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Modelling Economies in Transition: An Introduction

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ABSTRACT

This paper considers the implications of structural breaks, such as have occurred in many transition economies, for econometric modelling based on the multivariate cointegration paradigm. It outlines recent developments on the identification of linear cointegrated systems, discusses some practical problems, and presents an extension to non-linear systems. This is followed by a discussion of the impact of structural breaks on the identification and estimation of such systems. Finally, it relates these issues to the other papers in this volume.

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1 Introduction

The economies of Europe have experienced significant changes in economic structure and economic policies pursued during the last three decades. Some economies have undergone substantial liberalisation of their labour, financial, and foreign exchange markets, including the privatisation of former state owned companies, with an example being the UK. Other economies of eastern and central Europe, for example Bulgaria, Poland, and the Czech Republic have moved from being centrally planned towards free market economies. Intermediate between these extremes are economies that have slowly adopted policies to liberalise their financial and foreign exchange markets, and introduce some degree of flexibility into their labour markets e.g. Italy and Spain. Any attempt to model sectors of these economies is likely to have to face the issue of how best to deal with these changes. Although the nature of the particular changes will influence the choice of model class to represent the relevant macroeconomic time series, the purpose of the modelling will also help determine an appropriate model choice. For example, Hendry and Mizon (1999) demonstrate that the best forecasting model in a non-stationary world is unlikely to be of value for economic policy analysis, with the converse being true when there are structural changes in sample. However, the reduced rank vector autoregression or cointegrated VAR (see e.g., Johansen, 1995b) has proved to be effective in the modelling of non-stationary macroeconomic time series, and it this class of model that we discuss and which is most commonly used in the other papers in this volume.

The plan of the paper is as follows. In section 2 we discuss the general identification problem in cointegrated linear systems, and consider the extension to non-linear systems, as well as some practical problems that arise in empirical implementation. Section 3 considers the situation in which non-stationarities arise from structural breaks and regime shifts as well as from unit roots. Consideration is given to the nature of these deterministic breaks, their detectability, and alternative ways to model them. Section 4 provides an overview of the papers in this volume, and the way they relate to each other and the overall methodology we have adopted. Section 5 contains general conclusions.

2 Economic Interpretation and Identification in Cointegrated VARs

Time series econometrics has been revolutionised by the developments in multivariate cointegration over the last fifteen years. Much of this advance has been achieved by the development of the statistical theory relevant for the analysis of cointegrated systems (see e.g., Johansen, 1995b). The issue of identification and economic interpretation of the parameters in such systems, though important, was not central to the early development of this theory. However, the Granger Representation Theorem (see Engle and Granger, 1987) by establishing that the vector autoregressive model (VAR) and the vector equilibrium correction model (VEqCM) are observationally equivalent, helped to bridge the gap between the statistical model and an economic interpretation of it. Subsequently much more attention has been paid to the development of models that are economically interpretable simplifications of the VAR. One approach is that which adopts the 'structural' VAR (SVAR) using economically interpretable restrictions to achieve identification, with further restrictions being testable as over-identifying hypotheses (see e.g., Davidson and Hall 1991, and for recent reviews Canova, 1995 and Pesaran and Smith, 1998). An alternative approach, developed in Hendry and Mizon 1993, exploits the fact that a statistically well-specified full-rank VAR is identified, and provides a valid basis against which to test simplification hypotheses such as those that arise from economic theory considerations and from the empirical evidence. We next introduce the notation to analyse the VAR, VEqCM, and the SVAR, and discuss identification in these models. However, before doing so we note that identification, especially as used in economics, has a number of different connotations, which it is important to distinguish. The primary meaning is associated with the global uniqueness of the maximum likelihood estimator for a particular parametric statistical model. Such a unique parameter estimator may not correspond precisely to a particular economic theory, and so another meaning concerns the establishment of a direct connection between the parameters and an economic theory - this provides an interpretation for the parameters. Identification is also used to mean the selection of one model rather than another, and thus implies the isolation or selection of the appropriate model for some purpose - the discussion of identification for ARMA models in Box and Jenkins (1976) provides an example.

2.1 Closed Linear Systems

For p lags on a vector of N variables $\mathbf{z}_t$ the closed VAR is:

$$\mathbf{z}_t = \sum_{j=1}^{p} \mathbf{D}_j \mathbf{z}_{t-j} + \delta + \varepsilon_t \quad \text{with} \quad \varepsilon_t \sim \mathsf{IN}_N(\mathbf{0}, \Sigma), \tag{1}$$

where $\mathbf{D}_j$ is an $N \times N$ matrix of autoregressive coefficients, and ε_t is a vector of N unobserved errors, which have a zero mean and constant covariance matrix Σ. Independently of whether the variables $\mathbf{z}_t$ are $\mathrm{I}(0)$ or $\mathrm{I}(1)$ the VAR in (1) can be re-parameterised as a VEqCM (see Johansen 1988, 1992c, and Hendry 1995):

$$\Delta\mathbf{z}_t = \sum_{j=1}^{p-1} \Gamma_j \Delta\mathbf{z}_{t-j} + \Pi\mathbf{z}_{t-1} + \delta + \varepsilon_t \tag{2}$$

where Δ is the first difference operator, $\Gamma_j = -\sum_{i=j+1}^{p} \mathbf{D}_i \quad (j = 1,2,\ldots p-1)$ are the short-run adjustment coefficient matrices and $\Pi = -(\mathbf{I}_N - \sum_{i=1}^{p} \mathbf{D}_i)$ is the long-run coefficient matrix. We note that although the model has been defined with an $N \times 1$ vector of intercepts δ, often there will be other deterministic variables included. For example, the deterministic variables might be $\delta + \alpha\lambda t + \mathbf{K}\mathbf{d}_t$ when δ is an intercept, $\alpha\lambda t$ a linear trend restricted to the cointegration space (trend is so restricted since few variables exhibit quadratic trend), and $\mathbf{d}_t$ some event specific dummy variables.

When Π has full rank N the variables $\mathbf{z}_t$ are $\mathrm{I}(0)$ and the parameters $(\Gamma_1, \ldots \Gamma_{p-1}, \Pi, \delta, \Sigma)$, or equivalently $(\mathbf{D}_1, \ldots \mathbf{D}_p, \delta, \Sigma)$, are all identified in that the maximum likelihood estimator of these parameters is unique. Since (1) and (2) are re-parameterisations of each other they are observationally equivalent, and the choice between them can be made on the basis of their interpretation. Indeed, an attraction of the parameterisation in (2) is the interpretation of its static long-run solution, $\mathsf{E}(\Pi\mathbf{z}_t + \delta) = 0$ as the equilibria of the system, with $(\Pi\mathbf{z}_t + \delta)$ being the disequilibria at time t. Economic theory is often informative about such equilibria. The short-run adjustment parameters Γ_j are also the subject of economic theory considerations concerning the time form of responses and speed of adjustment, though these are typically less precise than the hypotheses concerning equilibria. However, the parameters of interest (ϕ say) may not be those of (1) or (2), and so the identification and estimation of ϕ has to be considered separately. In the SVAR approach to modelling the parameters of interest ϕ are a function of $(\mathbf{A}_0, \mathbf{A}_1, \ldots \mathbf{A}_p, \mathbf{c}, \mathbf{\Phi})$ in:

$$\mathbf{A}_0\mathbf{z}_t = \sum_{j=1}^{p} \mathbf{A}_j\mathbf{z}_{t-j} + \mathbf{c} + \mathbf{u}_t \quad \text{with} \quad \mathbf{u}_t \sim \mathrm{IN}_N(\mathbf{0}, \mathbf{\Phi}), \tag{3}$$

When $\mathbf{A}_0$ is an $N{\times}N$ non-singular matrix, $\mathbf{A}_0\mathbf{z}_t$ represents N linear combinations of the N variables in $\mathbf{z}_t$ that characterise their determination in the context of some economic theory, especially those in which the simultaneous determination of $\mathbf{z}_t$ is a feature. It is well known that $(\mathbf{A}_0, \mathbf{A}_1, \ldots \mathbf{A}_p, \mathbf{c}, \mathbf{\Phi})$ are not identified. In the class of linear dynamic systems the VAR in (1), or equivalently the VEqCM in (2), characterise the distribution $\mathbf{z}_t | \mathbf{Z}_{t-1}$ and thus provide the reduced form of (3). This implies that:

$$\mathbf{D}_j = \mathbf{A}_0^{-1}\mathbf{A}_j \ (j = 1,2,\ldots p), \ \boldsymbol{\delta} = \mathbf{A}_0^{-1}\mathbf{c}, \text{ and } \mathbf{\Sigma} = \mathbf{A}_0^{-1}\mathbf{\Phi}(\mathbf{A}_0^{-1})', \tag{4}$$

and leads to the conventional discussion of identification in simultaneous equations models in which restrictions on $(\mathbf{A}_0, \mathbf{A}_1, \ldots \mathbf{A}_p, \mathbf{c}, \mathbf{\Phi})$ are required for (4) to have a unique solution for $(\mathbf{A}_0, \mathbf{A}_1, \ldots \mathbf{A}_p, \mathbf{c}, \mathbf{\Phi})$ in terms of $(\mathbf{D}_1, \ldots \mathbf{D}_p, \boldsymbol{\delta}, \mathbf{\Sigma})$ or $(\mathbf{\Gamma}_1, \ldots \mathbf{\Gamma}_{p-1}, \mathbf{\Pi}, \boldsymbol{\delta}, \mathbf{\Sigma})$, (see e.g., Johnston, 1972, Greene, 1991). All sets of restrictions on $(\mathbf{A}_0, \mathbf{A}_1, \ldots \mathbf{A}_p, \mathbf{c}, \mathbf{\Phi})$ that achieve just-identification are observationally equivalent; each being observationally equivalent to the unrestricted reduced form parameterised via $(\mathbf{D}_1, \ldots \mathbf{D}_p, \boldsymbol{\delta}, \mathbf{\Sigma})$ or $(\mathbf{\Gamma}_1, \ldots \mathbf{\Gamma}_{p-1}, \mathbf{\Pi}, \boldsymbol{\delta}, \mathbf{\Sigma})$. Hence the possibility of discriminating between alternative SVAR's on the basis of empirical evidence only arises when they are over-identified; in which case the over-identifying restrictions are testable against the unrestricted reduced form. We note that when the variables contained in $\mathbf{z}_t$ are chosen carefully (perhaps after transformation) to ensure that the parameters of the (1) are such that the parameters of (economic) interest are recoverable from them, and that the VAR is congruent, then the approach to model simplification in Hendry and Mizon

(1993) operates with identified parameters throughout and results in the selected model having unique, interpretable, and relevant parameters.

A further identification problem arises when the variables being modelled are $I(1)$, but satisfy $r < N$ cointegrating relationships $\beta'\mathbf{z}_t$ that are $I(0)$. This is often the case for macroeconomic time series, as is illustrated in many of the papers in this volume. In this case the rank of Π is r a feature that can be incorporated into the model by defining $\Pi = \alpha\beta'$ with α and β being $N\times r$ matrices of rank r thus leading to the reduced rank or cointegrated VEqCM:

$$\Delta\mathbf{z}_t = \sum_{j=1}^{p-1} \Gamma_j\, \Delta\mathbf{z}_{t-j} + \alpha\beta'\mathbf{z}_{t-1} + \delta + \varepsilon_t \tag{5}$$

Note though that α and β are not identified since $\alpha\beta' = \alpha^{+}\beta^{+\prime} = \alpha\mathbf{P}\mathbf{P}^{-1}\beta'$ for any non-singular $r\times r$ matrix $\mathbf{P}$ (rotation). Hence in the reduced rank case, with the reduced rank imposed, neither the VAR in (1) nor the VEqCM in (2) is identified. In particular, it is necessary to determine r, and identify α and β, and there are many routes in which this might be achieved in practice - see Figure 1 in Greenslade, Hall and Henry (1998) for a diagrammatic representation of the possibilities.

Since r is not known *a priori* its value has to be determined empirically, and this provides one possible starting point. A commonly adopted procedure is the maximum likelihood one developed by Johansen (1988), which employs likelihood ratio criteria for determining r, and for a given choice of r yields a unique estimate of Π of rank r. Since the short-run adjustment coefficients Γ_j $(j = 1,2,\ldots p-1)$ and the error covariance matrix Σ are identified, unique unrestricted maximum likelihood estimates of these parameters are available for a given value of r. The Johansen procedure also produces unique estimates of α and β satisfying $\Pi = \alpha\beta'$ as a result of imposing the restriction that the resulting β be orthogonal (they are in fact eigenvectors for a canonical correlation problem). Important features of these β estimates include: spanning the space of all just-identified cointegrating vectors, being neutral about normalization, and not having any zero restrictions on the elements of β. However, it is likely that such β will be in a form that does not have a clear economic interpretation. Thus although identification associates a unique point in parameter space with maximisation of the likelihood, relevance and interpretability of the resulting estimates are important as well (see section 2.5 of Hendry, 1995a). In order to achieve relevance and interpretability for the phenomenon being analysed alternative sets of parameter restrictions can be considered, and an important source of information for these restrictions is economic theory related to the aspect of the economy under study. For example, there are usually economic theories, which imply the existence of long run relationships or equilibria amongst the variables $\mathbf{z}_t$, $\mathsf{E}(\beta^{*\prime}\mathbf{z}_t) = \mu$ say, with $(\beta^{*\prime}\mathbf{z}_t - \mu)$ being interpretable as the disequilibria at time t. In so far as these equilibria consist of r linear combinations of the variables in $\mathbf{z}_t$ with sufficient restrictions on the Nr elements of β to identify them, they can be represented in the Johansen framework, with any over-identifying restrictions (i.e., restrictions beyond the r^2 required for normalization and just-identification) testable using a standard likelihood ratio test statistic as implemented for example in CATS and PcFiml (see Hansen and Juselius, 1994 and Doornik and Hendry, 1997 respectively). Equally hypotheses concerning the long run weak exogeneity of some variables within $\mathbf{z}_t$ for the coefficients in the long run equilibria β^* correspond to particular zero restrictions on the long run adjustment coefficients α (see e.g., Johansen 1992b).

Phillips (1991) presented an alternative approach to identifying β by partitioning the set of variables $\mathbf{z}_t$ into $\mathbf{z}_{1,t}$ and $\mathbf{z}_{2,t}$, and giving β the form $(-\mathrm{I}_r, \mathbf{B})$ to yield a block recursive structure:

$$\begin{aligned} \mathbf{z}_{1,t} &= \mathbf{B}\mathbf{z}_{2,t} + \mathbf{v}_{1,t} \\ \Delta \mathbf{z}_{2,t} &= \mathbf{v}_{2,t} \end{aligned} \qquad (6)$$

in which $\mathbf{z}_{1,t}$ is a vector of r variables, $\mathbf{z}_{2,t}$ is an $(N-r)$ vector, and $\mathbf{v}_{1,t}$ and $\mathbf{v}_{1,t}$ are independent I(0) processes which in general are temporally dependent. The assumptions underlying (6) are sufficient to exactly identify the system, but this is achieved by assuming that r is known, by imposing a very restricted block-recursive structure which may not often correspond to a relevant economic theory, and the outcome is not invariant to the ordering of the variables within $\mathbf{z}_t$. Saikkonen (1993) discusses the complete identification of a VEqCM that has a similar partition of $\mathbf{z}_t$ into

$\mathbf{z}_{1,t}$ and $\mathbf{z}_{2,t}$ to that in the Phillips (1991) system. Note though that **for a given value of** r both the Johansen and Phillips approaches to exactly identifying β are observationally equivalent, so that empirical evidence cannot distinguish between them. It is often argued that this is precisely when economic theory and economic interpretation can help, and provides motivation for the SVAR approach. The question of identifying the parameters of a SVAR which has $r < N$ cointegrating vectors has received increased attention, and there is now a reasonably complete understanding of the process of identifying a structural cointegrated system - see inter alia Johansen (1994), Johansen (1995a), and Robertson and Wickens (1994). In such cases (3), written in VEqCM format, becomes:

$$\mathbf{A}_0 \Delta \mathbf{z}_t = \sum_{j=1}^{p-1} \mathbf{C}_j \, \Delta \mathbf{z}_{t-j} + \mathbf{A}^{*} \beta' \mathbf{z}_{t-1} + \mathbf{c} + \mathbf{u}_t \qquad (7)$$

with $\mathbf{C}_j = \mathbf{A}_0 \Gamma_j \quad (j = 1,2,\ldots p-1)$ and $\mathbf{A}^{*} = \mathbf{A}_0 \alpha$. Conditional on having chosen the cointegrating rank r it is necessary to consider the identification of the contemporaneous coefficients $\mathbf{A}_0$ and the long run coefficients β, and these are essentially separate issues in that there are no mathematical links between restrictions on $\mathbf{A}_0$ and those on β. In particular, since a Π matrix of rank r is identified and satisfies $\Pi = \alpha\beta' = \mathbf{A}_0^{-1} \mathbf{A}^{*} \beta'$, it follows that restrictions are required to identify β even if $\mathbf{A}_0$ were known. Conversely, restrictions on β have no mathematical implication for the restrictions on $\mathbf{A}_0$. It remains possible though that the economic interpretation of a restricted set of cointegrating vectors $\beta' \mathbf{z}_t$ may have implications for the nature of restrictions on $\mathbf{A}_0$ that will be economically interesting, particularly when $\mathbf{A}^{*}$ is restricted via α. Mathematical, and possibly economic, linkages do exist between restrictions on the adjustment coefficients α and those required to identify β - see Doornik and Hendry (1997).

The formal identification of β is the main subject of Johansen and Juselius (1992) and Pesaran and Shin (1997) where it was demonstrated that a necessary condition for exact identification is that there are $k = r^2$ restrictions. Johansen (1995a) and Pesaran and Shin (1997) also give a necessary and sufficient rank condition for exact identification, which for example rules out dependence amongst the r^2 restrictions. In general if the number of available restrictions is $k < r^2$ the system is under-identified, if $k = r^2$ the system is exactly identified, and when $k > r^2$ the system is over-identified, and subject to the rank condition being satisfied the over-identifying restrictions are testable.

Whatever the restrictions are that achieve exact identification of $\mathbf{A}_0$ and β, for a given value of r the resulting just-identified SVAR in (7) is observationally equivalent to the cointegrated VEqCM of (5), and to the triangular system of (6) for the same ordering of variables in $\mathbf{z}_t$. In other words any of these approaches could be used to obtain a value for the unrestricted likelihood function for the rank r cointegrated VAR, which then forms the basis for comparison with restricted likelihood values corresponding to over-identified systems. Given that all just-identified systems are observationally equivalent it is important to realise that in specifying an over-identified system it is not necessary to maintain any particular set of just-identifying restrictions. For example, although PcFiml uses the Johansen procedure to obtain estimates of just-identified β and the corresponding unrestricted likelihood value, when asking for the specification of an over-identified model to be estimated and its over-identifying restrictions tested, it does not impose orthogonality between the columns of β even though that was imposed in obtaining the just-identified estimates of β. This contrasts with the procedure of Phillips (1991) which maintains the just-identifying restrictions of the triangular representation when testing over-identifying restrictions. Hence if the particular triangular representation for a given ordering of variables in $\mathbf{z}_t$ is not data coherent when used in conjunction with the over-identifying restrictions the latter are likely to be rejected. Similarly, if there are inappropriate restrictions (particularly zero restrictions) imposed on the elements of β within a SVAR to just-identify it, then this may cause problems in the subsequent analysis of over-identifying restrictions if the just-identifying restrictions are maintained. The just-identified cointegrated VEqCM of (5), when estimated by the Johansen procedure, on the other hand is invariant to normalization, has imposed no zero restrictions on the elements of β, and is a suitable benchmark against which to test any over-identifying restrictions.

2.2 Open Linear Systems

In cases where the vector $\mathbf{z}_t$ is partitioned into $\mathbf{z}'_t = (\mathbf{y}'_t, \mathbf{x}'_t)$ with $\mathbf{y}_t$ an $M\times 1$ vector of endogenous variables and $\mathbf{x}_t$ is a $Q\times 1$ vector of exogenous variables $(N = M + Q)$, the closed SVAR will be replaced by an open SVAR that does not model, but conditions on, the exogenous variables $\mathbf{x}_t$. Inference based on such an open SVAR will be appropriate when the conditioning variables $\mathbf{x}_t$ are weakly exogenous for the parameters of interest (see Engle, Hendry and Richard 1983), and this is a testable hypothesis relative to the VAR in (1). Note though that there is no generic test for weak exogeneity, but the hypothesis of long-run weak exogeneity is readily testable via zero restrictions on α (see e.g., Hendry and Mizon, 1993). In addition, when N is large (e.g., $N > 8$) it can be extremely difficult to successfully model all N variables, and it will often be convenient to condition on a subset of them, particularly those that satisfy economic concepts of exogeneity (as opposed to being weakly exogenous) and are the most difficult to model - this is often true of variables which characterise economic policy or are instruments used in implementing policy such as tax rates and *repo* interest rates.

2.3 Non-linear Systems

In this section we discuss a natural extension of the identification conditions mentioned above to the case of non-linear cointegrated systems. This is particularly relevant given that some of the work in the PHARE project has adopted highly non-linear model formulations.

Wegge (1965), Fisher (1966), Rothenberg (1971), Bowden (1973), and Richmond (1976) have studied the problem of identification in the non-linear case in full rank systems. The basic approach is simply to require sufficient restrictions on the general model to ensure that the information matrix has full rank. Thus given our definition of $\mathbf{z}_t$ and defining θ to be the vector of all parameters in the model we may define the joint density of $Z_T = (\mathbf{z}_1, \mathbf{z}_2, \ldots \mathbf{z}_T)'$ to be $f(Z_T | \theta)$. Then the information matrix may be defined as $I(\theta) = -E(\partial^2 \ln f/\partial\theta\partial\theta')$. Further, let $\mathbf{g}(\theta)$ be a set of q known functions of θ with continuous partial derivatives, which are such that $\mathbf{g}(\theta_0) = 0$ holds for the true value θ_0 of θ, and define $G(\theta) = \partial g/\partial\theta'$ as the

Jacobian matrix. Then following Rothenberg (1971), further define $\mathbf{V}(\theta)' = (\mathrm{I}(\theta)', G(\theta)')$ and enables the following theorem to be stated.

Theorem

Suppose θ^+ is a regular point of both $G(\theta)$ and $\mathbf{V}(\theta)$. Then θ^+ is locally identifiable if and only if $\mathbf{V}(\theta^+)$ has full column rank N.

Note that this can only establish local identification in the non-linear case. However there is a simpler and useful special case in which the adjustment parameters are linear but the long-run relationships are non-linear (Escribano, 1985, contains an early analysis of a model of this type). In this case when there are r cointegrating relationships, but they are non non-linear such that $\mathbf{w}_t = \mathbf{h}(\beta, \mathbf{z}_t)$ is an $\mathrm{I}(0)$ process, and all the adjustment is linear, we can rewrite (7) as:

$$\mathbf{A}_0 \Delta \mathbf{z}_t = \sum_{j=1}^{p-1} \mathbf{C}_j \Delta \mathbf{z}_{t-j} + \mathbf{A}^* \mathbf{w}_{t-1} + \mathbf{c} + \mathbf{u}_t \, .$$

It is then possible to concentrate the likelihood function with respect to β, and identification may be undertaken in two separate stages. This is a useful case since theory is often informative about long-run relationships or equilibria which are non-linear, but less informative about the form of adjustment and so this may often be left in linear form (this is exactly the form of model used by Hall and Nixon in this volume).

2.4 Testing Over-identifying Restrictions

Whether the restrictions considered are motivated by economic theory, or as empirical simplifications that do not contradict economic theory, the over-identifying restrictions are testable using standard likelihood ratio test statistics. Indeed, once the cointegrating rank of the system r has been determined and α and β identified, hypotheses concerning the short-run adjustment coefficients $\Gamma_j \ (j = 1,2,\ldots p-1)$, the long-run adjustment coefficients α, and the coefficients of the long-run equilibria β, are all testable using standard likelihood ratio test statistics. Based on results in Phillips (1991) and Johansen (1991) showing that these likelihood ratio test statistics have limiting central χ^2 null distributions with degrees of freedom equal to the number of over-identifying restrictions, Pesaran and Shin (1997) demonstrated that this result applies for non-linear as well as linear restrictions. There is however an important question as to the performance of these tests in finite samples, which has been recently addressed in Greenslade *et al.* (1998). These authors point out that in small samples the order in which restrictions are imposed on the VEqCM may have major effects on both the power and the size of the tests for particular hypotheses. In particular, Greenslade *et al.* (1998) argue that imposing long-run weak exogeneity restrictions and restricting the short-run adjustment coefficients can improve the size of the tests on the over-identifying restrictions of the long-run coefficients β enormously. They offer a set of Monte Carlo experiments to support this assertion. Johansen (1999) has also shown that Bartlett adjusted likelihood ratio statistics have improved finite sample performance.

2.5 Over-identification, Structural Breaks, and Observational Equivalence

Although it is well known that there are many models that are just-identified and observationally equivalent to the unrestricted reduced form of the system, it is less widely appreciated that for a given number of over-identifying restrictions k there is in general a set of models each of which is observationally equivalent to the restricted reduced form of the system (see Hendry and Mizon, 1993 for further discussion of this point). Although there can be two or more observationally equivalent over-identified models of degree k the mechanism that generates the data is unique, and so there will be at most one model that directly characterises this and is invariant to extensions of information. Such a model was defined in Hendry (1995) as capturing the structure of the economy. Hence, the presence of structural breaks in some models implies that they cannot be structural. Thus the existence of structural breaks can provide a very powerful means of discriminating between otherwise equivalent models, and perhaps discovering structure. Seen in this light structural breaks provide valuable information, rather than being simply a nuisance. Lu and Mizon (1999) discuss this point in more detail, and provide some illustrations using simulated data. In the next section we discuss: the nature of structural breaks and regime shifts, methods for detecting their presence, and alternative approaches to developing models that are invariant.

3 Structural Breaks and Regime Shifts in Cointegrated VARs

The model in (5) is in $I(0)$ space when correctly formulated, thus inference concerning its parameters $(\Gamma_1, \ldots \Gamma_{p-1}, \Pi, \alpha, \beta, \Sigma)$ can be conducted using conventional procedures. It thus provides a convenient framework within which to consider structural breaks and regime shifts. Consider a steady state in which $\mathsf{E}(\Delta \mathbf{z}_t) = \gamma$ and $\mathsf{E}(\beta' \mathbf{z}_t) = \mu + \lambda t \; \forall t$, with δ is replaced by $(\Xi\gamma - \alpha\mu - \alpha\lambda(t-1) + \mathbf{K}\mathbf{d}_t)$, when $\beta'\gamma = 0$ for identification, $\Xi = (\mathbf{I}_N - \sum_{j=1}^{p-1} \Gamma_i)$, $\alpha\lambda t$ is a linear deterministic trend restricted to lie in the cointegrating space, and $\mathbf{d}_t$ contains event specific dummy variables. Hence (5) can be re-written as:

$$(\Delta \mathbf{z}_t - \gamma) = \sum_{j=1}^{p-1} \Gamma_j (\Delta \mathbf{z}_{t-j} - \gamma) + \alpha(\beta' \mathbf{z}_{t-1} - \mu - \lambda(t-1)) + \mathbf{K}\mathbf{d}_t + \varepsilon_t \tag{8}$$

In this formulation each of $(\Delta \mathbf{z}_{t-i} - \gamma)$ and $(\beta' \mathbf{z}_{t-1} - \mu - \lambda(t-1))$ is $I(0)$ and has a zero mean, and it makes clear the sources of growth, namely drift in $\mathbf{z}_t$ via γ, and deterministic trend in the equilibrium mean $\mathsf{E}(\beta' \mathbf{z}_t)$ via λ. Depending on their nature (impulse or step change) the event specific dummies induce step or trend change behaviour in $\mathbf{z}_t$.

As mentioned in section 2.2, there are occasions when it is convenient to condition on a subset of the variables $\mathbf{z}_t$. Consider the conditional/marginal factorisation of (8):

$$\begin{aligned}
(\Delta \mathbf{y}_t - \gamma_1) &= \Upsilon(\Delta \mathbf{x}_t - \gamma_2) + \sum_{j=1}^{p-1} (\Gamma_{11,j} - \Upsilon\Gamma_{21,j})(\Delta \mathbf{y}_{t-j} - \gamma_1) \\
&\quad + \sum_{j=1}^{p-1} (\Gamma_{12,j} - \Upsilon\Gamma_{22,j})(\Delta \mathbf{x}_{t-j} - \gamma_2) + (\mathbf{K}_1 - \Upsilon\mathbf{K}_2)\mathbf{d}_t \\
&\quad + (\alpha_1 - \Upsilon\alpha_2)(\beta' \mathbf{z}_{t-1} - \mu - \lambda(t-1)) + \nu_t \\
&= \Upsilon(\Delta \mathbf{x}_t - \gamma_2) + \sum_{j=1}^{p-1} \Psi_j (\Delta \mathbf{z}_{t-j} - \gamma) \\
&\quad + \psi(\beta' \mathbf{z}_{t-1} - \mu - \lambda(t-1)) + \kappa \mathbf{d}_t + \nu_t
\end{aligned} \tag{9}$$

and

$$
\begin{aligned}
(\Delta \mathbf{x}_t - \gamma_2) &= \sum_{j=1}^{p-1} \Gamma_{21,j}(\Delta \mathbf{y}_{t-j} - \gamma_1) \quad + \sum_{j=1}^{p-1} \Gamma_{22,j}(\Delta \mathbf{x}_{t-j} - \gamma_2) \\
&\quad + \alpha_2(\beta' \mathbf{z}_{t-1} - \mu - \lambda(t-1)) + \mathbf{K}_2 \mathbf{d}_t + \varepsilon_{2,t} \\
&= \sum_{j=1}^{p-1} \Gamma_{2,j}(\Delta \mathbf{z}_{t-j} - \gamma) \quad + \alpha_2(\beta' \mathbf{z}_{t-1} - \mu - \lambda(t-1)) \\
&\quad + \mathbf{K}_2 \mathbf{d}_t + \varepsilon_{2,t}
\end{aligned} \tag{10}
$$

with $\varepsilon_{2,t} \sim \mathrm{IN}_M(\mathbf{0}, \Omega)$, $\nu_t \sim \mathrm{IN}_Q(\mathbf{0}, \Sigma_{22})$, $\mathsf{E}(\varepsilon_{2,t}\nu_t') = 0$ when $\Upsilon = \Sigma_{11}\Sigma_{22}^{-1}$, $\Omega = (\Sigma_{11} - \Upsilon\Sigma_{22}\Upsilon')$, $\alpha' = (\alpha_1', \alpha_2')$, $\mathbf{K}' = (\mathbf{K}_1', \mathbf{K}_2')$, and $\Gamma_l = (\Gamma_{ij,l})$ for $i,j = 1,2$ and $l = 1,2,\ldots p-1$.

For this factorisation modelling the variables $\mathbf{y}_t$ conditional on $\mathbf{x}_t$ involves the open system (9), whereas the joint modelling of the conditioning variables $\mathbf{x}_t$ involves the marginal system (10).

3.1 Types of Change

When models of the type in (8), or its conditional-marginal factorisation in (9) and (10), are used for data on economies that have experienced substantial change, or are in transition, some parameter non-constancies are likely particularly if no event-specific dummy variables are included so that $\mathbf{K} = 0$ *a priori*. The parameters $(\Upsilon, \Psi_1, \cdots \Psi_{p-1}, \psi, \beta, \mu, \lambda, \gamma, \kappa, \Omega)$ of the conditional system (9) when they change suffer structural breaks as defined by Hendry and Mizon (1998) and Hendry and Mizon (1999). Often such structural breaks arise from events that are external to the sector of the economy that is being modelled, and further are changes about which agents are unlikely to form expectations - rational or otherwise. The OPEC oil price increase in 1973, the speculative attack on the Italian lira and the British pound by Soros funds leading to both currencies leaving the ERM in September 1992, and the opening of the Berlin Wall in 1990, are examples of unknown and external changes that caused structural breaks in many econometric models. Also, as envisaged by Lucas (1976), observed or expected changes in economic policy might induce structural breaks in econometric models that are conditioned on the policy variables in question. Whether structural breaks are the result of external and unpredictable events, or of the Lucas critique, is an empirical issue. We note that Ericsson and Irons (1996) found little evidence of the Lucas critique being empirically relevant. Another type of change that can be represented in econometric modelling by parameter shifts consists of announced changes in (economic) policy, and institutional changes, which by their nature are known. Examples of this type of change include switches from Keynesian to monetarist policies, and the introduction of interest bearing current accounts, and of credit cards. When $\mathbf{x}_t$ consists of policy instruments, shifts in the parameters ($\Gamma_{2,1}, \cdots \Gamma_{2,p-1}, \gamma, \alpha, \beta, \mu, \lambda, \mathbf{K}_2, \Sigma_{22}$) represent such policy and institutional regime shifts.

The European economies studied in this volume have experienced significant changes in economic structure and economic policies pursued during the last three decades. Some economies have undergone substantial liberalisation of their labour, financial, and foreign exchange markets. Other economies of eastern and central Europe have moved from being centrally planned towards free market economies. Intermediate between these extremes are economies that have slowly adopted policies to liberalise their financial and foreign exchange markets, and introduce some degree of flexibility into their labour markets. How are such changes likely to exhibit themselves in the class of econometric model given by (8)? Firstly, there might be changes in the underlying equilibria arising from changes in the number of equilibria r and/or in the parameters of $\beta'\mathbf{z}_t$ and $\mathsf{E}(\beta'\mathbf{z}_t)$, namely β, μ, and λ. In the context of (8) these are perhaps the most dramatic kind of change, though another important change arises when there is a shift in $\mathsf{E}(\Delta\mathbf{z}_t) = \gamma$ the mean growth rate of $\mathbf{z}_t$. The long-run characteristics of the system are determined by r and the parameters namely β, μ, λ, and γ, and changes in these are in an important sense fundamental. Indeed, were there to be changes in β, then the model cannot be capturing the underlying structure of the sector of the economy being modelled which by definition is invariant. The adjustment of the system in response to such fundamental changes is determined by its dynamic characteristics, namely the short-run adjustment parameters Γ_j of the closed system and Ψ_j of the open system, together with the corresponding long run adjustment coefficients α and ψ respectively. We note further that changes in α or ψ can reflect changes in the weak exogeneity status of variables for the long-run parameters β.

All the above parameter changes have involved first moments of the variables, and modelling such changes is the focus of attention in the papers in this volume. However, it is clear that changes in second and higher order moments are possible, as for example a change in the innovation error variance-covariance matrix Σ. More complicated changes in second moments are possible, and particularly relevant for high frequency financial data - ARCH (see e.g., Engle, 1995) and stochastic volatility models (see e.g., Kim, Shephard and Chib 1998) aim to account for this phenomenon.

3.2 Implications and Detectability of Change

The existence of regime shifts and structural breaks in econometric models has many consequences for the use of such models. For example, forecasting in the presence of structural breaks is often treacherous. Using the taxonomy of forecast errors in Clements and Hendry (1996a), Hendry and Doornik (1997) establish that deterministic shifts are the primary source of systematic forecast failure in econometric models. In particular, changes in intercepts and deterministic trends have more pernicious effects on forecasts than changes in slope parameters, which are often coefficients of zero-mean variables. Whereas changes in the coefficients of zero-mean variables (e.g., Γ_j and α in (8) and Υ, Ψ_j, and ψ have limited effects on forecasts, shifts in equilibrium means and growth rates (e.g., μ and γ) can induce dramatic forecast failure. Similarly, Hendry and Doornik (1997) and Hendry (1999) establish that, even quite large changes in the coefficients of zero-mean variables (Γ_j , α, Υ, Ψ_j and ψ), are not easily detected using standard tests for forecast failure. Shifts in equilibrium means and growth rates (μ, λ and γ) on the other hand are easily detected. Nevertheless, there exist devices that can robustify forecasting models against such breaks, provided they have occurred prior to forecasting (see e.g., Clements and Hendry 1996b, and Hendry and Clements 1998). These robust devices include modelling differences of variables rather than their levels (i.e., $\Delta\mathbf{z}_t$ or $\Delta^2\mathbf{z}_t$ rather than $\mathbf{z}_t$), so that for example a deterministic shift in $\mathbf{z}_t$ is replaced by a single blip in $\Delta\mathbf{z}_t$, or a change in the deterministic growth rate of $\mathbf{z}_t$ becomes a blip in $\Delta^2\mathbf{z}_t$. Another such device, with a long history of usage in macroeconomic forecasting (see e.g., Klein 1971), is intercept-correcting which adjusts forecasts back towards the unconditional mean, thus correcting any systematic forecast error. However, no methods are robust to unanticipated breaks that occur after forecasting, and Clements and Hendry (1998) show that those same 'robustified' devices do not offset post-forecasting breaks. Although these robust forecasting devices can help mitigate forecast failures, they will often not have useful policy implications.

For policy analysis using econometric models to be valuable reliable estimates of policy responses are required. Since forecasting devices robust to in-sample structural breaks usually do not include policy responses; the best forecasting device will often be uninformative for policy analysis. Conversely, despite having experienced forecast failures from pre-forecasting structural breaks, econometric systems which do embody the relevant policy effects can be valuable for estimating the likely effects of policy regime shifts. Consequently, when both structural breaks and regime shifts occur, neither forecasting devices nor econometric models alone are adequate for forecasting. This led Hendry and Mizon (1999) to explore the advantages of combining information from both sources. Hence the presence of structural breaks and regime shifts cannot be ignored with impunity when forecasting or undertaking economic policy analysis.

A major objective of economic and econometric analysis is to find invariant relationships that accurately characterise observed economic activity. Though this is a relatively easy task in a stationary world, there are many non-stationarities, in the form of unit roots and deterministic shifts, in the data generating process for typically observed macroeconomic time series data. However, rather than being a nuisance, such deterministic shifts can be invaluable for model evaluation. Indeed, one of the most powerful ways of discriminating between regime-specific invariance and genuine structure (which is relevant both within and across regimes), is to isolate the models that are invariant to substantial change in an economy. The recent changes that have taken place in the economies of central and Eastern Europe, as well as those in Western Europe, provide us with valuable information for discovering structure.

3.3 Modelling Strategies in the Presence of Structural Breaks

The existence of non-constant parameters in many classes of statistical and econometric model has led to the development of alternative strategies for accommodating these changes. Many of the models developed to deal with evolving systems, for which linear representations are not always adequate, are non-linear, and Granger and Teräsvirta (1993) provides an overview of this literature. Maddala and Kim (1998) provide an alternative and more detailed account in the context of cointegrated models.

One of the earliest classes of model developed was that in which coefficients are treated as random variables, or the parameters as time-varying - see Chow (1984) for a review. In effect, such models are attempting to find invariance at a higher level than that of linear coefficients. For many of the issues analysed in this volume it is not appropriate to adopt models that have linear random coefficients, for example, following random walks. A similar approach is that of structural time series modelling, described for example in Harvey (1989) and Harvey (1993), and implemented in the software written by Koopman, Harvey, Doornik and Shephard (1995), based on use of the Kalman filter. Another class of model has a number of different regimes (typically no more than three), between which there is frequent switching. The Markov switching model of Hamilton (1989), and the further developments in Krolzig (1997), are examples in this category. When there is not frequent switching between regimes and a more gradual transition from one regime to another, the class of smooth transition autoregressive models is likely to be relevant - see Teräsvirta (1994) for example.

Many of the papers in this volume are concerned with at most two regimes, with a single transition from one to the other. Hence we will confine attention to the issue of parameter non-constancy in the class of VAR models in (1) and VEqCM's in (8). Within this class of model parameter non-constancy can arise from using a restricted information set, both in terms of the number of variables and the number of lags included. For example, omitting an important variable such as an interest rate or an exchange rate in modelling wage and price determination in an open economy can often result in parameter non-constancy and predictive failure. Conversely, coefficients of some variables at long lags may appear significant, not because the relevant dynamic structure implies long lags, but as a result of the power such coefficients have to capture regime shifts or structural breaks. Care is therefore required in the interpretation of apparently significant estimated coefficients at long lags. When there are strong reasons to believe that the particular sector of an economy being modelled has experienced an important shift, modelling this sector as separate regimes pre and post the shift will often be appropriate. This approach though does not incorporate in the model an explanation of the regime shift, but rather regards it as an exogenously given event. The unification of Germany, moves from fixed to floating exchange rate regimes, and the shift from central planning to more market-oriented economies in Central and Eastern Europe, provide examples. Important changes, which are less severe in their consequences than those leading to regime shifts, can be accommodated by the introduction of impulse and step dummy variables into the model. This volume contains a number of papers that have successfully adopted the latter two approaches.

4 Overview of the Volume

The papers presented in this volume deal with varying but related problems of modelling economies in transition. Although the authors were very much aware of differences between the economic systems, which were the subject of each modelling exercise, they employed a common theoretical and methodological approach. Consequently an overall unifying theme of this volume is the analysis of cointegrated structural models. The difference between studies lies principally in the degree of structure, which was imposed in each case and in the amount, and speed of the structural change, which has gone on in each country.

Each study in this volume also started from a consensus view on the economic background of the empirical analyses. Thus, as a starting point, it was postulated that real earnings depend on productivity and unemployment whilst themselves they can affect productivity and also inflation if their increases are not compensated by productivity growth. Unemployment should react negatively to both real wage and productivity increases. The wage indexation mechanism was generally considered as one of the main determinants of inflation and its persistence; with its impact depending on the existant indexation structure.

We define the term “transition” in a wide sense. It is not limited to the transition from centrally planned to competitive market structure alone, but also covers the transition from Keynesian to monetarist economic systems, which was observed during the 1980’s in many European countries. There are also more unusual forms of transition such as the unification of Germany, which resulted in a very high unemployment rate in East Germany; similar to the rates observed in the CEEC. All these ‘transitions’ characteristically exhibit a single shift from one regime to the other, and thus have considerable consequences for the functioning of the whole system. As a result, the long-run, as well as the short-run, relationships are subject to change. Consequently the class of model applied for such an unstable world must allow for changes in parameters defining the equilibria, and even in the number of cointegrated vectors.

The overview of statistical modelling of cointegration with special emphasis on mathematical formulation of the model and derivation of estimators and test statistics is given in Johansen's paper. The idea of common trends in the cointegrated VAR is then developed, and the methodological base for cointegration analysis is provided. This paper contains a summary of the asymptotic theory applied to the test statistics and estimators, as well as a brief discussion of two applications in the context of cointegration - namely rational expectations, and arbitrage pricing theory. This paper gives a unifying methodological basis to the rest of the empirical work reported in other papers.

Marcellino and Mizon analyse the labour market characteristics of the UK, Italy, and Poland, paying attention and testing the possibility that there has been a substantial shift in each country's labour market. Other issues raised in the analysis include; the relationship between unemployment and productivity, inflation and real earnings, and the wage-price nexus. The relevance of these issues is clear when it is noted that the high rate of unemployment is one of the biggest problems currently facing European Union governments. The authors find clear evidence for there being a major change, (occurring around 1979/80), in the underlying nature of the unemployment-inflation and wage-price relationships in each of these economies. However, they find no empirical support for the hypothesis of there being a common structure across the countries in the two regimes or the transition between them. Noting the substantial economic and political differences between these three countries, and the very different economic policies adopted by their governments, the lack of commonalities in the changes that have taken place is understandable.

In analysing the UK economy, Hall and Nixon were mainly concerned with the explanation of the specific properties of British unemployment, especially the long-run implications of capital-labour substitution for wage setting behaviour. They found that the natural rate of unemployment is a function of factor shares, and so the monetarist economic policies adopted in many Western countries may have led to an increase in equilibrium unemployment. The testing of this hypothesis was done in the context of a complete structural model of the UK supply side, where production was modelled using a dynamic flexible functional form. Much attention was devoted to the issue of imposing all the relevant cross-equation restrictions based on economic theory, prior to estimating in levels form an internally consistent system for production, factor demands, and wage and price determination. Firstly, the authors tested for cointegration within this non-linear model in each equation separately. They then estimated the full dynamic, non-linear, model, and tested the validity of the over-identifying restrictions. The authors also used the model's *ex ante* forecasts to test some more complicated hypotheses about long-run behaviour of the complete supply-side system.

Welfe argues those cost-push factors, and especially wage pressures, were the major sources of inflation during the 1990's in Poland, following the introduction of the stabilisation program and the indexation mechanism. Two different, aggregate models of inflation were built. The first was specified entirely in growth rates (first differences of logs), and thus excludes completely the long-run level relationships. However, it allows for the analysis of turning points (shifts). The failure to reject the hypothesis that prices provide a threshold effect, led to consideration of an endogenous switching model, which was estimated by Bayesian methods. The results indicate the presence of some positive threshold of inflation. However, a possible limitation is that the value of the threshold was kept constant throughout the sample, and this threshold was used to define periods characterised by the highest inflation rates. To address this problem, the second model uses the SVAR approach and aims to identify the long-run relationships between wages, prices, labour productivity, and unemployment. The results indicate that the influence of unemployment on wages is insignificant, while wage costs are the dominant factor leading to inflation. It is worth noting that the estimated long-run price elasticity of wages is (close to) unity, as postulated by theory.

The contribution by Hansen explores the insider-outsider hypothesis in the context of the German economy before and after unification. The paper analyses how the West German system reacted to the shock, and questioned whether the unified Germany has a stable labour market structure. The results show that the West German labour market can be described by means of five variables with two cointegration relations, which satisfy standard labour demand and wage setting restrictions. Due to significant effects of the rate of unemployment for wage setting there is no strong evidence for insider behaviour before the unification, while the adjustment to equilibrium of real wages is still slow. Results for the period, which includes unification, confirm that the restrictions imposed for the earlier period are consistent with the complete data set if the coefficient of the unemployment in the wage equation is allowed to increase. The latter means that the unification shock is well understood in wage bargaining, although the speed of adjustment towards equilibrium may be too sluggish. There is no evidence that after unification wages were set mainly in favour of West German employees.

5 Conclusion

This paper has set out the methodological framework that unified the research activities of the ACE project on *Inflation and Unemployment in Economies in Transition,* and which underlies all the papers in this volume. In doing so it has outlined the broad structure of the cointegrated VAR approach to modelling macroeconomic time series. In particular, it has shown how such systems can be identified when they have linear or non-linear cointegrating vectors, and how they may be reduced to a structural VAR embodying testable economic hypotheses. It has further considered the implications of structural breaks within this framework. Finally it has related the papers in this volume to the overall modelling strategy, and briefly discussed their results. Indeed, the papers in the volume provide an illustration of the feasibility and value of adopting this framework when modelling the labour sector of economies that have experienced substantial change.

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