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Europe quo vadis? - Regional Questions at the Turn of the Century

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Structural Change in Regional Economies: a Varying Coefficients Econometric Modeling Approach

Abstract:

The traditional tools of econometric analysis are generally founded on the assumption that the structure of the economy is stable. This is a very restrictive hypothesis when modeling regional economies, since it is difficult to assume that future behaviour will be similar to that observed in the past.

The objective of the present work is to investigate the effects of structural change on the process of building regional econometric models, examining how this process deviates from standard econometric practice when the structural relationships are not constant. The method will be illustrated by specifying and estimating an econometric model for Extremadura, a region in SW Spain which has undergone profound changes in its economic structure in the last decade.

JEL classification: C22, C51, R15.

Keywords: structural change, cointegration, regional econometric models, varying parameters, Kalman filter.

1. INTRODUCTION

The construction of an econometric model on the regional scale typically has the aim of obtaining inferences about the region's present or future from historical data. In many cases these inferences concern the internal workings of the economy (intersectorial relationships, dependence on the exterior, efficiency, etc.) or the analysis of the impact of certain economic policy decisions (prices, subsidies, employment policies, etc.). On other occasions, the goal pursued is to forecast some variable of interest (production, employment, investment, etc.). In either case, for the inferences to be valid the models that are constructed have to be "stable" in the sense that it is assumed that the future will be similar to the past, with the implication that the parameters of the econometric model have to be constant. If the hypothesis of constant parameters is not satisfied in practice [1], whatever inference obtained from them and whatever economic policy implication derived from the model will be biased. In particular, the out-of-sample simulations and forecasts will be greatly affected, so that the usefulness of the model as a valid instrument on which to base economic policy decisions will be questionable.

In the case of regional economies, the problem of instability becomes even more acute than in the case of a national economy, since the impact of external or internal shocks is much greater than for the country as a whole. Thus, the agricultural sector (for climatological reasons) or the industrial sector (for reasons of location policies or of the production policies of large industrial companies) are prototypical economic activities in whose econometric modeling it is difficult to assume that the structure of the system characterizing them will be stable in the future.

The objective of the present work consists in analyzing how the presence of structural changes affects the process of constructing econometric models at the regional scale. The statistical framework used is cointegration theory (Engle and Granger, 1987), which combines in its basic econometric specification the relationships of long-term equilibrium suggested by economic theory with the process of (short-term) adjustment to equilibrium of the said relationships by way of error correction mechanisms. In the long term, and given the interpretation of the cointegration relationships, there is a foreseeable presence of a small number of break points, so that the working hypothesis is that the said changes can be modeled through the introduction of dummy variables[2] . In the short term, this hypothesis is replaced by another more general hypothesis in which the parameters of the model can vary continuously over time. The alternative hypothesis to constancy is that the parameters are stochastic[3] and vary according to a (multivariate) random walk model. The scope of the resulting model will then include all types of structural changes (sharp or smooth) which have taken place during the sample period.

The work is organized as follows. Section 2 gives details of the basic results concerning the structural specification that serves as the basis for the later empirical analysis. The specification of the econometric model for the region of Extremadura is conditioned both by the limitations on the statistical information available at a regional level in Spain and by the ultimate purpose of the model which is basically predictive. Section 3 is a methodological discussion, including the statistical model used for all the equations of the econometric model which is to be constructed in Section 4. In Section 4, there is first an analysis of the stochastic properties of the series that are used. The next step is the specification of the long-term equations of the seven economic sectors considered, analyzing in each case the stability of the estimated relationships. If necessary in the light of the results of the stability tests, this relationship is re-estimated, the short-term error correction mechanisms are modeled, the hypothesis of parametric stability is again tested, and each equation is re-specified if necessary, following the methodological considerations of Section 3. Section 5 gathers certain conclusions that can be drawn from our work.

2. THE BASIC STRUCTURAL MODEL

The econometric model that has been constructed (which will be described in detail in Section 4) is based on the fundamental ideas of the "economic base" models (see, for example, Treyz, 1993). For statistical reasons (in Spain there are no sufficiently long[4] regional time series disaggregated from the perspective of demand), the proposed economic model performs a supply-side sectorial disaggregation. Specifically, the regional production is divided into seven economic sectors: agriculture, energy, manufacturing industries, construction, sales-oriented services (except transport and communications), transport and communications, and non-sales-oriented services. The endogenous variables to be explained (and forecasted) are given by the production of each of these sectors, measured by the gross value added at market prices in 1986 constant pesetas (which we shall represent as W).

Following the line of argument of the economic base models, one can distinguish between basic and non-basic or local sectors, the former being those whose production supplies the national or supranational market, and the latter whose production is sold in the regional market. For the former, the level of activity is fundamentally determined by external factors, so that the standard specification is of the form:

$$E[W^b | IE^b, IR_1^b, IR_2^b, ...] = \alpha_b + \beta_{b_0} IE^b + \sum_{i=1}^{n_b} \beta_{b_i} IR_i^b \quad \textbf{(1)}$$

where $E[.|.]$ represents the conditional expectation value, the superindex refers to the basic sector *b*, *IE* is an external indicator which measures the evolution of the cycle of the national market of sector *b* [5], and *IR* are regional indicators which complement the basic specification (including variables which measure the advantages of siting the sector *b* in the region, as well as others which explain the intersectorial relationships within the regional ambit). For the local sectors, the standard equation is:

$$E[W^l | IR, IR_1^l, IR_2^l, ...] = \alpha_l + \beta_{l_0} IR + \sum_{i=1}^{n_l} \beta_{l_i} IR_i^l \quad \textbf{(2)}$$

where now *IR* is an indicator of the level of total internal demand in the region[6], and IR^l are regional indicators which complement the basic relationships (including variables which reflect the relationships of the local sector with basic activities of the region).

As in practice, however, there do not exist any purely national or regional markets, the productive sectors are usually mixed in the sense that part of their activity is determined by factors that are exogenous to the region and part by endogenous circumstances[7]. This fact implies that the relationships that will be specified for the different sectors will be a mixture of Equations [1] and [2].

3. METHODOLOGY

3.1 The Econometric Specification

The functional structure of the equations making up the regional econometric model is based on the modern theory of cointegration (Engle and Granger, 1987), which has shown itself to be a very useful tool for regional economy specialists because of its applicability in different econometric modeling ambits. Specifically, "standard" equations are proposed in the form of error correction mechanisms (ECM), in which a long-term equilibrium relationships is set up between the explanatory (not necessarily exogenous) and the explained (endogenous) variables, at the same time as allowing the existence of short-term deviations with respect to this equilibrium situation through the introduction of dynamic terms.

The basic structure that we propose for the model equations is a variant of the traditional ECM, and is given by the expression

$$\Delta W_t^s = \gamma_0 + \sum_{i=1}^{p} \gamma_{1i} \Delta W_{t-i}^s + \sum_{i=0}^{q} \boldsymbol{\gamma}_{2i} \Delta \boldsymbol{Z}_{t-i}^s - \alpha(W_{t-1}^s - \boldsymbol{\beta}' \boldsymbol{X}_{t-1}^s) + u_t^s \qquad \textbf{(3)}$$

where W^s represents the gross value added of sector s (in logarithms), $\boldsymbol{X}^s$ is a vector of explanatory variables (generally also in logarithms) which cointegrate with the dependent variable W^s [8], and $\boldsymbol{Z}^s$ is a vector of variables which explain (together with the lagged values of the dependent variable) the short-term deviations of the situation from equilibrium (amongst the components of $\boldsymbol{Z}^s$ may be found some of the variables of the vector $\boldsymbol{X}^s$ [9]).

3.2 The Statistical Model

In the specification [3] we have assumed that the parameters are fixed, i.e., that the structural relationships are stable in the short and long term. To relax this hypothesis, we shall distinguish two cases according as to whether the structural instability is present in the long or the short term.

Equation [3] assumes that the linear combination W_t^s-$\boldsymbol{\beta}'\boldsymbol{X}_t^s$ of the integrated variables has a stationary distribution. There exists, however, the possibility of a more general type of cointegration allowing the cointegration vector $\boldsymbol{\beta}$ to change at some point of the sample period[10]. The standard cointegration null hypothesis implies the model

$$W_t^s = \beta_0 + \boldsymbol{\beta}_1' \boldsymbol{X}_t^s + e_t^s \qquad \textbf{(4)}$$

where W^s_t and $\boldsymbol{X}^s_t$ are I(1) and e^s_t is I(0). If relationship [4] is stable, the parameters β_0 and $\boldsymbol{\beta}_1$ must be constant (time invariant). But if there exists structural instability the said parameters will remain constant over some period of time to change subsequently (β_0 or some component of the vector $\boldsymbol{\beta}_1$) to a new level, yielding another equilibrium relationship with different values for the slope or the ordinate at the origin. This change may be definitive, but it may also happen that after a certain period of time one returns to the original situation or passes to another equilibrium state characterized by a new set of coefficients. If we assume that the change in the parameters is discrete, the structural change can be modeled by introducing a fictitious variables of the type

$$\varphi_{[t_0,t_1]}(t) = \begin{cases} 0 & si \quad t<t_0 \\ 1 & si \quad t_0 \le t \le t_1 \\ 0 & si \quad t>t_1 \end{cases} \qquad \textbf{(5)}$$

where t_0 denotes the breaking point of the cointegration relationships and t_1 the point of return to the initial situation, with $1<t_0\leq t_1\leq T$. In the most general case where structural change implies a modification of both the ordinate at the origin and the slopes, the cointegration relationships with structural change is given by

$$W_t^s=\beta_{01}+\beta_{02}\varphi_{[t_0,t_1]}(t)+\boldsymbol{\beta}_{11}^{/}\boldsymbol{X}_t^s+\boldsymbol{\beta}_{12}^{/}\boldsymbol{X}_t^s\varphi_{[t_0,t_1]}(t)+e_t^s \qquad \textbf{(6)}$$

where β_{01} and β_{02} represent, respectively, the ordinates at the origin before and after the structural change, and $\boldsymbol{\beta}_{11}$ and $\boldsymbol{\beta}_{12}$ the slope coefficients in the cointegration relationships before and after the change of regime.

The above model can be generalized to allow more than one breaking point by simply introducing additional fictitious variables. In any case, the change of regime that is being considered is entirely discrete. This hypothesis could be relaxed by way of two different mechanisms. One (which will be used in the present work for the case of modeling the short-term relationships) consists in allowing the coefficients of Equation [4] to follow a stochastic process of the type $\boldsymbol{\beta}_t=\boldsymbol{\Psi}\boldsymbol{\beta}_{t-1}+\boldsymbol{\xi}_t$ such that the parameters change continuously throughout the sample period[11]. The other alternative is to model the structural change parametrically, allowing the parameters to change gradually from one stable system to another through some function Υ. This is the model proposed by Lin and Teräsvirta (1994), who consider a specification of the type[12]

$$W_t^s=\boldsymbol{\beta}_1^{/}\boldsymbol{X}_t^s+\boldsymbol{\beta}_2^{/}\boldsymbol{X}_t^s\Upsilon(\boldsymbol{Z}_t,\boldsymbol{\theta})+e_t^s \qquad \textbf{(7)}$$

where $\Upsilon(\boldsymbol{Z}_t,\boldsymbol{\theta})$ is a transition function which allows the model to change from the state $E[W_t^s|\boldsymbol{X}_t^s]=\boldsymbol{\beta}_1'\boldsymbol{X}_t^s$ to $E[W_t^s|\boldsymbol{X}_t^s]=(\boldsymbol{\beta}_1+\boldsymbol{\beta}_2)'\boldsymbol{X}_t^s$ through the variables $\boldsymbol{Z}_t$. If $\boldsymbol{Z}_t$=t and $\Upsilon=\varphi_{[to,t1]}$ we recover the discrete case, but other function types will make the model of change from one regime to another more flexible (see Lin and Teräsvirta, 1994 for an analysis of different choices).

With respect to the parameters of the error correction mechanism[13], and expecting that in the short term there may exist major instabilities, we formulated a model that is adaptable to any type of change (sharp or smooth) that may occur. In this sense, the deterministic model of fictitious variables is a priori too rigid, it being more advisable to use a stochastic model[14] which allows greater flexibility in the temporal evolution of parameters. Re-specifying the model [3] such that all the parameters $(\gamma_0,\gamma_{11},...,\gamma_{1p},\boldsymbol{\gamma}_{20},..,\boldsymbol{\gamma}_{2q},-\alpha)$ appear in the vector $\boldsymbol{\alpha}$, and all the explanatory variables appear in the vector $\boldsymbol{H}^s$ [15], the structure of the equations of the model that

we propose is the following

$$\begin{aligned} \Delta W_t^s &= \boldsymbol{\alpha}_t' \boldsymbol{H}_t^s + u_t^s \\ \boldsymbol{\alpha}_t &= \boldsymbol{\alpha}_{t-1} + \boldsymbol{\eta}_t \end{aligned} \qquad \textbf{(8)}$$

where we assume that the errors u_t^s are normally distributed with zero mean and constant variance σ^2 and are mutually independent, and that $\boldsymbol{\eta}_t$ is a vector of normal random variables with zero mean and covariance matrix $\sigma^2\mathbf{P}=\mathbf{Q}$ whose distribution is independent of that of the errors u_t^s and of the vector $\boldsymbol{\alpha}_0$. The fist equation of the system [8] is known as the measurement equation and the second as the transition equation, which describes the temporal evolution of the parameter vector of interest, $\boldsymbol{\alpha}_t$, now known as state vector (and its components state variables). In our application, we shall assume that the matrix **Q** (known as the dispersion matrix) is diagonal, i.e., the state variables are not allowed to interact amongst themselves which would involve non-zero off-diagonal elements. The case $\mathbf{Q}=\mathbf{0}$, of course, reduces to the constant parameter model [3].

The specification [8] assumes that the parameter vector $\boldsymbol{\alpha}_t$ follows a (trendless) random walk type of multivariate distribution which, as it is not stationary, evolves with time such that all the structural changes that have taken place during the sample period can be included. Obviously, other stochastic models for $\boldsymbol{\alpha}_t$ can be put forward, depending on the a priori level of information that one possesses on the form, timing, and speed of the structural change[16]. In our case, given that we lack the said information[17], we preferred to use a random walk model as alternative hypothesis (quite customary elsewhere).

4. THE EMPIRICAL RESULTS

4.1 Statistical Sources and Stochastic Properties of the Data

Most of the variables used in the work were obtained from the HISPALINK and HISPADAT[18] data bases. The former is a historical collection of data for the period 1970-85 (see Hispalink, 1993). The latter covers the period 1986-95 and consists of official data from the National Statistics Institute (INE), and forecasts from 1996 to 1999 (see Hispalink, 1997). In particular, all variables referring to gross added value (by sector and at the regional and national levels) were obtained from the said bases. The rest was constructed from different sources of regional or national statistical information[19]. In general, the data that are to be used cover 1970 to 1995 (the last year for which official figures are available), although the length of the series is reduced in some sectors due to the inexistence of disaggregated series for several years at the

beginning of the period under consideration. One of the goals of the present work is to analyze the out-of-sample forecasting performance of the model, and for this reason we retain the years 1994 and 1995 to make ex-post predictions. The sample period used in the following estimates is therefore generally that between 1970 and 1993.

Before continuing with the estimate of the equations of the model, we need to analyze the order of integrability of all the exogenous and endogenous variables that appear in the model (see footnote 9). For this purpose we used the augmented Dickey-Fuller (ADF) test (Dickey and Fuller, 1981) which is based on the following regression equations (the null hypothesis being H_0:{x_t~I(d)} and the alternative hypothesis H_1:{x_t~I(d-1)}):

$$\begin{aligned} \Delta^d x_t &= \alpha_0 + \alpha_1 \Delta^{d-1} x_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta^d x_{t-j} + \varepsilon_t \\ \Delta^d x_t &= \alpha_0^* + \alpha_1^* \Delta^{d-1} x_{t-1} + \alpha_2^* \Delta^{d-1} t + \sum_{j=1}^{p} \gamma_j^* \Delta^d x_{t-j} + \varepsilon_t^* \end{aligned} \qquad \textbf{(9)}$$

where the errors are assumed to be Gaussian "white noise" type perturbations. The t statistics of α_1 or α_1^* are the values used to test the hypothesis that this coefficient is zero or significantly different from zero.

Taking into account, however, previous experiences with the study of the stochastic properties of Spanish macroeconomic series (Andrés *et al.*, 1990; Molinas *et al.*, 1991), as well as the nature of the regional series themselves (with frequent break points[20]), we also considered the more general version of the ADF test to include the possibility of the existence of segmented deterministic trends in the mean (Rappoport & Reichlin, 1989). In this case the mean can be written as

$$\mu_t = \begin{cases} c_1 + b_1 t \;\; para \;\; t \le t_1^* \\ c_2 + b_2 t \;\; para \;\; t_1^* \le t \le t_2^* \\ \qquad \ldots \\ c_n + b_n t \;\; para \;\; t \ge t_{n-1}^* \end{cases} \qquad \textbf{(10)}$$

where t_i^* are the points when there is a break in the trend. The ADF test then takes the form

$$\Delta^d x_t = \alpha_1 \Delta^{d-1} x_{t-1} - \alpha_1 \Delta^{d-1} mu_{t-1} + a^{**}(L) \Delta^d x_{t-1} + a^*(L) \Delta^d mu_t + \varepsilon_t \qquad \textbf{(11)}$$

where the lag polynomials a^*(L) and a^{**}(L) are related (imposing the normalization constraint

$a^*(0)=1$) through the equation $a^*(L)=1+a^{**}(L)L$. The regression equation [11] is made operative by replacing μ_t by the expression

$$\mu_t = c_1 + b_1 t + \sum_{i=2}^{n} (c_i - c_{i-1}) D_{i-1,t} + \sum_{i=2}^{n} (b_i - b_{i-1}) D_{i-1,t} t \qquad \textbf{(12)}$$

where the dummy variable $D_{i,t}$ takes the value unity from the instant t_i^*.

The results of applying the tests described above to the variables under study are listed in Table 1[21]. The conclusion to be drawn is that all of the series, except one [which can be considered as I(0)], can be regarded as I(1) variables, some of them with a single deterministic trend, and the rest of the variables with various segmented trends in the mean.

4.2 The Estimation of the Model

Following the philosophy of the two-stage method proposed by Engle and Granger (1987), we firstly estimated the long-term relationships of the type [4] using the theoretical arguments outlined in Section 2. In all cases the cointegration test of Engle and Granger (which uses the unit root test of Dickey-Fuller, 1979) was applied to determine whether the variables involved in the regression were cointegrated. The result was that the null hypothesis of the existence of a unit root in the residuals (i.e., the absence of cointegration) was not rejected in most of the tests.

In the light of this evidence, and taking into account the results of Campos *et al.* (1996) (which show not only that the presence of structural change in stationary series may lead to spurious unit roots[22], but also that such breaks affect considerably the power of cointegration tests in general, and in particular of the two-stage procedure of Engle and Granger[23]), we applied a series of parametric stability tests to detect the presence of break points in each of the equations of long-term behaviour[24].

Specifically, we applied various tests based on the calculation of Wald sequential statistics, $F_T(\delta)$, which test the null stability hypothesis against the alternative of the existence of some break point in the observation t_0 (more precisely, in the fraction $\delta_0=t_0/T$ of the sample). The exact localization of the possible points of intersection is not known a priori, so that the statistics $F_T(t/T)$ are calculated for all the points of the sample[25], and then some functional of these statistics is constructed. The three functionals considered[26] in our work are: the Quandt (1960) likelihood ratio statistic

$$SupF = \sup_{\delta \in [\delta_1, \delta_2]} F_T(\delta) \qquad \textbf{(13)}$$

the mean statistic proposed by Andrews and Ploberger (1994) and Hansen (1992)

$$MeanF = \int_{\delta_0}^{\delta_1} F_t(\delta)\, d\delta \qquad \textbf{(14)}$$

and the average exponential statistic proposed by Andrews and Ploberger (1994)

$$ExpF = \ln \left\{ \int_{\delta_0}^{\delta_1} \exp\left(\frac{1}{2} F_t(\delta) \right) d\delta \right\} \qquad \textbf{(15)}$$

In all cases of applying these three tests to the model's long-term equations except one, the values of the tests surpassed the critical values corresponding to a 1% significance level. (In the sole exception, the null hypothesis was rejected at a 5% level.)

The following step consisted in introducing the fictitious variables needed to approximate the structural changes detected through the sequential application of Wald statistics[27], i.e., cointegration relationships of type [6] were estimated with the same number of functions $\varphi_{[to,t1]}$ as break points detected. The results for each of the sectors considered are listed in Table 2. We shall now make some remarks concerning the said results.

Firstly, for each of the estimated equations, the stability tests *SupF*, *MeanF*, and *ExpF* were again applied, with the result that the null hypothesis of stability of the estimated parameters was not rejected in any case.

Secondly, as can be seen in the table, the *DF* statistic rejects in all cases the presence of a unit root in each equation's estimated errors, i.e., the linear combinations of the variables of each model are stationary and, therefore, the said relationships can be interpreted as long-term cointegration or equilibrium equations with changes of regime.

Thirdly, one observes that there are few break points[28], with at most three structural changes per equation (in the cases of the industrial, the transport and communications, and the non-sales-oriented service sectors). Also, in several cases there are simultaneously changes in the level and in the slopes of the model, with two sectors (those of sales-oriented and non-sales-oriented services) in which different regimes were detected, with two break points with change both in the slope and in the intercept.

After the estimation of the long-term relationships, the second step of the procedure of

Engle and Granger (1987) consists in estimating the short-term equations given by [3], with the expression in parentheses (which would now be of the type [6] with various functions $\varphi_{[to,t1]}$) replaced by the estimated errors that are derived from Table 2.

As in the case of the long-term model, firstly we estimated the individual equations without dummy variables, and then performed the three stability tests *SupF*, *MeanF*, and *ExpF*. Except for the case of the transport and communications sector, in the rest of the sectors the three statistics rejected simultaneously the null hypothesis of stability of the parameters of the error correction mechanisms.

Two options were considered to take into account the presence of structural change. One was to use, as in the long-term case, dummy variables to pick up the effect of the said structural changes. The other consisted in modeling the structural break by proposing as alternative the adaptive model represented by Equations [8]. These two options are developed in the following.

The information provided by the Wald statistics used in the stability tests (complemented with a graphical analysis of each dependent variable and of the residuals estimated from the initial model) aided in identifying the dummy variables to be introduced into each of the equations of short-term behaviour. The final result is the set of regressions presented in Table 3.

With reference to the number of fictitious variables introduced, two remarks should be made. Firstly, the number of fictitious variables of each equation is determined by the values of the stability tests: with the variables that are introduced (and only with these) one attains stability for the error correction mechanisms in the sense that none of the three proposed statistics surpasses the corresponding threshold. Secondly, and as was to be expected, the number of break points that appear is far greater than in the long-term relationships, pointing to the presence of greater instability in the short-term relationships than in the equilibrium equations.

With respect to the estimate of Equations [8] for each of the seven sectors considered[29], the estimation technique used is based on the recursive application of the Kalman filter (Kalman, 1960). At each instant t, and given the observations $\Delta W_1^s,...,\Delta W_t^s$, the interest is centred on estimating the vector $\boldsymbol{\alpha}_t$ using the model equation [8]. Under the normality hypotheses established for the errors of the model (and for the initial state vector $\boldsymbol{\alpha}_0$), and assuming that σ^2, $\mathbf{P}$ (or $\mathbf{Q}$), and the mean and covariance matrix of $\boldsymbol{\alpha}_0$ are known[30], the optimal estimator of $\boldsymbol{\alpha}_t$ using the information I_s available up to instant s, is given by the conditional expectation of $\boldsymbol{\alpha}_t$ taking I_s as known, which we will denote by $E[\boldsymbol{\alpha}_t| I_s]=\mathbf{a}_{t|s}$ [31]; and the optimal estimator for the covariance matrix of $\boldsymbol{\alpha}_t$ using the available information I_s will be given by $Cov[\boldsymbol{\alpha}_t| I_s]=\Sigma_{t|s}$.

The Kalman filter recursions for t=0, 1, 2, ... are given by the following equations (see,

for example, Lütkepohl, 1993):

$$\begin{aligned} \boldsymbol{a}_{t+1|t} &= \boldsymbol{a}_{t|t} \\ \Sigma_{t+1|t} &= \Sigma_{t|t} + \boldsymbol{Q} \\ \boldsymbol{a}_{t+1|t+1} &= \boldsymbol{a}_{t+1|t} + k_{t+1}[\Delta W^{s}_{t+1} - \boldsymbol{\alpha}^{/}_{t+1|t}\boldsymbol{H}^{s}_{t+1}] \\ \Sigma_{t+1|t+1} &= \Sigma_{t+1|t} - k_{t+1}\boldsymbol{H}^{s}_{t+1}\Sigma_{t+1|t} \end{aligned} \qquad \textbf{(16)}$$

where

$$k_{t+1} = [\boldsymbol{H}^{s}_{t+1}\Sigma_{t+1|t}\boldsymbol{H}^{/}_{t+1} + \sigma^2]^{-1}\Sigma_{t+1|t}\boldsymbol{H}^{/}_{t+1} \qquad \textbf{(17)}$$

Then the *n*-period forward forecast of $\Delta W_{t+n}{}^{s}$ will be given by

$$\Delta W^{s}_{t+n|t} = \boldsymbol{\alpha}^{/}_{t|t}\boldsymbol{H}^{s}_{t+n} \qquad \textbf{(18)}$$

In the application of the above formulae, there are a number of problems which it is necessary to address referring to the set of parameters which are assumed as known a priori. Specifically, since a "random walk" parametric variation model was specified for $\boldsymbol{\alpha}_t$, there do not exist any automatic values (such as the mean or unconditional covariance matrix) for the values $\mathbf{a}_{0|0}$ and $\Sigma_{0|0}$. Neither are the elements of the matrix **Q** nor the parameter σ^2 known. The latter is the least problematic since it may be estimated by maximum likelihood, isolating it from the rest of the parameters (Chow, 1984, p. 1222).

With respect to the initialization values of the Kalman filter, one may use the first *K* observations, with *K* being the dimension of the state vector (Harvey, 1981, 1989), or an a priori "diffuse" value (Ansley and Kohn, 1983), or they can be estimated by maximum likelihood together with the rest of the model's parameters (Chow, 1984). In the present case, we used another alternative, fixing (as is done in Hackl and Westlund, 1996) the elements $\mathbf{a}_{0|0}$ andy $\Sigma_{0|0}$ at the MCO values obtained by estimating the model with constant parameters (and without dummy variables) over the complete sample period[32].

With respect to the elements of the matrix **Q** (known as hyperparameters), there are two possible routes. They may be estimated by a maximum likelihood method (Chow, 1984), or fixed beforehand as proposed again by Hackl and Westlund (1996), to avoid problems of lack of identification and large oscillations in the estimates of the state variable parameters. In the present case, we followed both options: estimating the elements q_{ii} of the matrix **Q** by maximum

likelihood, and fixing their values beforehand by taking **Q**=**I** [33].

4.3 Ex-post Forecasting Analysis

In this subsection, we compare the out-of-sample forecasting performance of the short-term models estimated in the previous section. As was noted in Subsection 4.1, all the models were estimated (in general) for the period 1970-93, leaving the years 1994 and 1995 as test beds on which to carry out an experiment of ex-post forecasting. Thus, the forecasts made with the structural models are based on real values of the explanatory variables, and the said forecasts were compared with the observed values of the endogenous variables for the years under consideration.

To measure the degree of goodness of the forecasts, we used four known statistics based on symmetric loss functions. In particular, we calculated the mean error (ME), the mean absolute error (MAE), the root mean square error (RMSE), and the inequality coefficient (U) of Theil (1966)[34].

The results of the ex-post simulation carried out with the three types of model used (**Q**=**0**, i.e., the model with fixed parameters and dummy variables; **Q**=**I**, i.e., fixing the hyperparameters beforehand; and **Q**=diag$\{q_{11,}q_{22},...,q_{KK}\}$, i.e., estimating by maximum likelihood the diagonal elements but setting the rest to zero) are presented in Table 4. Some remarks should be made concerning the values that appear in the table.

Firstly, one observes from the comparison of the statistics that there is no clear dominance of the variable parameter models over the fixed parameter model: the results are improved relative to the latter in three of the seven sectors, and similar in another two cases.

Secondly, the model in which the hyperparameters are estimated yields generally poorer results than the model where they are fixed beforehand. This may be due to the problems of identification of such parameters caused by the small sample size.

Lastly, with respect to the U statistic of Theil (1966), in the case of the energy sector, the three models' forecasts are systematically worse than the naive random walk model. In the remaining sectors, the three models' forecasts generally perform better than the random walk model, although it is again observed that the variable parameter model, in which the diagonal elements q_{ii} of the dispersion matrix **Q** are estimated, yields worse results than the other two models.

5. CONCLUSIONS

The stability of an econometric model is a basic requisite for it to be used for forecasting or inferential purposes. At a regional scale, however, the parametric instability of the models estimated with historical data is quite usual, since "changes in regime" are frequent at this level. Hence, the forecasting experiments that are carried out to simulate the impact of different economic policy decisions will only have meaning if the (necessary) condition of stability is satisfied.

There is a large literature concerning tests of the hypothesis of parametric stability (see, for example, Stock and Watson, 1996, and references therein), but there have been far fewer contributions on the question of how to model structural change once it has been detected. Often, a significant stability test result indicates some type of faulty specification, so that the next step is to try alternative specifications. In other cases, the parameters are allowed to vary throughout the sample, the commonest method used being to introduce dummy variables that interact with the original variables to allow changes in the slopes and/or the intercept. An alternative is to allow the regression coefficients to vary at random, endowing the resulting model with a greater flexibility than in the fictitious variables case.

In the present work we used the two approaches to construct an econometric model for one of the Spanish regions, Extremadura, which has undergone profound changes in its economic structure since the beginning of the 1980's. The econometric framework was cointegration theory, which distinguishes long-term economic relationships from short-term dynamics by introducing error correction models which form the analytical basis of the econometric model constructed in this work.

The results show firstly that in the case of Extremadura structural instability exists in both the long and the short terms. The standard econometric methods are therefore not applicable in our case. In the long term, the introduction of fictitious variables is sufficient to pick up the changes in regime that occurred during the sample period. In the short term, however, the instability is far greater, so that it seems more reasonable to make the model of change more flexible, allowing the parameters to vary at random. In the forecasting analysis that was carried out, the variable parameter model yielded results that were similar to those of the fixed parameter (with dummy variables) models, but we consider that the time varying parameter model has the advantage of incorporating future uncertainty about the values of the parameters into the forecasts, which is not the case with the dummy variables model.

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FOOTNOTES:

1 Different factors may be the cause of the parametric instability. Amongst them, we would emphasize: a) bad specification of the model equations, originated by the deficient contribution of economic theory in proposing causes and the functional form linking the same with the dependent variable; b) appearance of changes in the economic policy instrument variables, which will alter the structure of the econometric model (Lucas, 1976); and c) unexpected general (supraregional) or specific (regional) shocks may also be the cause of instability in the model equations.

2 In Canarella *et al.* (1990), the presence of long-term structural instability is modeled in a time-varying parameter approach. As pointed out by Hall (1994), however, it would be difficult to give an economic interpretation to the resulting long-term model.

3 Modeling by means of dummy variables implies that the alternative to the hypothesis of constancy is a deterministic change of the parameters in time. Apart from this difference, the stochastic approach proposed here assumes that the change in the parameters is "smooth", as against the hypothesis of discrete (and hence sharp) change of the dummy variable approach.

4 The HISPALINK and HISPADAT data bases, from which have been taken the main macromagnitudes used in the present work, only contain regional final demand series for the period 1986-93.

5 The coefficient β_{lo} will then measure the sensitivity of the sector's regional production to variations in the added value of the same at the national scale.

6 In this case, the coefficient β_{lo} will measure the sensitivity of the sector to changes in regional income.

7 Which is no obstacle to the basic or local character prevailing in each particular sector.

8 $W^s=\boldsymbol{\beta}\mathbf{X}^s$ will then be the long-term equilibrium relationship, and $W_t^s-\boldsymbol{\beta}'\boldsymbol{X}_t^s$ will measure the deviations from equilibrium at each instant *t*.

9 The proposed specification assumes that all the variables that appear (in levels) in [3] are I(1).

10 For simplicity of exposition, we assume that there exists only one break point.

11 If $\boldsymbol{\Psi}=\mathbf{I}$ we shall have a "random walk" model, and if $\boldsymbol{\Psi}=\mathbf{B}=\text{diag}\{b_1, b_2,..., b_K\}$ the "return to normality" model. At least from the long-term perspective, both models - considered overall - are "excessively" general, in the sense that they presuppose a continuous structural change (from one period to the next, all the coefficients change). In this sense, at least in the return to normality model, the coefficients converge (if the parameters b_i satisfy the constraint of being less than unity in absolute value) to a stationary state, $\boldsymbol{\beta}^*$, which would be that having the interpretation of cointegration vector.

12 Which Lin and Teräsvirta denominate the smooth transition model (STR). The model [7] also assumes that there only exists one break point in the long-term relationship, but it can be readily generalized to the case of more than one structural change.

13 In this work we shall follow a strategy of two-stage estimation (Engle and Granger, 1987), in which at the first stage the cointegration relationship is estimated, and then the econometric model mechanism is estimated by introducing the (lagged) residues of the long-term relationship estimated in the first stage.

14 One could also use a model of the STR type as in Wolters *et al.* (1996).

15 Replacing the cointegration vector $W^s_{t-1}-\boldsymbol{\beta}'X^s_{t-1}$ by the value of the estimated residuals.

16 See Section 4 of Hall (1994).

17 And the small number of observations, which limits any type of generalization.

18 In Otero *et al.* (1996), there is a discussion of the origin, content, and methods used in the construction of the two data bases.

19 In Ramajo and Márquez (1996), a detailed analysis is given of the different sources of information used to a greater or lesser degree in our work.

20 Due not only to structural changes such as those analyzed in the present work, but also to problems "in origin", such as changes of basis, redefinition of variables, measurement errors caused by the application of distribution methods, use of approximate deflators, etc.

21 For reasons of space, details will not be given of each of the regressions carried out.

22 A result which had already been demonstrated analytically by Perron (1989) and shown empirically by Hendry and Neale (1991).

23 It has to be taken into account that the standard cointegration tests assume that the cointegration vector is time invariant under the alternative hypothesis.

24 Indeed, on applying the cointegration tests proposed by Gregory and Hansen (1996), which allow the possibility

of changes of regime, the result was that in all cases the null hypothesis of no cointegration was rejected (in this case, as against the alternative of cointegration in the presence of a possible break point). Since, however, the tests of Gregory and Hansen only permit one break point (the procedure they use also allows the point to be identified), and given the possibility that a greater number exist in our case, the process of structural change modeling has been continued.

25 As noted by Andrews (1993), one can not use all the points $t/T\in[0,1]$, since in this case the tests will diverge to infinity, so that he proposes using the region $\Im=[\delta_1,\delta_2]=[0.15,0.85]$.

26 The asymptotic distributions of these statisitcs are discussed in Andrews (1993) and in Andrews and Ploberger (1994), being in all cases non-standard (functionals of multi-dimensional Brownian motions).

27 We are aware of some of the problems that this approach might lead to. Firstly, MCO estimation is not efficient, and the significance tests do not have the standard asymptotic distributions under the hypothesis of cointegration with changes of regime. Secondly, under the null hypothesis of parameter stability, and given that the break points are unknown a priori, the Wald statistics constructed also have non-standard distributions. The results of Hansen (1992) and Quintos and Phillips (1993) might be useful in resolving these problems.

28 Which in a certain sense justifies the dummy variable approach, because of its ready implementation as against other more elaborate alternatives.

29 Although it was not in principle necessary to re-estimate the model corresponding to the transport and communications sector, which is stable in the short term, it was also included in this phase in order to compare the results of the fixed parameter and the varying parameter models.

30 Which we shall denote by $\mathbf{a}_{0|0}$ and $\Sigma_{0|0}$, respectively.

31 When $s=t$, the evaluation of $\mathbf{a}_{t|t}$ is known as filtering, and when $s>t$ as smoothing.

32 We recognize that, if one wants to fix the initial values instead of estimating them, either of the other two initialization methods considered is more orthodox and correct than that used here. In one of the cases, the small sample size did not allow us to discard certain observations at the start of the sample period. In the second, we believed it advisable to give an initial value by incorporating information a priori, since with so few observations a diffuse initialization value could give rise to trajectories with large fluctuations from one period to the next originated by a poor choice of the initial point.

33 Wolff (1987) considers a range of matrices of the type $\mathbf{Q}=\gamma\Sigma_{0|0}$ for γ=0.0, 0.01, 0.05, 0.10 and 0.25, representing in order lesser to greater variability in the state variables and interaction amongst the same.

34 The U statistic is the ratio of the RMSE of the forecasts obtained with the estimated model and the random walk model. This statistic has a straightforward interpretation: if U<1 the forecasts of the model are better than the naive forecasts, and if U>1 they are worse.

TABLES

TABLE 1: Results of the unit root tests

	vs			vs			
	ADF			**ADF**			
VARIABLE	PGD	*t*	VC	PGD	*t*	VC	Result
LVAES	CT,0	-5,49	-3,61	C,0	-2,46	-2,98	I(1)
LVAEX	N,1	-5,58	-1,95	C,0	-3,08	-3,73*	I(1)
LVOLAG	N,0	-5,95	-1.95	CT,0	-2,52	-3,60	I(1)
LVEES	CT,0	-5,21	-3,61	C,0	-2,69	-2,98	I(1)
LVEEX	N,0	-6,17	-1,95	CT,0	-2,44	-3,60	I(1)
LPRECIP	N,2	-4,58	-1,95	C,0	-4,20	-2,98	I(0)
LENERG	RR,2	-6,58	-4,08	RR,2	-2,48	-4,08	I(1)
LVBEX	C,0	-5,59	-2,99	CT,0	-2,72	-3,60	I(1)
LVABES	C,1	-3,09	-2,99	CT,1	-3,93	-4,39*	I(1)
LVABEX	C,0	-5,17	-3,00	CT,0	-2,40	-3,62	I(1)
LVIEX	N,0	-5,21	-1,95	C,0	-2,42	-2,98	I(1)
LVIES	N,0	-3,20	-1,95	CT,1	-2,77	-3,62	I(1)
LVNOAEEX	C,0	-5,07	-3,00	CT,0	-1,95	-3,62	I(1)
LVLEX	N,1	-2,35	-1,95	RR,3	-4,26	-4,76	I(1)
LVLES	RR,2	-4,14	-4,08	CT,1	-2,85	-3,62	I(1)
LVZEX	N,0	-3,11	-1,95	CT,1	-2,28	-3,62	I(1)
LVZES	RR,2	-4,39	-4,08	CT,1	-3,33	-3,62	I(1)
LVGEX	RR,2	-7,55	-4,08	RR,2	-3,72	-4,08	I(1)
LVGES	RR,2	-4,16	-4,08	CT,1	-2,86	-3,61	I(1)

Notes:
-The notation used to represent the variables is:
L indicates that logarithms have been taken.
V denotes the gross added value in constant 1986 pesetas. V is accompanied by a single letter to indicate the sector of reference (VA: agriculture; VE: energy; VB: construction; VI: manufacturing industry; VL: sales-oriented services (except transport and communications); VZ: transport and communications; VG: non-sales-oriented services); or of various letters for each of the sectorial groups considered (VAB: total gross added value; VNOAE: total non-agricultural and non-energy gross added value).
The termination of each variable specifies whether it is a national (ES) or regional Extremadura (EX) variable.
VOLAG is the volume of reservoir water at the end of each year in Extremadura.
PRECIP is the mean volume of precipitation recorded in Extremadura.
ENERG is the gross electrical energy production in Extremadura.
- The column PGD specifies the data generation process which was considered for the variable in question. Thus the letter N indicates that the estimated auxiliary regression included no deterministic component; C indicates the admission of a constant term; and CT denotes the presence of a constant and a deterministic linear trend. The number immediately following these letters (separated by a comma) indicates the number of lags introduced into the ADF test. The letters RR mark the consideration of segmented deterministic trends in the mean for the ADF test (the number of segments being indicated by the corresponding digit).
- In the following column, *t* is the t-statistic used in each of the ADF tests.
- VC is the tabulated critical value at a 5% significance level for each of the tests obtained from MacKinnon (1991) or Rappoport and Reichlin (1989). In the case of working at 1%, an * is placed beside the corresponding figure.

TABLE 2: Results of the estimation of the cointegration relationships (MCO with dummy variables).

AGRICULTURE
ESTIMATED EQUATION (t = 1971,...,1993)
$LVAEX_t = -2,431 + 0,951\ LVAES_t + 0,231\ D7175_t$ (-0,422) (2,388) (2,685)
SPECIFICATION TESTS
R^2 = 0,279; Durbin-Watson = 1,457; ADF = -3,191; *SupF* = 7,784; *MeanF* = 2,041; *ExpF* = 1,797.
ENERGY
ESTIMATED EQUATION (t = 1970,...,1993)
$LVEEX_t = -5,070 + 1,050\ LVEES_t + 0,561\ D7073_t + 1,087D8493_t$ (-0,743) (2,193) (2,652) (6,014)
SPECIFICATION TESTS
R^2 = 0,899; Durbin-Watson = 2,153; ADF = -4,868; *SupF* = 7,090; *MeanF* = 2,326; *ExpF* = 1,767.
CONSTRUCTION
ESTIMATED EQUATION (t = 1972,...,1993)
$LVBEX_t = -2,366 + 1,011\ LVABEX_t - 19,077\ D7079_t + 1,446\ D7079*LVABES_t$ (-1,154) (6,547) (-2,545) (2,511)
SPECIFICATION TESTS
R^2 = 0,941; Durbin-Watson = 1,829; ADF = -4,515; *SupF* = 7,948; *MeanF* = 1,745; *ExpF* = 1,618.
MANUFACTURING INDUSTRY
ESTIMATED EQUATION (t = 1970,...,1993)
$LVIEX_t = -7,994 + 1,182\ LVIES_t + 0,206\ F78_t + 0,223\ F80_t + 0,104\ D8285_t$ (-5,158) (12,039) (2,551) (2,767) (2,400)
SPECIFICATION TESTS
R^2 = 0,894; Durbin-Watson = 1,995; ADF = -4,476; *SupF* = 5,472; *MeanF* = 3,612; *ExpF* = 1,937.
SALES-ORIENTED SERVICES
ESTIMATED EQUATION (t = 1972,...,1993)
$LVLEX_t = -5,366 + 1,083\ LVLES_t + 43,937\ D8185_t - 2,724\ D8185*LVLES_t - 11,327\ D8693_t$ (-2,016) (6,556) (5,657) (-5,673) (-3,192) $+ 0,682\ D8695*LVLES_t$ (3,118)
SPECIFICATION TESTS
R^2 = 0,962; Durbin-Watson = 2,351; ADF = -5,586; *SupF* = 9,726*; *MeanF* = 2,983; *ExpF* = 2,464.
TRANSPORT AND COMUNICATIONS
ESTIMATED EQUATION (t = 1972,...,1993)
$LVZEX_t = -5,269 + 0,722\ LVZES_t + 0,399\ LVLEX_t + 0,257\ F72_t + 17,686\ D8084_t$ (-6,491) (16,269) (4,422) (7,038) (2,505) $- 1,232\ D8084*LVZES_t - 0,131\ D8889_t$ (-2,493) (-5,414)
SPECIFICATION TESTS
R^2 = 0,980; Durbin-Watson = 1,988; ADF = -4,335; *SupF* = 7,559; *MeanF* = 3,168; *ExpF* = 2,224.
NON-SALES-ORIENTED SERVICES
ECUACIÓN ESTIMADA (t = 1971,...,1993)
$LVGEX_t = -3,636 + 1,004\ LVGES_t + 3,645\ D7078_t - 0,257\ D7078*LVGES_t - 10,452\ D7985_t$ (-4,066) (17,219) (3,447) (-3,683) (-7,288) $+ 0,689\ D7985*LVGES_t + 0,024\ D9293_t$ (7,289) (2,225)
SPECIFICATION TESTS
R^2 = 0,998; Durbin-Watson = 2,682; ADF = -6,395; *SupF* = 2,603; *MeanF* = 0,513; *ExpF* = 0,340;

Notes:
- In parentheses, beneath the estimated coefficients, appear the t** statistics.
- * denotes rejection of the null hypothesis at 10%, ** at 5%, and *** at 1%.

TABLE 3: Results of the estimation of the error correction models (MCO with dummy variables)

AGRICULTURE
ESTIMATED EQUATION (t = 1972,...,1993)
$DLVAEX_t = 0{,}053 - 0{,}923\ _{t-1} + 1{,}005\ DLVAES_t + 0{,}171\ DLVOLAG_{t-1} + 0{,}214\ DD7175_t$ (4,358) (-8,426) (4,908) (3,337) (4,498) $- 0{,}130\ D7375_t - 0{,}400\ F83_t - 0{,}197\ D9293_t$ (-4,386) (-8,353) (-5,057)
SPECIFICATION TESTS
$R^2 = 0{,}998$; Adjusted $R^2 = 0{,}922$; Durbin-Watson: DW = 2,214; Jarque-Bera: P-val. = 0,361; Breusch-Godfrey: a) [AR(1)] P-val: 0,221; b) [AR(2)] P-val: 0,356; Ljung-Box (p=6): P-val.= 0,829; ARCH: a) [ARCH(1)] P-val: 0,201; a) [ARCH(2)] P-val: 0,600; White: P-val. = 0,761; *SupF* = 7,279; *MeanF* = 3,918; *ExpF* = 2,350.
ENERGY
ESTIMATED EQUATION (t = 1972,...,1993)
$DLVEEX_t = -0{,}142 - 0{,}771\ _{t-1} + 2{,}574\ DLVEES_t + 0{,}407\ DLPRECIP_t + 1{,}089\ DLENERG_t$ (-2,623)(-6,024) (2,876) (5,751) (10,128) $- 0{,}307\ D7476_t + 0{,}590\ F77_t + 0{,}526\ F82_t + 0{,}171\ D8892_t$ (-3,954) (4,343) (3,814) (2,697)
SPECIFICATION TESTS
$R^2 = 0{,}951$; Adjusted $R^2 = 0{,}922$; Durbin-Watson: DW = 2,384; Jarque-Bera: P-val. = 0,874; Breusch-Godfrey: a) [AR(1)] P-val: 0,083; b) [AR(2)] P-val: 0,073; Ljung-Box (p=6): P-val.= 0,065; ARCH: a) [ARCH(1)] P-val: 0,118; a) [ARCH(2)] P-val: 0,264; White: P-val. = 0,304; *SupF* = 10,621; *MeanF* = 5,270; *ExpF* = 3,757.
CONSTRUCTION
ESTIMATED EQUATION (t = 1973,...,1993)
$DLVBEX_t = 0{,}012 - 0{,}448\ _{t-1} + 0{,}892\ DLVABEX_t + 0{,}661\ DLVABES_t + 0{,}143\ DLVBEX_{t-1}$ (2,559)(-7,792) (9,230) (4,604) (3,398) $- 0{,}083\ D7475_t + 0{,}066\ D7677_t + 0{,}033\ F81_t + 0{,}068\ F83_t - 0{,}183\ F85_t + 0{,}113\ F86_t$ (-9,965) (8,060) (2,333) (5,744) (-16,459) (8,532) $- 0{,}250\ F88_t - 0{,}050\ F89_t$ (-21,128) (-3,298)
SPECIFICATION TESTS
$R^2 = 0{,}994$; Adjusted $R^2 = 0{,}985$; Durbin-Watson: DW = 2,305; Jarque-Bera: P-val. = 0,746; Breusch-Godfrey: a) [AR(1)] P-val: 0,304; b) [AR(2)] P-val: 0,159; Ljung-Box (p=6): P-val.= 0,331; ARCH: a) [ARCH(1)] P-val: 0,629; a) [ARCH(2)] P-val: 0,253; White: P-val. = 0,521; *SupF* = 14,859; *MeanF* = 7,278; *ExpF* = 5,183.
MANUFACTURING INDUSTRY
ESTIMATED EQUATION (t = 1973,...,1993)
$DLVIEX_t = -0{,}061 - 0{,}555\ _{t-1} + 2{,}262\ DLVNOAEEX_t + 0{,}231\ DLVIEX_{t-1} + 0{,}171\ DLVAEX_t$ (-2,830)(-2,337) (5,103) (1,968) (2,107) $+ 0{,}111\ DF80_t + 0{,}128\ DD8285_t + 0{,}142\ F77_t + 0{,}152\ F85_t$ (2,352) (2,857) (2,084) (2,452)
SPECIFICATION TESTS
$R^2 = 0{,}873$; Adjusted $R^2 = 0{,}788$; Durbin-Watson: DW = 2,051; Jarque-Bera: P-val. = 0,564; Breusch-Godfrey: a) [AR(1)] P-val: 0,506; b) [AR(2)] P-val: 0,799; Ljung-Box (p=6): P-val.= 0,942; ARCH: a) [ARCH(1)] P-val: 0,308; a) [ARCH(2)] P-val: 0,446; White: P-val. = 0,611; *SupF* = 14,247; *MeanF* = 8,141*; *ExpF* = 5,723*.

SALES-ORIENTED SERVICES
ESTIMATED EQUATION (t = 1973,...,1993)
$DLVLEX_t = 0{,}006 - 0{,}836\ _{t-1} + 0{,}420\ DLVNOAE_t + 0{,}924\ DLVLES_t - 0{,}059\ F77_t$ (0,815)(-3,801) (2,253) (3,617) (-3,440) $- 0{,}054\ D8085_t - 0{,}036\ F90_t$ (-5,665) (-2,134)
SPECIFICATION TESTS
R^2 = 0,938; Adjusted R^2 = 0,911; Durbin-Watson: DW = 2,365; Jarque-Bera: P-val. = 0,542; Breusch-Godfrey: a) [AR(1)] P-val: 0,166; b) [AR(2)] P-val: 0,015; Ljung-Box (p=6): P-val.= 0,324; ARCH: a) [ARCH(1)] P-val: 0,957; a) [ARCH(2)] P-val: 0,166; White: P-val. = 0,157; *SupF* = 3,350; *MeanF* = 2,059; *ExpF* = 1,132.
TRANSPORT AND COMUNICATIONS
ESTIMATED EQUATION (t = 1974,...,1993)
$DLVZEX_t = 0{,}008 - 0{,}690\ _{t-1} + 0{,}222\ DLVLEX_t + 0{,}348\ DLVZES_t + 0{,}245\ DLVZEX_{t-1}$ (1,123)(-4,123) (3,054) (1,967) (3,142) $+20{,}551\ DD8084_t - 0{,}082\ DD8889_t - 1{,}435\ D(D8084LVZES)_t$ (7,676) (-7,614) (-7,658)
SPECIFICATION TESTS
R^2 = 0,944; Adjusted R^2 = 0,912; Durbin-Watson: DW = 1,756; Jarque-Bera: P-val. = 0,865; Breusch-Godfrey: a) [AR(1)] P-val: 0,741; b) [AR(2)] P-val: 0,184; Ljung-Box (p=6): P-val.= 0,781; ARCH: a) [ARCH(1)] P-val: 0,147; a) [ARCH(2)] P-val: 0,159; White: P-val. = 0,540; *SupF* = 5,486; *MeanF* = 2,539; *ExpF* = 1,580.
NON-SALES-ORIENTED SERVICES
ESTIMATED EQUATION (t = 1972,...,1993)
$DLVGEX_t = -0{,}001 - 0{,}701\ _{t-1} + 0{,}974\ DLVGES_t + 2{,}796\ DD7078_t - 0{,}205\ D(D7078LVGES)_t$ (-0.194)(-7,689) (13,952) (4,335) (-4,797) $- 17{,}353\ DD7985_t + 1{,}141\ D(D7985LVGES)_t + 0{,}027\ DD9293_t + 0{,}015\ F80_t - 0{,}040\ F82_t$ (-17,291) (14,593) (6,191) (3,112) (-7,621) $- 0{,}041\ F84_t + 0{,}021\ F88_t$ (-8,694) (4,808)
SPECIFICATION TESTS
R^2 = 0,991; Adjusted R^2 = 0,982; Durbin-Watson: DW = 2,345; Jarque-Bera: P-val. = 0,645; Breusch-Godfrey: a) [AR(1)] P-val: 0,191; b) [AR(2)] P-val: 0,012; Ljung-Box (p=6): P-val.= 0,209; ARCH: a) [ARCH(1)] P-val: 0,076; a) [ARCH(2)] P-val: 0,167; White: P-val. = 0,154; *SupF* = 6,164; *MeanF* = 3,407; *ExpF* = 2,025.

Notes:
- In parentheses, beneath the estimated coefficients, appear the t** statistics.
- * denotes rejection of the null hypothesis at 10%, ** at 5%, and *** at 1%.

TABLE 4: Statistics on the forecasting performance of the models in the ex-post simulation

MODEL	ME	MAE	RMSE	U
Agriculture				
Q=0	-0,065	0,065	0,075	0,587
Q=I	0,063	0,063	0,085	0,667
Q diag. estimated	-0,060	0,066	0,089	0,698
Energy				
Q=0	0,010	0,042	0,043	3,165
Q=I	-0,026	0,045	0,052	3,800
Q diag. estimated	-0,084	0,084	0,109	7,932
Construction				
Q=0	-0,016	0,016	0,017	0,898
Q=I	-0,021	0,021	0,021	1,128
Q diag. estimated	0,012	0,013	0,014	0,723
Manufacturing Industry				
Q=0	0,026	0,026	0,028	1,075
Q=I	0,012	0,024	0,027	1,012
Q diag. estimated	0,009	0,009	0,013	0,494
Sales-oriented Services				
Q=0	0,011	0,011	0,013	0,664
Q=I	0,016	0,016	0,017	0,874
Q diag. estimated	0,041	0,041	0,042	2,152
Transport y comunications				
Q=0	0,006	0,017	0,018	0,569
Q=I	-0,002	0,017	0,017	0,549
Q diag. estimated	-0,040	0,040	0,048	1,547
Non-sales-oriented Services				
Q=0	-0,001	0,003	0,003	0,135
Q=I	-0,003	0,003	0,004	0,172
Q diag. estimated	-0,021	0,021	0,026	1,127

Notes:
ME= mean error; MAE=mean absolute error; RMSE=root means square error; U= coefficient of inequality of Theil (1996).

REFERENCES

Andrés J.A., Escribano C., Molinas C. and Taguas D., 1990, *La inversión en España: Modelización con Restricciones de Equilibrio* (Barcelona-Madrid: Antoni Bosch and Instituto de Estudios Fiscales).

Andrews D., 1993, Tests for Parameter Instability and Structural Change With Unknown Change Point, *Econometrica*, 61, 821-856.

Andrews D. and Ploberger W., 1994, Optimal Tests When a Nuisance Parameter Is Present Only Under the Alternative, *Econometrica*, 62, 1383-1414.

Ansley C. and Kohn R., 1983, State Space Models and Diffuse Initial Conditions, I: Filtering and Likelihood, Technical Report 13, University of Chicago, Statistics Research Center, Graduate School of Business.

Campos J., Ericsson N. and Hendry D.F., 1996, Cointegration test in the presence of structural breaks, *Journal of Econometrics*, 70, 187-220.

Canarella G., Pollard S. and Lai K., 1990, Cointegration between exchange rates and relative price: another view, *European Economic Review*, 34, 1303-1322.

Chow G., 1984, Random and Changing Coefficient Models, in: Z. Griliches and M. Intriligator, eds., *Handbook of Econometrics* (Vol. II) (Amsterdam: North-Holland), 1213-1245.

Dickey D. and Fuller W., 1979, Distribution of the estimators for autoregressive time series with a unit root, *Journal of the American Statistical Association*, 84, 427-431.

Dickey D. and Fuller W. , 1981, Likelihood ratio statistics for autoregressive time series with a unit root, *Econometrica*, 50, 1057-1072.

Engle R. and Granger C., 1987, Co-integration and error correction: representation, estimation and testing, *Econometrica*, 55, 251-276.

Gregory A. and Hansen B., 1996, Residual-based tests for cointegration in models with regime shifts, *Journal of Econometrics*, 70, 99-126.

Hackl P. and Westlund A., 1996, Demand for international telecommunication. Time-varying price elasticity, *Journal of Econometrics*, 70, 243-260.

Hall S., 1994, Modelling economies in transition, Centre for Economic Forecasting, London Business School.

Hansen B., 1992, Tests for Parameter Instability in Regressions with I(1) Processes, *Journal of Business and Economic Statistics*, 10, 321-336.

Harvey A., 1981, *The Econometric Analysis of Time Series* (Oxford: Philip Allan Publishers Limited).

Harvey A., 1989, *Forecasting, Structural Time Series Models, and the Kalman Filter* (Cambridge: Cambridge University Press).

Hendry D.F. and Neale A., 1991, A Monte Carlo study of the effects of structural breaks on tests for unit roots, in: P. Hackl and A. Westlund, eds., *Economic structural change: analysis and forecasting* (Berlin: Springer-Verlag), 95-119.

Hispalink, 1993, Banco de datos multirregional HISPALINK (Madrid: Consejo Superior de Cámaras de Comercio, Industria y Navegación de España).

Hispalink, 1997, HISPADAT. Banco de datos (Madrid: Instituto L.R. Klein and Consejo Superior de Cámaras de Comercio, Industria y Navegación de España).

Kalman R., 1960, A new approach to linear filtering and prediction problems, *ASME Journal of Basic Engineering*, 82, 35-45.

Lin C-F. and Teräsvirta T., 1994, Testing the constancy of regression parameters against continuous structural change, *Journal of Econometrics*, 62, 211-228.

Lucas R.E., 1976, Econometric Policy Evaluation: A Critique, *Carnegie-Rochester Conference Series on Public Policy*, 1, 19-46.

Lütkepohl H., 1993, *Introduction to Multiple Time Series Analysis* (New York: Springer-

Verlag).

McKinnon J., 1991, Critical values for cointegration tests in long-run economic relationships, en: R. Engle and C. Granger, eds., *Readings in Cointegration* (New York: Oxford University Press), 267-276.

Molinas C., Sebastián M. and Zabalza A.,1991, *La economía española: una perspectiva macroeconómica* (Barcelona-Madrid: Antoni Bosch and Instituto de Estudios Fiscales).

Otero J.M., Isla F., Trujillo F., Fernández A. and López P., 1996, Modelización económica regional: el proyecto Hispalink-Andalucía, *Boletín Económico de Andalucía*, 21, 49-66.

Perron P., 1989, The great crash, the oil price shock, and the unit root hypothesis, *Econometrica*, 57, 1361-1401.

Quandt R., 1960, Tests of the Hypothesis That a Linear Regression System Obeys Two Separate Regimes, *Journal of the American Statistical Association*, 55, 324-330.

Quintos C. and Phillips P., 1993, Parameter constancy in cointegrating regresssions, *Empirical Ecoonomics*, 18, 675-703.

Ramajo J. and Márquez M.A., 1996, Elaboración de indicadores sintéticos para el seguimiento de la coyuntura económica de Extremadura, Monográfico de la Consejería de Economía, Industria y Hacienda, Junta de Extremadura.

Rappoport P. and Reichlin L., 1989, Segmented trends and non-stationary time series, *The Economic Journal*, 99, 168-177.

Stock J. and Watson M., 1996, Evidence on structural instability in macroeconomic time series relations, *Journal of Business and Economic Statistics*, 14, 11-30.

Theil H., 1966, *Applied Economic Forecasting* (Amsterdam: North-Holland).

Treyz G., 1993, *Regional economic modeling. A Systematic Approach to Economic Forecasting and Policy* (Boston: Kluwer Academic Publishers).

Wolf C., 1987, Time-Varying Parameters and the Out-of-Sample Forecasting Performance of Structural Exchange Rate Models, *Journal of Business and Economic Statistics*, 5, 87-97.

Wolters J., Teräsvirta T. and Lütkepohl H., 1996, Modelling the Demand for M3 in the Unified Germany, Discussion Paper 23, Humboldt Universität zu Berlin.