4 Inflation modelling in the euro area

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1 Introduction

Controlling inflation, at least in the long run, is widely regarded as the primary, and sometimes the only, goal of monetary policy. To this aim, in many countries central banks have explicitly adopted inflation-targeting strategies, setting precise quantitative targets for the monetary authorities' actions. Though not an inflation targeter, the European Central Bank (ECB) adopted a monetary policy strategy aimed at maintaining an annual inflation rate below 2% over a medium-term horizon (ECB 1999). This strategy is based on an announced reference value for M3 money growth and on the outlook of price developments in the euro area. The analysis of the behaviour of monetary aggregates and their components relies on a number of tools recently summarised in ECB (2001). The aim of this chapter is to provide an empirical investigation of the interrelationships among money, prices, interest rates and output in the euro area with a particular focus on the behaviour of the inflation rate over a long-run horizon. In fact, one of the main open issues in inflation analysis stems from the fact that short-run fluctuations of the observed inflation rate may be due to only temporary disturbances to which monetary policy should not respond. How to construct a reliable empirical measure of the underlying, long-run trend of inflation – '*core*' inflation – has therefore become a crucial issue in monetary policy design.

Core inflation series have been constructed following different methodologies (see Wynne 1999 for a thorough overview and assessment of different measures). Some measures are obtained from the cross-sectional distribution of individual price items, either by excluding from the price index some categories of goods (such as energy and food items) which are believed to be high-variance components, or by computing more efficient, 'limited influence'

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estimators of the central tendency of the distribution, such as the (weighted) median popularised by Bryan and Cecchetti (1994) and Cecchetti (1997) for the USA. Other measures are derived from univariate statistical techniques, such as simple moving averages computed over a variable time span (from 3–6 up to 36 months) or more sophisticated methodologies (i.e. unobserved component models, or the one-sided low-pass filter proposed by Cogley 2002). Finally, Quah and Vahey (1995) applied to the UK a bivariate structural vector autoregressive (SVAR) approach to core inflation estimation based on long-run output neutrality of permanent shocks to the inflation rate.

We propose a different, explicitly *forward-looking*, measure of core inflation, based on (appropriately estimated and tested) long-run relations among major macroeconomic variables. This measure may provide useful information in the light of the 'two-pillar' monetary policy strategy of the ECB, which considers: (i) the deviations of M3 growth from a reference value (a money growth indicator), and (ii) a broadly based assessment of the outlook for future price developments in the euro area as a whole (ECB 1999, 2000). This framework is motivated by the (alleged) close long-run relationship between money growth and inflation. Recent results have provided some evidence of stable long-run relationships among money, output, interest rates and inflation over the last two decades for the EMU countries (Brand and Cassola 2000, Gerlach and Svensson 2001, Golinelli and Pastorello 2002). We use such information to construct a forward-looking measure of core inflation consistent with the long-run features of the euro area macroeconomy.

To this aim, we consider a multivariate framework, capturing the dynamic interactions among the inflation rate, real money balances, short- and long-term interest rates and output, extending the analysis in Bagliano, Golinelli and Morana (2002). A stylised macroeconomic model is set up in section 2 to provide a theoretical rationale for the potential long-run relationships among those variables. The existence of valid cointegrating relations is then explored using euro area data for the 1979–2001 period. The problem of structural breaks in the behaviour of the long-term real interest rate is addressed by means of a Markov-switching model for the real rate. In order to decompose observed inflation into a non-stationary (stochastic) trend component, capturing the effect of permanent shocks only, and a stationary transitory element, we adopt a common trends approach. The permanent, 'core' inflation component bears the interpretation of the long-run inflation forecast conditional on an information set including several important macroeconomic variables. The main advantage of this measure of core inflation lies in its forward-looking nature, capturing the long-term element of the inflation process (of particular interest from the monetary policy perspective) consistent with the long-run properties of the macroeconomic system. Section 3 describes the common trends methodology and presents empirical results. Several properties of the estimated core inflation process are then assessed, namely its relative volatility with respect to observed inflation and its ability to forecast future headline inflation rates. Further features of the permanent-transitory decomposition of the inflation rate are analysed in section 4, where the nature of the non-core inflation fluctuations and the convergence of the observed rate to the core inflation rate are discussed. Finally, our main message is summarised in the concluding section 5: the ECB should take into proper account a forward-looking measure of the core inflation rate consistent with its whole monetary policy framework, based on strong and stable long-run relationships between inflation and other major macroeconomic variables.

2 Long-run analysis of a small-scale macro system

To organise thinking about the long-run relationships among inflation, output, money and interest rates we start with a general equation for inflation determination, nesting a traditional backward-looking Phillips curve, whereby inflation is mainly determined by the 'output gap', and a P^* model (see Hallman, Porter and Small 1991), which assumes that inflation dynamics is governed by the 'price gap'. The latter model has recently received strong support for the euro area from Gerlach and Svensson (2001). Ignoring additional dynamic terms and exogenous variables, the equation for the inflation rate is of the form:

$$\pi_t = \pi^e_{t,t-1} + \alpha_y(y_{t-1} - y^*_{t-1}) + \alpha_m(p_{t-1} - p^*_{t-1}) + \varepsilon^\pi_t, \tag{4.1}$$

where π_t is the annualised inflation rate in quarter t ($\pi_t \equiv 4(p_t - p_{t-1})$) and $\pi^e_{t,t-1}$ is the expected inflation rate as of quarter $t-1$, y—y^* measures the output gap, with y^* denoting potential output, and $p - p^*$ is the 'price gap', the key determinant of inflation in the P^* model, to be more precisely defined below. Finally, ε^π represents a random shock to inflation. The empirical specification of equation (4.1) requires us to model inflationary expectations. As in other studies which use a backward-looking Phillips curve (e.g. Taylor 1999, Rudebusch and Svensson 1999, Staiger, Stock and Watson 2001), the expected inflation rate $\pi^e_{t,t-1}$ is set equal to π_{t-1}.[1] Therefore we get:

$$\Delta\pi_t = \alpha_y(y_{t-1} - y^*_{t-1}) + \alpha_m(p_{t-1} - p^*_{t-1}) + \varepsilon^\pi_t. \tag{4.2}$$

Moreover, we assume:

$$y^*_t = \beta^y_0 + y^*_{t-1} + \varepsilon^y_t \tag{4.3}$$

$$m_t - p_t = \beta^m_0 + \beta^m_1\, y_t + \beta^m_2\,(l_t - s_t) + \varepsilon^m_t \tag{4.4}$$

$$p^*_t = m_t - \left[\beta^m_0 + \beta^m_1\, y^*_t + \beta^m_2(s^*_t - l^*_t)\right] \tag{4.5}$$

[1] Gerlach and Svensson (2001) adopt a different specification, setting $\pi^e_{t,t-1}$ as a weighted average of π_{t-1} and of the central bank's inflation objective.

$$l_t = \beta_0^f + \pi_{t+1,t}^e + \varepsilon_t^f \tag{4.6}$$

$$l_t = \beta_0^s + s_t + \varepsilon_t^s. \tag{4.7}$$

In (4.3) potential output follows a random walk. Real money demand is specified by (4.4), where the long–short interest rate differential ($l - s$) proxies the opportunity cost of money holdings. Equation (4.5) defines p_t^* as the price level consistent with the current money stock, potential output and long-run equilibrium values for the short and long interest rates (s^* and l^*), according to the P^* model. Finally, (4.6) and (4.7) capture a Fisher parity and a term structure relation respectively. All structural parameters (βs) are positive and the εs are random shocks. In a long-run equilibrium, the following relations hold:

$$\begin{aligned}
y &= y^* \\
\pi &= \pi^e \\
l^* &= \beta_0^f + \pi \\
l^* &= \beta_0^s + s^* \\
m^* - p^* &= \beta_0 + \beta_1^m\, y^*,
\end{aligned}$$

where m^* denotes long-run equilibrium nominal money balances and $\beta_0 \equiv \beta_0^m + \beta_2^m\, \beta_0^s$.

In the above framework, the inflation rate and output are non-stationary, $I(1)$, and the output gap is stationary, $I(0)$. Moreover, the long-term interest rate is $I(1)$ and cointegrated with the inflation rate, so that $l - \pi$ is $I(0)$, and the short-term rate is $I(1)$ and cointegrated with the long rate, so as to make $l - s$ stationary. From (4.4) real money balances are $I(1)$ and cointegrated with output; if the cointegration parameter $\beta_1^m \neq 1$, also money velocity is non-stationary. Then, the first step of our empirical analysis looks at the integration and cointegration properties of the series, to check their consistency with the above macroeconomic framework.

In order to proceed with the empirical analysis, we need euro-area variables over a time span pre-dating the launch of the euro at the beginning of 1999. For the pre-euro period (up to 1998Q4) aggregate variables for the euro area were constructed by aggregating the historical data of the twelve current member countries. This approach is based on the assumption that the artificial euro-area data before monetary union are appropriate for analysing and forecasting the area-wide behaviour under EMU.[2]

[2] Despite this caveat, the aggregation route was followed by several other recent studies: Gerlach and Svensson (2001) and Galí, Gertler and López-Salido (2001) recently used area-aggregated data to study the EMU inflation rate, and Golinelli and Pastorello (2003) find some results in favour of the statistical poolability of single-country money demand functions. The latter results are partly supported by Dedola, Gaiotti and Silipo (2001), who find that the area-wide money

In the present analysis, we use quarterly variables at an area-wide level over the 1978Q4−2001Q3 period. We measure (the log of) real money balances $(m - p)$ by the (log of the) index of nominal M3 (published by the ECB) deflated by the (log of the) Harmonised Index of Consumer Prices (HICP) used by the ECB; output (y) is measured by (the log of) real GDP, the nominal short and long-term interest rates (s and l) are the T-bill and the government bond rates, the inflation rate (π) is the annualised quarterly rate of change of the HICP, and the output gap ($ygap \equiv y - y^*$) is measured by the rate of capacity utilisation in the manufacturing sector measured by the OECD.[3]

The results of unit-root Dickey-Fuller ADF tests reported in table 4.1 are clear-cut: with the only exception of *ygap*, which is stationary, all the variables of interest are first order integrated. Moreover, the lower part of the table reports ADF test statistics for a number of additional variables: if the (null) unit-root hypothesis is rejected, then the corresponding $I(1)$ series are cointegrated with a $(1, -1)$ cointegrating vector. The results show that money velocity is $I(1)$ even when a linear trend is allowed in the specification of the test, the term interest rate differential is stationary (short and long-term rates are cointegrated), whereas the short- and long-term real interest rates are not stationary. As a whole, the evidence is consistent with the features of the above theoretical framework, except for the behaviour of the real interest rate series.

The missing Fisher parity relation deserves more careful scrutiny. To this aim, the lower panel of figure 4.1 plots the long-term real (ex-post) interest rate and the term interest rate differential for the euro area over the whole 1978Q4−2001Q3 period. While the interest rate differential fluctuates quite persistently around a constant mean, the real long-term interest rate shows a much lower mean for the sub-periods 1978–81 and 1997–2001, possibly suggesting that the non-stationarity detected by the ADF test is spurious, and due to a neglected structural change in the constant term of the Fisher parity relation.[4] For example, the introduction of the single monetary policy explicitly aimed at a price stability objective may have reduced inflation uncertainty and therefore the inflation risk premium embodied in the level of the

demand equation is not significantly affected by aggregation bias. Brand and Cassola (2000) and Coenen and Vega (2001) also study money demand only at an area-wide level. On the other side, Marcellino, Stock and Watson (2003), and Espasa, Albacete and Senra (2002) provide evidence against the use of aggregate models and prefer to forecast a number of euro-area variables at country level. Against this view, Bodo, Golinelli and Parigi (2000) show that the area-wide model is better than single country models in forecasting industrial production. Finally, a completely different approach is followed by Rudebusch and Svensson (2002), who use a model estimated on US data to discuss euro-area policy issues.

3 The data used in the empirical analysis are updated from Golinelli and Pastorello (2002). The data set is available for downloading at http://www.spbo.unibo.it/pais/golinelli/macro.htm, where further details on the sources are also provided.

4 Moreover, the Hansen (1992) instability test confirms the presence of instability in the mean real interest rate at the 5% significance level ($L_c = 1.40$).

Table 4.1. *Unit root ADF tests, 1978Q4−2001Q3*

Variable	ADF	k	Model
$m - p$	−3.09	2	c,t
$\Delta(m-p)$	−3.73**	1	c
y	−1.56	0	c,t
Δy	−5.65**	0	c
s	−1.29	1	c
Δs	−5.44**	0	c
l	−1.05	1	c
Δl	−5.15**	0	c
π	−1.36	1	c
$\Delta\pi$	−9.49**	1	c
$ygap$	−3.63**	2	c
$y-(m-p)$	−2.60	1	c,t
$s - \pi$	−1.88	2	c
$l - \pi$	−2.39	1	c
$l - s$	−3.40*	1	c

Notes: * and ** denote rejection of the null hypothesis of unit root at the 5% and 1% level, respectively. MacKinnon critical values are: −2.89 (5%) and −3.50 (1%) for models with constant only (c); −3.46 and −4.06 for models with constant and trend (c, t). k denotes the number of lags in the test, selected following the general-to-specific procedure advocated by Ng and Perron (1995) with $k_{\max} = 5$.

long-term interest rate over the last part of the sample. Then, instead of equation (4.6), a more appropriate specification of the Fisher relation for the euro area could be the following, allowing for changes in the mean real interest rate:

$$l_t = \beta_0^f(r_t) + \pi_t + \varepsilon_t^f, \tag{4.8}$$

where r_t is a random variable indexing the risk premium regime.

Structural change in the real interest rate has been investigated by means of a Markov-switching model (Hamilton, 1989), allowing us to detect potential break points endogenously, with no a priori assumption concerning their number and timing. Table 4.2 summarises the main features of the estimated Markov-switching model. According to the LR and specification tests, a two-regime model for the intercept in (4.8), with a first-order autoregressive term, can be

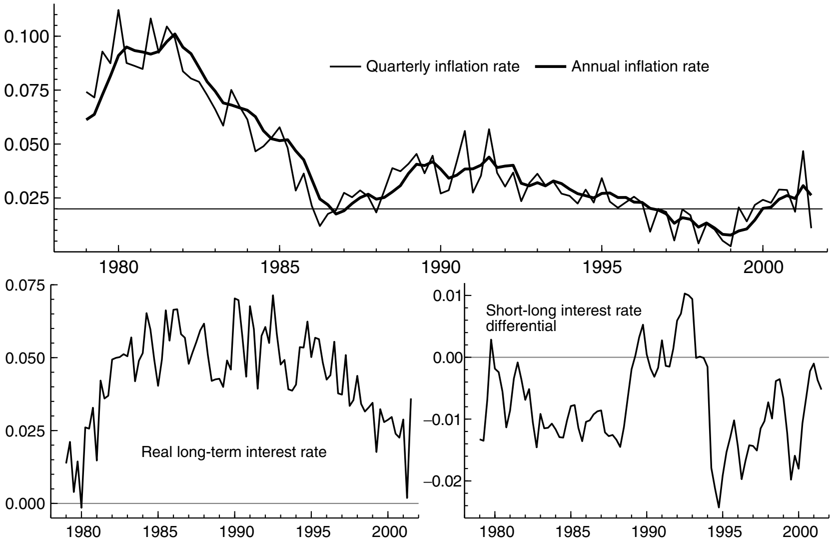

Figure 4.1 Quarterly and annual euro-area inflation rates, real long-term interest rate and short-long interest rate differential 1978Q4–2001Q3
Note: The quarterly inflation rate π is annualized; the annual inflation rate is computed $\sum_{i=0}^{3} \pi_{t-i}$; as the real long-term interest rate is obtained as $l - \pi$, the interest rate differential is computed as $s - l$.

selected, suggesting that the persistence in the real interest rate is not fully explained by the break process only.[5] As shown in the table, the estimated mean ex-post real interest rate is 2.6% in the 'low' regime and 5.2% in the 'high' regime. The estimated mean real interest rate is plotted in the upper panel of figure 4.2, together with the observed rate. In the lower panel of the figure, the estimated smoothed probabilities of the two regimes are shown: the 'low' real interest rate regime ends in 1981Q3 and starts again in 1997Q3, suggesting that the fall in the risk premium pre-dated the introduction of the common monetary policy in 1999,[6] whereas the 'high' real rate regime spans the 1981Q4–1997Q2 period. This finding points to an important contribution of

[5] The p-value of the LR test for the null of a single regime model against the two-regime model (computed as in Davies, 1987 to account for the non-standard asymptotic distribution of the test), is 0.002. The p-value of the test for two against three regimes is 1. Similar results are obtained by using the Perron (1997) DF test with endogenous break point: over the period 1981Q4 to 2001Q3 the long-term real interest rate is stationary with a break in 1997Q2 (the test statistic is −6.5 against the 1% critical value of −5.77).

[6] On the other hand, if the reference date is the Maastricht Treaty (February 1992), our findings are consistent with a lagged adjustment of the risk premium. The reduction may have taken place once the macroeconomic convergence in the euro area and the compatibility with the Maastricht parameters were unambiguous.

Table 4.2. *Regime switching analysis of the long-term real interest rate.*

	Regime 1	Regime 2
Regime 1	0.952	0.016
Regime 2	0.048	0.984
Mean	2.58 (0.21)	5.19 (0.13)
Duration (quarters)	21	61
Number of observations	29	63

Notes: The first four rows of the table report the transition matrix ($p_{ij} = \Pr\{r(t) = i\ /\ r(t-1) = j\}$). Mean denotes the estimated ex-post real interest rate in the two regimes. Duration denotes the average duration of each regime in quarters. The number of observations in each regime is reported in the last row.

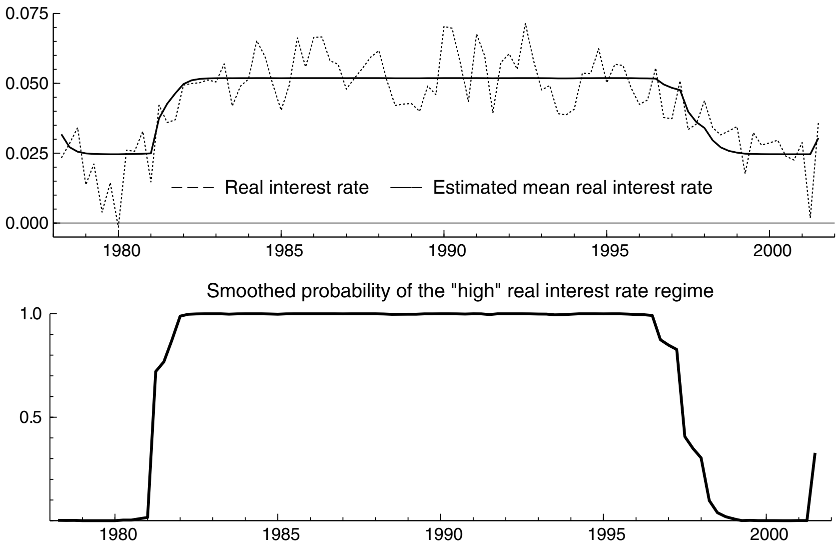

Figure 4.2 Markov-switching model of real interest rate

monetary unification to economic growth, through a reduced cost of investment financing.

The existence of two different regimes in the real interest rate behaviour has relevant consequences for the long-run empirical modelling of our set of six variables of interest ($m - p$, y, s, l, π and $ygap$). In fact, when tests for the cointegration rank and the forecasting ability are performed on a VAR(3) system over the 1981Q4−1997Q2 period only (identified above by the Markov-switching model as the 'high' real interest rate regime), the Johansen (1995) trace statistics support the existence of four cointegrating relationships at the 10% significance level. However, the one-step (ex-post) parameter constancy forecast test over the period 1998–2001 reveals strong evidence of a significant shift and, accordingly, the cointegration test over the full sample detects fewer than four cointegrating relationships. In short, the extension of the sample period leads to forecast failure and missing cointegration owing to parameter instability.

In order to capture the structural change in the long-run Fisher relation detected above, we include in the basic VAR system a step dummy variable (*RP*) taking the value of 1 during the 'high' real rate regime (1981Q4−1997Q2), and 0 in the 'low' rate regime (1978Q4−1981Q3 and 1997Q3−2001Q3). Prior to presenting the results, the next subsection shows how the standard methodology is extended to include a dummy variable in the cointegrating space.

2.1 Methodology

The standard vector error-correction mechanism (VECM) representation of the model, controlling for a linear trend in the level of the variables, can be written as

$$\mathbf{\Pi}^*(L)\,\Delta\mathbf{x}_t = \boldsymbol{\nu} + \mathbf{\Pi}(1)\,\mathbf{x}_{t-1} + \varepsilon_t, \tag{4.9}$$

where $\mathbf{x}_t$ is the vector of $n\,I(1)$ cointegrated variables of interest, $\boldsymbol{\nu}$ is the vector of intercept terms, $\varepsilon_t \sim NID\,(0, \Sigma)$; $\mathbf{\Pi}(L) = \mathbf{I}_n - \sum_{i=1}^{p} \mathbf{\Pi}_i L^i$, $\mathbf{\Pi}^*(L) = \mathbf{I}_n - \sum_{i=1}^{p-1} \mathbf{\Pi}_i^* L^i$ and $\mathbf{\Pi}_i^* = -\sum_{j=i+1}^{p} \mathbf{\Pi}_j$ $(i = 1, \ldots, p-1)$. If there are $0 < k < n$ cointegration relationships among the variables, $\mathbf{\Pi}(1)$ is of reduced rank k and can be expressed as the product of two $(n \times k)$ matrices: $\mathbf{\Pi}(1) = \boldsymbol{\alpha}\boldsymbol{\beta}'$, where $\boldsymbol{\beta}$ contains the cointegrating vectors, such that $\boldsymbol{\beta}'\mathbf{x}_t$ are stationary linear combinations of the $I(1)$ variables, and $\boldsymbol{\alpha}$ is the matrix of factor loadings. When one of the cointegrating vectors (i.e. the kth vector) contains a switching intercept modelled by dummy variables, it is possible to rewrite the $\boldsymbol{\beta}$ matrix as

$$\underset{(n+q)\times k}{\bar{\boldsymbol{\beta}}} = \begin{pmatrix} \multicolumn{2}{c}{\underset{n\times k}{\bar{\boldsymbol{\beta}}}} \\ \underset{q\times(k-1)}{\bar{\mathbf{0}}} & \underset{q\times 1}{\bar{\boldsymbol{\beta}}^*} \end{pmatrix},$$

where $\boldsymbol{\beta}^*$ is the $q \times 1$ subvector containing the parameters of the q deterministic variables in the kth cointegrating vector. If there are q regimes, $q-1$ regimes may be normalised relative to the qth regime; this amounts to measuring the switches relative to a constant intercept term, therefore requiring a constant term and $q-1$ intervention dummies. The VECM representation can then be

Table 4.3. *Cointegration parameter estimates*

	Loading coeff. (α)				Restricted cointegrating vectors (β')					
					$m{-}p$	y	s	l	π	$ygap$
$m{-}p$	−0.091	0	0	0						
	(0.023)									
y	0	−0.176	0	0	1	−1.583	0	0	0	0
		(0.094)				(0.026)				
s	0	0	0	0.055	0	0	1	−1	0	0
				(0.024)						
l	0	0.129	−0.144	0	0	0	0	1	−1	0
		(0.056)	(0.046)							
π	0.213	0	0.378	0.151	0	0	0	0	0	1
	(0.061)		(0.138)	(0.062)						
$ygap$	0	0	0	−0.136						
				(0.034)						

Overidentifying restrictions test: $\chi^2(22) = 21.6$ (p-value 0.48)

rewritten as

$$\mathbf{\Pi}^*(L)\,\Delta\mathbf{x}_t = \boldsymbol{\nu} + \boldsymbol{\alpha}\,\bar{\boldsymbol{\beta}}'\bar{\mathbf{x}}_{t-1} + \varepsilon_t, \tag{4.10}$$

where $\bar{\mathbf{x}}_t' = (\mathbf{x}_t'\ 1\ \mathbf{d}_t')$ and $\mathbf{d}_t$ is a $(q-1) \times 1$ subvector including the $q-1$ intervention dummies. Denoting the last column of $\boldsymbol{\alpha}$ by $\boldsymbol{\alpha}_k$, equation (4.10) can be expressed in an estimable form as:

$$\mathbf{\Pi}^*(L)\Delta\mathbf{x}_t = \boldsymbol{\nu}^* + \boldsymbol{\alpha}_k\,\boldsymbol{\beta}_2^{*\prime}\,\mathbf{d}_{t-1} + \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{x}_{t-1} + \varepsilon_t, \tag{4.11}$$

where $\boldsymbol{\nu}^* = \boldsymbol{\nu} + \boldsymbol{\alpha}_k\beta_1^*$, β_1^* and $\boldsymbol{\beta}_2^*$ denote respectively the first and the last $q-1$ elements of $\boldsymbol{\beta}^*$, and $\mathbf{d}_t$ contains the $q-1$ intervention dummies. In practice the model can be estimated leaving the deterministic components unrestricted.

2.2 Long-run results

The previously estimated VAR(3) system is then extended to include the (unrestricted) dummy variable *RP* to capture regime shifts in the long-term real interest rate behaviour. The estimation period now spans the full sample, from 1978Q4 to 2001Q3. Diagnostic tests on the whole system do not detect any sign of autocorrelation (supporting the choice of a three-lag specification) and heteroscedasticity. Only some residual non-normality is detected in the *ygap* equation.

Since formal Johansen's (1995) tests for the cointegration rank cannot be used owing to the presence of the *RP* dummy variable, we rely on visual inspection

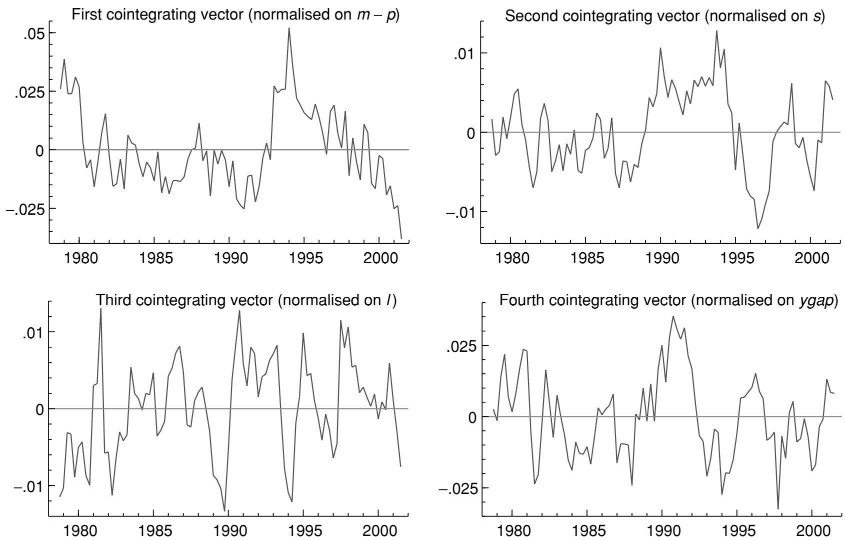

Figure 4.3 Restricted cointegrating vectors

and proceed under the assumption that there exist four valid cointegrating relationships among the variables in **x**. In accord with the theoretical framework illustrated above, we interpret such relationships as a long-run money demand function involving $m - p$ and y, a term structure equation between s and l, a Fisher parity relation linking l and π, and a long-run constant rate of capacity utilisation (stationarity of *ygap*). The LR test of the resulting set of seven overidentifying restrictions on the coefficients of $\boldsymbol{\beta}$ yields a $\chi^2(7)$ statistic of 9.1, with a corresponding p-value of 0.25, strongly supporting the chosen identification scheme. If additional zero restrictions are imposed on the loading parameter in $\boldsymbol{\alpha}$ we obtain a $\chi^2(22)$ test statistic of 21.6 with a p-value of 0.48. The restricted loading factors and cointegration parameter estimates are reported in table 4.3, and the four (restricted) cointegrating vectors are shown in figure 4.3. The money demand long-run elasticity to income is very precisely estimated and considerably larger than unity (in line with the results in Gerlach and Svensson 2001), explaining the $I(1)$ feature of money velocity mentioned above.[7] Recursive estimation over the 1995–2001 subperiod shows that this elasticity

[7] When the term structure and the Fisher long-run relations are estimated without imposing $(1, -1)$ cointegrating vectors, the following results are obtained:

$$l = \underset{(0.069)}{0.951} s, \quad \underset{(0.074)}{1.048} \pi,$$

supporting the imposed restrictions.

is remarkably stable over time. The estimated loading parameters show that positive deviations from the equilibrium relation between $m - p$ and y cause a strong upward pressure on inflation and output and an error-correcting reaction of real money balances. An increase of the short-term interest rate relative to the long rate determines a negative reaction of output and an equilibrating response of the long-term rate. The long-term interest rate exhibits error-correcting behaviour also in response to positive deviations from the Fisher parity relation with the inflation rate. Finally, increases in the capacity utilisation rate have a positive impact on inflation (a 'Phillips curve') and on the short-term interest rate (a 'Taylor rule' effect). The whole set of overidentifying restrictions on the loading factors and the cointegrating vector parameters is never rejected at the 5% significance level when the system is estimated recursively from 1995.

3 Permanent and transitory components of inflation

The long-run (cointegration) properties of the data analysed in the previous section may then be used to disentangle the short- and long-run ('*core*') components of the variables analysed, as shown by Stock and Watson (1988) and Gonzalo and Granger (1995). To this aim, we apply the common trends methodology of King et al. (1991) and Mellander, Vredin and Warne (1992) to our small-scale macroeconomic system and focus in particular on the inflation rate. In this context, core inflation is interpreted as the long-run forecast of the inflation rate conditional on the information contained in the variables of the system and consistent with the long-run cointegration properties of the data. A similar definition of core inflation is adopted by Cogley and Sargent (2001) in their analysis of the dynamic behaviour of post-war US inflation. Moreover, in a multivariate system, structural shocks are likely to be identified more precisely than, for example, in the bivariate approach of Quah and Vahey (1995), and the forecast error variance decomposition can yield meaningful information about the dynamic effects of different disturbances on the inflation process. The rest of this section outlines and applies this econometric methodology to euro-area data.

3.1 Econometric methodology

As in Mellander, Vredin and Warne (1992) and Warne (1993), the cointegrated VAR in (4.11) can be inverted to yield the following stationary Wold representation for $\Delta\mathbf{x}_t$ (henceforth, deterministic terms, including the constant vector $\boldsymbol{\nu}^*$ and the dummy variable vector $\mathbf{d}$ capturing different real interest rate regimes are omitted for ease of exposition):

$$\Delta\mathbf{x}_t = \mathbf{C}(L)\,\varepsilon_t, \tag{4.12}$$

where $\mathbf{C}(L) = \mathbf{I} + \mathbf{C}_1 L + \mathbf{C}_2 L^2 + \ldots$ with $\sum_{j=0}^{\infty} j \mid \mathbf{C}_j \mid < \infty$. From the representation in (4.12) the following expression for the levels of the variables can be derived by recursive substitution:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{C}(1) \sum_{j=0}^{t-1} \varepsilon_{t-j} + \mathbf{C}^*(L)\varepsilon_t, \tag{4.13}$$

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 $\mathbf{x}$ and $\mathbf{x}_0$ is the initial observation in the sample.

In order to obtain an economically meaningful interpretation of the dynamics of the variables of interest from the reduced form representations in (4.12) and (4.13), the vector of reduced form disturbances ε must be transformed into a vector of underlying, 'structural' shocks, some with *permanent* effects on the level of $\mathbf{x}$ and some with only *transitory* effects. Let us denote this vector of i.i.d. structural disturbances as $\boldsymbol{\varphi}_t \equiv \binom{\boldsymbol{\psi}_t}{\boldsymbol{\nu}_t}$, where $\boldsymbol{\psi}$ and $\boldsymbol{\nu}$ are subvectors of $n-k$ and k elements, respectively. The structural form for the first difference of $\mathbf{x}_t$ is:

$$\Delta \mathbf{x}_t = \boldsymbol{\Gamma}(L)\boldsymbol{\varphi}_t \tag{4.14}$$

where $\boldsymbol{\Gamma}(L) = \boldsymbol{\Gamma}_0 + \boldsymbol{\Gamma}_1 L + \ldots$. Since the first element of $\mathbf{C}(L)$ in (4.12) is $\mathbf{I}$, equating the first term of the right-hand sides of (4.12) and (4.14) yields the following relationship between the reduced form and the structural shocks:

$$\varepsilon_t = \boldsymbol{\Gamma}_0 \boldsymbol{\varphi}_t, \tag{4.15}$$

where $\boldsymbol{\Gamma}_0$ is an invertible matrix. Hence, comparison of (4.14) and (4.12) 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 $\boldsymbol{\psi}_t$ as the permanent shocks and the elements of $\boldsymbol{\nu}_t$ as the transitory disturbances, the following restriction on the long-run matrix $\boldsymbol{\Gamma}(1)$ must be imposed:

$$\boldsymbol{\Gamma}(1) = (\boldsymbol{\Gamma}_g\ \mathbf{0}), \tag{4.16}$$

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

From (4.14) the structural form representation for the endogenous variables in levels is derived as

$$\mathbf{x}_t = \mathbf{x}_0 + \boldsymbol{\Gamma}(1) \sum_{j=0}^{t-1} \boldsymbol{\varphi}_{t-j} + \boldsymbol{\Gamma}^*(L)\boldsymbol{\varphi}_t = \mathbf{x}_0 + \boldsymbol{\Gamma}_g \sum_{j=0}^{t-1} \boldsymbol{\psi}_{t-j} + \boldsymbol{\Gamma}^*(L)\boldsymbol{\varphi}_t, \tag{4.17}$$

where the partition of ϕ and the restriction in (4.16) have been used and $\mathbf{\Gamma}^*(L)$ is defined analogously to $\mathbf{C}^*(L)$ in (4.13). The permanent part in (4.17), $\sum_{j=0}^{t-1} \boldsymbol{\psi}_{t-j}$, may be expressed as an $(n-k)$-vector random walk $\boldsymbol{\tau}$ with innovations $\boldsymbol{\psi}$:

$$\boldsymbol{\tau}_t = \boldsymbol{\tau}_{t-1} + \boldsymbol{\psi}_t = \boldsymbol{\tau}_0 + \sum_{j=0}^{t-1} \boldsymbol{\psi}_{t-j}. \tag{4.18}$$

Using (4.18) in (4.17), we finally obtain the common trend representation of Stock and Watson (1988) for $\mathbf{x}_t$:

$$\begin{aligned} \mathbf{x}_t &= \underbrace{\mathbf{x}_0 + \mathbf{\Gamma}_g \boldsymbol{\tau}_t} + \underbrace{\mathbf{\Gamma}^*(L)\boldsymbol{\varphi}_t} \\ \Rightarrow \mathbf{x}_t &= \qquad \mathbf{x}_t^c \qquad + \qquad \mathbf{x}_t^{nc}, \end{aligned} \tag{4.19}$$

where $\mathbf{x}_t^c$ and $\mathbf{x}_t^{nc}$ correspond to the 'trend' and 'cycle' components in the Beveridge–Nelson–Stock–Watson decomposition of $\mathbf{x}_t$. According to (4.19) the trend behaviour of the variables is determined by the permanent disturbances only, whereas the cyclical component is determined by all innovations in the system, both permanent and transitory. This implies that permanent innovations also induce transitory dynamics.

As shown in detail by Stock and Watson (1988), King et al. (1991) and Warne (1993), the identification of separate permanent shocks requires a sufficient number of restrictions on the long-run impact matrix $\mathbf{\Gamma}_g$ in (4.19). Part of these restrictions are provided by the cointegrating relations and the consistent estimation of $\mathbf{C}(1)$; additional ones are suggested by economic theory (e.g. long-run neutrality assumptions). Finally, having estimated $\mathbf{\Gamma}_g$, the behaviour of the variables in $\mathbf{x}_t$ due to the permanent disturbances only, interpreted as the long-run forecast of $\mathbf{x}_t$, may be computed as $\mathbf{x}_0 + \mathbf{\Gamma}_g \boldsymbol{\tau}_t$. Formally, such a long-run forecast can be expressed as

$$\lim_{h \to \infty} E_t\, \mathbf{x}_{t+h} = \mathbf{x}_0 + \mathbf{\Gamma}_g \boldsymbol{\tau}_t, \tag{4.20}$$

capturing the values to which the series are expected to converge once the effect of the transitory shocks have died out (Cogley and Sargent 2001). Moreover, from the moving average representation in (4.14), impulse responses and forecast error variance decompositions may be calculated to gauge the relative importance of permanent and transitory innovations in determining fluctuations of the endogenous variables.

3.2 *Results*

In our common trends framework, the existence of four cointegrating vectors in the six-variable system implies the presence of two sources of shocks having permanent effects on at least some of the variables in $\mathbf{x}'$ ($m - p$, y, s, l, π and *ygap*). As previously mentioned, the four (restricted) cointegrating vectors provide a set of restrictions that can be used to identify the elements of $\mathbf{\Gamma}_g$ in (4.19). However, one additional restriction is needed to achieve identification. To this aim, we make the following assumption on the nature of the two permanent shocks in the system: we consider a *real* shock (ψ_r) and a *nominal* disturbance (ψ_n). The permanent part (4.18) of the common trends representation is then given by the following bivariate random walk:

$$\begin{pmatrix} \tau_r \\ \tau_n \end{pmatrix}_t = \begin{pmatrix} \mu_r \\ \mu_n \end{pmatrix} + \begin{pmatrix} \tau_r \\ \tau_n \end{pmatrix}_{t-1} + \begin{pmatrix} \psi_r \\ \psi_n \end{pmatrix}_t, \tag{4.21}$$

where μ is a vector of constant drift terms. Consistent with the theoretical framework sketched in section 2, as an additional restriction we assume that 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 $\mathbf{\Gamma}_g$, this neutrality assumption implies $\gamma_{22} = 0$. Given the long-run relationship linking output and real money balances only, an implication of the long-run neutrality restriction is that the nominal trend does not have a long-run impact on real money balances as well ($\gamma_{12} = 0$). In addition, the same long-run money demand relation implies that the response of $m - p$ to the real permanent shock (γ_{11}) is given by $\beta_1^m \gamma_{21}$, with the estimated value of β_1^m being 1.583 (see results in table 4.3). Moreover, the cointegration properties of the interest rates and inflation also imply that $\gamma_{31} = \gamma_{41} = \gamma_{51}$ and $\gamma_{32} = \gamma_{42} = \gamma_{52}$: in the long-run a permanent disturbance (either real or nominal) has the same effect on s, l and π. Finally, since the output gap is a stationary variable and therefore not affected by permanent shocks in the long run, we have $\gamma_{61} = \gamma_{62} = 0$. The common trends representation of the variables in levels (4.19) becomes therefore the following:

$$\begin{pmatrix} m-p \\ y \\ s \\ l \\ \pi \\ ygap \end{pmatrix}_t = \begin{pmatrix} m-p \\ y \\ s \\ l \\ \pi \\ ygap \end{pmatrix}_0 + \begin{pmatrix} \beta_1^m \gamma_{21} & 0 \\ \gamma_{21} & 0 \\ \gamma_{31} & \gamma_{32} \\ \gamma_{31} & \gamma_{32} \\ \gamma_{31} & \gamma_{32} \\ 0 & 0 \end{pmatrix} \begin{pmatrix} \tau_r \\ \tau_n \end{pmatrix}_t + \mathbf{\Gamma}^*(L) \begin{pmatrix} \psi_r \\ \psi_n \\ \nu_1 \\ \nu_2 \\ \nu_3 \\ \nu_4 \end{pmatrix}_t, \tag{4.22}$$

where the ν_is ($i = 1,2,3,4$) are 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.

The estimated core inflation series from the common trends model is then computed as $\hat{\pi}_t^c = \pi_0 + \hat{\gamma}_{31}\hat{\tau}_{r,t} + \hat{\gamma}_{32}\hat{\tau}_{n,t}$. Such a measure captures the

Table 4.4. *Common trends model*

	Long-run effects ($\mathbf{\Gamma}_g$)		Long-run(∞) forecast error variance explained by:	
Variable	ψ_r	ψ_n	ψ_r	ψ_n
$m-p$	0.980* (0.395)	0	1	1
y	0.619* (0.249)	0	1	0
s	1.104 (0.177)	0.384** (0.085)	0.069 (0.225)	0.931** (0.225)
l	0.104 (0.177)	0.384** (0.085)	0.069 (0.225)	0.931** (0.225)
π	0.104 (0.177)	0.384** (0.085)	0.069 (0.225)	0.931** (0.225)
$ygap$	0	0	0	0

Notes: ψ_r and ψ_n denote the real and nominal permanent shocks respectively; asymptotic standard errors in parentheses; * and ** denote statistical significance at the 5% and 1% level, respectively.

long-run effects on inflation of the two identified permanent disturbances and bears the interpretation of the (conditional) forecast of the inflation rate over a long-term (infinite) horizon, when all transitory fluctuations in the inflation rate have vanished.

The main results from the estimation of the common trends model are shown in table 4.4, where the estimated elements of the long-run impact matrix $\mathbf{\Gamma}_g$ (with asymptotic standard errors in parentheses) and the long-run forecast error variance decomposition of all variables are reported. The estimated long-run effects of permanent shocks show that the real shock (ψ_r), which is the only determinant of the long-run behaviour of real money balances and output, plays only a marginal role in explaining the long-run features of the two interest rates and the inflation rate, which are dominated by nominal disturbances (ψ_n). This finding supports the separation of the long-run properties of real money balances and output on the one hand and nominal interest rates and inflation on the other (as noted also by Cassola and Morana 2002 in a larger system of euro-area variables). Therefore, the measure of core inflation derived from the common trends model is almost entirely explained by the nominal trend. This conclusion is supported also by the result of the forecast error variance decomposition reported in table 4.4, showing that in the long run more than 90% of the inflation rate variability is attributable to the nominal permanent disturbance.

The upper panel of figure 4.4 plots the estimated core inflation series, π^c, the measured HICP inflation, and the 'non-core' inflation rate ($\pi^{nc} \equiv \pi - \pi^c$),

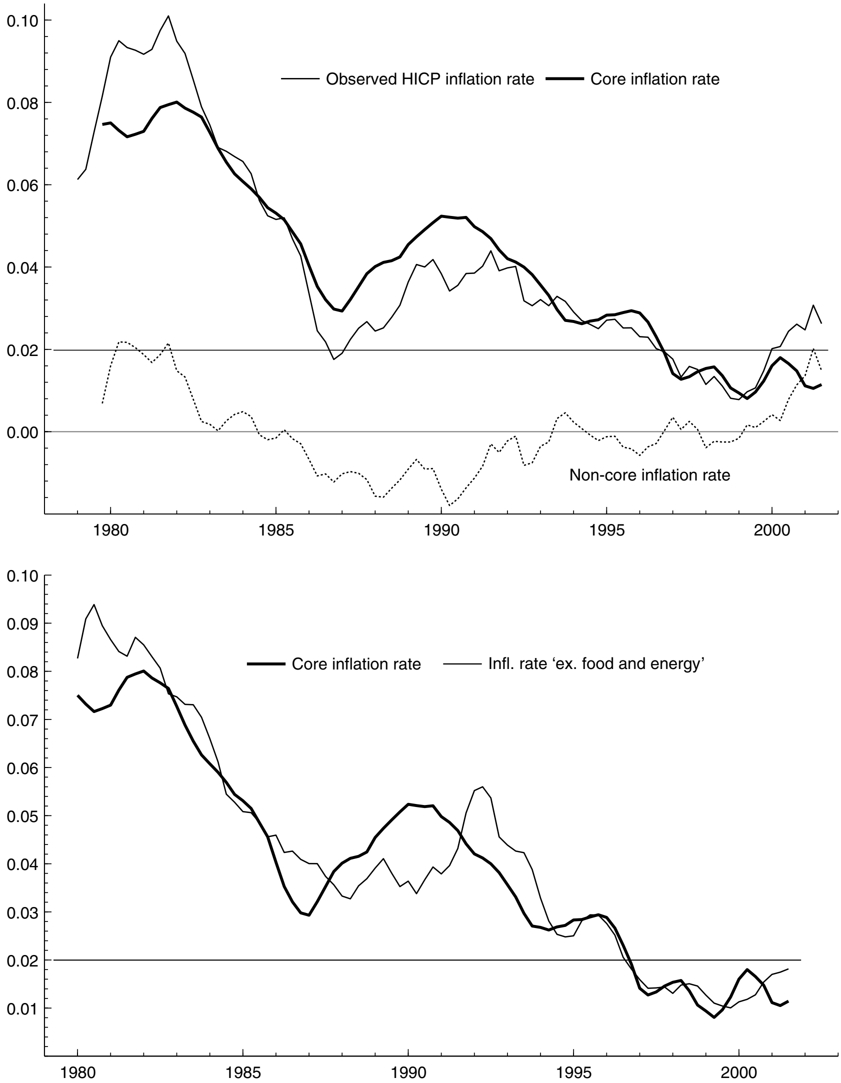

Figure 4.4 Observed annual inflation rates and common trend core and non-core inflation measure

all expressed as annual rates (four-quarter lagged moving averages) over the whole 1979–2001 period. In the lower panel, the common trend measure of core inflation is compared with a widely used measure of the underlying trend in the inflation rate, namely the rate of change of the CPI price level excluding 'food and energy' goods. As shown in the figure, in the 1980s the core inflation rate shows more limited fluctuations, ranging from 3% to 8%, with respect to both observed inflation measures, which vary widely between 2% and 10%. In particular, core inflation displays a lower peak during the oil-shock episode of the early 1980s (around 8% against 9–10% observed inflation rate), whereas this pattern is reversed during the counter-shock in the mid-1980s. Starting in the early 1990s, the various inflation rates show more similar behaviour, though with some notable exceptions, namely in 1991, when the core rate began to decrease rapidly in the face of broadly stable (HICP) or increasing ('ex food and energy') actual inflation. Then, all inflation measures declined below the 2% level at around the same time in the second half of 1996.

Of particular interest is the relative behaviour of the actual and core inflation series since the introduction of the euro in January 1999. Initially, the core and the HICP rates increased from around 1% in early 1999 up to around 2% in mid-2000 (in 2000Q2 the core inflation rate was at 1.8% and the HICP rate at 2.1%). Such an increase is commonly attributed to the sharp rise in oil prices, since the consumer price inflation rate 'excluding food and energy items' remained stable within a 1–1.2% range. However, the forward-looking, common trends measure of core inflation signals that the long-run inflation forecast as of 2000Q2 was very close to the HICP observed inflation, even though the 'ex food and energy' index showed a lower and stable inflation rate. This evidence can lend some support to the prudent monetary policy attitude of the ECB in 1999 and 2000 in the management of policy interest rates. From 2000Q3, the behaviour of the estimated core inflation rate started to diverge from that of the two observed rates. While the HICP continued to increase up to 3.1% in 2001Q2 before going back to 2.6% in the following quarter, and the CPI 'ex food and energy' rate reached 1.8% in 2001Q3, the core inflation rate declined during the second half of 2000 and stabilised at 1.1% in 2001. The increase in inflation observed in 2001 does not, then, necessarily signal higher long-term inflation prospects.

To give reliable information for policy use, a core inflation measure must possess some desirable properties, as stressed by Bryan and Cecchetti (1994) and Wynne (1999). First, the estimated core inflation series should display lower variability and higher persistence than actual inflation. As noted above, the common trends measure of core inflation portrayed in figure 4.4 is less volatile than measured consumer price inflation. The smoothing property of the estimated core inflation is further illustrated in panel A of table 4.5, which reports correlation coefficients among changes in the quarterly and annual (four-quarter moving average) inflation rates, including observed inflation and the common trends core and non-core measures, denoted by $\Delta\pi^c$ and $\Delta\pi^{nc}$ respectively, with $\Delta\pi \equiv \Delta\pi^c + \Delta\pi^{nc}$. Standard deviations in percentage points are shown on the diagonal. These latter statistics show that there is a remarkable difference in variability between the core and the non-core component: standard

Table 4.5. *Assessment of the common trend core inflation measure*

A. *Correlations*

	$\Delta\pi$	$\Delta\pi^c$	$\Delta\pi^{nc}$	$\Delta\pi^c_{NFE}$	$\Delta\pi^{nc}_{NFE}$
	Quarterly inflation rates: 1979Q2–2001Q3				
$\Delta\pi$	0.290				
$\Delta\pi^c$	0.384	0.100			
$\Delta\pi^{nc}$	0.938	0.042	0.268		
$\Delta\pi^c_{NFE}$	0.044	−0.033	0.060	0.259	
$\Delta\pi^{nc}_{NFE}$	0.723	0.315	0.675	−0.647	0.380
	Annual inflation rates: 1980Q1–2001Q3				
$\Delta\pi$	0.358				
$\Delta\pi^c$	0.509	0.210			
$\Delta\pi^{nc}$	0.806	−0.099	0.306		
$\Delta\pi^c_{NFE}$	0.478	0.120	0.470	0.293	
$\Delta\pi^{nc}_{NFE}$	0.639	0.434	0.440	−0.371	0.335

B. *Results from bivariate VAR systems*: 1979Q4–2001Q3

Equation for:	*F* test (*p*-value) on 2 lags of: $\Delta\pi$	$\Delta\pi^c$	$\Delta\pi^c{}_{NFE}$	Coefficient estimate on: $(\pi - \pi^c)_{t-1}$	$(\pi - \pi^c_{NFE})_{t-1}$
$\Delta\pi$	0.009**	0.643		-0.215^{*}	
				(0.107)	
$\Delta\pi^c$	0.941	0.816		0.013	
				(0.045)	
$\Delta\pi$	0.037^{*}		0.761		−0.165
					(0.126)
$\Delta\ \pi^c{}_{NFE}$	0.579		0.192		0.332^{**}
					(0.109)

Notes: $\Delta\pi$ denotes the first difference of the measured HICP inflation rate; $\Delta\pi^c$ and $\Delta\pi^c_{NFE}$ denote the first differences of the common trend measure of core inflation and of the CPI 'excluding food and energy' inflation rates respectively; $\Delta\pi^{nc}$ and $\Delta\pi^{nc}_{NFE}$ are the associated non-core inflation changes, defined as $\Delta\pi^{nc} = \Delta\pi - \Delta\pi^c$ and $\Delta\pi^{nc}_{NFE} = \Delta\pi - \Delta\pi^c_{NFE}$. The figures on the main diagonals in panel A are standard deviations in percentage points (quarterly inflation rates are not annualised). * and ** denote statistical significance at the 5% and 1% levels, respectively.

deviations are 0.10 and 0.27 for $\Delta\pi^c$ and $\Delta\pi^{nc}$ respectively in quarterly data (0.21 and 0.31 in annual data), with a standard deviation of changes in the observed inflation rate of 0.29 (0.36). We also note that quarterly changes in observed inflation are much more closely correlated with changes in the non-core component (the correlation coefficient is 0.94) than with changes in the estimated core rate (0.38), and that there is a very low correlation between core and non-core inflation changes (0.04 in quarterly and −0.10 in annual data).

Panel A of Table 4.5 also reports standard deviations and correlations of the change in the CPI inflation rate 'excluding food and energy' goods, $\Delta\pi_{NFE}$ (the associated transitory inflation component is denoted by $\Delta\pi^{nc}_{NFE} \equiv \Delta\pi - \Delta\pi^{c}_{NFE}$). The standard deviations of changes in both inflation components obtained from the ex-food and energy price level are large (0.26 for $\Delta\pi^{c}_{NFE}$ and 0.38 for $\Delta\pi^{nc}_{NFE}$ in quarterly data), suggesting that this inflation indicator does not possess the smoothing property displayed by the common trends core inflation measure.

A second desirable property of a core inflation measure is the ability to forecast future headline inflation rates. The long-run forecasting power of our common trends measure is warranted, since it is estimated as the long-run conditional forecast of inflation. This property can be formally assessed by means of a bivariate VAR system including the observed inflation rate and core inflation π^c. As argued by Freeman (1998), the integration and cointegration properties of the inflation series require an error-correction representation to perform appropriate Granger-causality tests. In fact, both π and π^c are non-stationary, $I(1)$ series, whereas the associated non-core component π^{nc} displays stationarity, which may be interpreted as evidence of cointegration between the core inflation measure and the actual inflation rate, since $\pi^{nc} \equiv \pi - \pi^c$. The specification of the bivariate system is then the following:

$$\Delta\pi_t = \delta_{10} + \sum_{i=1}^{2} \delta_{11}(i)\Delta\pi_{t-i} + \sum_{i=1}^{2} \delta_{12}(i)\Delta\pi^c_{t-i} + \rho_\pi(\pi - \pi^c)_{t-1} + u_{1t}$$

$$\Delta\pi^c_t = \delta_{20} + \sum_{i=1}^{2} \delta_{21}(i)\Delta\pi_{t-i} + \sum_{i=1}^{2} \delta_{22}(i)\Delta\pi^c_{t-i} + \rho_c(\pi - \pi^c)_{t-1} + u_{2t}, \tag{4.23}$$

where two lags are sufficient to eliminate residual serial correlation. Panel B of table 4.5 reports the results of the F-tests on each block of lagged regressors and the coefficient estimates of the error-correction coefficients ρ_π and ρ_c. Although lags of $\Delta\pi^c$ do not have additional predictive power for the actual inflation rate, a sizeable and significant error-correction coefficient ρ_π (−0.22) is estimated, showing a tendency of actual inflation to adjust to the core component, whereas no adjustment is detected in the behaviour of π^c. We also estimated the bivariate system in (4.23) with π^c_{NFE} in the place of π^c. The ex-food and energy inflation measure does not show any strong additional predictive power for the observed inflation rate. Moreover, the positive and strongly significant estimated error-correction coefficient on $(\pi - \pi^c_{NFE})_{t-1}$ suggests that past values of the inflation rate above the 'underlying' component measured by π^c_{NFE} cause an increase in π^c_{NFE} itself, reflecting the transmission of transitory shocks to the permanent component of inflation and casting some doubts on the usefulness of this measure as an indicator of the long-run inflation trend.

4 A closer look at the properties of inflation components

The common trends model applied in the preceding section decomposes observed inflation into a long-run, core component and a transitory, non-core element. In this section we analyse several features of this decomposition, starting from the sources of temporary fluctuations in the inflation rate captured by the non-core component. Then, we investigate how long it takes for the inflation rate to converge to the core inflation rate, interpreted as a long-run inflation forecast. Finally, we compare the estimated core inflation rate with the inflation forecast at various horizons obtained from a structural dynamic model encompassing the VAR.

4.1 The nature of the cyclical inflation component

By construction, the common trends core inflation measure embeds only the information contained in the permanent shocks hitting the system, abstracting from the more volatile dynamics generated by transitory shocks. However, the latter disturbances may not be the only sources of inflation fluctuations around the core component. In fact, an important property of the Beveridge–Nelson–Stock–Watson decomposition is that the 'cyclical' (here interpreted as the 'non-core') component π^{nc} is explained not only by transitory shocks, but also by permanent shocks. Proietti (1997) has proposed a methodology to disentangle in cyclical fluctuations the contribution of permanent shocks from the effect of transitory disturbances. Following Cassola and Morana (2002), a similar decomposition of the cycles can be obtained by rewriting the vector of cyclical components $\mathbf{x}^{nc}$ as

$$\mathbf{x}^{nc}_t = \mathbf{\Gamma}^*(L)\,\boldsymbol{\varphi}_t = \mathbf{\Gamma}^*_1(L)\,\boldsymbol{\psi}_t + \mathbf{\Gamma}^*_2(L)\,\boldsymbol{v}_t. \tag{4.24}$$

The vector $\mathbf{\Gamma}^*_1(L)\boldsymbol{\psi}_t$ gives the contribution of permanent innovations to the overall cycle (henceforth referred to as the 'dynamics *along* the attractor', DAA), while the vector $\mathbf{\Gamma}^*_2(L)\boldsymbol{v}_t$ measures the contribution of the transitory innovations to the overall cycle ('dynamics *towards* the attractor', DTA).

The latter kind of short-run dynamics have the error-correction process as generator and, therefore, are disequilibrium fluctuations, while the dynamics along the attractor may be related to the overshooting of the variables

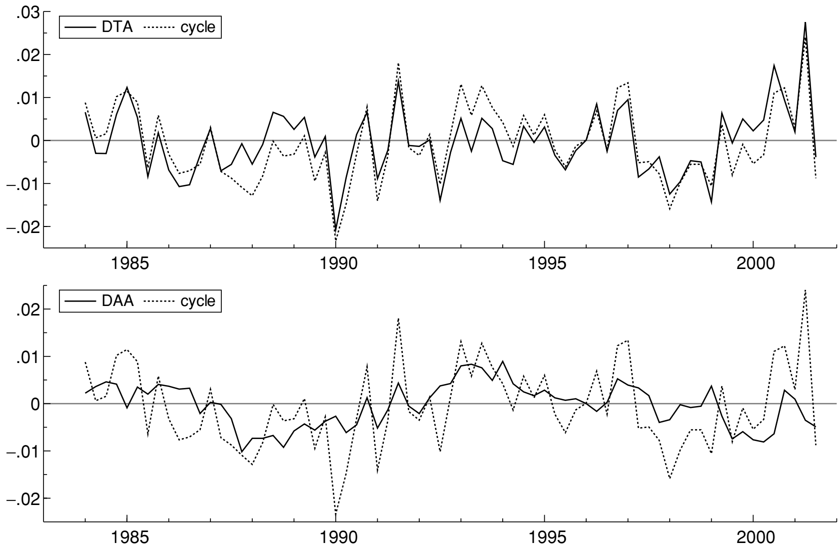

Figure 4.5 Non-core quarterly inflation rate 1984Q1−2001Q3
Note: The quarterly non-core inflation rate is decomposed into 'dynamics towards the attractor' (DTA) and the 'dynamics along the attarctor' (DAA) components.

to permanent innovations, i.e. they are the transitional dynamics which take place following a shock to the common trend. Since along the attractor the cointegration relationships are satisfied, the DAA adjustment captures equilibrium fluctuations. This distinction is of particular interest here since it allows us to attribute deviations of observed inflation from its core rate to the effects of transitory shocks and to the overshooting of the system to permanent shocks.

The decomposition of the non-core quarterly inflation rate into the DTA and DAA components is plotted in figure 4.5. After some experimentation we concluded that twenty lags are sufficient to reconstruct the cyclical components, so that our analysis focuses on the period starting in 1984Q1. As shown in the figure, both cyclical components are important determinants of the short-run inflation dynamics, with the DTA capturing most of the fluctuations. Over the reconstruction period the DTA explain about 50% of the unconditional variance of non-core inflation, while the contribution of the DAA is 38%.[8] Of the latter proportion, 49% is explained by the real permanent shock ψ_r and 32% by the nominal permanent disturbance ψ_n.

[8] The fractions of variance need not to sum to one, since the orthogonality of structural shocks holds only on the entire estimation period, 1978Q1−2001Q3.

4.2 Convergence to the core inflation rate

Our proposed measure of core inflation bears the interpretation of long-run inflation forecast, i.e. $\pi_t^c = \lim_{h\to\infty} E_t \pi_{t+h}$. Although a long-run perspective is consistent with the monetary policymakers' ability to influence the price level, an infinite horizon is not literally appropriate for the purposes of policy analysis; for example, the ECB price stability objective is explicitly referred to as a 'medium-term' horizon. Then, for the common trend measure of core inflation to provide useful information to policymakers on the consistency of current inflation developments with their longer-term price stability goal, it is important to assess how long it takes for transitory and permanent shocks to exhaust their effects on the non-core inflation component π^{nc}, i.e. how long it takes for the observed inflation rate π to converge to the long-run forecast π^c.

In order to provide some empirical evidence on this issue, we estimated the impulse response functions of the non-core inflation rate to the various structural disturbances. Figure 4.6 shows the impulse responses of the non-core inflation rate to the real and nominal permanent shocks, to a composite permanent shock, i.e. the sum of the two permanent shocks, capturing the 'dynamics along the attractor', and to a composite transitory shock, i.e. the sum of the four transitory shocks, capturing the 'dynamics towards the attractor'.

As shown in the figure, both composite disturbances have short-lived effects on non-core inflation; in particular, transitory shocks tend to be inflationary whereas permanent disturbances tend to be deflationary, and complete convergence to the reference value is achieved within six and twenty quarters for the DTA shock and the DAA shock, respectively (consistent with the result of the decomposition of the overall short-run inflation fluctuations in the previous subsection). As far as the DAA composite disturbance is concerned, the response of non-core inflation is dominated by the reaction to the real permanent shock, with inflation falling as productivity increases. According to the estimated significance bands (one standard error), the responses of π^{nc} are not statistically different from zero after only a few quarters (one and six quarters for DTA and DAA shocks, respectively), suggesting that the overall inflation rate quickly reverts to its long-run, core component. The empirical evidence therefore supports the proposed core inflation measure as a potentially useful indicator of long-run inflation prospects over a horizon appropriate for monetary policy evaluation.

4.3 Forecasting inflation from a structural dynamic model

Finally, we compare the estimated core inflation from the common trends model with the forecast of a structural econometric model (SEM) derived from the cointegrated VAR system previously estimated. Starting from the cointegrated

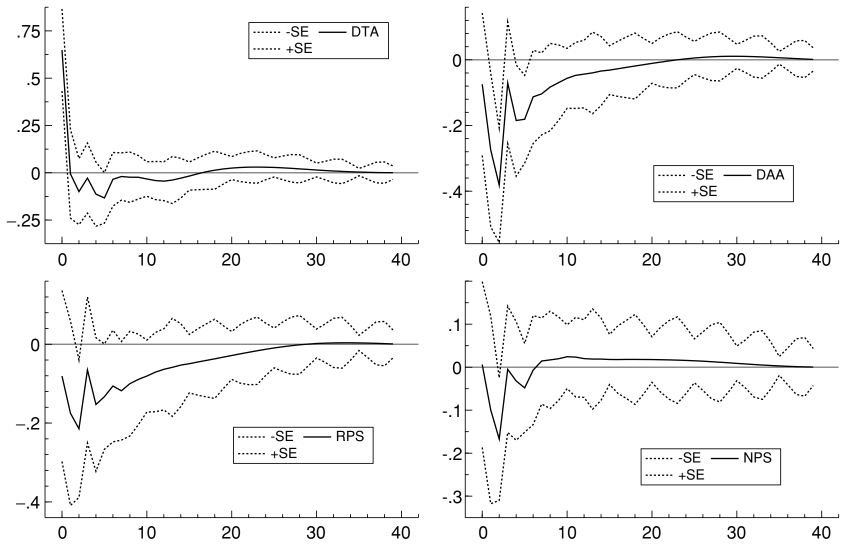

Figure 4.6 Responses of the non-core quarterly inflation rate to shocks.
Note: The upper panels show the impulse response functions of the non-core quarterly inflation rate π^{nc} to composite transitory shocks (DTA), and to composite permanent shocks (DAA). The lower panels show the impulse response functions of π^{nc} to the real permanent shock ψ_r (RPS), and to the nominal permanent disturbance ψ_n (NPS). One-standard error confidence bands have been computed by Monte Carlo simulations, with 1000 replications.

VAR set up in section 2.2, we followed a 'general-to-specific' modelling strategy (Hendry and Mizon, 1993). Zero restrictions were imposed in successive steps on several lags of the endogenous variables in the six equations of the system; after each step a test of the overidentifying restrictions was performed, supporting the restrictions imposed. The FIML estimates of the final specification of the SEM are shown in table 4.6, where deterministic terms are not reported for brevity.

The system diagnostic tests show the data congruence of the SEM and recursive one-step and break-point Chow tests support parameter stability. The tracking of the model is good (the residual standard errors are relatively small) and the test of the whole set of overidentifying restrictions has a p-value of 0.96 (beside the zero restrictions, one additional parameter restriction is imposed in the Δy equation, with a p-value of 0.53). Moreover, the residual correlation matrix shows low coefficients (usually lower than 0.3), suggesting the success of the modelling strategy. Finally, the static long-run real interest rate estimates reported in the bottom part of table 4.6. are consistent with both the Markov-switching results in section 2.2 and with those in Gerlach and Schnabel (2000).

Table 4.6. *The structural dynamic model (FIML estimates)*

	Equation for:					
	$\Delta(m-p)$	Δy	Δs	Δl	$\Delta\pi$	$\Delta ygap$
$\Delta(m-p)_{t-1}$	0.416	–	–	–	–	–
	(0.089)					
$\Delta(m-p)_{t-2}$	0.219	0.351	–	0.160	–	–
	(0.088)	(0.090)		(0.059)		
Δy_{t-1}	–	–	0.116	–	–	0.362
			(0.064)		(0.093)	
Δy_{t-2}	0.118	–	–	–	–	0.239
	(0.062)					(0.097)
Δs_{t-1}	–	–	0.191	–	–	–
			(0.099)			
Δs_{t-2}	−0.150	0.197	−0.244	−0.174	–	−0.318
	(0.080)	(–)	(0.094)	(0.076)		(0.140)
Δl_{t-1}	−0.340	–	0.380	0.623	–	–
	(0.090)		(0.121)	(0.090)		
Δl_{t-2}	–	–	–	–	–	0.419
						(0.156)
$\Delta\pi_{t-1}$	0.093	–	0.072		−0.117	−0.337
	(0.041)		(0.037)	(0.046)	(0.122)	
$\Delta\pi_{t-2}$	–	0.197	–	−0.082	−0.191	–
		(0.043)		(0.032)	(0.091)	
$\Delta ygap_{t-1}$	–	0.196	–	–	–	0.223
		(0.079)				(0.093)
$\Delta ygap_{t-2}$	–	–	–	0.095	–	0.261
				(0.047)		(0.088)
$[(m-p)-1.58y_{t-1}$	−0.079	–	–	–	0.185	–
	(0.020)				(0.050)	
$(s-l)_{t-1}$	–	−0.106	–	0.127	–	–
		(0.072)		(0.045)		
$(l-\pi)_{t-1}$	–	–	–	−0.153	0.370	–
				(0.041)	(0.110)	
$ygap_{t-1}$	–	–	0.068	–	0.176	−0.110
			(0.020)		(0.050)	(0.027)
St. error regression	0.0036	0.0047	0.0039	0.0032	0.0094	0.0048

Misspecification tests:

AR(5) F: 1.08 [0.28] Heter. F: 0.97 [0.65] Norm. $\chi^2(2)$: 23.4 [0.03]

Tests of overidentifying restrictions:

Zero restrictions $\chi^2(63)$: 44.7 [0.96] Other restrictions $\chi^2(1)$: 0.4 [0.53]

Static long-run:

Δy^* (annualised): 0.0249 (0.003), $\beta_0^f = l^* - \pi^*$, low-rate regime: 0.0331 (0.0025)

difference high–low rate regimes: 0.0191 (0. 0028)

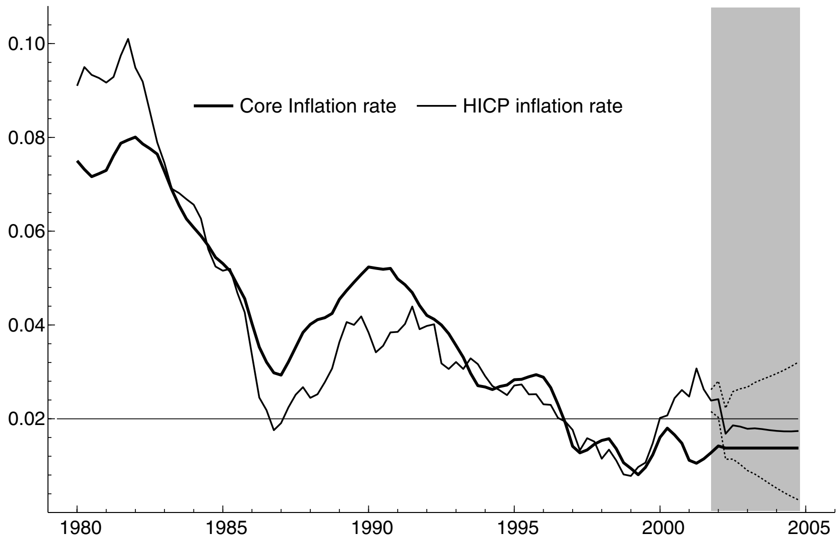

Figure 4.7 Core inflation rate and HICP inflation rate with forecast from a multiple-equation structural dynamic model
Note: One-standard forecast error bands are shown. The shaded area indicates the forecast period: 2001Q4−2004Q4.

Turning to the inflation forecasting issue, figure 4.7 displays the annual HICP inflation rate with point forecast values for the period 2001Q4−2004Q4 from the estimated SEM. The inflation rate is forecast to decline rapidly from the 2.6% level reached in 2001Q3 and stabilise in the 1.7–1.8% range from 2002Q2. Over a two- to three- year horizon these values are broadly consistent with the long-run inflation forecast measured by the common trends core inflation series, which predicts an annual inflation rate around 1.4%.[9]

5 Conclusions

A common trends model has been used to estimate the underlying, ‘core’ inflation behaviour for the euro area from 1978 to 2001. In this framework core inflation is interpreted as the long-run forecast of inflation conditional on the information contained in money growth, output fluctuations and movements of the term structure of interest rates.

[9] The relatively wide one-standard forecast error bands, computed taking into account the error variance only, show that, despite the overall good statistical performance of the econometric model, forecasting accuracy is still insufficient to make point inflation forecast from the SEM a reliable guide for policymakers.

A price stability-oriented monetary policy has to be forward looking and respond only to shocks having long-lasting effects on the inflation rate. The common trends core inflation measure may be useful for monetary policy purposes since it embodies long-run economic restrictions strongly supported by the data and bears the interpretation of a long-run forecast, affected only by permanent disturbances to the inflation rate.

Our empirical exercise on euro-area data shows that purely transitory shocks have short-lived effects on the inflation rate and the estimated core measure captures the permanent component of inflation fluctuations over a medium-term horizon consistent with the monetary policy strategy of the European Central Bank. An important implication of our results is that deviations of core inflation, rather than actual inflation, from the price stability objective convey the appropriate signals for policy action. This conclusion partly contrasts with a large body of the monetary policy literature, where policy behaviour is modelled by means of a standard Taylor rule.[10]

As a final word of caution, we observe that a core inflation rate estimated from a common trend model depends 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 chapter, featuring long-run relationships between real money balances, output, inflation and interest rates, seems an useful benchmark to evaluate the properties of other measures of core inflation currently used in the monetary policy debate. As a first step in this direction we compared the smoothing and forecasting properties of the common trends core inflation with those of the 'ex food and energy' CPI inflation rate. The comparison lends support to our core inflation measure as a more reliable indicator of the long-run inflation trend.

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[10] A first application of a modified monetary policy rule which determines short-term interest rates according to the deviations of core inflation from the price stability objective is provided by Cassola and Morana (2002).

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