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European Monetary Union

2002/41



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Printed in Italy in December 2002
Università degli Studi dell'Insubria
Via Ravasi 2, 21100 Varese, Italy

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On the determinants of inflation in Italy: evidence of cost-push effects before the European Monetary Union

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December 2002

Abstract

This paper provides evidence on price markup and inflation dynamics in Italy over the period 1970-1998. We investigate the price mark-up on imported and labor costs and its relation to inflation, using cointegration techniques. It is found that, despite different policy regimes across decades, the relation between the markup and inflation is remarkably stable, and useful in predicting inflation dynamics. This evidence clarifies that pure monetarist theories of inflation are not able to account for the price dynamics in Italy from the seventies up to stage III of the European Monetary Union.

Keywords: Inflation dynamics, Mark-up model, Monetary policy, Phillips curve, Cointegration, Equilibrium correction model.

JEL classification: C32, E00, E31.

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1 Introduction

The explicit mission of the European Central Bank (ECB) is concerned with price stability in the Euro area. Several papers attempt to analyze the sources and nature of inflation on aggregate time series of the Euro area, see e.g. Gali et al. (2001) and reference therein. The empirical analysis of the Euro area is usually performed on backward-aggregated data. Unfortunately the aggregation process may blur the behavioral relations connecting prices and other macroeconomic variables at the country level: each country was subject to different monetary authorities and economic policies, in addition to being potentially characterized by a different production structure.

In this paper we analyze inflation data for a single country, Italy; this renders the behavioral assumptions less questionable than at the Euro aggregate level, and makes it possible to single out possible structural breaks as the effects of national economic policy rather than artifacts of aggregation. The economic history of Italy in the last three decades suggests the presence of several monetary policy regimes. This allows to test whether these regimes induced a different structure on inflation dynamics.

Inflation dynamics are usually analyzed by some form of Phillips curve, connecting inflation to a cyclical indicator, see e.g. Stock and Watson (1999). More advanced models view inflation as the result of interactions of aggregate demand and supply; see Gali et al. (2001) for a model incorporating effect of marginal costs. In this paper we concentrate on the supply-side determinants of inflation, and specifically on the relation between markup and inflation.

Several authors have recently postulated a (negative) relation between inflation and the markup in the long run, see Banerjee et al. (2001) and reference therein. The underlying idea is that inflation may represent a cost for firms even in the long run either because higher inflation leads to greater competition and thus to a reduction of the markup, or because of the difficulties faced by price-setting firms in adjusting prices in an inflationary environment with incomplete information.

This type of relation has important implications. It implies that, for given level of productivity, inflation is positively related to the real wage; if unemployment is partly dependent on the real wage, this in turn implies a non-vertical long-run Phillips curve.¹ The existence of such a relation is obviously an empirical question; cointegration techniques have been recently employed in this area by Banerjee et al. (2001), who report results for the case of Australia and by Banerjee and Russell (2000) who investigate the issue in the G7 economies.

In this paper we apply the same tools on Italian quarterly data over the last three decades preceding the European Monetary Union (EMU). Despite very different monetary regimes in each decade of the sample period, see Fratianni and Spinelli (2001) and Bertocco (2002), we find a surprisingly stable long-run relation connecting the Italian price index, unit labor cost and import prices. This finding is at variance with the claim that only the stance of monetary policy affects inflation, typical of

¹This relation would also explain the negative correlation between stock returns and inflation, where stock returns reflect firms' profitability. Moreover this would explain why firms prefer a low rather than high level of inflation. For more details and a model of the labor market which implies a negative relation between inflation and the markup, we refer to Banerjee et al. (2001) and reference therein.

the interpretation given by the monetary historians Fratianni and Spinelli (2001).

They claimed that ‘fiscal dominance’ in the ’70s was the main cause of the growth of monetary aggregates and (hence) of the price level. In the ’80s, when Italy joined the EMS, the monetary authority reached independence from the Treasury and pursued the stabilization of the exchange rate; this policy was ineffective in curbing inflation and led to the over-evaluation of the exchange rate, which caused the devaluation of 1992. Fratianni and Spinelli (2001) attribute the success in curbing inflation to the monetary policy implemented in the rest of ’90s, which aimed at the explicit control of monetary aggregates.

This view has been criticized. Bertocco (2002) noted that in order for the monetarist view to be effective, the policy instrument should be a (totally exogenous) supply of money, the disequilibria in the money market should cause changes in aggregate demand, and that these in turn should only influence prices. He argued that all these assumptions are questionable for the present sample.

Moreover he gave evidence from the Bank of Italy yearly reports of the wide recourse to cost-push explanations by all the Governors in the period, who claimed that inflation dynamics cannot be explained just by the behavior of monetary authorities and that monetary policy affects inflation also through factors that influence production costs; they all suggested the joint use of monetary, fiscal and income policy measures in order to curb inflation.

The investigation of the connection between markup and inflation in this paper sheds some light in the importance of supply-side effects on price dynamics. Thanks to the structure of cointegration models, we are also able to estimate the influence of deviations from the stable long-run relation on predictions of price fluctuations. It is found that the forecast of inflation is deeply affected by deviations from this stable long-run relation. While this does not exclude the possible importance of demand-side determinants, it shows the potential explanatory power of supply-side cost-push effects in the analysis and forecast of inflation.

The rest of the paper is organized as follows. Section 2 presents the quarterly data used in the application; Section 3 reports a simple markup model of price-setting for the Italian aggregate price level and of the dynamics of inflation. Section 4 reports tests and estimation results over the period 1970-1998. Section 5 discusses policy implications of the present empirical findings on the relation between monetary aggregates and inflation. Section 6 reports conclusions and an Appendix completes the paper.

2 A preliminary look at the data

We consider quarterly data on log prices p_t , log unit labor cost ulc_t , and log import prices pm_t , covering the period 1970:2 - 1998:4. p_t is measured as the log GDP deflator with base year 1995. ulc_t is the log nominal Unit Labour Costs of the manufacturing sector; in particular $ulc_t = (w_t - prod_t)$ where w_t is the log nominal wage per person in the manufacturing sector and $prod_t$ is the log output per person in the manufacturing sector; pm_t is the log tariff-adjusted total import price index (including energy) with 1995 as base year.²

²The data source is ISTAT; the data have been kindly provided by Prometeia Associazione, Bologna.

The three prices series p_t , ulc_t and pm_t , are plotted in Fig. 1. It can be observed that the time series are very smooth and increasing. Inflation $\pi_t = \Delta p_t$ is calculated here as the quarterly growth rate of prices p_t ; π_t is graphed in Fig. 2. Here Δ indicates the first difference operator, $\Delta = 1 - L$, L is the backward operator, $Lx_t = x_{t-1}$.

Inflation itself appears to be a non-stationary process within this sample period. The level of inflation increased to double digits in the 1970's, and slowly decreased in the 1980's and 1990's, showing high persistence.

Taking prices p_t as a numeraire, we calculated relative prices, i.e. the log of the ratios of unit labor cost and import prices to the price level, $ulc_t - p_t$ and $pm_t - p_t$. The relative prices and the inflation rate $\pi_t = \Delta p_t$ are graphed in Fig. 2 in levels and first differences. It can be observed that $ulc_t - p_t$, $pm_t - p_t$ and π_t appear stationary in first differences, while they are possibly non-stationary in levels.

Many models may account for the non-stationarity of $ulc_t - p_t$, $pm_t - p_t$ and π_t ; these variables may e.g. be integrated of order one, $I(1)$, or have a time-varying average. In this paper we assume that the various time series are either $I(1)$ or stationary.³

In the present sample inflation has a wide variation, over 30 percentage points. Very different levels of inflation are observed in association with different levels of markup, and this is an ideal situation to investigate and measure the degree of association between the two.

The markup is usually defined in terms of (the negative of) $ulc_t - p_t$, within a closed economy framework. Italy is on the contrary a small open economy, highly dependent on import for raw material and energy in particular; this suggests the possible importance of the markup on import prices, (the negative of) $pm_t - p_t$. Instead of choosing one of the possible markups we consider an average of the two, following De Brouwer and Ericsson (1998).

Other factor costs could be considered, such as the unit cost of capital. Because these are difficult to measure, we subsequently introduce a time trend as a proxy for these other possible factor costs which are not observed. The model we employ also accounts for price rigidities characterizing the Italian economy during the 1970-1998 period. The model is described in the following section.

3 A markup model

In this section we introduce a simple model of the markup and discuss its implications on the long-run of the system as well as on its adjustment behavior.

³This assumption is by no means an hypothesis on the Data Generating Process (DGP) of inflation for any country, any frequency of observation, any sample size. Asymptotics should not be understood as postulating that the effective economy is left at work for an increasing number of quarters; in the present case this would mean observing data about the Italian economy during the EMU, a different epoch from the one of the last three decades of the century considered here. Asymptotics should be understood instead as a technical thought-experiment, which is used to approximate the sampling distribution of estimators and test statistics. In this thought-experiment the DGP that is assumed to have generated the observed sample is left at work for an increasing number of periods. In other words the finding that 'inflation is $I(1)$, prices are $I(2)$ ' is not a statement about the actual economy in the next century, but an useful approximation for the observed sample period.

3.1 A simple model

We consider a markup model in the tradition of e.g. Franz and Gordon (1993)⁴; however, the standard cost-push model is suitably modified to take into account the relationship among the markup and inflation postulated by more recent theories. Following Banerjee and Russell (2000), the main hypothesis is that in an open economy framework, in the long run the domestic general price level is a markup over total unit costs net of the cost of inflation.

Assuming linear homogeneity, the relation linking the consumer price level to its determinants in the long run can be formulated as

$$z_t = p_t - \gamma ulc_t - \delta pm_t - \eta \pi_t = q_t - \eta \pi_t \quad (1)$$

where z_t is the retail markup over costs at time t net of the cost of inflation, q_t is the ‘gross’ markup and $\pi_t = \Delta p_t$ is the inflation rate. The coefficients $\gamma \geq 0, \delta \geq 0$ satisfy the homogeneity restriction $\gamma + \delta = 1$ and can be interpreted as the elasticities of the price level with respect to unit labour costs and import costs, whereas $\eta \geq 0$ is the inflation cost coefficient, i.e. a measure of the impact that inflation exerts on firms’ markup.⁵

We assume the validity of the homogeneity restriction $\gamma + \delta = 1$; in Appendix A2 we discuss a test of this restriction and find ample support to this hypothesis. We also find the inflation cost coefficient η to be significant.

Note that (1) implicitly describes the influence of exchange rates on prices through the price of imports pm_t ; in fact

$$pm_t = ex_t + pm_t^f$$

where pm_t^f represents a (logged) index of total import prices in foreign currency (e.g. in U.S. dollars) at time t and ex_t is the (logged) nominal exchange rate, home versus foreign. Note that $pm_t - p_t = ex_t - pm_t^f - p_t$ can be interpreted as a measure of the real exchange rate.

The following two subsections describe the econometric implications on the long run and on the adjustment towards the long run.

3.2 Long-run implications

Under the assumption of homogeneity $\gamma + \delta = 1$ the reference equation (1) can be re-written in the form

$$z_t = \gamma(p_t - ulc_t) + \delta(p_t - pm_t) - \eta \pi_t. \quad (2)$$

This can be verified by substituting p_t with $(\gamma + \delta)p_t$ in (1).

As discussed in Section 2, π_t , $p_t - ulc_t$, and $p_t - pm_t$ appear to be integrated of order 1, I(1). In order for the retail markup to be stable, eq. (2) must be a cointegrating relation in the sense of Engle and Granger, see Johansen (1996). The relation is of the form

$$(p_t - ulc_t) + \beta_1(p_t - pm_t) + \beta_2 \pi_t = e_t \quad (3)$$

⁴Forward-looking formulations of this kind of models (see e.g. Galí and Gertler, 1999) are not considered here in order to simplify estimation issues.

⁵See Banerjee and Russell (2000) for details.

with $\beta_1 = \delta/\gamma$, $\beta_2 = -\eta/\gamma$ and where $e_t = z_t/\gamma$ is stationary for the markup model to hold in the long run⁶. The cointegration relation (3) implies that shocks involving the markup z_t or e_t have a transitory nature, i.e. tend to dissipate over time. Aggregate demand shocks could well be responsible for the fluctuations of e_t over the business cycle; in the present paper we do not disentangle explicitly supply- and demand-side fluctuations. However we interpret adjustment of inflation to this long-run relation as evidence of cost-push effects.

The econometric investigation of the markup model of inflation dynamics as a long run model can be based on testing the occurrence of the cointegration relation (3) and of the dynamic adjustment towards this relation.⁷

3.3 Dynamic adjustment

Given the cointegration relation (3), the structure of cointegrated VAR models allows to investigate the adjustment behavior of the growth rates of π_t , $ulc_t - p_t$, $pm_t - p_t$. This adjustment behavior is an example of the equilibrium correction model as originally introduced by Davidson et al. (1978).

Following Gruen et al. (1999), we interpret the following price adjustment equation as a ‘accelerationist’-type Phillips curve:

$$\Delta\pi_t = \omega_1\Delta(ulc_t - p_t) + \omega_2\Delta(pm_t - p_t) - ae_{t-1} + lags + \xi_{1t} \quad (4)$$

where ξ_{1t} is a White Noise term and ‘lags’ includes possible lags of $\Delta\pi_t$, $\Delta(ulc_t - p_t)$ and $\Delta(pm_t - p_t)$; e_{t-1} is defined as the deviation from the equilibrium relation in eq. (3). Formally eq. (4) describes the conditional model of $\Delta\pi_t$ given $\Delta(ulc_t - p_t)$ and $\Delta(pm_t - p_t)$, reflecting an expected simultaneous causality from costs of inputs to the growth rate of inflation.⁸

The contemporaneous coefficients ω_1 and ω_2 describe the within-quarter instantaneous effect of imported inflation and labor cost. One may expect imported inflation to have instantaneous effects, i.e. $\omega_2 > 0$, while unit labor cost may take one or more quarters to influence the growth rate of inflation; this would be consistent with $\omega_1 = 0$. This is what is found empirically to be the case.

The long-run adjustment is described by the a coefficient. The a coefficient describes how the inflation acceleration rate adjusts to the (lagged) disequilibrium among the price level and cost factors, $e_t = (p_t - ulc_t) + \beta_1(p_t - pm_t) + \beta_2\pi_t$.

The complementary part of the dynamic system for $(ulc_t - p_t)$, $(pm_t - p_t)$ and π_t consists of the two marginal equations for $(ulc_t - p_t)$, $(pm_t - p_t)$ given the past of the system (see the Appendix). In particular the following error correcting equations hold

$$\Delta(ulc_t - p_t) = \alpha_2 e_{t-1} + lags + \varepsilon_{2t} \quad (5)$$

$$\Delta(pm_t - p_t) = \alpha_3 e_{t-1} + lags + \varepsilon_{3t} \quad (6)$$

⁶Observe that the relation (3) has been normalized with respect to the real unit labour costs just for convenience; different normalizations could be applied.

⁷Observe that (3) can be regarded as a polynomial cointegration relations by recalling that $\pi_t = \Delta p_t$, see e.g. Banerjee et al. (2001). See also the Appendix for a discussion of the issues related to the possible presence of I(2) components in X_t .

⁸Observe that De Brower and Ericsson (1995) found that inflation was best described as a stationary process for their data; accordingly their equilibrium correction formulation reflected this finding and is different from eq. (4). Also Gruen et al. (1999) treat inflation as a stationary process.

Eq. (5) is a real wage equation,⁹ whereas eq. (6) is a real imports-price equation. The coefficients α_2 and α_3 measure the adjustment of real factor costs to the equilibrium relation. Given that Italy is a small open economy, one expects α_3 to be 0, reflecting the insensitivity of import prices to disequilibria associated with Italian prices.¹⁰

On the contrary one may expect α_2 to measure the sensitivity of real wages to disequilibria. Within the sample period, a wage indexation system ('scala mobile') has been in place before the currency crisis of 1992. One may thus expect different reactions of real wages to disequilibria before and after 1992. This can be evaluated by testing for possible breaks in the wage equation, especially in the α_2 coefficient.

The system (4) (5) (6) is a simultaneous system of equations, with a triangular form. The price-adjustment is the only equation to have simultaneous effects. The following section reports empirical results.

4 Econometric results

We estimated a VAR model of order 3 for $X_t = (\pi_t, ulc_t - p_t, pm_t - p_t)'$, over the sample 1971:1 - 1998:4 (112 observations). The deterministic component was specified as $D_t = (d_{74:1t}, t)'$, where $d_{74:1t}$ is a dummy variable for the quarter 1974:1, corresponding to the first oil shock. The linear trend was restricted to lie in the cointegration space. Observe that a trend term in the cointegration relation (3) accounts for other unobserved factor costs affecting (1) such as e.g. the cost of capital per unit of output. Calculations were performed in PcGive 10.0, see Doornik and Hendry (2001).

A set of mis-specification tests on the residuals of the model were performed. The AR 1-5 test on the individual series and at the system level did not signal any residual autocorrelation. The Jarque Bera normality test signaled significant departure from normality for the residuals of the equations of π_t and $ulc_t - p_t$. The absence of normality, however, does not affect the asymptotic inference on the cointegration relations of the system, see Gonzalo (1994). We thus retained this system as correctly specified.

The largest roots of the unrestricted companion matrix were found to be 0.9378, 0.8367, $0.4806 \pm 0.2933i$, $0.2952 \pm 0.4154i$ with moduli 0.9386, 0.8417, 0.5630, 0.5096. This suggests the possible presence of two roots at $z = 1$. We tested for cointegration rank within the restricted-trend model which implies no quadratic trends in the data; the resulting LR trace test are reported in Table 1. The test suggest a cointegration rank equal to 1, as expected.

The largest roots of the companion matrix of the cointegrated model with 1 cointegrating vectors were 1 (twice), $0.4352 \pm 0.3673i$, -0.4296 , $0.2569 \pm 0.3280i$, with moduli 1 (twice), 0.5695, 0.4296, 0.4166. We also performed the I(2) tests in Paruolo (1996) in order to check the homogeneity restriction; these tests did not signal any presence of I(2)-ness.

⁹Recall that $(ulc_t - p_t) = (w_t - prod_t - p_t)$ where w_t is the log nominal wage per person at time t and $prod_t$ is the log output per person, i.e. productivity.

¹⁰In econometric terms this corresponds to the weak exogeneity of the variable $(pm_t - p_t)$ with respect to the cointegrating parameters in (3), see e.g. Hendry (1995).

The long-run relations (3) was estimated as

$$\widehat{e}_t = (p_t - ulc_t) + 0.237(p_t - pm_t) + 9.909\pi_t - 0.00393t, \quad (7)$$

where all coefficients were significant. The time series of \widehat{e}_t is plotted in Fig. 3 along with residuals of the VAR system. In terms of the theoretical reference equation (1) the estimates in (7) imply the following numerical values of parameters¹¹

$$\widehat{z}_t = p_t - 0.81ulc_t - 0.19pm_t - 0.0032t - 8.01\pi_t$$

We next proceeded to estimate a simultaneous system of adjustment equations as described in Section 3.3, fixing the long-run coefficients to their estimated values. After simplification of insignificant coefficients, we obtained the specification listed in Table 2. The LR test of over-identifying restrictions gave a $\chi^2(14)$ test statistic of 20.126 with a p -value of 0.1262, giving support to the reduction. It can be seen that α_3 was not significantly different from 0, as expected. The equilibrium correction term is highly significant in the inflation equation as well as in the wage equation. The fit of the equations is reported in Fig. 4.

In order to investigate the parameter constancy of the model, Fig. 5 and Fig. 6 report the 1-step recursive residuals with associated 95% confidence band and the 1-step recursive Chow tests (see Doornik and Hendry, 2001) performed on the parameters of the vector equilibrium correction representation (VEqC) of the VAR, obtained after fixing the cointegration relation (7). The N-down and N-up versions of the Chow tests gave fewer rejections, and are not reported for brevity.

It should be stressed that the inflation equation is very stable over the rest of the full sample. At the system level, moreover, the parameters of the model appear substantially stable. The stability of this relation supports the cost-push hypothesis over different monetary policy regimes, and it is at variance with the monetarist explanation of inflation.

4.1 Inflation dynamics

The estimated inflation dynamic equation was

$$\begin{aligned} \Delta\pi_t = & 0.10\Delta(pm_t - p_t) - 0.058\widehat{e}_{t-1} + 0.036\Delta(pm_{t-2} - p_{t-2}) + \\ & + 0.018 + 0.028d_{74:1t} + \widehat{\xi}_{1t} \end{aligned}$$

This equations shows a significant simultaneous effect of imported inflation from $\Delta(pm_t - p_t)$, as well as a lagged effect of the same variable. Unit labor cost enters the equation only through the long-run markup (disequilibrium) term \widehat{e}_t . A constant and the dummy for the first oil shock complete the specification. The contemporaneous effect of $\Delta(ulc_t - p_t)$ was found to be insignificant, as expected.

¹¹Banerjee and Russell (2000) obtained different results on a similar specification by using quarterly seasonally adjusted data from 1972:1 to 1997:1 taken from the June 1997 OECD Data Compendium. In particular, in Banerjee and Russell (2000) the elasticities of the price level to unit labour costs and import prices are respectively 0.72 and 0.28, whereas the inflation cost coefficient is equal to -8.04 . These differences are due to several factors: a different sample, a different seasonal treatment of the data, different specification of the linear trend; moreover in Banerjee and Russell (2000) the price index was measured as the private consumption implicit deflator at factor cost.

Common wisdom contends that Italy experienced many policy regimes during the 70s, 80s and 90s, see e.g. Sarcinelli (1995) and Bertocco (2002). This equation is surprisingly stable over the whole sample period across the three main monetary policy regimes characterizing the three decades. Also during the currency crisis of 92 the relation appears stable.

In order to evaluate the stability of this relation out of sample, we also performed 1-step ahead forecasts for the next 8 quarters. The point forecasts as well as the forecast interval are reported in Fig.7. They show no apparent sign of breakdown of the relation, supporting the view that this relation may have been unaltered by the new very different institutional setup where the monetary authority has been moved to the ECB.

4.2 Wage equation

The estimated wage equation was

$$\Delta(ulc_t - p_t) = 0.12\widehat{\varepsilon}_{t-1} - 0.53\Delta\pi_{t-1} - 0.56\Delta\pi_{t-2} - 0.042 - 0.039d_{74:1t} + \widehat{\varepsilon}_{2t}$$

This equation shows a delayed negative short-run effect of the acceleration rate on the growth of real unit labor costs which may be due e.g. to rigidities in the nominal wage in the short run. By simple algebra, the wage equation can be re-written as

$$\begin{aligned} \Delta ulc_t = & \pi_t - 0.53\pi_{t-1} - 0.03\pi_{t-2} + 0.56\pi_{t-3} \\ & - 0.042 + 0.12\widehat{\varepsilon}_{t-1} - 0.039d_{74:1t} + \widehat{\varepsilon}_{2t} \end{aligned}$$

where contemporaneous and delayed values of the inflation rate may proxy the effect of the expected inflation rate, $E_t\pi_{t+1}$, on real wage growth. The adjustment effect to the disequilibria in the long run relation are of the expected sign. They confirm a wage-price adjustment to the same long-run relation.

We also tested stability of this equation; it appears reasonably stable over most of the sample, possibly showing signs of instability at the very end of the sample. This seems to confirm that, also during the currency devaluation of 1992 which prompted the end of the ‘scala mobile’, the adjustment from prices to unit labor cost was stable.

We also used the estimated equation to forecast $ulc_t - p_t$ out of sample. Fig. 7 reports 1-step ahead forecasts. This exercise did not show signs of mis-prediction.

4.3 Imported price equation

The estimated import price equation was

$$\Delta(pm_t - p_t) = 0.37\Delta(pm_{t-1} - p_{t-1}) - 0.003 + 0.187d_{74:1t} + \widehat{\varepsilon}_{3t}.$$

As expected, the price of imports does not depend on national variables, and it does not adjust to the internal price equilibrium relation. This relation shows weak signs of instability around the currency crises of 1986 and 1992, as shown by the graphs of recursive residuals and Chow tests. However also the forecast of this equation does not show signs of instability out of sample, see Fig. 7.

Overall the results obtained from the (cointegrated) error correcting model appear to well describe the quarterly dynamics of prices in the last thirty years. This

evidence shows that pure monetarist theories are not able to account for the inflation mechanisms Italy experienced from the seventies up to Stage III of European Monetary Union, as argued in detail in the following section.

5 Policy implications

The empirical evidence presented in the previous sections has important implications regarding the effectiveness of various economic policies aimed at curbing inflation. We list several implications, not necessarily in order of importance.

First of all the empirical analysis reveals that the accommodating attitude of monetary authorities cannot fully account for the stages of inflation expansion experienced by Italy in the seventies. We found in fact that a stable long-run relation connects prices and marginal costs, and that inflation error-corrects towards this equilibrium. While the present analysis cannot disentangle the relative importance of supply- versus demand-side factors in inflation dynamics, it does give clear evidence of cost-push, supply-side effects in the long run.

The monetarist tenet is usually summarized in the statement that the decisions of firms and trade union are irrelevant for inflation in the long run provided they do not affect the growth of money, see e.g. Friedman (1969). This can be reconciled with the evidence on the cointegrating equation (7) only if some monetary aggregate has a common trend with $(ulc_t - p_t) + 0.237(pm_t - p_t)$, which is not expected.

Secondly, the empirical evidence suggests that monetary policy affects the inflation rate through its impact on factors influencing production costs (e.g., foreign exchange rate and the level of aggregate demand). Note however, that this effect does not directly involve monetary aggregates.

According to the model estimated above, production costs influence inflation; because income and fiscal policies are usually designed to affect production costs, manoeuvres aimed at curbing inflation based on monetary, income and fiscal policies are likely to have a relevant (combined) effect on the growth of prices.

This finding stresses the importance of policies adopted in September 1992 by the Amato government, who – taking advantage of the currency crisis – succeeded in adopting a 92,000 billion lire manoeuvre aimed at setting public accounts to equilibrium. In subsequent years, similar manoeuvres made it possible to progressively reduce the public deficit/GDP ratio from over 10 percent in 1991 to below 3 percent as required by the Maastricht Treaty. In July 1993, an agreement containing income policy measures was reached by the government with trade unions. This agreement completed the one signed in July 1992, which had eliminated the wage indexation system based on the so-called ‘scala mobile’ (cost-of-living adjustment).

Third, these findings cast doubts on the new classic macroeconomics view that the decisions of firms, trade unions and public sector are influenced by the strength of the anti-inflationary commitment of the central bank, see e.g. Cukierman (1992) for a significant example of this approach. At first sight, this claim would appear to be confirmed by the pattern of inflation in Italy in the three decades of the 20th century, with high levels of inflation and expansionist monetary policy in the ‘70s, moderate inflation and exchange rate targeting in the ‘80s and strict monetary policy and more stable inflation in the ‘90s; see Bertocco (2002) and Fratianni and Spinelli (2001).

However, this description does not explain why, in the face of very different monetary policies, firms and trade unions had a remarkably stable behavior supporting the estimated cointegrating relation (1). The decisions of firms, trade unions and public sector do not appear to be influenced by the strength of the anti-inflationary commitment of the central bank except through the realized current inflation rate level.

Moreover, common wisdom contends that economic policies were more effective in the '90 than in the '80. Policies in the '80 were only based on monetary instruments, while the policies of the '90 included also income and fiscal interventions. This observation contrasts with the interpretation of inflation as a pure monetary phenomenon. The anti-inflationary stance of monetary policy was very clear-cut in the '80, following the adhesion of Italy to the EMS in 1979. The target of a stable exchange rate was an element of discipline for firms which could no longer rely on devaluation, and had to protect their competitiveness through the control of costs. The monetary authorities protected the exchange rate by using interest rates.

Hence, according to Cukierman (1992), the strong anti-inflationary commitment of the monetary authorities should have had a significant impact on the behavior of workers and of the public sector. This did not happen. On the contrary, the absence of significant income and fiscal policy measures forced the monetary authority to raise interest rates to particularly high levels. This led to the disequilibria that are the basic cause of the devaluation of the lira in September 1992, see e.g. Bertocco (2002) and reference therein. Paradoxically, in coincidence with the devaluation, which marked the failure of the anti-inflationary policy based on the stabilization of the foreign exchange rate, the conditions were laid for the adoption of the income and fiscal policy measures which made it possible to reduce inflation notwithstanding the devaluation of the lira.

The experience in the '80s points out the limits of an anti-inflationary manoeuvre exclusively based on monetary policy; on the other hand, the experience in the '90 shows that an anti-inflationary manoeuvre based on the simultaneous recourse to monetary policy, fiscal policy and income policy was effective.

The presence of a stable long run relation connecting the Italian price index, unit labor cost and import prices, appears consistent with this different characteristics of the anti-inflationary manoeuvre realized in Italy in '80 and '90; a mix of monetary, fiscal and income policies has combined effect in the model estimated in the previous sections.

6 Conclusions

Empirical evidence based on cointegration techniques and error correction models has been used to estimate the relationship between markup and inflation dynamics. The results obtained in the paper suggest that a cost-push model can be used to fit a stable relation across different monetary regimes during the 1970-1998 period. The proposed model does not rule out that short run deviations in the markup might be due to shocks stemming from aggregate demand and in particular from monetary disturbances. However, this evidence points that pure monetarist theories of inflation are not able to account for the inflation mechanisms Italy experienced from the seventies up to Stage III of European Monetary Union.

Appendix

A1: The VAR model

Let $X_t = (X_{1t}, X_{2t}, X_{3t})'$; we consider the VAR model

$$A(L)X_t = \mu_t + \varepsilon_t \quad (8)$$

where k is the lag length, $A(L) = I - \sum_{i=1}^k A_i L^i$, A_1, \dots, A_k are 3×3 matrices of parameters, $\mu_t = \Phi D_t + \mu$, μ is a 3×1 vector of constants, D_t contains additional deterministic variables, Φ is a matrix of parameters associated with deterministic variables and $\varepsilon_t = (\varepsilon_{1t}, \varepsilon_{2t}, \varepsilon_{3t})'$ is a White Noise with 3×3 covariance matrix $\Omega = (\sigma_{ij})$. Let $\Pi := -A(1)$; if the system is I(1) then $\Pi = \alpha\beta'$ and the system can be rewritten in the Vector Equilibrium Correction (VEqC) form

$$\Delta X_t = \alpha\beta' X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu_t + \varepsilon_t \quad (9)$$

where $\Gamma_i = -(A_{i+1} + \dots + A_k)$, and $e_t = \beta' X_t$ are the stationary cointegrating relations. The coefficients in α measure the adjustment of variables to the long run equilibria.

The joint system can be decomposed into a conditional model for ΔX_{1t} given the $\Delta X_{2t}, \Delta X_{3t}$ with equation

$$\Delta X_{1t} = \omega_1 \Delta X_{2t} + \omega_2 \Delta X_{3t} + \alpha\beta' X_{t-1} + \sum_{i=1}^{k-1} \Gamma_{1i.2} \Delta X_{t-i} + \mu_{1.2t} + \varepsilon_{1.2t} \quad (10)$$

where $\omega = (\omega_1, \omega_2) = (\sigma_{12}, \sigma_{13})\Omega_{22}^{-1}$, $(\alpha, \Gamma_{1i.2}, \mu_{1.2t}, \varepsilon_{1.2t}) = (1, -\omega)(\alpha, \Gamma_i, \mu, \varepsilon_t)$,

$$\Omega_{22} = \begin{pmatrix} \sigma_{22} & \sigma_{23} \\ \sigma_{23} & \sigma_{33} \end{pmatrix}$$

and a marginal system with equations

$$\Delta X_{2t} = \alpha_2 \beta' X_{t-1} + \sum_{i=1}^{k-1} \Gamma_{2i} \Delta X_{t-i} + \mu_{2t} + \varepsilon_{2t} \quad (11)$$

$$\Delta X_{3t} = \alpha_3 \beta' X_{t-1} + \sum_{i=1}^{k-1} \Gamma_{3i} \Delta X_{t-i} + \mu_{3t} + \varepsilon_{3t} \quad (12)$$

For a reference on the marginal-conditional decomposition see Johansen (1996), Chapter 8. Eq. (10) gives eq. (4), while system (11) (12) gives (5) and (6).

Further restrictions on the coefficients in (10) (11) (12) correspond to a simultaneous system of adjustment equations. In the empirical analysis we first tested for cointegration, estimated the long-run relation giving e_t , and then estimated a simultaneous system of adjustment equations by FIML for e_t fixed at \hat{e}_t .

A2: I(2)-ness in the data

A reasonable doubt may arise about the validity of the homogeneity restriction introduced in Section 3.1. If $X_t = (\Delta p_t, ulc_t - p_t, pm_t - p_t)'$ is an admissible nominal-to-real transformation of $Y_t = (ulc_t, pm_t, p_t)'$, in the sense of Kongsted (1998, 1999), then X_t must be an I(1) system. Tests on the presence of I(2) in $X_t = (\Delta p_t, ulc_t - p_t, pm_t - p_t)'$ are test of homogeneity restriction, i.e. the nominal-to-real transformation. The I(2) tests in Paruolo (1996) have been performed on the system X_t , and no evidence of I(2) was found. This showed the admissibility of the nominal-to-real transformation, i.e. of the homogeneity restriction.

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Figures

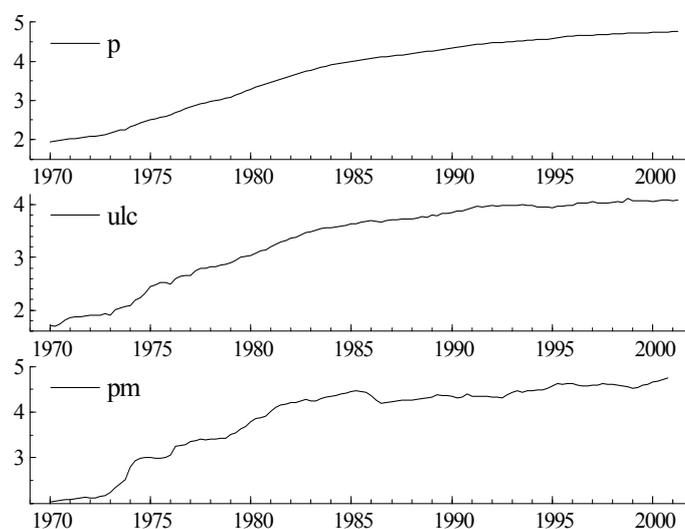


Figure 1. Levels of p_t , ulc_t , and pm_t ; 1970:1 - 2000:4.

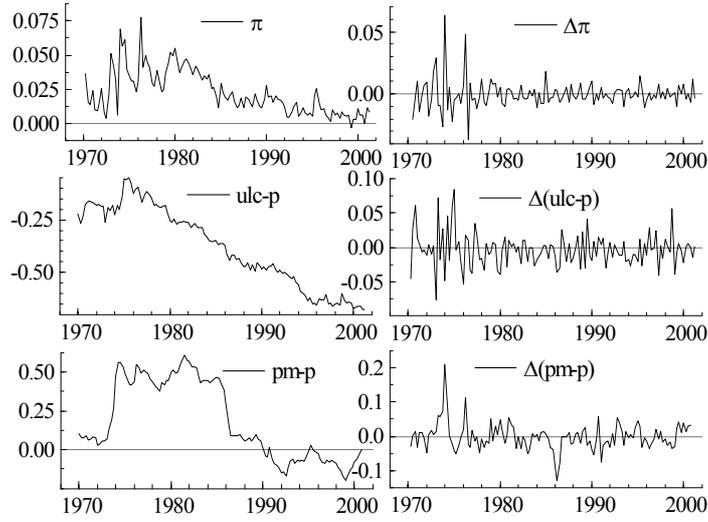


Figure 2. Levels (left panel) and first differences (right panel) of π_t , $ulc_t - p_t$, and $pm_t - p_t$; 1970:2 - 2000:4.

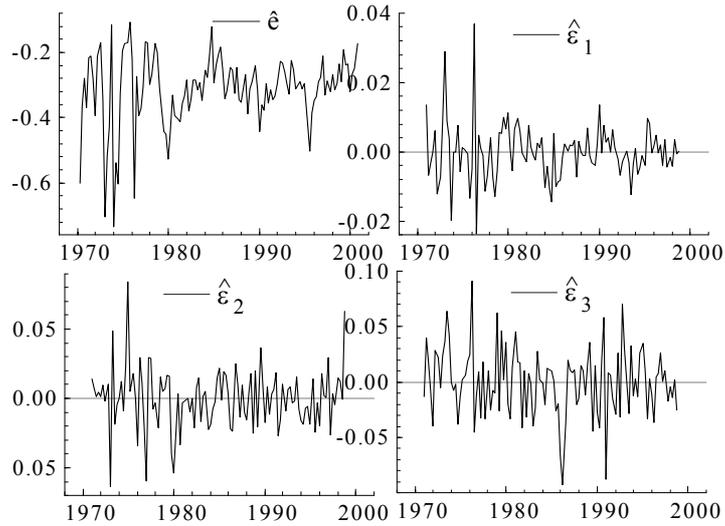


Figure 3. Estimated disequilibrium term \hat{e}_t and residuals \hat{e}_{it} , $i = 1, 2, 3$ from the cointegrated VAR with 1 cointegrating vector, see Eq. (3); 1971:1-1998:4.

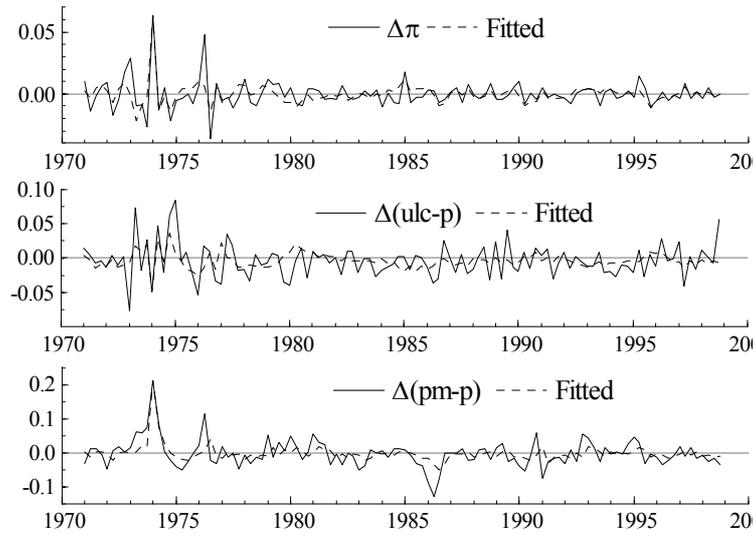


Figure 4. Actual and fitted values; 1971:1-1998:4.

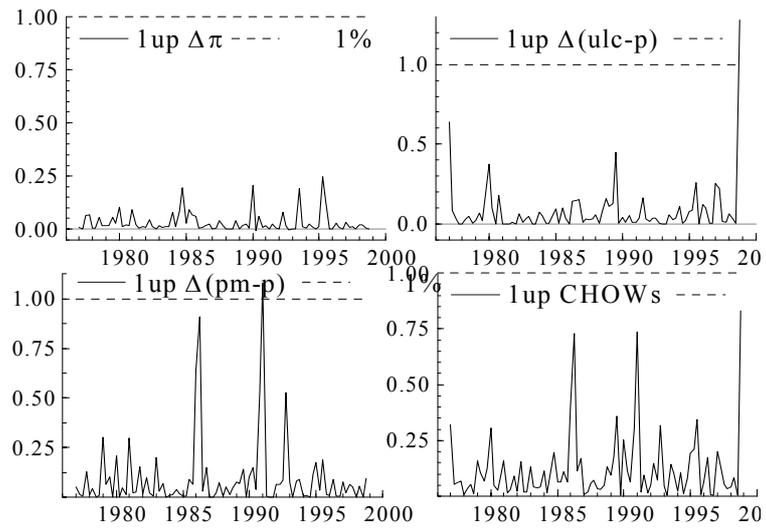


Figure 5. 1-step Chow tests for the single equations and for the system as a whole; 1971:1-1998:4.

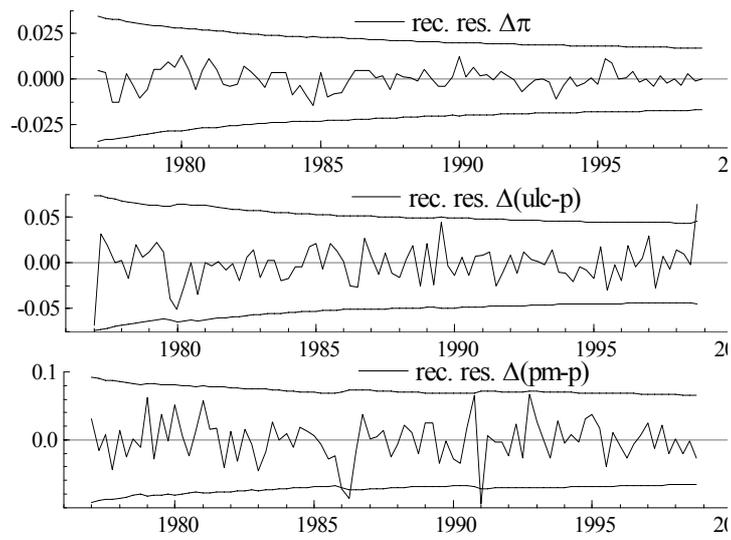


Fig 6 1-step recursive residuals VEqM; 1971:1-1998:4.

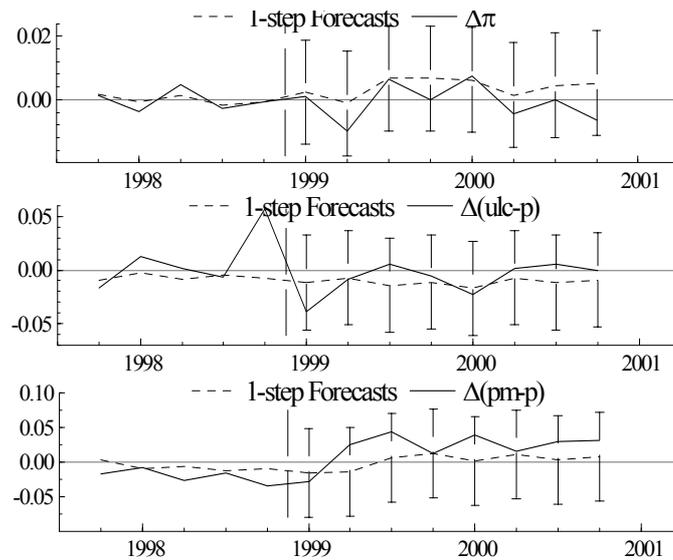


Fig 7 1-step ahead forecasts and forecast 95% confidence interval; forecast period: 1999:1-2000:4.

Tables

j	Trace test	p -value
0	64.738	0.000
1	18.905	0.292
2	7.1964	0.334

Table 1. Cointegration test in the restricted trend I(1) model, trace statistic; $H_0 : r \leq j$; 1971:1-1998:4.

regressor	Inflation coefficient	equation std.error	$\Delta\pi_t$ t -value	p -value
$\Delta(pm_t - p_t)$	0.100742	0.06064	1.66	0.100
$\Delta(pm_{t-2} - p_{t-2})$	0.0363442	0.02185	1.66	0.099
e_{t-1}	-0.0576095	0.008161	7.06	0.000
Constant	0.0182333	0.002809	6.49	0.000
$d_{74:1}$	0.0279269	0.01568	1.78	0.078
eq. SE	0.00822958			

regressor	wage coefficient	equation std.error	$\Delta(ulc_t - p_t)$ t -value	p -value
$\Delta\pi_{t-1}$	-0.530549	0.2698	-1.97	0.052
$\Delta\pi_{t-2}$	-0.558366	0.2194	-2.55	0.012
e_{t-1}	0.121972	0.02748	-4.44	0.000
Constant	-0.0422377	0.008998	-4.69	0.000
$d_{74:1}$	-0.0390766	0.02291	-1.71	0.091
eq. SE	0.0220006			

regressor	import price coefficient	equation std.error	$\Delta(pm_t - p_t)$ t -value	p -value
$\Delta(pm_{t-1} - p_{t-1})$	0.370656	0.07678	4.83	0.000
Constant	-0.00329831	0.003050	-1.08	0.282
$d_{74:1}$	0.187006	0.03274	5.71	0.000
eq. SE	0.0320658			

Table 2: Simultaneous system of adjustment equations, estimated by FIML; 1971:1-1998:4.