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FORECASTING INFLATION IN THE EURO AREA USING MONTHLY TIME SERIES MODELS AND QUARTERLY ECONOMETRIC MODELS

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Abstract

Economic agents and financial authorities require frequent updates to a path of accurate inflation forecasts and need forecasts to include an explanation of the factors by which they are determined. This paper studies how to approach this need, developing a method for analysing inflation in the euro area, measured according to HICP. Time series models using the most recent information on prices and an important functional and geographically disaggregation can provide monthly forecasts which are reasonably accurate, but they do not provide an explanation of the factors by which the forecast is determined. In this respect, it is important to enlarge the data set used considering explanatory variables and build congruent econometric models including variables which, following previous works by D. Hendry, capture disequilibria on different markets, goods and services, labour, monetary and international. The final result of this work shows that combining the forecasts from a monthly time series vector model, constructed on price subindexes from a disaggregation of the HICP by countries and sectors, with the forecasts derived from a quarterly econometric vector model on aggregate inflation and other economic variables, very accurate forecasts are obtained. Both vector models are specified including empirical cointegration restrictions, which in the first case capture the constrains necessary present between the trends of the price subindexes and in the second approximate the long-run restrictions postulated by economic theory.

Keywords: block-diagonal disaggregate monthly vector model, cointegration restrictions, congruent models, forecast combination, thick modelling, correction of outliers.

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1. INTRODUCTION

In a previous paper, Espasa and Albacete (2004), we study inflation in the euro area by disaggregating the HICP in five regions, one for each of the four largest countries and a fifth one grouping all remaining countries, and two sectors, core and residual; the core sector, following Espasa et al. (1987), includes prices for processed food, non-energy industrial goods and services and the residual sector, the prices for energy and unprocessed food. We provide evidence that forecasting these ten components and then aggregating the results one gets better forecasts for the year-on-year rate of growth of HICP than forecasting directly the aggregate. This result holds for all horizons considered in that paper, one to twelve months.

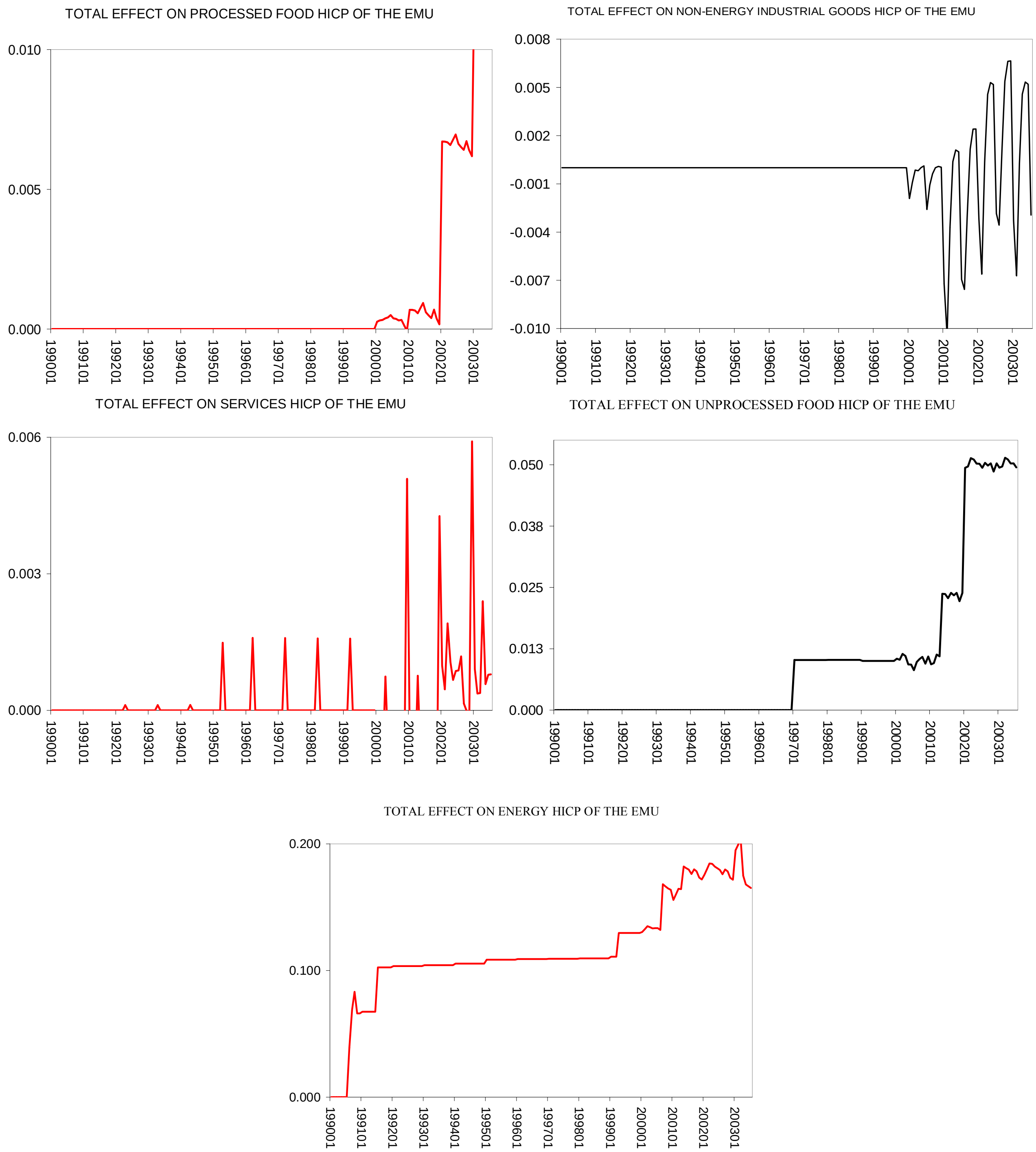
This evidence contrasts with the results of Hubrich (2003), Den Reijer & Vlaar (2003) and Benalal et al. (2004) where the authors found that disaggregating does not improve the forecast of the aggregate or just get a minor improvement for only short horizons.

Following Hendry´s terminology it can be said that considering disaggregated information one does not reduce the predictability of the aggregate. Questions related to the building of the corresponding models and the way the analysts tackle them will finally lead or not to get better forecast accuracy by disaggregating.

Two aspects are important in this respect when study the inflation in the euro area and both refer to the approach taken in building the model for components. One corresponds to the existence of cointegrated relationships between the components. Due to the fact that the markets corresponding to the mentioned sectors have been and are in an accelerated process of integration and they have been affected by a convergence process which ended up in 1999 with a monetary union, the trends of these price subindexes must share some long-run restrictions coming from the integration attained between them and the common monetary policy. Therefore, we expect that between these prices there will be some cointegration relationships. But the cointegration will not be full in the sense that in the vector of *n* price sub-indexes there would not be *(n-1)* cointegration relationships, implying that all prices are not driven by a unique common factor trend. We expect some cointegration restrictions and also several common factor trends. In this situation when modelling the vector of components it would matter to consider a simultaneous model with cointegration restrictions. The inclusion of them is one of the differences of the work in Espasa and Albacete (2004) with respect to the other mentioned papers.

A second aspect refers to the fact that there exists special events affecting inflation data in the euro area like: the introduction of sales prices – by different countries at different moments in the sample period - when calculating the HICP; the introduction of the euro; tax changes in the case of administered prices such as tobacco, gas or electricity prices; abrupt changes in international oil prices; adverse weather conditions and epidemics affecting the prices of non-processed food, etc. In these circumstances if one does not properly take into account these special events in the modelling process, this is going to affect more negatively the results of the disaggregate model than those of the aggregate one, generating a bias in favour of the latter. In this respect Espasa and Albacete (2004) give precise details of how they deal with all the special events mentioned above. Since most of them occur in the post-sample period used for forecasting evaluation, in that paper these effects, which are causing structural breaks, are taken as known in all models in the forecast evaluation. Figure 1, taken from Albacete (2004) shows these effects on five sectors which compose the total HICP in the EMU. Other recent works consider seasonally adjusted data and do not handle these relevant special events.

Figure 1: *Effects derived from special events affecting the five sectorial sub-indexes that compose the total HICP in the euro area.*



Inflation forecasts are useful if they are accurate and include information about the factors determining them. In our previous work we have shown that a substantial accuracy level can be obtained in path forecasts from one to twelve months. But a monthly time series vector model for the components does not include causal-based explanatory variables. In this paper we propose a causal-based model for aggregate inflation and study its forecast accuracy and analyse the convenience of combining forecasts from this econometric model with those from the disaggregated-vector model.

In the reminder of this paper we proceed as follows. In section 2 we present the context in which the econometric model is formulated and we build a quarterly model for the aggregate inflation in the euro area. Section 3 compares the forecasting performance at a quarterly level of the disaggregated-vector model and the aggregate-econometric one and studies a strategy to use both models to give the most accurate forecasts with an explanation of the factors determining them. Finally, section 4 concludes.

2. AN AGGREGATE-ECONOMETRIC MODEL FOR INFLATION IN THE EURO AREA.

Monthly forecasts derived from time series models, including the most recent information on prices and an important functional and geographical disaggregation, can be reasonably accurate, see for instance Espasa and Albacete (2004), but they do not provide an explanation of the factors by which the forecast is determined. In this respect, it is important to advance in the data set used and consider explanatory variables showing a causal relationship with inflation, based on economic theory, and then build congruent econometric models, based on this economic theory and according to the data available. We formulate models on a quarterly basis, since factors determining inflation such as unit labour costs are only observed quarterly.

To specify the factors determining inflation in the euro area we follow Hendry (2001) and consider variables which capture disequilibria on different markets, goods and services, labour, monetary and international, thus contemplating the most relevant theories when analysing inflation, such as the mark up models.

In order to select the explanatory variables is also important to consider the results derived from Stock & Watson (1999), for the case of inflation in USA, and Banerjee et al. (2003), for the case of inflation in the euro area, which conclude that real variables show a greater predictive capacity than monetary ones.

The following subsection describes the explanatory variables included in the congruent model built for the aggregate inflation in the euro area.

2.1 DESCRIPTION OF THE EXPLICATIVE VARIABLES

Following Hendry (2001), the disequilibrium between supply and demand on the goods and services market is represented by the output gap, defined as the ratio between actual and potential output. The model also includes the monetary aggregate M3 and the import deflactor to relate inflation whit monetary and international markets. Another important variable in our model is the unit labour cost, which is a factor determining the long run behaviour of prices according with the mark up theory and has been considered in papers as De Brouwer & Ericsson (1998), Banerjee et al. (2001), Galí et al. (2001), Bowdler & Jansen (2004), etc. Finally, the disequilibrium between supply and demand on the labour market can be captured by the rate of unemployment, but this variable is not available for the euro area from the beginning of the sample period employed in this paper. The above mentioned excess demand variable acts also as a proxy for the unemployment rate. In fact, different authors use a proxy of this type when specifying a Phillips curve model for inflation. Besides, the disequilibrium in the labour market could also be held on changes in unit labour costs, which is included in our model. In any case we also introduce the rate of growth of unemployment, which turns out to be exogenous as also occurs in Hendry (2001).

The sample considered goes from the first quarter of 1991 to the second quarter of 2003. Series prior to 1991 do not reflect the ESA95 conventions. Besides, series for the unified Germany are available only since 1991. Table 1 shows the variables used to construct a measure of output gap and all the other explanatory variables.

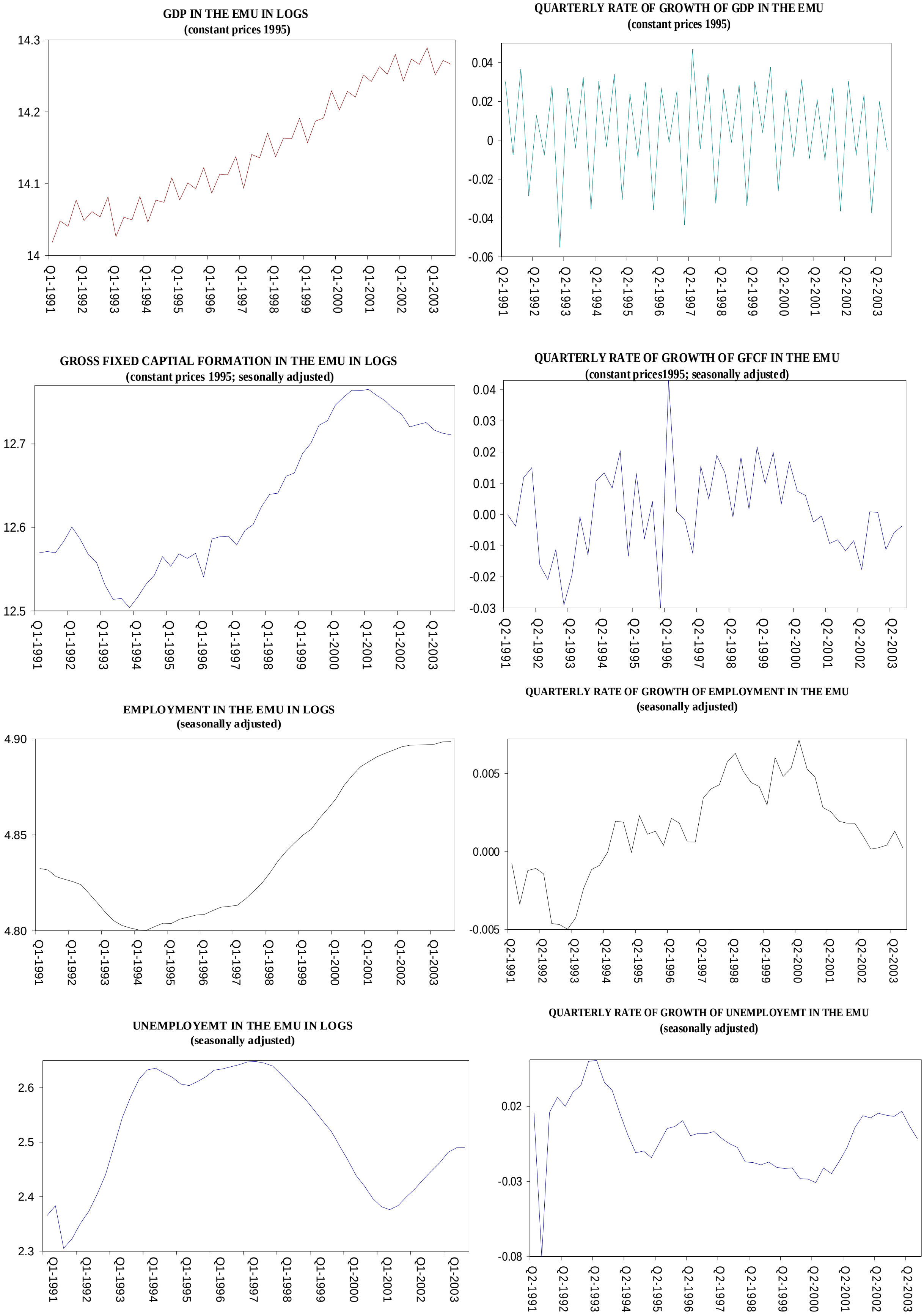
Table 1: *Variables used in the congruent quarterly models*

Dependent Variable	**Description**	**Unit**	**Source**	**Seasonally Adjusted**
HICP	Harmonised Index of Consumer Price	Index, 1996=100	Eurostat	NO
Explicative Variables	**Description**	**Unit**	**Source**	**Seasonally Adjusted**
GDP	Gross Domestic Product, at market price, constant prices 1995	Millions of euros (ecu before 1999)	Eurostat	NO
GFCF	Gross fixed Capital Formation constant prices 1995	Millions of euros (ecu before 1999)	Eurostat	YES
L	Employment	Millions of persons	ECB	YES
U	Unemployment	Millions of persons	Eurostat	YES
ULC	Unit Labour Costs	Index, 2000=100	ECB	YES
M3	Monetary Aggregate M3, stocks end of quarter	Billions of euros	OECD 1991(1)-1997(2) Eurostat 1997(3)-2003(2)	YES (also adjusted by working days)
DM	Import Deflactor	Index, 1995=100	Eurostat	NO

Graphs of these variables in logarithms and their quarter-on-quarter rates of growth can be found in figures 2 and 3.

Figure 2: *HICP, DM, ULC and M3 in logarithms and the quarterly rates of growth.*

Figure 3: *GDP, GFCF, L and U in logarithms and the quarterly rates of growth.*



Following Hendry (2001), the output gap measure has been constructed using a Cobb-Douglas production function with constant returns of scale. In this function the shares of labour and capital inputs are equal to the elasticities of output with respect to these inputs. The values of these elasticities are approximately 0.6 and 0.4 under the traditional assumption of the literature, Barro & Sala-i-Martin (1995). The capital stock is determined in a recursive way with an assumed depreciation rate of 6% per annum, Barro & Sala-i-Martin (1995). Technological progress is modelled as a linear time trend, according to Jones (1995a, 1995b).

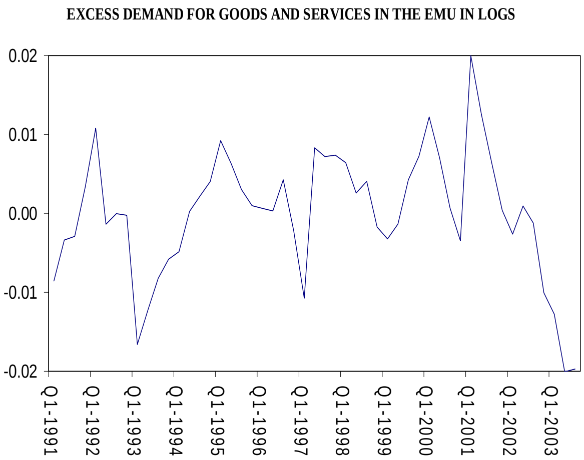
Therefore, the production function shows the following expression:

$PY_t = A_t L_t^{\alpha} K_t^{\beta}$, where PY is the potential production, A is the technological progress, L is the labour input, K is the capital input, α is the elasticity of output with respect to the labour input and β is the elasticity of output with respect to the capital input. Under the assumption of constant returns of scale it is verified that $\alpha + \beta = 1$. With all of this we get $PY_t = A_t L_t (K_t/L_t)^{\beta}$ and taking logarithms we obtain $py_t = a_t + l_t + \beta(k_t-l_t)$, where lower-case letters denote the original variables in logarithms. The impact of technological progress is estimated by regressing ($y_t - l_t - \beta(k_t-l_t)$) on a constant, linear trend and centered seasonal dummies, where y_t is the actual GDP in logarithms.

Consequently, $\hat{a}_t = \hat{\mu} + \hat{\lambda} t + \hat{\gamma}_1 s_{1t} + \hat{\gamma}_2 s_{2t} + \hat{\gamma}_3 s_{3t}$; where s_{it} is a dummy, which takes the value 1 in the i-quarter, minus 1 in the fourth quarter and 0 in the remaining quarters. Therefore, the excess demand for goods and services is determined by $y_t^d = y_t - l_t - \beta(k_t - l_t) - \hat{a}_t$. Following Hendry (2001), the variable $\beta(k_t-l_t) + \hat{a}_t$ represents a measure of potential capacity cap_t. Consequently, $y_t^d = y_t - l_t - cap_t$.

Figure 4 shows the graphs of the excess demand for goods and services (ED) in logarithms and the corresponding quarter-on-quarter rate of growth.

Figure 4: *Excess demand for goods and services in logarithms and its quarter-on-quarter rate of growth in the euro area.*



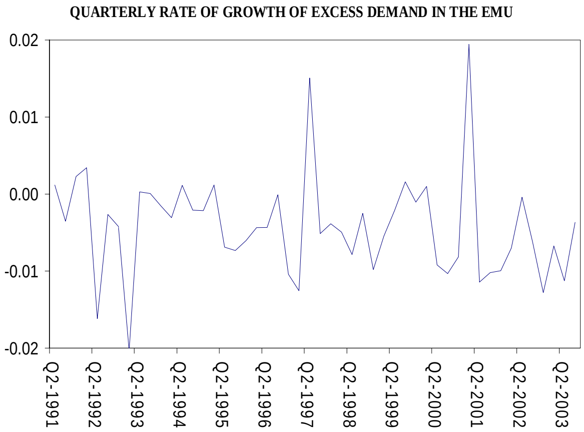


Table 2 lists the augmented Dickey-Fuller (1981) statistics for the variables in logs and for their first differences.

Table 2: *ADF statistics for testing for a unit root in the HICP and the explicative variables in the congruent model.*

Null Hypothesis	HICP	ULC	U	M3	DM	ED
I(1)	-2.53 (-0.01)	-0.20 (0.00)	-1.66 (-0.05)	1.57 (0.01)	-1.27 (-0.05)	-2.67 (-0.33)
I(2)	-5.82** (-0.93)	-5.34** (-0.71)	-1.39 (-0.08)	-5.70** (-0.84)	-4.56** (-0.68)	-7.40** (-1.12)

Notes:

(1) Here and elsewhere in this paper, asterisks * and ** denote rejection at the 5% and 1% critical values. The critical values for this table are calculated from MacKinnon (1991) and are -2.93 at the 5% and -3.58 at the 1%.

(2) The alternative used includes a constant and centered seasonal dummies for HICP and DM. Like the remaining variables are seasonally adjusted only a constant is included in the regression.

(3) Series are taken in logs.

(4) Values reported in parentheses are the estimated coefficient on the lagged variable x_{t-1}.

Empirically, all variables appear to be integrated of order 1 - I(1)- with the hypothesis of a second unit root being rejected, excluding unemployment level that appears to be integrated of order 2, I(2). For this reason we are going to work with the rate of growth of unemployment level, defined as the first difference of the logarithm, obtaining then a vector composed by I(1) variables. The result that yields unemployment level as I(2) variable could be related to the fact of considering the number of unemployed instead of the unemployment rate. It is also possible that the presence of certain outliers under the linearity hypothesis, which is the traditional assumption for the integration and cointegration tests, determines the rejection of the I(1) hypothesis in favour of I(2) in the case of unemployment level. Nevertheless, this is not a relevant inconvenience because this variable results to be exogenous, as it is shown below.

The following section describes the estimation of a vector equilibrium correction model for the vector composed by these six variables: harmonised index of consumer prices, unit labour costs, unemployment rate of growth, monetary aggregate M3, import deflactor and excess demand.

2.2 ESTIMATION OF A CONGRUENT ECONOMETRIC MODEL FOR INFLATION IN THE EURO AREA

The sample available goes from the first quarter of 1991 to the second quarter of 2003, but there was some sort of instability at the beginning of the sample period and recursive estimations have been applied. Stability appears after the third quarter of 1993 and this is the initial date that has been considered for the estimation of the congruent econometric model. For forecasting purposes the model is re-estimated considering the fourth quarter of 1999 as the last observation.

The greatest eigenvalue and trace eigenvalue statistics, corresponding to the Johansens' (1988, 1991) cointegration analysis, reject the null of two cointegration relationships in favour of at least three cointegrating relationships. This last hypothesis is not rejected in favour of a hypothesis with more than three cointegration restrictions. So, this analysis indicates that there exist three long-run relationships between the aggregate consumer price index and the other economic explicative variables. In the estimation of the cointegration relationships a dummy corresponding to the first quarter of 2001, q1, has been included. This variable picks up the effect of the incorporation of Greece to the EMU. The estimated restricted cointegration relationships can be expressed as:

$$1)\ \ln(HICP) - 1.31\ln(ULC) + 0.03\Delta\ln(U) - \underset{(0.008)}{0.14\ln(M3)} + \underset{(0.031)}{0.06\ln(DM)} - \underset{(0.087)}{1.31n(ED)} + \underset{(0.002)}{0.04E01q1}$$

$$2)\ -5.25\ln(HICP) + \ln(ULC) - 1.28\Delta\ln(U) + \underset{(0.110)}{1.38\ln(M3)} + \underset{(0.416)}{0.71\ln(DM)} + \underset{(1.180)}{2.92\ln(ED)} + \underset{(0.033)}{0.04E01q1}$$

$$3)\ 0.78\ln(HICP) - 1.34\ln(ULC) + \Delta\ln(U) + \underset{(0.032)}{0.15\ln(M3)} - \underset{(0.123)}{0.12\ln(DM)} + \underset{(0.349)}{0.17\ln(ED)} - \underset{(0.010)}{0.01E01q1}.$$

where standard errors of estimates are in brackets. Imposing more restrictions, which are not rejected by the data we get:

$$1)\ \ln(HICP) - 1.31\ln(ULC) + 0.03\Delta\ln(U) - \underset{(0.008)}{0.14\ln(M3)} - 1.31og(ED) + \underset{(0.002)}{0.04E01q1}$$

$$2)\ -5.25\ln(HICP) + \ln(ULC) - 1.28\Delta\ln(U) + 1.28\ln(M3) + 5.25\ln(ED)$$

$$3)\ \ln(HICP) - \ln(ULC) + \Delta\ln(U).$$

$$(0.091)$$

The coefficients of the variables show the expected sign. So, in the three cointegration restrictions it is observed the following relations: a) the coefficient of unemployment rate of growth presents the same sign that the corresponding to the consumer prices according to the negative relation established by the Phillips curve; b) the coefficient of unit labour costs shows the contrary sign that the coefficient of consumer prices in line with mark-up models. In the two first long-run restrictions, the coefficients corresponding to the monetary aggregate and excess demand also present a contrary sign that the coefficient of consumer prices. These results also appear in other works, such as Hendry (2001) and Dreger (2002). With all of this, it is obtained that consumer prices depend positively on unit labour cost, monetary mass and excess demand, and negatively on unemployment rate.

The exogeneity tests show that only the unemployment rate of growth could be considered as exogenous variable.

The final estimated vector equilibrium correction model (VEqCM) is shown in table 3. CI_{1t}, CI_{2t} and CI_{3t} represent the previous restricted cointegration relationships. In the equations corresponding to the HICP and imports deflactor, centered seasonal dummies are included. In the HICP equation another set of dummies variables is also included in order to pick up the effect derived from the introduction of sales prices in the calculation of the HICP(1). Intervention analysis for the remaining endogenous variables is also elaborated with the objective of capturing only the most relevant outliers affecting each serie. Appropriate dummies are included in the corresponding equations through shift variables with the following dates: a) first quarter of 2003 in the HICP serie; b) second quarter of 1997 in ULC equation; c) third quarter of 1997 and first quarter of 2001 in M3 serie; d) second quarter of 1999, first and fourth quarters of 2000 and 2001 and first quarter of 2003 in the equation of the imports deflactor; e) first quarter of 1993, second quarter of 1997 and first quarter of 2001 in excess demand serie.

(1) These variables are constructed by aggregating the sales effects on the HICP of Belgium (since 2000) and Italy and Spain (since 2001) calculated in each case for the five sectors mentioned in figure 1. For more details see Espasa and Albacete (2004b).

Table 3: *Congruent VEqCM Model*

$$
\begin{pmatrix}
1+0.25L^2-0.33L^3 & -0.11L^3 & 0 & 0 & -0.06L^2 & 0.05L^2 \\
0 & 1 & 0 & -0.16L^2 & -0.07L^2 & 0 \\
0.30L^3 & -0.45L^2 & 1-0.26L^2 & 0 & 0 & -0.57L^2 \\
0.77L & -0.44L-0.73L^2-0.79L^3 & 0.14L-0.21L^2 & 1+0.26L^3 & -0.26L & -0.77L-0.57L^2-0.41L^3 \\
-1.18L+1.65L^2 & -0.22L^2-0.82L^3 & 0.11L+0.49L^2 & 0.30L^2+0.16L^3 & 1-0.45L^3 & 0.47L^2-0.47L^3 \\
-0.36L & 0.20L^2 & -0.11L^3 & 0 & 0 & 1+0.22L^2
\end{pmatrix}
\begin{pmatrix}
\Delta ln(HICP)t \\ \Delta ln(ULC)t \\ \Delta(\Delta ln(U))t \\ \Delta ln(M3)t \\ \Delta ln(DM)t \\ \Delta ln(ED)t
\end{pmatrix} =
$$

$$
\begin{pmatrix} 0.0052 \\ 0.0036 \\ -0.0007 \\ 0.0133 \\ 0.0025 \\ -0.0004 \end{pmatrix} +
\begin{pmatrix}
0.00 & 0.02 & 0.05 \\
0.20 & 0.00 & 0.00 \\
0.00 & 0.00 & 0.00 \\
0.54 & -0.06 & 0.00 \\
-1.27 & 0.00 & 0.00 \\
0.00 & -0.02 & -0.16
\end{pmatrix}
\begin{pmatrix} CI1t\text{-}1+2.23 \\ CI2t\text{-}1+8.96 \\ CI3t\text{-}1\text{-}0.04 \end{pmatrix} +
\begin{matrix} \Delta \textit{centered} \\ \textit{seasonal} \\ \textit{dummies} \\ \textit{for } \Delta lhicp \textit{ y } \Delta ldm \end{matrix} +
\begin{matrix} \textit{Dummies for sales} \\ \textit{in } \Delta lhicp \end{matrix} +
\begin{matrix} \Delta other \\ dummies \end{matrix} +
\begin{pmatrix} a1t \\ a2t \\ a3t \\ a4t \\ a5t \\ a6t \end{pmatrix}
$$

Vector normality $\chi^2(12)=$ 10.395 [0.5814].

This model picks up a considerable amount of transitory dynamics besides the long-run restrictions. Inflation in the short run is determined by lagged inflation and changes in unit labour costs, import prices and excess demand. The cointegration restrictions do not enter in the equation corresponding to the unemployment rate. This result coincides with the one derived from exogeneity tests. The first cointegration relationship do not affect to inflation equation. The residuals do not reject the normality hypothesis. The cross correlograms shown that only the contemporaneous correlations are significant. These correlations appear in table 4.

Table 4: *Residual contemporaneous correlations derived from congruent VEqCM*

	Δln(HICP)	Δln(ULC)	Δ(Δln(U))	Δln(M3)	Δln(DM)	Δln(ED)
Δln(HICP)	1					
Δln(ULC)	0.04	1				
Δ(Δln(U))	-0.04	0.76	1			
Δln(M3)	-0.33	0.38	0.35	1		
Δln(DM)	0.74	-0.20	-0.17	-0.25	1	
Δln(ED)	-0.01	-0.67	-0.46	0.05	0.53	1

As shown table 4, the highest residual contemporaneous correlations hold between unit labour costs and unemployment rate of growth, both variables representatives of labour market, and between consumer and import prices. Table 5 shows residual standard deviations derived from the congruent VEqCM model.

Table 5: *Residual standard deviations derived from the congruent VEqCM **(%)***	
$\Delta\ln(HICP)$	0.14
$\Delta\ln(UCL)$	0. 41
$\Delta(\Delta\ln(U))$	0. 72
$\Delta\ln(M3)$	0. 36
$\Delta\ln(DM)$	0. 66
$\Delta\ln(ED)$	0. 39

The residual standard deviation for inflation equation derived from the quarterly econometric vector model, 0.14%, is very similar to the one obtained through the monthly time series model in Espasa and Albacete (2004), 0.12%.

Instead of estimating the cointegration relationships in the construction of the econometric model it is also possible, Hendry (2001), to impose long-term restrictions as established by economic theory, such as the quantitative theory of money from which monetary deviations from the nominal GDP are obtained, and the mark-up theory, according to which prices are determined in the long term as the margin over unit labour costs. We have constructed a model of this kind but it does not improve the sample fit and obtains worse forecasts than the previous vector model which estimates the cointegration restrictions between prices and other economic variables.

2.3 ANALYSIS OF FACTORS DETERMINIG INFLATION IN THE EURO AREA

The above congruent model leads to an analysis of inflation according to its determining factors, of which we can distinguish four classes: (1) transient dynamic factors including lagged inflation, variations in unit labour costs, in the excess demand and in import prices; (2) long-term disequilibria, consisting of empirical cointegration relations between aggregate prices and other economic variables; (3) factors including the effect of dummy variables capturing deterministic seasonality and sales prices in HICP construction since the start of the year 2000; and finally 4) a residual factor.

Table 6: *Contributions[*] to the mean quaterly inflation rate in the euro area derived from the congruent VEqCM*										
Period	**Mean quaterly inflation rate (%)**	**Historic mean quaterly inflation rate 1992Q3-2003Q2 (%)**	**Lagged inflation**	**Changes in unit labour costs**	**Changes in import prices**	**Changes in excess demand**	**Second cointegration relationship f(ULC, M3, ED and unemployment rate)**	**Third cointegration relationship f(ULC and unemployment rate)**	**Dummy variables (deterministic seasonality, sales and E03q1)**	**Residual Factor**
1993	0.79	0.52	0.03	0.09	0.00	0.01	0.16	-0.03	-0.01	0.02
1994	0.62	0.52	0.02	-0.02	0.01	-0.02	0.17	-0.09	-0.01	0.02
1995	0.59	0.52	0.00	-0.03	0.02	-0.01	0.13	-0.09	-0.01	0.06
1996	0.49	0.52	0.00	0.01	0.00	0.01	-0.02	-0.01	-0.01	-0.02
1997	0.38	0.52	0.02	0.00	-0.01	-0.01	-0.09	0.00	-0.01	-0.04
1998	0.20	0.52	-0.04	-0.08	-0.03	0.00	-0.04	0.00	-0.01	-0.12
1999	0.38	0.52	-0.04	-0.01	-0.03	0.00	-0.05	-0.04	-0.01	0.02
2000	0.61	0.52	-0.01	-0.03	0.10	-0.01	0.02	-0.02	-0.03	0.05
2001	0.52	0.52	-0.06	0.01	0.03	-0.01	-0.03	0.06	-0.02	0.01
2002	0.57	0.52	0.03	0.04	-0.05	0.01	-0.19	0.16	-0.01	0.05
2003 (Q1y Q2)	0.65	0.52	0.05	-0.01	-0.05	0.03	-0.32	0.18	0.32	-0.06
* values in percentage points										

From this classification of factors, we can calculate their effect on inflation at any given time, now or in the future, and interpret the monetary policy followed or obtain a possible pattern for its future implementation. Table 6 shows that, given the behaviour of other variables appearing in the second cointegration relationship, from 1993 to 1995 the monetary policies applied by the different central banks ended pushing inflation up in the euro area, whereas the policies applied in the following years had constraining effects. Of special interest is the 2002-2003 period of low interest rate policy, which having a pushing up effect on inflation, it turned to be insufficient to compensate for the reducing effect of the lack of demand. This suggests that a loosen monetary policy could had been applied.

3. FORECASTING PERFORMANCE OF THE AGGREGATE-ECONOMETRIC MODEL

Espasa & Albacete (2004), comparing the forecasting performance of the different monthly time series models, show that a good monthly strategy for forecasting inflation in the euro area consists of disaggregating the total HICP into ten components, two sectors – core and residual – and five geographic areas, constructing a VEqCM model including a diagonal block constraint between the equations corresponding to core and residual price indexes, and including Brent crude oil prices as a leading indicator only for horizons 1 and 2. One result of interest in that work consists of obtaining empirical evidence that international crude oil prices do not increase forecasting accuracy – which they actually diminish – in horizons of over two months, both when the indicator is forecast using univariate models or by means of the prices on current future markets.

Table 7 compares the forecasting performance at a quarterly level of the disaggregated-vector model proposed in Espasa and Albacete (2004) and the aggregate-econometric model presented in the previous section.

Table 7: *Root Mean Squared Forecast Error for the year-on-year inflation rate in quarterly forecasts for the euro area. Forecast Period 2000(I)-2003(II).*						
Quarters ahead	***Quarterly econometric vector model***	***Monthly block-diagonal vector model***			***Forecasts Combination***	
		First three months unknown (a)	*First month known (a)*	*Two first months known (a)*	*First three months unknown (a)*	*Two first months known (a)*
1	0.12	0.12	0.06	**0.02**	0.09	0.06
2	0.18	0.18	0.17	0.13	0.14	**0.11**
3	0.21	0.21	0.20	0.19	0.17	**0.15**
4	0.28	0.24	0.21	**0.18**	0.22	0.20

(a) Numbers of months of the first quarter of the forecasting path for which the inflation rate has been observed.

Numbers in bold type correspond to the least value.

Table 7 shows that the gains in forecasting quarterly inflation rates using observations corresponding to some of the months in the first quarter appearing on the forecasting path are considerable. Therefore, the highest available level of disaggregation over time is always preferable. In this type of forecast the largest possible amount of recent information on the variable of interest seems to be more important than less recent data concerning explanatory variables, because immediate endogenous lags include a lot of information on more remote past of such variables.

Combining the results derived from the block-diagonal disaggregate monthly vector model, on a quarterly basis, with the forecasts derived from the aggregate quarterly econometric vector model including empirical long-term constraints, the combined forecast has always smaller RMSE if no inflation data have been observed for any month of the first quarter in the forecasting path. But if the monthly model can include data of the two first months of this quarter, then the combination with econometric forecasts not always improves the results.

An alternative to the combination of forecasts mentioned above could be formulating monthly models by adding the cointegration relationships between prices and the other economic variables, derived from the quarterly econometric vector model, to the monthly disaggregate model as exogenous variables. For this propose the cointegration relationships are interpolated at monthly level. The monthly models, besides contemplating heterogeneous inflation in different sectors, will thus also be able to contemplate the impact of different factors determining inflation, which could be different for the different components. But this last procedure does not obtain an improvement over the forecasts obtained with the monthly model which does not include such cointegration relationships.

From the combined forecasts of table 7, the contributions of the different economic variables to inflation according to the econometric model should be adjusted to obtain a precise explanation of the factors determining the final forecasts. Likewise, we will have to adjust the forecasts of the different price sub-indices by country and sector to provide a sector and geographic map of the estimated future values of inflation in the euro area. We thus obtain congruence between the geo-sectorial breakdown of inflation – which is necessary in any case to increase forecasting accuracy – and the contributions of the economic factors determining inflation forecasts. This is important, because the two sources of additional detail about future inflation are useful. The former informs of the nuclei (sectors through different countries) of more or less inflationist tension, and this is of interest for economic diagnosis and policy. The latter provides an estimation of the factors determining the inflation forecasts required by the authorities to design monetary policy and by economic agents to better assess inflation forecasts and, particularly, to form more accurate expectations related to changes in monetary policy.

4. CONCLUSIONS

In this work a unique econometric model, using the most frequent monthly data on price indexes, aimed at forecasting as accurately as possible while including appropriate estimates of the determining factors has not been possible. However, these type of results are demanded by the authorities and economic agents, and the paper shows that they can be obtained by separate models, combining their results and adjusting the partial contributions of the sectors and economic factors. It is important to emphasise that the vector formulation with cointegration restrictions is important in both models.

The results of this paper are in line of thick modelling advocated by Granger and Jeon (2004).

The above results lead to a proposed method for forecasting inflation in the euro area which can be applied to other indicators or macroeconomic variables. The basic points of this method are:

(a) Use the greatest possible level of disaggregation over time if the quality of the data on such a level is acceptable and it is feasible to construct appropriate econometric models.

(b) Using a disaggregated, both functionally and geographically, data set, including the long-term constraints between the components in the vector model.

(c) Simplifying the model by block-diagonal formulations, so that the number of parameters required is compatible with the size of the sample available.

(d) Including specific and general indicators when explaining the different components of the aggregate phenomenon.

(e) Combining forecasts from different models if it increases the accuracy of the forecasting paths, normally constructed for the current and following year.

(f) If the above forecasts are not based on a congruent econometric model, they should be combined with those provided by a model of this kind, possibly on an aggregate formulation of the variable of interest, to obtain an economic explanation, followed by adjusting the partial contributions of the components and economic factors.

We can conclude by saying that the method proposed in this work is based on the principle of progressive augmentation of the relevant data set with an appropriate econometric analysis in each case, which is determined by the forecasting improvements derived from such a progressive approach.

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