', *Brooking Papers on Economic Activity*, n. 1, pp. 1-65.
- O. Blanchard (1983), 'Price Asynchronization and Price-Level Inertia', in *New Keynesian Economics*, N.G. Mankiw and D. Romer eds, vol.1, The MIT Press, Cambridge, Massachusetts, London, England, (1991).
- _____ and Fisher (1989), *Lectures on Macroeconomics*, Cambridge, Mass: MIT Press.
- G. Caselli and L. Marinelli (1993) 'Italian GNP Growth 1890-1992: a Unit Root or Segmented Trend Representation', *Materiali di Discussione*, n. 93, Dipartimento Economia Politica, Universita'di Modena.
- J. Darby and S. Wren-Lewis (1993), 'Is there a Cointegrating Vector for UK Wages?', *Journal of Economic Studies*, vol. 20, n. 1/2, pp. 87-115.
- L.A. Dicks-Mireaux and J.C.R. Dow (1959), 'The Determinants of Wage Inflation: United Kingdom, 1946-1956', *J.Royal Stat. Society*, Series A, Part II, 122, pp. 145-74.
- D. Dickey - W.A. Fuller (1979), 'Distribution of the Estimation for Autoregressive Time Series with a Unit Root', *Journal of American Statistical Association*, 74, pp. 427-31.
- _____ and W.A. Fuller (1981), 'Likelihood ratio statistics for Autoregressive Time series with a Unit Root', *Econometrica*, 49, pp. 1057-72
- R.F. Engle - C.W.J. Granger (1987), 'Cointegration and Error Correction: Representation, Estimation and Testing', *Econometrica*, 251.
- I. Fisher (1926), 'A Statistical Relation between Unemployment and Price Changes', *Int. Lab. Rev.*, June, 13(6), pp. 185-92, reprinted in *Journal of Political Economy*, I, March/April, 81(2), pp. 496-502.
- S. Fisher (1977), 'Long Term Contracts, Rational Expectations, and the Optimal Money Supply Rule', *Journal of Political Economy*, 85, n.1, pp.191-205.
- M. Friedman (1968), 'The Role of Monetary Theory', *American Economic Review*, 58, pp. 1-17.
- C.W.J. Granger (1981), 'Some Properties of Time Series Data and their Use in Econometric Model Specification', *Journal of Econometrics*, 16, pp. 121-30.

- S.G. Hall (1986), 'An Application of The Granger and Engle Two-Step Estimation Procedure to United Kingdom Aggregate Wage Data', *Oxford Bulletin of Economics and Statistics*, 48, 3.
- _____(1989), 'Maximum Likelihood Estimation of Cointegration Vector: An Example of The Johansen Procedure', *Oxford Bulletin of Economics and Statistics*, 51, pp. 229- 238.
- S. Hylleberg and G.E. Mizon (1990), 'Cointegration and Error Correction Mechanism', *The Economic Journal*, 99, pp. 113-125.
- S. Johansen (1988), 'Statistical Analysis of Cointegration Vectors', *Journal of Economics Dynamics and Control*, 12.
- ____and K. Juselius (1990), 'Maximum Likelihood Estimation and Inference on Cointegration- with Applications to The Demand for Money', *Oxford Bulletin of Economics and Statistics*, 52.
- ____and K. Juselius (1992), 'Testing Structural Hypothesis in a Multivariate Cointegration Analysis of The PPP and The UIP for UK', *Journal of Econometrics*, 53, pp. 211-44.
- K. Juselius (1991), 'Domestic and Foreign Effects on Prices in an Open Economy', *Discussion Paper 91-05*, Institute of Economics, University of Copenhagen
- L.R. Klein and R.J. Ball (1959), 'Some Econometrics of the Determination of Absolute Prices and Wages', *Econ. Journal*, Sept. 69, pp. 465-82.
- E. Kuh (1967), 'A Productivity Theory of Wage Level: An Alternative to the Phillips Curve', *Rev. Econ. Stud.*, Oct., 34, pp. 333-60.
- R. Layard, S. Nickell and R. Jackman (1991), *Unemployment*, The Alden Press, Oxford.
- R.G. Lipsey (1960), 'The Relation between Unemployment and the Rate of Change of Money Wage Rates in the United Kingdom 1862-1957: a further Analysis', *Economica*, 27, pp. 1-31.
- W.F. Mitchell (1993), 'Testing for Unit Roots and Persistence in OECD Unemployment Rates', *Applied Economics*, 25, pp. 1489-1501.
- R. Nymoen (1989), 'Modelling Wages in the Small Open Economy: an Error Correction Model of Norwegian Manufacturing Wages', *Oxford Bulletin of Economics and Statistics*, 51, 3.
- M. Osterwald - Lenum (1992), 'Recalculated and extended tables of the asymptotic distribution of some important maximum likelihood cointegration test statistics', *Oxford Bulletin of economic and Statistic*, 54, pp. 461-472.
- P. Onofri and B. Salituro (1985), 'Inflazione e Politiche di Stabilizzazione in Italia: 1960-1984', *Politica Economica*, n.2, Agosto, pp. 167-196.
- P. Perron (1988), 'Trends and Random Walks in Macroeconomic Time Series. Further Evidence from a New Approach', *Journal of Economic Dynamics Control*, 12.
- _____(1989), 'The Great Crash, the Oil Price Shock and the Unit Root Hypothesis', *Econometrica*, 57, pp. 1361-1401.
- _____(1993), 'The HUMP-Shaped Behavior of Macroeconomic Fluctuations', *Empirical Economics*, 18, pp. 707-27.

- A.W.H. Phillips (1958), 'The Relation between Unemployment and the Rate of Change of Money Wage Rates in the United Kingdom, 1861-1957', *Economica*, 25, pp. 283-299.
- _____ (1959), 'Wage Changes and Unemployment in Australia, 1947-1958', The economic Society of Australia and New Zealand (Victoria Branch) Economic Monograph, n.14.
- P.C. Phillips - P. Perron (1988), 'Testing for Unit Roots in Time Series Regression', *Biometrika*, 75, pp. 335-346.
- B. Raj (1992), 'International Evidence of Persistence in Output in the Presence of an Episodic Change', *Journal of Applied Econometrics*, n.7, pp. 281-293.
- P. Rappoport and L. Reichlin (1990), 'Segmented Trends and Non-Stationary Time Series', *The Economic Journal*, 99, pp. 168-177.
- S.E. Said - D.A. Dickey (1984), 'Testing for Unit Roots in Autoregressive-Moving Average Models of Unknown Order', *Biometrika*, 71, pp. 599-604.
- P. Samuelson and R. Solow (1960), 'The Analytics of an Anti-Inflation Policy', *American Economic Review*, 50, pp. 177-194.
- N.J. Simler and A.Tella (1968), 'Labor Reserves and the Phillips Curve', *Rev. Econ. Stud.*, Feb., 50, pp. 32-49.
- J.H. Stock (1987), 'Asymptotic Properties of Least-Squares Estimators of Co-integrating Vectors', *Econometrica*, 58, pp. 1035-56.
- J.B. Taylor (1979), 'Staggered Wage Setting in a Macro Model', *American Economic Review*, 69, May, pp. 108-113.
- J.Tobin (1972), 'Inflation and Unemployment', *American Economic Review*, March, 62(1), pp. 1-18.
- J. Vanderkamp (1972), 'Wage Adjustment, Productivity and Price Change Expectations', *Rev.Econ. Stud.*, Jan., 39(1), pp. 61-72.
- M. Zenezini (1986), 'I Salari e la Curva di Phillips: Alcune Considerazioni sull'Esperienza Italiana e un Commento ad un Articolo di Onofri e Salituro', *Politica Economica*, n.3, Dicembre, pp. 401-433.

Appendix

Fig. 11. Scatter plot of wages inflation versus unemployment rate: 1960-1990

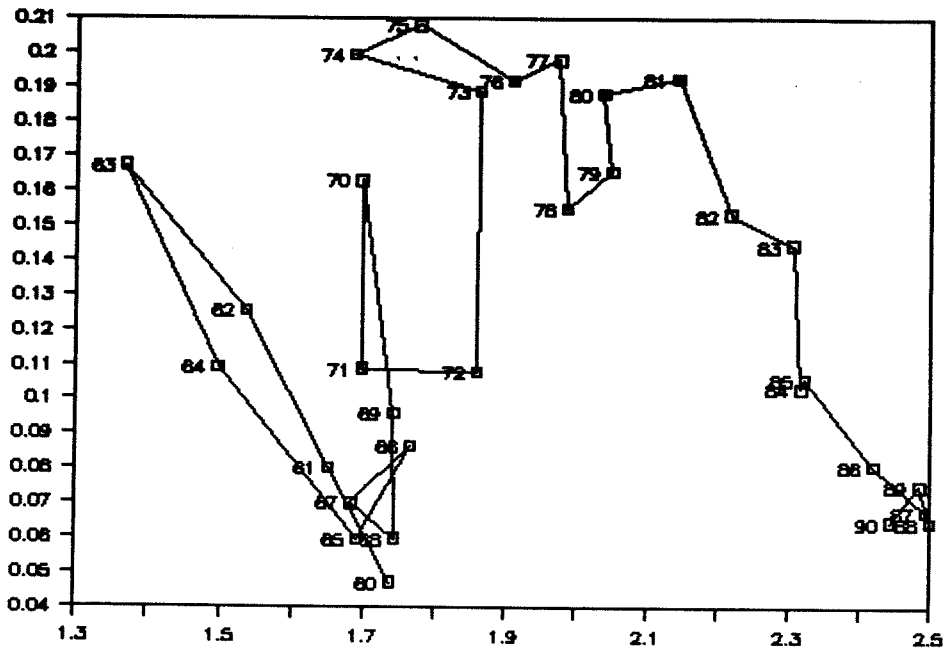
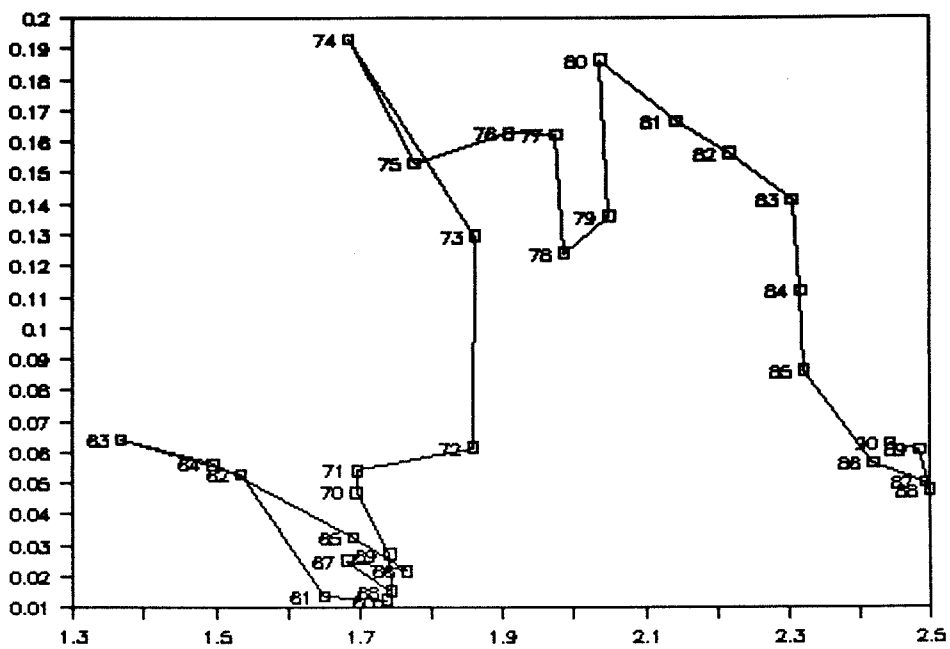


Fig. 12. Scatter plot of price inflation versus unemployment rate: 1960-1990



Materiali di discussione

1. Maria Cristina Marcuzzo [1985] "Joan Violet Robinson (1903-1983)", pp.134.
2. Sergio Lugaesi [1986] "Le imposte nelle teorie del sovrappiù", pp.26.
3. Massimo D'Angelillo e Leonardo Paggi [1986] "PCI e socialdemocrazie europee. Quale riformismo?", pp.158.
4. Gian Paolo Caselli e Gabriele Pastrello [1986] "Un suggerimento hobsoniano su terziario e occupazione: il caso degli Stati Uniti 1960/1983", pp.52.
5. Paolo Bosi e Paolo Silvestri [1986] "La distribuzione per aree disciplinari dei fondi destinati ai Dipartimenti, Istituti e Centri dell'Università di Modena: una proposta di riforma", pp.25.
6. Marco Lippi [1986] "Aggregation and Dynamics in One-Equation Econometric Models", pp.64.
7. Paolo Silvestri [1986] "Le tasse scolastiche e universitarie nella Legge Finanziaria 1986", pp.41.
8. Mario Forni [1986] "Storie familiari e storie di proprietà. Itinerari sociali nell'agricoltura italiana del dopoguerra", pp.165.
9. Sergio Paba [1986] "Gruppi strategici e concentrazione nell'industria europea degli elettrodomestici bianchi", pp.56.
10. Nerio Naldi [1986] "L'efficienza marginale del capitale nel breve periodo", pp.54.
11. Fernando Vianello [1986] "Labour Theory of Value", pp.31.
12. Piero Ganugi [1986] "Risparmio forzato e politica monetaria negli economisti italiani tra le due guerre", pp.40.
13. Maria Cristina Marcuzzo e Annalisa Rosselli [1986] "The Theory of the Gold Standard and Ricardo's Standard Commodity", pp.30.
14. Giovanni Solinas [1986] "Mercati del lavoro locali e carriere di lavoro giovanili", pp.66.
15. Giovanni Bonifati [1986] "Saggio dell'interesse e domanda effettiva. Osservazioni sul capitolo 17 della General Theory", pp.42.
16. Marina Murat [1986] "Between old and new classical macroeconomics: notes on Leijonhufvud's notion of full information equilibrium", pp.20.
17. Sebastiano Brusco e Giovanni Solinas [1986] "Mobilità occupazionale e disoccupazione in Emilia Romagna", pp.48.
18. Mario Forni [1986] "Aggregazione ed esogeneità", pp.13.
19. Sergio Lugaesi [1987] "Redistribuzione del reddito, consumi e occupazione", pp. 17.
20. Fiorenzo Sperotto [1987] "L'immagine neopopulista di *mercato debole* nel primo dibattito sovietico sulla pianificazione", pp. 34.
21. M. Cecilia Guerra [1987] "Benefici tributari del regime misto per i dividendi proposto dalla Commissione Sarcinelli: una nota critica", pp 9.
22. Leonardo Paggi [1987] "Contemporary Europe and Modern America: Theories of Modernity in Comparative Perspective", pp. 38.
23. Fernando Vianello [1987] "A Critique of Professor Goodwin's 'Critique of Sraffa'", pp. 12.
24. Fernando Vianello [1987] "Effective Demand and the Rate of Profits: Some Thoughts on Marx,

- mobile e l'“appiattimento” delle retribuzioni in una ricerca” pp. 120
75. Enrico Giovannetti [1990] “Crisi e mercato del lavoro in un distretto industriale: il bacino delle ceramiche. Sez. I” pp. 150
 76. Enrico Giovannetti [1990] “Crisi e mercato del lavoro in un distretto industriale: il bacino delle ceramiche. Sez. II” pp. 145
 77. Antonietta Bassetti e Costanza Torricelli [1990] “Il portafoglio ottimo come soluzione di un gioco bargaining” pp. 15
 78. Antonietta Bassetti e Costanza Torricelli [1990] “Una riqualificazione dell'approccio bargaining alla selezioni di portafoglio” pp. 4
 79. Mario Forni [1990] “Una nota sull'errore di aggregazione” pp. 6
 80. Francesca Bergamini [1991] “Alcune considerazioni sulle soluzioni di un gioco bargaining” pp. 21
 81. Michele Grillo e Michele Polo [1991] “Political exchange and the allocation of surplus: a model of two-party competition” pp. 34
 82. Gian Paolo Caselli e Gabriele Pastrello [1991] “The 1990 Polish Recession: a Case of Truncated Multiplier Process” pp. 26
 83. Gian Paolo Caselli e Gabriele Pastrello [1991] “Polish firms: Pricate Vices Public Virtues” pp. 20
 84. Sebastiano Brusco e Sergio Paba [1991] “Conessioni, competenze e capacità concorrenziale nell'industria della Sardegna” pp. 25
 85. Claudio Girmaldi, Rony Hamaui, Nicola Rossi [1991] ”Non marketable assets and households' portfolio choices: a case study of Italy” pp. 38
 86. Giulio Righi, Massimo Baldini, Alessandra Brambilla [1991] ”Le misure degli effetti redistributivi delle imposte indirette: confronto tra modelli alternativi” pp. 47
 87. Roberto Fanfani, Luca Lanini [1991] ”Innovazione e servizi nello sviluppo della meccanizzazione agricola in Italia” pp. 35
 88. Antonella Caiumi e Roberto Golinelli [1992] ”Stima e applicazioni di un sistema di domanda Almost Ideal per l'economia italiana” pp. 34
 89. Maria Cristina Marcuzzo [1992] ”La relazione salari-occupazione tra rigidità reali e rigidità nominali” pp. 30
 90. Mario Biagioli [1992] ”Employee financial participation in enterprise results in Italy” pp. 50
 91. Mario Biagioli [1992] ”Wage structure, relative prices and international competitiveness” pp. 50
 92. Paolo Silvestri e Giovanni Solinas [1993] ”Abbandoni, esiti e carriera scolastica. Uno studio sugli studenti iscritti alla Facoltà di Economia e Commercio dell'Università di Modena nell'anno accademico 1990/91” pp. 30
 93. Gian Paolo Caselli e Luca Marinelli [1993] ”Italian GNP growth 1890-1992: a unit root or segmented trend representation?” pp. 25
 94. Angela Politi [1993] ”La rivoluzione fraintesa. I partigiani emiliani tra liberazione e guerra fredda, 1945-1955” pp. 55
 95. Alberto Rinaldi [1993] “Lo sviluppo dell'industria metalmeccanica in provincia di Modena: 1945-1990” pp. 70
 96. Paolo Emilio Mistrulli [1993] “Debito pubblico, intermediari finanziari e tassi d'interesse: il caso italiano” pp. 30
 97. Barbara Pistori [1993] “Modelling disaggregate and aggregate labour demand equations. Coin-

tegration analysis of a labour demand function for the Main Sectors of the Italian Economy: 1950-1990" pp. 45

98. Giovanni Bonifati [1993] "Progresso tecnico e accumulazione di conoscenza nella teoria neoclassica della crescita endogena. Una analisi critica del modello di Romer" pp. 50

