EUI WORKING PAPERS IN ECONOMICS

EUI Working Paper ECO No. 93/22

Evaluating a Real Business Cycle Model

F. Canova, M. Finn
and
A. R. Pagan

European University Institute, Florence

Please note
As from January 1990 the EUI Working Paper Series is divided into six sub-series, each sub-series is numbered individually (e.g. EUI Working Paper **LAW** No. 90/1).

EUROPEAN UNIVERSITY INSTITUTE, FLORENCE

ECONOMICS DEPARTMENT

EUI Working Paper **ECO** No. 93/22

Evaluating a Real Business Cycle Model

F. CANOVA, M. FINN
and
A. R. PAGAN

BADIA FIESOLANA, SAN DOMENICO (FI)

Printed in Italy in May 1993
European University Institute
Badia Fiesolana
I – 50016 San Domenico (FI)
Italy

EVALUATING A REAL BUSINESS CYCLE MODEL

F. Canova*
European University Institute

M. Finn
Federal Reserve Bank of Richmond

A.R. Pagan
Australian National University

May 5, 1993

Abstract

This paper provides a method to formally evaluate models with calibrated parameters. We examine whether the restricted VAR representation implied by the model is consistent with the data. Two types of restrictions are analyzed: one due to the fact that there are less forcing processes than endogenous variables and one obtained factoring out the first type of restrictions. We also propose a way to evaluate the performance of a model relative to a competitor. An application to the model analyzed by Burnside, Eichenbaum and Rebelo (1992) is considered.

*We would like to thank John Robertson, Peter Hartley, Robert King and Steve Durlauf for their comments on an earlier version of this paper.

1. Introduction

In the last decade real business cycles (RBC) models have gone from the preliminary explorations of Long and Plosser (1983) and Kydland and Prescott (1982) to well developed and tested models such as Burnside, Eichenbaum and Rebelo (1990) and McGrattan (1991). Early models could be regarded as "idealized", in the sense adopted in the philosophical literature summarized in Hoover (1991a), in that they were "simplifications that were designed to isolate an **essential** core"; in this instance attempts to capture the characteristics of fluctuations within industrial economies. Given such an objective it was appropriate that the method employed to determine whether the "essence" of an economy had been captured or not was the method of "stylized facts". In this procedure a certain number of key "facts" are identified and subsequently used to gauge the performance of the model. Thus Long and Plosser concentrated upon the idea that business cycles generated co–movements between de–trended variables, and they asked whether it was possible to obtain such a feature with the very simple RBC model that they had constructed. Others have been somewhat more precise, asking if the variances and covariances between variables such as output, consumption and real wages observed in the U.S. economy agreed with the predictions of their model. In extensions of this early work, *e.g.* King, Plosser and Rebelo (1988), a similar strategy was adopted to that of Prescott and Kydland, but with a more extensive range of stylized facts to be explained.

Stylized facts are obviously a good way of evaluating idealized models. By their very nature the latter models are not meant to provide a complete description of any time series such as consumption or output, but rather attempt to emulate a few of the major characteristics of those variables. Nevertheless, even with such a limited objective, there still remains an important practical problem of determining just how well the models are emulating reality, and this necessitates the development of some "metric" for that task. Because RBC models are explicitly stochastic a number of measures have been proposed that involve computing standard errors for the model predictions, either by analytic means or by computer based simulation *e.g.* Gregory and Smith (1991), Canova (1990).

Early comparisons of model projections with stylized facts revealed that the models did not adequately account for the latter. Perhaps the most striking failures were the correlation of productivity with hours worked and of government consumption with the Solow residual. Stimulated by this fact researchers in the area began to develop the models in a number of different directions, with the aim of getting a better match with the stylized facts. As discussed by Hoover (1991a) this development can be thought of as "concretizing" the idealized models so as to make for a better correspondence with the "real world". "Concretization" has now been

performed in many different directions and there has been substantial success in clearing up some of the striking failures of the early models.

The developments described above are reminiscent of early work with macroeconometric models. Initially, the desire was to explain some very broad characteristics of the data. As ambitions rose and simple models were replaced by large scale ones in an attempt to capture real world complexities, it was necessary to devise tests of the latter that were much more demanding, so as to try to isolate where the deficiencies of the models lay. It seems appropriate therefore that the attempts at evaluation via stylized facts, which has characterized most RBC studies to date, should also be replaced by more demanding and comprehensive tests, particularly since these models are progressively "concretized" in order to account for specific "stylized facts". What makes this task different to the older econometric literature is that RBC models are models with a great deal of internal dependence, and it is very hard to evaluate the components separately; one is inevitably faced with the need to work with the whole model. Consequently, many of the "single equation" tests that have been used so effectively when evaluating large scale macroeconometric models are difficult to apply, since one could not make a modification to a "part" of the RBC model without affecting it somewhere else. Complete model evaluation methods are the logical way to proceed.

This paper is an attempt to do the requisite analysis along these lines. It is well known that RBC models involve a VAR in the variables — see Long and Plosser (1983) for example. Furthermore, as we observe in section 2, it is a highly restricted VAR. Thus, just as for the rational expectations models considered by Sargent (1978), it seems as if a sensible way to evaluate the models is to test the restrictions on the implied VAR. Although the idea is straightforward, one has to be somewhat more cautious. Frequently, the driving forces in these models are integrated, and the VAR is actually a vector ECM, due to there being a smaller number of driving forces than variables being explained. If the driving forces are integrated, analysis suggests that there are two types of restrictions that might be tested. First, there are the co–integrating restrictions stemming from the fact that there are generally more variables to be modelled than there are independent integrated forcing processes. Second, there are restrictions upon the dynamics which apply to the system written after factoring out the co–integrating relations. Section 2 develops these ideas.

Section 3 of the paper takes a particular RBC model, that due to Burnside et al. and applies the ideas developed earlier to it. This model was chosen because there have been a number of concretizing steps taken to make it emulate the real world, although there remains some doubt over whether it actually agrees with a comprehensive range of stylized facts. Our claim is that consideration of the two types of restrictions described above, and a determination of whether they are acceptable, can be a very useful input into a modelling exercise. In particular, such

information can highlight deficiencies in the models and may suggest suitable re-specifications. In our example, we find that the BER model is strongly rejected by the data and we enquire into what changes might be made to the model to produce a VAR that more closely appoximates what is seen in the data. Finally, in section 4 we ask the question of how well the RBC model functions relative to a simple model such as a multiplier–accelerator mechanism and discuss whether the latter is any more successful in reproducing the VAR than the RBC model is. The viewpoint of this section is that, ultimately, the relevant question pertains to the relative rather than absolute quality of a model. Such comparisons are also likely to yield better information about potential respecifications. Finally, section 5 concludes with some suggestions about how the RBC model might be modified to produce a better fit to the data.

2. Testing the Restrictions of an RBC Model

Define y_t as the (qx1) vector of variables of interest, z_t as the (nx1) vector of controlled and uncontrolled state variables, and x_t as the (px1) vector of exogenous or forcing variables (the uncontrolled states). Most RBC models can be regarded as conforming to a linear structure of the form

$$y_t = Az_t \tag{1}$$

$$z_t = Fz_{t-1} + Ge_t, \tag{2}$$

where e_t are the innovations into the forcing variables and G is a vector showing how these innovations impinge upon the state variables. Generally G is a matrix that only has rank p *i.e.* there are more state variables than there are stochastic elements in (2). The linearity of the system stems from the fact that these systems are frequently solved by either linearizing the Euler equations around the steady state, as in King, Plosser, and Rebelo (1988), or solving the Riccati equation associated with the linear/quadratic control problem, a method employed by McGratten. It is possible to argue that (1) and (2) are more general than they might appear to be in that some types of non–linearities might be accomodated *e.g.* z_t might be functions of state variables. Some of the solution methods, such as Marcet (1989) or Chow (1992a), can allow this interpretation. Higher order dynamics can also be incorporated, but, since the application given later has first order dynamics, the discussion will focus upon the special case.

An especially important characteristic of many RBC models is that F and A are functions of a smaller number of parameters such as utility and production function parameters, and the latter are typically selected by some "calibration" strategy. It is hard to be precise about exactly what the latter is as it ranges from

selecting parameter values gleaned from micro or macro studies — estimated either by sample averages or by methods such as GMM and FIML — to "guesstimates". We will simply assume that A and F have precise numercial values assigned to them, so that an RBC model is both a set of relationships as in (1) and (2) and a specific set of values for the parameters A and F. Of course, this is true of any macroeconomic model. Nevertheless, one might argue over whether the parameter values should be taken as capturing the "essence" of an economy or are simply concretizing assumptions *i.e.* perhaps what should be tested is the general format in (1) and (2) rather than the particular structure coming from specific values of A and F. As an example of the difference, suppose y_t was consumption and z_t was output. Then (1) can be interpreted as either saying that the average propensity to consume is **exactly** A or that the average propensity to consume is simply some unknown **constant** A. Although there are some testable implications of the latter viewpoint, they are obviously very weak, and it is likely that many models would yield such a prediction *e.g.* there were many early consumption relations that were not inter–temporal but which would imply constancy of the consumption ratio. Hence, as a way of distinguishing between different theories, it seems necessary to maintain that the numbers assigned to A and F are parts of the model. One could plausibly argue against this strategy if A and F were estimated directly from data, but since they are functions of a much smaller number of "deep parameters", the power of the RBC model presumably derives from just this fact. Indeed that seems to have been an essential ingredient in the original arguments put forth for such models in Long and Plosser (1983) and by Prescott (1991) who imposes parameter values as a consequence of steady state relations.

It is necesssary to distinguish between two scenarios for (1) and (2) depending upon the nature of the forcing variables x_t. In many applications of RBC models x_t are made I(1) processes, generally independent of each other, *i.e.*

$$x_t = x_{t-1} + e_t. \tag{3}$$

Under this specification the structure of F is $F = \begin{bmatrix} \gamma & \delta \\ 0 & I_p \end{bmatrix}$ so that p of the eigenvalues of F are unity while the remaining $(n-p)$ are the eigenvalues of γ. In RBC models the latter are less than unity, implying that there must be $(n-p)$ co–integrating vectors among the z_t. Defining the elements of z_t which exclude x_t as z_{1t}, if the z_{1t} are I(1) then it follows immediately that the co–integrating vectors are $[\ (I-\gamma)\ -\delta]$; alternatively, if any of the z_{1t} is I(0), the corresponding row, $(\gamma_1\ \delta_1)$, must be a co–integrating vector. Identifying z_t with some observed data this would be a **first prediction** of the RBC model. It is also apparent that there are some Granger Causality predictions which stem from (3).

Equation (1) predicts that an exact relation should hold between y_t and z_t. Such an exact relation is unlikely to be observed with any set of data and it is

important to weaken (1) so as to allow it to be non–exact. The most appropriate extension would seem to be to assume that $y_t - Az_t$ is an I(0) process. There are two arguments one might make in favour of this stance. The first is that the RBC model aims at capturing the essential mechanisms at work in the economy, and, **prima facie**, this suggests that what is left out should be distinguishable as something of less importance than what is retained. When z_t is integrated it is natural therefore to think that what has been ignored should be non–integrated. Second, if one thought of observed data as being different from the model constructs due to measurement error, it is natural to make the measurement error an I(0) process when the variable being incorrectly measured is I(1). Therefore, in terms of either argument, $(I_q - A)$ should be a set of co–integrating vectors, and this is a **second testable implication** of an RBC model. Note that what we have is not only the requirement that y_t and z_t be co–integrated but that they be co–integrated with the numerical values assigned to A.

A second set of restrictions implied by RBC models involves the dynamic structure, or what will be termed the "non–cointegrating restrictions". To derive these write (1) and (2) as

$$\Delta y_t = A\Delta z_t, \tag{4}$$

where

$$\Delta z_t = (F - I)z_{t-1} + Ge_t \tag{5}$$

$$= \Pi z_{t-1} + Ge_t \tag{6}$$

$$= \alpha\beta' z_{t-1} + Ge_t \tag{7}$$

and β are the co–integrating vectors existing among the z_t. Substituting (6) into (5) yields

$$\Delta y_t = A\alpha\beta' z_{t-1} + AGe_t \tag{8}$$

and, forming the co–integrating error $v_t = \beta' z_{t-1}$, we have

$$\Delta y_t = A\alpha v_{t-1} + AGe_t, \tag{9}$$

which is a relation solely between I(0) variables. Defining $w_t' = (\Delta y_t' \; v_t')$ the VAR in w_t implied by an RBC model therefore has two characteristics. First, unless y_t is a state variable, Δy_{t-1} is excluded from it. Second, the coefficients of v_{t-1} are given by $A\alpha$. These are the **third set of testable predictions**, and they concern the non–cointegrating restrictions. Notice that the restrictions stem from the dynamic nature of the model, provided we have previously accepted that the co–integrating restrictions are valid ones.

When the forcing variables are not I(1) the distinction between the two types of restrictions ceases to be valid. In these cases, although (1) would still be a restriction, it would be very hard to test it, as any variables left out are of the

same order of integration, zero, and one would be faced with the prospect of doing a regression in the presence of specification error. Hence, in these cases, it is logical to combine the two directly, substituting (1) into (2) to get

$$y_t = AFz_{t-1} + AGe_t. \tag{10}$$

Viewed as a VAR, now in the I(0) variables $\bar{w}_t' = (y_t' \; z_t')$, one finds a similar set of restrictions to the non–cointegrating set found above. Specifically, y_{t-1} does not appear in the VAR and the coefficients of z_{t-1} should be AF.

Basically, the argument for testing the restrictions upon the VAR advanced above is that it may be possible to identify suitable re–specifications of the RBC model in the event that rejections of the restrictions are encountered. For example, if the prediction that Δy_{t-1} is excluded from the VAR is false, attention is immediately directed to how the RBC model might be modified so as to induce such a variable into the implied VAR. The VAR is therefore being used as a "reduced form" and, indeed, the evaluation strategy being followed here is the modern equivalent of the classical precepts laid down by the Cowles Commission researchers when testing the structural equation restrictions upon the reduced form — see Byron (1974). All that has changed is the substitution of the reduced form by its time series construct, the VAR. This idea has been mentioned or exploited by a number of authors *e.g.* Spanos (1986), Monfort and Rabemananjara (1990) and Hendry and Mizon (1990); the latter being the most complete treatment in that it allows for variables to be either I(1) or I(0).

Although the systems approach to testing set out above is an attractive one, there may be advantages to focussing upon more restricted implications of the RBC model. One of these is the nature of the final equations for y_t *i.e.*if y_t is a scalar, finding the ARMA process

$$C(L)y_t = D(L)\epsilon_t \tag{11}$$

implied by (1) and (2). Comparing this derived equation to the ARMA models estimated from the data may be used to indicate how good a representation the RBC model is. Tinbergen (1939) was an early user of the final equations for summarizing the properties of a system, and the idea was subsequently formalized and utilized in Zellner and Palm (1974), Wallis (1977) and Evans (1988). Cogley and Nason (1992) apply the idea to a variety of RBC models, showing that, with the exception of the Burnside et al. model, such models do not reproduce the higher order autocorrelation features of GDP data for the U.S. Obviously, such a comparison may be extremely valuable in revealing how well the system mimics the data on selected variables. Its principal disadvantage is that the information gleaned from such a comparison may be extremely difficult to use in re–specifying any RBC model, simply because $C(L)$ and $D(L)$ will inevitably be complex functions of all parts of the original model.

A related procedure, after making y_t a vector, is to determine the VAR in y_t alone i.e. to reduce the VAR in y_t, z_t to one in terms of y_t alone. Such a construct may be of interest because of our familiarity with many bivariate and trivariate relations. For example, if y_t is composed of net investment and output, the accelerator mechanism is a well known bivariate relation linking those two variables, and it might therefore be profitable to enquire into whether there is an accelerator mechanism at work within RBC models. To perform this task requires a number of steps. First, after computing the autocovariances of y_t from (1) and (2), an approximating VAR can be fitted by solving the multivariate version of the Yule–Walker equations. This would lead to

$$C(L)y_t = \epsilon_t. \tag{12}$$

Second, suppose that y_t was bivariate with elements y_{1t} and y_{2t}. To investigate relations like the accelerator necessitates relating y_{1t} to y_{2t} as well as their past histories. The error term in ϵ_t has to be decomposed to isolate the contemporaneous effect. To this end let the VAR in (11) be re–expressed as

$$y_{1t} = C_1(L)y_{t-1} + \epsilon_{1t} \tag{13}$$

$$y_{2t} = C_2(L)y_{t-1} + \epsilon_{2t}. \tag{14}$$

Owing to the linear structure ϵ_{1t} can be written as $\epsilon_{1t} = \rho\epsilon_{2t} + \eta_{1t}$, where $\rho = \sigma_{22}^{-1}\sigma_{12}$, $\sigma_{ij} = E(\epsilon_{it}\epsilon_{jt})$, and η_{1t} is an innovation with respect to $\{y_{t-j}, y_{2t}\}_{j=1}^{\infty}$. Consequently,

$$y_{1t} = [C_1(L) - \rho C_2(L)]y_{t-1} + \rho y_{2t} + \eta_{1t} \tag{15}$$

gives the desired relationship. In (13) the polynomials $C_j(L)$ and the correlation coefficient ρ are by–products from fitting the VAR (11) to the autocovariances of y_t coming from (1) and (2). Operationally, one simply has to decide upon the order of the approximating VAR. Of course, the relation under study might also be a trivariate one *e.g.* if y_t contains real money, interest rates and output, a "money demand function" could be elicited. Perhaps the main use of this device is when comparisons are made between RBC and alternative business cycle models such as multiplier–accelerator, as conversion of the RBC model to resemble the alternative model allows an easier assessment of the relative performance of the two contenders. Another use is if one wants to compute quantities such as the Kullback–Liebler Information Criterion (KLIC) in order to compare models. Because there are fewer shocks than variables in most RBC models, the density for z_t would be singular, and hence the KLIC is not defined. However, by restricting attention to a VAR system whose order equals the number of shocks one can define the KLIC for such a system.

Although what should be tested when evaluating RBC models seems to be fairly clear, exactly how it is to be done is much more controversial. The source of the controversy resides in the fact that the variables y_t and z_t in the model may not

be accurately measured by data *i.e.* there are errors in variables. When testing the co–integrating restrictions such a difficulty can be ignored, provided that the errors are I(0), but the same cannot be said for tests of the non–cointegrating restrictions. Here what is being tested is whether the coefficients of $v_{t-1} = \beta' z_{t-1}$ in (9) have the values predicted by the RBC model. But if the errors in z_t and y_t are linearly related to z_{t-1}, the observed value could validly deviate from that predicted by the RBC model. Without some statement about the mapping of the errors into z_{t-1}, it would therefore be impossible to follow the testing strategy outlined above. Within the literature on calibrated models, this point appears to be regarded as the critical one that prohibits formal econometric testing — Kydland and Prescott (1991) and Watson (1990). There is little that can be said about this objection. It could be applied to any model and, taken to its extreme, would result in nothing being testable. If it is adopted the only consistent attitude would seem to be one in which all quantitative modelling was eschewed. However, such consistency is a rare phenomenon; it is not uncommon to find proponents of RBC models rejecting competitive scenarios as incompatible with the data but failing to apply the same test to their preferred approach on the grounds that the models are too idealized. For example, Kydland and Prescott (1991) regurgitate the Lucas–Sargent criticism that large scale Keynesian models of the 1970s were inadequate due to a failure to correctly predict the observed unemployment–inflation correlations of that decade, but immediately exempt RBC models from a similar test by stating that "the issue of how confident we are in the econometric answer is a subtle one which cannot be resolved by computing some measure of how well the model economy mimics historical data. The degree of confidence in the answer depends on the confidence that is placed in the economic theory being used". (1991, *p.* 171).

The only way out of this morass is to place some constraint upon the relationship of any errors in variables to z_{t-1}. Traditionally, this has been to insist upon the errors being white noise. Such errors in y_t would result in a white noise disturbance for (9), whereas a similar assumption for errors in z_t would create an MA(1) disturbance. In the first instance estimation and testing would proceed in the normal way; in the second some form of instrumental variables estimation would need to be performed to allow for the correlation between z_{t-1} and the MA(1) disturbance. Of course, the disturbance in (9) could be uncorrelated with z_{t-1} under weaker conditions than white noise in the errors in variables. The situation is reminiscent of rational expectations modelling where forward looking behavior creates disturbance terms that are MA's but which are still orthogonal to any regressors that appear in agents' information sets. If this extension is envisaged allowance needs to be made for the effects of such serial correlation upon inferences by adopting robust measures of the variances of estimators.

If the errors in variables are to be allowed to be functions of z_{t-1} it may still be

possible to find some measures of fit of the model to data, even though inference is highly unlikely. This is Watson's (1990) approach. He takes the deviation between model output and data to be an "error", u_t, and then finds an expression for it when the objective is to reproduce the autocovariance function (a.c.f.) of the data. Thus, distinguishing data by means of an asterisk, $y_t^* = y_t + u_t = AFz_{t-1} + AGe_t + u_t$, and the task is to determine u_t. For convenience in exposition it will be assumed that z_t is perfectly measured and that y_t is a scalar. Approximating the observed a.c.f. of y_t^* with a VAR in y_t^* and z_t gives $y_t^* = C_1(L)y_{t-1}^* + C_2(L)z_{t-1} + \epsilon_t$. By equating the two expressions for y_t^*, u_t is found to be $u_t = C_1(L)y_{t-1}^* + (C_2(L) - AF)z_{t-1} - AGe_t + \epsilon_t$, and this choice of u_t means that the augmented model output reproduces the a.c.f. of the data (at least up to the chosen order of VAR). Watson's proposal is then to compute an "R^2", equal to $1 - (\text{var}(u_t)/\text{var}(y_t^*))$, as a measure of fit of the model. As it stands this latter measure is indeterminate as the var(u_t) depends upon an unknown, the covariance of ϵ_t with u_{t-j}. Because this is a free parameter, Watson proposes to choose it such that var(u_t) is minimized. To see how this is done take $C_1(L) = c_1 > 0$. Then the smallest value of σ_u^2 occurs when $cov(\epsilon_t u_{t-1})$ attains its largest negative value $-\sigma_\epsilon \sigma_u$(this corresponding to a correlation between the two variables y_t^* and $y_t$of -1). A low "R^2" would presumably be taken as indicating that there is much left unexplained by the RBC model. In practice there are significant complications coming from the fact that y_t will generally be a vector, as the variance of u_t will become a matrix and there is no longer a unique measure of fit.

Watson's idea is certainly ingenious and, given the concern expressed about the idealized nature of these models, has to be useful information for anyone wishing to assess them. However, one cannot escape the feeling that the criterion has to be augmented with supplementary information. One problem that arises is the decision to take the minimum value of σ_u^2 as the basis of the "R^2". This is arbitrary, as many values of cov($\epsilon_t u_{t-1}$) would reproduce the a.c.f.. of y_t^*, and it is unclear why the one minimizing σ_u^2 is to be preferred. Obviously a model with a low R^2 would not be satisfactory, but it is conceivable that a high R^2 could be produced solely due to the particular selection made for cov($\epsilon_t u_{t-1}$), while other choices of this parameter may produce low R^2. Since the parameter, $cov(\epsilon_t u_{t-1})$, has nothing to do with the model, and is essentially "unidentified", it would seem misleading to conclude from the evidence of a high R^2 that the RBC model was satisfactory. At the very least it would seem important that the R^2 be provided for the values of cov($\epsilon_t u_{t-1}$) that both maximize and minimize R^2. If this range is narrow, and the minimum R^2 is a high one, it might be appropriate to conclude that, **prima facie**, the RBC model provides a satisfactory description of the data.

A second problem with the measure is that it does not provide information that may be useful in re–specifying the model. The variance of u_t may be large for a variety of reasons; — a high σ_ϵ^2, a large gap between $C_2(L)$ and A, a large value

for $C_1(L)$ etc., but this information is lost in the aggregative measure. However, our attitude towards the model is likely to be significantly affected by which one of these is the principal contributor. If it was due to a high value of $C_1(L)$, we would be led to enquire into whether the RBC model might be re–specified so as to induce the variable y_{t-1} into the VAR. In contrast, if it was a consequence of a large value for σ_ϵ^2, we are less likely to feel that there is something inadequate in the idealized model, as this parameter represents the extent to which variables exogeneous to the model are unpredictable, and all models would have a similar deficiency *e.g.* a Keynesian model also has to make some assumption about how government expenditure is to evolve over time.

3. Evaluating an RBC Model

The model chosen for the evaluation exercise is due to Burnside, Eichenbaum and Rebelo (1990) (BER). It represents a modification of that described in Christiano and Eichenbaum (1992). Appendix 1 presents the principal equations underlying it. The controlled state variables are the capital stock and employment and the uncontrolled states are the technology and government expenditure shocks. When measured as deviations from a steady state growth path these variables are designated as k_t, n_t, a_t, and g_t respectively. Other variables explained by the model, also as deviations from steady state, are output (y_t), private consumption (c_t), and investment (i_t). An assumption of the model is that the forcing factors are AR(1) processes. Parameter values for the model were estimated by BER from data over 1955/3 to 1984/1 using various moment conditions.

To evaluate the model BER compared the numerical values of selected variable correlations predicted by the model with the estimated values from the data. The vector of discrepencies can be formally compared with zero using the J–test of over–identifying restrictions. The principal comparison BER made involved the cross correlation of productivity and hours worked at L leads and lags. When all the sample was used there was strong rejection if $L = 2$ (the p value of the test being .001). This outcome encouraged them to split the sample at 1969/1 and to perform validation of the model on two different samples. They then concluded that the model seemed satisfactory for the first period (p-value= .278) but not for the second period (p–value= .001). Because of this diversity of outcomes the discussion below concentrates upon the two sub–samples separately. We also avoid the emphasis upon the relation between productivity and hours that characterizes BER's evaluation work, as an important ingredient of the way in which their model manages to emulate the data is by making the assumption that the employment data is subject to errors of measurement. That modification seems to be very important to their success along the productivity/ hours correlation dimension, even though it

is hard to think of it as part of a "model".

a) Sample Period 1955/3 to 1969/4

As reviewed in the preceeding section any RBC model makes a number of predictions, either about the co–integrating vectors expected to hold between variables or the dynamic behaviour of the variables. Our strategy will be to determine if the predictions made by the RBC model are consistent with the data.

A first item to check is whether the assumption made pertaining to the evolution of the uncontrolled states is valid. BER's point estimates for the AR(1) coefficients of a_t and g_t are .87 and .94 respectively. Although these are different from unity the ADF tests recorded in Table 1 point to the fact that the hypothesis of the series being integrated is accepted fairly easily. Furthermore, the correlation between the residuals from the AR(1)'s fitted to g_t and a_t is only .12, which suggests the processes are uncorrelated, as specified in BER's model. Based on this outcome, and the evidence of integration for k_t and n_t in Table 1, it is anticipated that the state vector comprising k_t, n_t, g_t and a_t should have two co–integrating vectors, as there are two common trends driving the RBC model (see the brief description of the main features of the model in the Appendix).

Using the parameter values provided by BER it is possible to compute $F = \begin{bmatrix} \gamma & \delta \\ 0 & I_p \end{bmatrix}$ in (2) and hence to derive the predicted co–integrating vectors among the four states viz. $[(I-\gamma) \ -\delta]$. Logically, there are two distinct questions here. One is whether there are two co–integrating vectors or not. Using a VAR(4), Johansen's likelihood ratio test (LR) for the hypothesis of r co–integrating vectors easily indicates that there are two (the test of $r = 1$ versus $r = 2$ gives $LR = 25.2$ while $r = 2$ versus $r = 3$ has $LR = 7.75$, where the critical values corresponding to the 5% significance levels are 21.0 and 14.0 respectively). Exactly the same conclusion is reached with Johansen's trace test. Thus the number of co–integrating vectors agrees with the model prediction. A more demanding test is to assess whether the predicted numeric values (.0435 -.0295 -.1434 .0062) and (.5627 1.0174 -1.5008 -.1974) are compatible with the data. For this query, a likelihood ratio test of the restrictions gives a value of 45.53 which, when referred to a $\chi^2(4)$, soundly rejects the constraint. Consequently, a basic property of the model is rejected. Figure 1 plots $.5627k_t + 1.0174n_t - 1.5008a_t - .1974g_t$, the projected second co–integrating error, and the lack of co–integration shows quite clearly (actually ADF tests applied to each co–integrating error separately shows that neither series is I(0)).

Table 1 Tests for Integration in Data		
Variable	**ADF(4) with trend**	**ADF(4) without trend**
k_t	-1.43	-1.80
n_t	-1.25	-2.97
a_t	-2.44	-2.42
g_t	-2.13	-2.12
c_t	-2.19	-2.18
y_t	-2.21	-2.31
i_t	-2.37	-2.88
crit. val.	-2.92	-3.50

In addition to the state variables being I(1), Table 1 shows that three "output" variables — consumption, output and investment — also possess this property. Therefore, RBC models conjecture that there are further co–integrating restrictions, now between the "outputs" and the states — see (1). King et al. (1992) and Neusser (1991) considered the long run implications of neoclassical growth models for relations **between** the three "output" variables above. In particular they argued that consumption and income and investment and income should be co–integrating pairs with co–integrating vectors (1 -1). In this model it is consumption, income and government expenditure which should be co–integrated, as well as investment and output. Johansen's tests indicate that the first of these relations is satisfied, but the likelihood and trace tests are in conflict over whether investment and output are co–integrated.[1] Moreover, for two reasons, tests of an RBC model performed in this way are rather weak. First, information is being discarded. The RBC model makes a direct prediction about the co–integrating relations between states and "outputs" but only an indirect one about the connection between "outputs". Thus we might have $c_t - z_t'\alpha_1$ and $y_t - z_t'\alpha_1$ both being I(0), so that $c_t - y_t$ is I(0), but the α_1 may not coincide with that indicated by the RBC model. Second, many models have the property that $c_t - y_t$ and $i_t - y_t$ are co–integrated, *e.g.* the multiplier accelerator model can be designed to produce this effect by an appropriate choice of ECM format, and therefore co–integration between "output" variables cannot be taken as validating the RBC viewpoint. In summary, what should be tested are the **direct** implications of the RBC model and not the **indirect** ones.

Choosing BER's parameter values, the specific co–integrating relations from equation (1) are

$$c_t = .55k_t + .017n_t + .348a_t - .07g_t + \eta_{ct} \tag{16}$$

$$y_t = .13k_t + .31n_t + 1.64a_t + .07g_t + \eta_{yt} \tag{17}$$

[1]Robert King has suggested that this failure may well be a consequence of the way in which data is constructed by BER.

$$i_t = -.65k_t + 1.12n_t + 5.45a_t - .24g_t + \eta_{it}, \tag{18}$$

and our objective is to test if the $\eta_{.t}$ are I(1). This could be done in one of two ways.[2] A first possibility is to apply an ADF test to the errors from (16), (17) and (18); since no parameters are being estimated many of the problems of using this test for co–integration are absent. An alternative is to use the fact that co–integration between variables means the existence of an ECM relationship — Engle and Granger (1985). Therefore, assuming (say) c_t and z_t are I(1) and co–integrated with vector $(1 - \alpha_1)$, an ECM of the form

$$\Delta c_t = a\Delta z_t'\alpha_1 + b(c_{t-1} - z_{t-1}'\alpha_1) \tag{19}$$

would connect c_t and z_t. If c_t and z_t are not co–integrated, $b = 0$ making the t–ratio for $H_0 : b = 0$ a suitable test of no co–integration. This test is proposed in Banerjee et al. (1986) and has been dubbed the "ECM test" by Kremers et al. (1992). The latter have argued that it has much better power than the ADF test whenever the latter imposes an invalid common factor restriction. Unfortunately, the distribution of the ECM test varies between the Dickey–Fuller density and the standard normal as $\text{var}[(a-1)\Delta z_t]/\text{var}(\Delta z_t)$ tends from zero to infinity. Because Δz_t in our situation is a vector it is difficult to determine exactly what the critical values are. One plan of action would be to be conservative and to adopt the DF critical values. Note that there are no tests of (17). The reason is that the unobserved variable a_t is effectively computed from data on y_t, k_t etc by inverting (17), and therefore η_{yt} is identically zero. Unless a separate estimate of a_t can be made it is therefore impossible to test this co–integrating restriction in an RBC model. [3]

Table 2
Tests of Co–integrating Relations in (16) and (18)

Variable	ADF(4) with trend	ADF(4) without trend
c_t	-1.71	-2.69
i_t	-2.66	-2.26
crit. val.	-2.92	-3.50

The evidence from Table 2 is that the co–integrating restrictions are most likely invalid. The problematic outcome is for consumption. Referred to an $N(0,1)$ random variable one would opt for co–integration, but this would not be true if

[2] A third method would be to employ Johansen's test, but the fact that the states do not have the co–integrating relations implied by the RBC model makes it more convenient to perform "single equation" tests.

[3] Provided a unit root is specified for the a_t process it would be possible to generate data on a_t using a random number genator and thereupon one could test (17). Smith (1990) advocates this approach when there is a latent variable.

the comparison was made with a 5% critical value from the DF density (-2.91). Nevertheless, Table 2 does hint at specification difficulties with the BER model. To see why the restrictions are being rejected it is useful to fit relations such as (16), (17) and (18) using the data to give

$$c_t = .89k_t + .12n_t + .56a_t - .03g_t \tag{20}$$

$$i_t = -.18k_t + .69n_t + 3.61a_t - .31g_t. \tag{21}$$

Comparing (16) and (18) with (20) and (21) it seems as if the weight given to k_t in the model is too low for both variables, whereas the influence of a_t is too low for consumption but far too high for investment. As the R^2 from the regressions in (20) and (21) are .94 and .93 respectively, provided the series are I(1) there is likely to be only small bias in the estimated co–integrating vectors.[4]

Although it seems unlikely, let us suppose that the co–integrating restrictions are satisfied. Then the third set of restrictions imposed by an RBC model are those relating to dynamics- equation (8). These involve testing if the coefficients of the co–integrating errors v_{t-1} are $A\alpha$ in the regression of Δy_t on v_{t-1}. A simple way to compute the statistic for such a test is to regress Δy_t on z_{t-1} and test if the coefficients are equal to Π(eq (6)). One has to be careful to refer the resulting test statistic to a $\chi^2(2)$ since the distribution of $\hat{\Pi}$ is singular owing to the co–integration, *i.e.* as v_{t-1} is a 2x1 vector, only two coefficients are really being tested. With Δy_t set to Δc_t and Δi_t, the test statistics are 3.5 and 98.3 respectively, showing that, although the dynamics of consumption seem to be accounted for, the investment dynamics are missed badly (there is some serial correlation in the regression for Δi_t but there is only a minor change in the value of the test statistic when computed robustly). Unlike the situation for co–integration tests, it is also possible to check the output dynamics, and the test statistic there is 28.58, again showing some problems with the model. Equations (22), (23) and (24) list the predicted dynamic relations along with the estimated relations (in brackets) for each of the series

$$\begin{aligned} \Delta c_t = &-.033k_{t-1} - .001n_{t-1} + .059a_{t-1} + .005g_{t-1} \\ &(-.100) \quad (-.038) \quad (.086) \quad (-.017) \end{aligned} \tag{22}$$

$$\begin{aligned} \Delta y_t = &-.18k_{t-1} - .309n_{t-1} - .266a_{t-1} + .055g_{t-1} \\ &(-.19) \quad (-.086) \quad (-.075) \quad (-.016) \end{aligned} \tag{23}$$

$$\begin{aligned} \Delta i_t = &-.603k_{t-1} - 1.16n_{t-1} + .877a_{t-1} + .24g_{t-1} \\ &(-.597) \quad (-.24) \quad (-.12) \quad (.03) \end{aligned} \tag{24}$$

[4] A more serious problem is that the parameters being estimated may not be identified. If there are only two stochastic trends then it is impossible to estimate the four parameters here as the number of identified parameters can be no larger than the number of trends.

As revealed by (22), (23) and (24) the major problem with the RBC model in its forecasts of dynamics is that it ascribes far too much weight to the productivity shock and lagged employment.

It is now appropriate to consider some objections that might be made to the above analysis. One of these is that the restrictions being tested are found by using the parameter values in BER and these are $\rho_a = .8691$ and $\rho_g = .938$ rather than the values of unity needed if we are to argue that the series are I(1). For this reason it is logically more correct to re–compute what the implied restrictions would be if unit roots are imposed upon the two forcing processes and to then test if the resulting restrictions are compatible with the data. This means that (16), (17) and (18) become

$$c_t = .55k_t + .017n_t + .888a_t - .14g_t + \eta_{ct} \tag{25}$$

$$y_t = .13k_t + .31n_t + 1.11a_t + .13g_t + \eta_{yt} \tag{26}$$

$$i_t = -.65k_t + 1.12n_t + 2.31a_t + .13g_t + \eta_{it}. \tag{27}$$

Doing so does not change any of the conclusions reached previously however. For example, the ADF tests for cointegration among the states now become -2.89 and -3.32 (with a $\chi^2(4) = 44.79$ when testing using Johansen's estimator), while ADF test values of -1.19 (c_t) and -2.64 (i_t) are found when directly testing the restrictions in (25) and (27). Tests of the dynamic restrictions yield $\chi^2(2)$ test statistics of 3.0 (c_t), 35.2 (y_t) and 118.4 (i_t).

Another objection to the analysis could be that the series are not integrated and that the power of the ADF test is low. There is some merit to this argument. If $\rho_a = .8691$, simulation of the ADF(4) test (with trend) for 58 observations shows that 55% of the time one gets an ADF test larger than -2.44 (the value of the ADF tests using the data on a_t). Hence one would falsely conclude that the series is I(1) 55% of the time. In the same vein, with $\rho_g = .93$, one would invalidly conclude there was a unit root 45% of the time (using the ADF value of -2.13 found from the data). Hence it may be more reasonable to conduct tests that assume the processes are I(0) rather than I(1). In this case we will test the restrictions from (9), *i.e.* that the coefficients of z_{t-1} are AF. Equations (28), (29), and (30) set out the theoretical coefficients for the "reduced" VAR.

$$c_t = .518k_{t-1} + .016n_{t-1} + .407a_{t-1} - .069g_{t-1} \tag{28}$$

$$y_t = -.045k_{t-1} - .001n_{t-1} + 1.907a_{t-1} + .128g_{t-1} \tag{29}$$

$$i_t = -1.255k_{t-1} - .039n_{t-1} + 6.322a_{t-1} + .003g_{t-1} \tag{30}$$

Corresponding empirical estimates (where we have added in missing terms from the VAR in (28), (29), and (30) if the t ratio was greater than 2) are

$$\underset{(2.5)}{c_t = .403k_{t-1}} + \underset{(.57)}{.026n_{t-1}} + \underset{(3.49)}{.411a_{t-1}} - \underset{(-1.8)}{.033g_{t-1}} + \underset{(2.6)}{.433c_{t-1}} \tag{31}$$

$$\underset{(-.43)}{y_t = -.057k_{t-1}} + \underset{(2.67)}{.222n_{t-1}} + \underset{(11.11)}{1.566a_{t-1}} + \underset{(1.57)}{.056g_{t-1}} \tag{32}$$

$$\underset{(-2.87)}{i_t = -.673k_{t-1}} + \underset{(.26)}{.047n_{t-1}} + \underset{(2.20)}{1.401a_{t-1}} - \underset{(-1.32)}{.106g_{t-1}} + \underset{(3.57)}{.579i_{t-1}} \tag{33}$$

The results in (31), (32) and (33) constitute a strong rejection of the restrictions implied by the RBC model. Testing that the parameters in (31), (32) and (33) equal those in (28), (29) and (30) gives χ^2 statistics of $\chi^2(5) = 104.1(c_t)$, $\chi^2(4) = 28.59(y_t)$, and $\chi^2(5) = 625.0(i_t)$. A comparison of the two sets of equations shows there are some variables missing from the former $-c_{t-1}$ in the c_t equation, and i_{t-1} in the i_t equation — and that the model accords productivity too great an influence in determining investment and output. Others have remarked upon such a "missing variable" feature, specifically for consumption, (Chow (1992b)), but a casual comparison of the equations emphasises that there are many factors responsible for the failure of the model to explain output and investment variations.

The outcomes observed above bring to the fore a question raised in the introduction; is the rejection being caused by the model or by the parameter values being supplied to it? That is, does there exist an RBC model of this form that would be compatible with the data but which had a different set of parameter values? It might be argued that the essence of the model is the type of functional forms fed in and not the values of the parameters chosen to calibrate it. Earlier we remarked why we feel that this view should be rejected, but it is worth exploring what would happen if we adopted it. One can say immediately that the non–zero coefficients seen for c_{t-1} and i_{t-1} in (31) and (33) cannot be matched by calibration changes, as the model design automatically assigns a zero coefficient to these variables. Only re–specification of the RBC model would change this fact. Some of the other parameters in (31), (32) and (33) can be modified by changing the calibration settings. By studying the sensitivity of (31), (32) and (33) to variations in the parameters of BER's model, it was found that we could improve the approximation by increasing ρ_a and reducing α. However, it was necessary to make ρ_a almost unity if the weight on the productivity variable was to be reduced to the required magnitude. This would mean that we are dealing with processes that are very close to being integrated and so it would be appropriate to test the co–integrating restrictions. As mentioned earlier however, these are rejected when we impose I(1) behaviour upon the forcing variables. Hence, it does not seem as if the essentials of the economy are captured by the BER formulation. It is worth emphasising here that the rejections

of the RBC model using the techniques above are far stronger than those encountered by BER, where what evidence there was against their model in this period was very mild. This fact emphasises that different types of information are being gathered by the different methods of evaluation.

(b) Sample period 1970/1 to 1984/1

In the second period there is evidence that the evolutionary pattern for the variables identified in the first period has changed. Looking first at the forcing processes, there is some doubt that they are now I(1). The ADF tests (with trend) are -3.40 and -2.85 for g_t and a_t respectively, while the ADF (without trend) for a_t of -2.86 is very close to the 5% critical value of -2.92. Examination of the estimates of the autoregressive parameters upon which the ADF test is based reveal them to be .47 (g_t) and .76 (a_t), below the values of .87 and .81 found in the first period. It seems very likely therefore that the processes are I(0); certainly one would only be comfortable with a single common trend, due to a_t, as the autoregressive parameter for g_t is far too low. Turning to the other series, here the evidence of I(1) behaviour is stronger, but even then the autoregressive parameter is (at best) just above .8.

What is to be done about these features? One possibility is to proceed with the tests outlined in the previous sub–section, maintaining that there is a single common trend. When this is done one encounters rejections of all the co–integrating restrictions. In the interests of economizing on space, and recognizing the doubt raised over the integration properties of the data, our preference has been to only report results derived under the assumption that the series are all I(0). This means that we perform and report the tests of the dynamic restrictions appearing in (9).

Equations (34), (35) and (36) provide the estimated equations, along with the predicted values of the coefficients in brackets (variables not entering the model VAR have been deleted if their t–ratio is less than 2, while estimated intercept terms have also been suppressed).

$$\begin{array}{l} c_t = -.112k_{t-1} - .071n_{t-1} - .109a_{t-1} - .001g_{t-1} + 1.153c_{t-1} \\ \quad\quad (-.535) \quad (.014) \quad\quad (.396) \quad\quad (-.021) \quad\quad (0) \end{array} \tag{34}$$

$$\begin{array}{l} y_t = -.685k_{t-1} - .153n_{t-1} - .113a_{t-1} - .151g_{t-1} + 1.053c_{t-1} + .297i_{t-1} \\ \quad\quad (-.034) \quad (-.001) \quad (1.913) \quad (.027) \quad\quad (0) \quad\quad\quad (0) \end{array} \tag{35}$$

$$\begin{array}{l} i_t = -1.371k_{t-1} - .537n_{t-1} - 1.340a_{t-1} - .342g_{t-1} + 1.970c_{t-1} + 1.033i_{t-1} \\ \quad\quad (-1.246) \quad (-.034) \quad (6.363) \quad (-.294) \quad\quad (0) \quad\quad\quad (0) \end{array} \tag{36}$$

The task is to determine whether the predicted and estimated parameters are significantly different from each other, and the resulting test statistics are $\chi^2(5) =$

225.2, $\chi^2(6) = 91.7$ and $\chi^2(6) = 638.6$ for c_t, y_t and i_t respectively. If only the coefficients of k_{t-1}, n_{t-1}, a_{t-1} and g_{t-1} are tested for having their predicted values, the corresponding $\chi^2(4)$ statistics would be 110.3, 37.18 and 151.1. As before, this constitutes a very strong rejection of the model, although an important difference from the previous period is that the prediction of zero coefficients for c_{t-1} and i_{t-1} in the equations is now wildly at variance with the data, indicating that the dynamic structure of the model seems to have undergone some major shifts in the period. Looking at the estimates in (34)-(36), the most striking feature is the fact that the technology shock a_t is estimated to have a **negative** impact on all variables in this period, which is in sharp contrast to the positive effect predicted by the model.

4. Comparing Models

As mentioned in the introduction it is perhaps more reasonable to evaluate a model by its performance relative to others than to impose an absolute standard. For this reason it was decided to effect a comparison of the RBC model with a stylized version of the type of macro model that was popular in the 1960s. This generally featured a consumption relation dynamically connecting consumption and output, as well as an accelerator mechanism for investment. Although money featured in such models as well, here it is excluded in the interest of retaining comparability with the RBC model; the idea being to work with the same variables as BER did, but to provide a "demand" rather than "supply" side account of developments in the U.S. economy.

Most of the models of the 1960s worked with levels of the variables and we therefore chose to do the same thing here. To make comparisons with the RBC model, the predictions of the latter had to be converted from deviations around steady state values back to levels. Levels of variables are distinguished by capital letters. The multiplier–accelerator model (MPA) that was fitted is given in the equations below. No experimentation with lag lengths etc. was undertaken; the idea was just to take a simple model and to see how well it performs on the same data set. Some of the regressors in the equations were insignificant but were nevertheless retained.

$$C_t = \alpha_1 C_{t-1} + \beta_1 Y_t + \beta_2 Y_{t-1} + c_1 + \phi_1 t \tag{37}$$

$$NI_t = \alpha_2 NI_{t-1} + \gamma_1 \Delta Y_t + \gamma_2 \Delta Y_{t-1} + c_2 + \phi_2 t \tag{38}$$

$$NI_t = I_t - \delta K_{t-1} \tag{39}$$

$$K_t = (1-\delta) K_{t-1} + I_t \tag{40}$$

$$G_t = \rho_g G_{t-1} + c_3 + \phi_3 t \tag{41}$$

$$Y_t = C_t + I_t + G_t + D_t \tag{42}$$

The variable D_t is needed to make the series on output satisfy the national income identity. It is always a small fraction of output Y_t and rarely reaches 1% of that variable, so that its introduction would not seem to produce any distorting factors. Table 3 gives the parameter estimates of the unknown parameters of the multiplier–accelerator model for each of the two periods. Estimation was done by OLS, as that was also the most common way of doing "calibration" at that time.

Table 3
Estimates for Multiplier–Accelerator Model (21)

Parameter	First period	Second Period
α_1	.695	.848
β_1	.135	.128
β_2	-.040	-.106
c_1	352.3	362.7
ϕ_1	2.32	1.279
α_2	.873	.942
γ_1	.267	.430
γ_2	.180	.065
c_2	18.65	25.86
ϕ_2	.523	-.214
δ	.020	.022
ρ_g	.934	.560
c_3	62.32	583.1
ϕ_3	.613	-.469

It is interesting to first ask whether the MPA model makes correct predictions of the VAR coefficients. As can be seen, the MPA model implies that the data should be a VAR(2), and the coefficients of each lag can be worked out by solving (37)-(42). The $\chi^2(10)$ statistics testing the adequacy of the model during the first period were 21.4, 23.4 and 42.3, for C_t, I_t and Y_t respectively,. The corresponding test statistics in the second period were 73.0, 112.0 and 162.7. Although the fact that we are working with levels, and hence potentially integrated data, makes the actual distribution of these "χ^2" statistics unlikely to be exactly that, their magnitude has to make one seriously question the MPA model as a good representation of the data. This conclusion is especially true of the second period, a feature that is consistent with the notion that "Keynesian" models broke down in the 1970s. If one takes the size of the χ^2 statistics as an index of how good the model is, then both the RBC and MPA models have noticably worse performance in the second period.

Figs 2–7 provide plots of the one step predictions of C_t, I_t and Y_t from both models for each of the time periods. The overall impression is that the MPA model

is more successful than the RBC model in tracking all series. Because the models are non–nested, imposing different restrictions upon the same VAR, one way to check the above impression is to enquire into whether the explanation of variables of interest given by the MPA model can be improved upon by using information from the RBC model. To this end we regress data on the variable being studied against the predictions of it made by both the MPA and RBC models; if the RBC model is correct then the coefficient on the predictions from the MPA model should be zero, and conversely. This test is in the spirit of Davidson and MacKinnon's (1981) J–test for non–nested models. Selecting C_t, I_t, and Y_t as the variables of interest the results are given below in (43), (44) and (45) for the first period and in (46), (47) and (48) for the second period, with t–ratios in brackets.

$$\underset{(2.15)}{C_t = .21C\hat{RBC}_t} + \underset{(8.08)}{.79C\hat{MPA}_t} \tag{43}$$

$$\underset{(1.74)}{I_t = .09I\hat{RBC}_t} + \underset{(15.94)}{.90\hat{I}MPA_t} \tag{44}$$

$$\underset{(2.41)}{Y_t = .18Y\hat{RBC}_t} + \underset{(11.94)}{.83\hat{Y}MPA_t} \tag{45}$$

$$\underset{(1.63)}{C_t = .14C\hat{RBC}_t} + \underset{(11.57)}{.88\hat{C}MPA_t} \tag{46}$$

$$\underset{(3.45)}{I_t = .15I\hat{RBC}_t} + \underset{(14.47)}{.82\hat{I}MPA_t} \tag{47}$$

$$\underset{(7.57)}{Y_t = .59Y\hat{RBC}_t} + \underset{(8.34)}{.52Y\hat{M}PA_t} \tag{48}$$

The evidence in the above equations is that the RBC model rarely adds a great deal to the explanatory power of the MPA model. Perhaps the most striking exception to this statement is in (48); it would seem that output in the second period cannot be satisfactorily explained by a pure demand side model like MPA. Given the oil–price shocks of the 1970s, part of which would be reflected in the ex–post measurements of productivity, such a conclusion may not be too surprising.

One way to understand the difference in the two models is to ask what the RBC model would look like if turned into an MPA type model. To do this we use the ideas in section 2 for reducing the VAR implied by the RBC model into bivariate VARs between the pairs (consumption, output) and (investment, output). Equations (49)-(52) give the implied bivariate VAR's for the two periods, with the estimated parameters in brackets.

First Period

$$c_t = \underset{(.79)}{1.08c_{t-1}} - \underset{(.13)}{.09c_{t-2}} + \underset{(.24)}{.18y_t} - \underset{(-.12)}{.19y_{t-1}} + \underset{(-.08)}{.07y_{t-2}} \tag{49}$$

$$i_t = .87i_{t-1} + .04i_{t-2} + 3.17y_t - 2.73y_{t-1} - .317y_{t-2} \tag{50}$$
$$(.94) \quad (-.10) \quad (.98) \quad (-.20) \quad (-.63)$$

Second Period

$$c_t = .99c_{t-1} - .05c_{t-2} + .21y_t - .21y_{t-1} + .07y_{t-2} \tag{51}$$
$$(1.12) \quad (-.16) \quad (.20) \quad (-.25) \quad (.05)$$

$$i_t = .73i_{t-1} + .03i_{t-2} + 3.31y_t - 2.39y_{t-1} - .24y_{t-2} \tag{52}$$
$$(.41) \quad (.12) \quad (1.81) \quad (-.55) \quad (-.23)$$

These equations encapsulate most of the information in figs 2–7 regarding the behaviour of consumption and investment. A succinct summary of the latter is that investment is much too volatile whilst consumption is too smooth. In terms of (49)-(52) the impact of output upon investment is seen to be too large, while the lag distribution of consumption response to income changes is longer for the RBC model. Another interesting feature of (49)-(52) is that the accelerator mechanism is very clear in the data of the first period but is not in evidence in the second (in the sense that the coefficients of y_t, y_{t-1} and y_{t-2} do not sum to zero). This provides one explanation for the MPA model's deterioration in performance during the second period.

5. Conclusion

This paper has set out a strategy for evaluating small linear models via the restrictions they impose upon the VAR in the variables they are meant to explain. Three types of restrictions were elicited. First, there are co–integrating restrictions implied among the state variables. Second, there are the co–integrating constraints existing between the state and "output" variables. Finally, there are restrictions upon the dynamics of the model when all variables are transformed to be I(0). It was recommended that evaluation should proceed by examining the constraints sequentially. The technology was then applied to an RBC model that had performed reasonably well when assessed relative to a set of "stylized facts"; failure on all three counts was evident, pointing to the need for some re–specification of the model.

A failure of the co–integrating restrictions is generally the hardest feature to rectify as some variables need to be added to the system. Candidates could be the effects of taxation upon capital accumulation, the impact of the external sector via terms of trade movements, or monetary factors. Although a complete study of this phenomenon is beyond the paper, understanding the source of the co–integration failures seems critical to determining what course of action should be followed. One useful piece of information is to try to determine whether there was a period in

which the co–integration implications of the model were valid. To this end figs 8 and 9 plot recursive estimates of the parameters used in Table 2 when concluding that there was a lack of co–integration between "output" and state variables. Fig 8 shows a plot of the coefficient of the lagged co–integrating error for consumption (CORC) that is the basis of the ADF test; a co–integration failure is signalled if this approaches unity. Fig 9 shows the coefficient of the lagged co–integrating error in investment (CORI) used in the "ECM" test for co–integration; values of this near zero suggest that co–integration does not hold. Both figures tell the same story; sometime around 1964 the RBC model's co–integrating predictions started to fail quite dramatically. As this was a period of major fiscal and monetary changes associated with the "Great Society" program, it suggests that the RBC model needs to be modified to capture the impact of such stimuli more directly.

Even if the co–integrating restrictions were made acceptable there also appears to be some difficulties with the "short run" responses within the model. Results presented in sections 3 and 4 make a strong case for introducing adjustment costs into investment in order to reduce the magnitude of its short run reponse to fluctuations in output. The opposite is true of consumption, where the impact of current income needs to be strengthened.

APPENDIX 1

The Burnside/ Eichenbaum/ Rebelo Model

Worker Utility

$$\ln(C_t^p) + \theta ln(T - \xi - W_t f)$$

T =Time endowment, C_t^p =private consumption, W_t =effort, f =hours worked per shift, ξ =fixed cost of work (in terms of hours of forgone leisure)

Non–Worker Utility

$$\ln(C_t^p) + \theta ln(T)$$

Cobb–Douglas Production Function

$$Y_t = Z_t K_t^{1-\alpha}(N_t W_t f)^\alpha$$

Z_t =Technology, N_t =fraction of agents who are workers (the number of agents is normalized to unity), K_t =Beginning of Period Capital Stock

Technology Change,

$$Z_t = \gamma^{\alpha t} A_t$$

Productivity Shock

$$\ln(A_t) = (1 - \rho_a)\ln(A) + \rho_a \ln(A_{t-1}) + \epsilon_t$$

Aggregate Resource Constraint

$$C_t^p + K_{t+1} - (1 - \delta)K_t + X_t \leq Y_t$$

X_t =Government Consumption

Fiscal Rule

$$X_t = \gamma_g^t G_t$$

$$\ln(G_t) = (1 - \rho_g)\ln(G) + \rho_g \ln(G_{t-1}) + \mu_t$$

It is assumed that a social planner maximizes

$$E_0 \sum_{t=0}^{\infty} \beta^t \{\ln(C_t^p) + \theta N_t \ln(T - \xi - W_t f) + \theta(1 - N_t)\ln(T)\}$$

subject to the constraints above and K_0 by choice of contingency plans for $\{C_t^p, K_{t+1}, N_t, W_t : t \geq 0\}$. E_0 is the time 0 conditional expectations operator, β is the subjective discount rate, $0 < \beta < 1$.

Certain transformations are made to the problem before it is solved. These are to express the variables as deviations from deterministic steady state growth paths. Thus $\bar{C}_t^p = C_t^p/\gamma^t$, $\bar{Y}_t = Y_t/\gamma^t$, $\bar{K}_t = K_t/\gamma^t$, $\bar{X}_t = X_t/\gamma^t$ which means that the constraints on the optimization can be reduced to

$$\gamma\bar{K}_{t+1} = A\bar{K}_t^{1-\alpha}(N_tW_tf)^\alpha - \bar{C}_t^p + (1-\delta)\bar{K}_t - \bar{X}_t$$

Table 1: Parameter Values for the BER Model

Parameter	Period 1	Period 2
δ	.0196	.0221
θ	.6593	.6504
ρ_a	.8691	.8815
σ_ϵ	.0042	.0067
$ln(A)$	8.4914	8.8733
$ln(\gamma_y)$	.0069	.0015
$ln(G)$	6.8090	7.1618
$ln(\gamma_g)$	.0073	-.0013
ρ_g	.938	.6618
σ_μ	.0143	.0115

while the optimand becomes

$$\sum_{t=0}^{\infty}\beta^t\log(\gamma^t) + E_0\sum_{t=0}^{\infty}\beta^t\{\ln(\overline{C_t^p}) + N_t\theta\ln(T-\xi-W_tf) + \theta(1-N_t)\ln(T)$$

Finally small letters indicate deviations of variables from steady states. Thus $a_t = \log(A_t/A)$.

The solutions to this problem after linearization of the Euler equations are laws of motion for the state variables k_t, n_t, a_t and g_t as well as linear relations connecting other variables such as y_t, c_t^p to these states. With the parameter values in the Table it is possible to compute numerical values for these relations and they are presented in the text. The parameter σ_v in the table arises from the assumption that there are measurement errors in hours worked.

REFERENCES

BANERJEE, A., J.J. DOLADO, D.F. HENDRY and G.W. SMITH (1986), "Exploring Equilibrium Relationships Through Static Models: Some Monte Carlo Evidence", *Oxford, Bulletin of Economics and Statistics*, 48, 3, 253–277.

BURNSIDE, C., M. EICHENBAUM and S. REBELO (1990), "Labor Hoarding and the Business Cycle" (mimeo, Northwestern University).

BYRON, R.P. (1974), "Testing Structural Specification Using the Unrestricted Reduced Form", *Econometrica*, 42, 869–883.

CANOVA, F. (1990), "Simulating General Equilibrium Dynamic Models Using Bayesian Techniques" (mimeo, University of Rochester).

CHOW, G.C. (1992a), "Dynamic Optimization without Dynamic Programming", *Economic Modelling*, 9, 3–9.

CHOW, G.C. (1992b), "Statistical Estimation and Testing of a Real Business Cycle Model", *Econometric Research Program Research Memorandum No 365*, Princeton University.

COGLEY, T. and J.M. NASON (1992), "Do Real Business Cycle Models Pass the Nelson–Plosser Test?" (mimeo, University of British Columbia).

CRISTIANO, L.J. and M.EICHENBAUM (1992), "Current Real–Business–Cycle Theories and Aggregate Labor Market Fluctuations", *American Economic Review*, 82, 430–450.

DAVIDSON, R. and J.G. MACKINNON (1981), "Several Tests for Model Specification in the Presence of Alternative Hypotheses", *Econometrica*, 49, 781–793.

EVANS, G.W. (1989), "Output and Unemployment Dynamics in the United States: 1950–1985", *Journal of Applied Econometrics*, 4, 213–237.

GREGORY, A.W. and G.W. SMITH (1991), "Statistical Aspects of Calibration in Macroeconomics", in G.S. Maddala, C.R. Rao and H.D. Vinod (eds.), *Handbook of Statistics, vol 11* (forthcoming).

HENDRY, D.F. and G.M. MIZON (1990), "Evaluating Dynamic Models by Encompassing the VAR", *Discussion Paper No 9011*, University of Southampton.

HOOVER, K.D. (1991a), "Six Queries About Idealization in an Empirical Context", *Ponzan Studies in the Philosophy of Science and the Humanities* (forthcoming).

HOOVER, K.D. (1991b), "Calibration and the Econometrics of the Macroeconomy" (mimeo, University of California at Davis).

KING, R.G., C. PLOSSER and S. REBELO (1988), "Production, Growth and Business Cycles I: The Basic Neoclassical Model", *Journal of Monetary Economics*, 21, 195–232.

KING, R.G., C. PLOSSER, J. STOCK and M. WATSON (1991), "Stochastic Trends and Economic Fluctuations", *American Economic Review*, 81, 819–846.

KREMERS, J.J.M., N.R. ERICSSON and J.J. DOLADO (1992), "The Power of Cointegration Tests", *International Finance Discussion Paper No 431*, Board of Governors of the Federal Reserve System.

KYDLAND, F. and E. PRESCOTT (1982), "Time to Build and Aggregate Fluctuations", *Econometrica*, 50, 1345–1370.

KYDLAND, F.E. and E.C. PRESCOTT (1991), "The Econometrics of the General Equilibrium Approach to Business Cycles", *Scandanavian Journal of Economics*, 93, 161–178.

LAIDLER, D. and B. BENTLEY (1983), "A Small Macro–model of the Post–War United States", *Manchester School*, 51, 317–340.

LONG, J.B. and C.I. PLOSSER (1983), "Real Business Cycles", *Journal of Political Economy*, 91, 39–69.

MARCET, A. (1989), "Solving Non–linear Stochastic Models by Parameterizing Expectations" (mimeo, Carnegie–Mellon University).

MCGRATTAN, E.B. (1991), "The Macroeconomic Effects of Distortionary Taxation", *Discussion Paper No 37*, Institute for Empirical Macroeconomics.

MONFORT, A. and R. RABEMANANJARA (1990), "From a VAR to a Structural Model, with an Application to the Wage Price Spiral", *Journal of Applied Econometrics*, 5, 203–227.

NEUSSER, K. (1991), "Testing the Long–Run Implications of the Neo–Classical Growth Model", *Journal of Monetary Economics*, 27, 3–37.

PRESCOTT, E.C. (1991), "Real Business Cycle Theory: What Have We Learned?" (unpublished lecture, Latin American Meeting of the Econometric Society, Punta del Este, Uruguay).

ROUWENHORST, K.G. (1991), "Time to Build and Aggregate Fluctuations: A Reconsideration", *Journal of Monetary Economics*, 27, 241–254.

SARGENT, T.J. (1978), "Estimation of Dynamic Labour Demand Schedules Under Rational Expectations", *Journal of Applied Econometrics*, 4, 213–237.

SINGLETON, K.J. (1988), "Economic Issues in the Analysis of Equilibrium Business Cycle Models", *Journal of Monetary Economics*, 21, 361–386.

SMITH, A.A. (1990), "Econometric Evaluation of a Real Business Cycle Model Using Simulation Methods" (Ch. 3, unpublished doctoral dissertation, Duke University).

SPANOS, A. (1986), *Statistical Foundations of Econometric Modelling* (Cambridge University Press, Cambridge).

TINBERGEN, J. (1939), *Statistical Testing of Business Cycle Theories, Vol I: A Method and Its Application to Investment Activity* (League of Nations: Geneva).

WALLIS, K.F. (1977), "Multiple Time Series Analysis and the Final Form of Econometric Models", *Econometrica*, 45, 1481–1497.

WATSON, M.W. (1990), "Measures of Fit for Calibrated Models" (mimeo, Northwestern University).

ZELLNER, A. and F. PALM (1974), "Time Series Analysis and Simultaneous Equation Econometric Models", *Journal of Econometrics*, 2, 17–54.

Fig 1: Plot of Second "Co-integrating" Error Among States

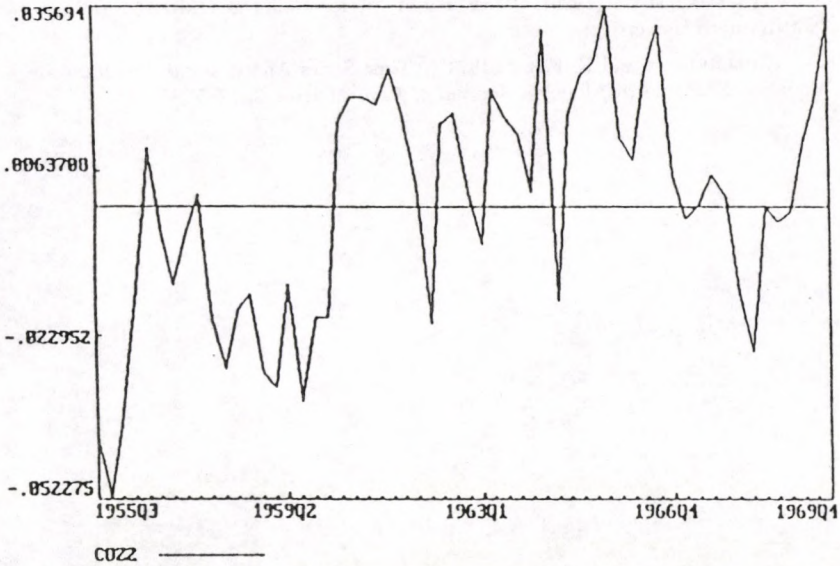


Fig 2 Consumption and Predictions of It from MPA and RBC Models

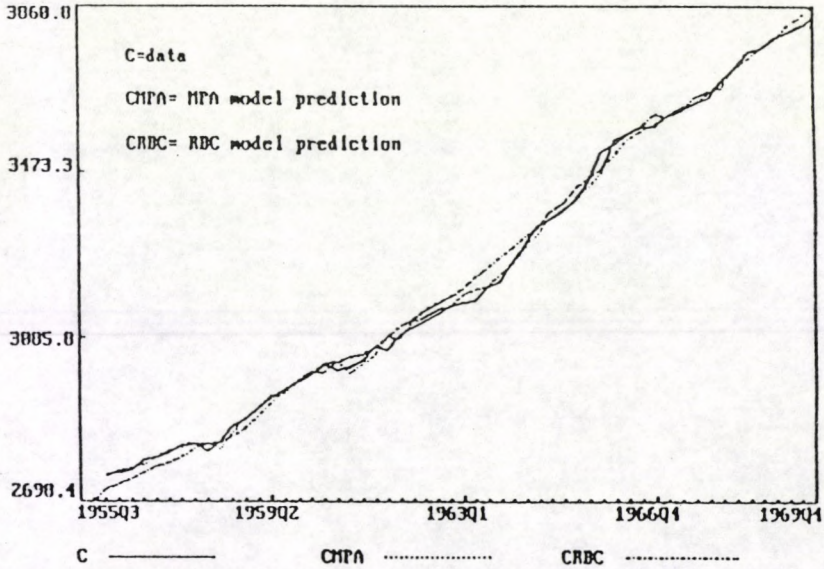


Fig 3: Investment and Predictions of It from MPA and RBC Models

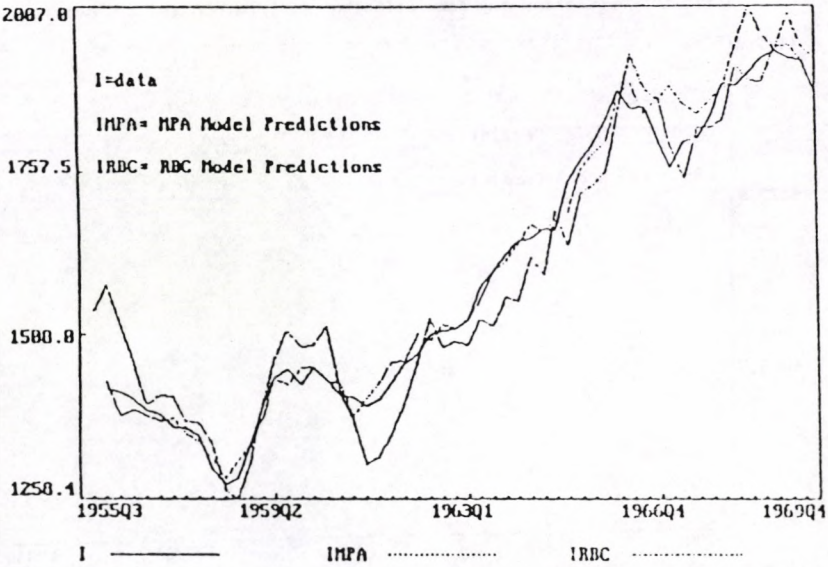


Fig 4: Output and Predictions of It from MPA and RBC Models

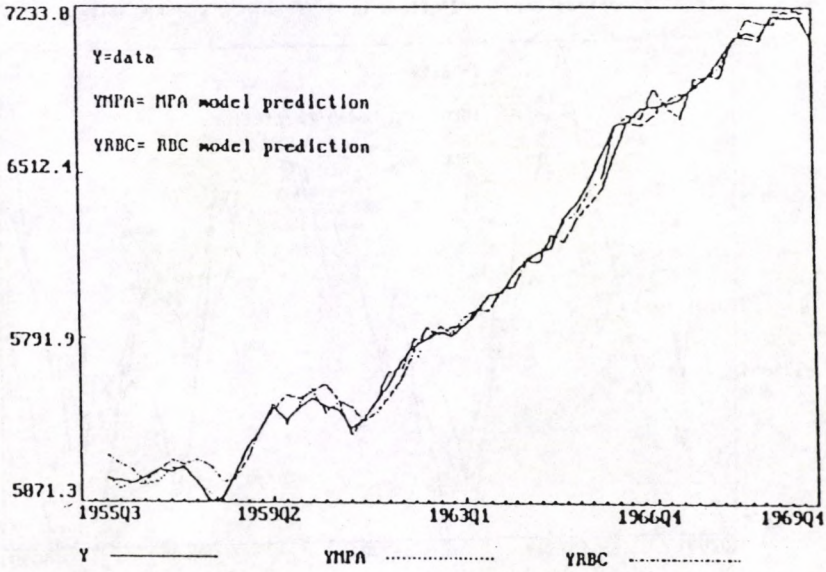


Fig 5: Consumption and Predictions of It from MPA and RBC Models

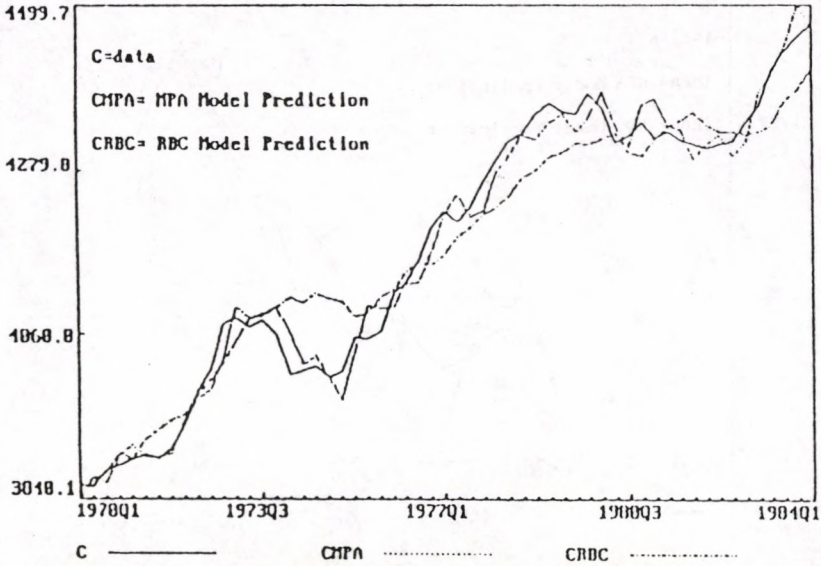


Fig 6: Investment and Predictions of It from MPA and RBC Models

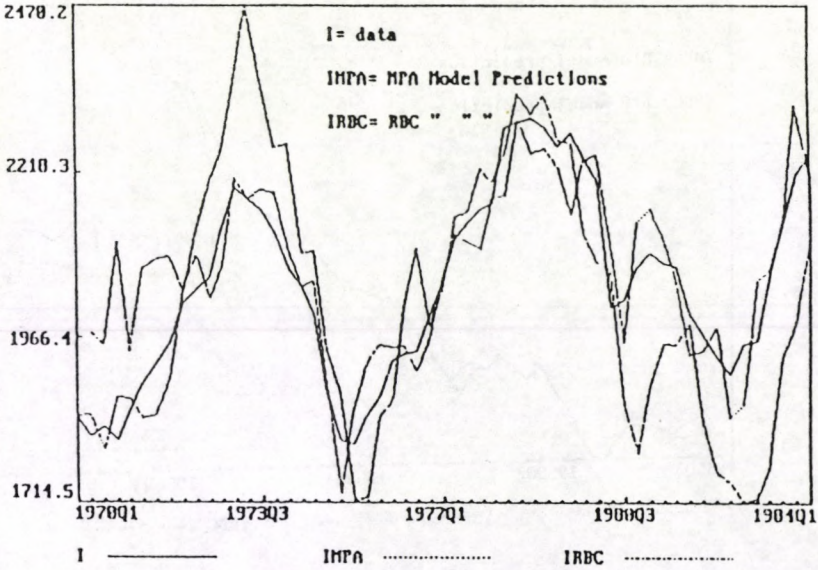


Fig 7: Output and Predictions of it from MPA and RBC Models

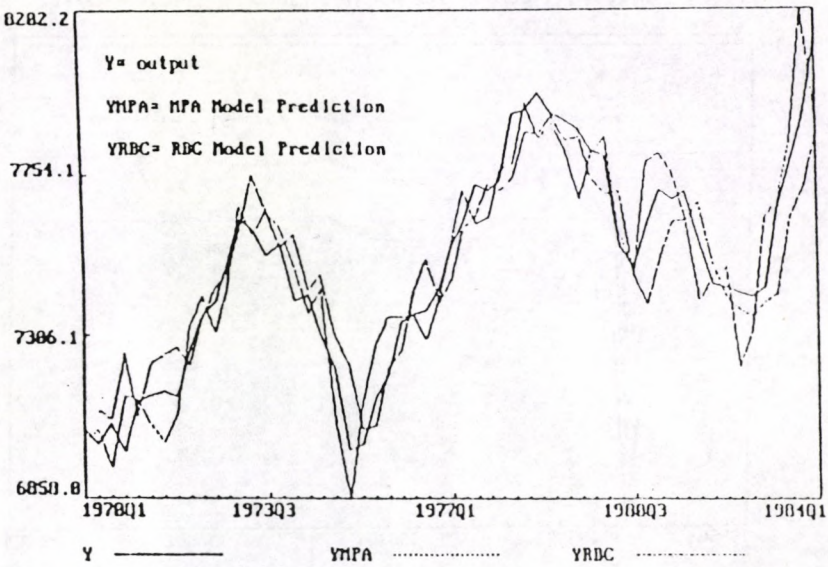


Fig 8: Recursive Estimate of AR(1) Parameter for CORC

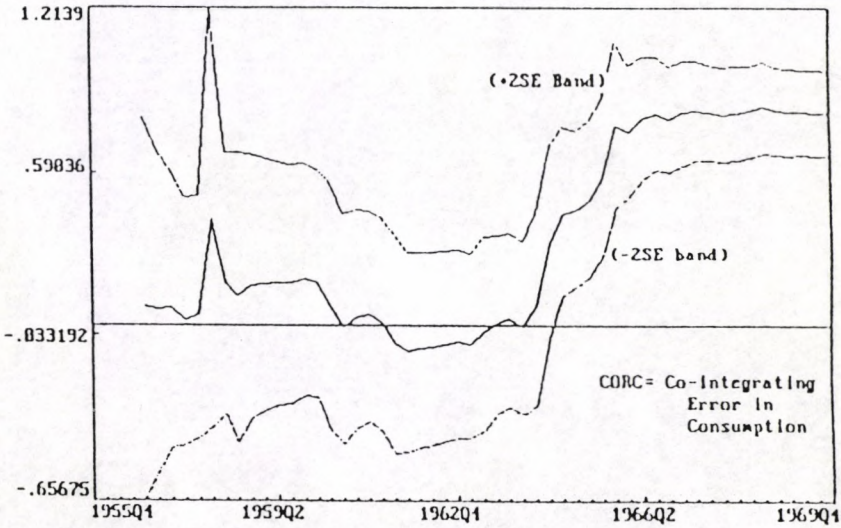
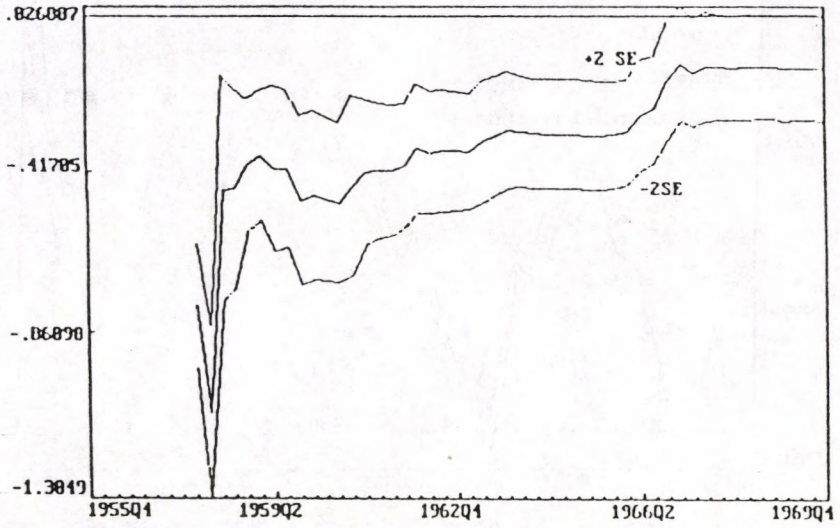


Fig 9 Recursive Estimate of the ECM Test Parameter for Investment



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

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

From Name . . . . . .
Address. . . . . .
. . . . . .
. . . . . .
. . . . . .
. . . . . .

❒ 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 1993/94
❒ Please send me the EUI Research Report

Please send me the following EUI 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 1990

ECO No. 90/1
Tamer BASAR and Mark SALMON
Credibility and the Value of Information Transmission in a Model of Monetary Policy and Inflation

ECO No. 90/2
Horst UNGERER
The EMS – The First Ten Years
Policies – Developments – Evolution

ECO No. 90/3
Peter J. HAMMOND
Interpersonal Comparisons of Utility: Why and how they are and should be made

ECO No. 90/4
Peter J. HAMMOND
A Revelation Principle for (Boundedly) Bayesian Rationalizable Strategies

ECO No. 90/5
Peter J. HAMMOND
Independence of Irrelevant Interpersonal Comparisons

ECO No. 90/6
Hal R. VARIAN
A Solution to the Problem of Externalities and Public Goods when Agents are Well-Informed

ECO No. 90/7
Hal R. VARIAN
Sequential Provision of Public Goods

ECO No. 90/8
T. BRIANZA, L. PHLIPS and J.F. RICHARD
Futures Markets, Speculation and Monopoly Pricing

ECO No. 90/9
Anthony B. ATKINSON/ John MICKLEWRIGHT
Unemployment Compensation and Labour Market Transition: A Critical Review

ECO No. 90/10
Peter J. HAMMOND
The Role of Information in Economics

ECO No. 90/11
Nicos M. CHRISTODOULAKIS
Debt Dynamics in a Small Open Economy

ECO No. 90/12
Stephen C. SMITH
On the Economic Rationale for Codetermination Law

ECO No. 90/13
Elettra AGLIARDI
Learning by Doing and Market Structures

ECO No. 90/14
Peter J. HAMMOND
Intertemporal Objectives

ECO No. 90/15
Andrew EVANS/Stephen MARTIN
Socially Acceptable Distortion of Competition: EC Policy on State Aid

ECO No. 90/16
Stephen MARTIN
Fringe Size and Cartel Stability

ECO No. 90/17
John MICKLEWRIGHT
Why Do Less Than a Quarter of the Unemployed in Britain Receive Unemployment Insurance?

ECO No. 90/18
Mrudula A. PATEL
Optimal Life Cycle Saving With Borrowing Constraints:
A Graphical Solution

ECO No. 90/19
Peter J. HAMMOND
Money Metric Measures of Individual and Social Welfare Allowing for Environmental Externalities

ECO No. 90/20
Louis PHLIPS/
Ronald M. HARSTAD
Oligopolistic Manipulation of Spot Markets and the Timing of Futures Market Speculation

* Working Paper out of print

ECO No. 90/21
Christian DUSTMANN
Earnings Adjustment of Temporary Migrants

ECO No. 90/22
John MICKLEWRIGHT
The Reform of Unemployment Compensation:
Choices for East and West

ECO No. 90/23
Joerg MAYER
U. S. Dollar and Deutschmark as Reserve Assets

ECO No. 90/24
Sheila MARNIE
Labour Market Reform in the USSR: Fact or Fiction?

ECO No. 90/25
Peter JENSEN/
Niels WESTERGÅRD-NIELSEN
Temporary Layoffs and the Duration of Unemployment: An Empirical Analysis

ECO No. 90/26
Stephan L. KALB
Market-Led Approaches to European Monetary Union in the Light of a Legal Restrictions Theory of Money

ECO No. 90/27
Robert J. WALDMANN
Implausible Results or Implausible Data? Anomalies in the Construction of Value Added Data and Implications for Estimates of Price-Cost Markups

ECO No. 90/28
Stephen MARTIN
Periodic Model Changes in Oligopoly

ECO No. 90/29
Nicos CHRISTODOULAKIS/
Martin WEALE
Imperfect Competition in an Open Economy

***

ECO No. 91/30
Steve ALPERN/Dennis J. SNOWER
Unemployment Through 'Learning From Experience'

ECO No. 91/31
David M. PRESCOTT/Thanasis STENGOS
Testing for Forecastible Nonlinear Dependence in Weekly Gold Rates of Return

ECO No. 91/32
Peter J. HAMMOND
Harsanyi's Utilitarian Theorem:
A Simpler Proof and Some Ethical Connotations

ECO No. 91/33
Anthony B. ATKINSON/
John MICKLEWRIGHT
Economic Transformation in Eastern Europe and the Distribution of Income*

ECO No. 91/34
Svend ALBAEK
On Nash and Stackelberg Equilibria when Costs are Private Information

ECO No. 91/35
Stephen MARTIN
Private and Social Incentives to Form R & D Joint Ventures

ECO No. 91/36
Louis PHLIPS
Manipulation of Crude Oil Futures

ECO No. 91/37
Xavier CALSAMIGLIA/Alan KIRMAN
A Unique Informationally Efficient and Decentralized Mechanism With Fair Outcomes

ECO No. 91/38
George S. ALOGOSKOUFIS/
Thanasis STENGOS
Testing for Nonlinear Dynamics in Historical Unemployment Series

ECO No. 91/39
Peter J. HAMMOND
The Moral Status of Profits and Other Rewards:
A Perspective From Modern Welfare Economics

* Working Paper out of print

ECO No. 91/40
Vincent BROUSSEAU/Alan KIRMAN
The Dynamics of Learning in Mis-Specified Models

ECO No. 91/41
Robert James WALDMANN
Assessing the Relative Sizes of Industry- and Nation Specific Shocks to Output

ECO No. 91/42
Thorsten HENS/Alan KIRMAN/Louis PHLIPS
Exchange Rates and Oligopoly

ECO No. 91/43
Peter J. HAMMOND
Consequentialist Decision Theory and Utilitarian Ethics

ECO No. 91/44
Stephen MARTIN
Endogenous Firm Efficiency in a Cournot Principal-Agent Model

ECO No. 91/45
Svend ALBAEK
Upstream or Downstream Information Sharing?

ECO No. 91/46
Thomas H. McCURDY/
Thanasis STENGOS
A Comparison of Risk-Premium Forecasts Implied by Parametric Versus Nonparametric Conditional Mean Estimators

ECO No. 91/47
Christian DUSTMANN
Temporary Migration and the Investment into Human Capital

ECO No. 91/48
Jean-Daniel GUIGOU
Should Bankruptcy Proceedings be Initiated by a Mixed Creditor/Shareholder?

ECO No. 91/49
Nick VRIEND
Market-Making and Decentralized Trade

ECO No. 91/50
Jeffrey L. COLES/Peter J. HAMMOND
Walrasian Equilibrium without Survival: Existence, Efficiency, and Remedial Policy

ECO No. 91/51
Frank CRITCHLEY/Paul MARRIOTT/
Mark SALMON
Preferred Point Geometry and Statistical Manifolds

ECO No. 91/52
Costanza TORRICELLI
The Influence of Futures on Spot Price Volatility in a Model for a Storable Commodity

ECO No. 91/53
Frank CRITCHLEY/Paul MARRIOTT/
Mark SALMON
Preferred Point Geometry and the Local Differential Geometry of the Kullback-Leibler Divergence

ECO No. 91/54
Peter MØLLGAARD/
Louis PHLIPS
Oil Futures and Strategic Stocks at Sea

ECO No. 91/55
Christian DUSTMANN/
John MICKLEWRIGHT
Benefits, Incentives and Uncertainty

ECO No. 91/56
John MICKLEWRIGHT/
Gianna GIANNELLI
Why do Women Married to Unemployed Men have Low Participation Rates?

ECO No. 91/57
John MICKLEWRIGHT
Income Support for the Unemployed in Hungary

ECO No. 91/58
Fabio CANOVA
Detrending and Business Cycle Facts

ECO No. 91/59
Fabio CANOVA/
Jane MARRINAN
Reconciling the Term Structure of Interest Rates with the Consumption Based ICAP Model

ECO No. 91/60
John FINGLETON
Inventory Holdings by a Monopolist Middleman

* Working Paper out of print

***

ECO No. 92/61
Sara CONNOLLY/John MICKLEWRIGHT/Stephen NICKELL
The Occupational Success of Young Men Who Left School at Sixteen

ECO No. 92/62
Pier Luigi SACCO
Noise Traders Permanence in Stock Markets: A Tâtonnement Approach. I: Informational Dynamics for the Two-Dimensional Case

ECO No. 92/63
Robert J. WALDMANN
Asymmetric Oligopolies

ECO No. 92/64
Robert J. WALDMANN /Stephen C. SMITH
A Partial Solution to the Financial Risk and Perverse Response Problems of Labour-Managed Firms: Industry-Average Performance Bonds

ECO No. 92/65
Agustín MARAVALL/Víctor GÓMEZ
Signal Extraction in ARIMA Time Series Program SEATS

ECO No. 92/66
Luigi BRIGHI
A Note on the Demand Theory of the Weak Axioms

ECO No. 92/67
Nikolaos GEORGANTZIS
The Effect of Mergers on Potential Competition under Economies or Diseconomies of Joint Production

ECO No. 92/68
Robert J. WALDMANN/ J. Bradford DE LONG
Interpreting Procyclical Productivity: Evidence from a Cross-Nation Cross-Industry Panel

ECO No. 92/69
Christian DUSTMANN/John MICKLEWRIGHT
Means-Tested Unemployment Benefit and Family Labour Supply: A Dynamic Analysis

ECO No. 92/70
Fabio CANOVA/Bruce E. HANSEN
Are Seasonal Patterns Constant Over Time? A Test for Seasonal Stability

ECO No. 92/71
Alessandra PELLONI
Long-Run Consequences of Finite Exchange Rate Bubbles

ECO No. 92/72
Jane MARRINAN
The Effects of Government Spending on Saving and Investment in an Open Economy

ECO No. 92/73
Fabio CANOVA and Jane MARRINAN
Profits, Risk and Uncertainty in Foreign Exchange Markets

ECO No. 92/74
Louis PHLIPS
Basing Point Pricing, Competition and Market Integration

ECO No. 92/75
Stephen MARTIN
Economic Efficiency and Concentration: Are Mergers a Fitting Response?

ECO No. 92/76
Luisa ZANCHI
The Inter-Industry Wage Structure: Empirical Evidence for Germany and a Comparison With the U.S. and Sweden

ECO NO. 92/77
Agustín MARAVALL
Stochastic Linear Trends: Models and Estimators

ECO No. 92/78
Fabio CANOVA
Three Tests for the Existence of Cycles in Time Series

ECO No. 92/79
Peter J. HAMMOND/Jaime SEMPERE
Limits to the Potential Gains from Market Integration and Other Supply-Side Policies

* Working Paper out of print

ECO No. 92/80
Víctor GÓMEZ and Agustín MARAVALL
Estimation, Prediction and Interpolation for Nonstationary Series with the Kalman Filter

ECO No. 92/81
Víctor GÓMEZ and Agustín MARAVALL
Time Series Regression with ARIMA Noise and Missing Observations Program TRAM

ECO No. 92/82
J. Bradford DE LONG/ Marco BECHT
"Excess Volatility" and the German Stock Market, 1876-1990

ECO No. 92/83
Alan KIRMAN/Louis PHLIPS
Exchange Rate Pass-Through and Market Structure

ECO No. 92/84
Christian DUSTMANN
Migration, Savings and Uncertainty

ECO No. 92/85
J. Bradford DE LONG
Productivity Growth and Machinery Investment: A Long-Run Look, 1870-1980

ECO NO. 92/86
Robert B. BARSKY and J. Bradford DE LONG
Why Does the Stock Market Fluctuate?

ECO No. 92/87
Anthony B. ATKINSON/John MICKLEWRIGHT
The Distribution of Income in Eastern Europe

ECO No.92/88
Agustín MARAVALL/Alexandre MATHIS
Encompassing Unvariate Models in Multivariate Time Series: A Case Study

ECO No. 92/89
Peter J. HAMMOND
Aspects of Rationalizable Behaviour

ECO 92/90
Alan P. KIRMAN/Robert J. WALDMANN
I Quit

ECO No. 92/91
Tilman EHRBECK
Rejecting Rational Expectations in Panel Data: Some New Evidence

ECO No. 92/92
Djordje Suvakovic OLGIN
Simulating Codetermination in a Cooperative Economy

ECO No. 92/93
Djordje Suvakovic OLGIN
On Rational Wage Maximisers

ECO No. 92/94
Christian DUSTMANN
Do We Stay or Not? Return Intentions of Temporary Migrants

ECO No. 92/95
Djordje Suvakovic OLGIN
A Case for a Well-Defined Negative Marxian Exploitation

ECO No. 92/96
Sarah J. JARVIS/John MICKLEWRIGHT
The Targeting of Family Allowance in Hungary

ECO No. 92/97
Agustín MARAVALL/Daniel PEÑA
Missing Observations and Additive Outliers in Time Series Models

ECO No. 92/98
Marco BECHT
Theory and Estimation of Individual and Social Welfare Measures: A Critical Survey

ECO No. 92/99
Louis PHLIPS and Ireneo Miguel MORAS
The AKZO Decision: A Case of Predatory Pricing?

ECO No. 92/100
Stephen MARTIN
Oligopoly Limit Pricing With Firm-Specific Cost Uncertainty

* Working Paper out of print

ECO No. 92/101
Fabio CANOVA/Eric GHYSELS
Changes in Seasonal Patterns: Are They Cyclical?

ECO No. 92/102
Fabio CANOVA
Price Smoothing Policies: A Welfare Analysis

***

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

* Working Paper out of print

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

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

* Working Paper out of print