EUI WORKING PAPERS IN ECONOMICS

EUI Working Paper ECO No. 95/10

Progressive Modelling of Macroeconomic Time Series: The LSE Methodology

GRAYHAM E. MIZON

European University Institute, Florence

EUROPEAN UNIVERSITY INSTITUTE, FLORENCE

ECONOMICS DEPARTMENT

EUI Working Paper ECO No. 95/10

Progressive Modelling of Macroeconomic Time Series: The LSE Methodology

GRAYHAM E. MIZON

BADIA FIESOLANA, SAN DOMENICO (FI)

Printed in Italy in April 1995
European University Institute
Badia Fiesolana
I – 50016 San Domenico (FI)
Italy

Progressive Modelling of Macroeconomic Time Series: The LSE Methodology.

Grayham E. Mizon

Economics Department, Southampton University, UK
&
Economics Department, European University Institute, Florence, Italy.

July 1994. Revised December 1994.

Abstract

A brief history of the development of the LSE econometric modelling methodology is presented. The essence of the methodology is the recognition that potentially valuable information for the analysis of any economic problem can come from numerous sources including, economic theory, the available sample of observations on the hypothetically relevant variables, knowledge of the economic history of the period under study, and from knowledge of the way in which the observed data are defined and measured, plus their relationship to the theory variables. The principal features of the methodology (especially congruence and encompassing for model evaluation, and general-to-specific as a modelling strategy) are then described, and their roles illustrated in the development of a simple model of the time series relationship between wages, prices, and unemployment in the UK between 1965 and 1993, from both a single equation and system perspective. It is found that single equation analysis yields inefficient inference relative to system analysis, and that there is a structural change which may reflect a change in economic policy.

Progressive Modelling of Macroeconomic Time Series: The LSE Methodology

Grayham E. Mizon

[illegible] Department, Southampton University, UK

[illegible] European University Institute, Florence, Italy

[illegible]

Abstract

[illegible]

1 Introduction

Econometric models, large and small, have played an increasingly important role in macroeconomic forecasting and policy analysis. However, there is a wide range of model types used for this purpose, including: simultaneous equations models in either reduced or structural form, vector autoregressive models (VAR), autoregressive-distributed lag models, autoregressive integrated moving average models, leading indicator models, and error correction models (ECM). Hendry, Pagan and Sargan [1984] discuss a typology for dynamic single equation models for time series variables, and Hendry [1993c] presents a typology for the various types of dynamic model used in the analysis of systems of equations. There is also a wide range of views about the appropriate way to develop and evaluate models. Sims [1980, 1992] advocates the use of VAR models, which can accurately represent the time series properties of data, whilst eschewing the reliance on "incredible indentifying restrictions" that characterizes the use of simultaneous equation models of the structural or Cowles Commission type. The potential value of structure (loosely defined) within the context of VAR models has led to the development of structural VAR models, and Canova [1993] provides a recent review of this literature. Leamer [1978, 1983] on the other hand has been critical of the use of non-Bayesian models that do not analyse formally the role and value of a priori information, especially when there is no checking of model sensitivity. Summers [1990], though aware of the important developments made in theoretical statistics and econometrics in this century, argues that too much emphasis is placed on the technical aspects of modelling, and not enough on the real issues which are concerned with the analysis of well established and fundamental relationships between economic variables. One approach to modelling that does not over-emphasize the role of model evaluation and statistical technique, is that associated with real business cycle analysis and the calibration of economic theory, rather than its evaluation. Kydland and Prescott [1982, 1993] have been pioneers in this field, and Canova, Finn and Pagan [1994] provide a critique.

The focus of this paper is an alternative approach to the modelling of economic time series that was originally developed at the LSE, though contributions and extensions to this methodology have subsequently been made by econometricians who have had no direct connection with LSE as an institution. The essence of this approach is the recognition that potentially valuable information for the analysis of any economic problem can come from numerous sources including, economic theory, the available sample of observations on the potentially relevant variables, knowledge of the economic history of the period under study, and from knowledge of the way in which the observed data are defined and measured, plus their relationship to the theory variables. In the development of econometric models it is therefore important that information from all these sources is exploited as fully as possible. Of course the marginal return to the exploitation of information from each source will vary across the sources, and the degree to which each source has been used already. However, attention here will be confined to consideration of the extent to which available information has been exploited, and the relationship between this and the evaluation and comparison of alternative models of the same phenomena.

The next section provides a brief history of the group of econometricians involved in the development of the LSE modelling methodology. This is followed by a section presenting the essential components of the LSE methodology, emphasizing the importance of evaluating and comparing alternative models within a statistically and economically coherent framework. This approach helps to ensure that econometric modelling is progressive by only abandoning models found to be inadequate in favour of models that have been demonstrated to be improvements. It is also progressive in the sense that it is not necessary to know the complete structure characterizing the relationship between economic variables prior to commencing an analysis of it. Rather it is possible to discover incrementally parts of the underlying structure as a result of careful analysis - see Hendry [1993b]. Section 4 contains an analysis of the time series relationship between wages, prices, and unemployment in the UK between 1965 and 1993, from both a single equation and system perspective. It is found that single equation analysis can yield misleading inferences relative to system analysis, and that there appears to be an important change in structure, which may reflect a change in economic policy. The final section provides conclusions.

2 Brief History of the LSE Methodology's Development

A distinctive feature of the British tradition in statistics is the high quality of applied work implementing, developing and stimulating theoretical research, with the early work of R.A. Fisher, J.B.S. Haldane, K. Pearson, and G.U. Yule, and the later contributions of G.A. Barnard, M.S. Bartlett, G.P.E. Box, D.R. Cox, M.G. Kendall, D.V. Lindley and E.S. Pearson being examples. That the LSE, as a leading university specializing in the social sciences, should follow and take advantage of this tradition was natural. Indeed, A.L. Bowley and R.G.D. Allen initiated the LSE's strength in statistics and economic statistics in particular, with Roy Allen teaching a course in econometrics as early as 1946/47 and 1947/48 - see Gilbert [1989]. From the 1950's the LSE fostered and enlarged it's group of statisticians engaged in social science applications, when the statistical research in most other institutions was concerned with medical and biological applications. As a result, the Statistics Department at LSE became pre-eminent in its chosen field of statistical research. By the mid 1960's R.G.D. Allen, D.J. Bartholemew, D. Brillinger, J. Durbin, J. Hajnal, M.G. Kendall, M.H. Quenouille, C.A.B. Smith, and A. Stuart had been amongst its members. Throughout this period time series analysis was an important area for research and teaching at the LSE, with Kendall, Quenouille and then Durbin being the early leaders in this field. This tradition in the Statistics Department was later to play an important role in the development of the LSE econometric methodology.

Although the Statistics Department had been responsible for the early teaching of economic statistics, and the newly emerging subject of econometrics, it was jointly with the Economics Department that the initiative was taken to expand the activities in econometrics. A.W.H. (Bill) Phillips, who had been in the Economics Department since the mid 1950's, used his knowledge of electrical engineering to study the dynamics of economic systems. This resulted in important research on continuous time autoregressive moving average models, an hydraulic model of the linkages between stocks and flows in the economy, introduction into economics of the concepts of integral, derivative and proportional control mechanisms, and the development of the empirical relationship between aggregate wages and unemployment - the Phillips curve. Phillips was joined by fellow New Zealander, Rex Bergstrom at the beginning of the 1960's, and he too developed statistical theory for the analysis of continuous time models as well as building small continuous time macroeconometric models (see Phillips [1992, 1993] for discussion of Bergstrom's contributions to econometrics). The major push in econometrics came in the period 1962-65, which resulted in the appointment of J. Denis Sargan, and Meghnad Desai, and Jan Tymes in the Economics Department, Kenneth F. Wallis in the Statistics Department, and the introduction of taught MSc courses in economics and econometrics. As an LSE student at the time I was aware that changes were taking place, and gradually realized that they were very important developments for the LSE and for at least one generation of students to follow.

By 1965 the "new" MSc courses in Economics and the more specialized course in

Mathematical Economics and Econometrics were introduced to provide a thorough training for professional economists, and in the next few years there were further staff recruitments e.g. Terence Gorman, Frank Hahn, Michio Morishima, and Amartya Sen in economics, and a considerable expansion in the number of PhD students working in theoretical economics and econometrics. Within the field of econometrics the first few cohorts of students supervised by Denis Sargan included: Ray Byron, Terry Charatsis, Emmanuel Dretakis, Toni Espasa, Tony Hall, David Hendry, Essie Maasoumi, William Mikhail, Grayham Mizon, Pravin Trivedi, Peter Phillips, Robin Rowley, Ross Williams, Cliff Wymer - see the fuller list including later cohorts given in Maasoumi [1988] volume 1. The principal supervisor of most PhD students in econometrics was Denis Sargan, and he had students working on an impressive range of topics: formulation, estimation, and testing of alternative dynamic model specifications; the use of the Lagrange multiplier testing principle for hypothesis testing in systems of equations; the treatment of missing observations in multivariate time series models; semiparametric estimation of systems using FIML and spectral estimates of the error covariance matrix; the development and use of nonlinear estimation methods; inference with continuous time models; the development of finite sample distribution theory and higher order asymptotic expansions to provide small sample approximations to distribution functions. A distinctive feature of much of the research conducted on these topics was the fact that it was embedded in applied econometric studies. Areas of application included models of: wages and prices; aggregate durable and non-durable consumer demand; consumer demand equations; aggregate production and factor demand behaviour, especially investment and inventory behaviour; and import and export determination.

Another important part in the development of the methodology was played by the visitors to LSE. Jean-François Richard, beginning with the year he spent at LSE as a visitor in 1973/4, but continuing through his collaboration in joint research with Hendry and Mizon, made major contributions in pointing out the important role of the joint distribution in the analysis of models of the relationship between observed variables, the role of the theory of reduction, and the formulation of precise definitions of exogeneity - Richard [1980] remains a tour de force in this area. The weekly Econometrics Workshop, associated with research programmes funded by the Social Science Research Council, was a focus for the development and discussion of research ideas internally, but it also benefited from the participation of longer term visitors such as Takeshi Amemiya, Ted Anderson (see his Foreword to Maasoumi [1988]), Rob Engle, Alain Monfort and Charles Nelson. Interaction with other econometricians came via periods of leave spent by members of the LSE econometrics group at the Australian National University, UC Berkeley, CORE, Yale, and particularly at UC San Diego. These visits provided valuable opportunities to prove and enrich ideas, and many lengthy, vigorous, and sometimes provocative conversations with Rob Engle, Clive Granger, Michel Mouchart, Adrian Pagan, Tom Rothenberg, and Hal White stimulated further developments and resulted in some joint papers.

By the end of the 1970's the LSE methodology was becoming more widely known as a result of journal publications, presentation of papers to conferences and in particular to the UK SSRC Econometrics Study Group, and via the expanding group of former students such as Gordon Anderson, Richard Blundell, Julia Campos, James Davidson, Neil Ericsson, Chris Gilbert, Mark Salmon and Aris Spanos. Also many of the estimation and testing procedures developed as an integral part of the methodology were implemented in the computer programme written by David Hendry, GIVE (Generalized Instrumental Variable Estimation), which was widely distributed and used. GIVE and other programmes in the AUTOREG library (see Hendry and Srba [1980]) were the mainframe precursors of the programmes PcGive and PcFiml (see Hendry [1989] and Doornik and Hendry [1992, 1993, 1994]) used in the empirical work reported in section 4 below.

Also at the end of the 1970's some members of the LSE econometrics group took up positions at other universities in the UK e.g. Wallis went to Warwick, Hendry to Oxford and Mizon to Southampton, and following this and the later retirement of Denis Sargan the methodology was much less clearly identified with the LSE. Indeed, the LSE has via Andrew Harvey (Statistics Department) and Peter Robinson (Economics Department) fostered other important aspects of econometrics, and the subsequent development of the methodology, especially for the analysis of systems of integrated-cointegrated variables, has been greatly influenced by the contributions of Peter Phillips (Yale), Søren Johansen (Copenhagen) and Katarina Juselius (Copenhagen) initially, and more recently by those of Peter Boswijk (Amsterdam) and Jean-Pierre Urbain (Limburg). Further, the initial development of the theory of encompassing by Hendry, Mizon and Richard (see Hendry and Richard [1982,1990], Mizon [1984] and Mizon and Richard [1986]) has been much extended and refined by the contributions of Jean-Pierre Florens (Toulouse), Neil Ericsson (Federal Reserve, Washington), and Maozu Lu (Southampton) amongst others. Gourieroux and Monfort [1992] provides a recent contribution. Hence there are many people who have contributed to the development of the LSE methodology, and many more now ensuring that it is refined, improved, and applied. The next section provides a brief description of its main components.

3 Components of the LSE Econometric Modelling Methodology

There have been many papers written on this topic (Ericsson, Campos and Tran [1993], Gilbert [1986, 1989], Hendry [1987], Hendry and Mizon [1990], Hendry and Richard [1982, 1983], Mizon [1991], Pagan [1987, 1989], Spanos [1990], and other papers in Granger [1990]). Hendry and Wallis [1984], which contains contributions by some of Sargan's former colleagues and students to mark his retirement, also provides an insight into the LSE methodology. Hence only a brief discussion of the main components of the LSE methodology will be presented in this section, together with comments on the nature of models and their relationship with the data generation process - DGP.

By way of illustration consider an artificial situation in which it is known that observations on y_t, x_t and z_t (which might be wages, prices and unemployment respectively) are independent drawings from a trivariate normal distribution with mean vector μ and covariance matrix $\mathbf{\Sigma}$. Hence the DGP for y_t, x_t and z_t is independent normal with mean vector μ and covariance matrix $\mathbf{\Sigma}$ with:

$$\mu = \begin{pmatrix} \mu_y \\ \mu_x \\ \mu_z \end{pmatrix} \quad \mathbf{\Sigma} = \begin{pmatrix} \sigma_{yy} & \sigma_{yx} & \sigma_{yz} \\ \sigma_{xy} & \sigma_{xx} & \sigma_{xz} \\ \sigma_{zy} & \sigma_{zx} & \sigma_{zz} \end{pmatrix} \tag{1}$$

i.e. the joint distribution of y_t, x_t and z_t, denoted $\mathsf{D}(y_t, x_t, z_t; \mu, \mathbf{\Sigma})$, has the form $\mathsf{D}(y_t, x_t, z_t; \mu, \mathbf{\Sigma}) = \mathsf{NI}(\mu, \mathbf{\Sigma})$. However, this joint density can be reparameterized as the product of conditional and marginal densities:

$$\begin{aligned} \mathsf{D}(y_t, x_t, z_t; \mu, \mathbf{\Sigma}) &= \mathsf{D}(x_t \mid z_t, y_t; \theta_x, \omega_{11}) \times \mathsf{D}(z_t \mid y_t; \theta_z, \omega_{22}) \times \mathsf{D}(y_t; \theta_y, \omega_{33}) \\ &= \mathsf{NI}(c_x + \beta y_t + \gamma z_t, \omega_{11}) \times \mathsf{NI}(c_z + \alpha y_t, \omega_{22}) \times \mathsf{NI}(c_y, \omega_{33}) \end{aligned}$$

This reparameterization is always possible, and yields an observationally equivalent representation of the DGP. Hence $\mathsf{NI}(\mu, \mathbf{\Sigma})$ and $\mathsf{NI}(\theta, \mathbf{\Omega})$ are observationally equivalent ways of presenting the trivariate normal distribution, when the mapping between the alternative parameterizations is given by:

$$\begin{aligned} \theta_x' &= (c_x, \beta, \gamma) \\ \theta_z' &= (c_z, \alpha) \\ \theta_y &= c_y \end{aligned} \qquad \mathbf{\Omega} = \begin{pmatrix} \omega_{11} & 0 & 0 \\ 0 & \omega_{22} & 0 \\ 0 & 0 & \omega_{33} \end{pmatrix}$$

with:

$$\begin{array}{lll} c_x = \mu_x - \beta\mu_y - \gamma\mu_z & c_z = \mu_z - \alpha\mu_y & c_y = \mu_y \\ \varphi\gamma = \sigma_{yy}\sigma_{zx} - \sigma_{zy}\sigma_{yx} & \varphi\beta = \sigma_{zz}\sigma_{yx} - \sigma_{yz}\sigma_{zx} & \alpha = \sigma_{zy}\sigma_{yy}^{-1} \\ \omega_{11} = \sigma_{xx} - \beta^2\sigma_{yy} - 2\beta\gamma\sigma_{yz} - \gamma^2\sigma_{zz} & \omega_{22} = \sigma_{zz} - \alpha^2\sigma_{yy} & \omega_{33} = \sigma_{yy} \end{array}$$

and $\varphi = \sigma_{yy}\sigma_{zz} - \sigma_{yz}^2$. Adopting the $(\theta, \mathbf{\Omega})$ parameterization the DGP can be described in the equations:

$$\begin{aligned} x_t &= c_x + \beta y_t + \gamma z_t + v_{1t} \\ z_t &= c_z + \alpha y_t + v_{2t} \\ y_t &= c_y + v_{3t} \end{aligned} \tag{2}$$

$$\mathbf{v}_t = (v_{1t}, v_{2t}, v_{3t})' \sim \mathsf{NI}(0, \mathbf{\Omega})$$

Now, let the equations (2) be the equations used to generate the data in a Monte Carlo simulation study, so that the investigator who writes down the model:

$$\mathsf{M}_1 : \; x_t = c_1 + \beta_1 y_t + \gamma_1 z_t + u_{1t} \tag{3}$$

with $u_{1t} \sim \mathsf{NI}\left(0, \sigma_{u_1}^2\right)$is clearly seen to have chosen a "true" model since it corresponds exactly to an equation of the DGP, with $c_1 = c_x, \beta_1{=}\beta, \gamma_1 = \gamma$ and $\sigma_{u_1}^2 = \omega_{11}$. Indeed, M_1 corresponds to $\mathsf{D}(x_t \mid z_t, y_t; \theta_x, \omega_{11})$ which is part of the DGP. Hence provided that the phenomenon of interest to the investigator can be analysed solely in terms of the parameters of M_1, and y_t and z_t are weakly exogenous variables for the parameters of interest, no problems will arise for that investigator. The fact that $\mathsf{D}(x_t \mid z_t, y_t; \theta_x, \omega_{11})$ is only part of the DGP so that M_1 is not exploiting all the information in the DGP, is the reason for requiring y_t and z_t to be weakly exogenous. If the parameters of interest to the investigator are denoted ϕ, then y_t and z_t will be weakly exogenous for ϕ provided that (i) ϕ can be recovered uniquely from θ_x and ω_{11}, and (ii) the values of θ_x and ω_{11} in the DGP are chosen independently of the values of the other parameters of the DGP, namely $\theta_z, \theta_y, \omega_{22}$ and ω_{33}. The requirement (ii) is satisfied for the DGP being discussed here provided that there are no a priori restrictions linking (θ_x', ω_{11}) to $(\theta_z', \theta_y, \omega_{22}, \omega_{33})$ (e.g. $\alpha = \beta$). Engle et al [1983] provide fuller definitions and discussion of this and other concepts of exogeneity.

Now consider another investigator who claims to be interested in the model:

$$\mathsf{M}_2 : \; y_t = c_2 + \alpha_2 x_t + \gamma_2 z_t + u_{2t} \tag{4}$$

with $u_{2t} \sim \mathsf{NI}\left(0, \sigma_{u_2}^2\right)$ this investigator is clearly not considering the "true" model since M_2 does not correspond to any of the equations of the DGP as given in (2)! It is though part of other valid parameterizations of the DGP:

$$\mathsf{D}(y_t, x_t, z_t; \mu, \mathbf{\Sigma}) = \mathsf{D}(y_t \mid x_t, z_t; \xi_y) \times \mathsf{D}(x_t \mid z_t; \xi_x) \times \mathsf{D}(z_t; \xi_z)$$

$$= \mathsf{D}(y_t \mid x_t, z_t; \xi_y) \times \mathsf{D}(z_t \mid x_t; \zeta_z) \times \mathsf{D}(x_t; \zeta_x)$$

and thus "true". Further, the investigator who chooses to analyse the model:

$$\mathsf{M}_3 : \quad y_t = c_3 + \alpha_3 x_t + u_{3t} \tag{5}$$

with $u_{3t} \sim \mathsf{NI}\left(0, \sigma^2_{u_3}\right)$ is clearly misguided in ignoring the influence of z_t on y_t and could be argued to be using a misspecified model, in addition to not having a "true" model. However, M_3 need be neither misspecified nor untrue. If the parameters of interest to the investigator are those of the distribution $y_t \mid x_t$ (*i.e.* ς in $\mathsf{D}(y_t \mid x_t; \varsigma)$) then M_3 is obtained as a reduction of the DGP given by $\mathsf{NI}(\mu, \mathbf{\Sigma})$, and as such is both true and correctly specified when $c_3 = \mu_y - \alpha_3\mu_x$ and $\alpha_3 = \sigma^{-1}_{xx}\sigma_{xy}$ with $\sigma^2_{u_3} = \sigma_{yy} - \alpha^2_3\sigma_{xx}$. Alternatively, if M_3 is thought to be a statement about the distribution $\mathsf{D}(y_t \mid x_t, z_t; \xi_y)$ then it will only be fully exploiting the available information if the conditional independence hypothesis $y_t \perp z_t \mid x_t$ (or $\gamma_2 = 0$ in M_2) is valid.

Two important points have been illustrated in this discussion. Firstly, all models can be obtained as reductions of the DGP for the variables they involve, and therefore their properties are implied by the DGP. In fact, the parameters of any model are appropriately interpreted as the pseudo true values (derived under the DGP) of the maximum likelihood estimators given the model specification. Secondly, it is important in all econometric studies to know what the parameters of interest are, or equivalently, what are the relevant statistical distribution and information set. Hence a model can only be relevant if it is defined with respect to an appropriate information set, and if the parameters of interest can be uniquely derived from its parameters. But this is not sufficient for efficient inference on the parameters of interest. It is also important that the model be congruent with the available information. For example, a necessary condition for a model like M_3 to be congruent, when the parameters of interest are functions of ξ_y, is $\gamma_2 = 0$. This is a testable hypothesis, and it is important that it be tested as a part of the evaluation of the model.

The importance of testing and evaluating econometric models is central to the LSE methodology. This was clear in Sargan [1964] in which tests for functional form (natural logarithm ln *versus* level), for appropriate dynamic specification (the order of dynamics and common factor restrictions - COMFAC), and instrument validity, were proposed and used. Subsequent papers which further developed this theme include Hendry and Mizon [1978], Hendry and Richard [1982, 1983], Mizon [1977a], and Hendry [1980] - this last paper concluding with the recommendation to "test, test, and test ". Hypothesis testing can be used both to assess the adequacy of the currently specified model (misspecification testing), and to explore the possibility for simplifying the current model without losing coherence with available information including the observed data (specification testing) - see Mizon [1977b] for discussion of this distinction. Both these forms of testing play important roles in the evaluation of econometric models an issue to which attention is now turned.

3.1 Criteria for Evaluation

If an econometric model is to be taken and used seriously then it's credentials must be established. Two important ways to do this are to demonstrate that the model is coherent with the available relevant information, and that it is at least as good as alternative models of the same phenomena. Hence congruence and encompassing are central to the LSE methodology.

3.1.1 Congruence

In assessing the coherence of a model with available information it has to be recognised that there are many sources of information, which can usefully be divided into four categories: (a) a priori theory; (b) observed sample information; (c) the properties of the measurement system; and (d) rival models.

(a) A Priori Theory That for econometric models the primary source for a priori theory is economics and economic theory is obvious and uncontentious. The relative importance of economic theory and statistical criteria in the development and evaluation of econometric models though has been the subject of much debate, with probably the most frequently cited example being the 'measurement without theory' debate between Koopmans and Vining (see Hendry and Morgan [1995] for a recent assessment of this debate). However, the tensions between VAR modelling, and Real Business Cycle and calibrated models are a current manifestation of the debate that still surrounds this issue. In crude terms the contrast is between data-driven modelling in the use of VAR's, and theory-driven modelling of either the structural type (á la Hansen and Sargent) or the calibrated real business cycle type. Canova [1994] presents a survey of VAR modelling, Ingram [1994] reviews structural modelling, and Kydland and Prescott [1994] present the case for calibration methods applied to models of the real business cycle, and Kim and Pagan [1994] discuss ways in which this latter class of model can be evaluated. In this context the LSE approach draws upon the strengths of each of these alternative approaches, though this is not the way that the LSE methodology evolved. Hence Hendry [1993b] argues that whilst it is desirable that econometric models be consistent with an economic theory, it is not essential. Indeed, if all models were required to have their origins in existing economic theory then the scope for developing new theories and learning from observation would be greatly reduced. Further, there is nothing that endows an economic theory with veracity a priori, and so coherence of an econometric model with an economic theory is neither necessary nor sufficient for it to be a good model. The practice of judging a model to be satisfactory if all the signs and magnitudes of its estimated parameters are consistent with an economic theory, and hence regarding the economic theory as being confirmed by the evidence, has little to recommend it - see Gilbert [1986] for a critique of this approach. Not the least reason for this, being the fact that it is possible for a number of alternative models to share this property (see Mizon [1989] for an illustration).

There is a sense though in which all econometric models result from the use of economic considerations, such as the choice of variables to include in the model, and of the functional forms to characterize the relationships between them. Congruence with economic theory in this category (which is similar to what Hendry [1993b] calls low level theory) is an essential feature for an econometric model. However, theories that imply very tight and specific forms for econometric models (e.g. models embodying rational expectations, intertemporal optimization, or Euler equation formulations) are higher level theories that embody testable hypotheses, rather than specifying essential characteristics for any econometric model of the phenomena being studied. The value and feasibility of testing economic theory against empirical observation are also much debated issues, as are the questions of what are appropriate interpretations and reactions to the results of statistical hypothesis tests. In particular, if a statistical test of an hypothesis is unfavourable, there are many reasons why this may not lead to the immediate rejection of the hypothesis, or the model incorporating the hypothesis. For example: (i) the statistical model providing the framework for the hypothesis test may yield invalid or misleading inferences as a result of being non-congruent with the sample information; (ii) there may be doubts about the suitability of the statistical model as an implementation of the economic theory; (iii) the observed data may not correspond closely with the latent variables of the economic theory, or may be subject to other measurement errors; or (iv) there may be no alternative models that dominate it.

(b) Sample Information The requirement that an econometric model be congruent with sample information is a necessary condition for statistical inferences based on it to be valid. Since there are distinct concepts associated with a model's congruence with respect to past, present and future sample information, each of these types of sample information will be considered separately.

(i) Relative Past Sample Information Past sample information consists of observations on lagged values of the variables in the modelling dataset. Hence if the t^{th}observation on the N variables in the dataset is denoted by $\mathbf{x}_t$ this constitutes the present (relative to the time index t) sample information, the relative past sample information is $\{\mathbf{x}_1, \mathbf{x}_2, ...\mathbf{x}_{t-1}\}$, and the relative future information is $\{\mathbf{x}_{t+1}, \mathbf{x}_{t+2}, \mathbf{x}_{t+3}, ...\mathbf{x}_T\}$ when T is the final observation available in the sample.

A model that is not congruent with past sample information will have errors that are correlated with lagged values of $\mathbf{x}_t$, and hence the errors will be serially correlated and at least partially explainable in terms of $\{\mathbf{x}_1, \mathbf{x}_2, ...\mathbf{x}_{t-1}\}$. For example, if the OLS estimator of δ in the static regression model:

$$y_t = \delta z_t + u_t \tag{6}$$

is used to make inferences, and possibly policy recommendations, involving the long run response of y_t to z_t when u_t is serially correlated, invalid and misleading inferences are likely to be the result. In particular, let the appropriate model to capture the relationship between y_t and z_t, and distinguish between long run and short run responses, be an autoregressive-distributed lag model ($\mathsf{AD}(1,1)$) of the form:

$$y_t = \alpha y_{t-1} + \beta z_t + \gamma z_{t-1} + \varepsilon_t \tag{7}$$

with:

$$z_t = \lambda z_{t-1} + v_t \tag{8}$$

when each of $\{\varepsilon_t\}$ and $\{v_t\}$ are identically independently distributed processes that are independent of each other with $\mid \alpha \mid < 1, \mid \lambda \mid < 1$, so that both y_t and z_t are stationary. Then in the static long run (*i.e.* when $y_t = y^*$ and $z_t = z^*\ \forall t$) the response of y_t to z_t is given by $\kappa = (\beta + \gamma)/(1 - \alpha)$, whereas the pseudo true value of the OLS estimator of δ is given by $\delta_0 = (\beta + \gamma\lambda)/(1 - \alpha\lambda)$. Hence inferences about the long run response of y_t to z_t based on the static regression model (6) will be invalid unless $\kappa = \delta_0$. Note that $\kappa = \delta_0$ if and only if $\gamma = -\alpha\beta$, which is the common factor restriction that ensures that the autoregressive distributed lag model (7) takes the form:

$$\begin{aligned} y_t &= \beta z_t + \epsilon_t \\ \epsilon_t &= \alpha\epsilon_{t-1} + \varepsilon_t \end{aligned} \tag{9}$$

so that all the dynamics is captured by a first order autoregressive error process. Also note that although $\lambda = 1$ appears to be another condition for $\kappa = \delta_0$, this possibility is ruled out by the assumption that z_t is stationary. That the static regression model (6) is not congruent with past sample information would be revealed by testing for the absence of serial correlation in the model residuals. The fact that congruence with information from this source can be tested via serial correlation tests, does not imply that when a model is found to have serially correlated residuals subsequent analysis should adopt estimation methods that allow for autoregressive (or even moving average) error processes since this will only be appropriate if $\gamma = -\alpha\beta$. Sargan [1964, 1980b] provides detailed discussion of this point, and further analysis is contained in Hendry and Mizon [1978], Hoover [1988], Mizon [1977a, 1993], Mizon and Hendry [1980] and Spanos [1988]. Mizon [1993] points out that there are also serious drawbacks to imposing invalid common factor restrictions in models for nonstationary variables, in particular integrated-cointegrated variables. This is illustrated in Section 4 below. In short, testing for congruence with past sample information is a way of checking the adequacy of the dynamic specification of the model (see the typology of univariate dynamic models in Hendry, Pagan and Sargan [1984] and that for multivariate dynamic models in Hendry [1993c]), but such tests will also have power against wrong functional form, regime shifts and other causes of parameter nonconstancy.

(ii) Relative Present Sample Information A model with errors that are not homoscedastic innovations with respect to the set of all current dated variables in the modeller's databank of relevant variables is not congruent, and can be improved by using information already available. For example, if a model's errors u_t are such that $\mathsf{E}(u_t\mathbf{x}_t) \neq 0$ then there is still information in $\mathbf{x}_t$ that is unexploited. This could arise from erroneously omitting variables from the model, or by conditioning on variables that are not weakly exogeneous for the parameters of interest. Another form of this type of noncongruence is heteroscedasticity in the errors (*i.e.* $\mathsf{E}(u_t^2) = \sigma_t^2 \neq \sigma^2 \forall t$), which implies that the distribution of the errors has features that are potentially explainable as functions of the $\mathbf{x}_t$. The distribution of the errors may also be skewed, leptokurtic, or platykurtic, and to the extent that these features of the error distribution are explainable as functions of the available data the model is noncongruent. Tests for omitted variables, for weak exogeneity (see Engle *et al* [1983], Ericsson [1993] and Richard [1980]), excess skewness or kurtosis in the error distribution (see Doornik and Hansen [1994]), and homoscedasticity (see e.g. Breusch and Pagan [1980] and White [1980]) are examples of ways in which a model's congruence with present sample information can be assessed.

(iii) Relative Future Sample Information For a model to be congruent with future sample information it must not suffer from predictive failure, or parameter nonconstancy. A model with characteristics that change with modest extensions to the estimation period is clearly failing to represent important features of the DGP, and is thus likely to yield poor inferences, whether these be parameter estimates, tests of hypotheses, or predictions. One of the essential characteristics for a model to have if it is to be valuable for inference, is that it captures fundamental invariants in the relationships between the variables under scrutiny *i.e.* represent the structure. A model with all its "parameters" transient is itself ephemeral, and inferences based on it are likely to appear whimsical when judged against subsequent observation. Indeed such a model could not be capturing underlying structure. A more durable model will have its parameter estimates approximately constant across varying estimation periods. Further, when a model is to be used for prediction of the likely effects of a change in policy (e.g. a change in tax rates, a move from fixed to floating exchange rates, or the introduction of a tight monetary policy), it is crucial that the model's parameters are invariant to the policy regime shift. Discussion of this issue has a long history, early contributors being Frisch and Haavelmo who in their analysis introduced the concept of autonomy (see Aldrich [1989]). Lucas [1976] presented a more recent view of the problem, which has spawned a large literature. Indeed, for many macroeconometricians the "Lucas critique" was sufficiently pernicious to have removed the credibility of much econometric modelling. However, if a model has conditioning variables that are weakly exogenous for the parameters of interest, and the latter are invariant to changes in the process generating the conditioning variables (the conditions for super exogeneity - see Engle *et al* [1983]), then the Lucas critique is rendered benign. An important implication of this statement is that the Lucas critique is brought into the realms of testable hypotheses as explained by Engle and Hendry [1993] and Favero and Hendry [1992] - see Ericsson [1993] for a recent appraisal. The main instruments of such testing are analysis of variance type parameter constancy and Chow [1960] prediction test statistics, as well as recursive estimation (see Brown, Durbin and Evans [1975], Hansen [1992], and Terasvirta [1970]).

(c) Measurement System It is important to know how variables are measured, and their specific properties. With this knowledge it is then possible to choose functional forms for models, or variable transformations, to ensure that the models will yield fitted and predicted values of the modelled variables that share these same properties. This is the requirement that the models be data admissible. For example, in modelling aggregate consumption which is known never to be negative, it makes sense to consider modelling the logarithm of consumption rather than its level in order to avoid generating negative fitted or predicted values. Similarly, a logarithmic or logit transformation of a variable like unemployment which is bounded between 0 and 1 (or equivalently between 0 and 100%) will guarantee that the model's fitted and predicted values share this property. Hence the congruence of a model with the measurement system requires it to be logically possible for the model to generate the observed and future data. This concept of congruence can be extended to incorporate the requirement that the model should be capable of representing empirically determined characteristics of the data such as stationarity, deterministic nonstationarity, integratedness, or seasonality. To choose models to be data admissible, and hence congruent with this source of information, is prudent rather than contentious, though the nature of the particular problem being analysed will determine the price of imprudence.

(d) Rival Models The importance of ensuring that a model is congruent with the information contained in rival models is twofold. Firstly, in economics there is usually no shortage of alternative theories for a given phenomenon. Secondly, it is possible for more than one model to be congruent with a priori theory, sample information and the properties of the measurement system. The fact that economics is a rich source of theories and hypotheses is a tribute to the economics profession's ingenuity and imagination, and only a problem if there is no serious attempt to discriminate between the competing models. Typically, alternative theories when implemented in econometric models use different information sets and possibly different functional forms, and are thus separate or nonnested models. It is this nonnested feature which enables more than one model to be congruent with respect to sample information - each can be congruent with respect to its **own** information set. Ericsson and Hendry [1989] analyse this issue and show that the corroboration of more than one model can imply the inadequacy of each, and Mizon [1989] provides an illustration. Hence to have a model which is congruent with respect to its own information set is a necessary condition for it to be exploiting that information, but it is not sufficient for it to be a dominant or encompassing model. An encompassing model is one which can account for the previous empirical findings which were thought to be relevant and adequate for the explanation of the variables and parameters of interest, and can explain the features of rival models being considered currently. The encompassing model renders other models inferentially redundant, and so is a dominant model.

The comparison of alternative models for the same phenomenon can be achieved by using one of the many statistics for testing nonnested hypotheses (e.g. Cox [1961,1962], Davidson and MacKinnon [1981] and Pesaran [1974]), but it is important to note that the statistical (size and power) properties of these test statistics are derived in the context of an embedding model - a point made clear by the encompassing interpretation of these test statistics (see Hendry and Richard [1989], Mizon [1984], and Mizon and Richard [1986]. Further, the use of nonnested test statistics in pairwise comparison of models can lead to apparent contradictions and is a non-transitive procedure. In particular, it is possible for the hypotheses $\mathsf{M}_1\mathcal{E}\mathsf{M}_2$ and $\mathsf{M}_2\mathcal{E}\mathsf{M}_3$ to be valid whilst the hypothesis $\mathsf{M}_1\mathcal{E}\mathsf{M}_3$ is invalid. Since this non-transitivity of encompassing arises essentially because the embedding model implicit in each of the three pairwise comparisons is different (often $\mathsf{M}_1\cup\mathsf{M}_2$, $\mathsf{M}_2\cup\mathsf{M}_3$ and $\mathsf{M}_1\cup\mathsf{M}_3$ respectively) the problem can be avoided by comparing all models relative to a general model that embeds or nests all of them. Letting M_c denote such an embedding model (i.e. $\mathsf{M}_i\subset\mathsf{M}_c\forall i$) it is then possible to test the nested hypothesis that M_i is an acceptable simplification of M_c for each i. If there is a model (M_1 say) that is a valid simplification of M_c then it parsimoniously encompasses M_c - denoted $\mathsf{M}_1\mathcal{E}_p\mathsf{M}_c$. In the population it is only possible for more than one model to parsimoniously encompass an embedding or completing model M_c if they are observationally equivalent, consequently, if $\mathsf{M}_1\mathcal{E}_p\mathsf{M}_c$ then $\mathsf{M}_1\mathcal{E}\mathsf{M}_j\forall j$. Therefore the requirement that a model be congruent with information contained in rival models is equivalent to that model parsimoniously encompassing a model that embeds all the rival models. This implies that the relevant sample information set for the determination of a congruent model, is the union of the information sets of the models implementing each competing theory. Unsurprisingly, this entails that if empirical evidence is to be used to evaluate and compare models rather than naively corroborate them, a modelling strategy that starts from a congruent general model and tests for valid simplifications or reductions of it, is likely to be an efficient way to find parsimonious yet durable representations of the salient features of the relationships between the variables of interest.

Finally, in this section on the evaluation of model congruence, note that testing for congruence with respect to sample information is misspecification testing, whereas testing for congruence with respect to rival model information is specification testing. In this context misspecification testing can be seen as a means of ensuring that there is a valid statistical basis for using specification testing to find a parsimonious encompassing model. Provided that the sample information set used ensures that all econometric models implementing relevant economic theories are nested within a general model that is congruent with the sample information (this is determined by misspecification testing or diagnostic checking of the general model), the model that parsimoniously encompasses the general model and is data admissible will be the dominant model that is congruent with information from all four sources (this is determined by specification testing or checking the validity of reductions from the congruent general model).

3.1.2 Encompassing

If parsimonious encompassing is an essential part of congruence what is the separate role for encompassing? As emphasized above, parsimonious encompassing is the property that a model be an acceptable reduction of a congruent embedding model, and as such is specification searching within a common general model. If after completing a modelling exercise and having determined which model is preferred further information becomes available, this raises the question as to whether the existing results are robust to this extension of the information set. The further information may take the form of new a priori theories which imply an extension of the information set to embrace the empirical models implementing these new theories. Testing the preferred model's ability to encompass the newly available rival models (which can be done using nonnested test statistics) is a form of misspecification testing. Conversely, the requirement that a model encompass its rivals ensures that the information set used for modelling is general enough to enable the empirical implementation of all the models. Hence each model is evaluated with respect to an information set more general than the minimum one required for its own implementation, thus achieving robustness to extensions of its information set in directions relevant for competing models.

3.2 Modelling Strategies

The previous section discussed the evaluation of econometric models, and emphasized the important roles of congruence and encompassing in this assessment. Indeed, the most important features of a model are its quality and suitability for its intended use, features which are assessed in the process of model evaluation and not endowed on it as a result of the process by which the model is developed. The route by which a model is discovered (including serendipity, brilliant intuition, and systematic application of a particular modelling strategy) does not determine model quality, though it does affect research efficiency - see Hendry and Mizon [1990] for further discussion of these points. Although a model with a good pedigree (e.g. derived from high quality economic theory by a leading practitioner) is more likely to be valuable than one without, such a pedigree is neither necessary nor sufficient for the discovery of a congruent and encompassing model. This is fortunate since otherwise there would be little or no scope for the discovery of new theories and classes of model.

Whilst there is no unique way to find congruent and encompassing models, some modelling strategies are more likely to do so, and to do so efficiently, than others. Of the many strategies that might be adopted in modelling attention here is confined to a contrast between 'specific-to-general' and 'general-to-specific' modelling - Granger [1990] contains discussions of other possibilities.

3.2.1 'Specific-to-General' Versus 'General-to-Specific' Modelling

In specific-to-general modelling the starting model is both very specific and simple, often implementing a particular narrowly defined economic theory. This approach it has been argued has the twin advantages of avoiding unnecessary generality (which results in loss of efficiency and can result in models being too finely tuned to the sample data), and guarantees the coherence of the model with economic theory thus ensuring that the model has an economic interpretation and is related to the phenomena being studied. An important issue with such modelling is the extent and means of model evaluation adopted. The Average Economic Regression approach described by Gilbert [1986] seeks corroboration of the underlying economic theory in terms of signs, magnitudes and significance of estimated parameters, and does not question the adequacy of the model as a whole relative to the available information. However, no matter how elegant and sophisticated an economic theory is there is no guarantee that the model implementing it will be statistically well specified, and yet without this inferences drawn from the model will be invalid. In short, the essence of the argument against specific-to-general modelling is that a priori theory is not endowed with empirical veracity and so econometric models must be scrupulously evaluated for congruence with information from sources additional to a priori theory, but tests performed in a non-congruent framework are themselves invalid in general. Further, it is possible to find more than one congruent model when the congruence of each model is with respect to its own implicit or explicit information set - see Mizon [1989] for an illustration.

Although it is possible to test the statistical adequacy of simple models, once a model is found to be inadequate, specific-to-general modelling is a seriously flawed strategy for discovering other models that retain economic theory consistency and also achieve coherence with the data. In this regard the alternative of a general-to-specific strategy has many advantages, such as those associated with the following arguments listed by Hendry and Doornik [1994]: "directed versus directionless strategies; validly interpreting intermediate test outcomes by avoiding later potential contradictions; escaping the *non sequitur* of accepting the alternative hypothesis when a test rejects the null; determining the baseline innovation error process on the available information; and circumventing the drawbacks of correcting manifest flaws when these appear as against commencing from a congruent model". Further discussion of these points can be found *inter alia* in Hendry [1983], and Mizon [1977b, 1989]. Hendry and Doornik [1994] provide ten further logical and methodological arguments in favour of general-to-specific modelling, and for commencing econometric analyses by modelling the joint density of all relevant variables.

Mizon [1993], in the context of a DGP given by the partial adjustment model:

$$y_t = \beta z_t + \alpha y_{t-1} + \varepsilon_t \tag{10}$$

with z_t generated by (8), illustrated the failure of a specific-to-general search starting from a static regression as in (6) to find an adequate model, whereas a general-to-specific strategy involving reduction from the congruent general model given by an AD$(1,1)$ model of the form (7) easily led to the selection of the DGP as the only model that is congruent and parsimoniously encompasses (7). In addition to being an efficient way to find a congruent encompassing model, the general-to-specific modelling strategy applied in Mizon [1993] was shown to avoid the statistically anomalous result of an algebraically more general model (9) failing to encompass the static regression (6) which is apparently nested within it. The failure of (9) to encompass the simpler model (6) arises because (9) involves the auxiliary COMFAC hypothesis $\alpha\beta + \gamma = 0$, which is not entailed in the specification of (6). However, adopting a general-to-specific strategy and only testing for simplifications within a congruent general model avoids this anomaly since (9) is clearly revealed to be non-congruent in that the hypothesis $\alpha\beta+\gamma = 0$ is rejected and (9) has serially correlated errors. Hendry and Doornik [1993] also contains a discussion of this anomaly and the use of a general-to-specific modelling strategy to avoid it.

Much of the early research done by members of the LSE School was concerned with single equation analysis, and in particular the choice of an appropriate dynamic specification. Single equation models of necessity involve conditioning on some variables, and hence weak exogeneity assertions. The necessity of ensuring that conditioning variables are weakly exogenous for the parameters of interest was always emphasized - see e.g. Hendry and Richard [1982, 1983]. The testing of weak exogeneity assertions also involves modelling from the general to the specific, and an important example of this point is provided by the analysis in Hendry [1993]. Case (d) of his analysis considered a situation in which a model of the conditional distribution $\mathsf{D}(y_t \mid z_t, y_{t-1}, z_{t-1})$ corresponded exactly to the an equation of the DGP. Indeed it embodied the conditional expectation $\mathsf{E}[y_t \mid z_t, y_{t-1}, z_{t-1}] = \beta z_t$, and yet inference on β based on OLS estimation was inefficient as a result of z_t not being weakly exogenous for β. This came about because the parameters β and σ_1^2 of the conditional distribution $\mathsf{D}(y_t \mid z_t, y_{t-1}, z_{t-1}; \beta, \sigma_1^2)$ were linked directly to the parameters of the marginal distribution $\mathsf{D}(z_t \mid y_{t-1}, z_{t-1}; \beta, \rho, \sigma_2^2)$ when the DGP had the form:

$$\begin{aligned} y_t &= \beta z_t + \epsilon_{1t} \\ \Delta z_t &= \rho\,(y_{t-1} - \beta z_{t-1}) + \epsilon_{2t} \end{aligned} \tag{11}$$

with

$$\begin{pmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{pmatrix} \sim \mathsf{NI}\left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \sigma_1^2 & 0 \\ 0 & \sigma_2^2 \end{pmatrix}\right] \tag{12}$$

This example, and others presented in Hendry [1993a] in the context of a bivariate cointegrated system, illustrate the crucial role of weak exogeneity in sustaining valid single-equation inference with cointegrated processes, despite the fact that the OLS estimator is super consistent in that context. This example indicates that whether or not variables are weakly exogenous for the parameters of interest depends on the properties of the joint distribution of all variables. Hence it is not surprising that members of the LSE School also paid attention to the issues of model congruence and encompassing, and the desirability of using a general-to-specific modelling strategy, in a systems context.

Although the importance of carefully considering the dynamic specification of single equation econometric models had been a crucial part of the LSE methodology since Sargan [1964], its relevance for systems of equations was also realized. Discussions of dynamic specification in systems is contained in Hendry [1971, 1974], Hendry and Anderson [1977] and Mizon [1977a]. However, the challenge to the econometric modelling of macroeconomic time series provided by the work of Box and Jenkins [1970] and Granger and Newbold [1974], stimulated other contributions. In particular, Prothero and Wallis [1976] analysed the univariate implications of systems of simultaneous equations, drawing attention to the particular form taken by the implied autoregressive integrated moving average (ARIMA) time series models - also see Zellner and Palm [1974]. These and other contributions, for which see the discussion in Hendry, Pagan and Sargan [1984], were concerned with the specification, estimation and testing of multivariate econometric models with respect to their own information sets, and did not raise the issue of one structural econometric model encompassing its rivals. Noting that the typical specification of a dynamic linear structural econometric model (SEM), based on the pioneering work of the Cowles Commission researchers (see Koopmans [1950]), has the generic form:

$$\begin{aligned} B\mathbf{y}_t + C\mathbf{z}_t + D\mathbf{x}_{t-1} &= \mathbf{u}_t \\ \mathbf{u}_t &\sim \mathsf{NI}(0,\ \Omega) \end{aligned} \tag{13}$$

when $\mathbf{y}_t$ are the endogenous variables being modelled, $\mathbf{z}_t$ are weakly exogenous variables which are not modelled, $\mathbf{x}'_{t-1} = (\mathbf{y}'_{t-1}, \mathbf{z}'_{t-1})$ are the lagged values of both endogenous and exogenous variables, and $\mathbf{u}_t$ is the vector of unobserved errors, it is clear that the underlying distribution is $\mathsf{D}(\mathbf{y}_t \mid \mathbf{z}_t, \mathbf{x}_{t-1})$ which corresponds to the unrestricted reduced form that in conventional notation can be written:

$$\begin{aligned} \mathbf{y}_t &= \Pi_z\mathbf{z}_t + \Pi_x\mathbf{x}_{t-1} + \mathbf{v}_t \\ \mathbf{v}_t &\sim \mathsf{NI}(0, \Sigma) \end{aligned} \tag{14}$$

As is well known, the parameters of (13) are unidentified without a priori restrictions, and so assuming that the specification of (13) asserts more restrictions than are required to achieve identification, it follows that the specification of (13) implies the hypothesis:

$$\begin{aligned} \Pi_z = -B^{-1}C, &\quad \Pi_x = -B^{-1}D \\ \Sigma &= B^{-1}\Omega(B^{-1})' \end{aligned} \tag{15}$$

which embodies the overidentifying restrictions. Hence an important part of the evaluation of a structural econometric model is a test of these overidentifying restrictions (see Anderson and Rubin [1950] and Sargan [1988]). Richard [1980] pointed out that the restrictions on Π_z and Π_x are equivalent to the following hypothesis concerning the conditional mean of $\mathbf{y}_t$:

$$B\mathsf{E}(\mathbf{y}_t \mid \mathbf{z}_t, \mathbf{x}_{t-1}) + C\mathbf{z}_t + D\mathbf{x}_{t-1} = 0 \; \forall \; t \tag{16}$$

and Mizon [1984] provided an encompassing interpretation of these and the covariance restrictions as the restrictions required for (13) to parsimoniously encompass (14). However, the identification of (13) relative to (14), and the parsimonious encompassing hypothesis (15) or (16), evaluate the SEM (13) relative to its own information set.

For linear dynamic systems, when there is more than one structural econometric model, and it is desired to test the exogeneity status of some variables in the system, an appropriate framework for evaluation is the VAR for the union of all variables in the structural models, plus any further variables required to achieve a congruent VAR. For the variables $\mathbf{x}_t$ a k^{th} order VAR takes the form:

$$\mathbf{x}_t = \sum\nolimits_{j=1}^{k} A_j \mathbf{x}_{t-j} + \varepsilon_t$$

which has proved to be a valuable way to model the joint distribution of $\mathbf{x}_t$ conditional on its history $\mathsf{D}(\mathbf{x}_t \mid \mathbf{x}_0, \mathbf{x}_1, ...\mathbf{x}_{t-1})$. Monfort and Rabemananjara [1990] for stationary $\mathbf{x}_t$ illustrated the use of a VAR as a general model within which to test the exogeneity status of a susbset of variables, and to test particular structural hypotheses. Hendry and Mizon [1993] proposed the use of a congruent VAR when $\mathbf{x}_t$ is nonstationary (e.g. $\mathbf{x}_t \sim \mathsf{I}(1)$ and/or the process generating $\mathbf{x}_t$ has deterministic nonstationarities such as trends and regime shifts) as a general statistical framework within which to test hypotheses about the dimension of cointegrating space following Johansen [1988], as well as hypotheses concerning dynamic simplification (e.g. Granger non-causality), weak exogeneity, and the ability of particular structural econometric models to parsimoniously encompass the VAR. Note that in this framework the evaluation of each SEM is relative to the information set supporting the VAR, and not just the SEM's own information set. Clements and Mizon [1991] and Hendry and Doornik [1994] contain applications of the general-to-specific modelling strategy for $\mathsf{I}(1)$ systems proposed by Hendry and Mizon [1993], and Section 4.2 provides an illustration using an updated databank for a subset of the variables modelled in Clements and Mizon [1991].

Just as the importance of evaluating models by checking their congruence and encompassing abilities, is not confined to single equation models, it is not limited to models for time series variables. It is just as important for cross section and panel data models that their congruence be tested, and that the encompassing properties of alternative models are assessed, rather than simply using empirical analysis to calibrate and confirm particular economic theories. Research devoted to furthering the development of model evaluation criteria and test statistics for microeconometric models, analogous to those developed for macroeconomic time series models, could yield high returns in terms of deepening and putting on a firmer basis our understanding of microeconomic relationships. Since general tools for estimation and hypothesis testing are available, the need is for a taxonomy of information sets relevant for the evaluation of microeconometric models, and the development of frameworks sufficiently general to enable the comparison of alternative models. Similarly, general-to-specific modelling strategies can be expected to be valid, effective and efficient ways to develop congruent and encompassing models in microeconometrics.

Though the description of the main components of the LSE methodology presented above is largely concerned with the theory of model evaluation and theoretical aspects of modelling strategies, it is important to note that a major part of the research that led to the development of the methodology was firmly embedded in applied econometric studies. Such empirical econometric studies, in addition to being valuable directly in tackling the issues on which they are focused (if they are successful), and providing a test bed to prove the worth of existing econometric techniques and modelling strategies, are a rich source of problems requiring new theoretical results. The next section contains an empirical study of the relationship between wages, prices and unemployment in the UK which provides an illustration of LSE-style modelling, and draws attention to problems concerning the inter-relationship between the modelling of nonstationary variables, regime shifts and forecasting.

4 An Illustration: Wages, Prices and Unemployment in the UK.

As a simple illustration of modelling (both single equation and system) within the spirit of the LSE methodology, three variables are taken from an updated data set containing the same variables as Clements and Mizon [1991]. Quarterly data for the three variables, ln wages w (strictly, this variable is defined as the ln of earnings per employee but will be referred to as wages), ln of the retail price index p, and ln of the percentage unemployment rate u, are available for the period 1965(1) to 1993(1). The data for w and u are seasonally adjusted,

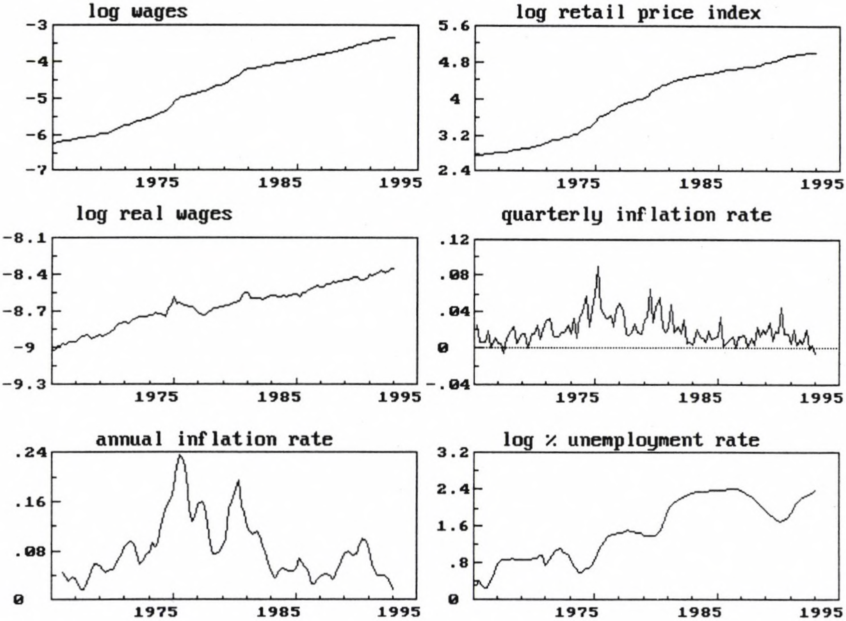

Figure 1: Data Plots 1965(1) - 1993(1)

but that for p and hence the inflation rate Δp are not. Precise definitions and sources are given in the Data Appendix.

Visual inspection of these graphs indicates that both w and p have strong positive trends of a similar magnitude, so that they might be modelled well as stationary deviations from a linear trend, or alternatively as variables with stochastic trends within a cointegrated-integrated system (see the discussion below and Banerjee et al [1993] for a fuller description). However, this is allowing the visually clear domination of trend in both w and p to dictate the choice of the potential components for analysing the structure of the series. u also has a strong upward trend, but the deviations from this exhibit cyclical movements which it will be important to capture in the modelling of these variables. That cyclical features, possibly resulting from changes in economic policies (e.g. fixed or floating exchange rates, accommodating or tight monetary policies) and 'autonomous' events (e.g. the movement in commodity prices

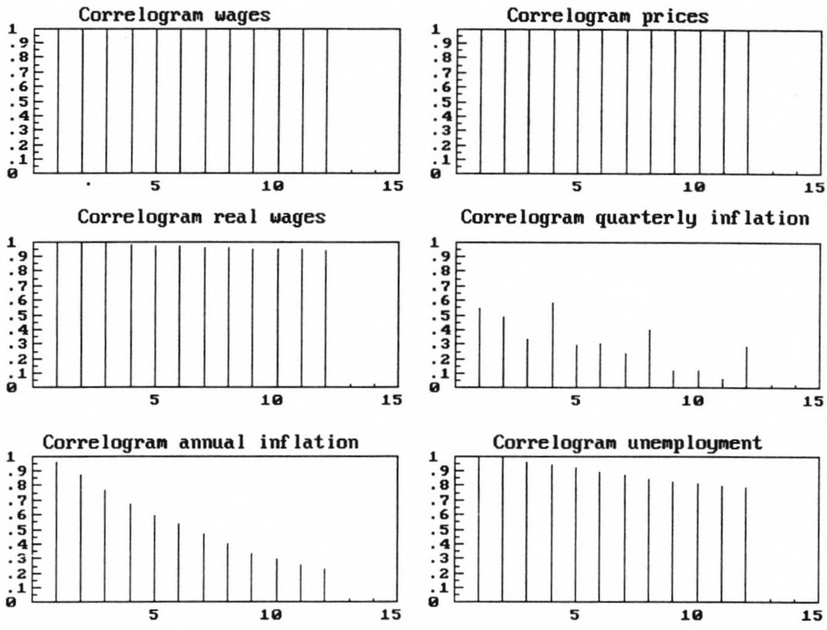

Figure 2: Correlograms of data

(including the price of crude oil) in the late 1960's and early 1970's) may be important is also suggested by inspection of the graphs of the ln of real wages $(w-p)$ and the quarterly and annual inflation rates Δp and $\Delta_4 p$ respectively, when $\Delta = (1-L)$ is the first difference operator and $\Delta_4 = (1-L^4)$ is the fourth difference operator.

Inspection of the correlograms for the variables $w, p, (w-p)$ and u reveals high first order serial correlation coefficients ($\cong 0.99$) with the higher order coefficients declining extremely slowly, and correspondingly their spectra have peaks at the zero frequency. This information is consistent with each of these series being nonstationary, and probably integrated of order one (i.e. $\mathsf{I}(1)$) which means that it is necessary to difference the variables in order to remove their stochastic nonstationarity and render them $\mathsf{I}(0)$. Indeed, the correlogram for Δp suggests that quarterly inflation is $\mathsf{I}(0)$, but with a strong seasonal pattern. The correlogram for $\Delta_4 p$ on the other hand exhibits no seasonal behaviour, and is consistent with this variable having a stationary first order autoregressive representation. As an illustration of the use of univariate test statistics for unit roots the following table provides the values of Augmented Dickey-Fuller test statistics which take the form of the 't' statistic for the hypothesis $\alpha = 0$ in the regression model for the generic variable y_t:

$$\Delta y_t = \alpha y_{t-1} + \mu + \textstyle\sum_{j=1}^{m} \beta_j \Delta y_{t-j} + \xi_t \tag{17}$$

Note that when $\alpha = 0$ equation (17) is a regression in the differences Δy_t corresponding to y_t being well modelled as a an $\mathsf{I}(1)$ process, and thus having a unit root in its autoregressive representation. The distribution of the 't' statistic when $\alpha = 0$ is nonstandard (i.e. neither Student's t, nor limiting normal), but tables of critical values were provided by Dickey and Fuller [1981]. Although, as pointed out by West [1988], the limiting distribution of this 't' statistic is normal when there is drift in the series ($\mu \neq 0$), the small sample distribution is better approximated by the Dickey-Fuller distribution unless μ/σ_ξ is "large" - see Hylleberg and Mizon [1989].

Augmented Dickey-Fuller statistics

	$t(ADF)$	lag length m
w	-1.27	2
u	-1.77	1
p	-1.31	1
$(w-p)$	-1.10	3
Δw	-3.30^{*}	1
Δp	-3.83^{**}	0
Δu	-4.95^{**}	0

In calculating these statistics the maximum lag length m was 5, and the reported results are for the 't' statistic corresponding to the longest significant lag within this maximum. In the calculation of the statistics for p and Δp only seasonal dummy variables were included in the regression (17). Statistics significant at 5% are denoted by * and those significant at 1% by **. On the basis of these Augmented Dickey-Fuller statistics all three variables w, p and u appear to be $\mathsf{I}(1)$ - the null hypothesis of $\mathsf{I}(0)$ being rejected against the alternative of $\mathsf{I}(1)$ and the null of $\mathsf{I}(1)$ not being rejected against the alternative of $\mathsf{I}(2)$ for each variable. In addition, the real wage $(w-p)$ is $\mathsf{I}(1)$ and so w and p do not cointegrate as the real wage. In fact, the hypothesis that each of these variables has a unit root was not rejected when tested against the alternative of trend stationarity, i.e. $\alpha = 0$ was not rejected even when a linear trend was added to the regression (17). Hence in the modelling of w, p and u it will be important to choose models that can represent their nonstationarity, with the possibility that all three variables form a cointegrating relationship i.e. there is some linear combination of w, p and u that is $\mathsf{I}(0)$. Another characteristic of wages and prices is that they are non negative, so the use of linear models in the logarithmic transformations of them is data admissible i.e. cannot produce negative fitted or predicted values. Similarly, the unemployment rate is bounded between 0% and 100%, so that a logit transformation might be appropriate for modelling it in a data admissible way. However, since all observations on the unemployment rate are below 13% the use of its logarithm should be data admissible, and this is borne out by noting that the graphs of u and the logit transformation of U are essentially identical when adjusted for their different means.

Having briefly analysed the univariate statistical properties of the variables to be analysed attention is now turned to a review of the Phillips curve literature in so far as it is relevant to the subsequent analysis.

4.1 Single Equation Analysis

Noting the emphasis placed above on adopting a general-to-specific modelling strategy, an explanation of why single equation analysis precedes system analysis is in order. Firstly, given that the primary purpose of the empirical analysis is to illustrate the LSE methodology, and that the methodology has important implications for system and single equation modelling it is relevant to illustrate both. Secondly, there are pedagogical benefits to presenting the simpler case before the general, and the intention is to exploit these.

Many past and present members of the LSE Economics Department have analysed models of wage and price determination (e.g. Desai [1975, 1984], Espasa [1975], Layard and Nickell [1985], Lipsey [1960], Nickell [1984], Phillips [1958], Sargan [1964, 1980a], and Vernon [1970]) and so it seems particularly appropriate to present the results of re-analysis of this relationship with the present data set. The first and most influential piece of research in this area was that of A.W.H. Phillips in which an empirical relationship representing a trade-off between money wages and unemployment was described. As Desai [1984, p.253] points out the Phillips curve "has been re-specified, questioned, rejected as being unstable and reinterpreted", and yet it is still evident in many models of aggregate labour market and wage and price behaviour. The Colston paper of

Sargan [1964], though primarily concerned with a number of methodological issues which had a major influence on LSE econometricians (see Hendry and Wallis [1984]), developed single equation models of wage and price determination with the distinction between equilibrium and disequilibrium behaviour explicit. Indeed, the foundations of the error correction model, which is now so prevalent in the analysis of nonstationary-integrated data, were laid in this paper. An additional feature of Sargan's paper was the extension of the Phillips framework to allow for the impact of productivity. The results presented below however are confined to describing relationships between wages, prices, and unemployment. This is done in order to keep the empirical modelling as simple as possible, whilst still providing an illustration of many of the important points made in the previous discussion of the LSE methodology. In order to increase the economic and statistical credibility, as well as the applicability, of the model it may well be necessary to enlarge the system to include productivity, hours of work, exchange rates and interest rates. In addition, it may be important to pay detailed attention to the changes that have taken place in the UK labour market since 1979 e.g. the increase in part-time working, the increase in self employed labour relative to employees, and the increased participation of women in the labour force. Thus the class of model considered in this section can be written in error correction form as:

$$\begin{aligned} \textstyle\sum_{j=0}^{k} \alpha_{1j} \Delta w_{t-j} &= (\alpha_0 + \gamma\beta_0) + \textstyle\sum_{j=0}^{k} (\alpha_{2j}\Delta p + \alpha_{3j}\Delta u)_{t-j} \\ & \quad -\gamma(w - \beta_1 p - \beta_2 u)_{t-1} + \lambda t + error \end{aligned} \tag{18}$$

when k is the maximum lag and with the normalization $\alpha_{10} = 1$. A linear trend is included in (18) as a proxy for omitted variables such as average labour productivity which are dominated by trend. The steady state solution of (18), in which Δw^*, Δp^* and Δu^* are the constant quarterly growth rates in nominal wages, retail prices and the unemployment rate respectively, is given by:

$$\mathsf{E}(w_t \mid p_t, u_t) = (\beta_0 + \beta_1 p_t + \beta_2 u_t) + (\lambda^* t + \alpha_1^* \Delta w^* - \alpha_0^* - \alpha_2^* \Delta p^* - \alpha_3^* \Delta u^*) \tag{19}$$

when:

$$\lambda^* = \lambda/\gamma \quad \alpha_0^* = \alpha_0/\gamma \quad \alpha_i^* = \left(\textstyle\sum_{j=0}^{k} \alpha_{ij}\right)/\gamma, \; i = 1, 2, 3 \tag{20}$$

The first term of (19) gives the static equilibrium response of w to the conditioning variables p and u and takes the form $\mathsf{E}(ecm_t \mid p_t, u_t) = 0$ when ecm_t is the observed disequilibrium defined as:

$$w_t - \beta_0 - \beta_1 p_t - \beta_2 u_t = ecm_t \tag{21}$$

Note that the defining characteristics of the static equilibrium are zero growth rates ($\Delta w^* = \Delta p^* = \Delta u^* = 0$) and no autonomous trend ($\lambda = 0$). When the growth rates are not zero equation (19) can be re-written as:

$$\begin{aligned} \mathsf{E}(w_t \mid p_t, u_t) &= (\beta_0 + \beta_1 p_t + \beta_2 u_t) + \lambda^* t + (\alpha_1^* \lambda^* - \alpha_0^*) \\ &\quad + (\alpha_1^* \beta_1 - \alpha_2^*)\Delta p^* + (\alpha_1^* \beta_2 - \alpha_3^*)\Delta u^* \end{aligned} \tag{22}$$

to ensure that the growth rate of nominal wages Δw^* is consistent with the equilibrium growth rate ($\Delta w^* = \beta_1 \Delta p^* + \beta_2 \Delta u^* + \lambda^*$) when for this single equation analysis p_t and u_t are presumed to be weakly exogenous variables. In the case that the disequilibrium $ecm_t \neq 0$ equation (18) can be interpreted as an equilibrium correction mechanism. Indeed, w responds to the disequilibrium (i.e. equilibrium corrects) with adjustment coefficient γ. With the equilibrium correction interpretation of (21) it is expected that $\beta_1 > 0$, and the hypothesis that real wages respond to the strength of demand in the labour market as measured by the unemployment rate is given by $\beta_1 = 1$ (the real wage hypothesis). Noting that w, p and u are $\mathsf{I}(1)$, and that $(w - p)$ also appears to be $\mathsf{I}(1)$, the real wage hypothesis can only hold if there is a cointegrating vector involving all three variables which has the sum of the coefficients of w and p equal to zero. The hypothesis that the strength of demand in the labour market has no effect on wages is given by $\beta_2 = 0$. This hypothesis does not preclude labour market effects on wages altogether since changes in unemployment Δu will influence w provided that $\alpha_3^* \neq 0$.

For OLS inference applied to single equations such as (18) to be valid it is important that the regressors are weakly exogenous for the parameters of interest, a point that was stressed by members of the LSE School in their discussions of single equation analysis - see for example Hendry and Richard [1982, 1983]. The importance of weak exogeneity as a necessary condition for efficient inference in this context implies that the testing of it is crucial. However, it is necessary to specify more about the joint distribution of all variables in the analysis in order to be able to test the hypothesis that a subset of them are weakly exogenous for the parameters of interest. Hence this issue will be addressed in Section 4.2 where a model of the joint determination of all variables is developed.

Preliminary analysis using the full sample of data to estimate an autoregressive-distributed lag model for w in terms of p and u with eight lags on all variables (corresponding to $k = 7$ in the error correction model (18)), plus a linear trend and seasonal dummy variables, revealed large outliers at the end of 1974 and beginning of 1975 when there was an 'explosion' of wages following the relaxation of wage and price restraint. Including the dummy variables: $D745$ (which takes the value 1 in 1974(3) and 1974(4), 2 in 1975(1) and zero elsewhere); and *Policy* (which takes the value 1 in 1968(2) and 1969(2) both of which were quarters when strong deflationary Budgets were introduced, and the value −1 in 1980(3) and 1980(4) to capture the effects of the recession induced by the Thatcher government's tight monetary policy, and zero elsewhere) produced a congruent model for the full sample which could be simplified by excluding the seasonal dummy variables and reducing the maximum lag length to 5. The following table gives the unrestricted estimates and their standard errors in parentheses for an ECM with $k = 4$.

Table 1. OLS estimates: general model for Δw_t 1966(2)-1993(1)

$\Delta w_t =$	-2.055	$+0.070\Delta w_{t-1}$	$+0.271\Delta w_{t-2}$	$+0.088\Delta w_{t-3}$	$-0.023\Delta w_{t-4}$
	(0.53)	(0.09)	(0.08)	(0.09)	(0.09)
	$+0.429\Delta p_t$	$+0.089\Delta p_{t-1}$	$-0.013\Delta p_{t-2}$	$+0.267\Delta p_{t-3}$	$+0.017\Delta p_{t-4}$
	(0.11)	(0.09)	(0.09)	(0.09)	(0.10)
	$-0.029\Delta u_t$	$+0.003\Delta u_{t-1}$	$-0.022\Delta u_{t-2}$	$+0.002\Delta u_{t-3}$	$+0.0001\Delta u_{t-4}$
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
	$-0.240 w_{t-1}$	$+0.204 p_{t-1}$	$+0.004 u_{t-1}$		
	(0.06)	(0.05)	(0.01)		
	$+0.002t$	$+0.038D745$	$-0.028Policy$		
	(0.0006)	(0.005)	(0.006)		

Diagnostic Statistics

$R^2 = 0.768$	$\sigma = 0.0099$	$V = 0.325$
$J = 2.581$	$AR(5,82) = 0.712$	$ARCH(4,79) = 0.642$
	$[p > 0.62]$	$[p > 0.63]$
$N = 0.042$	$H(40,46) = 0.756$	$R(1,86) = 0.851$
$[p > 0.98]$	$[p > 0.82]$	$[p > 0.36]$

The reported test statistics are: $AR(5,.)$ a Lagrange multiplier (LM) test statistic for fifth order serial correlation in the residuals which under the null of no serial correlation has a $\chi^2(5)$ distribution which is presented in the form that has an approximate $F(5,.)$ null distribution; $ARCH(4,.)$ an LM statistic for testing fourth order autocorrelated squared residuals which under the null of no autoregressive conditional heteroscedasticity (no $ARCH$) has a $\chi^2(4)$ distribution that is presented in the form that has an approximate $F(4,.)$ null distribution; $R(1,.)$ the regression specification test ($RESET$) which tests the null of correct specification against the alternative that the residuals are correlated with the squared fitted values of the regressand and has an $F(1,.)$ null distribution; $H(.,.)$ a statistic for testing the null hypothesis of unconditional homoscedasticity against the alternative that the residuals are correlated with the model's regressors and their squares, and has an approximate $F(.,.)$ null distribution; N a new statistic (see Doornik and Hansen [1994]) for testing the null of normal skewness and kurtosis which has a $\chi^2(2)$ under the null that the residuals are normally distributed; V the residual variance instability and J the joint (variance and regression coefficients) instability test statistics based on the cumulative backward score statistics of Hansen [1992]. When appropriate the p-value is reported in [.] after a statistic. Doornik and Hendry [1994] give further details of these statistics.

On the basis of these statistics there is no evidence of misspecification. This is substantiated by the following graphs in Figure 3 of actual and fitted values, the residuals, the residual correlogram, and the residual frequency plot and density.

Further, re-estimation of the model by recursive least squares, starting from an initial sample of 50 observations and sequentially increasing the sample from 50 to 108, revealed no serious indications of overall nonconstancy. The graphs of the recursively computed 1-step ahead residuals bordered by $\pm$ twice their standard errors, and the 1-step ahead Chow, breakpoint-F and forecast-F test statistics in figure 4 confirm this, though there is a suggestion of some non-constancy around 1984 which was the year of the miners' strike. Doornik and Hendry [1992, 1994] provide the definitions of these statistics, as well as discussing their role in modelling. However, the recursively computed estimates of the *constant* and the coefficients of w_{t-1}, p_{t-1} and u_{t-1} show some evidence of change, starting in 1979 and stabilizing at new values after 1984.

Hence this general model appears to be reasonably congruent with the available information and so forms a valid basis for testing for further simplifications, and testing other hypotheses. The real wage hypothesis $\beta_1 = 1$ is rejected at 1% significance since the test statistic for the hypothesis that the coefficients of w_{t-1} and p_{t-1} sum to zero is $F(1,87) = 5.00$ [$p > 0.028$]. However, the hypothesis that $\beta_3 = 0$ is not rejected ($F(1,87) = 0.403$ [$p > 0.53$]), so that wages appear to be unaffected by the level of unemployment. For this model the solved ecm_t with autonomous growth added is:

$$ecm_t = w_t \underset{(0.12)}{+8.57} \underset{(0.05)}{-0.85p_t} \underset{(0.02)}{-0.02u_t} \underset{(0.001)}{-0.008t} \tag{23}$$

which describes a highly significant (the Wald test statistic for the hypothesis

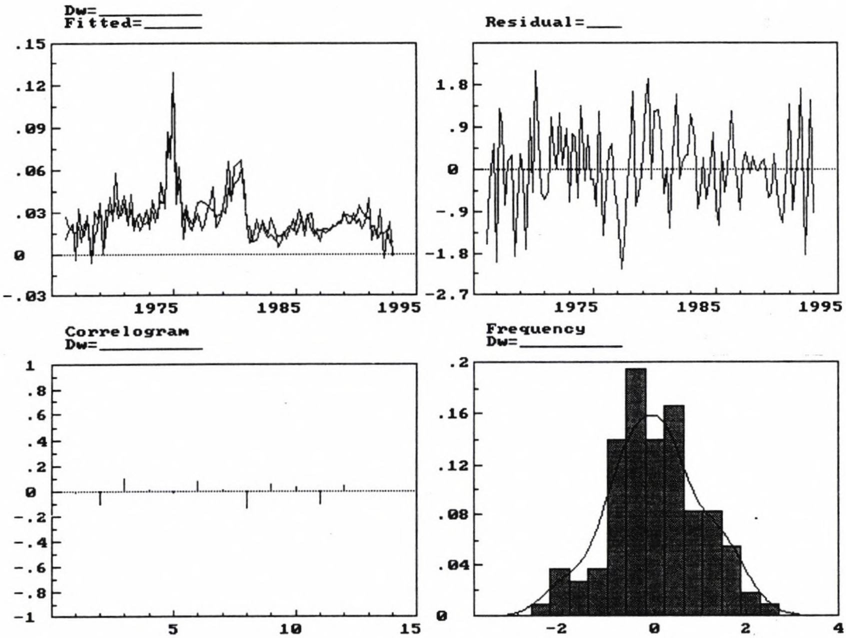

Figure 3: OLS Wage Equation 1966(2) - 1993(1).

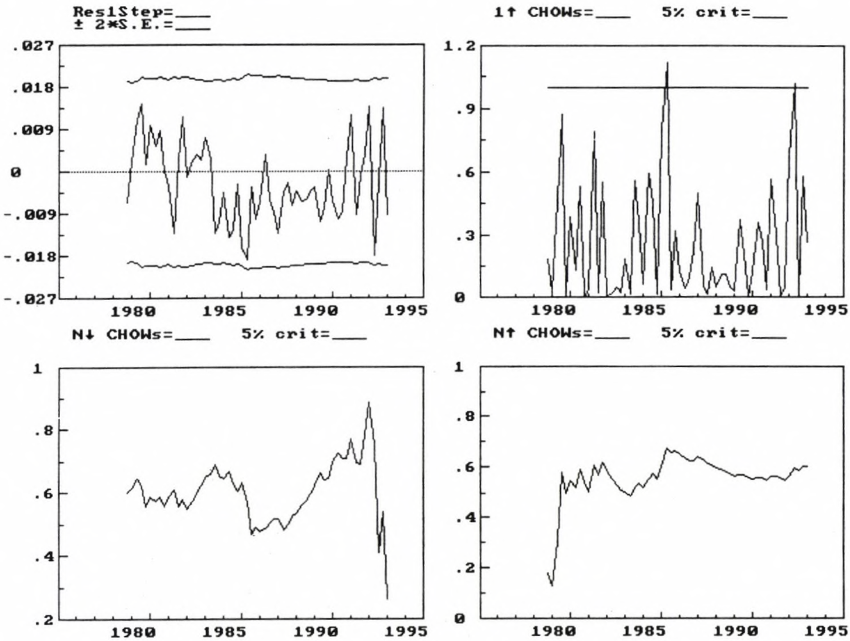

Figure 4: Recursive Graphics

that all coefficients including those of the 2 dummy variables but excluding the constant are zero $\chi^2(5) = 47155$ has a p-value of zero) equilibrium correction mechanism between wages, prices, unemployment and trend. However, unemployment u appears to have no role in the ecm_t, which also corresponds to a mechanism for determining nominal, rather than real, wages. The fact that $\beta_1 = 1$ is rejected, but $\beta_3 = 0$ is not, suggests that w and p might cointegrate when a linear trend and the dummy variables $D745$ and $Policy$ are included in the modelling. If this is the case though, the cointegrating vector would have the form $(1, -\kappa)$ with $\kappa \neq 1$ since the real wage hypothesis is rejected. However, these comments are predicated on the assumption that p and u are weakly exogenous variables for the parameters of interest in the wage equation, which the analysis in Section 4.2 shows to be doubtful.

The model reported in Table 1 clearly can be simplified, and the testing of a series of reductions from this congruent general model led to the following simple model that parsimoniously encompasses the general model ($F(10, 87) = 0.50$ [$p > 0.89$]). Not only does the model in Table 2 parsimoniously encompass that in Table 1, it is congruent as indicated by the diagnostic statistics, and the graphical analysis in Figure 5.

Table 2. OLS estimates reduced Δw model: 1966(1) - 1993(1)

$\Delta w_t =$	$0.284\Delta w_{t-2}$	$+0.499\Delta p_t$	$+0.260\Delta p_{t-3}$	$-0.031\Delta u_t$	$+0.0017t$
	(0.07)	(0.09)	(0.08)	(0.02)	(0.0004)
-1.80	$-0.210w_{t-1}$	$+0.181p_{t-1}$	$-0.0001u_{t-1}$	$+0.038D745$	$-0.03Policy$
(0.37)	(0.04)	(0.04)	(0.005)	(0.004)	(0.005)

Diagnostic Statistics

$R^2 = 0.753$	$\sigma = 0.0097$	$V = 0.299$
$J = 1.960$	$AR(5, 93) = 0.965$	$ARCH(4, 90) = 1.152$
	[$p > 0.44$]	[$p > 0.34$]
$N = 0.228$	$H(20, 77) = 1.345$	$R(1, 97) = 0.621$
[$p > 0.89$]	[$p > 0.18$]	[$p > 0.43$]

In addition, this reduced model still indicates that there is no role for the level of unemployment in the explanation of nominal wages, and so Table 3 presents the results for the final simplification which eliminates u (but not Δu) from the model.

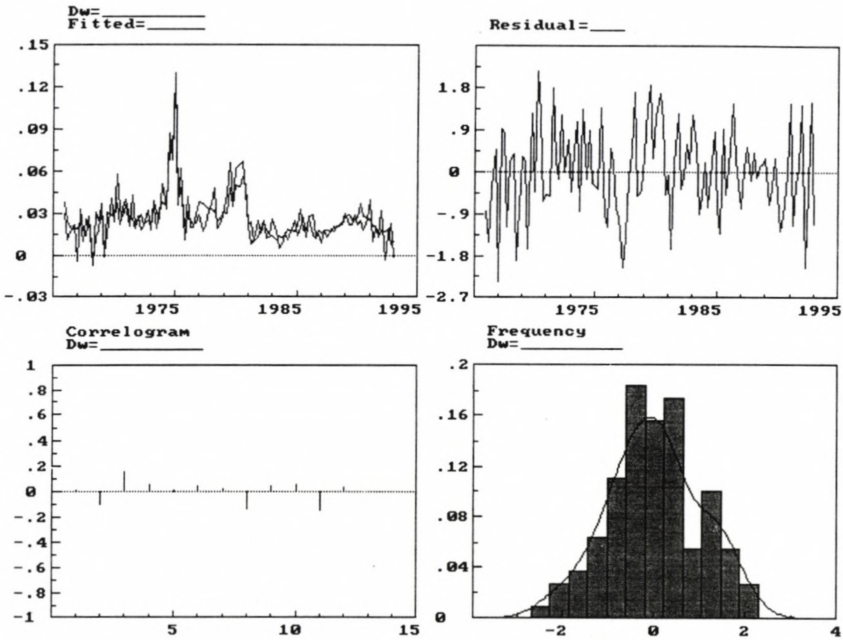

Figure 5. OLS Reduced Wage Equation 1966(1) - 1993(1).

Table 3. OLS estimates final Δw model: 1966(1) - 1993(1)

$$\Delta w_t = \underset{(0.07)}{0.285\Delta w_{t-2}} \quad \underset{(0.08)}{+0.500\Delta p_t} \quad \underset{(0.08)}{+0.260\Delta p_{t-3}} \quad \underset{(0.02)}{-0.030\Delta u_t} \quad \underset{(0.0004)}{+0.0017t}$$

$$\underset{(0.34)}{-1.801} \quad \underset{(0.04)}{-0.210w_{t-1}} \quad \underset{(0.04)}{+0.181p_{t-1}} \quad \underset{(0.004)}{+0.038D745} \quad \underset{(0.005)}{-0.03Policy}$$

Diagnostic Statistics

$R^2 = 0.753$	$\sigma = 0.0096$	$V = 0.298$
$J = 1.680$	$AR(5,94) = 0.941$	$ARCH(4,91) = 1.161$
	$[p > 0.46]$	$[p > 0.33]$
$N = 0.233$	$H(18,80) = 1.347$	$R(1,98) = 0.608$
$[p > 0.89]$	$[p > 0.18]$	$[p > 0.44]$

Not surprisingly this remains a congruent model, and the hypothesis that it parsimoniously encompasses the general model of Table 1 is not rejected since the relevant statistic is $F(11,87) = 0.462$ $[p > 0.92]$. The long run solution (with the dummy variables $D745$ and $Policy$, which were included in the estimation, omitted) of this final single equation model takes the form:

$$ecm_t = w_t \quad \underset{(0.11)}{+8.573} \quad \underset{(0.05)}{-0.852p_t} \quad \underset{(0.001)}{-0.008t} \tag{24}$$

This equilibrium correction mechanism is essentially the same as that estimated from the general model reported in Table 1, and has the same interpretation. In particular, the real wage hypothesis is rejected, and in the longer term the level of the unemployment variable has no effect on wages. This latter result is surprising and will be investigated further in Section 4.2. The estimated steady state corresponding to equation (22) for this model has the form:

$$\mathsf{E}(w_t \mid p_t, u_t) = -8.61 + 0.85p_t + 0.008t - 0.68\Delta p^* + 0.14\Delta u^* \tag{25}$$

By construction this steady state equation is consistent with the growth of wages implied by the long run equation (24). An implication of this steady state equation for w is that the equilibrium value of nominal wages is smaller (larger) ceteris paribus the larger is the steady state inflation rate when the latter is positive (negative). Since the rate of inflation is almost always positive this finding is consistent with the view that higher rates of positive inflation are undesirable. The other implication of the steady state equation is that the equilibrium value of nominal wages is larger (smaller) the larger is the growth rate of the percentage unemployment rate when it is positive (negative). Hence the equilibrium value of the nominal wage ceteris paribus is associated positively with increases in the unemployment rate.

Given the simplicity of this illustration there are not many sophisticated economic models that can be obtained from this data set and considered as alternative explanations of the determination of nominal wages. However, in the spirit of illustration the following models can be used as examples of alternatives corresponding to four of the commonly used single equation model types:

$$\begin{array}{ll} \mathsf{M}_{w1} & \Delta w_t = c_1 + \beta_1 \Delta p_t + \gamma_1 (w_{t-1} - \kappa_1 p_{t-1}) + \nu_{1t} \\ \mathsf{M}_{w2} & \Delta w_t = c_2 + \gamma_2 (w_{t-1} - \kappa_2 u_{t-1}) + \nu_{2t} \\ \mathsf{M}_{w3} & (w_t - p_t) = c_3 + \lambda_3 t + \nu_{3t} \\ \mathsf{M}_{w4} & \Delta w_t = c_4 + \nu_{4t} \end{array} \tag{26}$$

These models are: M_{w1} a first order error correction model with wages adjusting towards a static equilibrium that depends on prices only (i.e. equation (18) with $k = 0$ and $\alpha_{30} = \beta_2 = \lambda = 0$); M_{w2} a first order partial adjustment model with target value of wages depending on unemployment only (*i.e.* equation (18) with $k = 0$ and $\alpha_{20} = \alpha_{30} = \beta_1 = \lambda = 0$); M_{w3} in which the logarithm of real wages is trend stationary (i.e. equation (18) with $k = 0, \alpha_{20} = \beta_1 = 1$ and $\alpha_{30} = \beta_2 = 0$); and finally a random walk with drift model M_{w4} (i.e. equation (18) with $k = 0, \alpha_{20} = \alpha_{30} = \gamma = \lambda = 0$). Hendry, Pagan and Sargan [1984] and Hendry [1994] contain more discussion of these and other single equation model types. With the exception of M_{w4} which is nested in M_{w1} and M_{w2}, these are nonnested models in the sense than none can be obtained as restricted versions of any of the other models, though of course they are all special cases of (18). Therefore these models can be compared by using nonnested test statistics such as those of Cox [1960, 1961], Ericsson [1983] and the complete parametric encompassing (CPE) statistic of Mizon [1984] and Mizon and Richard [1986]. However, as was emphasized in Section 3 above, when the alternative models to be considered are all special cases of the congruent general model, so that none of them imply an extension of the dataset or class of models being used, the principle of parsimonious encompassing provides an appropriate method for comparing the models. Further, the use of nonnested test statistics to make pairwise comparisons of rival models makes little sense unless each model is at least congruent with respect to its own information set (i.e. the minimum information set needed to support the model). When the four models in (26) were estimated with the dummy variables $D745$ and *Policy* included in each, they all had serially correlated residuals and hence are probably dynamically misspecified, and models M_{w2},M_{w3} and M_{w4} had significant ARCH effects and non-normality in their residuals, as well as having significant RESET test statistics. In addition, M_{w2} and M_{w3} exhibited signs of parameter non-constancy in the variance and joint variance and regression coefficient stability test statistics V and J. Hence there is strong evidence that the four models in (26) are not coherent with their own information sets, and so are inadequate. Despite this, M_{w2} and M_{w4} $\mathcal{E}$ M_{w3} in pairwise comparisons as the following statistics indicate:

Encompassing test statistics for M_{w2}, $\mathsf{M}_{w4}\mathcal{E}\mathsf{M}_{w3}$

	Cox	Ericsson	CPE
M_{w2}	$N(0,1) = -0.82$	$N(0,1) = 0.78$	$F(2,105) = 1.53$
M_{w4}	$N(0,1) = -0.78$	$N(0,1) = 0.74$	$F(2,106) = 0.33$

This serves to illustrate the dangers of using such test statistics when the models being compared are non-congruent, for both M_{w2}, M_{w4} are inadequate models when evaluated relative to their own information sets and relative to that of the congruent general model (and neither model parsimoniously encompasses the reduced general model reported in Table 2).

Given that each of the models in (26) is nested within the congruent model reported in Table 2 (this model rather than the final model of Table 3 is used in order to include u_{t-1} in the analysis) it is possible to test the validity of the implied reductions from that model, and the corresponding parsimonious encompassing test statistics are reported below:

Parsimonious encompassing test statistics

M_{w1}	M_{w2}	M_{w3}	M_{w4}
$F(5,98) = 9.15$	$F(6,98) = 11.12$	$F(7,98) = 10184.0$	$F(10,98) = 29.86$

Hence each of the four models fails to parsimoniously encompass the general model (all the test statistics have p values of zero) and so they are each inadequate characterizations of the relationship between w_t, p_t and u_t.

At this stage the model that performs most satisfactorily is that reported in Table 3, and so before turning to system analysis of the relationships between w_t, p_t and u_t the forecasting ability of this model will be evaluated. Firstly, this model was re-estimated with sample data for 1966(1) to 1989(3) and then used to forecast Δw from 1989(4) to 1993(1). The outcome is reasonable as indicated by the graphs in figure 6. In addition, the χ^2 prediction test statistic at $\chi^2(14) = 23.01$ [$p > 0.06$] and the Chow [1960] prediction test statistic

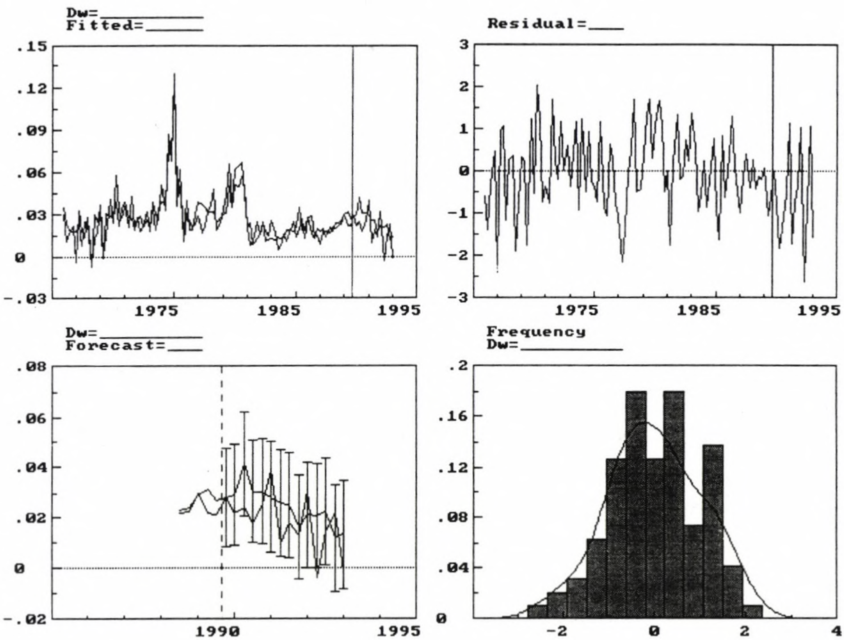

Figure 6. Forecasts 1989(4) - 1993(1).

at $F(14, 85) = 1.35$ [$p > 0.20$] do not reject the null of parameter constancy, though this result is very marginal for the χ^2 prediction test statistic. Indeed, Lu and Mizon [1989] showed that the implicit null hypothesis of the χ^2 prediction statistic is $(\mathbf{x}_t'\Delta\beta)^2 + (\sigma_2^2 - \sigma_1^2) = 0 \;\forall\, t \in [(T+1), (T+H)]$, and that of the Chow statistic is known to be $X_2\Delta\beta = 0$, when for the generic linear regression model $y_t = \mathbf{x}_t'\beta + u_t$ the data for the sample period ($t = 1, 2, ...T$ which is denoted by subscript 1) and prediction period ($t = (T+1), (T+2), ...(T+H)$ which is denoted by subscript 2) can be written as:

$$\begin{pmatrix} y_1 \\ y_2 \end{pmatrix} = \begin{pmatrix} X_1 & 0 \\ X_2 & I_H \end{pmatrix} \begin{pmatrix} \beta_1 \\ X_2\Delta\beta \end{pmatrix} + \begin{pmatrix} u_1 \\ u_2 \end{pmatrix} \tag{27}$$

when $\Delta\beta = (\beta_2 - \beta_1)$ and with:

$$\mathsf{E}\begin{pmatrix} u_1 \\ u_2 \end{pmatrix} = \begin{pmatrix} 0 \\ 0 \end{pmatrix} \quad \text{and} \quad \mathsf{V}\begin{pmatrix} u_1 \\ u_2 \end{pmatrix} = \begin{pmatrix} \sigma_1^2 I_T & 0 \\ 0 & \sigma_2^2 I_H \end{pmatrix} \tag{28}$$

Hence the χ^2 prediction test statistic is sensitive to both changes in the error variance and in the regression coefficients via changes in the conditional mean of y_t, whereas the Chow prediction statistic is powerful against changes in the conditional mean only - a constant error variance ($\sigma_1^2 = \sigma_2^2$) is part of the maintained hypothesis for the Chow statistic. Clearly though, if the error variance is not constant and/or the error distribution is otherwise nonstationary, the Chow statistic will not have a central $F(H, T + H - k)$ distribution under the hypothesis $X_2\Delta\beta = 0$. Bearing in mind that 1989(4) saw the tightening of monetary policy following the UK joining the European Monetary Mechanism the finding that there might have been a change in the model's error variance is not surprising, even though it is a clear limitation of the model. However, when secondly the model was re-estimated with data for the period 1966(1) to 1979(2), and then used to forecast Δw over the period 1979(3) to 1993(1) the result was much less satisfactory. The χ^2prediction test statistic over this period takes the value $\chi^2(55) = 326.2$ [$p > 0.00$] which strongly indicates parameter non-constancy in the model. The Chow prediction statistic on the other hand takes the value $F(55, 44) = 0.962$ [$p > 0.56$] and so still does not reject the null of parameter constancy - but there is evidence that the error does not have a stationary distribution throughout the period 1966(1) to 1993(1). Inspection of the graphs in Figure 7 does reveal the poor forecasting performance of the model for the period 1979(3) to 1993(1).

The estimates for the final model over the sample period 1966(1) to 1979(2) are given in Table 4 and these are noticeably different from those for the full sample in Table 3.

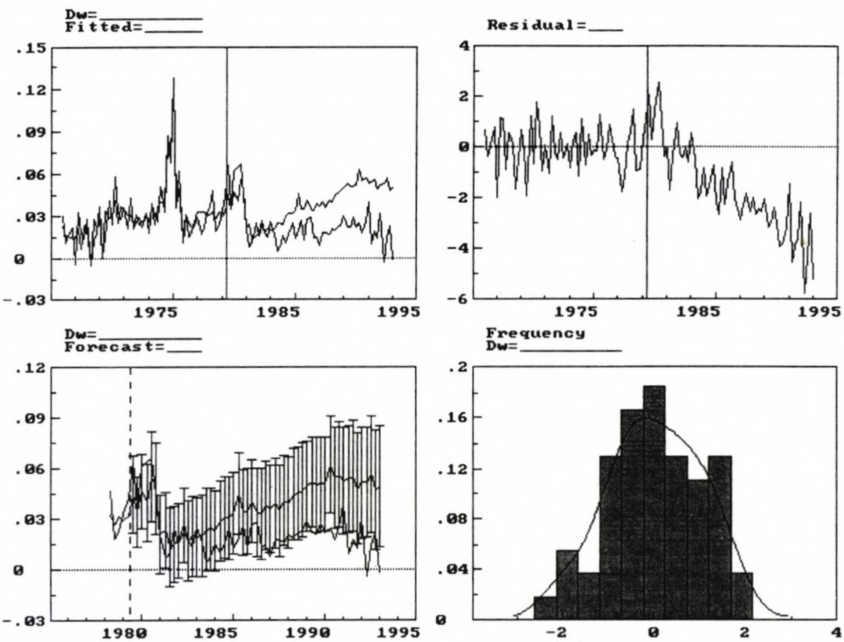

Figure 7. Forecasts 1979(3) - 1993(1).

Table 4. OLS estimates final Δw model: 1966(1) - 1979(2)

$$\Delta w_t = \underset{(0.08)}{0.304\Delta w_{t-2}} \; \underset{(0.13)}{+0.360\Delta p_t} \; \underset{(0.12)}{+0.077\Delta p_{t-3}} \; \underset{(0.02)}{-0.007\Delta u_t} \; \underset{(0.0008)}{+0.003t}$$
$$\underset{(0.52)}{-1.73} \; \underset{(0.06)}{-0.212w_{t-1}} \; \underset{(0.05)}{+0.145p_{t-1}} \; \underset{(0.005)}{+0.039D745} \; \underset{(0.007)}{-0.03Policy}$$

Diagnostic Statistics

$R^2 = 0.823$	$\sigma = 0.0097$	$V = 0.176$
$J = 0.759$	$AR(5,39) = 2.599^*$	$ARCH(4,36) = 0.377$
	$[p > 0.04]$	$[p > 0.82]$
$N = 0.930$	$H(17,26) = 1.210$	$R(1,43) = 0.107$
$[p > 0.63]$	$[p > 0.32]$	$[p > 0.74]$

As a consequence the estimated steady state solution of the model over the period 1966(1) to 1979(2) is:

$$\mathsf{E}(w_t \mid p_t, u_t) = -8.19 + 0.682p_t + 0.014t - 0.18\Delta p^* \qquad (29)$$

which differs substantially from that obtained for the full sample (25). The coefficient of p_t is even further away from the real wage hypothesis value of unity, the response of w_t to the steady state inflation rate Δp^* is a quarter of its full sample value, and there is no role for Δu^* since its estimated coefficient in the steady state solution is -0.03 and not significantly different from zero. Hence, although the full sample estimates of the general model (given in Table 1) appeared to be congruent, and the simplified model presented in Table 3 an acceptable reduction of it which also inferentially dominates the simple alternative models in (26), there is strong evidence that the simplified model is not constant pre- and post-1979(3). This is consistent with the suggestion that the recursively estimated constant and the coefficients of w_{t-1}, p_{t-1} and u_{t-1} changed their values in the 1980's. In fact, when the general model is re-estimated for the sample 1966(2) - 1979(2) it does indeed exhibit evidence of a change, as is indicated by its steady state solutions for the full sample and the sample pre-1979(3):

General model steady state solutions.

1966(2) – 1993(1)

$$\mathsf{E}(w_t \mid p_t, u_t) = -8.57 + 0.851p_t + 0.016u_t + 0.008t + 1.19\Delta p^* - 0.23\Delta u^*$$

1966(2) – 1979(2)

$$\mathsf{E}(w_t \mid p_t, u_t) = -8.12 + 0.66p_t - 0.08u_t + 0.017t - 0.318\Delta p^* + 0.46\Delta u^*$$

That there might be a change in the determination of wages within the class of linear regression models involving w_t, p_t, u_t and t pre- and post-1980 is not a surprise. There have been important changes affecting the labour market such as the changes in, labour productivity, the prevailing rate of unemployment, hours of work, the amount of part-time working, and the participation of women, as well as changes in the method of collection of some labour market statistics. In addition, there have been important changes in the rate of inflation, interest rates and exchange rates, often associated with changes in government economic policies. Indeed, the Thatcher government which came into power in May 1979 encouraged small businesses and self employment; encouraged the employment of cheaper labour especially female labour which led to more part-time working; and induced a severe recession with a very tight monetary policy. These aspects of policy are consistent with the observed data exhibiting, a decrease in the rate of inflation Δp, and a rapid increase in unemployment - both these features are clearly evident in Figure 1.

Therefore there appears to be evidence that the single equation models developed above for the explanation of w_t in terms of p_t, u_t and t have not captured a constant underlying structure. Note that this has arisen despite the fact that the full sample estimated models appeared to be congruent, with neither of the Hansen [1992] instability statistics V and J indicating any structural break around 1979-1980. Possible reasons for this include lack of cointegration between w_t, p_t, and u_t, and the invalidity of the assumption that p_t and u_t are weakly exogeneous variables. These issues are related and best analysed in the context of system modelling, to which attention is now turned.

4.2 System Analysis

A more comprehensive analysis of the relationships between the variables under scrutiny can be undertaken by modelling their joint distribution. In particular, this allows issues concerning the presence and nature of regime shifts and structural breaks, and their links (if any) to changes in the exogeneity status of variables to be explored, and enables hypotheses about the precise form of the relationships between the variables to be tested.

The class of system to be considered in this section is a three dimensional VAR of the form:

$$\mathbf{x}_t = \sum_{j=1}^{k} A_j \mathbf{x}_{t-j} + \Phi D_t + \varepsilon_t \tag{30}$$

when $\mathbf{x}_t' = (w_t, p_t, u_t)$, D_t contains deterministic variables such as a constant, trend, seasonal and other dummy variables, and ε_t is a three dimensional error vector which is independently distributed with mean zero and covariance matrix Σ. It is assumed that the roots of det $\left(I - \sum_{j=1}^{k} A_j L^j\right) = 0$ lie on or outside the unit circle so that there are no explosive roots, and that k is finite so that moving average error processes do not fall into the category of model being considered. The initial conditions $\mathbf{x}_{1-k}, \mathbf{x}_{2-k}, ...\mathbf{x}_0$ are assumed fixed, and the parameters $(A_1, ...A_k, \Phi, \Sigma)$ are required to be constant in order for the adopted inferential procedures to be valid. An observationally equivalent parametrization of (30) is provided by the following vector equilibrium correction model (VECM):

$$\Delta \mathbf{x}_t \quad = \quad \sum_{j=1}^{k-1} \Pi_j \Delta \mathbf{x}_{t-j} + \Pi \mathbf{x}_{t-k} + \Phi D_t + \varepsilon_t \tag{31}$$

in which $\Pi = \left(I - \sum_{j=1}^{k} A_j\right)$ is the static equilibrium response matrix, and $\Pi_j = \left(I - \sum_{i=1}^{j} A_i\right)$ for $j = 1, 2, ...(k-1)$ are the interim multiplier matrices. In the case that $\mathbf{x}_t \sim \mathsf{I}(1)$ the rank of Π is determined by the number of cointegrating vectors. Letting the rank of Π be r and α and β be $3 \times r$ matrices of rank r such that $\Pi = \alpha\beta'$ yields the reduced rank VAR system:

$$\Delta \mathbf{x}_t \quad = \quad \sum_{j=1}^{k-1} \Pi_j \Delta \mathbf{x}_{t-j} + \alpha\beta' \mathbf{x}_{t-k} + \Phi D_t + \varepsilon_t \tag{32}$$

in which $\beta' \mathbf{x}_t$ are the r cointegrating vectors which are $\mathsf{I}(0)$ and α contains the adjustment coefficients. The Johansen [1988] maximum likelihood procedure, as modified in Johansen and Juselius [1990] to allow for dummy variables, enables the empirical determination of r provided that the systems in (30) and (31) are well specified and in particular have innovation homoscedastic errors and constant parameters.

In the light of the single equation analysis presented above an unrestricted VAR for $w, p,$ and u with 5 lags on each variable (corresponding to (30) with $k = 5$) and D_t containing an unrestricted constant vector and linear trend was estimated for the full sample period 1966(2) to 1993(1). The results strongly indicated that this system was non-congruent, with each equation exhibiting serially correlated, heteroscedastic and non-normal residuals, with $ARCH$ effects present as well. Inspection of the residuals revealed many apparent outliers, and recursive estimation showed many parameter estimates to be non-constant. The sample period contains many important changes in economic policy and institutional arrangements, and so it may well be difficult to obtain a constant parameter VAR without some specific allowance for these events. For example: there were sharp movements in commodity prices generally in 1966 and in oil prices at the end of 1973, in 1979(3) and in 1986; there were pre-election booms 1972/3, 1974, 1979, 1986/7, 1991; floating exchange rates were introduced in 1971 and exchange controls were abandoned in 1979; the rate of value added tax (VAT) was increased from 8% to 15% in July 1979; periods of very tight monetary policy were introduced in 1980 and 1989 which had major effects on income, employment and unemployment; in 1984 there was the year long miners' strike; the introduction of the Business Expansion Scheme in 1980 affected the ratio of self-employed to employees in the total employed; there were changes in the method of collection of some key statistics such as unemployment and the numbers self-employed. Noting that many of these changes affected the labour market and the rate of inflation, suggests that the relationship between wages, prices and unemployment may not be an easy one to capture without some allowance for their effect.

From inspection of the residuals for the unrestricted VAR with $k = 5$ and reference to the relevant economic history, it was decided to include the following 4 dummy variables: $D793$ which takes the value 1 in 1979(3) and 0 elsewhere to allow for the increase in value added tax and the second oil price hike; $D745$ which takes the value 1 in 1974(3) and 1974(4), the value 2 in 1975(1), and zero elsewhere to capture the wage explosion following the relaxation of wage and price controls; *Budget* which takes the values 1 and -1 respectively in the second and third quarters of every year and zero in all other quarters, and represents fixed seasonal effects and especially the changes in excise duties and other tax rates, announced in the annual Budget, on prices and the rate of inflation; and *Expansion* which takes the value 1 in 1966(3) -1 in 1967(2), 1972(3), and 1974(3) highlighting quarters in which fiscal policy was used to stimulate(+)/dampen($-$) the economy, and takes the values 2 in 1971(2) -2 in 1971(1) to represents the effect of the announced cut of 50% in Selective Employment Tax. Since these four dummy variables should not have a long run effect on any of the modelled variables w_t, p_t and u_t they are entered unrestrictedly into the VAR when the reduced rank maximum likelihood procedure of Johansen [1988] is applied it. However, there is no evidence of quadratic trend in any of the modelled variables and so the linear deterministic trend is restricted to lie in the cointegration space.

When an unrestricted VAR parameterized as (31) with $k = 4$, a constant and the four dummy variables plus a linear trend entered as described above, was estimated over the period 1966(1) to 1993(1) it produced a plausible and more-or-less congruent system. Figure 8, which gives the graphs of the actual and fitted values for $\Delta w_t, \Delta p_t$ and Δu_t, and the corresponding residuals (scaled), illustrates this point. The system provides a high degree of explanation for the changes $\Delta w_t, \Delta p_t$ and Δu_t in the three modelled variables w_t, p_t and u_t (it is of course even more impressive in terms of the levels of the variables), and

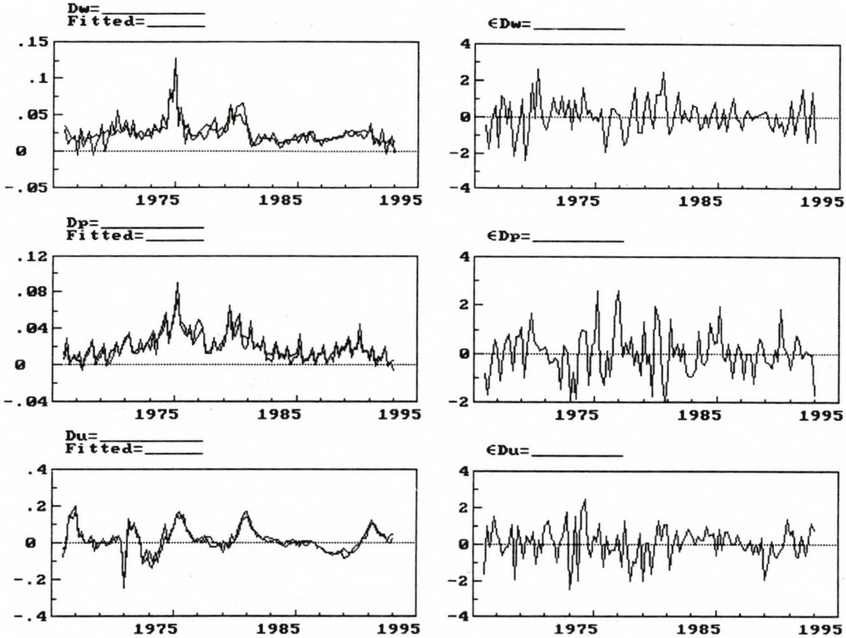

Figure 8. System Actual, Fitted and Residual Values.

there does not appear to be any strong heteroscedasticity or serial dependence in the residuals. Table 5 provides descriptive statistics for the estimated system: first single equation residual standard deviations $\hat{\sigma}$, serial correlation AR, autoregressive heteroscedasticity $ARCH$, heteroscedasticity H, and normality N test statistics as defined for Table 1; and second test statistics for vector autoregressive residuals $vecAR$, vector heteroscedasticity $vecH$, and finally vector normality $vecN$ (see Doornik and Hendry [1994] for more details of these test statistics).

The only evidence of noncongruence comes from the AR test statistic for the Δp_t equation and the $vecAR$ statistic both of which are significant at 1%.

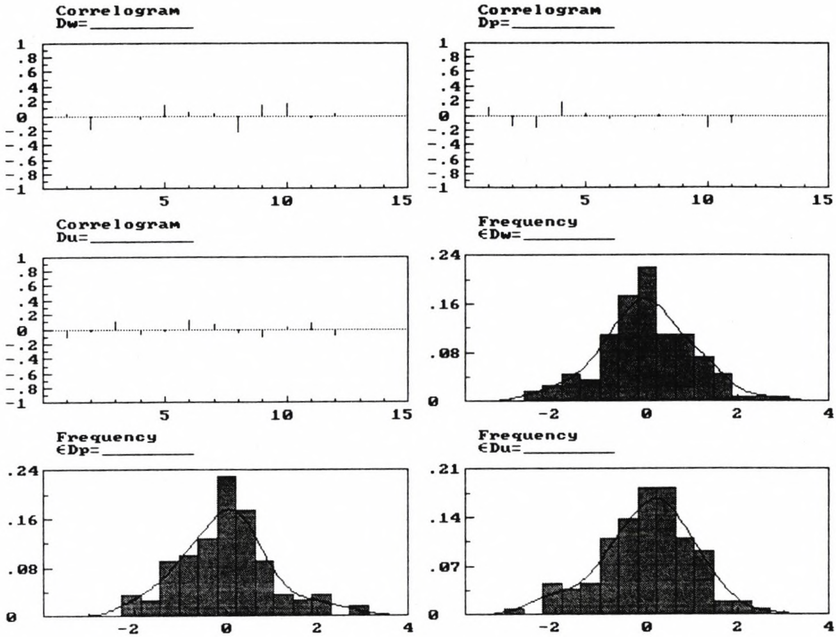

Figure 9. System Residual Correlograms and Frequency Plots.

Table 5. System Diagnostic Statistics			
	Δw_t	Δp_t	Δu_t
$\hat{\sigma}$	1.08%	0.72%	2.97%
$AR(5,86)$	1.98	2.53*	0.68
$ARCH(4,83)$	2.16	0.12	2.47
$H(26,64)$	1.31	1.15	1.04
$N(2)$	1.67	3.08	2.58
$vecAR(45,220)$	1.59*		
$vecH(156,354)$	0.93		
$vecN(6)$	7.20		

Further evidence on the approximate congruence of the system is provided by the residual correlograms and frequency plots given in Figure 9.

Though there is some evidence of residual serial correlation in the equations for Δw_t and Δp_t it is not strong, and experimentation revealed that inappropriate lag length k does not appear to be the cause of this problem. In fact, inspection of the residual plots in Figure 8 suggests that the difficulty may still lie in the system's inability to represent the many changes in government policies towards wages and prices, interest rates and exchange rates, and unemployment during the sample period. However, rather than introduce more event specific dummy variables into the analysis, which would increase the risk of the system being too finely tuned to the particular sample data and hence even more likely to suffer from predictive failure, it was decided to continue with the present set of dummy variables. The other alternative of extending the information set to include interest and exchange rates is not pursued here since the present modelling is intended to be illustrative rather than definitive, but it is the subject of further research.

Table 6. Residual Correlations

	w_t	p_t	u_t
p_t	0.14	1.0	–
u_t	0.08	0.03	1.0

System Dynamics

	w_t	p_t	u_t
$F_{k=3}(3,89)$	0.30	3.42*	1.08
$F_{k=2}(3,89)$	1.79	1.32	2.89*
$F_{k=1}(3,89)$	4.59**	3.25**	11.84**
$\mid \lambda_\pi \mid$	0.29	0.04	0.00

Inspection of the residual correlations in Table 6 suggests that there is a modest correlation between w_t and p_t, but the correlations between w_t and u_t and p_t and u_t are negligible. The hypothesis that the order of lag required for each variable is i can be tested using the statistics $F_{k=i}$, and on the basis of these it appears that there is scope for simplifying the dynamics, but that 4 lags are required for p_t consistent with it being seasonal. The estimated long run matrix Π has eigenvalues whose moduli are given by $\mid \lambda_\pi \mid$, and these indicate that there are probably 2 zero eigenvalues suggesting that $r = 1$. This is confirmed by the results of applying the Johansen [1988] maximum likelihood procedure to estimate the dimension of cointegrating space r which are reported in Table 7.

Table 7: Cointegration Statistics			
r	1	2	3
l	1438	1445	1446
μ	0.28	0.12	0.02
Max	36.17**	14.08	2.05
$Trace$	52.29**	16.13	2.05

The eigenvalues μ_i which are involved in the maximization of the log-likelihood function with respect to β in order to obtain estimated cointegrating vectors (see Johansen [1988] or Banerjee *et al* [1993] for details) are small. However, the largest eigenvalue $\mu_1 = 0.28$ is significantly different from zero on the basis of the maximum eigenvalue ($Max = -T \ln(1-\mu_r)$) and trace ($Trace = -T \sum_{i=1}^{r} \ln(1-\mu_i)$) test statistics. r is the dimension of cointegrating space and l (which is defined as $-T/2 \sum_{i=1}^{r} \ln(1-\mu_i)$) is the corresponding value of the log-likelihood function apart from a constant. Hence the value of 1445.87 for the unrestricted log likelihood function is reported as 1446 for $r = 3$. The Max and $Trace$ test statistics are not adjusted for degrees of freedom as suggested by Reimers [1992], since the results in Kostial [1994] indicate a tendency for them to underestimate the dimension of cointegrating space even when unadjusted. The critical values used for the Max and $Trace$ statistics are given in Osterman-Lenum [1992]. On the basis of the statistics in Table 7 it is concluded that there is one cointegrating vector ($r = 1$), and an estimate of it is given by $\hat{\beta}_1'$ in Table 8.

Table 8. Adjustment Coefficients α_i and Eigenvectors β_i'

$\hat{\alpha}_i$	$i=1$	$i=2$	$i=3$	$\hat{\beta}_i'$	w	p	u	t
w	−0.16	−0.00	0.00	$i=1$	1	−0.88	−0.017	−0.008
p	0.16	−0.00	0.00	$i=2$	−14.85	1	7.57	0.29
u	−0.25	−0.01	−0.00	$i=3$	−0.51	−1.11	1	0.007

This cointegration analysis with the linear trend t restricted to be in the cointegration space, yields an estimated cointegrating vector $\hat{\beta}_1'\mathbf{x}_t = w_t - 0.88p_t - 0.017u_t - 0.008t$ which implies that wages in a long run equilibrium ($\mathsf{E}(\hat{\beta}_1'\mathbf{x}_t) = 0$) increase by 3% per annum after the influence of prices and unemployment have been taken into account. Figure 10 gives the graphs of the disequilibria or cointegrating vectors $\hat{\beta}_i'\mathbf{x}_t$, the "actual" ($x_{it}$) and "fitted" $\left(\sum_{j \neq i} \hat{\beta}_{ij} x_{jt}\right)$ values for $\mathbf{x}_t' = (w_t, p_t, u_t)$, and the recursively estimated eigenvalues (see Hansen and Johansen [1993]) each after having partialled out the full sample short run dynamics and the unrestricted variables (i.e. the constant and the dummies

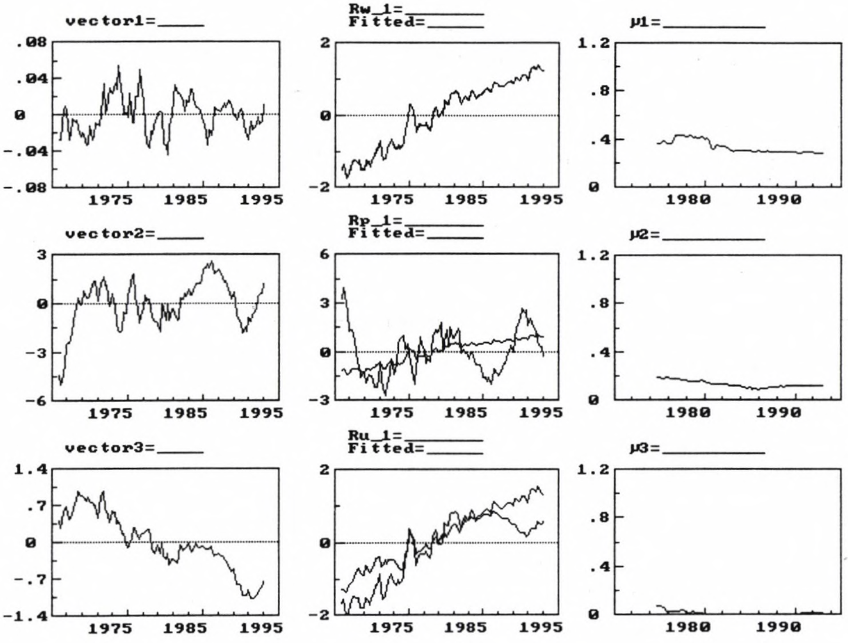

Figure 10. Cointegrating Vectors and Recursive Eigenvalues

Budget, *Expansion*, $D745$ and $D793$). The first cointegrating vector appears to be $\mathsf{I}(0)$, but the other two vectors are clearly nonstationary. There is a very close correspondence between the "actual" and the "fitted" values for w_t consistent with the disequilibrium rarely being over 4%. Note though that the estimated coefficient of u_t in the error correction has the opposite sign to that estimated by Clements and Mizon [1991], implying that ceteris paribus wages increase with increases in unemployment and so are counter cyclical! However, Clements and Mizon [1991] included average labour productivity in their analysis, and found evidence of a long run positive association between productivity and unemployment.

Note that these inferences concerning the integration-cointegration properties of the system require the system parameters to be constant if they are to be valid, whereas conventional test statistics for parameter constancy (e.g. system analogues of the analysis of variance and Chow statistics mentioned in Section

4.1) have known distributions for $\mathsf{I}(0)$ rather than $\mathsf{I}(1)$ systems. Although constancy tests for $\mathsf{I}(1)$ variables are being developed, recursive estimation already provides a valuable check on parameter constancy. It is therefore reassuring that the recursive estimates of the eigenvalues are essentially constant. In particular, the largest μ_1 is approximately 0.4 for sample sizes 35 to 109, with only a suggestion of nonconstancy around 1979 and 1980. Of the other two eigenvalues neither of which is significantly different from zero in the full sample, μ_2 take values declining from 0.2 to 0.1, and μ_3 is uniformly close to zero for all sample sizes. Therefore, it was decided to proceed to analyse the system further on the assumption that there is a single cointegrating vector.

Although the single cointegrating vector is already identified it is of interest to test overidentifying restrictions on it such as: (i) the absence of an unemployment effect on wages and prices ($\beta_{13} = 0$); (ii) a long run real wage equilibrium ($\beta_{12} = -1$); and (iii) no long run trend in wages and prices ($\beta_{14} = 0$). In addition, necessary conditions for p_t and u_t to be weakly exogenous for the parameters of the long run wage-price equation - ($\alpha_{21} = 0$) and ($\alpha_{31} = 0$) respectively - can be tested (see Boswijk [1992], Hendry and Mizon [1993], Johansen [1992a, 1992b] and Urbain [1992]). The test statistics are conventional likelihood ratio statistics, since the hypotheses are linear on an $\mathsf{I}(0)$ parameterization of the system (N.B. if $\mathbf{x}_t \sim \mathsf{I}(1)$, $\Delta\mathbf{x}_t$ and $\beta'\mathbf{x}_t \sim \mathsf{I}(0)$ when there is cointegration), and so they have limiting χ^2 distributions with the degrees of freedom equal to the number of independent restrictions being tested.

Table 9. Overidentifying and Weak Exogeneity Tests

Hypothesis	Statistic	p-value
$\beta_{13} = 0$	$\chi_2(1) = 0.57$	0.45
$\beta_{14} = 0$	$\chi^2(1) = 17.23^{**}$	0.00
$\beta_{12} = -1$	$\chi^2(1) = 5.46^{*}$	0.02

Hypothesis	Statistic	p-value
$\alpha_{11} = 0$	$\chi^2(1) = 9.31^{**}$	0.00
$\alpha_{21} = 0$	$\chi^2(1) = 14.49^{**}$	0.00
$\alpha_{31} = 0$	$\chi^2(1) = 2.08$	0.15
$\alpha_{11} + \alpha_{21} = 0$	$\chi^2(1) = 0.01$	0.93

Hypothesis	Statistic	p-value
$\alpha_{31} = 0,\ \beta_{13} = 0$	$\chi^2(2) = 5.48$	0.07
$\alpha_{31} = 0,\ \beta_{12} = -1$	$\chi^2(2) = 12.92^{**}$	0.00
$\alpha_{11} + \alpha_{21} = 0,\ \beta_{13} = 0$	$\chi^2(2) = 0.67$	0.72
$\alpha_{11} + \alpha_{21} = 0,\ \beta_{13} = 0,\ \alpha_{31} = 0$	$\chi^2(3) = 5.48$	0.14

The hypotheses that the long run equilibrium is one for real wages ($\beta_{12} = -1$), that prices are weakly exogenous for a long run wage equation ($\alpha_{21} = 0$), that wages are weakly exogenous for a long run price equation ($\alpha_{11} = 0$), and that there is no trend in the long run wage-price equation ($\beta_{14} = 0$), are all rejected. However, the hypotheses that u_t is weakly exogeneous for the parameters of the long run wage-price equation ($\alpha_{31} = 0$), that there is no unemployment effect in the long run wage-price equation ($\beta_{13} = 0$), and that the adjustment coefficient for wages is equal in magnitude but opposite in sign to that of prices ($\alpha_{11} + \alpha_{12} = 0$), are not rejected separately or jointly. Hence, the full sample estimates of this system are reasonably congruent, and consistent with the existence of a long run wage-price equilibrium of the form:

$$\mathsf{E}(w_t - 0.90p_t) - 0.008t = 0 \tag{33}$$

in which the p_t coefficient estimate of 0.9 is significantly different from unity. The estimated adjustment coefficient when the restrictions $\alpha_{11} + \alpha_{12} = 0, \alpha_{31} = 0$ and $\beta_{13} = 0$ are imposed is $\hat{\alpha}_{11} = -0.15$, which implies that both wages and prices adjust slowly to their equilibrium. Note that although u_t is weakly exogenous for the parameters of (33), it does not itself play a role in the long run equilibrium. In fact, in this system u_t is determined independently of contemporaneous wages and prices, depending only on lagged values of wage and price inflation rates. This finding is consistent with u_t being a target of economic policy with at most weak links to the levels of wages and prices, instead being influenced by fiscal and monetary policies. Further, since the long run wage-price disequilibrium is involved in the determination of both w_t and p_t these variables require joint modelling in order to make efficient inferences. In particular, this casts doubt on the OLS results obtained in Section 4.1, though there is a remarkable similarity between (24) and (33).

The reduced rank ($r = 1$) system can be mapped from $\mathsf{I}(1)$ to $\mathsf{I}(0)$ space by adoption of the re-parameterization

$$\Delta \mathbf{x}_t = \textstyle\sum_{j=1}^{3} \Pi_j \Delta \mathbf{x}_{t-j} + \alpha \cdot ecm_{t-1} + \Phi D_t + \varepsilon_t \tag{34}$$

in which $ecm_t = w_t - 0.88p_t - 0.017u_t - 0.008t$ and the trend is excluded from D_t. Estimation and evaluation of (34) revealed the maintenance of the original system's congruence. This is illustrated by the graphs of actual and fitted values of $\Delta w_t, \Delta p_t$ and Δu_t, and the associated residual correlograms and frequency plots in Figure 11.

As a result of imposing a rank of one on the cointegrating space (which the statistics in Table 7 indicated to be appropriate), and dropping the trend (other than its presence in the cointegrating vector ecm_t), the $\mathsf{I}(0)$ system has 9 fewer parameters than the original system and so will be referred to as a parsimonious VAR (PVAR). Although there is further scope for simplifying this PVAR since a number of the estimated coefficients do not appear to be significantly

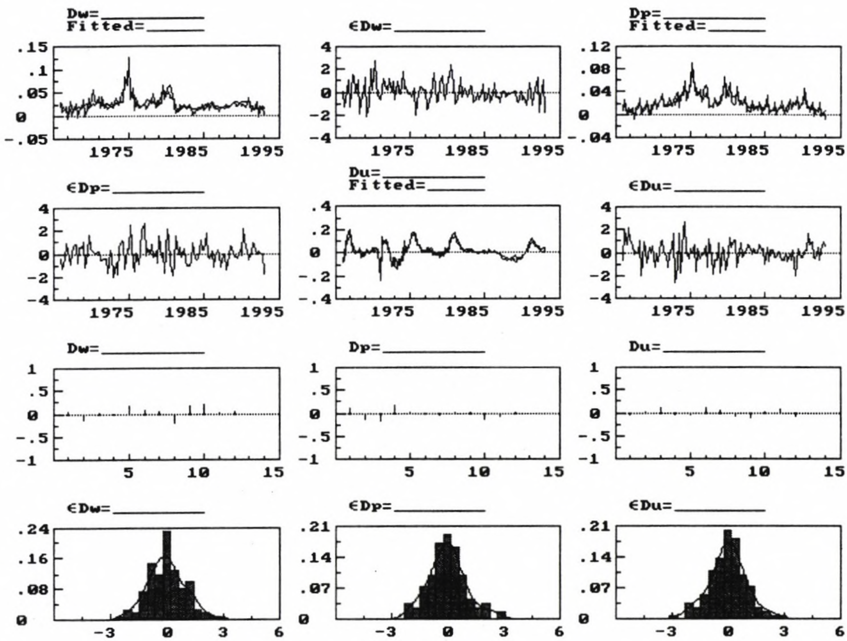

Figure 11. I(0) System (PVAR) Actual, Fitted and Residual Plots.

different from zero, the PVAR is retained in this form as a framework within which to compare alternative models by checking their ability to parsimoniously encompass the PVAR. The advantages of testing for simple models being able to parsimoniously encompass the PVAR are: (i) reducing the risk of using models that are not robust to changes in the sample information (over-parameterized models can be too finely tuned to the peculiarities of a particular sample), (ii) increasing the chance that developed models are invariant to regime changes and thus capturing autonomous relationships rather than ephemeral occurrences, and (iii) allowing the evaluation of models of particular interest (e.g. those implementing specific economic theories).

The first model to be considered is an empirical simplification of the PVAR which retains the joint determination of w_t and p_t via the error correction ecm_t, but simplifies the dynamics and restricts the presence of the dummy variables. Table 10 provides the FIML estimates of the simplified model.

Table 10. Simplified Model FIML Estimates

$$
\begin{aligned}
\Delta w_t &= \underset{(0.14)}{0.32\Delta p_t} \; \underset{(0.07)}{+0.33\Delta w_{t-2}} \; \underset{(0.08)}{+0.30\Delta p_{t-3}} \; \underset{(0.02)}{-0.04\Delta u_{t-2}} \\
&\quad \underset{(0.06)}{-0.13 ecm_{t-1}} \; \underset{(0.005)}{+0.04 D745} \; \underset{(0.01)}{+0.03 D793} \; \underset{(0.48)}{-1.10} \\
\Delta p_t &= \underset{(0.06)}{0.37\Delta p_{t-1}} \; \underset{(0.01)}{-0.04\Delta u_{t-2}} \; \underset{(0.02)}{+0.18 ecm_{t-1}} \; \underset{(0.001)}{+0.01 Budget} \\
&\quad \underset{(0.003)}{+0.01 D745} \; \underset{(0.007)}{+0.05 D793} \; \underset{(0.21)}{+1.59} \\
\Delta u_t &= \underset{(0.17)}{0.76\Delta w_{t-1}} \; \underset{(0.06)}{+0.79\Delta u_{t-1}} \; \underset{(0.06)}{+0.12\Delta u_{t-2}} \; \underset{(0.06)}{-0.14\Delta u_{t-3}} \\
&\quad \underset{(0.01)}{+0.13 Expansion} \; \underset{(0.005)}{-0.01} \\
ecm_t &\equiv ecm_{t-1} + \Delta w_t - 0.88\Delta p_t - 0.017\Delta u_t - 0.008
\end{aligned}
$$

In the simplified model u_t is weakly exogenous for the parameters of the two equations that jointly determine w_t and p_t, and the long run wage-price equilibrium is given by $\mathsf{E}(ecm_t) = \mathsf{E}(w_t - 0.88p_t - 0.017u_t) - 0.008t = 0$. The additional model information given in Table 11 provides further evidence that the simplified model, apart from increased serial correlation in the residuals for Δw_t and Δp_t and more $ARCH$ effects in those for Δu_t, performs just as well as the congruent system.

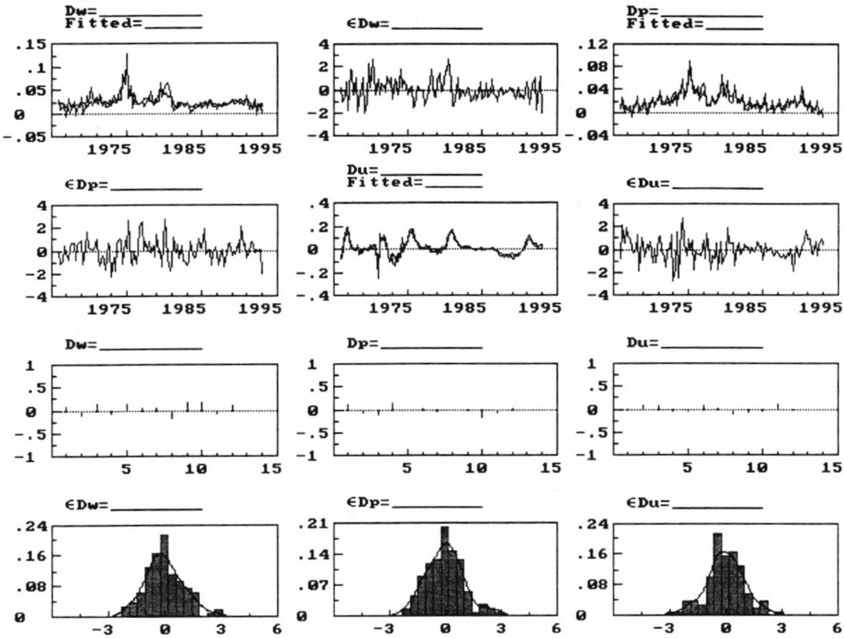

Figure 12. Simplified Model Actual, Fitted and Residual Plots.

Table 11. Simplified Model Diagnostic Statistics

	Δw_t	Δp_t	Δu_t
$\hat{\sigma}$	1.08%	0.71%	3.10%
$AR(5,89)$	3.85^{**}	3.33^{**}	2.04
$ARCH(4,86)$	1.94	0.05	3.79^{**}
$H(27,66)$	0.93	1.34	1.10
$N(2)$	2.23	3.02	2.43

Vector Tests		Residual Correlations	w_t	p_t	u_t
$vecAR(45,253)$	1.34	p_t	−0.01	1.0	–
$vecH(156,413)$	1.04	u_t	0.11	0.07	1.0
$vecN(6)$	7.45				

The 24 restrictions implied by the simplified model relative to the PVAR are not rejected by a likelihood ratio test ($\chi^2(24) = 24.69$ [$p > 0.42$]), so that the simplified model parsimoniously encompasses the PVAR. Since there is evidence of serial correlation and $ARCH$ effects in the residuals the validity of this parsimonious encompassing test is called into question. However, comparison of Figure 12 with Figure 8 reveals the similarity between the congruent system and the simplified model and so lends some support (albeit weak support) to the use of this likelihood ratio test statistic.

The second model considered for encompassing comparison within the framework of the PVAR is a VAR for the differences $\Delta w_t, \Delta p_t$ and Δu_t (DVAR). This model corresponds to (34) with $\alpha = 0$ so that the disequilibrium ecm_t is ignored. As a system, a VAR in differences would be (31) with $\Pi = 0$ so that the long run or zero frequency information in the data is ignored, and represents a class of model that has been popular in the time series analysis of nonstationary data particularly since the work of Box and Jenkins [1970]. In addition to the three restrictions in $\alpha = 0$ there are a further 6 from restricting the number of dummy variables included in each equation as in the simplified model. The likelihood ratio test statistic takes the value $\chi^2(9) = 45.06$ $[p = 0.00]$ so that the DVAR does not parsimoniously encompass the PVAR. Note that relative to the test statistics reported in Table 9 the present test is of the joint hypothesis $\alpha_{i1} = 0 \quad i = 1, 2, 3$ plus the zero restrictions on the coefficients of the dummy variables. Hence, the fact that the hypothesis $\alpha_{21} = 0$ was rejected contributes to, but does not entirely explain, the rejection of the DVAR model. The DVAR also fails to parsimoniously encompass the PVAR when the dummy variables are unrestricted - $\chi^2(3) = 33.86$ $[p = 0.00]$ and so within-sample the zero frequency information contained in the cointegrating vector ecm_t has a valuable role in the modelling w_t, p_t and u_t.

Recalling the evidence of a regime shift in 1979 for the single equation modelling in Section 4.1, it is relevant to test the parameter constancy of the otherwise congruent system, the PVAR, the simplified model, and the DVAR. Table 12 presents the one-step ahead forecast test statistics from estimating the system and models with data up to 1979(2) and checking their ability to forecast over the 55 quarters in the period 1979(3)-1993(1). The three statistics are $F_i(165, T-k) = e'\Psi_i^{-1}e/165 \quad i = 1, 2, 3$ when $e = vec(E), E$ is a 55×3 matrix of forecast errors for the 3 modelled variables, the Ψ_i's are alternative estimates of the asymptotic covariance matrix of e, T $(= 53)$ the sample size, and k the average number of estimated parameters per equation. Ψ_1 is the sample estimate of the innovation covariance matrix Σ, Ψ_2 is an estimate of the forecast error covariance matrix that allows for parameter uncertainty as well as the innovation variance, and Ψ_3 is an estimate of the forecast error covariance matrix that allows for innovation variance, parameter uncertainty, and the covariance between the forecast errors - see Doornik and Hendry [1994] for details. The three statistics $F_i(165, 53-k)$ $i = 1,2,3$ have approximate central F distributions with degrees of freedom 165 and $(53-k)$ under the null hypothesis of parameter constancy.

Table 12. System & Model 1-Step Ahead Forecast Statistics 1979(3)-1993(1)

VAR				PVAR			
$F_1(165,37)$	$F_2(165,37)$	$F_3(165,37)$		$F_1(165,40)$	$F_2(165,40)$	$F_3(165,40)$	
8.45**	2.72**	1.05		1.08	0.90	0.85	
$p = 0.00$	$p = 0.00$	$p > 0.45$		$p > 0.40$	$p > 0.69$	$p > 0.77$	
	Δw_t	Δp_t	Δu_t		Δw_t	Δp_t	Δu_t
Forecast Error				Forecast Error			
Mean	−0.033	−0.002	0.032	Mean	−0.003	0.0001	−0.002
SD	0.035	0.010	0.041	SD	0.012	0.010	0.031
Simplified Model				DVAR			
$F_1(165,48)$	$F_2(165,48)$	$F_3(165,48)$		$F_1(165,43)$	$F_2(165,43)$	$F_3(165,43)$	
0.92	0.85	*na*		1.10	0.93	*na*	
$p > 0.66$	$p > 0.77$	*na*		$p > 0.37$	$p > 0.64$	*na*	
	Δw_t	Δp_t	Δu_t		Δw_t	Δp_t	Δu_t
Forecast Error				Forecast Error			
Mean	−0.003	0.0001	−0.001	Mean	−0.003	−0.001	−0.001
SD	0.012	0.009	0.023		0.013	0.010	0.030

The forecast performance of the VAR is poor (it is only $F_3(165, 37)$ that does not reject parameter constancy) as the graphs in Figure 13 make clear, especially for Δw_t which has the actual and forecast values diverging disconcertingly, and the forecasts lying outside the confidence interval given by ± 2 forecast error standard deviations. The difficulties of the full rank VAR are even more pronounced when the forecasts are presented for the levels of the variables w_t, p_t and u_t.

However, the PVAR, the simplified model, and the DVAR all forecast much better than the VAR, as evidenced by the statistics in Table 12. On the basis of the parameter constancy test statistics and the means and standard deviations of the forecast errors given in Table 12 the best overall performance is found in the simplified model, which is parsimonious and retains the zero frequency information in the disequilibrium ecm_t. The improved quality of these forecasts relative to those from the VAR is shown in Figure 14 which gives the actual, fitted and forecast values for the simplified model.

Indeed, the difficulty lies in the different information used by the different models in forecasting the variables w_t, p_t and u_t. The VAR uses the sample information on the levels of w_t, p_t and u_t and their history, but wrongly treats the VAR as having full rank, and also ignores the regime shift in 1979. The DVAR uses sample information on the differences $\Delta w_t, \Delta p_t$ and Δu_t and their history,

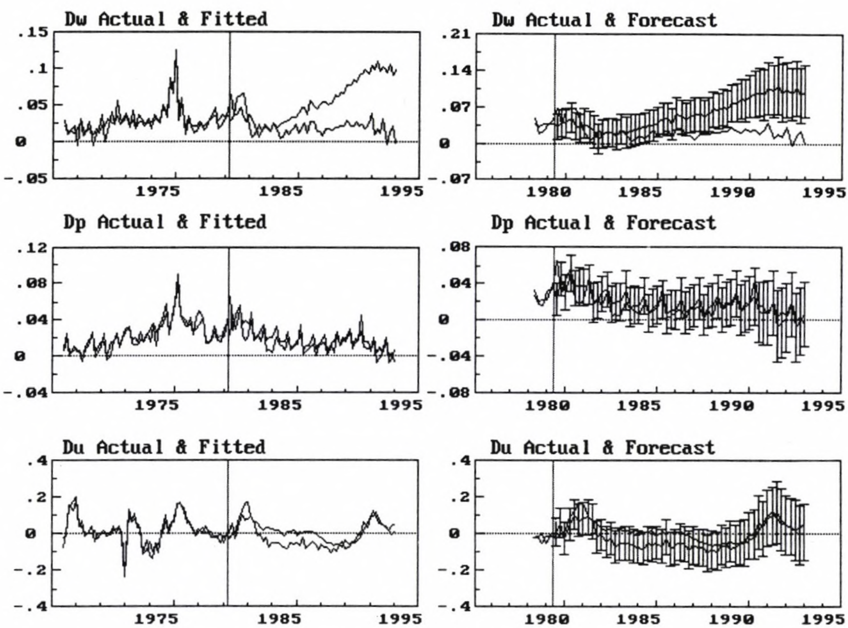

Figure 13. VAR Actual, Fitted and Forecasts

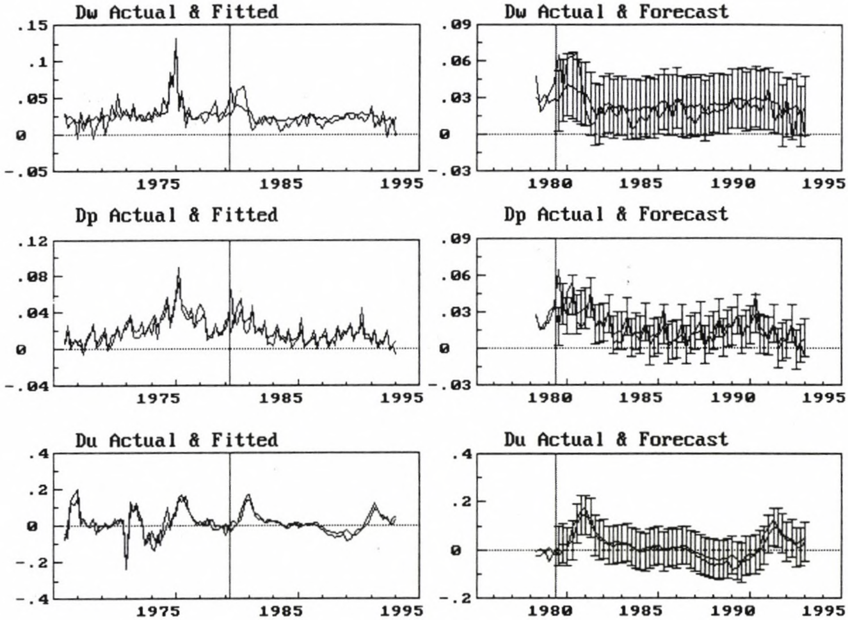

Figure 14. Simplified Model Actual, Fitted and Forecasts

but ignores sample information on the levels of these variables, the fact that the VAR has reduced rank, and that there is a regime shift. The PVAR (i.e. the reduced rank VAR mapped into I(0) space) and the simplified model on the other hand implicitly use sample information on the levels of the variables and their history, as well as the fact that the VAR does not have full rank, but ignore the regime shift. However, the latter two models forecast well because they use a full sample estimate of ecm_t which thus reflects the regime shift, and so keeps the forecasts on track. If instead ecm_t were replaced by the estimate of the cointegrating vector using data from 1966(1) to 1979(2) only, the forecast performance of all the models (other than the VAR in differences) deteriorates to be similar to that of the VAR. Table 13 gives the split sample unrestricted (ecm_t) estimates of the cointegrating vector, plus restricted estimates (ecm_t^*) satisfying the non-rejected overidentifying restrictions discussed below.

Table 13. Split Sample Estimates of the Cointegrating Vector

1966(1) - 1993(1)

$$ecm_t = w_t - 0.88p_t - 0.017u_t - 0.008t$$
$$ecm_t^* = w_t - 0.90p_t - 0.008t$$

1966(1) - 1979(2)

$$ecm1_t = w_t - 0.75p_t + 0.08u_t - 0.014t$$
$$ecm1_t^* = w_t - 0.75p_t + 0.06u_t - 0.014t$$

1979(3) - 1993(1)

$$ecm2_t = w_t - 0.94p_t + 0.06u_t - 0.006t$$
$$ecm2_t^* = w_t - p_t - 0.006t$$

Comparing $ecm1_t$ with ecm_t there is a 15% decrease in the coefficient of p_t, a 571% increase in the coefficient of u_t with a resultant change in sign (though neither estimate of this coefficient is significantly different from zero), and a 75% increase in the coefficient of the trend. Note that the response of wages and prices to unemployment in the period 1966(1) - 1979(2) is pro-cyclical, whereas that for the full sample and for the period after 1979(2) is counter-cyclical. Not surprisingly in view of this big difference between ecm_t and $ecm1_t$, the cointegrating vector estimated with data for the period 1979(3) to 1993(1) $ecm2_t$ is very different from $ecm1_t$ - see Table 13. In fact, though the Johansen Max and $Trace$ statistics clearly indicated one cointegrating vector for the periods 1966(1) - 1993(1) and 1966(1) - 1979(2), they only marginally rejected the hypothesis of no cointegrating vectors for the period 1979(3) - 1993(1). Therefore there is evidence of a fundamental change in the underlying relationship between w_t, p_t and u_t pre- and post- 1979(3). Indeed, there appears to have been a change in the exogeneity status of prices as the statistics in Table 14 indicate.

Table 14. Split Sample Overidentifying and Weak Exogeneity Tests

1966(1) - 1979(2)

Hypothesis	Statistic	p-value	Hypothesis	Statistic	p-value
$\beta_{12} = 0$	$\chi^2(1) = 19.64^{**}$	0.00	$\alpha_{11} = 0$	$\chi^2(1) = 5.78^{*}$	0.02
$\beta_{13} = 0$	$\chi_2(1) = 7.48^{**}$	0.01	$\alpha_{21} = 0$	$\chi^2(1) = 8.25^{**}$	0.00
$\beta_{14} = 0$	$\chi^2(1) = 14.61^{**}$	0.00	$\alpha_{31} = 0$	$\chi^2(1) = 1.69$	0.19
$\beta_{12} = -1$	$\chi^2(1) = 19.64^{**}$	0.00	$\alpha_{11} + \alpha_{21} = 0$	$\chi^2(1) = 2.10$	0.15

1979(3) -1993(1)

Hypothesis	Statistic	p-value	Hypothesis	Statistic	p-value
$\beta_{12} = 0$	$\chi^2(1) = 6.40^{*}$	0.01	$\alpha_{11} = 0$	$\chi^2(1) = 5.26^{*}$	0.02
$\beta_{13} = 0$	$\chi_2(1) = 0.87$	0.35	$\alpha_{21} = 0$	$\chi^2(1) = 0.00$	0.98
$\beta_{14} = 0$	$\chi^2(1) = 2.29$	0.13	$\alpha_{31} = 0$	$\chi^2(1) = 2.52$	0.11
$\beta_{12} = -1$	$\chi^2(1) = 0.08$	0.78	$\alpha_{11} + \alpha_{21} = 0$	$\chi^2(1) = 4.07^{*}$	0.04

The restriction $\alpha_{31} = 0$ (a necessary condition for u_t to be weakly exogenous for the parameters of the long run wage-price equation) is not rejected for any period, and so within this trivariate system unemployment appears to be weakly exogenous for the parameters of the one long run equilibrium equation throughout the sample period. However, the corresponding hypothesis for prices ($\alpha_{21} = 0$), whilst being rejected for the full sample and for the period up to 1979(2), is not rejected for the post-1979(2) period. This result is consistent with prices and wages, whilst exhibiting autonomous growth, being jointly determined conditional on unemployment up to 1979(2), but with prices (and hence inflation) being a target of government policy after 1979(3) determined separately from contemporaneous wages and unemployment. Such an interpretation is coherent with the Conservative government elected in May 1979 adopting an economic policy aimed at reducing radically the rate of inflation, pursuant to which it introduced a tight monetary policy, instigated new labour market arrangements, and was willing to accept the attendant sharp changes in unemployment. Indeed, the pre-1979(3) error correction when estimated subject to the non-rejected restrictions $\alpha_{31} = 0$ and $\alpha_{11} + \alpha_{21} = 0$ (the test statistic for this joint hypothesis is $\chi^2(2) = 2.10$ [$p > 0.35$]) takes the form of $ecm1_t^*$ in Table 14, which is very different from the post-1979(2) error correction $ecm2_t^*$ estimated subject to the restrictions $\alpha_{31} = 0, \alpha_{21} = 0, \beta_{13} = 0$ and $\beta_{12} = -1$ ($\chi^2(4) = 8.40$ [$p > 0.08$]). Hence there is evidence that u_t plays no role in the determination of the long run equilibrium values of w_t and p_t after

1979(2), and that w_t adjusts fully to p_t and in addition has an autonomous trend. Finally, notice that ecm_t and $ecm2_t$ (ecm_t^* and $ecm2_t^*$) are very similar and noticeably different from $ecm1_t$ ($ecm1_t^*$). This helps to understand why the one-step ahead forecasts for the period 1979(3) - 1993(1) from the PVAR and the simplified model both of which have ecm_t as an explanatory variable, are better than the corresponding forecasts from the VAR. In fact, if the single equation model for wages is re-parameterized to be in I (0) space, it too has good one-step ahead forecasts. In particular, if $w_{t-1}, p_{t-1}, u_{t-1}$ and t in the reduced model reported in Table 2, or w_{t-1}, p_{t-1} and t in the final model reported in Table 3, are replaced by ecm_{t-1} the resulting forecasts for Δw_t are very similar to those for Δw_t of the simplified model shown in Figure 14. However, the fact that p_t is not weakly exogenous for the parameters of the long run equilibrium $\mathsf{E}(ecm_t) = 0$ means that more efficient inferences should result from jointly modelling w_t and p_t, rather than using single equation OLS analysis. In fact, the forecasts for w_t from the system modelling should be preferable to those from the single equation modelling. This is borne out in practice as can be seen from comparison of the forecast error mean and standard deviation from the single equation analysis (-0.0167 and 0.0168 respectively) with those for the simplified model reported in Table 12.

Although the above empirical analysis illustrates the importance of testing for weak exogeneity, and modelling the joint distribution of variables in its absence, it is important to note that reasonable forecasts for the period 1966(1) - 1993(1) were only obtained by using the full sample estimate of the disequilibrium or error correction ecm_t, which would not be possible in practice for ex ante forecasting. The only model considered in this paper that would have been successful in ex ante forecasting is the DVAR. This highlights the fact that the presence of regime shifts in integrated-cointegrated systems for macroeconomic time series presents a substantial challenge to econometric modelling. Differencing time series that are subject to regime shifts does not account for the shifts even when they might be identifiable as changes in structure resulting from changes in government economic policies, but joint modelling of the differences in a VAR does enable the generation of multi-step forecasts that converge on the unconditional means of the series (see Hendry and Clements [1994]). Forecasts from econometric models that include variables which have not been suitably differenced to transform regime shifts from step or trend changes into impulses or blips which have no duration (such as ecm_t in the above illustration) may require intercept corrections or other adjustments to keep them on track when there are regime shifts. Seen in this light the use of intercept corrections can be interpreted as a means of exploiting past forecast errors to keep present forecasts on track (see Clements and Hendry [1994]).

It is instructive also to note that the long run disequilibrium variable ecm_t, despite the fact that it is commonly called an error correction mechanism, does not error-correct when it does not incorporate an explanation for regime shifts that are present in the system being modelled. Hence, ecm_t and $\mathsf{E}(ecm_t) = 0$ have been referred to as long run disequilibrium and long run equilibrium respectively. Ideally, econometric models should provide explanations for structural changes and account for autonomous shifts that affect the variables being modelled, and thus avoid the necessity of differencing such changes out of the system in order to produce reliable forecasts. Although this statement describes an enormous challenge to econometric modellers, those who overcome the challenge are likely to develop econometric models with great potential for use in economic policy analysis and forecasting. To rise to this challenge in the modelling of wages, prices, unemployment and other related variables such as productivity over the period 1966(1) - 1993(1) in the UK, it is likely that account will have to be taken of the important changes in the labour market structure, in exchange rate regimes, and government policies towards inflation and unemployment. The fact that the models presented in this section, have a single deterministic trend to approximate the effects of technical change, changes in productivity and external factors, and do not have long run equilibria with price homogeneity (except for $ecm2_t^*$ in Table 13) indicates that there is ample scope for further modelling. This important task lies beyond the scope of the present paper which has concentrated on the presentation of illustrative, rather than definitive, models of the relationship between wages, prices and unemployment.

5 Conclusions

The LSE modelling methodology developed rapidly during the 1960's and 1970's, and has had a significant effect on econometric modelling at large, as described in Gilbert [1986]. However, no vital methodology is static, but rather evolves with important developments. This is certainly true of the LSE methodology, and as suggested by Pagan [1992] a more accurate name today might be the LSE-Copenhagen-San Diego-Yale methodology - though even this ignores the valuable contributions of researchers in institutions in other geographical locations. The econometric modelling methodology associated with the LSE has achieved much, is still evolving and being refined, and provides a powerful tool for reaping the benefits of scientific econometric modelling. Though ultimately the model is the message, the system (*i.e.* the congruent representation of the joint distribution of the observed variables of relevance) is the appropriate statistical framework for developing and evaluating such econometric models. The legacy of this methodology includes the widely accepted need to rigorously evaluate econometric models by checking their congruence with available information and their ability to encompass rival models, as well as the advocacy of the general-to-specific modelling strategy as an efficient and efficacious modelling strategy.

Econometric concepts that provide important objectives to be achieved in the LSE methodology are homoscedastic innovation error processes, weak, strong and super exogeneity, encompassing, and parameter constancy. Failure to achieve these objectives in practice can result in econometric models that are likely to yield misleading conclusions from economic policy analyses, and produce unreliable forecasts. Indeed one of the most important challenges in the econometric modelling of macroeconomic time series is to produce multi-step forecasts that are at least as accurate as those generated from multivariate time series models such as DVAR's. This is a goal to which many of those espoused to the LSE methodology aspire. The simple illustration of empirical modelling presented above demonstrates the seriousness of this challenge, but also suggests that there is hope of future success. In addition, the illustration highlighted the important effect changes in economic policy can have on the inter-relationships between macroeconomic variables, especially on the underlying equilibrating mechanisms and the exogeneity status of variables within them. The change in economic policy from monitoring the rate of unemployment, to reducing radically the rate of inflation, had major effects on the UK economy post 1979.

The views presented in this paper on the nature and role of econometric modelling are mine. However, since I am associated with what has become known as the LSE methodology (or sometimes Professor Hendry's methodology see Gilbert (1986), Ericsson et al (1990)) this is consistent with my brief to write a paper describing the LSE methodology. I should like to thank my many friends and colleagues of the "LSE School" for the intellectual stimulation they have provided over the last quarter of a century. In particular, I wish to thank Mike Clements, David Hendry, Maozu Lu, and Jean-François Richard for allowing me to draw on our joint research without wishing to hold them responsible for the result. I am also grateful to David Hendry for providing many valuable suggestions whilst the paper was in gestation, and to Mike Clements and Kevin Hoover for providing detailed and constructive comments on an earlier version of the paper. Financial support for this research from the ESRC under grant R000231184, and from the EUI Research Council, is gratefully acknowledged.

6 Data Appendix

The quarterly data set runs from 1965(1) to 1993(1). The precise estimation periods used vary according to the number of lags in the specification. All series are seasonally adjusted except for the price series. The wage variable w is defined as $\log(WS/EE * AVH)$, where WS is wages and salaries, EE is employees in employment and AVH is a measure of average weekly hours in the manufacturing sector. The price variable p is the $\log(P)$ when P is the retail price index, so that Δp is the quarterly aggregate inflation rate. Finally, $u = \log(U)$, U being the % unemployment rate.

The precise definitions and sources are:

WS	Wages, salaries and forces pay. £mn. CSO mnemonic AIJB	ETAS for 65:1 to 90:4 ET for 91:1 to 93:1
EE	Employees in employment 000's, whole UK CSO mnemonic BCAJ	ETAS for 65:1 to 89:3 ET for 89:4 to 93:1
AVH	Index of average weekly hours worked per operative in manufacturing (1980=100)	EG, 65:1 to 93:1
P	Retail price index, all items (1987=100) CSO mnemonics FRAG and CHAW	ETAS for 65:1 to 92:2 ET for 92:3 to 93:1
U	Unemployment rate (UK) % CSO mnemonic BCJE	ETAS for 65:1 to 91:1 ET for 91:2 to 93:1

ETAS refers to Economic Trends (Annual Supplement) 1993 edition; ET refers to Economic Trends, August 1993 edition, both compiled by the Central Statistical Office (CSO). EG refers to the Employment Gazette Historical Supplements (No. 1, April 1985 Vol. 93 No. 4 and No. 2, Oct. 1987 Vol. 95 No. 10) and from various issues of the Gazette published by the Department of Employment.

References

[1] Aldrich, J. (1989), "Autonomy", *Oxford Economic Papers*, 41,15-34.

[2] Anderson, T.W. and Rubin, H. (1950), "The Asymptotic Properties of Estimates of the Parameters of a Single Equation in a Complete System of Stochastic Equations", *Annals of Mathematical Statistics*, 21, 570-582.

[3] Banerjee, A., Dolado, J.J., Galbraith, J.W. and Hendry, D.F. (1993), *Co-integration, Error-Correction, and the Econometric Analysis of Non-Stationary Data*, Oxford: Oxford University Press.

[4] Boswijk, H.P. (1992), *Cointegration, Identification and Exogeneity*, Amsterdam: Thesis Publishers.

[5] Box, G.E.P. and Jenkins, G.M. (1970), *Time Series Analysis Forecasting and Control*, Holden-Day: San Francisco.

[6] Brown, R.L., Durbin, J. and Evans, J.M. (1975), "Techniques for Testing the Constancy of Regression Relationships Over Time", *Journal of the Royal Statistical Society*, Series B, 37, 149-192.

[7] Campos, J., Ericsson, N.R. and Hendry, D.F. (1993), "Cointegration Tests in the Presence of Structural Breaks", International Finance Discussion Papers, No. 440, Board of Governors of the Federal Reserve System, February 1993.

[8] Canova, F. (1993), "The Economics of VAR Models", chapter 3 in K.D. Hoover (ed.), *Macroeconometrics: Developments, Tensions and Prospects*, Kluwer Academic Publishers.

[9] Canova, F., Finn, M. and Pagan, A.R. (1994), "Evaluating a Real Business Cycle Model", in C.P. Hargreaves (ed.), *Nonstationary Time Series Analyses and Cointegration,* Oxford: Oxford University Press, 225-255.

[10] Clements, M.P. and Hendry, D.F. (1994), "Towards a Theory of Economic Forecasting", in C.P. Hargreaves, (ed.), *Nonstationary Time Series Analyses and Cointegration*, Oxford: Oxford University Press, 9-52.

[11] Clements, M.P. and Mizon, G.E. (1991), "Empirical Analysis of Macroeconomic Time Series: VAR and Structural Models", *European Economic Review*, 35, 887-932.

[12] Cox, D.R. (1961), "Tests of Separate Families of Hypotheses", in Neyman, J. (ed.), *Proceedings of the Fourth Berkeley Symposium on Mathematical Statistics and Probability*, 1, 105-123, Berkeley:University of California Press.

[13] Cox, D.R. (1962), "Further Results of Tests of Separate Families of Hypotheses", *Journal of the Royal Statistical Society*, Series B, 24, 406-424.

[14] Desai, M.J. (1975), "The Phillips Curve: A Revisionist Interpretation", *Economica*, 42, 1-19.

[15] Desai, M.J. (1984), "Wages, Prices and Unemployment a Quarter of a Century After the Phillips Curve", chapter 9 in Hendry and Wallis (1984) op.cit., 251-273.

[16] Dickey, D.A. and Fuller, W.A.(1981), "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", *Econometrica*, 49, 1657-1672.

[17] Doornik, J.A. and Hansen, H. (1994), "A Practical Test for Univariate and Multivariate Normality", mimeo., Nuffield College, Oxford, March 1994.

[18] Doornik, J.A. and Hendry, D.F. (1992), *PC-GIVE Version 7: An Interactive Econometric Modelling System*, Institute of Economics and Statistics, University of Oxford.

[19] Doornik, J.A. and Hendry, D.F. (1993), *PC-FIML Version 7: An Interactive Econometric Modelling System*, Institute of Economics and Statistics, University of Oxford.

[20] Doornik, J.A. and Hendry, D.F. (1994), *PC-GIVE and PC-FIML Professional, Version 8: An Interactive Econometric Modelling System*, London: International Thomson.

[21] Engle, R.F. and Hendry, D.F. (1993), "Testing Super Exogeneity and Invariance in Regression Models", *Journal of Econometrics*, 56, 119-140.

[22] Engle, R.F., Hendry, D.F. and Richard, J-F. (1983), "Exogeneity", *Econometrica*, 51, 277-304.

[23] Ericsson, N.R. (1983), "Asymptotic Properties of Instrumental Variables Statistics for Testing Non-nested Hypotheses", *Review of Economic Studies*, 50, 287-303.

[24] Ericsson, N.R. (1993), "Cointegration, Exogeneity, and Policy Analysis: Representation, Estimation and Testing", *Journal of Policy Modeling*, 14, 251-280.

[25] Ericsson, N.R., Campos, J., and Tran, H-A. (1990), "PC-GIVE and David Hendry's Econometric Methodology", *Revista de Econometria*, 10, 7-117.

[26] Ericsson, N.R. and Hendry D.F. (1989), "Encompassing and Rational Expectations: How Sequential Corroboration Can Imply Refutation", International Finance Discussion Paper, No.354, Board of Governors of the Federal Reserve System, June 1989.

[27] Espasa, A. (1975), "A Wage-Earnings-Prices Inflation Model for the United Kingdom 1950-1970: Its Specification and Estimation by Classical and Spectral Methods", PhD thesis, London School of Economics.

[28] Favero, C. and Hendry, D.F. (1992), "Testing the Lucas Critique", *Econometric Reviews*, 11, 265-306.

[29] Gilbert, C.L. (1986), "Professor Hendry's Econometric Methodology", *Oxford Bulletin of Economics and Statistics*, 48, 283-307.

[30] Gilbert, C.L. (1989), "LSE and the British Approach to Time Series Econometrics", *Oxford Economic Papers*, 41, 108-128.

[31] Gourieroux, C., Monfort, A. (1992), "Testing, Encompassing and Simulating Dynamic Econometric Models", mimeo. INSEE.

[32] Granger, C.W.J. (1990) (ed.), *Modelling Economic Series. Readings in Econometric Methodology*, Oxford: Oxford University Press.

[33] Granger, C.W.J. and Newbold, P. (1974), "Spurious Regressions in Econometrics", *Journal of Econometrics*, 2, 111-120.

[34] Hansen, B.E. (1992), "Testing for Parameter Instability in Linear Models", *Journal of Policy Modeling,* 14, 4, 517-533.

[35] Hansen, H. and Johansen, S. (1993), "Recursive Estimation in Cointegrated VAR Models", Discussion Paper, Institute of Mathematical Statistics, University of Copenhagen.

[36] Hendry, D.F. (1971), "Maximum Likelihood Estimation of Systems of Simultaneous Regression Equations with Errors Generated by a Vector Autoregressive Process", *International Economic Review*, 12, 257-272, with correction in 15, p260.

[37] Hendry, D.F. (1974), "Stochastic Specification in an Aggregate Demand Model of the United Kingdom", *Econometrica*, 42, 559-578.

[38] Hendry, D.F. (1980), "Econometrics: Alchemy or Science?", *Economica*, 47, 387-406.

[39] Hendry, D.F. (1983), "Econometric Modelling: the Consumption Function in Retrospect", *Scottish Journal of Political Economy*, 30, 193-220.

[40] Hendry, D.F. (1987), "Econometric Methodology: A Personal Perspective", chapter 10 in Bewley, T.F. (ed.), *Advances in Econometrics*. Cambridge: Cambridge University Press, 29-48.

[41] Hendry, D.F. (1989), *PC GIVE: An Interactive Econometric Modelling System*, Oxford: Oxford Institute of Economics and Statistics.

[42] Hendry, D.F. (1993a), "On the Interactions of Unit Roots and Exogeneity", forthcoming in *Econometric Reviews.*

[43] Hendry, D.F. (1993b), " The Roles of Economic Theory and Econometrics in the Analysis of Economic Time Series", unpublished paper, Institute of Economics and Statistics, Oxford. Presented to the European Meeting of the Econometric Society, Uppsala, September 1993.

[44] Hendry, D.F. (1994), *Dynamic Econometrics*, Oxford: Oxford University Press.

[45] Hendry, D.F. and Anderson, G.J. (1977) "Testing Dynamic Specification in Small Simultaneous Systems: An Application to a Model of Building Society Behaviour in the United Kingdom", in Intriligator, M.D. (ed.), *Frontiers in Quantitative Economics Volume IIIa*, North Holland: Amsterdam, 361-383.

[46] Hendry, D.F. and Clements, M.P. (1994), "Can Econometrics Improve Economic Forecasting?", *Swiss Journal of Economics and Statistics,* 130, 267-298.

[47] Hendry, D.F. and Doornik, J.A. (1994), " Modelling Linear Dynamic Econometric Systems", *Scottish Journal of Political Economy*, 41, 1-33.

[48] Hendry, D.F. and Mizon, G.E. (1978), "Serial Correlation as a Convenient Simplification, not a Nuisance: a Comment on a Study of the Demand for Money by the Bank of England", *Economic Journal*, 88, 549-63.

[49] Hendry, D.F. and Mizon, G.E. (1990), "Procrustean Econometrics: or Stretching and Squeezing Data", chapter 7 in *Modelling Economic Series: Readings in Econometric Methodology*, edited by Granger, C.W.J. Oxford: Oxford University Press, 121-136.

[50] Hendry, D.F. and Mizon, G.E. (1993), "Evaluation of Dynamic Econometric Models by Encompassing the VAR", chapter 18 in *Models, Methods and Applications of Econometrics, Essays in Honor of Rex Bergstrom*, Phillips, P.C.B. (ed.), Oxford: Basil Blackwell, 272-300.

[51] Hendry, D.F. and Morgan, M.S. (eds.), (1995), *The Foundations of Econometric Analysis*, Cambridge: Cambridge University Press. Forthcoming.

[52] Hendry, D.F., Pagan, A.R. and Sargan, J.D. (1984), "Dynamic Specification", chapter 18 in *Handbook of Econometrics*, Volume 2, edited by Griliches, Z. and Intrilligator, M.D. Amsterdam: North-Holland, 1023-1100.

[53] Hendry, D.F. and Richard, J-F. (1982), "On the Formulation of Empirical Models in Dynamic Econometrics", *Journal of Econometrics*, 20, 3-33.

[54] Hendry, D.F. and Richard, J-F. (1983), "The Econometric Analysis of Economic Time Series", *International Statistical Review*, 51, 111-163.

[55] Hendry, D.F. and Richard, J-F. (1989), "Recent Developments in the Theory of Encompassing", in *Contributions to Operations Research and Econometrics. The Twentieth Anniversary of CORE*, Cornet, B. and Tulkens, H. (eds.), Cambridge, Mass.: MIT Press, 393-440.

[56] Hendry, D.F. and Srba, F. (1980), "AUTOREG: A Computer Program Library for Dynamic Econometric Models with Autoregressive Errors", *Journal of Econometrics*, 12, 85-102.

[57] Hendry, D.F. and Wallis, K.F. (eds.), (1984), *Econometrics and Quantitative Economics*, Oxford: Basil Blackwell.

[58] Hylleberg, S. and Mizon, G.E. (1989), "A Note on the Distribution of the Least Squares Estimator of a Random Walk with Drift", *Economics Letters*, 29,225-230.

[59] Johansen, S. (1988), "Statistical Analysis of Cointegration Vectors", *Journal of Economic Dynamics and Control*, 12, 231-54.

[60] Johansen, S. (1992a), "Cointegration in Partial Systems and the Efficiency of Single Equation Analysis", *Journal of Econometrics*, 52, 389-402.

[61] Johansen, S. (1992b), "Testing for Weak Exogeneity and the Order of Cointegration in the UK Money Demand Data", *Journal of Policy Modeling*, 14, 313-334.

[62] Johansen, S. and Juselius, K. (1990), "Maximum Likelihood Estimation and Inference on Cointegration - with Applications to the Demand for Money", *Oxford Bulletin of Economics and Statistics*, 52, 169-210.

[63] Kim, K. and Pagan, A.R. (1994), "The Econometric Analysis of Calibrated Macroeconomic Models", mimeo. Australian National University, March 1994.

[64] Koopmans, T.C. (ed.), (1950), *Statistical Inference in Dynamic Economic Models*, Cowles Commission Monograph No. 10, New York: Wiley.

[65] Kostial, K. (1994), "The Role of the Signal-Noise Matrix in the Estimation of Cointegrated Systems", European University Institute, Economics Department Discussion Paper No. EC94/27. Presented to the European Meeting of the Econometric Society, Maastricht, September 1994.

[66] Kydland, F. and Prescott, E. (1982), "Time to Build and Aggregate Fluctuations", *Econometrica,* 50, 1345-1370.

[67] Kydland, F. and Prescott, E. (1994), "The Econometrics of the General Equilibrium Approach to Business Cycles", chapter 5 in K.D. Hoover (ed.), *Macroeconomics: Developments, Tensions and Prospects*, Kluwer Academic Publishers.

[68] Layard, R. and Nickell, S.J. (1985), "The Causes of British Unemployment", *National Institute Economic Review*, 111, 62-85.

[69] Leamer, E.E. (1978) *Specification Searches: Ad Hoc Inference with Non-Experimental Data*, John Wiley: New York.

[70] Leamer, E.E. (1983), "Let's Take the Con Out of Econometrics", *American Economic Review*, 73, 31-44.

[71] Lipsey, R.G. (1960), "The Relation Between Unemployment and the Rate of Change of Money Wage Rates in the United Kingdom 1862-1957: a Further Analysis", *Economica,* 27, 1-31.

[72] Lu, M. and Mizon, G.E. (1991), "Forecast Encompassing and Model Evaluation", in Hackl, P. and Westlund, A. (eds.), *Economic Structural Change: Analysis and Forecasting*, Vienna: Springer-Verlag, 123-138.

[73] Lucas, R.E. (1976), "Econometric policy evaluation: a critique" in *The Phillips Curve and Labor Markets* ed. K. Brunner and A.H. Meltzer, Amsterdam: North Holland Publishing Company (Carnegie- Rochester Conference Series on Public Policy, Vol.1.), 19-46.

[74] Maasoumi, E. (1988), *Contributions to Econometrics: John Denis Sargan*, Volumes 1 and 2, Cambridge: Cambridge University Press.

[75] Mizon, G.E. (1977a), "Model Selection Procedures in Dynamic Models", chapter 4 in *Studies in Modern Economic Analysis*, edited by M.J. Artis and A.R. Nobay, Oxford: Basil Blackwell, 97-120.

[76] Mizon, G.E. (1977b), "Inferential Procedures in Nonlinear Models: An Application in a UK Cross Sectional Study of Factor Substitution and Returns to Scale", *Econometrica*, 45, 1221-1242.

[77] Mizon, G.E. (1984), "The Encompassing Approach in Econometrics", chapter 6 in D.F. Hendry and K.F. Wallis (eds) *Econometrics and Quantitative Economics*, Oxford: Basil Blackwell, 135-172.

[78] Mizon, G.E. (1991), "The Role of Measurement and Testing in Economics", chapter 28 in *Companion to Contemporary Economic Thought*, edited by M. Bleaney, D. Greenaway and I M Stewart, London: Routledge, 574-592.

[79] Mizon, G.E., (1993), "A Simple Message for Autocorrelation Correctors: DON'T", forthcoming in *Journal of Econometrics*.

[80] Mizon, G.E. and Hendry, D.F. (1980), "An Empirical Application and Monte Carlo Analysis of Tests of Dynamic Specification", *Review of Economic Studies*, 49, 21-45.

[81] Mizon, G.E. and Richard, J-F. (1986), "The Encompassing Principle and its Application to Non-Nested Hypothesis Tests", *Econometrica*, 54, 657-678.

[82] Monfort, A. and Rabemananjara, R. (1990), "From a VAR Model to a Structural Model: with an Application to the Wage-Price Spiral", *Journal of Applied Econometrics*, 5, 203-227.

[83] Nickell, S.J. (1984), "The Modelling of Wages and Employment", chapter 2 in Hendry and Wallis (1984) op. cit., 13-35.

[84] Osterwald-Lenum, M.(1992),"A Note With Quantiles of the Asymptotic Distribution of the ML Cointegration Rank Test Statistics", *Oxford Bulletin of Economics and Statistics,* 54, 461-472.

[85] Pagan, A.R. (1987), "Three Econometric Methodologies: A Critical Appraisal", *Journal of Economic Surveys*, 1, 3-24, reprinted in C.W.J. Granger (1990), op. cit., 97-120.

[86] Pagan, A.R. (1989), "Twenty Years After: Econometrics, 1966-1986", in *Contributions to Operations Research and Econometrics . The Twentieth Anniversary of CORE*, edited by B. Cornet and H. Tulkens. Cambridge, Mass.: MIT Press, 319-383.

[87] Pagan, A.R. (1992), "Three Econometric Methodologies: An Update", mimeo. Australian National University.

[88] Phillips, A.W.H. (1958), "The Relation Between unemployment and the Rate of Change of Money Wages Rates in the United Kingdom 1861-1957", *Economica*, 25, 283-299.

[89] Phillips, P.C.B. (1988), "The *ET* Interview; A.R. Bergstrom", *Econometric Theory*, 4, 301-327.

[90] Phillips, P.C.B. (ed.), (1993a), *Models, Methods, and Applications of Econometrics. Essays in Honor of A.R. Bergstrom*, Cambridge: Basil Blackwell.

[91] Phillips, P.C.B. (1993b), "Rex Bergstrom's Career and Research", in Phillips (1993a) op. cit., 3-8.

[92] Prothero, D.L. and Wallis, K.F. (1976), "Modelling Macro-Économic Time Series (with discussion)", *Journal of the Royal Statistical Society*, Series A, 139, 468-500.

[93] Reimers, H.-E. (1992), "Comparisons of Tests for Multivariate Cointegration", forthcoming *Statistical Papers*.

[94] Richard, J-F. (1980), "Models with Several Regimes and Changes in Exogeneity", *Review of Economic Studies*, 47, 1-20.

[95] Sargan, J.D. (1964), "Wages and Prices in the United Kingdom: A Study in Econometric Methodology", in Hart, P.E. Mills, G. and Whitaker, J.K. (eds.), *Econometric Analysis for National Economic Planning*, Butterworths, London. Reprinted in Hendry and Wallis (1984) op. cit., 275-314.

[96] Sargan, J.D. (1980a), "A Model of Wage-Price Inflation", *Review of Economic Studies*, 47, 97-112.

[97] Sargan, J.D. (1980b), "Some Tests of Dynamic Specification for a Single Equation", *Econometrica*, 48, 879-897.

[98] Sargan, J.D. (1988), *Lectures on Advanced Econometric Theory*, Oxford: Basil Blackwell.

[99] Sims, C.A. (1980), "Macroeconomics and Reality", *Econometrica*, 48, 1-48. Reprinted in Granger (1990) op. cit., 137-189.

[100] Sims, C.A. (1992), "Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy", *European Economic Review*, 36, 975-1000.

[101] Spanos, A. (1986), *Statistical Foundations of Econometric Modelling*, Cambridge: Cambridge University Press.

[102] Summers, L.H. (1991), "The Scientific Illusion in Empirical Macroeconomics", *Scandinavian Journal of Economics,* 93, 129-148.

[103] Terasvirta, T. (1970), "Stepwise Regression and Economic Forecasting", Economic Studies Monograph No.31, Helsinki: Finnish Economics Association.

[104] Urbain, J.P. (1992), "On Weak Exogeneity in Error Correction Models", *Oxford Bulletin of Economics and Statistics*, 54, 187-208.

[105] Vernon, K. (1970), "An Econometric Study of Wage and Price Inflation in the United Kingdom in the Post War Period", PhD thesis London School of Economics.

[106] West, K.D. (1988), "Asymptotic Normality When Regressors Have a Unit Root", *Econometrica,* 56, 1397-1417.

[107] Zellner, A. and Palm, F. (1974), "Time Series Analysis and Simultaneous Equation Econometric Models", *Journal of Econometrics*, 2, 17-54.

EUI WORKING PAPERS

EUI Working Papers are published and distributed by the
European University Institute, Florence

Copies can be obtained free of charge
– depending on the availability of stocks – from:

The Publications Officer
European University Institute
Badia Fiesolana
I-50016 San Domenico di Fiesole (FI)
Italy

Please use order form overleaf

Publications of the European University Institute

Department of Economics Working Paper Series

To Department of Economics **WP**
European University Institute
Badia Fiesolana
I-50016 San Domenico di Fiesole (FI)
E-mail: publish@datacomm.iue.it
Italy

From Name . . . .

Address. . . . .

. . . .

. . . .

. . . .

. . . .

(Please print)

❐ Please enter/confirm my name on EUI Economics Dept. Mailing List
❐ Please send me a complete list of EUI Working Papers
❐ Please send me a complete list of EUI book publications
❐ Please send me the EUI brochure Academic Year 1995/96

Please send me the following EUI **ECO** Working Paper(s):

No, Author . . . .
Title: . . . .
No, Author . . . .
Title: . . . .
No, Author . . . .
Title: . . . .
No, Author . . . .
Title: . . . .

Date . . . . Signature . . . .

Working Papers of the Department of Economics Published since 1993

ECO No. 93/1
Carlo GRILLENZONI
Forecasting Unstable and Non-Stationary Time Series

ECO No. 93/2
Carlo GRILLENZONI
Multilinear Models for Nonlinear Time Series

ECO No. 93/3
Ronald M. HARSTAD/Louis PHLIPS
Futures Market Contracting When You Don't Know Who the Optimists Are

ECO No. 93/4
Alan KIRMAN/Louis PHLIPS
Empirical Studies of Product Markets

ECO No. 93/5
Grayham E. MIZON
Empirical Analysis of Time Series: Illustrations with Simulated Data

ECO No. 93/6
Tilman EHRBECK
Optimally Combining Individual Forecasts From Panel Data

ECO NO. 93/7
Víctor GÓMEZ/Agustín MARAVALL
Initializing the Kalman Filter with Incompletely Specified Initial Conditions

ECO No. 93/8
Frederic PALOMINO
Informed Speculation: Small Markets Against Large Markets

ECO NO. 93/9
Stephen MARTIN
Beyond Prices Versus Quantities

ECO No. 93/10
José María LABEAGA/Angel LÓPEZ
A Flexible Demand System and VAT Simulations from Spanish Microdata

ECO No. 93/11
Maozu LU/Grayham E. MIZON
The Encompassing Principle and Specification Tests

ECO No. 93/12
Louis PHLIPS/Peter MØLLGAARD
Oil Stocks as a Squeeze Preventing Mechanism: Is Self-Regulation Possible?

ECO No. 93/13
Pieter HASEKAMP
Disinflation Policy and Credibility: The Role of Conventions

ECO No. 93/14
Louis PHLIPS
Price Leadership and Conscious Parallelism: A Survey

ECO No. 93/15
Agustín MARAVALL
Short-Term Analysis of Macroeconomic Time Series *

ECO No. 93/16
Philip Hans FRANSES/Niels HALDRUP
The Effects of Additive Outliers on Tests for Unit Roots and Cointegration

ECO No. 93/17
Fabio CANOVA/Jane MARRINAN
Predicting Excess Returns in Financial Markets

ECO No. 93/18
Iñigo HERGUERA
Exchange Rate Fluctuations, Market Structure and the Pass-through Relationship

ECO No. 93/19
Agustín MARAVALL
Use and Misuse of Unobserved Components in Economic Forecasting

ECO No. 93/20
Torben HOLVAD/Jens Leth HOUGAARD
Measuring Technical Input Efficiency for Similar Production Units: A Survey of the Non-Parametric Approach

*out of print

ECO No. 93/21
Stephen MARTIN/Louis PHLIPS
Product Differentiation, Market Structure and Exchange Rate Passthrough

ECO No 93/22
F. CANOVA/M. FINN/A. R. PAGAN
Evaluating a Real Business Cycle Model

ECO No 93/23
Fabio CANOVA
Statistical Inference in Calibrated Models

ECO No 93/24
Gilles TEYSSIÈRE
Matching Processes in the Labour Market in Marseilles. An Econometric Study

ECO No 93/25
Fabio CANOVA
Sources and Propagation of International Business Cycles: Common Shocks or Transmission?

ECO No. 93/26
Marco BECHT/Carlos RAMÍREZ
Financial Capitalism in Pre-World War I Germany: The Role of the Universal Banks in the Financing of German Mining Companies 1906-1912

ECO No. 93/27
Isabelle MARET
Two Parametric Models of Demand, Structure of Market Demand from Heterogeneity

ECO No. 93/28
Stephen MARTIN
Vertical Product Differentiation, Intra-industry Trade, and Infant Industry Protection

ECO No. 93/29
J. Humberto LOPEZ
Testing for Unit Roots with the k-th Autocorrelation Coefficient

ECO No. 93/30
Paola VALBONESI
Modelling Interactions Between State and Private Sector in a "Previously" Centrally Planned Economy

ECO No. 93/31
Enrique ALBEROLA ILA/J. Humberto LOPEZ/Vicente ORTS RIOS
An Application of the Kalman Filter to the Spanish Experience in a Target Zone (1989-92)

ECO No. 93/32
Fabio CANOVA/Morten O. RAVN
International Consumption Risk Sharing

ECO No. 93/33
Morten Overgaard RAVN
International Business Cycles: How much can Standard Theory Account for?

ECO No. 93/34
Agustín MARAVALL
Unobserved Components in Economic Time Series *

ECO No. 93/35
Sheila MARNIE/John MICKLEWRIGHT
Poverty in Pre-Reform Uzbekistan: What do Official Data Really Reveal? *

ECO No. 93/36
Torben HOLVAD/Jens Leth HOUGAARD
Measuring Technical Input Efficiency for Similar Production Units: 80 Danish Hospitals

ECO No. 93/37
Grayham E. MIZON
A Simple Message for Autocorrelation Correctors: DON'T

ECO No. 93/38
Barbara BOEHNLEIN
The Impact of Product Differentiation on Collusive Equilibria and Multimarket Contact

ECO No. 93/39
H. Peter MØLLGAARD
Bargaining and Efficiency in a Speculative Forward Market

***

*out of print

ECO No. 94/1
Robert WALDMANN
Cooperatives With Privately Optimal Price Indexed Debt Increase Membership When Demand Increases

ECO No. 94/2
Tilman EHRBECK/Robert WALDMANN
Can Forecasters' Motives Explain Rejection of the Rational Expectations Hypothesis?

ECO No. 94/3
Alessandra PELLONI
Public Policy in a Two Sector Model of Endogenous Growth *

ECO No. 94/4
David F. HENDRY
On the Interactions of Unit Roots and Exogeneity

ECO No. 94/5
Bernadette GOVAERTS/David F. HENDRY/Jean-François RICHARD
Encompassing in Stationary Linear Dynamic Models

ECO No. 94/6
Luigi ERMINI/Dongkoo CHANG
Testing the Joint Hypothesis of Rationality and Neutrality under Seasonal Cointegration: The Case of Korea

ECO No. 94/7
Gabriele FIORENTINI/Agustín MARAVALL
Unobserved Components in ARCH Models: An Application to Seasonal Adjustment *

ECO No. 94/8
Niels HALDRUP/Mark SALMON
Polynomially Cointegrated Systems and their Representations: A Synthesis

ECO No. 94/9
Mariusz TAMBORSKI
Currency Option Pricing with Stochastic Interest Rates and Transaction Costs: A Theoretical Model

ECO No. 94/10
Mariusz TAMBORSKI
Are Standard Deviations Implied in Currency Option Prices Good Predictors of Future Exchange Rate Volatility?

ECO No. 94/11
John MICKLEWRIGHT/Gyula NAGY
How Does the Hungarian Unemployment Insurance System Really Work? *

ECO No. 94/12
Frank CRITCHLEY/Paul MARRIOTT/Mark SALMON
An Elementary Account of Amari's Expected Geometry

ECO No. 94/13
Domenico Junior MARCHETTI
Procyclical Productivity, Externalities and Labor Hoarding: A Reexamination of Evidence from U.S. Manufacturing

ECO No. 94/14
Giovanni NERO
A Structural Model of Intra-European Airline Competition

ECO No. 94/15
Stephen MARTIN
Oligopoly Limit Pricing: Strategic Substitutes, Strategic Complements

ECO No. 94/16
Ed HOPKINS
Learning and Evolution in a Heterogeneous Population

ECO No. 94/17
Berthold HERRENDORF
Seigniorage, Optimal Taxation, and Time Consistency: A Review

ECO No. 94/18
Frederic PALOMINO
Noise Trading in Small Markets *

ECO No. 94/19
Alexander SCHRADER
Vertical Foreclosure, Tax Spinning and Oil Taxation in Oligopoly

ECO No. 94/20
Andrzej BANIAK/Louis PHLIPS
La Pléiade and Exchange Rate Pass-Through

ECO No. 94/21
Mark SALMON
Bounded Rationality and Learning; Procedural Learning

*out of print

ECO No. 94/22
Isabelle MARET
Heterogeneity and Dynamics of Temporary Equilibria: Short-Run Versus Long-Run Stability

ECO No. 94/23
Nikolaos GEORGANTZIS
Short-Run and Long-Run Cournot Equilibria in Multiproduct Industries

ECO No. 94/24
Alexander SCHRADER
Vertical Mergers and Market Foreclosure: Comment

ECO No. 94/25
Jeroen HINLOOPEN
Subsidising Cooperative and Non-Cooperative R&D in Duopoly with Spillovers

ECO No. 94/26
Debora DI GIOACCHINO
The Evolution of Cooperation: Robustness to Mistakes and Mutation

ECO No. 94/27
Kristina KOSTIAL
The Role of the Signal-Noise Ratio in Cointegrated Systems

ECO No. 94/28
Agustín MARAVALL/Víctor GÓMEZ
Program SEATS "Signal Extraction in ARIMA Time Series" - Instructions for the User

ECO No. 94/29
Luigi ERMINI
A Discrete-Time Consumption-CAP Model under Durability of Goods, Habit Formation and Temporal Aggregation

ECO No. 94/30
Debora DI GIOACCHINO
Learning to Drink Beer by Mistake

ECO No. 94/31
Víctor GÓMEZ/Agustín MARAVALL
Program TRAMO "Time Series Regression with ARIMA Noise, Missing Observations, and Outliers" - Instructions for the User

ECO No. 94/32
Ákos VALENTINYI
How Financial Development and Inflation may Affect Growth

ECO No. 94/33
Stephen MARTIN
European Community Food Processing Industries

ECO No. 94/34
Agustín MARAVALL/Christophe PLANAS
Estimation Error and the Specification of Unobserved Component Models

ECO No. 94/35
Robbin HERRING
The "Divergent Beliefs" Hypothesis and the "Contract Zone" in Final Offer Arbitration

ECO No. 94/36
Robbin HERRING
Hiring Quality Labour

ECO No. 94/37
Angel J. UBIDE
Is there Consumption Risk Sharing in the EEC?

ECO No. 94/38
Berthold HERRENDORF
Credible Purchases of Credibility Through Exchange Rate Pegging: An Optimal Taxation Framework

ECO No. 94/39
Enrique ALBEROLA ILA
How Long Can a Honeymoon Last? Institutional and Fundamental Beliefs in the Collapse of a Target Zone

ECO No. 94/40
Robert WALDMANN
Inequality, Economic Growth and the Debt Crisis

ECO No. 94/41
John MICKLEWRIGHT/ Gyula NAGY
Flows to and from Insured Unemployment in Hungary

*out of print

ECO No. 94/42
Barbara BOEHNLEIN
The Soda-ash Market in Europe: Collusive and Competitive Equilibria With and Without Foreign Entry

ECO No. 94/43
Hans-Theo NORMANN
Stackelberg Warfare as an Equilibrium Choice in a Game with Reputation Effects

ECO No. 94/44
Giorgio CALZOLARI/Gabriele FIORENTINI
Conditional Heteroskedasticity in Nonlinear Simultaneous Equations

ECO No. 94/45
Frank CRITCHLEY/Paul MARRIOTT/Mark SALMON
On the Differential Geometry of the Wald Test with Nonlinear Restrictions

ECO No. 94/46
Renzo G. AVESANI/Giampiero M. GALLO/Mark SALMON
On the Evolution of Credibility and Flexible Exchange Rate Target Zones

***

ECO No. 95/1
Paul PEZANIS-CHRISTOU
Experimental Results in Asymmetric Auctions - The 'Low-Ball' Effect

ECO No. 95/2
Jeroen HINLOOPEN/Rien WAGENVOORT
Robust Estimation: An Example

ECO No. 95/3
Giampiero M. GALLO/Barbara PACINI
Risk-related Asymmetries in Foreign Exchange Markets

ECO No. 95/4
Santanu ROY/Rien WAGENVOORT
Risk Preference and Indirect Utility in Portfolio Choice Problems

ECO No. 95/5
Giovanni NERO
Third Package and Noncooperative Collusion in the European Airline Industry

ECO No. 95/6
Renzo G. AVESANI/Giampiero M. GALLO/Mark SALMON
On the Nature of Commitment in Flexible Target Zones and the Measurement of Credibility: The 1993 ERM Crisis

ECO No. 95/7
John MICKLEWRIGHT/Gyula NAGY
Unemployment Insurance and Incentives in Hungary

ECO No. 95/8
Kristina KOSTIAL
The Fully Modified OLS Estimator as a System Estimator: A Monte-Carlo Analysis

ECO No. 95/9
Günther REHME
Redistribution, Wealth Tax Competition and Capital Flight in Growing Economies

ECO No. 95/10
Grayham E. MIZON
Progressive Modelling of Macroeconomic Time Series: The LSE Methodology

*out of print