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MEASURING CORE INFLATION IN ITALY

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In this paper the long-run trend in *CPI* inflation (*core inflation*) for Italy is estimated over the 1962-1997 period within the framework of a multivariate common trends model. In this framework core inflation is directly linked to money and wage growth and interpreted as the long-run forecast of inflation. This measure displays several desirable properties: lower variability than observed inflation, forecasting power, robustness to the estimation sample and economic interpretability.

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INTRODUCTION

In the recent debate about monetary policy targets, one prominent view favours the direct formulation of the central bank's objective in terms of the ultimate policy goal, price stability. Inflation targeting policies, setting precise quantitative targets for monetary authorities, have been advocated and recently implemented in several countries (see Bernanke - Laubach - Mishkin - Posen (1999) for a cross-country assessment and a review of the main implementation issues). Though the essence of inflation-targeting policies can be simply stated, qualifications are needed once it is recognized that observed

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inflation may fluctuate in the short-run due to only temporary disturbances of both real and nominal nature.

In fact, monetary policy should try and control the underlying, trend behaviour of inflation, not reacting to the short-run noisy fluctuations in the measured inflation rate. Moreover, the relatively long time needed for a monetary policy action to affect real activity and inflation suggests that an appropriate target should be expressed in terms of a (relatively long-run) forecast of inflation rather than in terms of the currently measured inflation rate (Svensson (1997)). How to construct a reliable empirical measure of the underlying, *core* inflation rate is therefore a crucial issue in monetary policy design, also for those monetary authorities, such as the European Central Bank, that have recently emphasized the goal of price stability, though not adopting explicitly an inflation-targeting policy.

Several methods for constructing measures of the underlying inflation rate have been proposed and implemented. The simplest one is the calculation of moving averages over a variable time span (from 3-6 months up to 36 months) or more sophisticated filters (e.g. the Kalman filter), in order to smooth and reduce the noise component in the inflation pattern.

Bryan - Cecchetti (1993, 1994) advocated the use of limited influence estimators, such as trimmed means or the (weighted) median, in the place of the conventional weighted mean calculated over the complete cross-sectional distribution of the individual price components. Given the asymmetry of the cross-sectional distribution of the individual price components, the median or a trimmed mean is likely to be a more efficient estimator of the central tendency of the distribution than the conventional mean. In addition, using trimmed means or the median may be seen as a better practice than excluding from the *CPI* some categories of goods, such as energy and food (at home), which are believed to be high-variance noise components. In fact, over time there is not guarantee that these goods may not turn into low-variance components[1]. However, those purely statistical measures do not directly link core inflation to what theory identifies as its main determinant: money growth.

In this paper we make this step explicitly by deriving a measure of core inflation from a small-scale macroeconomic common trends model, including (some of) the main determinants of the inflationary process, namely

[1] In its latest *Annual Report*, the Bank of Italy explicitly indicated the changes in the *CPI* price level excluding food and energy goods as its preferred measure of the underlying inflation rate (Bank of Italy (1999)). The Bryan-Cecchetti methodology has been applied to Italian data by Gavosto - Sabbatini (1997) over the 1978-1994 period.

money and wage growth and the behaviour of commodity prices as a proxy for supply-side factors. This methodology is applied to the Italian inflation data over a sample period spanning more than three decades, from 1962 to 1997.

The main advantages of this common trend measure of core inflation over purely statistical alternatives are twofold: first, it is a *forward-looking* measure, capturing the long-term element in the inflation process (according to the long-run properties of the macroeconomic system, that will be appropriately estimated and tested); second, it has some *theoretical basis*, linking inflation to money and wage growth[2].

The rest of the paper is organised as follows. In Section 1 we outline the common trends approach to core inflation estimation and report the empirical results in Section 2. The properties of the estimated core inflation process and its suitability for monetary policy purposes are discussed in Section 3. Section 4 briefly concludes.

1. THE COMMON TRENDS APPROACH TO CORE INFLATION ESTIMATION

To construct a forward-looking measure of the core inflation process, we adopt the framework of the common trends model of Stock - Watson (1988), King - Plosser - Stock - Watson (1991), Mellander - Vredin - Warne (1992) and Warne (1993).

In this framework inflation and output are analysed in a system including several other macroeconomic quantities that we see as important determinants of the inflation rate, e.g. nominal money and wage growth. The existence of long-run equilibrium (cointegrating) relationships among the (non-stationary) variables reduces the number of independent disturbances having permanent effects on the level of the series. The common trends representation of the system allows to decompose the variables into a non-stationary (stochastic) trend component and a stationary transitory element. The former component captures the effect of only the permanent shocks and bears the interpretation of long-run forecast of the endogenous variables in the system. In the case of consumer prices, the common trends representation may provide the appropriate measure of the long-run trend inflation that monetary authorities should regard as their main target variable. In the

[2] In a thorough comparison of several measures of core inflation, WYNNE (1999) has recently argued that only measures derived from *VAR* systems can possess these two desirable properties.

next two sub-sections we summarize the econometric methodology employed.

1.1 *The econometric methodology*

Following Warne (1993) and Gallo - Kempf (1996), let us start from the unrestricted $VAR(p)$ representation of a vector $\mathbf{x}_t$ of n $I(1)$ variables of interest, written in levels and in the Engle - Granger (1987) error-correction ($VECM$) form:

$$\mathbf{x}_t = \mathbf{\Pi}(L)\mathbf{x}_{t-1} + \varepsilon_t \tag{2.1}$$

$$\Delta\mathbf{x}_t = \mathbf{\Pi}^*(L)\Delta\mathbf{x}_{t-1} + \mathbf{\Pi}(1)\mathbf{x}_{t-1} + \varepsilon_t \tag{2.2}$$

where ε_t is a vector of identically and independently distributed, serially uncorrelated disturbances with zero mean and variance-covariance matrix $\mathbf{\Sigma}$, $\mathbf{\Pi}(L) = \mathbf{\Pi}_1 + \mathbf{\Pi}_2 L + ... + \mathbf{\Pi}_p L^{p-1}$, $\mathbf{\Pi}(1) = \sum_{i=1}^{p} \mathbf{\Pi}_i$, $\mathbf{\Pi}^*(L) = \mathbf{\Pi}_1^* + \mathbf{\Pi}_2^* L + ... + \mathbf{\Pi}_{p-1}^* L^{p-2}$ and $\mathbf{\Pi}_i^* = -\mathbf{I} + \mathbf{\Pi}_1 + ... + \mathbf{\Pi}_i$ $(i = 1, ..., p-1)$[3].

If there are $0 < r < n$ cointegrating relations among the variables, $\mathbf{\Pi}(1)$ is of reduced rank r and can be expressed as the product of two $(n \times r)$ matrices: $\mathbf{\Pi}(1) = \alpha\beta'$, where β contains the cointegrating vectors, such that $\beta'\mathbf{x}_t$ are stationary linear combinations of the $I(1)$ variables, and α is a matrix of factor loadings. The resulting cointegrated VAR is then:

$$\Delta\mathbf{x}_t = \mathbf{\Pi}^*(L)\Delta\mathbf{x}_{t-1} + \alpha\beta'\mathbf{x}_{t-1} + \varepsilon_t \tag{2.3}$$

Following the procedure set out in Mellander - Vredin - Warne (1992), the cointegrated VAR in (3) can be inverted to yield the following stationary Wold representation for $\Delta\mathbf{x}_t$:

$$\Delta\mathbf{x}_t = \mathbf{C}(L)\varepsilon_t \tag{2.4}$$

where $\mathbf{C}(L) = \mathbf{I} + \mathbf{C}_1 L + \mathbf{C}_2 L^2 + ...$ with $\sum_{j=0}^{\infty} j \mid \mathbf{C}_j \mid < \infty$. From the representation in (2.4) the following expression for the levels of the variables can be derived by recursive substitution:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{C}(1)\sum_{j=0}^{t-1} \varepsilon_{t-j} + \mathbf{C}^*(L)\varepsilon_t \tag{2.5}$$

where $\mathbf{C}^*(L) = \sum_{j=0}^{\infty} \mathbf{C}_j^* L^j$ with $\mathbf{C}_j^* = -\sum_{i=j+1}^{\infty} \mathbf{C}_i$. $\mathbf{C}(1)$ captures the long-run effect of the reduced form disturbances in ε on the variables in $\mathbf{x}$. The existence of r cointegrating vectors implies that the long-run matrix $\mathbf{C}(1)$ has rank $n - r \equiv k$ and $\beta'\mathbf{C}(1) = \mathbf{0}$.

[3] For ease of exposition, we do not include a constant term in (2.1) and (2.2), that would add a deterministic time trend in the representation for the levels below.

In order to obtain an economically meaningful interpretation of the dynamics of the variables of interest from the reduced form representations in (2.4) and (2.5), the vector of reduced form disturbances ε must be transformed into a vector of underlying, "structural" shocks, some of which with *permanent* effects on the level of $\mathbf{x}$ and some with only *transitory* effects on the short-run dynamics. Let us denote this vector of structural disturbances as $\varphi_t \equiv \begin{pmatrix} \psi_t \\ \nu_t \end{pmatrix}$, where ψ and ν are subvectors of k and r elements respectively. The structural form for the first difference of $\mathbf{x}_t$ is:

$$\Delta \mathbf{x}_t = \boldsymbol{\Gamma}(L)\varphi_t \tag{2.6}$$

where $\boldsymbol{\Gamma}(L) = \boldsymbol{\Gamma}_0 + \boldsymbol{\Gamma}_1 L + \ldots$ and the previously defined vector φ_t is identically and independently distributed, serially uncorrelated, with zero mean and an identity variance-covariance matrix. The relationship between the reduced form and the structural shocks is given by:

$$\varepsilon_t = \boldsymbol{\Gamma}_0 \varphi_t \tag{2.7}$$

where $\boldsymbol{\Gamma}_0$ is an invertible matrix. Hence, comparison of (2.6) and (2.4) shows that

$$\mathbf{C}(L)\boldsymbol{\Gamma}_0 = \boldsymbol{\Gamma}(L)$$

implying that $\mathbf{C}_i\boldsymbol{\Gamma}_0 = \boldsymbol{\Gamma}_i$ $(\forall i > 0)$ and $\mathbf{C}(1)\boldsymbol{\Gamma}_0 = \boldsymbol{\Gamma}(1)$. In order to identify the elements of ψ_t as the permanent shocks and the elements of ν_t as the transitory disturbances the following restriction on the long-run matrix $\boldsymbol{\Gamma}(1)$ must be imposed:

$$\boldsymbol{\Gamma}(1) = \begin{pmatrix} \boldsymbol{\Gamma}_g & \mathbf{0} \end{pmatrix} \tag{2.8}$$

with $\boldsymbol{\Gamma}_g$ an $n \times k$ submatrix. The disturbances in ψ_t are then allowed to have long-run effects on (at least some of) the variables in $\mathbf{x}_t$, whereas the shocks in ν_t are restricted to have only transitory effects.

From (2.6), the structural form representation for the endogenous variables in levels is derived as

$$\begin{aligned} \mathbf{x}_t &= \mathbf{x}_0 + \boldsymbol{\Gamma}(1)\sum_{j=0}^{t-1}\varphi_{t-j} + \boldsymbol{\Gamma}^*(L)\varphi_t = \\ &= \mathbf{x}_0 + \boldsymbol{\Gamma}_g \sum_{j=0}^{t-1}\psi_{t-j} + \boldsymbol{\Gamma}^*(L)\varphi_t \end{aligned} \tag{2.9}$$

where the partition of φ and the restriction in (2.8) have been used and $\boldsymbol{\Gamma}^*(L)$ is defined analogously to $\mathbf{C}^*(L)$ in (2.5). The permanent part in (2.9),

$\sum_{j=0}^{t-1} \psi_{t-j}$, may be expressed as a k-vector random walk with innovations ψ:

$$\begin{aligned}\tau_t &= \tau_{t-1} + \psi_t = \\ &= \tau_0 + \sum_{j=0}^{t-1} \psi_{t-j}\end{aligned} \tag{2.10}$$

Using (2.10) in (2.9) we finally obtain the common trend representation for $\mathbf{x}_t$:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{\Gamma}_g \tau_t + \mathbf{\Gamma}^*(L)\varphi_t \tag{2.11}$$

The identification of separate permanent shocks requires restrictions on the long-run impact matrix $\mathbf{\Gamma}_g$ in the common trend model (2.11). Moreover, separate transitory shocks can be identified by making assumptions on their contemporaneous impact on the endogenous variables (captured by the elements of the last r columns of $\mathbf{\Gamma}_0$). Under these restrictions all shocks can be given an economic interpretation.

To estimate the $n \times k$ matrix $\mathbf{\Gamma}_g$, we need (at least) nk restrictions on its elements. Cointegration implies:

$$\beta' \mathbf{\Gamma}_g = 0 \tag{2.12}$$

(since $\beta' \mathbf{\Gamma}(1) = \beta' C(1) \mathbf{\Gamma}_0 = 0$), yielding rk linear restrictions. Moreover, from (5) and (10) we find that $\mathbf{C}(1)\varepsilon_t = \mathbf{\Gamma}_g \psi_t$. Hence, since $E(\psi_t \psi_t') = \mathbf{I}$ and $\mathbf{C}(1)$ has reduced rank k, an additional $k(k+1)/2$ restrictions on the elements of $\mathbf{\Gamma}_g$ are provided by:

$$\mathbf{C}(1)\Sigma C(1)' = \mathbf{\Gamma}_g \mathbf{\Gamma}_g' \tag{2.13}$$

The remaining $k(k-1)/2$ restrictions needed for (exact) identification of $\mathbf{\Gamma}_g$ have to be derived from economic theory. The elements of $\mathbf{C}(1)$ and $\mathbf{\Sigma}$ can be consistently estimated from the *VAR* model and $\mathbf{\Gamma}_g$ can be obtained from imposition of a sufficient number of restrictions. The structural permanent shocks can then be constructed using their relation with the *VAR* residuals: $\mathbf{C}(1)\varepsilon_t = \mathbf{\Gamma}_g \psi_t$. We get: $\psi_t = (\mathbf{\Gamma}_g \mathbf{\Gamma}_g')^{-1} \mathbf{\Gamma}_g' \mathbf{C}(1)\varepsilon_t$. The behaviour of the variables in $\mathbf{x}_t$ due to the permanent disturbances, interpreted as the long-run forecast of $\mathbf{x}_t$, may then be computed as $\mathbf{x}_0 + \mathbf{\Gamma}_g \sum_{j=0}^{t-1} \psi_{t-j}$.

Estimation of the common trend model is discussed in detail in Stock - Watson (1988), Mellander - Vredin - Warne (1992) and Warne (1993). From the moving average representation in (2.6) impulse response and forecast error variance decompositions may be calculated with respect to permanent and transitory innovations.

1.2 *Modelling Structural Instability*

A major issue in the specification of the *VAR* model is structural stability, especially when, as in our case, the sample spans several decades. The empirical analysis of the *VAR* over the whole sample shows that instability is mainly due to shifts in the variance-covariance matrix of residuals, concentrated in periods of high volatility of commodity and domestic prices. Dummy variables could be added to the *VAR* specification to take care of outlier observations and account for shifts in the residual variance pattern. In this paper we take a different approach and try to endogenously detecting and modelling structural breaks, adopting a time-varying specification governed by a Markov-switching mechanism. With respect to the inclusion of dummy variables from the start, this approach is preferable since it does not require a priori assumptions about the exact timing of breaks and their actual specification in terms of impulse or step dummies.

The approach relies on the idea that the observations can be partitioned in several groups according to their generating regime. Since in our empirical analysis we find that a two-regime process yields a sufficiently stable description of the data, the following methodological discussion refers to only two regimes, differing in terms of both conditional mean and variance of the generating mechanism, with smooth transition between conditional means and perfect correlation of regimes across the variables.

Given our n $I(1)$ endogenous, cointegrated variables subject to regime shifts, the conditional probability density of the observed time series vector $\mathbf{x}_t$ may be written as:

$$p(\mathbf{x}_t \mid \mathbf{X}_{t-1}, s_t) = \begin{Bmatrix} f(\mathbf{x}_t \mid \mathbf{X}_{t-1}, \theta_1) & \text{if } s_t = 1 \\ f(\mathbf{x}_t \mid \mathbf{X}_{t-1}, \theta_2) & \text{if } s_t = 2 \end{Bmatrix}$$

where $s_t = 1,2$ indicates the feasible regimes, θ_m is the parameter vector in regime $m = 1,2$ and $\mathbf{X}_{t-1}$ are past observations of $\mathbf{x}$. The regime s_t can be modelled according to the following discrete-state homogeneous Markov-chain generating mechanism:

$$p\left(s_t \mid \{s_{t-j}\}_{j=1}^{\infty}, \mathbf{X}_{t-1}\right) = p(s_t \mid s_{t-1})$$

The probability of being in a given regime at date t depends exclusively on the prevailing regime one period earlier, according to transition probabilities to be estimated.

The error-correction representation of the *VAR* model in (2) is then extended to allow for regime dependence of both the intercept and the error variance-covariance matrix, resulting in a Markov-switching vector error-correction (*MS-VECM*) model of the form:

$$\Delta \mathbf{x}_t = \nu(s_t) + \mathbf{\Pi}^*(L)\Delta \mathbf{x}_{t-1} + \mathbf{\Pi}(1)\mathbf{x}_{t-1} + \varepsilon_t \tag{2.14}$$

where ν is an intercept vector and $\varepsilon_t \sim NID(\mathbf{0}, \mathbf{\Sigma}(s_t))$. Following Krolzig (1997), estimation is carried out in two stages. In the first stage the long-run equilibrium relationships are estimated by means of the Johansen (1988, 1991) approach ignoring the presence of different regimes. In the second stage the estimated error-correction terms are included as exogenous regressors in the specification, and the *MS-VECM* (2.14) is estimated via the Expectation-Maximization (*EM*) algorithm[4]. This procedure yields estimates of the intercept vectors ν and the matrices $\mathbf{\Sigma}$ pertaining to the two regimes, and also an estimate of the probability that each observation is generated by each of the two regimes. On the basis of these probabilities we partition the sample between periods belonging to the first regime and observations generated by the second regime.

The common trend analysis is then carried out, along the lines set out in the previous sub-section, with reference to the regime prevailing in most of the sample (in our case comprising around 85% of the observations). To this aim, we impose in the common trend estimation, the variance-covariance matrix obtained for the "dominant" regime and include in the *VECM* a set of dummy variables corresponding to the dates belonging to the alternative regime.

2 Empirical results

2.1 *Data*

We apply the estimation method described above to a small-scale macroeconomic system including, beside the inflation rate, a measure of economic activity and three main determinants of inflation behaviour, namely money and wage growth rates and oil prices. All data are monthly, seasonally adjusted (with the exception of oil prices) over the period 1962(9)-1997(12). The five variables in the system are then the following:

π: price inflation measured as the monthly rate of change of the (all-item) *CPI index;*

y: the log of the industrial production index;

m: the rate of growth of nominal *M2;*

[4] The robustness of the Johansen estimator with respect to the possible omission of regime shifts has been defended theoretically in a number of previous studies (see Krolzig (1997) and the references therein). In the following analysis we provide additional empirical support for this view.

w: the rate of growth of nominal wages for the entire economy;
oil: the log of the price of oil in Italian liras[5].

Fig. 1 shows the behaviour of the monthly and annual *CPI* inflation rates, the latter computed as a twelve-month lagged moving average of the monthly rate. The variables are chosen so as to capture both domestic and foreign effects on the inflation rate. Although, for most of the time span considered, Italy has participated to agreements aimed at stabilising exchange rate fluctuations (so that nominal money growth and inflation were strongly related to exchange rate dynamics), the frequent devaluations undertaken over the 1980s and the large fluctuation bands employed for a long time during the *EMS* experience, may suggest that the Italian inflation process had

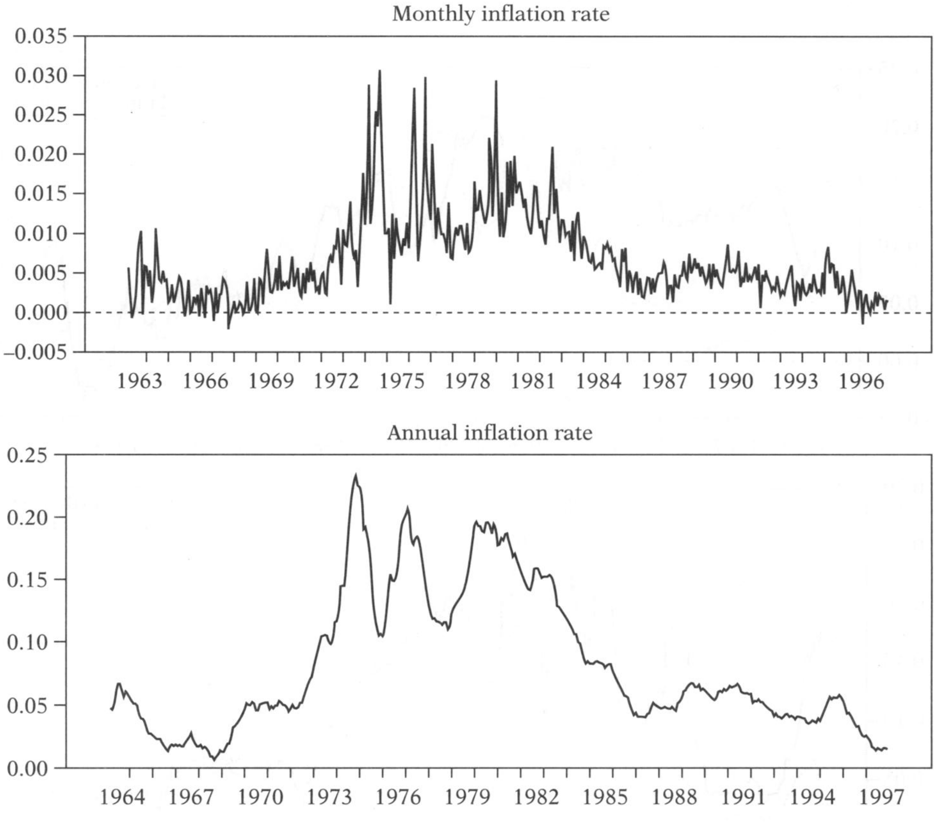

Fig. 1 - The Italian Inflation Rate 1963-1997

[5] Data sources are: *ISTAT* (π, y and w), BANK OF ITALY (m) and DATASTREAM (*oil*).

a strong domestic component. It is true, however, that with the convergence process and the macroeconomic harmonisation leading to the European Monetary Union, the role of domestic factors may have lost some importance in the more recent period and will continue to do so in the future. Moreover, the mechanism of wage indexation (abolished only few years ago) has been an important determinant of the wage-price inflation spirals occurred in Italy in the 1970s and 1980s, suggesting that a joint modelling of wage inflation, price inflation and nominal money growth may be appropriate for core inflation estimation over the time span considered. The behaviour of the annual growth rates of nominal *M*2 and nominal wages is shown in Fig. 2 together with the annual rate of inflation to highlight the long-run comovement of the series.

Foreign effects on the inflation rate have been explicitly introduced in

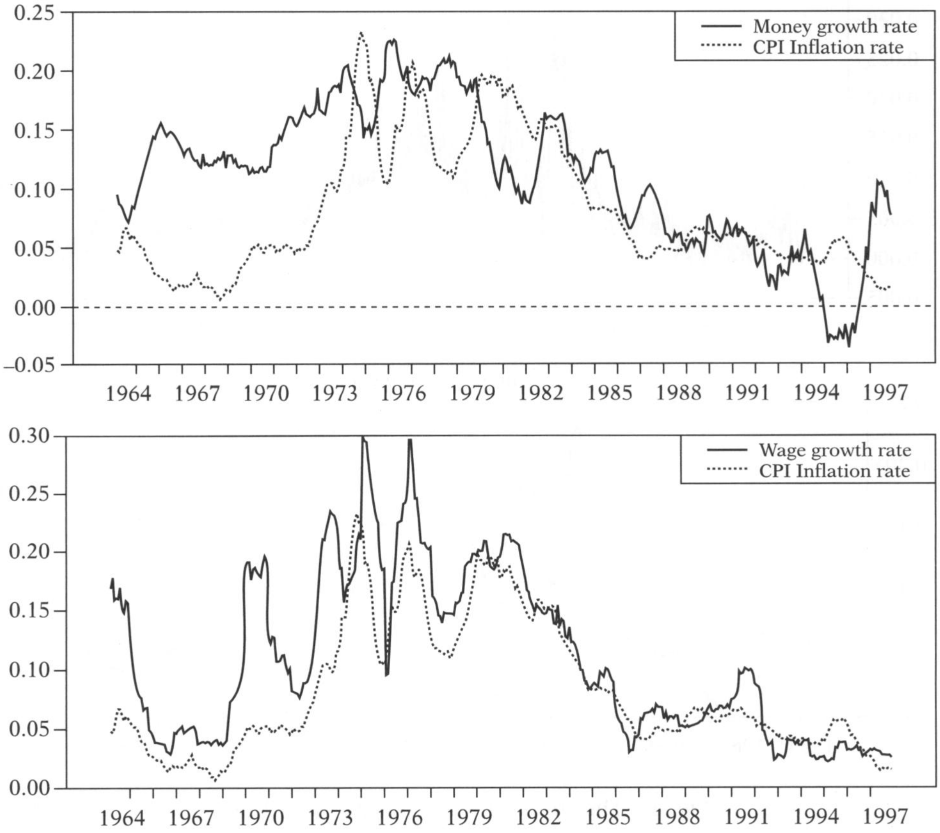

Fig. 2 - Annual Growth Rates of M2, Wages and CPI Price Level.

the model through the oil price variable, capturing a major source of supply side shocks. This should provide valuable additional information about the determinants of long-run inflation. Furthermore, the choice of the growth rate of the *CPI* as a measure of price inflation (instead of the *GDP* deflator, for example) allows for the contribution of external factors to explain core inflation dynamics.

Another feature of the relatively long sample analyzed is the presence of different monetary policy regimes, which in principle can alter the relation between money growth, inflation and output. In particular, two clear policy regime changes occurred in 1969-70 and in 1981. At the end of 1969, following a generalized increase in foreign interest rates and a surge in world commodity prices, the Bank of Italy abandoned the interest rate stabilization policy closely followed in the sixties. Broad money growth gradually become an (intermediate) monetary policy target, although the autonomous monetary base creation through financing of public deficits was responsible for the frequent target misses over the period. The close link between public expenditure and monetary growth was severed at the beginning of the eighties, when the central bank started to gain independence from the Treasury in the management of monetary policy.

Those changes in the prevailing policy regime, alongwith the different nature (domestic and foreign) of the determinants of the inflation process throughout the sample period, are a potential cause of instability of the *VAR* system. Therefore, prior to applying the common trends methodology to derive the core inflation process, we tackle the stability issue.

2.2 *VAR Specification*

The time-series properties of the series are established by means of standard unit-root tests, showing that all variables can be modelled as $I(1)$ processes[6]. The vector of endogenous variables is then $\mathbf{x}_t = (oil_t y_t w_t m_t \pi_t)'$. The oil price is modelled as endogenous only to allow for the shock analysis in the common trends model of the next subsection.

A preliminary recursive analysis on the *VAR* with six lags shows several episodes of instability mainly concentrated in the oil shock periods (mid-

[6] In particular, an *ADF* test (with 6 lags) on the *CPI* inflation rate yields a statistic of -2.28 (with a 5% critical value of -2.87). When a (not significant) time trend term is added the statistic is -2.35 (-3.47). The $I(1)$ nature of π is not a general feature of the inflation rate but clearly depends on the sample chosen. In particular, as suggested by a referee, a successful inflation-targeting policy should induce stationarity in the inflation rate. This is not the case for Italy in the period under analysis.

'70s, early '80s and mid-'80s), characterized by high volatility in some of the series, not entirely captured by the variables included in the system.

To reach a sufficiently stable specification of the system we adopt the methodology described in Section 1.2, allowing the data in the sample to be generated by different generating mechanisms or "regimes". Such regimes differ in terms of both conditional means and residual variance-covariance matrices.

Following Krolzig (1997) we first carry out the cointegration analysis on the *VAR* ignoring the possibility of different regimes and find the presence of two long-run relationships. Then, we use such (unrestricted) cointegrating vectors in the Markov-switching vector-error-correction form (16), allowing initially up to five different regimes[7]. Our results support the existence of only two distinct regimes, whose main features are reported in Table 1. Since in one regime the residual variance of all equations (with the only exception of the equation for Δm) is larger, we denote the two regimes as a "low-variance " and a "high-variance" regime. The table reports the different estimated intercept and residual variance-covariance terms in the two regimes, together with the estimated transition probabilities from one regime to the other. Such probabilities show that the low-variance regime is more persistent than the high-variance one (the relevant probabilities being 0.92 and 0.55 respectively). To the low-variance regime belong 364 of the 424 months in the 1962(9)-1997(12) period analysed, with the remaining 60 observations pertaining to the high-variance regime and mainly concentrated in the oil shock periods[8]. Once the existence of these two regimes is allowed for, the resulting *VAR* is reasonably stable throughout the sample period, as shown by the *CUSUM* tests in Fig. 3, that do not suggest any further evidence of structural breaks. Moreover, the autocorrelation and normality tests (performed on the standardized residuals) reported in Table 1 do not provide evidence of mis-specification[9].

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[7] Estimation of the *MS-VECM* has been carried out using the *MSVAR* routine for *Ox* by H.-M. Krolzig.

[8] The dating of the high-variance regime is the following: 1962(12)-1963(1), 1963(8), 1964(2), 1969(3), 1969(12)-1970(1), 1970(3)-1970(4), 1973(3), 1973(7), 1973(11)-1974(3), 1974(6)-1974(9), 1974(12)-1975(4), 1976(1)-1976(4), 1976(8)-1976(9), 1977(1), 1979(4), 1979(6)-1979(8), 1979(12)-1980(1), 1980(10), 1986(1)-1986(5), 1986(7)-1986(8), 1988(10)-1988(12), 1990(7)-1991(1), 1991(4), 1991(11), 1994(4), 1994(8).

[9] When the two-regime *MS-VECM* specification is tested against a constant-parameter *VAR*, a likelihood-ratio test yields a clear rejection of the latter. The corresponding test-statistic for the 22 restrictions (15 for the variances and covariances, 5 for the constant terms and 2 for the transition probabilities) is $\chi^2(22) = 594.10$ (*p*-value: 0.000).

TABLE 1 - *Instability Analysis*

TRANSITION MATRIX

	from low-variance regime	from high-variance regime
to low–variance regime	0.917	0.447
to high-variance regime	0.083	0.553

COEFFICIENTS: INTERCEPTS
$v(s1)$: Low-variance regime, $v(s2)$: High-variance regime

	Δoil	Δy	Δw	Δm	$\Delta \pi$
$v(s1)$	0.0227 (0.0146)	−0.0153 (0.0063)	0.0104 (0.0021)	0.0003 (0.0012)	− 0.0044 (0.0014)
$v(s2)$	0.0391 (0.0231)	−0.0063 (0.0072)	0.0141 (0.0028)	−0.0002 (0.0012)	− 0.0018 (0.0017)

COEFFICIENTS: VARIANCE-COVARIANCE MATRIX
Low–variance (High-variance) regime

	Δoil	Δy	Δw	Δm	$\Delta \pi$
Δoil	0.0013 (0.0242)	–	–	–	–
Δy	−0.00003 (0.0011)	0.0006 (0.0007)	–	–	–
Δw	7.5×10^{-7} (−0.0001)	-8.9×10^{-6} (0.0001)	0.00002 (0.0002)	–	–
Δm	0.00003 (−0.00002)	0.00002 (−0.00002)	4.5×10^{-7} (1.1×10^{-6})	0.00002 (0.00001)	–
$\Delta \pi$	7.3×10^{-6} (−0.00008)	-5.2×10^{-7} (0.00003)	1.4×10^{-6} (1.7×10^{-6})	1.4×10^{-7} (-6.5×10^{-6})	4.8×10^{-6} (0.00003)

SPECIFICATION TESTS
(p-values in brackets)

	Δoil	Δy	Δw	Δm	$\Delta \pi$
AR (1 – 6)	2.49 (0.7780)	0.73 (0.9808)	1.57 (0.9048)	6.35 (0.2739)	6.75 (0.2396)
Normality $\chi^2(2)$	6.52 (0.0383)	5.21 (0.0736)	5.06 (0.0797)	5.51 (0.0637)	4.03 (0.1333)

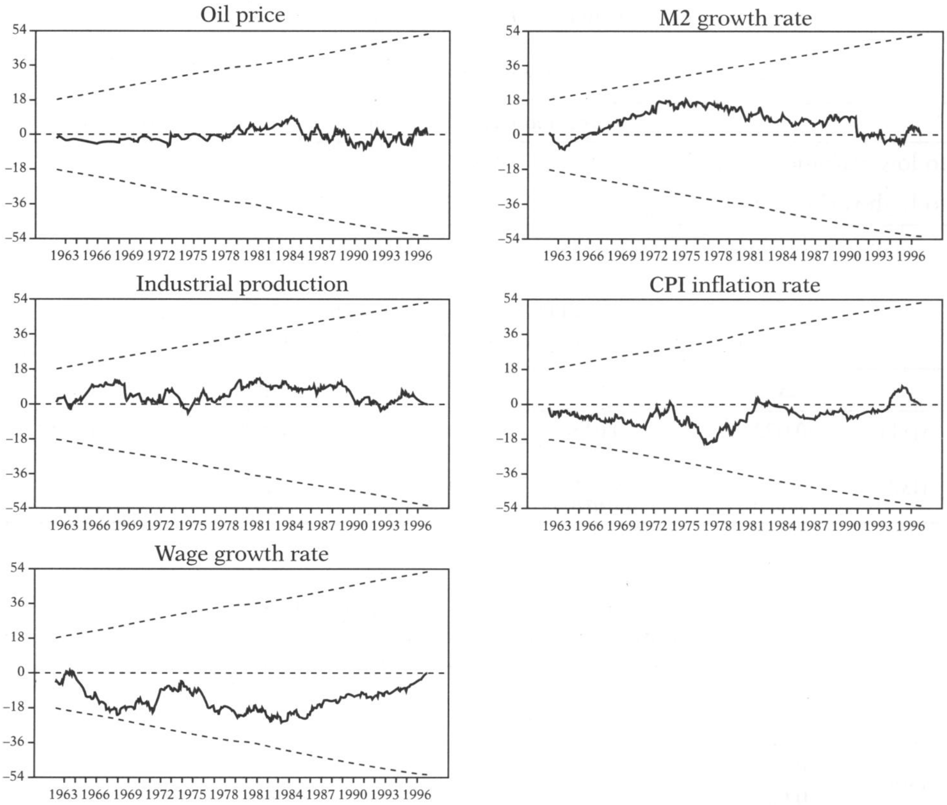

Fig. 3 - *CUSUM* Tests (with 10% Confidence Intervals).

A series of dummy variables corresponding to the dates of the high-variance regime is then constructed and included in the five-variable *VAR* system with six lags, on which the whole common trends analysis is carried out.

2.3 *Cointegration Analysis*

The first step in the common trends methodology is the determination of the long-run properties of the system. By including dummy variables corresponding to the high-variance regime, we restrict attention in the following analysis to the low-variance regime, spanning around 85% of the observations inthe sample.

The main results from the cointegration analysis are reported in Table 2. Formal tests support the existence of two valid cointegrating vectors. The coefficients of the cointegrating vectors are shown in the middle panel of the table, normalized on π and w for convenience. To gain some intuition on the

TABLE 2 - *Cointegration Analysis*

COINTEGRATION TESTS

Eigenvalue:	0.222	0.081	0.033	0.013	0.004
Hypothesis:	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$	$r \leq 4$
λ_{MAX}	106.8**	35.9**	14.3	5.7	1.6
95% crit. value	33.5	27.1	21.0	14.1	3.8
λ_{TRACE}	164.4**	57.6**	21.6	7.4	1.6
95% crit. value	68.5	47.2	29.7	15.4	3.8

r denotes the number of valid cointegrating vectors;
** denotes significance at the 1% level.

ESTIMATED COINTEGRATING VECTORS
(β' matrix; cointegrating vectors normalized on π and w respectively)

	oil	y	w	m	π
β'_1	−0.003	0.009	0.882	−1.330	1
β'_2	0.0021	−0.0013	1	0.048	−1.478

RESTRICTED COINTEGRATING VECTORS
(restricted β' matrix; cointegrating vectors normalized on π and w respectively; standard errors in parentheses)

	oil	y	w	m	π
β'_1	−0.0012 (0.0004)	0	0	−0.631 (0.088)	1
β'_2	0.0017 (0.0003)	0	1	0	−1.383 (0.078)

Likelihood-ratio tests: $\chi^2(2) = 1.21$ (p-value: 0.55)

economic nature of the underlying long-run relationships among the endogenous variables, various sets of restrictions on the cointegrating vectors have been imposed and tested. The last panel of the table shows the estimated coefficients obtained when the two vectors are restricted to capture long-run relations among π, m, *oil* and w, π, *oil* respectively. This set of restrictions is clearly not rejected by the likelihood ratio test (with a p-value of 0.55).

The estimates of the first cointegrating vector suggest that money growth and the oil price are the main long-run determinants of the inflation rate, whereas the second vector shows that nominal wage growth reacts more than proportionally to inflation and negatively to the oil price. The resulting error-correction terms ($\beta' \mathbf{x}_{t-1}$ in (2.3)) could then be interpreted as "excess" inflation and wage growth with respect to their long-run equilibrium levels[10]. The pattern of statistical significance of the estimated loading coefficients (the elements of matrix α in (2.3)) supports this interpretation. Excess inflation ($\beta_1' \mathbf{x}_{t-1}$) determines a negative response of inflation and a positive reaction of nominal money growth, suggesting a bi-directional causality between money growth and inflation, whereas excess wage growth ($\beta_2' \mathbf{x}_{t-1}$) induces a negative, error-correcting, adjustment of nominal wage growth only. Finally, it is interesting to notice that the inflation-money growth relationship enters the equation for Δy with a negative sign, suggesting a short-run inverse relationship between inflation and output growth.

The two restricted cointegrating vectors are incorporated in the common trends model specified and estimated in the following section[11].

2.4 *Common Trends Estimation*

In the common trends model in (2.11), the presence of two cointegrating relationships among the five endogenous variables in $\mathbf{x}_t$ implies the existence of three distinct sources of shocks having permanent effects on (at least some of) the elements of $\mathbf{x}_t$. The restrictions in (2.12) and (2.13) yield twelve equations which can be used to obtain the fifteen elements of the

[10] When the formal restrictions of long-run proportional relationships between inflation and money and wage growth are imposed (yielding the two restricted vectors $\pi = m - \beta_{11}\, oil$ and $w = \pi - \beta_{12}\, oil$), a clear rejection is obtained (the p-value for the associated test is 0.0001).

[11] If the same set of cointegrating restrictions is imposed on the *VAR* system with no dummies (used in the preliminary stage of the procedure described in the previous subsection), the following results are obtained:

$$\pi = \underset{(0.096)}{0.686}\, m + \underset{(0.0004)}{0.0017}\, oil$$

$$w = \underset{(0.099)}{1.266}\, \pi - \underset{(0.0003)}{0.0019}\, oil$$

with a likelihood-ratio test yielding $\chi^2(2) = 1.54$ (p-value: 0.46). The restrictions are not rejected and the coefficients are very close to those reported in Table 2, showing that the two estimated long-run relationships are strong features of the data. Moreover, this result suggests that the cointegration analysis is not substantially affected by a potentially mis-specified short-run structure, supporting the use of the Johansen procedure in the first step of the *MS-VECM* specification advocated by KROLZIG (1997) and adopted in the previous subsection.

long-run impact matrix $\mathbf{\Gamma}_g$. Three additional restrictions are then needed for (exact) identification of the common trends. To achieve identification, we make the following assumptions on the nature of the three permanent shocks in the system: we consider a *foreign real* shock (ψ_f), motivated by the huge oil price movements in the sample period, a *domestic real* shock (ψ_r), and a *nominal* disturbance (ψ_n). The specification adopted for the permanent part of the common trend representation in (10) is the following trivariate random walk:

$$\begin{pmatrix} \tau_f \\ \tau_r \\ \tau_n \end{pmatrix}_t = \begin{pmatrix} \mu_f \\ \mu_r \\ \mu_n \end{pmatrix} + \begin{pmatrix} \tau_f \\ \tau_r \\ \tau_n \end{pmatrix}_{t-1} + \begin{pmatrix} \psi_f \\ \psi_r \\ \psi_n \end{pmatrix}_t \tag{3.1}$$

where μ is a vector of constant drift terms, added to the model in estimation.

The three additional restrictions on the elements of $\mathbf{\Gamma}_g$ needed for identification are consistent with the above assumptions on the economic nature of the common trends. We assume that both the domestic real shocks and the nominal disturbances do not have long-run effects on the oil price, and that domestic output is not affected in the long-run by the nominal shock (a long-run neutrality assumption). Letting γ_{ij} denote the generic element of $\mathbf{\Gamma}_g$, the two assumptions above imply $\gamma_{12} = \gamma_{13} = 0$ and $\gamma_{23} = 0$ respectively. The common trends representation of the variables in levels is therefore the following:

$$\begin{pmatrix} oil \\ y \\ w \\ m \\ \pi \end{pmatrix}_t = \begin{pmatrix} oil \\ y \\ w \\ m \\ \pi \end{pmatrix}_0 + \begin{pmatrix} \gamma_{11} & 0 & 0 \\ \gamma_{21} & \gamma_{22} & 0 \\ \gamma_{31} & \gamma_{32} & \gamma_{33} \\ \gamma_{41} & \gamma_{42} & \gamma_{43} \\ \gamma_{51} & \gamma_{52} & \gamma_{53} \end{pmatrix} \begin{pmatrix} \tau_f \\ \tau_r \\ \tau_n \end{pmatrix}_t + \mathbf{\Gamma}^*(L) \begin{pmatrix} \psi_f \\ \psi_r \\ \psi_n \\ \nu_1 \\ \nu_2 \end{pmatrix}_t \tag{3.2}$$

where ν_1 and ν_2 are two purely transitory disturbances (uncorrelated with the permanent shocks) to which, given the main focus of our analysis, we do not attribute any structural economic interpretation.

Estimation of the common trends model is performed following the approach set out in the methodological section and the main results are shown in Table 3, where the estimated elements of the long-run impact matrix $\mathbf{\Gamma}_g$ and the long-run forecast error variance decomposition of the variables are reported (with asymptotic standard errors in parentheses)[12].

[12] Estimation has been carried out using the *CT* routine for *RATS* by A. WARNE and H. HANSEN.

TABLE 3 - *The Estimated Common Trends Model*

LONG-RUN EFFECTS OF PERMANENT SHOCKS (MATRIX Γ_g)
(asymptotic standard errors in parentheses; ** denote statistical significance at the 1% level)

Variable	Shock ψ_f	ψ_r	ψ_n
oil	0.0267**	0	0
	(0.0023)	(–)	(–)
y	–0.0043	0.0114**	0
	(0.0023)	(0.0019)	(–)
w	0.0003**	0.0003	0.0005**
	(0.00008)	(0.0001)	(0.00003)
m	0.0003**	0.0003	0.0006**
	(0.00009)	(0.0001)	(0.00004)
π	0.0002**	0.0002**	0.0004**
	(0.00005)	(0.00008)	(0.00003)

LONG-RUN (∞) FORECAST ERROR VARIANCE DECOMPOSITION

Variables	Shock ψ_f	ψ_r	ψ_n
oil	1	0	0
	(–)	(–)	(–)
y	0124	0.876	0
	(0.113)	(0.113)	(–)
w	0.148	0.191	0.661
	(0.083)	(0.121)	(0.093)
m	0.149	0.191	0.661
	(0.083)	(0.121)	(0.093)
π	0.195	0.180	0.625
	(0.090)	(0.117)	(0.087)

A number of features can be noticed from the estimated model. First of all, the foreign shock has permanent, positive and statistically significant effects not only on the oil price but also on inflation and on the nominal variables (money and wage growth rates), whereas the negative effect on output is not significant. According to the forecast error variance decomposition results, the foreign shock explains about 20% of inflation variability in the long run. The domestic real shock appears to have a similar permanent effect on the inflation rate; its contribution to the long-run inflation variance is also of similar magnitude (18%), though less precisely estimated. Finally, the nominal disturbance has long-run effects on π, m and w of a very close magnitude, suggesting that the behaviour of these variables is consistent with the assumed long-run neutrality of the nominal shock. This shock explains 62% of long-run inflation variability and a similar fraction of the variance of nominal wage and money growth (66%).

The main important conclusion to be drawn from the results in Table 3 is that the estimated core inflation from the common trends model, interpreted as the long-run inflation forecast due to all permanent shocks identified in the system, does not coincide simply with the output-neutral inflation, due exclusively to permanent nominal disturbances. A relevant share of long-run inflation movements are attributable to the foreign and to the real domestic shocks, which jointly explain about 38% of the long-run forecast error of the inflation rate.

The estimated core inflation series from the common trends model, computed as $\hat{\pi}_t^c = \pi_0 + \hat{\gamma}_{51}\hat{\tau}_{f,t} + \hat{\gamma}_{52}\hat{\tau}_{r,t} + \hat{\gamma}_{53}\hat{\tau}_{n,t}$, is shown in Fig. 4 together with the observed inflation rate and the transitory, "non-core" component simply computed as $\hat{\pi}_t^{nc} = \pi_t - \hat{\pi}_t^c$. All series are expressed as annual rates for the 1964(3)-1997(12) period.

As expected, core inflation is lower than observed inflation during the negative oil-shock episodes of the mid-1970s and early 1980s, whereas the non-core component closely follows the actual inflation rate. This pattern is reversed during the counter-shock of the mid-1980s. Core inflation is very close to the observed rate in the recent disinflation period, especially from 1989 to 1993, and does not follow the temporary increase in measured price growth occurred in 1995. However, the sharp decline in the observed inflation rate in 1996-1997 is not matched by a similar decrease of the core inflation series, which in fact shows an opposite behaviour, with a sharp rise from an annual rate of around 3% to 5% in the first half of 1997, due to positive realizations of the nominal disturbance. Only towards the end of the year core inflation seems to start converging again to the lower observed rate. According to our specification of the common trends model, therefore, the long-run forecast of the inflation rate during 1997 was well above the low measured inflation rates. This evidence could justify the prudent attitude of the Italian monetary authorities in cutting interest rates in the face of steadily low inflation rates.

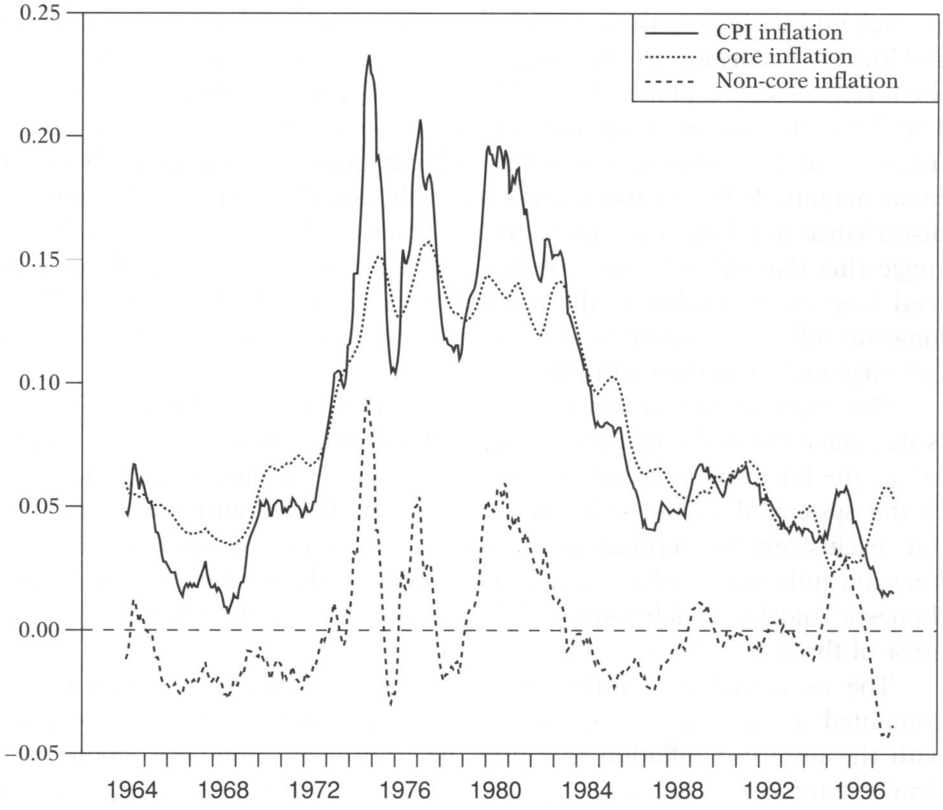

Fig. 4 - Annual Measured CPI Inflation and Estimated Core and Non-Core Inflation Rates from the Common-Trends Model.

To complete the picture, Fig. 5 displays the estimated core component (long-run forecast) of the nominal money and wage growth, obtained as $\hat{m}_t^c = m_0 + \hat{\gamma}_{41}\hat{\tau}_{f,t} + \hat{\gamma}_{42}\hat{\tau}_{r,t} + \hat{\gamma}_{43}\hat{\tau}_{n,t}$ and $\hat{w}_t^c = w_0 + \hat{\gamma}_{31}\hat{\tau}_{f,t} + \hat{\gamma}_{32}\hat{\tau}_{r,t} + \hat{\gamma}_{33}\hat{\tau}_{n,t}$ respectively. As for the recent increase in the money growth rate, from slightly negative values in mid-1996 up to an annual rate of around 10% in the first months of 1997, the estimated core money growth shows that the long-run forecast of m is around 5% at the end of 1997, suggesting that a substantial part of the increase in money growth is due to transitory disturbances. Finally, the analysis of the impulse response functions of the inflation rate to the three permanent shocks shows that convergence to the long-run response is entirely achieved within a period ranging from two to three years.

To summarize our results, we note that the estimated core inflation, interpreted as a long-run forecast, shows that purely nominal, output-neutral

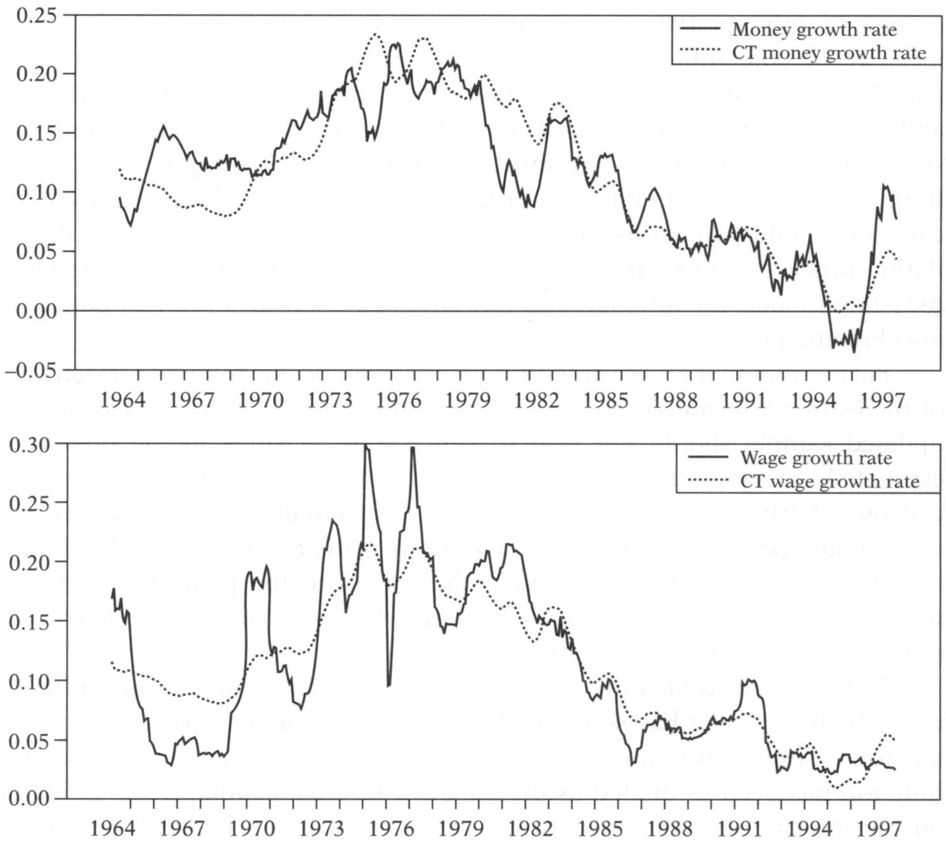

Fig. 5 - Permanent Component of Annual Money and Wage Growth Rates.

shocks, though important in explaining long-run inflation variability, are not the only determinant of trend inflation. Indeed, almost 40% of inflation variability is attributable to supply-side disturbances of both foreign and domestic nature.

5. ASSESSMENT OF THE ESTIMATED CORE INFLATION PROCESS

Core inflation is an unobserved process and its estimation is unavoidably subject to some degree of arbitrariness. Then, to give reliable information for policy use, a core inflation measure must display some desirable properties.

Bryan - Cecchetti (1994) define core inflation as "the long-run, or persistent, component of the measured price index, which is tied in some way to money growth" (p. 197), and point to a number of desirable features that a measure of core inflation should possess to match its theoretical counterpart. First, the estimated core inflation series should display lower variability and higher persistence than actual inflation. In principle, core inflation should be less sensitive to extreme observations, smoothing out peaks and troughs of the observed inflation process. Second, core inflation should forecast actual inflation, so as to be used to extrapolate trends in the actual inflation process or to be included as an additional explanatory variable in forecasting models for inflation and other related quantities (on this point, see also Freeman (1998)).

Furthermore, the core inflation process should be robust to extensions of the estimation sample. Re-estimation of the core inflation process over an updated sample should not lead to substantial revisions in the previously estimated core measure. Finally, a measure of core inflation that is grounded on a fairly general view of the main determinants of the inflation process, should perhaps be preferred to a purely statistical measure. The theoretical framework can provide useful information to the policymaker on the relative importance of the various economic forces in shaping the behaviour of the core component of inflation.

The common trends methodology yields a measure of the core inflation rate which naturally has some of the features mentioned above. In particular, forecasting power for the actual inflation rate is warranted, since core inflation has been estimated as the long-run forecast of inflation conditional on the information contained in the variables included in the *VAR* system. Moreover, the trend component of the series has been modelled as a multivariate random walk, therefore exhibiting a high degree of persistence. Finally, the relatively large information set embedded in our multivariate framework should lead to an effective decomposition of the observed inflation into a permanent and a transitory component, also allowing a deeper understanding of the structural determinants of the core inflation process.

However, the smoothing and robustness properties of the estimated core inflation require further investigation. As already shown in Fig. 4, the core component of inflation obtained from the common trends model displays less variability than the observed inflation rate, especially during the oil-shock episodes in the 1970s and 1980s. As a more formal assessment of the issue, we report in Table 4 some summary statistics on the distribution of the monthly changes in the observed inflation rate ($\Delta\pi$) and in the two (core and non-core) estimated components, such that $\Delta\pi = \Delta\pi^{c}_{CT} + \Delta\pi^{nc}_{CT}$ (the subscript *CT* indicates that the permanent and transitory components have been estimated with the common trends model). Fig. 6 displays the density functions of the changes in the observed inflation rate and of the estimated $\Delta\pi^{c}_{CT}$ and $\Delta\pi^{nc}_{CT}$ compo-

TABLE 4 - *Assessment of the Estimated Core Inflation*

SUMMARY STATISTICS FOR MONTHLY INFLATION CHANGES

	$\Delta\pi$	$\Delta\pi^{c}_{CT}$	$\Delta\pi^{nc}_{CT}$	$\Delta\pi^{c}_{QV}$
Mean (%)	−0.001	−0.001	0.0002	0.0001
St. dev. (%)	0.350	0.054	0.333	0.347
Skewness	−0.296	−0.019	−0.393	−0.211
Excess kurtosis	3.500	0.347	3.805	3.467
Normality $\chi^2(2)$	104.8	2.89	111.4	107.8
(p-value)	(0.00)	(0.24)	(0.00)	(0.00)

CORRELATION COEFFICIENTS

	$\Delta\pi$	$\Delta\pi^{c}_{CT}$	$\Delta\pi^{nc}_{CT}$
$\Delta\pi$	-	-	-
$\Delta\pi^{c}_{CT}$	0.381	-	-
$\Delta\pi^{nc}_{CT}$	0.989	0.241	-
$\Delta\pi^{c}_{QV}$	0.985	0.404	0.969

The sample period is 1964(4)-1997(12). $\Delta\pi^{c}_{CT}$ and $\Delta\pi^{nc}_{CT}$ denote the monthly change in the core and non-core inflation rates estimated with the common tremds model. $\Delta\pi^{c}_{QV}$ denotes the monthly change in the core inflation rate obtained using the Quah-Vahey procedure discussed in Section 5. Kurtosis statistics are expressed as deviations from 3, which characterizes a normal distribution.

nents. As expected, the standard deviation of core inflation changes (0.05 percentage points) is much lower than that of observed inflation changes (0.35), very close to the variability of the non-core component (0.33). Moreover, the shape of the distribution of $\Delta\pi^{c}_{CT}$ does not show any relevant deviation from normality (a formal test based on skewness and kurtosis measures does not reject normality with a *p*-value of 0.24), showing that the estimated core component is not heavily affected by extreme realizations.

For completeness, we compare the properties of the core inflation derived from the common trends model with those of the core series obtained from a simpler approach (originally applied to the UK by Quah - Vahey (1995)) based on a two-variable *VAR* system, including only the industrial production index and the inflation rate. Within this simpler (and less informative) framework, the identification of the permanent component of the inflation process relies on the assumption that shocks to the core inflation rate have no long-run effect on output, whereas disturbances to the non-core inflation component have only transitory effects on the inflation rate but permanently affect output. These joint restrictions are imposed on the long-run multiplier matrix of the bivariate *VAR*, following the procedure introduced by Blanchard - Quah (1989)[13].

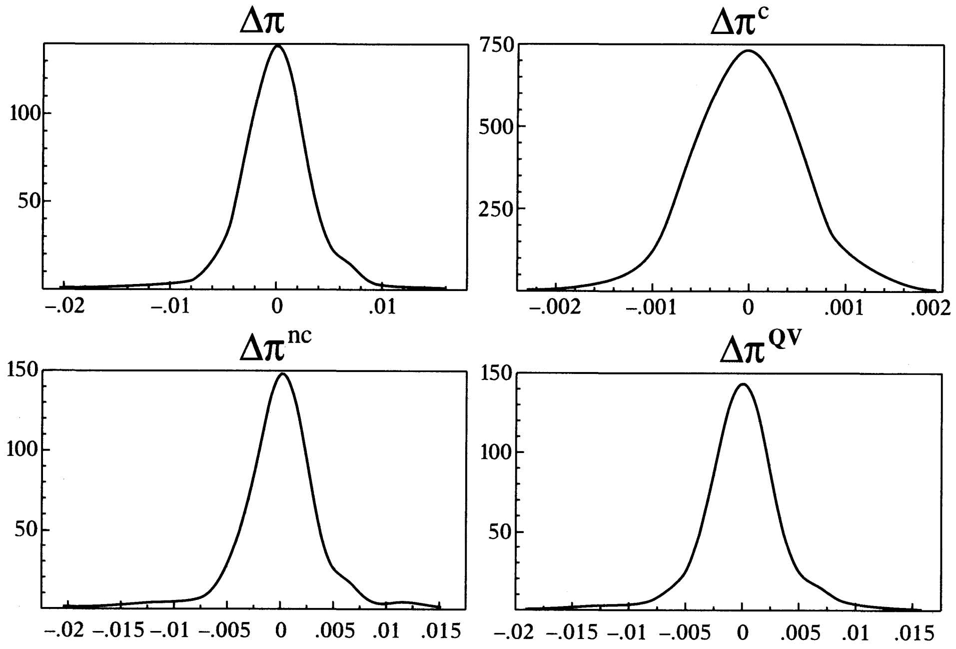

Fig. 6 - Density Functions of $\Delta\pi$, $\Delta\pi^c$, $\Delta\pi^{nc}$ and $\Delta\pi^{QV}$

When this procedure is implemented on our data, further support is obtained for the more informative common trends approach illustrated above, explicitly including money growth, wage growth and a supply-side variable as inflation determinants[14]. Firstly, formal tests strongly reject the overall

[13] An Appendix (available upon request) provides a detailed account of the methodology implemented.

[14] To make the results from the Quah-Vahey approach comparable with the common trends model, a series of dummy variables has been added to the bivariate (Δy, $\Delta\pi$) system in order to achieve stability over the sample period. The dummies are: 1962(12)-1963(1), 1970(3)-1970(4), 1973(11)-1974(3), 1974(6)-1974(9), 1974(12)-1975(4), 1976(1)-1976(4), 1979(12)-1980(1), 1994(4). No term in levels is added, since there is no evidence of cointegration between output and inflation, which allows the implementation of the Quah-Vahey methodology.

long-run identification scheme, showing that this methodology is not successful in separating the underlying (core) inflation process from the noise generated by disturbances with no long-run effect on the measured inflation rate[15]. Secondly, Table 4 shows that monthly changes in the core inflation obtained using the Quah-Vahey procedure ($\Delta\pi^c_{QV}$) almost replicate the features of observed inflation change, displaying a very close standard deviation and an almost perfect correlation (0.98, whereas the correlation between $\Delta\pi^c_{CT}$ and $\Delta\pi$ is lower than 0.4). The huge deviation from normality detected by the test reported in the table and displayed in the lower panel of Fig. 6 confirm that the Quah-Vahey core inflation does not correspond to the theoretical concept of underlying inflation that it is meant to measure.

As a last step in the assessment of our common trends core inflation measure, we investigated the sensitivity of the estimated core inflation process to the updating of the estimation sample. To this aim, the same common trends model has been fitted to a sample ending in 1994(12) and the resulting core inflation estimate is then compared, over the same time span 1964(3)-1994(12), with that previously obtained from a sample including the years 1995-1997. The wide fluctuations in the observed inflation rate over this three-year period, due to the exchange rate crisis and the surge in commodity prices in 1995 and to the subsequent sharp disinflation, should guarantee a harsh robustness test.

The discrepancies between the core inflation processes estimated using the longer and the shorter sample have then been evaluated by means of the root mean square error (*RMSE*) measure and Theil (1961) inequality coefficient criteria. These statistics are usually employed the evaluate the forecasting performance of a model, but in the present context we use them to assess the sensitivity of our results to a (limited but meaningful) extension of the estimation sample. Both the *RMSE* and the inequality coefficient (U) criteria give indications about the average deviation between the two core inflation series. A measure close to zero for these two criteria guarantees that the two series are close to each other. In addition, the Theil U coefficient can be used to gain some insights into the three sources of any possible deviation of the two series from each other, namely bias, variance, and covariance. The bias component (U^m) is a measure of the systematic error, whereas the variance component (U^v) gives an indication about whether the two series share a common variability; both these components ideally should be zero.

[15] The appropriate likelihood test of the over-identifying restrictions yields a $\chi^2(1)$ value of 37.47, with an associated p-value smaller than 0.001.

Finally, the covariance component (U^c) may be interpreted as a measure of the unsystematic error. Ideally, the covariance component should account fully for the deviations between the two series, so that $U^c = 1$. When computed for the core inflation series estimated over the 1964-1994 and 1964-1997 sample periods, the above statistics are:

RMSE	U	U^m	U^v	U^c
0.0020	0.0126	0.1232	0.2318	0.6450

According to all measures (and to the graphical evidence in Fig. 7, where the two core inflation series are plotted along with the observed inflation rate), the two series are remarkably close to each other, supporting the conclusion that the estimated core inflation process is robust to this extension of the sample period.

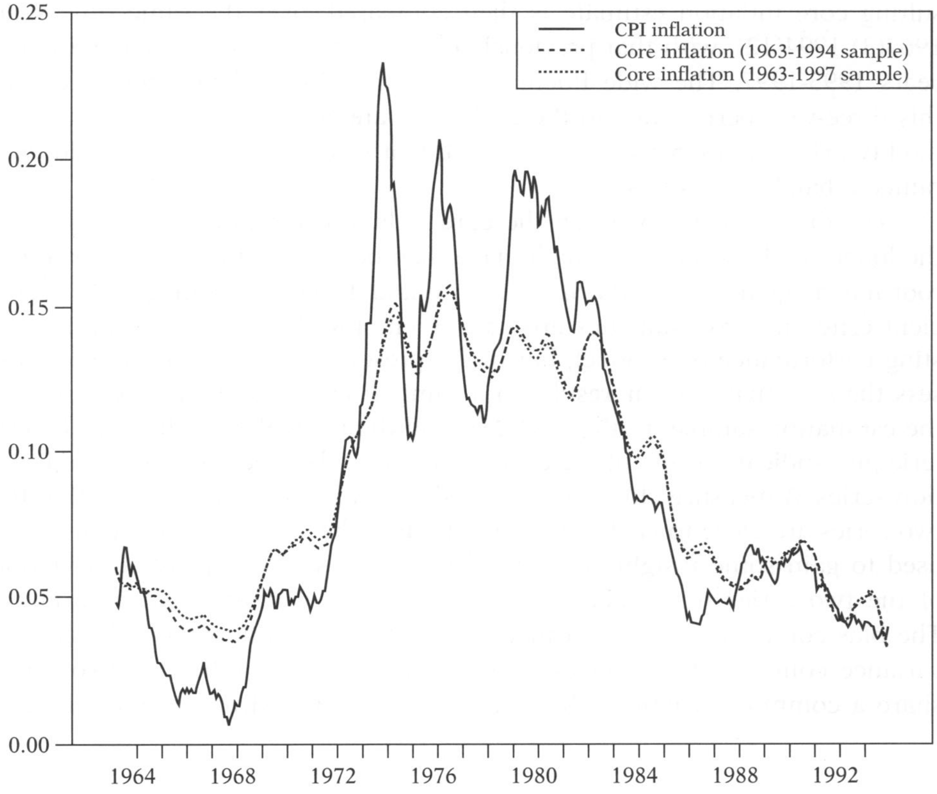

Fig. 7 - Observed and Core Annual Inflation Rates Estimated Over the 1963-1994 and 1963-1997 Periods.

Conclusions

A common trends model has been used to estimate the underlying, core inflation behaviour for Italy over a long period lasting more than thirty years. In this framework core inflation is interpreted as the long-run forecast of inflation. Estimating a common trends model allows for the decomposition of observed inflation into an underlying trend (core) and a transitory component. In our specification, including output, inflation and three main determinants of inflationary pressures (money and wage growth, and the oil price), core inflation behaviour is attributable not only to permanent nominal disturbances, but also to foreign (supply) and real domestic shocks.

This measure of the long-run inflation trend displays several desirable properties (lower variability than observed inflation, forecasting power, robustness to the estimation sample, economic interpretability) that make it attractive from a monetary policy perspective.

Of course, a core inflation rate estimated from a common trends model does depend on the specification of the system in terms of variables included, sample period, dynamic specification, and other modelling choices. However, the core inflation series obtained from the small-scale macroeconomic model used in this paper, featuring long-run links between inflation and its main macroeconomic determinants, seems an useful benchmark against which to evaluate the properties of other purely statistical and atheoretical measures of core inflation.

REFERENCES

Bank of Italy (1999), *Annual Report for 1998*, Rome, Bank of Italy.

Bernanke B.S. - Laubach T. - Mishkin F.S. - Posen A.S. (1999), *Inflation Targeting. Lessons from the International Experience*, Princeton, Princeton University Press.

Blanchard O. - Quah D. (1989), «The dynamic effects of aggregate supply and demand disturbances», *American Economic Review*, vol. 79, pp. 655-673.

Bryan M.F. - Cecchetti S.G. (1993), «The consumer price index as a measure of inflation», *Economic Review*, Federal Reserve Bank of Cleveland, 4, pp. 291-320.

—— - —— (1994), «Measuring core inflation», in *Monetary Policy*, ed. by N.G. Mankiw, Chicago, University of Chicago Press, pp. 195-215.

Engle R.F. - Granger C.W.J. (1987), «Cointegration and error correction: representation, estimation and testing», *Econometrica*, vol. 55, pp. 251-276.

Freeman D.G. (1998), «Do core inflation measures help forecast inflation?», *Economics Letters*, vol. 58, pp. 143-147.

Gallo G.M. - Kempf H. (1996), «Common cycles and economic fluctuations», European University Institute, mimeo.

GAVOSTO A. - SABBATINI R. (1997), «Misure di 'core inflation»', *Politica Economica, vol. 13, pp. 67-85.*

JOHANSEN S. (1988), «Statistical analysis of cointegrating vectors», *Journal of Economic Dynamics and Control*, vol.12, pp. 231-254.

—— - —— (1991), «Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models», *Econometrica*, vol.59, pp. 1551-1580.

KING R.J. - PLOSSER C. - STOCK J. - WATSON M. (1991), «Stochastic trends and economic fluctuations», *American Economic Review*, vol. 81, pp. 280-315.

KROLZIG M.-H. (1997), «Statistical analysis of cointegrated VAR processes with Markovian regime shifts», Oxford University, mimeo.

MELLANDER E. - VREDIN A. - WARNE A. (1992), «Stochastic trends and economic fluctuations in a small open economy», *Journal of Applied Econometrics*, vol. 7, pp. 369-394.

QUAH D. - VAHEY S.P. (1995), «Measuring core inflation», *Economic Journal*, vol. 105, pp. 1130-1144.

STOCK J.H. - WATSON M.W. (1988), «Testing for common trends», *Journal of the American Statistical Association*, vol. 83, pp. 1097-1107.

SVENSSON L.E.O. (1997), «Inflation forecast targeting: implementing and monitoring inflation targets», *European Economic Review*, vol. 41, pp. 1111-1146.

THEIL H. (1961), *Economic Forecasts and Policy*, Amsterdam, North Holland.

WARNE A. (1993), «A common trends model: identification, estimation and inference», Seminar Paper No. 555, IIES, Stockholm University.

WYNNE M.A. (1999), «Core inflation: a review of some conceptual issues», European Central Bank Working Paper n. 5, Frankfurt.