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A common trends model of UK core inflation*

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Abstract. In this paper the long-run trend in *RPI* inflation (*core inflation*) for the UK over the 1961–1997 period is estimated within the framework of a multivariate common trends model which extends the bivariate *VAR* approach of Quah and Vahey (1995). In this context core inflation is directly linked to money and wage growth and interpreted as the long-run forecast of inflation from a small-scale, cointegrated macroeconomic system.

Key words: core inflation, common trend, monetary policy.

JEL classification: C32, E31, E52.

1. Introduction

In recent years, price stability has been widely adopted as the main final goal of monetary policy, albeit with a different emphasis across countries. Inflation targeting policies, setting precise quantitative targets for monetary authorities, have been implemented by several central banks during the '90s.[1] Other countries have adopted price stability as a primary objective even without adhering to an explicit inflation targeting strategy. In all cases, though the essence of policies aimed at controlling inflation can be simply stated, qualifications are needed once it is recognized that observed inflation may fluctuate in the short-run due to only temporary disturbances of both real and nominal nature, with no or little impact on medium- to long-term inflation prospects. Therefore, as

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[1] Bernanke, Laubach, Mishkin and Posen (1999) provide a thorough cross-country assessment and a review of the main implementation issues raised by inflation targeting.

argued by Cecchetti (1997), short-run changes in the observed inflation rate should be carefully analyzed in order to extract the long-run, trend component of inflation, commonly referred to as the "underlying" or "*core*" inflation rate.

Accordingly, the empirical study of inflation has become a crucial issue in monetary policy analysis, with the aim of distinguishing persistent sources of inflationary pressures from only transient fluctuations in the inflation rate. Several measures of the core inflation rate have been put forward and used in practical monetary policy conduct.[2] Some are based on purely statistical methodologies, yielding estimators of the inflation rate alternative to the conventional mean of sectoral price changes. Other measures are based on econometric techniques aimed at decomposing economic time-series into permanent and transitory components. In particular, Quah and Vahey (1995) applied to the UK a bivariate structural vector autoregression (*VAR*) approach to core inflation estimation based on the assumption of long-run output neutrality of permanent shocks to the inflation rate.

The present paper extends the bivariate, output-inflation setting of Quah and Vahey (1995) to a multivariate framework. In this context, we interpret UK core inflation as the long-term inflation forecast obtained from a small-scale common trends model, built around (appropriately tested) long-run equilibrium relations between the inflation rate and two main sources of inflationary pressures, namely the money and wage growth rates. In so doing, we follow the lead of Bryan and Cecchetti (1994), who define core inflation as the long-run, persistent, component of the measured inflation rate, "which is tied in some way to monetary growth" (p. 197). Also Quah and Vahey (1995) argue that it would be informative to allow for more variables in the VAR system used to estimate core inflation. In particular, one should consider "monetary and labour cost variables, allowing an assessment of the sources of inflationary pressures" (p. 1143).

The remainder of the paper is organised as follows. In Section 2 we briefly discuss different approaches to core inflation measurement, putting our common trends model into perspective. Section 3 describes the common trends methodology and its application to the UK inflation over a long time period (1961–1997). The estimated core inflation rate is obtained and discussed. Conclusions are provided in the final Section 4.

2. Measures of core inflation

In the last decade some interesting contributions have been made to core inflation measuring. Two main approaches have been followed in the literature and both interpret core inflation as the underlying trend in the inflationary process. However, while the first approach focuses on the cross-sectional distribution of the consumer price index (*CPI*) components, the second concentrates directly on the aggregate price level series.

[2] See the collection of papers published by the Bank for International Settlements (1999). Measures of the core inflation component have been adopted by some central banks as an explicit objective for monetary policy. For example, a concept of core inflation is adopted as operational target of policy in Canada (Johnson, 1999); in Australia, a core inflation rate was the official target in the 1996 *Statement on the conduct of monetary policy* (Cockerell, 1999); finally, the explicit targeting of core inflation is advocated for New Zealand (Cassino, Drew and McCaw, 1999).

The first approach, mainly due to Bryan and Cecchetti (1993, 1994), relies on 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 skewness of the distribution of price changes due to the asymmetric behaviour of price-setters, the median or a trimmed mean are likely to be more robust measures of central tendency than the conventional mean. In addition, the use of 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 no guarantee that these goods may not turn into low-variance components.

The second approach applies various techniques to the aggregate price change series to measure the core inflation component. For example, univariate techniques, such as simple moving averages calculated over a variable time span (from 3–6 up to 36 months) or more sophisticated methodologies (e.g. unobserved component models), are used to smooth the noise component in the inflation pattern.

Insights from the econometric literature on the decomposition of economic time-series into permanent and transitory components have also been used to measure core inflation. Starting with the seminal work of Beveridge and Nelson (1981), different approaches to the permanent-transitory decomposition have been proposed (see Quah (1992) for a general treatment of the identification of permanent versus transitory components in time-series). Blanchard and Quah (1989) have shown how a trend-cycle decomposition may be attained for non cointegrated $I(1)$ variables by constraining the long-run multiplier matrix obtained from their *VAR* representations. Quah and Vahey (1995) applied this methodology to UK data for the 1969–1994 period, estimating the core inflation component from a *VAR* model including industrial production and inflation. In their framework core inflation is defined as that component of the observed inflation rate that has no long-run effect on output.

In this paper we extend the work of Quah and Vahey (1995) by considering a multivariate framework. The information set used to construct an estimate of core inflation is enlarged to include other macroeconomic variables that can play an important role in determining the long-run inflation rate, namely the growth rate of the nominal money stock and the growth rate of nominal wages. In statistical terms such long-run relationships may be conceived in terms of (suitably tested) cointegrating vectors. The cointegration properties of the system may then be used to disentangle the short- and long-run components of the variables analyzed, as shown by Stock and Watson (1988) and Gonzalo and Granger (1995). To this aim, we apply the common trends methodology of King, Plosser, Stock and Watson (1991) and Mellander, Vredin and Warne (1993) to a small-scale macroeconomic system including the inflation rate, output, money and wage growth and the oil price as a supply-side variable.[3] In this context, core inflation is interpreted as the long-run forecast of inflation conditional onto the information contained in the system's variables and consistent with the long-run cointegration properties of the data (Evans and Reichlin,

[3] Blix (1997) uses a common trends framework to implement the Quah-Vahey long-run identification scheme for several countries (including the UK), but with no extension to other macroeconomic variables.

1994).[4] A similar definition of core inflation is adopted by Cogley and Sargent (2000) in their analysis of the dynamic behaviour of post-war US inflation.[5] Moreover, in a multivariate system, structural shocks are likely to be identified more precisely and the forecast error variance decomposition can yield meaningful information about the dynamic effects of different disturbances onto the inflation process. The econometric methodology used is illustrated in detail in the next section.

3. The common trends approach to inflation decomposition

To construct our measure of the core inflation process, we adopt the framework of the common trends model of King, Plosser, Stock and Watson (1991) and Mellander, Vredin and Warne (1992) applied to a system of non-stationary variables. The existence of long-run cointegrating relationships 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 and a stationary transitory element. The former component captures the effect of only permanent shocks and bears the interpretation of long-run forecast of the endogenous variables in the system, including inflation. In the next two subsections we present the econometric methodology employed in the study and the empirical results.

3.1. Methodology

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 error-correction $(VECM)$ form:

$$\mathbf{x}_t = \boldsymbol{\Pi}(L)\mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \tag{3.1}$$

$$\Delta\mathbf{x}_t = \boldsymbol{\Pi}^*(L)\Delta\mathbf{x}_{t-1} + \boldsymbol{\Pi}(1)\mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \tag{3.2}$$

where $\boldsymbol{\varepsilon}_t$ is a vector of identically and independently distributed, serially uncorrelated disturbances with zero mean and variance-covariance matrix $\boldsymbol{\Sigma}$, $\boldsymbol{\Pi}(L) = \boldsymbol{\Pi}_1 + \boldsymbol{\Pi}_2 L + \cdots + \boldsymbol{\Pi}_p L^{p-1}$, $\boldsymbol{\Pi}(1) = \sum_{i=1}^{p} \boldsymbol{\Pi}_i$, $\boldsymbol{\Pi}^*(L) = \boldsymbol{\Pi}_1^* + \boldsymbol{\Pi}_2^* L + \cdots + \boldsymbol{\Pi}_{p-1}^* L^{p-2}$ and $\boldsymbol{\Pi}_i^* = -\mathbf{I} + \boldsymbol{\Pi}_1 + \cdots + \boldsymbol{\Pi}_i$ $(i = 1, \ldots, p-1)$.[6]

If there are $0 < r < n$ cointegrating relations among the variables, $\boldsymbol{\Pi}(1)$ is of reduced rank r and can be expressed as the product of two $(n \times r)$ matrices: $\boldsymbol{\Pi}(1) = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\beta}$ contains the cointegrating vectors, such that $\boldsymbol{\beta}'\mathbf{x}_t$ are sta-

[4] Evans and Reichlin (1994) show that the Quah-Vahey approach yields a measure of core inflation that cannot bear the interpretation of a long-run forecast for the inflation series.

[5] Assuming that the central bank adjusts interest rates so that the inflation rate converges to the policy target in the long-run, Cogley and Sargent (2000) interpret core inflation as the monetary authorities' inflation target. We thank a referee for bringing the Cogley-Sargent paper to our attention.

[6] For ease of exposition, we do not include a constant term in (3.1) and (3.2), that would add a deterministic time trend in the representation for the levels below.

tionary linear combinations of the $I(1)$ variables, and $\boldsymbol{\alpha}$ is a matrix of factor loadings. The resulting cointegrated *VAR* is then:

$$\Delta \mathbf{x}_t = \boldsymbol{\Pi}^*(L)\Delta \mathbf{x}_{t-1} + \boldsymbol{\alpha}\boldsymbol{\beta}'\mathbf{x}_{t-1} + \boldsymbol{\varepsilon}_t \tag{3.3}$$

Following the procedure set out in Mellander, Vredin and Warne (1992), the cointegrated *VAR* in (3.3) can be inverted to yield the following stationary Wold representation for $\Delta \mathbf{x}_t$:

$$\Delta \mathbf{x}_t = \mathbf{C}(L)\boldsymbol{\varepsilon}_t \tag{3.4}$$

where $\mathbf{C}(L) = \mathbf{I} + \mathbf{C}_1 L + \mathbf{C}_2 L^2 + \cdots$ with $\sum_{j=0}^{\infty} j|\mathbf{C}_j| < \infty$. From the representation in (3.4) the following expression for the levels of the variables can be derived by recursive substitution:

$$\mathbf{x}_t = \mathbf{x}_0 + \mathbf{C}(1)\sum_{j=0}^{t-1} \boldsymbol{\varepsilon}_{t-j} + \mathbf{C}^*(L)\boldsymbol{\varepsilon}_t \tag{3.5}$$

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 $\boldsymbol{\varepsilon}$ 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 $\boldsymbol{\beta}'\mathbf{C}(1) = \mathbf{0}$.

In order to obtain an economically meaningful interpretation of the dynamics of the variables of interest from the reduced form representations in (3.4) and (3.5), the vector of reduced form disturbances $\boldsymbol{\varepsilon}$ must be transformed into a vector of underlying, "structural" shocks, some with *permanent* and some with only *transitory* effects on the level of $\mathbf{x}$. Let us denote this vector of structural disturbances as $\boldsymbol{\varphi}_t \equiv \begin{pmatrix} \boldsymbol{\psi}_t \\ \boldsymbol{\nu}_t \end{pmatrix}$, where $\boldsymbol{\psi}$ and $\boldsymbol{\nu}$ are subvectors of k and r elements respectively. The structural form for the first difference of $\mathbf{x}_t$ is:

$$\Delta \mathbf{x}_t = \boldsymbol{\Gamma}(L)\boldsymbol{\varphi}_t \tag{3.6}$$

where $\boldsymbol{\Gamma}(L) = \boldsymbol{\Gamma}_0 + \boldsymbol{\Gamma}_1 L + \cdots$ and the previously defined vector $\boldsymbol{\varphi}_t$ 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:

$$\boldsymbol{\varepsilon}_t = \boldsymbol{\Gamma}_0 \boldsymbol{\varphi}_t \tag{3.7}$$

where $\boldsymbol{\Gamma}_0$ is an invertible matrix. Hence, comparison of (3.6) and (3.4) 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 $\boldsymbol{\psi}_t$ 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:

$$\boldsymbol{\Gamma}(1) = (\boldsymbol{\Gamma}_g \quad \mathbf{0}) \tag{3.8}$$

with $\boldsymbol{\Gamma}_g$ an $n \times k$ 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 $\boldsymbol{v}_t$ are restricted to have only transitory effects.

From (3.6), the structural form representation for the endogenous variables in levels is derived as

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

$$= \mathbf{x}_0 + \boldsymbol{\Gamma}_g\sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j} + \boldsymbol{\Gamma}^*(L)\boldsymbol{\varphi}_t \tag{3.9}$$

where the partition of $\boldsymbol{\varphi}$ and the restriction in (3.8) have been used and $\boldsymbol{\Gamma}^*(L)$ is defined analogously to $\mathbf{C}^*(L)$ in (3.5). The permanent part in (3.9), $\sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j}$, may be expressed as a k-vector random walk with innovations $\boldsymbol{\psi}$:

$$\boldsymbol{\tau}_t = \boldsymbol{\tau}_{t-1} + \boldsymbol{\psi}_t$$

$$= \boldsymbol{\tau}_0 + \sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j} \tag{3.10}$$

Using (3.10) in (3.9) we finally obtain the common trend representation for $\mathbf{x}_t$:

$$\mathbf{x}_t = \mathbf{x}_0 + \boldsymbol{\Gamma}_g\boldsymbol{\tau}_t + \boldsymbol{\Gamma}^*(L)\boldsymbol{\varphi}_t \tag{3.11}$$

The identification of separate permanent shocks requires restrictions on the long-run impact matrix $\boldsymbol{\Gamma}_g$ in the common trend model (3.11). 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 $\boldsymbol{\Gamma}_0$). Under these restrictions all shocks can be given an economic interpretation.

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

$$\boldsymbol{\beta}'\boldsymbol{\Gamma}_g = \mathbf{0} \tag{3.12}$$

(since $\boldsymbol{\beta}'\boldsymbol{\Gamma}(1) = \boldsymbol{\beta}'\mathbf{C}(1)\boldsymbol{\Gamma}_0 = \mathbf{0}$), yielding kr linear restrictions. Moreover, from (3.5) and (3.9) we find that $\mathbf{C}(1)\boldsymbol{\varepsilon}_t = \boldsymbol{\Gamma}_g\boldsymbol{\psi}_t$. Hence, since $E(\boldsymbol{\psi}_t\boldsymbol{\psi}_t') = \mathbf{I}$ and $\mathbf{C}(1)$ has reduced rank k, an additional $k(k+1)/2$ restrictions on the elements of $\boldsymbol{\Gamma}_g$ are provided by:

$$\mathbf{C}(1)\boldsymbol{\Sigma}\mathbf{C}(1)' = \boldsymbol{\Gamma}_g\boldsymbol{\Gamma}_g' \tag{3.13}$$

The remaining $k(k-1)/2$ restrictions needed for (exact) identification of $\boldsymbol{\Gamma}_g$ have to be derived from economic theory. The elements of $\mathbf{C}(1)$ and $\boldsymbol{\Sigma}$ can be consistently estimated from the VAR model and $\boldsymbol{\Gamma}_g$ 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: $\mathbf{C}(1)\boldsymbol{\varepsilon}_t = \boldsymbol{\Gamma}_g\boldsymbol{\psi}_t$. We get: $\boldsymbol{\psi}_t = (\boldsymbol{\Gamma}_g\boldsymbol{\Gamma}_g')^{-1}\boldsymbol{\Gamma}_g'\mathbf{C}(1)\boldsymbol{\varepsilon}_t$. 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 $\mathbf{x}_0 + \boldsymbol{\Gamma}_g\sum_{j=0}^{t-1}\boldsymbol{\psi}_{t-j}$.

Estimation of the common trend model is discussed in detail in King, Plosser, Stock and Watson (1991), Mellander, Vredin and Warne (1992) and

Warne (1993). From the moving average representation in (3.6) impulse responses to permanent and transitory innovations and forecast error variance decompositions may be derived.

3.2. Empirical results

We apply the methodology described above to monthly data for the UK over the period 1961–1997. In the empirical analysis we consider five variables, namely the log of the price of oil in US dollars (oil), the log of the industrial production index (y), the monthly rate of change of nominal wages for the whole economy (w), the monthly rate of change of nominal $M0$ (m) and price inflation measured by the monthly rate of change of the retail price index, $RPI(\pi)$.[7] Figure 1 displays the monthly and annual inflation rates starting in January 1962 (the annual inflation rate in month t, is computed as $\sum_{i=0}^{11} \pi_{t-i}$).

For most of the time span considered, a quasi-flexible exchange rate regime has been working in the UK. This could suggest that nominal money growth and inflation may have a strong domestic component. The decision of the UK to postpone the participation to the Euro area will allow internal factors still to play a key role in determining the short- and medium-run effects of monetary policy. The behaviour of the annual growth rates of nominal $M0$ and nominal wages (computed as the annual inflation rate above) are shown in Figure 2 together with the annual rate of RPI inflation to highlight the long-run comovement of the series. The selection of the growth rate of the RPI as a measure of price inflation allows for the contribution of external factors to explain core inflation dynamics, since the RPI include prices of imported as well as domestic goods. We explicitly introduce in the system the oil price as a shock variable which is expected to be important in particular for the short-run inflation dynamics in the sample. Finally, the industrial production index is employed to include in the model a measure of real activity.

As a first step of the empirical analysis, standard unit-root tests show that all variables can be considered as $I(1)$ processes.[8] The issue of non-stationarity of the variables is further explored below, within the framework of the cointegrated system. The vector of endogenous variables is then specified as $\mathbf{x}_t = (oil_t\ y_t\ w_t\ m_t\ \pi_t)'$. For the oil price, the assumption of endogeneity is functional to the shock analysis undertaken subsequently. The main rationale for the inclusion of oil in the system is to evaluate the response of the other variables to a major source of supply side shocks. This should provide valuable additional information about the determinants of long-run inflation.

Cointegration analysis has been carried out using Johansen's (1988, 1995) Maximum Likelihood approach over the period 1961(2)–1997(12). Twelve lags of each variable have been included in the short-run specification of the model on the basis of diagnostic tests of dynamic specification, showing that a twelve-lag dynamic structure is capable of eliminating all residual serial correlation.

[7] Data are taken from *Datastream* and, with the exception of *oil*, seasonally adjusted using the standard X12-ARIMA methodology.

[8] Various unit-root tests have been used. The results from Augmented Dickey-Fuller (ADF) tests support the non-stationarity of all variables at the 5% significance level. Kwiatkowski-Phillips-Schmidt-Shin (KPSS) tests reject the null hypothesis of *stationarity* at the 5% level for y, oil, w and π, and at the 10% level in the case of m. Consistently with the above results, the non-stationarity of money *velocity* cannot be rejected over our sample (the ADF(6) tests statistic is -0.41 with a 5% critical value of -3.42).

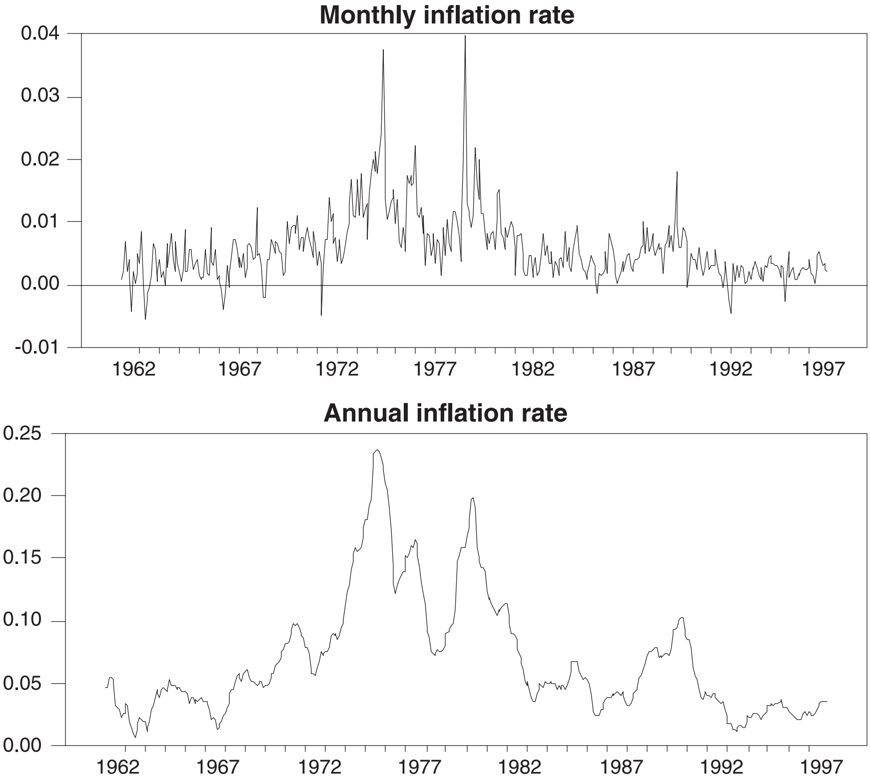

Fig. 1. The UK inflation rate

In Table 1 we report the results of the cointegration analysis. The cointegration tests clearly suggest the existence of two valid cointegrating vectors.[9] The restrictions that only w and π belong to the first vector and only m and π belong to the second are tested and not rejected, obtaining a $\chi^2(4)$ statistic of 4.3, with a corresponding p-value of 0.36. When the more stringent restrictions that the coefficients on the inflation rate are -1 in both vectors are imposed, the obtained p-value is 0.12, again supporting the restrictions. Therefore the two restricted cointegrating vectors capture the stationarity of the rate of growth of real wages and of real money balances, shown in Figure 3, and such restricted vectors are imposed in the following analysis of the common

[9] Further diagnostic tests on the VAR system detected some sign of residual non-normality, mainly concentrated in the equation for the oil price. In principle, deviations from normality could affect the outcome of the Johansen procedure determining the cointegration rank of the system. However, this is not a serious concern in our case for two reasons. First, the detected deviations from normality are due to excess kurtosis and not skewness; as shown by Gonzalo (1994), the Johansen procedure is fairly robust against this kind of non-normality. Second, the results of the Johansen procedure reported above hold also when the main cause of non-normality is eliminated and a four-variable system (excluding the oil price) is estimated. In this case the results strongly support a cointegration rank $r = 2$ (confirming those obtained for the full, five-variable system); in addition, also the restrictions on the cointegrating vectors that we are going to use in the common trends procedure are not rejected.

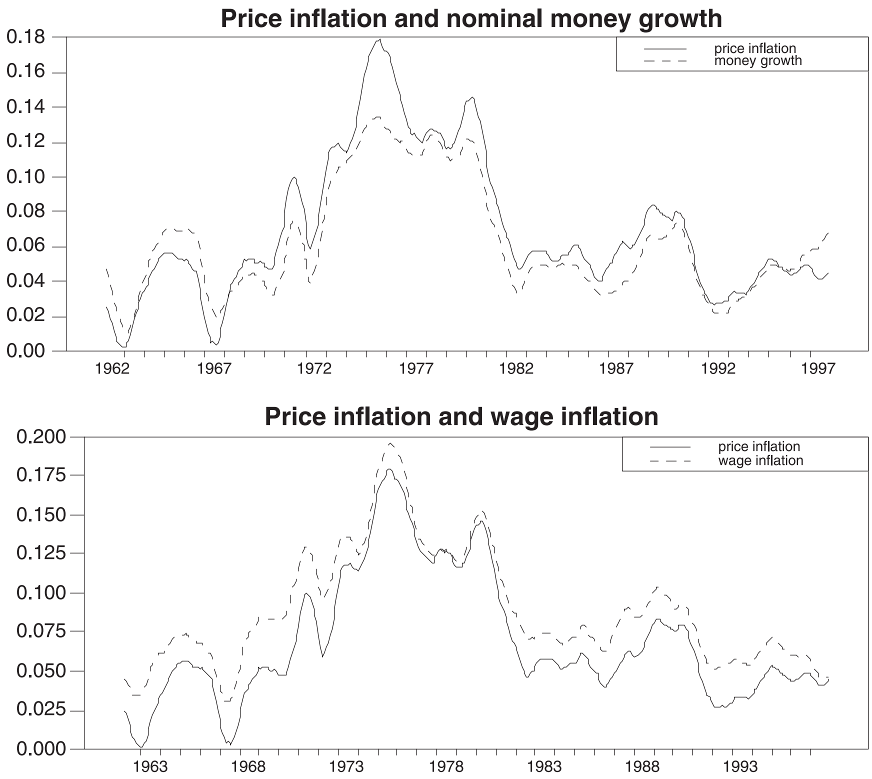

Fig. 2. Nominal money growth, nominal wage growth, and measured *RPI* inflation rate

trends model. Conditional on a cointegration rank $r = 2$, stationarity tests on each variable are carried out to further check the evidence of non-stationarity from unit-root tests mentioned above. The results reported in Table 1 strongly reject stationarity for each variable in the system. The result of non-stationarity of the inflation rate in our sample is not uncommon in the literature. Also Quah and Vahey (1995) found evidence of non-stationarity of the UK observed inflation process, though over a shorter sample period (1969–1994). For the US inflation, tests carried out by Fuhrer and Moore (1995), Stock and Watson (1999) and Ireland (1999) support the result of non-stationarity.[10]

In the common trends model in (3.11), the presence of two cointegrating relationships among the five endogenous variables in $\mathbf{x}_t$ implies that there are three distinct sources of shocks having permanent effects on (at least some of) the elements of $\mathbf{x}_t$. The restrictions in (3.12) and (3.13) yield twelve equations which can be used to obtain the fifteen elements of the long-run impact matrix $\boldsymbol{\Gamma}_g$. Three additional restrictions are needed for (exact) identification of the common trends. To achieve identification, we make the following assumptions

[10] Ireland (1999) provides also a theoretical model yielding a non-stationary inflation rate (due to assumption of a unit root in the natural rate of unemployment, to which the inflation rate is related in the long-run). The construction of a theoretical model aimed at explaining the non-stationarity features of the data is beyond the scope of the present paper.

Table 1. Cointegration analysis

Cointegration tests

Eigenvalue:	0.142	0.074	0.051
Hypothesis:	$r = 0$	$r \leq 1$	$r \leq 2$
λ_{MAX}	58.7**	29.6*	19.8
95% crit. value	33.5	27.1	21.0
λ_{TRACE}	115.0**	56.3**	25.7
95% crit. value	68.5	47.2	29.7

r denotes the number of valid cointegrating vectors;
** and * denote significance at the 1% and 5% level respectively.

Stationarity tests
(tests are conducted on each variable, conditional on a cointegration rank $r = 2$)

	oil	*y*	*w*	*m*	π
$\chi^2(3)$ p-value	32.7 (0.000)	62.3 (0.000)	10.3 (0.016)	14.9 (0.002)	12.4 (0.006)

Restricted cointegrating vectors
($\boldsymbol{\beta}'$ matrix; cointegrating vectors normalized on w and m respectively)

	oil	*y*	*w*	*m*	π
$\boldsymbol{\beta}_1'$	0	0	1	0	−0.778 (0.071)
$\boldsymbol{\beta}_2'$	0	0	0	1	−0.740 (0.118)

Likelihood-ratio test: $\chi^2(4) = 4.33$ [0.36]

	oil	*y*	*w*	*m*	π
$\boldsymbol{\beta}_1'$	0	0	1	0	−1
$\boldsymbol{\beta}_2'$	0	0	0	1	−1

Likelihood-ratio test: $\chi^2(6) = 10.08$ [0.12]

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 (domestic) *nominal* disturbance (ψ_n). The permanent part of the common trend representation in (3.9) is then the following three-variate random walk:

$$\begin{pmatrix} \tau_f \\ \tau_r \\ \tau_n \end{pmatrix}_t = \begin{pmatrix} \mu_f \\ \mu_r \\ \mu_n \end{pmatrix} + \begin{pmatrix} \tau_f \\ \tau_r \\ \tau_n \end{pmatrix}_{t-1} + \begin{pmatrix} \psi_f \\ \psi_r \\ \psi_n \end{pmatrix}_t \tag{3.14}$$

where $\boldsymbol{\mu}$ is a vector of constant drift terms, added to the model in estimation. The three additional restrictions on the elements of $\boldsymbol{\Gamma}_g$ needed for identifica-

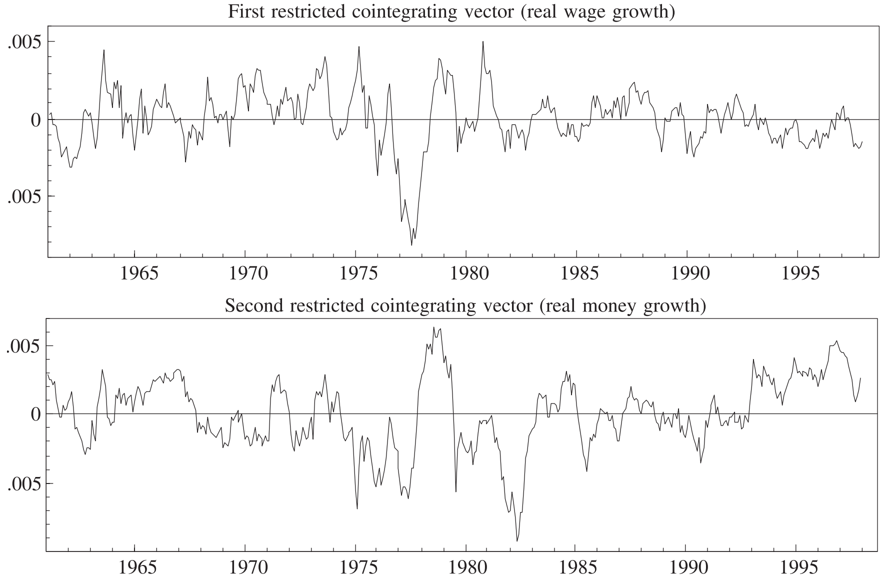

Fig. 3. Restricted cointegrating vectors

tion are consistent with the above assumptions on the economic nature of the common trends. We assume that both the domestic real and 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 $\boldsymbol{\Gamma}_g$, 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:

$$\begin{pmatrix} oil \\ y \\ w \\ m \\ \pi \end{pmatrix}_t = \begin{pmatrix} oil \\ y \\ w \\ m \\ \pi \end{pmatrix}_0 + \begin{pmatrix} \gamma_{11} & 0 & 0 \\ \gamma_{21} & \gamma_{22} & 0 \\ \gamma_{31} & \gamma_{32} & \gamma_{33} \\ \gamma_{41} & \gamma_{42} & \gamma_{43} \\ \gamma_{51} & \gamma_{52} & \gamma_{53} \end{pmatrix} \begin{pmatrix} \tau_f \\ \tau_r \\ \tau_n \end{pmatrix}_t + \boldsymbol{\Gamma}^*(L) \begin{pmatrix} \psi_f \\ \psi_r \\ \psi_n \\ v_1 \\ v_2 \end{pmatrix}_t \tag{3.15}$$

where v_1 and v_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.[11]

Estimation of the common trends model is performed following the methodology set out in Warne (1993) and the main results are shown in Table 2, where the estimated elements of the long-run impact matrix $\boldsymbol{\Gamma}_g$ and the long-run forecast error variance decomposition of the variables are reported (with asymptotic standard errors in parentheses).[12]

[11] As noted by a referee, the general criticisms of Faust and Leeper (1997) to the long-run restrictions of the Blanchard-Quah (and Quah-Vahey) type apply also to the common trends model employed here. However, our use of an enlarged set of variables may help in disentangling various sources of shocks, whereby mitigating at least one of the Faust-Leeper main points.

[12] Estimation has been carried out using the *CT* Rats routine of A.Warne and H. Hansen.

Table 2. The estimated common trends model

Long-run effects of permanent shocks (matrix $\boldsymbol{\Gamma}_g$)
(asymptotic standard errors in parentheses; * denote stat. significance at the 5% level)

Variable	Shock		
	ψ_f	ψ_r	ψ_n
oil	0.0829*	0	0
	(0.0157)	(–)	(–)
y	−0.0037	0.0111*	0
	(0.0031)	(0.0019)	(–)
w	0.0002	0.0003*	0.0006*
	(0.0002)	(0.0001)	(0.0001)
m	0.0002	0.0003*	0.0006*
	(0.0002)	(0.0001)	(0.0001)
π	0.0002	0.0003*	0.0006*
	(0.0002)	(0.0001)	(0.0001)

Long-run (∞) forecast error variance decomposition
(asymptotic standard errors in parentheses)

Variables	Shock		
	ψ_f	ψ_r	ψ_n
oil	1	0	0
	(–)	(–)	(–)
y	0.098	0.902	0
	(0.144)	(0.142)	(–)
w	0.129	0.198	0.673
	(0.149)	(0.149)	(0.150)
m	0.129	0.198	0.673
	(0.149)	(0.149)	(0.150)
π	0.129	0.198	0.673
	(0.149)	(0.149)	(0.150)

A number of features can be noticed. First, the foreign shock does not play a relevant role in explaining the long-run features of either nominal or real variables. It has a permanent positive effect on the nominal variables and a negative effect on output, but they are not statistically significant. Second, the real internal shock has a statistically significant long-run effect on the nominal variables as well as on output. In particular, the shock which permanently increases industrial production tends to permanently increase price and wage inflation and the nominal money growth rate.[13] Finally, the nominal disturbance significantly affects wage and price inflation and the nominal money growth rate in the long-run. The forecast error variance decomposition results give some ad-

[13] The identification assumptions in (3.15) allow for a broad interpretation of the permanent shock ψ_r, which captures all underlying disturbances (including technology shocks) with permanent output effects not related to movements in the oil price. Such disturbances determine long-run increases in production accompanied by upward movements also in nominal wage and money growth (possibly due to wage pressures and accommodating monetary responses), which are positively linked to the inflation rate in the long run.

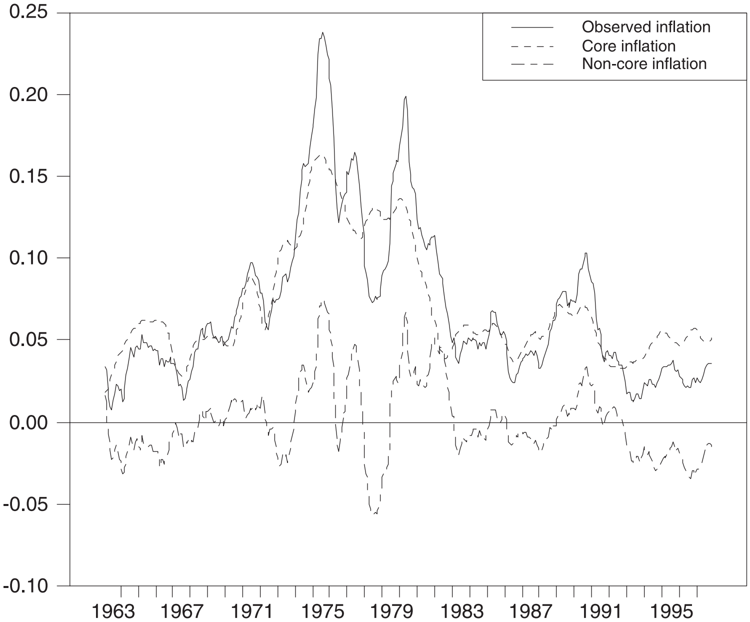

Fig. 4. Observed inflation rate and core and non-core estimated inflation rates from the common trends model

ditional insight into the long-run analysis, though in many cases the relatively large asymptotic standard errors do not allow sharp inferences. In particular, the internal real shock, which explains almost entirely the variability of output (90%), accounts for about 20% of the variability in the nominal variables. The foreign shock explains only about 10% of the long-run variability of output. Finally, the nominal shock accounts for around 67% of the variability in the nominal variables.

The estimated monthly core inflation series from the common trends models is then computed as $\hat{\pi}_t^c = \pi_0 + \hat{\gamma}_{51}\hat{\tau}_{f,t} + \hat{\gamma}_{52}\hat{\tau}_{r,t} + \hat{\gamma}_{53}\hat{\tau}_{n,t}$. Figure 4 displays the estimated annual core inflation rate (expressed, for each month t, as $\sum_{i=0}^{11} \hat{\pi}_{t-i}^c$), together with the measured annual inflation rate. The difference between the observed inflation rate and the estimated core rate, named "non-core inflation" is also added to the picture, capturing the transitory component of the inflation process.

As expected, core inflation is lower than observed inflation during the negative "oil shock" episodes of the mid-'70s and early '80s, whereas the non-core inflation component closely follows the observed rate. During the counter-shock of the mid-'80s the relation between the actual and the core inflation rates is consistently reversed, with the latter above the former. Finally, according to our specification of the common trends model, the estimated core inflation rate lies above the measured inflation rate over the final part of the sample, from 1993 to 1997. Relatively large positive real shocks in 1993–1994 and positive nominal shocks concentrated in 1994 are mainly responsible for keeping the core inflation rate above the actual rate throughout the 1993–1997 period.

Overall, two general features of the estimated core inflation series can be noted. First, core inflation is smoother than observed inflation: the standard deviation of the monthly changes in the core inflation rate is only 0.07%, compared with 0.38% of changes in the actual inflation rate (changes in the non-core inflation component have a standard deviation of 0.37%, very close to actual inflation). Second, monthly changes in observed inflation are much more correlated with changes in the non-core component (the correlation coefficient is 0.98) than with changes in the estimated core rate (0.13).[14]

From the above results, some final observations can be made. First, the estimated core inflation, interpreted as the long-run inflation forecast, is not entirely determined by output-neutral (nominal) disturbances (ψ_n), given the statistical significance of the long-run impact of the domestic real shock (ψ_r). However, the role of the latter disturbance should not be overstated since its contribution to the long-run variability of inflation (20%) is not precisely estimated. Second, given the the relatively long time needed for monetary policy actions to affect real activity and inflation, a measure of the long-run inflation forecast may be useful for the design of forward-looking monetary policy rules, although the optimality of (purely) forecast-based policy rules is an open issue.[15]

Finally, we also implemented on our sample the identification scheme of Quah and Vahey (1995), using a bivariate *VAR* system with inflation and industrial production. As already mentioned, core inflation is defined here as that component of the observed inflation rate that does not affect output in the long-run. Core-inflationary shocks are then identified by imposing a zero restriction on their long-run output effect and the core inflation series is constructed using only this kind of disturbances. This scheme yields, as an additional identifying condition, "the property that non-core inflationary shocks have no permanent impact on the measured rate of inflation. This allows a further assessment of the validity of [the Quah-Vahey] approach. If the data do not support the hypothesis that non-core disturbances have little sustained impact on measured inflation then our identification procedure is dubious." (Quah and Vahey, 1995, p. 1135). When the additional restriction of no long-run effect of non-core inflationary shocks on the observed inflation rate is imposed on the data, a formal test yields a strong rejection, casting serious doubts on the appropriateness of the overall Quah-Vahey identification scheme for our sample.[16]

4. Conclusions

A common trends model has been used to estimate the underlying, "core" inflation behaviour for the UK over the 1961–1997 period. Estimating a common trends model allows for the decomposition of observed inflation into an

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[14] The core and non-core components of inflation display a correlation close to zero (−0.05).

[15] Forecasts of inflation over a medium-term horizon play an important role as an intermediate policy target in the "inflation-forecast targeting" rules described by Svensson (1999). Levin, Wieland and Williams (1999) do not advocate the formulation of monetary policy rules in terms of inflation forecast on the basis of simulations showing that, for the US economy, such forecast-based rules hardly improve the inflation-output variability trade-off with respect to rules formulated in terms of observed variables. Woodford (2000) summarizes the theoretical arguments against purely forward-looking policy rules.

[16] The likelihood test for the over-identifying restriction imposed on the *VAR* delivers a $\chi^2(1)$ value of 63.7, with an associated p-value smaller than 0.001.

underlying trend (core) and a purely transitory component. Extending the bivariate *VAR* system analyzed by Quah and Vahey (1995), our model includes, besides inflation and output, also three main determinants of inflationary pressures, namely money growth, wage growth and oil price movements. In this framework core inflation is interpreted as the long-run forecast of the inflation rate consistent with the long-run (cointegrating) relationships liking money growth, wage growth and inflation.

Although the core inflation behaviour obtained from the common trends model does depend on the specification of the system (in terms of variables included, sample period, dynamic specification and other modelling choices), we suggest that an interpretation of core inflation in terms of long-run inflation forecast from a small-scale multivariate system of the kind estimated here may yield valuable information on the relative importance of various sources of inflationary disturbances.

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